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Session I
Stock markets

Overview: market structure issues in market liquidity

Maureen O'Hara[1]

The behaviour of prices and even the viability of markets depend on the ability of the trading mechanism to match the trading desires of sellers and buyers. This matching process involves the provision of market liquidity. The role of the market maker in providing liquidity is widely recognised, but liquidity can also arise from other aspects of the trading mechanism. In particular, rules and market practices governing the trading process, such as how trading orders are submitted and what trading information must be disclosed, can affect the creation of liquidity. This raises the question of whether changes in market structure can enhance the provision of liquidity. Is there a "Golconda exchange" that provides optimal liquidity?

What is microstructure?

Issues related to market liquidity are part of a broader analysis of the microstructure of markets. Market microstructure refers to the study of the process and outcomes of exchanging assets under a specific set of rules. While much of economics abstracts from the mechanics of trading, microstructure theory focuses on how specific trading mechanisms affect the price formation process.[2]

Much of the microstructure literature has focused on the *price-setting problem* confronting market intermediaries. The Walrasian auctioneer provides the simplest (and oldest) characterisation of the price-setting process. The auctioneer announces a potential trading range, and traders determine their optimal order at that price. If there are imbalances in traders' demands and supplies, a new potential price is suggested, and traders then revise any orders. No trading takes place until a market-clearing price is found. The London gold fixing loosely resembles the Walrasian framework, but most other markets differ dramatically. In particular, specific market participants play roles far removed from the passive one of the auctioneer. Demsetz (1968) was one of the first economists to analyse how the behaviour of traders affects the formation of prices. Demsetz argued that while a trader willing to wait might trade at the single price envisioned in the Walrasian framework, a trader not wanting to wait could pay a price for immediacy, ie liquidity. This results in two equilibrium prices. Moreover, since the size of the price concession needed to trade immediately depends on the number of traders, the structure of the market could affect the cost of immediacy and thus the market-clearing price.

The price-setting problem examined by Demsetz has been investigated more formally using inventory-based models. These models view the trading process as a matching problem in which the market maker - or price-setting agent - must use prices to balance supply and demand across time. There are several distinct approaches to modelling how prices are set by market makers: Garman (1976) focused on the nature of order flow; Stoll (1978) and Ho and Stoll (1981) examined the optimisation problem facing dealers; and Cohen, Maier, Schwartz and Whitcomb (1981) analysed the effects of multiple providers of immediacy. Common to each of these approaches are uncertainties in order flow, which can result in inventory problems for the market maker and execution problems for traders.

An alternative approach to modelling the behaviour of prices focuses on the *learning problem* confronting market intermediaries. Starting with Kyle (1984, 1985), Glosten and Milgrom (1985) and Easley and O'Hara (1987), market structure research has given greater attention to the effect of asymmetric information on market prices. If some traders have superior information about the underlying value of an asset, their trades could reveal what this underlying value is and so affect the behaviour of prices.

The key to extracting information from order flows is Bayesian learning. Each trader has a prior belief about the true value V of an asset. Traders observe some data, say a trade, and then calculate the probability that V equals their prior belief given that these data have been observed. This conditional

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1 Johnson Graduate School of Management, Cornell University and President-elect of the American Finance Association. Special thanks to Philip Wooldridge at the Bank for International Settlements for transcribing this presentation.

2 For a survey of the literature, see O'Hara (1995) or Madhavan (2000). Lyons (forthcoming) provides a comprehensive review of the microstructure of foreign exchange markets.

probability incorporates the new information that traders learned from observing the data, and is hence their posterior belief about V (Graph 1). The posterior then becomes the new prior, more data are observed, and the updating process continues.

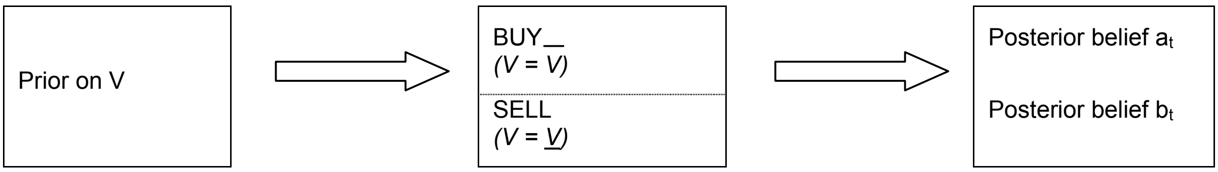
Graph 1

Bayesian learning

In information-based models, the solution to this learning problem determines the prices set by market makers. The ask price a_t equals the expected value of V given that a trader wishes to buy, and depends on the conditional probability that V is either lower ($V = \underline{V}$) or higher ($V = \overline{V}$) than the market maker's prior belief given that a trader wishes to buy. The bid price b_t is defined similarly given that a trader wishes to sell. An important characteristic of these prices is that they explicitly depend on the probability of a sale or buy (Graph 2). If uninformed traders are assumed equally likely to buy or sell whatever the information, good news ($V = \overline{V}$) will result in an excess of buy orders as informed traders decide to buy. Likewise, bad news ($V = \underline{V}$) will result in an excess of sell orders as informed traders decide to sell.

Graph 2

Dealer pricing



What we have learned

The information-based approach has greatly enhanced our understanding of the behaviour of markets and by extension the nature of market liquidity. Perhaps the greatest insight of this approach is how information affects quotes and spreads. Information-based models highlight the role of market parameters such as the size of the market or the ratio of large to small trades in the adjustment of prices. This in turn provides an explanation for the existence of bid-ask spreads even in competitive markets, without reference to explicit transactions or inventory costs. Inventory-based explanations of the bid-ask spread are problematic because empirical evidence of inventory effects in financial markets is weak.

Another important conclusion is that prices ultimately converge to their true, full-information value; in the limit markets are strong-form efficient.[3] This follows from the Bayesian learning process. It is not entirely clear, however, what market efficiency means in a dynamic setting. Given that some traders have superior information, prices along the adjustment path do not exhibit strong-form efficiency, and indeed there can be very great differences in the speed with which prices move toward full-information levels. Markets with greater volume, for example, adjust faster (in clock time) to information. The time between trades, in particular the tendency for transactions to cluster, also appears to affect the adjustment of prices.

The time varying process by which transactions arrive has important implications for econometric modelling of market volatility. Generalised Autoregressive Conditional Heteroscedasticity (GARCH) models and Autoregressive Conditional Duration (ACD) models have come to be widely used for analysing price and transactions data, respectively.

3 Following the categorisations of the efficient market hypothesis used by Fama (1970), weak-form efficiency assumes that security prices fully reflect all security-market information, semi-strong form efficiency assumes that security prices fully reflect all publicly available information, and strong-form efficiency assumes that security prices fully reflect all information from public and private sources.

Finally, much has been learned about the information contained in specific trades. Different types of trades seem to have different information content. Similarly, trades in different markets seem to have different information content.

What we still do not know

For all that we have learned, there remain several puzzling issues concerning the trading process. Foremost is what determines volume. While empirical research has identified a strong link between volume and price movements, it is not obvious why this should be so. Volume may simply be a consequence of the trading process; whereas individual trades cause prices to change, volume per se may not affect prices. Or as seems more likely, volume could reveal underlying information, and thus be a component in the learning process. Pfleiderer (1984), Campbell et al (1991), Harris and Raviv (1993), Blume et al (1994), and Wang (1994) have examined this informational role.

A second set of issues revolves around what the uninformed traders are doing. It is the uninformed traders who provide the liquidity to the informed, and so understanding their behaviour can provide substantial insight and intuition into the trading process. Information-based microstructure models typically assume that uninformed traders do not act strategically. Yet, if it is profitable for informed traders to time their trades, then it must be profitable for uninformed traders to do so as well. Admati and Pfleiderer (1988, 1989), Foster and Viswanathan (1990), Seppi (1990) and Spiegel and Subrahmanyam (1992) among others have applied a game-theoretic approach to modelling the decisions of uninformed traders. A common outcome with this approach, however, is the occurrence of multiple equilibria.

Another open question is what traders can learn from other pieces of market data, such as prices. Neither sequential trade models such as Glosten and Milgrom (1985) nor batch trading models such as Kyle (1985) allow traders to learn anything from the movement of prices that is not already in their information set. But in actual asset markets the price elasticity of prices appears to be important. Technical analysis of market data is widespread in markets, with elaborate trading strategies devised to respond to the pattern of prices.

Finally, microstructure theory has not yet convincingly addressed how the existence of more than one liquidity provider in more than one market setting affects the price adjustment process. Much of the literature assumes the existence of a single market-clearing agent. However, alternative mechanisms could arise that divert order flow away from the specialist. Multi-market linkages introduce complex and often conflicting effects on market liquidity and trading behaviour. Indeed, it is not even obvious whether a segmented market equilibrium is sustainable. Current models of liquidity, for example, suggest that securities markets may have an inherent disposition toward being natural monopolies. Further research in this area is particularly important given the rapid increase in the number of electronic exchanges in recent years.

Market structures

Markets are currently structured in a myriad of ways, and new market-clearing mechanisms are arising with surprising frequency. All trading in a particular security can be directed to a single specialist, who is expected to make a market in that security. The New York Stock Exchange (NYSE) is the best known example of such a market structure (Table 1). Alternatively, dealers can compete for trades, buying and selling securities for their own account. Traditionally dealers competed in a central location, such as the London Stock Exchange or NASDAQ, but competition need not be centralised. Bonds, for example, trade primarily through bilateral negotiations between dealers and customers. A still third trading mechanism is the automatic matching of orders through an electronic broker. Today the majority of trading in the global foreign exchange market takes place over electronic exchanges such as Reuters and Electronic Broking System (EBS).

Table 1

Market structures

	Specialist	Dealer	Electronic
Equity	New York Stock Exchange	NASDAQ London Stock Exchange	Stock Exchange of Hong Kong Instinet Paris Bourse
Bond		Bond dealers	Tradenet EUREX
Foreign exchange		FX brokers	Reuters EBS

Actual markets do not conform to simple structures. Indeed, they typically involve more than one structure. What is important, therefore, is not the operation of any specific trading mechanism, but rather the rules by which trades occur. These rules dictate what can be traded, who can trade, when and how orders can be submitted, who may see or handle the order, and how orders are processed. The rules determine how market structures work, and thus how prices are formed.

Since rules can affect the behaviour of prices, liquidity might also naturally depend on how a market is structured. Indeed, liquidity concerns may dictate the structure of the market. Drawing on the extensive body of research investigating the interaction between market structure and liquidity, the remainder of this paper focuses on two critical issues in the creation of liquidity: the impact of limit orders, and the effects of transparency.

Limit orders

A wide variety of order types are found in securities markets. The most familiar type is a market order to buy or sell one round lot at the prevailing price. Other orders, such as "market-at-close", "fill-or-kill" and "immediate-or-cancel" allow traders to control the timing, quantity or execution of their trades. By far the most common alternative type of order is a limit order specifying a price and a quantity at which a trade is to transact. Limit orders specify a price either above the current ask or below the current bid and await the movement of prices to become active. If the market is rising, the upward price movement triggers limit orders to sell; if the market is falling, the downward movement triggers limit orders to buy. Limit orders thus provide liquidity to the market.

Limit order traders receive a better price than they would have if they had submitted a market order, but face the risk of non-execution and a winner's curse problem. Whereas a market order executes with certainty, limit orders await the movement of prices to become active, ie a limit order is held in a "book" until either a matching order is entered or the order is cancelled. Moreover, because once posted their prices do not respond to the arrival of new information, limit orders are more likely to be executed when they are mispriced. Foucault (1999) finds that in deciding whether to submit a market order or post a limit order, traders' main consideration is the volatility of an asset. In a volatile market, the probability of mispricing an asset is higher, and so limit order traders quote relatively wide bid-ask spreads. This raises the cost of market order trading, thereby increasing the incentive to use limit orders rather than market orders. But as a result of fewer market orders, the execution risk associated with limit orders increases.

Order size may also influence investors' choice between market and limit orders. Seppi (1997) concludes that small retail and large institutional investors prefer hybrid markets such as the NYSE, where specialists compete with limit orders to execute market orders.[4] Mid-size investors, on the other hand, might prefer pure limit order markets such as electronic exchanges. According to Seppi, specialists will undercut limit order prices at the margin. Such undercutting lowers the probability that limit orders will execute, thus resulting in reduced depth in the book. Evidence in Sofianos (1995) of a

4 In hybrid markets, the ability of limit orders to compete with market makers depends on priority rules. Limit orders to sell at prices at or below the price at which the specialist proposes to sell, or limit orders to buy at or above the specialist's bid price, typically have priority for execution.

U-shaped relationship between specialists' total revenue and trade size suggests that specialists do indeed provide relatively more liquidity to small and large trades.

The composition of order flows is a dynamic process, with investors' preferred order type changing in response to developments over time. Goldstein and Kavajecz (2000) examine the behaviour of liquidity providers on the NYSE during periods of extreme volatility. They find that following a precipitous drop in equity prices, traders abandoned limit orders in favour of floor brokers. In particular, whereas specialists maintained narrow spreads and normal depth, liquidity drained out of the limit order book. Similarly, in foreign exchange markets, trading tends to move from electronic order-matching markets to dealer markets during periods of market stress. Such dynamics raise the question of whether dealer markets handle information more efficiently than pure limit order markets.

Another issue relating to limit orders is whether they can provide enough liquidity for every type of trade. The experience of limit order markets suggests not. For example, on the NYSE, the Toronto Stock Exchange and other exchanges with features of limit order markets, a substantial proportion of block trades - trades of 10,000 shares or more - are submitted to block traders or "upstairs market makers", who form a syndicate of buyers to take the other side of the trade. One reason for using block traders rather than limit orders is that large transactions might be interpreted as signalling new information, and so move prices against the seller. Limit order systems are constantly evolving as new technologies are developed, and indeed OptiMark designed an electronic trading system that was supposed to minimise the impact that large orders had on price. OptiMark's system ensured that orders remained anonymous until executed in full and was initially lauded as presaging the transformation of institutional trading. Despite the system's advantages, however, it was poorly received by brokers and OptiMark ran into financial difficulties in mid-2000.

Finally, there is the question of how much information about the limit order book is optimal. On the NYSE and a number of other exchanges, orders held in the specialist's book are not common knowledge, although the specialist may choose to allow traders to view the book. By contrast, on electronic exchanges the order limit book is usually transparent. Madhavan and Panchapagesan (2000) find that on the NYSE the ability to observe the evolution of the book conveys valuable information to the specialist. In particular, specialists use information from the order book to set a more efficient opening price than the price that would prevail if all orders - both market and limit orders - were considered. Coppejans and Domowitz (1999) examine a pure limit order market and conclude that the trading process is influenced only by the flow of orders, not the stock of orders on the book. The book is not irrelevant; flows, after all, are changes in stocks. But in a market with an open book, the book per se does not appear to contain information on the value of the asset being traded. While helping us to understand how price formation occurs in actual markets, the results of these empirical studies do not imply that one particular market structure provides for more efficient price discovery than another. The experimental methods discussed below offer more meaningful insight into such hypothetical questions.

Transparency

As the information-based microstructure models demonstrated, the information available in the trading process can affect the trading strategies of market participants. It thus follows that the market equilibrium depends on the degree of transparency, ie the ability of market participants to observe the information in the trading process. Consider the previous discussion of the limit order book. If the book were known only to the market maker (as on the NYSE), then the market maker, as well as the informed and uninformed traders, would behave differently than if the book were common knowledge (as in the market examined by Coppejans and Domowitz).

The openness of the book is but one of many differences in the degree of transparency across markets. The breadth of trade data reported and even the timeliness of the reported data can also differ tremendously. Some markets such as bond dealers provide only pre-trade information, meaning that quote data are made available but not transactions data. Other markets require post-trade transparency, ensuring that the price and quantity of trades are observable. The NYSE and NASDAQ, for example, are required to report immediately all quotes and trades. At the other extreme, trades handled "off board" - trades executed outside of the United States after US markets close - need not even be acknowledged.

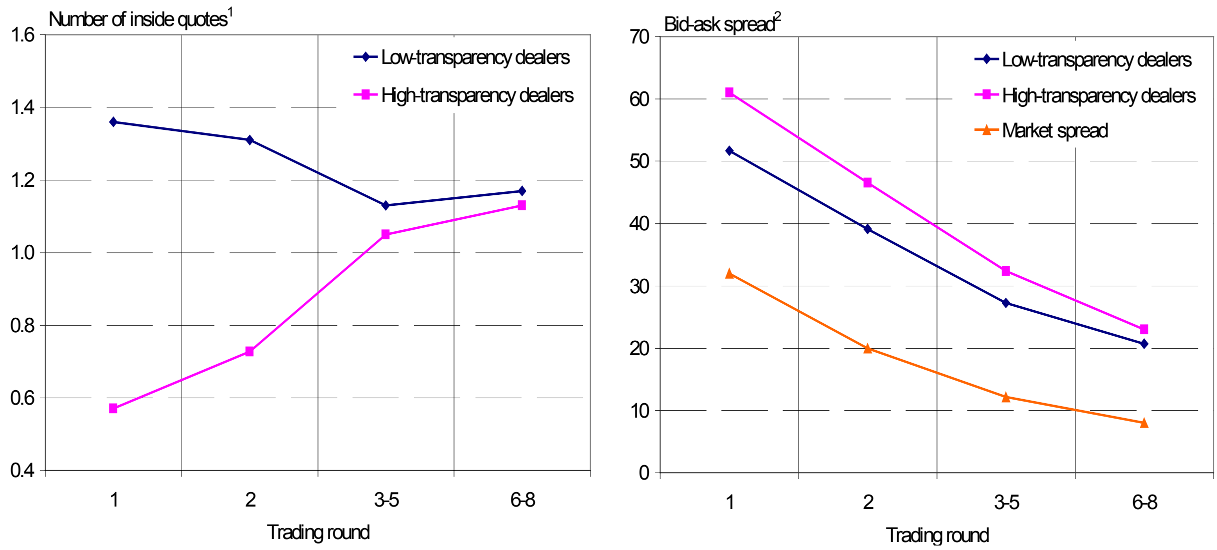
Differences in transparency may play a significant role in the creation of liquidity. As a factor in traders' strategic decisions, transparency can influence their willingness to participate in the trading process. In the United Kingdom, for example, the Financial Services Authority allows the reporting of large trades to be delayed for a period of time because it believes that immediate disclosure would expose market makers to undue risk as they unwound their positions and so discourage them from providing liquidity. Transparency is also a crucial consideration in the competition among markets for trading volume, and thus in the prospects for further fragmentation of liquidity.

Bloomfield and O'Hara (1999, 2000) use laboratory experiments to address some of these issues. Their experiments include multiple dealers operating under varying degrees of transparency and traders with differing trade motivations. A key finding is that low-transparency dealers are more likely to set the highest bid and the lowest ask (inside quotes) in early rounds of trading, in order to capture more order flow (Graph 3). The information learned from the order flow allows low-transparency dealers to quote narrower spreads than their more transparent competitors and to avoid money-losing trades. This informational advantage declines with repeated rounds of trading because low-transparency dealers reveal their information through their choices of quotes. Moreover, as trade progresses and individual dealers learn from trade outcomes, spreads for all dealers decline (Graph 3).

Trading gains follow a pattern similar to spreads. Wide spreads in early rounds result in large gains because traders in need of liquidity are forced to buy at high prices and to sell at low prices (Graph 4). Gains then decline in concert with the decline in spreads. Notably, neither high- nor low-transparency dealers earn money at outside quotes, ie bids that are lower than the highest bid and asks that are higher than the lowest ask. Even though trades at outside quotes are executed at more favourable prices, dealer profits are eliminated by the higher likelihood of transacting with an informed trader. At inside quotes, the proportion of total trades coming from informed traders is approximately 10%, but at outside quotes, the proportion rises to 70% (Graph 4). Liquidity traders' preference to transact at the best available quote results in this higher degree of adverse selection at outside quotes.

Interestingly, traders do not behave strategically in these experiments. Concern about the possible impact of a trade in one round on prices in future rounds might be expected to lead traders to pay a premium to conceal their trades by trading with low-transparency dealers. Recall that this concern was one of the motivations behind the design of the OptiMark trading system. As Graph 4 shows, however, informed traders are as equally likely to transact with low-transparency dealers as with high-transparency dealers. In this setting, transparency seems to have a greater impact on dealer behaviour than on trader behaviour.

Graph 3
Quote setting behaviour

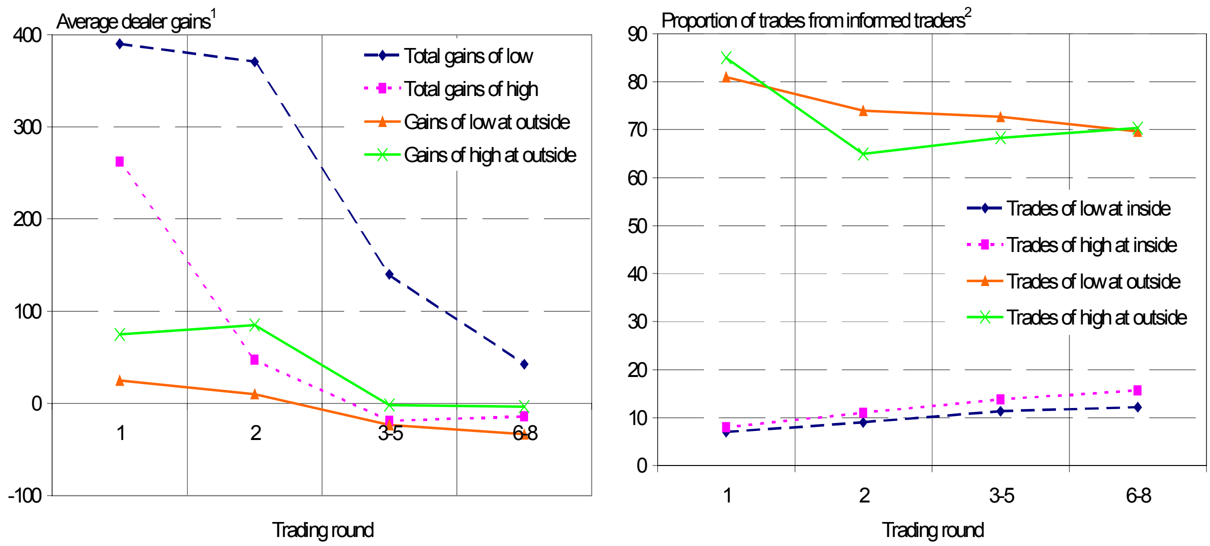


[1] Average number of inside quotes (the highest bid or the lowest ask) set by dealers. [2] Average bid-ask spread quoted for each security. The market spread is defined as the lowest ask minus the highest bid.

Source: Bloomfield and O'Hara (2000).

Graph 4

Dealer gains and adverse selection



[1] Average gains of low- and high-transparency dealers over all trades (total gains) and for trades at outside quotes (gains at outside). [2] Proportion of total trades coming from informed traders. Total trades are segregated by dealer type (low or high transparency) and quote form (inside or outside quote).

Source: Bloomfield and O'Hara (2000).

In conclusion, economic experiments provide disquieting evidence that transparent markets may be less liquid than markets with weaker reporting requirements. Transparency reduces the information content of specific trades and so reduces dealers' incentive to compete for orders. As a result, bid-ask spreads in transparent markets tend to be wider than those in less transparent markets. This accords with the experience in actual markets. Spreads on Instinet, for example, are frequently narrower than those on NASDAQ. The "Golconda exchange" may be less transparent than some of the markets that currently dominate global trading.

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Events that shook the market

Ray C Fair[1]

Abstract

Tick data on the S&P 500 futures contract and newswire searches are used to match events to large one-to five-minute stock price changes. Sixty-nine events that led to large stock price changes are identified between 1982 and 1999, 53 of which are directly or indirectly related to monetary policy. Many large stock price changes have no events associated with them.

1. Introduction

Although it is obvious that stock prices respond to events, it is not easy to match particular events to particular changes in stock prices. For example, Cutler, Poterba and Summers (1989) chose the 50 largest daily changes in the S&P 500 index from 1946 through 1987 and attempted to find an explanation of each change in the next day's *New York Times*. They found few cases in which it could be said with any confidence that a particular event led to the change. A problem with studies like this is that the daily interval may be too long, since many events can take place in a 24-hour period.

In this paper tick data on the S&P 500 futures contract and newswire searches are used to match events to stock price changes. The tick data are used to create one-to five-minute price changes. Although it is somewhat arbitrary what one takes as a "large" price change, for purposes of this study "large" is taken to be a one-to five-minute change greater than or equal to 0.75% in absolute value. The standard deviation of the 1,918,678 one-minute price changes computed in this study is 0.048%, and the standard deviation of the 1,688,955 five-minute price changes is 0.112%. A change of 0.75% is thus a very large change.

Given each large change, newswires were searched to see if an event could be found that led to the change. Table 1, which is at the end of this paper, lists the large price changes and the events that were found. This paper is essentially a discussion of Table 1. There are 4,417 trading days in the dataset (between 21 April 1982 and 29 October 1999), and in 220 of these days at least one large price change occurred, ie a one-to five-minute change greater than or equal to 0.75% in absolute value. Events were found for 69 of these days.

Knowledge of the 69 events in Table 1 may prove useful in other studies. Each of these events is big in that it changed the total value of US equities by a large amount rapidly. This information may be useful in examining changes in individual stock prices, both absolute and relative to price changes of other stocks. From a macroeconomic perspective, the events are macro shocks, and knowledge of these shocks may be useful in examining various macroeconomic questions.

It is important to stress that this study is purely descriptive. No attempt is made to explain why a particular event led to the large price change, why other similar events did not lead to large price changes, why many large price changes have no events associated with them, and so on. The main contribution of this paper is simply to list the 69 events.[2]

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[2] There does not appear to be other studies in which events have been identified in the way done in this paper. Mitchell and Mulherin (1994) and Berry and Howe (1994) examine the effects of the amount of news per unit of time on stock prices and trading volume. Niederhoffer (1971) examines the effects of the world events on daily stock prices. Boyd, Hu and Jagannathan (1999) examine daily S&P 500 changes around days in which there is an employment announcement. French and Roll (1986) examine the volatility of individual stock prices during trading and non-trading hours. Wood, McInsh and Ord (1985) examine the behaviour of a minute-by-minute market return index. Harris (1986) examines the behaviour of portfolio returns over 15-minute intervals. A number of studies have examined the effects of announcements on *daily* changes in stock prices, and these studies are discussed in Section 4.

It is also important to stress that with a very few exceptions it is almost certain that each of the 69 events listed in Table 1 caused the particular price change. The events can thus be interpreted as "facts". For example, it is almost certain that the five-minute price decrease of 0.79% on 16 July 1982 was essentially all due to the 4.10 pm money supply announcement (see line 8, Table 1). There would likely have been, of course, a price change had there been no announcement, since the price generally changes each minute, but with a standard deviation of 0.112%, a typical price change is very small relative to a change of 0.79%. For all intents and purposes one can attribute all of the price change to the money supply announcement.

A way of thinking about the events is the following. Consider asking stock brokers a few minutes after the occurrence of one of the price changes in Table 1 that is associated with an event what led, if anything, to the change. The main point here is that almost without exception the brokers would say the event. Some events may have been missed - more will be said about this later - but there is little doubt that each of the 69 events chosen led to the particular price change.

The construction of Table 1 is discussed in Section 2, and the results are discussed in Sections 3 and 4.

2. The construction of Table 1

The price of an S&P 500 futures contract follows closely the value of the S&P 500 index. Since the S&P 500 index includes most US stocks by market value, the price of an S&P 500 futures contract is a good indicator of the total value of US equities. Tick data are available for the S&P 500 futures contracts from April 1982 on.[3] For "Regular Trading Hours" (RTH) the tick data per day begin at 10.00 am prior to 30 September 1985, and at 9.30 am after that.[4] The RTH data end at 4.15pm, which is 15 minutes after the regular market has closed. Beginning in 1994 the contracts were traded after hours on the GLOBEX market, and tick data are available for these trades as well. These data begin at 4.30 pm and end at 9.15 am the next day. The GLOBEX market is closed Friday night and all day Saturday. It opens at 6.30 pm Sunday night.

For this study the RTH data begin in 21 April 1982 and end in 29 October 1999. Data are missing for the last half of December 1991 - the 1991 data end 13 December. The GLOBEX data begin in 4 January 1994 and end in 29 October 1999. Data are missing for the last half of 1998 - the 1998 GLOBEX data end 31 July. Many government announcements of macroeconomic data occur at 8.30 am, and since the GLOBEX market is open at this time, it can respond immediately to these announcements. Had the GLOBEX market been in existence back to 1982 and tick data been available, it is likely that many more large price changes and associated events would have been found. It is also likely that a number of large price changes and associated events would have been found in GLOBEX data for the last half of 1998 had the data been available.

The one-minute price change was taken to be the price of the last trade in the current minute interval less the price of the last trade in the previous minute interval (all changes in % terms). The two-minute price change was taken to be the price of the last trade in the current minute interval less the price of the last trade in the minute interval two minutes ago, and so on through five-minute price change.

Table 1 lists the following (a large change is always a change greater than or equal to 0.75% in absolute value): (1) all large one-minute price changes; (2) all large two-minute price changes except when at least one of the two one-minute price changes is large; (3) all large three-minute price changes except when at least one of the two two-minute price changes is large; (4) all large four-minute price changes except when at least one of the four one-minute price changes is large or at least one of the three two-minute price changes is large or at least on of the two three-minute price changes is large, and; (5) all large five-minute price changes except when at least one of the five one-minute price changes is large or at least one of the four two-minute price changes is large or at least one of the three three-minute price changes is large or at least one of the two four-minute price changes is large. This procedure finds all the large one- through five-minute price changes without

[3] The tick data were purchased from the Futures Industry Institute, which obtains the data from the Chicago Mercantile Exchange.

[4] All times in this paper are Eastern even though the RTH and GLOBEX markets are in the Central time zone.

duplication. The most actively traded contract on the particular day was used for these calculations. As can be seen from Table 1, there were a total of 1,159 changes chosen.

The “end time” in Table 1 is the time at the end of the *k*-minute change, where *k* ranges from 1 to 5. “Vol.” is the total number of ticks in the *k*-minute interval, and “ave. vol.” is the average number of ticks per minute.

The next step was to see which event, if any, led to the large and rapid change. The Dow Jones Interactive service on the internet was used for this purpose. This service allows one to search for news reports by time of day. The following four news services were searched: *Dow Jones News Service, Associated Press Newswire, New York Times* and *Wall Street Journal.*

For example, the first case in Table 1 is for 24 June 1982, where at 3.28 pm the price had fallen by 0.85% in the last five minutes. For this case the news services were searched for news reports between 3.00 pm and 4.00 pm to see what happened about 3.23 pm that led to the large change. In this case no news report was found that seemed likely to have led to change.

In the next case in Table 1 an event was found, which was the 4.10 pm announcement that M1 was down $2.3 billion. In the two minutes following the annoucement the price rose 0.82%. Although the regular stock market is closed at 4.00 pm, the RTH market does not close until 4.15 pm, and so the RTH market has time to respond to the money supply announcements.

In some cases an event was found that seemed almost surely to have led to the price change, but for which no exact time could be found. In these cases “?time” is used in Table 1 to denote that the exact time of the event was not found. For the 9 October 1990 change it is not completely clear that the Brazil event in fact led to the change, and this is indicated by a “(?)” in the table. For the 1 August 1997 change is unclear which of the three events listed led to the change, and this is also indicated by a “(?)” in the table.

An important government announcement each month is the employment report. This report is released at 8.30 am, and it contains data from both the household survey and the establishment survey. The main variable of interest from the household survey is the unemployment rate, and the main two variables of interest from the establishment survey are the number of jobs (called “payrolls”) and average hourly earnings. The variable that gets the most attention is the payroll variable, and so the payroll announcement is listed in Table 1. The “event” is, however, the entire employment report.

To save space in Table 1, not all large changes following an initial large change are listed for a particular day, especially on highly volatile days. When some changes are omitted, it is always indicated how many changes are omitted.[5]

3. Discussion of Table 1

Although, as discussed in Section 1, it is almost certain that each of the 69 events listed in Table 1 caused the particular price change, it may be that some events have been missed (aside from the missing RTH and GLOBEX data). The most likely error is an event for which there was no news report. Less likely is a news report that was listed in the search but that was not noticed as an important event. The number of events missed is likely to be small, probably fewer than ten. Remember, however, that many more large price changes and events would likely have been found had the GLOBEX market been in existence prior to 1994.

Assuming that the number of events that have been missed is small, Table 1 shows that there are many large price changes that are not due to identifiable events. There are, for example, no events associated with any of the large number of large price changes in October 1987. Regarding the price changes with no events, consider the thought experiment about stockbrokers mentioned in Section 1. For the price changes with no associated events in Table 1, what would stockbrokers say a few minutes after the change? The argument here is that except for the few events that might have been missed in the newswire searches, the brokers would not come up with a unique event. Some might

5 A complete table of all the changes is available. This table in pdf format is on the website mentioned in the introductory footnote. Click “Table 1A” near the bottom of the home page of the website for the table.

say there was no event, and some might mention something non-specific like “profit taking”, “renewed confidence”, “interest rate fears” and the like.

It should be stressed that the events that have been found are not necessarily surprises in the sense of an actual value differing from an expected value, although most of them probably are. Consider, for example, a payroll announcement. Say that market participants believe that there are three possible outcomes regarding the payroll change: 100,000, 300,000 and 500,000 jobs. Assume that market participants weight each possibility equally, so that the expected value is 300,000. Assume also that the participants expect that the Fed will leave the funds rate unchanged if the outcome is 100,000 or 300,000, but raise the funds rate if the outcome is 500,000. Assume finally that participants expect the S&P 500 price to be 1,430 if there is no funds rate change and 1,400 if there is one. The expected value of the price is thus 1,420, which if the participants are risk neutral will be the price before the announcement. In this case even if the actual payroll value is equal to the expected value (300,000), the stock price will change (from 1,420 to 1,430). Simply relieving uncertainty may thus change stock prices even if the announced value is equal to the expected value. The events that have been found are thus not necessarily surprises.

The main results from Table 1 are the following. First, the breakdown of the 69 events is:

- Twenty-two events are money supply or interest rate announcements or testimony by monetary authorities. In 1982 the focus was on money supply announcements, and after that it was on the federal funds rate;
- fourteen events are payroll announcements (employment reports);
- eleven events are CPI or PPI or employment cost index announcements;
- six events concern other macroeconomic announcements (NAPM, retail sales, durable goods, new homes);
- five events concern Iraq;
- four events concern Congressional issues;
- three events concern Brazil or Mexico;
- one event is fear of Larry Summers.

The 31 non-monetary macroeconomic announcements (payroll, CPI, PPI, employment cost and other) are indirectly related to monetary policy in that these announcements may change people’s expectations about future monetary policy. If, for example, there is a large payroll increase, people may think it is more likely that the Fed will tighten in the future because of fear of inflation. If these 31 announcements are added to the 22 direct monetary policy events, this gives 53 of the 69 events that are directly or indirectly related to monetary policy.

Second, the largest response by far was to the cut in the federal funds rate at 3.14 pm on 15 October 1998. The first five one-minute price changes following this announcement were 0.89, 1.00. 1.00, 1.29 and 1.00%. This is roughly a 5% increase in five minutes. The announcement of this rate cut was unusual in that it did not follow a normally scheduled FOMC meeting.

Third, the large price changes are not close to being spread evenly across years. Between 1982 and 1993, before the introduction of the GLOBEX market, the number of days of large price changes per year are respectively: 43, 2, 2, 0, 12, 33, 6, 4, 18, 8, 3 and 1. Between 1994 and 1999 the number of days are respectively: 5, 0, 12, 26, 22 (GLOBEX data for the last half of 1998 missing) and 23 (through October).

Finally, as noted above, many large price changes have no events associated with them.

4. Implications for other studies

It seems clear that no simple model of stock price determination can explain the facts in Table 1. There have, for example, been hundreds of important macroeconomic announcements between 1982 and 1999, and only a small fraction have led to a large stock price change. An adequate model would need to explain why the particular events in Table 1 led to large price changes, while many other seemingly similar events did not. There is also the problem from a model building perspective that there are many large price changes for which there appear to be no obvious causes.

A number of statistical studies have examined the effects of announcements on *daily* changes in stock prices (ie the change from the close of one day to the close of the previous day). The daily % change in a stock index is regressed on estimates of the "surprise" components of announcements, and the components are tested for their statistical significance. The surprise component of an announcement is the difference between the announced value and an estimate of its expected value. The expected value is usually either taken from a survey or to be a prediction from an autoregressive equation.

This literature generally finds that surprise monetary announcements are significant, but little else seems to matter. Schwert (1981), Pearce and Roley (1985) and Hardouvelis (1987) find surprise monetary announcements significant, and McQueen and Roley (1993) find inflation surprises sometimes significant after controlling for different stages of the business cycle. Jain (1988) finds surprise monetary and CPI announcements significant. The results in Table 1 suggest that if anything is to be found significant in explaining stock prices it is likely to be monetary announcements, which is what the literature tends to find. The "facts" in Table 1 thus provide some support to the statistical results using daily data, but they also suggest that an adequate model of stock price determination is likely to be more complicated than the models that have been used so far for the statistical tests.

As noted in Section 1, Cutler, Poterba and Summers (1989) chose the 50 largest daily changes in the S&P 500 index from the 1946 through 1987 and attempted to find an explanation of each change in the next day's *New York Times*. Of the 50 changes, 17 occurred between 1982 and 1987, which are years included in Table 1. It is interesting that five of these 17 changes occurred on days not listed in Table 1, in other words, on days in which there was not at least one large one- to five-minute price change. Of the 12 changes that occurred on days that are listed in Table 1, none of the price changes has an even associated with it. Table 2 lists the 12 changes and the *New York Times* explanation that Cutler, Poterba and Summers (1989, Table 4) found. It is clear that none of the explanations in Table 2 are obvious causes of the stock price changes. The results in Table 1 are consistent with this in that no events could be found to explain the large price changes on these days.

Haugen, Talmor and Torous (1991) examine daily changes in the Dow Jones Industrial Average between 1897 and 1988. They compute a measure of volatility using the daily data and choose periods of increased and decreased volatility. For the 217 periods of increased volatility that were chosen, they identified events for 28 of them. For the 224 periods of decreased volatility, they identified 18 events. Again, it is difficult with daily data to find events, which is probably the main reason they found so few events over such a long period of time.

Table 2

Twelve large daily S&P 500 price changes

Day	Percent change	*New York Times* explanation
08.17.82	4.76	Interest rates decline.
08.20.82	3.54	Congress passes Reagan tax bill; prime rate falls.
11.30.82	3.22	"Analysts were at a loss to explain why the Dow jumped so dramatically in the last two hours
09.11.86	– 4.81	Foreign governments refuse to lower interest rates; crackdown on triple witching announced
10.16.87	– 5.16	Fear of trade deficit; fear of higher interest rates; tension with Iran.
10.19.87	– 20.47	Worry over dollar decline and trade deficit; fear of US not supporting dollar.
10.20.87	5.33	Investors looking for "quality stocks".
10.21.87	9.10	Interest rates continue to fall; deficit talks in Washington; bargain hunting.
10.22.87	– 3.92	Iranian attack on Kuwait oil terminal; fall in markets overseas; analysts predict lower prices.
10.26.87	– 8.28	Fear of budget deficits; margin calls; reaction to falling foreign stocks.
10.29.87	4.46	Deficit reduction talks begin; durable goods orders increase; rallies overseas.
10.30.87	3.33	Dollar stabilises; increase in prices abroad.

Taken from Table 4 in Cutler, Poterba, and Summers (1989).

Fleming and Remolona (1997) (FR) examine five-minute price changes for the five year US Treasury note for the period 23 August 1993 - 19 August 1994. They chose the 25 largest five-minute price changes over this period, and they find that each of these changes was preceded by a macroeconomic announcement. Of these 25 changes, 17 are on days for which S&P 500 futures data exist. Data for these 17 days are presented in Table 3. The five-minute bond price change is presented (taken from Table 3 in Fleming and Remolona (1997)) along with the five-minute S&P 500 futures price change.[6]

The stock price changes in Table 3 are in general fairly large, although not nearly as large as 0.75%, the Table 1 cutoff. It is remarkable that in every case except the last one the bond and stock price changes are in the same direction. The direction is the same in the last case if the same time 1.45-1.50pm is used, but not if 1.16-1.21 pm is used for the stock price. As FR point out (p32), bond and stock prices need not move in the same direction following an announcement, since stock prices depend on expectations of both earnings and interest rates, whereas bond prices depend only on expectations of interest rates. The fact that they do move in the same direction suggests that the announcements mostly affect interest rate expectations.

Finally, Gwilym, McMillan and Speight (1999) examine five-minute stock price changes for the UK market using FTSE-100 data. The data are for the 24 January 1992 - 30 June 1995 period. Among other things, their data show that trading volume is higher around announcement times than otherwise. The results in Table 1 are consistent with this conclusion. For example, note that in general volume is quite high around the 8.30 am announcements in the table.

5. Conclusion

As mentioned in the Introduction, this study is purely descriptive. By focusing on very short time intervals, it has been possible to associate particular events with particular stock price changes, something which is generally not possible to do using daily data. Sixty-nine events have been identified between 1982 and 1999 that led to a one- to five-minute S&P 500 futures price change greater than or equal to 0.75% in absolute value. Knowledge of these events may prove useful in both macroeconomic studies and studies of individual stock prices.

The results in Table 1 suggest that stock price determination is complicated. Many large price changes correspond to no obvious events, and so many large changes appear to have no easy explanation. Also, of the hundreds of fairly similar announcements that have taken place between 1982 and 1999, only a few have led to large price changes (ie those in Table 1), and it does not appear easy to explain why some do and some do not.

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6 In some cases slightly different time intervals from the FR intervals were used. For the 8.30 am announcements, 8.29-8.34 am instead of 8.30-8.35 am was used, since at least in the S&P 500 futures data an 8.30 am announcement affects the 8.30 am price. For 3 June 1994, 8.30 am employment announcement, FR used 8.40-8.45 am and this was also done here. There was very little change in the price before 8.40 am. For the 2.26 pm announcement of the Federal funds rate on 17 May 1994 FR used 2.35-2.40 pm. In Table 3 both the stock price changes for 2.25-2.30 and 2.35-2.40 pm are presented. Finally, for the 1.17 pm announcement of the Federal funds rate on 16 August 1994, FR used 1.45-1.50 pm, and in Table 3 both the stock price changes for 1.16-1:21 pm and 1.45-1.50 pm are presented.

Table 3
Five-minute bond and stock price changes

Day	Bond interval	Bond change	Stock interval	Stock change	Announcement
01.07.94	8.30-8.35 am	0.282	8.28-8.33 am	0.07	8.30 am: Employment
02.04.94	8.30-8.35 am	0.315	8.29-8.34 am	0.15	8.30 am: Employment
02.04.94	11.05-11.10 am	– 0.259	11.04-11.09 am	– 0.09	11.05 am: Federal funds rate
02.11.94	8.30-8.35 am	0.223	8.29-8.34 am	0.31	8.30 am: PPI, retail sales
04.13.94	8.30-8.35 am	0.224	8.29-8.34 am	0.10	8.30 am: CPI, retail sales
05.06.94	8.30-8.35 am	– 0.536	8.28-8.33 am	– 0.14	8.30 am: Employment
05.11.94	1.40-1.45 pm	– 0.223	1.40-1.45 pm	– 0.37	1.42 pm: 10-year-note auction results
05.12.94	8.30-8.35 am	0.384	8.29-8.34 am	0.43	8.30 am: PPI, retail sales
05.17.94	2.35-2.40 pm	0.221	2.25-2.30 pm	0.33	2.26 pm: Federal funds rate
			2.35-2.40 pm	0.00	
05.27.94	8.30-8.35 am	– 0.343	8.29-8.34 am	0.20	8.30 am: GDP
06.03.94	8.40-8.45 am	– 0.265	8.40-8.45 am	– 0.23	8.30 am: Employment
07.08.94	8.30-8.35 am	– 0.440	8.29-8.34 am	– 0.30	8.30 am: Employment
07.12.94	8.30-8.35 am	0.222	8.29-8.34 am	0.28	8.30 am: PPI
07.14.94	8.30-8.35 am	0.253	8.29-8.34 am	0.11	8.30 am: Retail sales
07.29.94	8.30-8.35 am	0.407	8.29-8.34 am	0.31	8.30 am: GDP
08.05.94	8.30-8.35 am	– 0.590	8.29-8.34 am	– 0.35	8.30 am: Employment
08.16.94	1.45-1.50 pm	– 0.266	1.16-1.21 pm	0.23	1.17 pm: Federal funds rate
			1.45-1.50pm	– 0.09	

Notes: No stock trades at 8.29 am on 1.07.94 and 5.06.94. Changes are percent changes. Bond results and announcement information taken from Table 3 in Fleming and Remolona (1997).

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Table 1

One- to five-minute price changes greater than 0.75% in absolute value

k-minute-change							Event
	day	**k**	**end time**	**size**	**vol.**	**ave. vol.**	
1	06.24.82	5	3.28 pm	– 0.85	23	4.6	none
2	06.25.82	2	4.11 pm	0.82	13	6.5	4.10 pm: M1 down $2.3 billion.
3	07.06.82	4	3.36 pm	0.79	25	6.3	none
4	07.08.82	5	3.09 pm	0.78	39	7.8	none
5	07.09.82	3	4.11 pm	0.86	21	7.0	4.10 pm: M1 down $3.7 billion.
6	07.09.82	3	4.12 pm	0.99	22	7.3	–
7	07.13.82	4	12.21 pm	0.77	17	4.3	12.20 pm: IBM profits $1.68 vs. $1.37 year ago.
8	07.16.82	5	4.14 pm	– 0.79	43	8.6	4.10 pm: M1 up $5.9 billion
9	08.11.82	4	2.36 pm	0.79	33	8.3	2.30 pm: retail sales up 1.0%
10	08.13.82	2	4.13 pm	0.78	20	10.0	4.10 pm: M1 up $2.0 billion.
11	08.17.82	4	2.20 pm	0.87	22	5.5	none
12	08.19.82	4	10.16 am	0.79	20	5.0	none
13	08.19.82	5	2.08 pm	– 0.88	32	6.4	?time: rumour a major US bank in trouble over Mexican loans.
14	08.19.82	2	2.09 pm	– 0.78	15	7.5	NY Fed denied rumour about 2.30 pm.
15	08.19.82	5	2.14 pm	0.79	46	9.2	–
16	08.19.82	4	2.29 pm	0.93	33	8.3	–
17	08.19.82	5	2.31 pm	0.79	38	7.6	–
18	08.19.82	5	2.32 pm	0.88	34	6.8	–
19	08.19.82	5	2.33 pm	0.93	31	6.2	–
20	08.19.82	4	2.34 pm	0.79	23	5.8	–
21	08.19.82	1	2.38 pm	– 0.88	13	13.0	–
22	08.19.82	1	2.40 pm	0.88	12	12.0	–
23	08.19.82	3	2.48 pm	– 0.78	21	17.0	–
24	08.19.82	3	2.49 pm	– 0.83	17	5.7	–
25	08.20.82	5	3.21 pm	0.77	29	5.8	none
26	08.23.82	5	3.36 pm	0.85	30	6.0	none
27	08.23.82	4	3.37 pm	0.76	24	6.0	–
28	08.23.82	5	3.40 pm	0.80	24	4.8	–
29	08.24.82	3	1.41 pm	– 0.78	17	5.7	1.40 pm: GM mid-August sales down to 81,597 from 134,949.
30	08.24.82	5	3.19 pm	– 0.77	42	8.4	none
31	08.24.82	5	3.22 pm	– 0.77	43	8.6	–
32	08.24.82	5	3.23 pm	– 0.82	41	8.2	–
33	09.02.82	3	2.56 pm	0.77	13	4.3	none
34	09.03.82	4	10.04 am	0.78	29	7.3	none
35	09.03.82	3	10.21 am	0.78	33	11.0	–
36	09.03.82	3	10.23 am	0.90	39	13.0	–
37	09.03.82	5	12.43 pm	0.78	22	4.4	–
38	09.03.82	5	1.07 pm	– 0.82	34	6.8	–
39	09.03.82	5	1.08 pm	– 0.78	33	6.6	–
40	09.03.82	3	1.39 pm	0.82	10	3.3	–
41	09.03.82	4	1.58 pm	– 0.82	20	5.0	–
42	09.03.82	5	3.25 pm	0.82	24	4.8	–
43	09.14.82	5	3.34 pm	– 0.88	32	6.4	3.27 pm: Rostenkowski said tax boost needed for defence budget.
44	09.14.82	4	3.35 pm	– 0.84	29	7.3	–
45	09.23.82	5	11.12 am	– 0.77	27	5.4	?time: Five Fed Presidents testify before Congress.
46	09.30.82	5	3.38 pm	– 0.79	36	7.2	none
47	10.01.82	3	4.12 pm	– 0.84	24	8.0	4.10 pm: M1 up $0.4 billion.
48	10.08.82	4	2.27 pm	– 0.77	17	4.3	none
49	10.08.82	1	2.28 pm	0.88	6	6.0	–
50	10.08.82	2	4.05 pm	0.85	19	9.5	–
51	10.08.82	2	4.07 pm	– 0.96	19	9.5	–
52	10.08.82	3	4.10 pm	0.77	37	12.3	4.10 pm: M1 down $2.7 billion.
53	10.08.82	3	4.11 pm	0.92	42	14.0	–
54	10.11.82	1	2.55 pm	– 0.82	4	4.0	none

k-minute-change							Event
	day	k	end time	size	vol.	ave. vol.	
55	10.11.82	4	3.46 pm	0.82	37	9.3	–
56	10.13.82	3	12.37 pm	– 0.78	22	7.3	none
57	10.13.82	3	2.09 pm	0.79	23	7.7	–
58	10.13.82	3	3.23 pm	0.86	24	8.0	–
59	10.13.82	3	3.24 pm	1.01	36	12.0	–
60	10.13.82	2	3.27 pm	0.98	23	11.5	–
61	10.13.82	2	3.29 pm	– 0.90	16	8.0	–
62	10.13.82	4	3.56 pm	0.79	18	4.5	–
63	10.22.82	4	4.13 pm	– 0.85	59	14.8	4.10 pm: M1 up $3.2 billion.
64	10.26.82	4	2.58 pm	– 0.78	23	5.8	none
65	10.26.82	3	2.59 pm	– 0.82	15	5.0	–
66	10.26.82	5	3.20 pm	0.82	36	7.2	–
67	10.27.82	5	2.59 pm	0.77	33	6.6	none
68	11.02.82	4	3.34 pm	– 0.83	32	8.0	none
69	11.05.82	2	11.46 am	– 0.8	17	8.5	none
70	11.05.82	2	11.47 am	– 0.98	11	5.5	–
71	11.05.82	3	3.14 pm	0.77	27	9.0	none
72	11.05.82	2	4.10 pm	– 0.77	12	6.0	4.10 pm: M1 up $2.7 billion.
73	11.05.82	2	4.11 pm	– 0.84	12	6.0	–
74	11.16.82	4	3.15 pm	0.81	35	8.8	?time: Larry Speaks reported to have said Fed will reduce discount rate. Denied at 3.33 pm.
75	11.16.82	4	3.16 pm	0.77	33	8.3	–
76	11.16.82	4	3.17 pm	0.77	34	8.5	–
77	11.22.82	5	10.17 am	– 0.77	37	7.4	none
78	11.30.82	5	2.53 pm	0.79	40	8.0	none
79	11.30.82	5	2.54 pm	0.82	38	7.6	–
80	12.01.82	5	2.27 pm	– 0.82	40	8.0	none
81	12.01.82	4	2.28 pm	– 0.82	38	9.5	–
82	12.01.82	5	2.32 pm	– 0.78	41	8.2	–
83	12.02.82	4	2.38 pm	– 0.75	24	6.0	2.30 pm: New home sales down 0.4%
84	12.06.82	4	3.26 pm	0.75	39	9.8	none
85	12.06.82	4	3.27 pm	0.86	43	10.8	–
86	12.07.82	5	2.55 pm	– 0.83	41	8.2	none
87	12.08.82	5	2.48 pm	– 0.77	32	6.4	none
88	12.09.82	3	3.24 pm	– 0.85	23	7.7	?time: Howard Baker withdrew capital gains bill.
89	12.14.82	5	1.38 pm	– 0.80	33	6.6	none
90	12.14.82	2	2.02 pm	– 0.77	9	4.5	–
91	12.14.82	4	3.27 pm	– 0.84	30	7.5	–
92	12.15.82	5	10.59 am	– 0.80	28	5.6	10.56 am: Murray Weidenbaum testified that deficit an obstacle to recovery.
93	12.17.82	4	12.00 pm	0.77	7	1.8	none
94	12.21.82	2	3.40 pm	0.77	18	9.0	none
95	12.22.82	5	3.21 pm	– 0.78	27	5.4	none
96	12.28.82	3	2.56 pm	– 0.80	18	6.0	none
97	01.04.83	5	3.02 pm	0.76	37	7.4	none
98	01.06.83	5	12.09 pm	0.80	50	10.0	none
99	05.30.84	4	2.37 pm	0.80	38	9.5	none
100	08.06.84	5	10.33 am	0.89	70	14.0	none
101	01.08.86	3	3.48 pm	– 0.79	48	16.0	none
102	01.08.86	3	3.49 pm	– 0.90	63	21.0	–
103	01.08.86	4	3.53 pm	0.77	63	15.8	–
104	01.17.86	4	9.58 am	– 0.76	38	9.5	9.54 am: IBM profits $4.36 vs. $3.55 year ago.
105	02.07.86	3	12.02 pm	– 1.03	47	15.7	12.00 pm: three judge panel ruled part of Gramm-Rudman law unconstitutional.
106	02.07.86	2	12.07 pm	0.75	33	16.5	–
107	09.11.86	4	11.36 am	– 0.88	59	14.8	none
108	09.11.86	4	11.53 am	0.78	58	14.5	–
109	09.12.86	4	10.05 am	– 0.76	33	8.3	none (nine more through 12.16 pm)
119	09.15.86	5	9.36 am	0.78	44	8.8	none

k-minute-change							Event
	day	k	end time	size	vol.	ave. vol.	
120	09.18.86	4	9.45 am	0.86	40	10.0	none
121	09.19.86	2	3.31 pm	– 0.78	36	18.0	none
122	09.25.86	4	11.01 am	– 0.76	46	11.5	none
123	09.25.86	5	11.03 am	– 0.83	50	10.0	–
124	10.03.86	5	11.08 am	– 0.87	63	12.6	none
125	10.03.86	2	11.09 am	– 0.76	21	10.5	–
126	10.03.86	3	11.11 am	– 0.80	25	8.3	–
127	10.03.86	3	11.17 am	0.76	28	9.3	–
128	10.06.86	2	10.06 am	0.75	31	15.5	none
129	12.11.86	4	11.15 am	– 0.95	46	11.5	none
130	12.11.86	2	11.16 am	– 0.79	28	14.0	–
131	01.23.87	4	2.35 pm	– 0.76	52	13.0	none (14 more during day)
146	03.09.87	5	9.35 am	– 0.86	74	14.8	
147	03.09.87	5	9.36 am	– 0.81	73	14.6	–
148	03.30.87	5	9.46 am	– 0.85	54	10.8	none
149	03.30.87	4	9.47 am	– 0.78	42	10.5	–
150	03.30.87	3	9.48 am	– 0.95	31	10.3	–
151	04.13.87	5	3.37 pm	– 0.81	53	10.6	none
152	04.27.87	4	12.26 pm	– 0.76	52	13.0	none
153	04.27.87	4	12.27 pm	0.78	43	10.8	–
154	04.27.87	4	12.42 pm	0.76	47	11.8	–
155	04.27.87	5	12.44 pm	0.80	67	13.4	–
156	04.27.87	5	12.45 pm	0.80	61	12.2	–
157	04.27.87	5	12.46 pm	0.87	66	13.2	–
158	05.11.87	4	4.00 pm	– 0.84	48	12.0	none
159	06.02.87	5	10.07 am	– 0.91	70	14.0	none
160	06.10.87	5	3.09 pm	– 0.75	45	9.0	none
161	06.30.87	2	4.14 pm	– 0.77	36	18.0	none
162	10.16.87	5	11.25 am	– 0.83	42	8.4	none (12 more during day)
175	10.19.87	1	9.33 am	0.76	7	7.0	none (123 more during day)
299	10.20.87	1	9.31 am	1.78	5	5.0	none (161 more during day)
461	10.21.87	1	9.31 am	1.26	7	7.0	none (82 more during day)
544	10.22.87	1	9.31 am	– 3.47	5	5.0	none (109 more during day)
654	10.23.87	1	9.31 am	– 1.23	4	4.0	none (33 more during day)
688	10.26.87	1	9.32 am	– 0.50	9	9.0	none (22 more during day)
711	10.27.87	1	9.31 am	0.84	9	9.0	none (22 more during day)
734	10.28.87	2	9.37 am	0.80	10	5.0	none (22 more during day)
757	10.29.87	1	9.32 am	1.27	6	6.0	none (7 more during day)
765	10.30.87	4	9.47 am	0.79	28	7.0	none (5 more during day)
771	11.02.87	4	9.34 am	– 0.78	38	9.5	none
772	11.02.87	5	9.36 am	– 0.78	50	10.0	–
773	11.03.87	2	9.35 am	0.80	19	9.5	none (18 more during day)
792	11.04.87	4	9.46 am	0.77	38	9.5	none
793	11.04.87	3	9.47 am	0.77	26	8.7	–
794	11.05.87	4	9.35 am	0.81	38	9.5	none
795	11.06.87	4	11.41 am	0.78	38	9.5	none
796	11.09.87	5	2.34 pm	– 0.83	52	10.4	none
797	11.10.87	5	9.35 am	– 0.88	48	9.6	none
798	11.10.87	4	9.36 am	– 0.88	39	9.8	–
799	11.10.87	5	10.10 am	– 0.82	51	10.2	–
800	11.19.87	2	4.02 pm	– 0.75	28	14.5	none

k-minute-change							Event
	day	k	end time	size	vol.	ave. vol.	
801	12.01.87	4	4.00 pm	– 0.77	64	16.0	none
802	12.08.87	4	3.58 pm	0.85	43	10.8	none
803	12.10.87	3	3.39 pm	– 0.76	38	12.7	none
804	12.14.87	4	9.37 am	0.78	65	16.3	none
805	12.15.87	5	3.31 pm	0.75	59	11.8	none
806	01.08.88	4	3.39 pm	– 0.76	60	15.0	none (4 more through 3.49 pm)
811	01.11.88	2	9.52 am	– 0.82	21	10.5	none (7 more through 10.16 am)
819	01.15.88	3	9.43 am	– 0.78	31	10.3	none
820	01.21.88	5	10.15 am	0.82	59	11.8	none
821	04.14.88	4	2.30 pm	– 0.86	58	14.5	none (6 more through 2.41 pm)
828	04.21.88	5	3.08 pm	– 0.81	64	12.8	none
829	10.13.89	3	3.04 pm	– 0.75	27	9.0	none (6 more through 3.44 pm)
836	10.16.89	1	9.31 am	1.20	10	10.0	none (18 more during day)
855	10.17.89	5	11.17 am	– 0.79	58	11.6	none
856	10.17.89	5	11.25 am	0.88	37	7.4	–
857	10.17.89	5	11.26 am	0.88	41	8.2	–
858	10.24.89	4	10.24 am	– 0.78	56	14.0	none (9 more during day)
868	01.12.90	3	9.33 am	– 0.86	22	7.3	none
869	01.12.90	4	2.36 pm	– 0.90	40	10.0	–
870	01.12.90	4	2.41 pm	0.84	47	11.8	–
871	01.24.90	3	9.39 am	– 0.83	25	8.3	none
872	01.24.90	3	9.40 am	– 0.77	24	8.0	–
873	01.24.90	1	9.41 am	0.77	7	7.0	–
874	07.23.90	5	10.30 am	– 0.75	58	11.6	none
875	07.23.90	4	10.32 am	– 0.77	49	12.3	–
876	07.23.90	2	10.33 am	– 1.10	19	9.5	–
877	07.23.90	5	10.54 am	0.91	43	8.6	–
878	08.03.90	4	9.47 am	– 0.85	31	7.8	?time: Iraq invaded Kuwait.
879	08.03.90	3	9.48 am	– 0.85	19	6.3	–
880	08.03.90	3	9.49 am	– 1.13	22	7.3	–
881	08.03.90	5	9.55 am	0.97	34	6.8	–
882	08.03.90	2	1.49 pm	– 1.11	14	7.0	–
883	08.03.90	4	2.01 pm	0.88	19	4.8	–
884	08.03.90	5	2.04 pm	0.86	28	5.6	–
885	08.06.90	5	10.27 am	0.84	31	6.2	none
886	08.17.90	3	12.16 pm	– 0.78	40	13.3	12.11 pm: Pentagon recommended maybe calling up reserves.
887	08.21.90	1	11.16 am	0.77	13	13.0	11.13 am: Iraq's Aziz says ready to discuss Gulf situation.
888	08.23.90	3	9.46 am	0.96	16	5.3	none
889	09.20.90	4	10.25 am	– 0.82	59	14.8	none
890	09.27.90	4	10.59 am	– 0.77	50	12.5	none
891	10.01.90	4	12.18 pm	0.76	43	10.8	none
892	10.01.90	4	12.19 pm	0.78	38	9.5	–
893	10.02.90	5	9.39 am	0.78	54	10.8	none
894	10.09.90	4	3.45 pm	– 0.79	47	11.8	3.39 pm: (?) Brazil's central bank president sees rescheduling needed.
895	10.10.90	3	3.28 pm	– 0.88	39	13.0	none
896	10.12.90	3	11.55 am	0.77	31	10.3	11.51 pm: opposition party in exile says Iraqi leaders considering Kuwait withdrawal.
897	10.15.90	3	10.45 am	– 0.82	42	14.0	none
898	10.15.90	2	10.46 am	– 0.82	29	14.5	–
899	12.04.90	5	3.33 pm	0.99	45	9.0	?time: British TV reports Iraq makes new offer on Kuwait.
900	12.04.90	2	3.34 pm	0.84	20	10.0	–

k-minute-change							Event
	day	k	end time	size	vol.	ave. vol.	
901	12.18.90	2	3.27 pm	0.84	29	14.5	none
902	12.18.90	3	3.29 pm	0.88	23	7.7	–
903	01.04.91	2	12.12 pm	0.83	20	10.0	none
904	01.09.91	2	1.57 pm	– 0.92	22	11.0	none
905	01.09.91	1	1.58 pm	– 1.50	10	10.0	–
906	01.09.91	1	1.59 pm	– 1.56	6	6.0	–
907	01.09.91	1	2.00 pm	0.94	5	5.0	–
909	01.09.91	2	2.04 pm	– 0.78	14	7.0	–
910	01.14.91	4	3.19 pm	0.84	44	11.0	none
911	01.16.91	1	12.43 pm	0.84	12	12.0	none
912	01.17.91	2	9.44 am	– 0.93	13	6.5	none
913	01.17.91	4	9.47 am	– 0.93	29	7.3	–
914	04.30.91	2	9.32 am	0.87	28	14.0	9.30 am: Fed cut discount rate to 5.5%.
915	05.10.91	5	3.25 pm	– 0.78	54	10.8	none
916	11.15.91	4	3.41 pm	– 0.81	34	8.5	none
917	11.15.91	4	3.46 pm	0.76	21	5.3	–
918	11.15.91	4	3.47 pm	0.76	27	6.8	–
919	11.19.91	5	10.47 am	– 0.81	57	11.4	none
920	01.02.92	5	4.05 pm	0.77	68	13.6	none
921	07.02.92	4	10.21 am	– 0.82	62	15.5	?time: Fed cut discount rate to 3.0% from 3.5%;
922	07.02.92	4	10.22 am	0.87	58	14.5	anemic employment report earlier.
923	10.05.92	5	10.21 am	– 0.78	57	11.4	none
924	10.05.92	3	10.26 am	– 0.83	33	11.0	–
925	10.05.92	3	10.27 am	– 0.78	24	8.0	–
926	10.05.92	3	11.16 am	– 0.86	27	9.0	–
927	02.16.93	5	10.43 am	– 0.79	55	11.0	none
928	03.02.94	1	5.15 am	– 1.52	4	4.0	none
929	03.31.94	5	10.55 am	– 0.78	52	10.4	none
930	03.31.94	5	10.57 am	– 0.76	45	9.0	–
931	03.31.94	4	11.07 am	0.75	27	6.8	–
932	10.13.94	4	8.33 am	0.75	70	17.5	8.30 am: PPI down 0.5; core up 0.1%
933	11.15.94	3	2.38 pm	– 0.83	38	12.7	2.37 pm: Fed raised discount rate to 4.75% from 4.0%
934	11.22.94	4	3.51 pm	– 0.81	47	11.8	none
935	02.26.96	5	3.27 pm	– 0.76	79	15.8	none
936	02.26.96	5	3.28 pm	– 0.76	70	14.0	–
937	03.08.96	4	8.33 am	– 0.79	110	27.5	8.30 am: payrolls up 705,000; largest increase in 12 years.
938	03.08.96	1	8.34 am	– 0.75	17	17.0	–
939	03.08.96	2	3.03 pm	– 0.80	21	10.5	none
940	03.08.96	4	3.27 pm	0.77	31	7.8	–
941	03.08.96	4	3.28 pm	0.77	26	6.5	–
942	04.10.96	4	3.43 pm	– 0.76	46	11.5	none
943	05.03.96	4	8.30 am	0.78	89	22.3	8.30 am: payrolls up 2,000
944	05.03.96	4	8.31 am	0.89	120	30.0	–
945	05.03.96	3	8.32 am	0.96	137	45.7	–
946	05.03.96	3	8.35 am	– 0.77	172	57.3	–
947	06.07.96	2	8.31 am	– 0.93	94	47.0	8.30 am: payrolls up 348,000
948	06.07.96	3	8.33 am	– 0.84	123	41.0	–
949	06.07.96	3	8.34 am	– 0.93	124	41.3	–
950	07.11.96	3	1.44 pm	– 0.79	20	6.7	none
951	07.15.96	4	3.29 pm	– 0.85	40	10.0	none
952	07.16.96	4	12.22 pm	– 0.84	40	10.0	none (7 more during day).

k-minute-change							Event
	day	k	end time	size	vol.	ave. vol.	
960	08.02.96	2	8.31 am	1.02	91	45.5	8.30 am: payrolls up 193,000.
961	09.06.96	1	8.30 am	0.81	61	61.0	8.30 am: payrolls up 250,000.
962	09.13.96	3	8.32 am	0.84	123	41.0	8.30 am: CPI up 0.1%; core up 0.1%. Also retail sales data.
963	12.11.96	5	2.12 am	0.80	31	6.2	none
964	01.23.97	4	3.46 pm	– 0.84	27	6.8	none
965	01.23.97	5	3.48 pm	– 0.76	35	7.0	–
966	01.23.97	5	3.55 pm	– 0.75	41	8.2	–
967	01.23.97	4	4.00 pm	0.76	37	9.3	–
968	01.29.97	3	8.31 am	0.76	91	30.3	8.30 am: durable goods down 1.7%.
969	02.05.97	3	3.40 pm	0.76	23	7.7	none
970	02.26.97	4	10.04 am	– 0.97	35	8.8	10.00 am: Greenspan testimony; angst about stock market.
971	02.26.97	2	10.05 am	– 0.86	14	7.0	–
972	03.07.97	1	8.30 am	– 0.81	48	48.0	8.30 am: payrolls up 339,000.
973	03.07.97	2	8.32 am	0.75	87	43.5	–
974	03.27.97	4	3.37 pm	– 0.87	30	7.5	none
975	03.27.97	4	3.45 pm	0.75	26	6.5	–
976	04.15.97	3	8.31 am	0.78	115	38.3	8.30 am: CPI up 0.1%; core up 0.2%.
977	04.15.97	3	8.32 am	0.99	182	60.7	–
978	04.29.97	2	8.31 am	1.09	106	53.0	8.30 am: employment cost index up 0.6% in first quarter.
979	05.20.97	4	2.17 pm	0.82	51	12.8	2.15 pm: Fed kept rates unchanged.
980	06.06.97	1	8.31 am	– 0.88	77	77.0	8.30 am: payrolls up 138,000.
981	07.09.97	4	3.38 pm	– 0.75	43	10.8	none
982	07.18.97	4	10.08 am	– 0.80	44	11.0	none
983	08.01.97	3	10.11 am	– 0.94	24	8.0	10.00 am: ? new orders up 1.2%; strong NAPM report; Michigan sentiment revised up.
984	08.08.97	2	8.29 am	– 0.89	39	19.5	none???
985	08.08.97	3	8.32 am	0.76	65	21.7	–
986	08.13.97	4	8.31 am	0.79	164	41.0	8.30 am: PPI down 0.1%; core down 0.1%. Also retail sales data.
987	08.13.97	4	8.32 am	0.85	226	56.5	–
988	08.13.97	4	8.33 am	1.04	265	66.3	–
989	08.13.97	3	8.34 am	0.84	192	64.0	–
990	08.13.97	5	10.22 am	– 0.82	64	12.8	none
991	08.13.97	3	10.23 am	– 0.77	38	12.7	–
992	08.13.97	3	10.24 am	– 0.77	34	11.3	–
993	08.22.97	5	3.53 pm	0.83	57	11.4	none
994	09.02.97	2	10.03 am	0.84	22	11.0	10.00 am: NAPM 56.8 vs. 58.6 last month.
995	10.03.97	2	8.31 am	– 0.77	97	48.5	8.30 am: payrolls up 215,000.
996	10.10.97	1	8.30 am	– 0.82	73	73.0	8.30 am: PPI up 0.5%; core up 0.4%.
997	10.27.97	3	1.52 pm	– 0.85	28	9.3	none
998	10.27.97	4	1.55 pm	– 0.80	42	10.5	–
999	10.27.97	2	1.56 pm	– 0.80	13	6.5	–
1000	10.27.97	2	3.12 pm	– 1.17	20	10.0	–
1001	10.27.97	2	3.13 pm	– 0.96	22	11.0	–
1002	10.27.97	4	3.23 pm	– 0.96	27	6.8	–
1003	10.27.97	3	3.24 pm	– 0.85	18	6.0	–
1004	10.28.97	3	6.48 am	–0.75	42	14.0	none
1005	10.28.97	4	9.42 am	0.82	27	6.8	none
1006	10.28.97	4	10.13 am	0.87	40	10.0	none (15 more during day)
1022	10.30.97	5	10.02 am	– 0.88	45	9.0	none
1023	11.07.97	5	9.35 am	-0.86	52	10.4	none

k-minute-change							Event
	day	k	end time	size	vol.	ave. vol.	
1024	11.07.97	5	9.40 am	0.76	44	8.8	–
1025	12.05.97	5	8.30 am	– 0.82	129	25.8	8.30 am: payrolls up 404,000.
1026	12.05.97	2	8.31 am	– 0.92	151	75.5	–
1027	12.23.97	5	3.56 pm	– 0.79	51	10.2	none
1028	04.30.98	2	8.31 am	0.87	154	77.0	8.30 am: employment cost index up 0.7% in first quarter. Also GDP data released.
1029	08.04.98	4	3.43 pm	– 0.93	39	9.8	none
1030	08.04.98	4	3.46 pm	– 0.80	34	8.5	–
1031	08.05.98	4	9.43 am	– 0.88	22	5.5	none
1032	08.05.98	4	3.45 pm	0.93	41	10.3	none
1033	08.27.98	3	11.38 am	– 0.84	29	9.7	none
1034	08.27.98	4	11.41 am	– 0.84	46	11.5	–
1035	08.27.98	5	2.47 pm	0.84	36	7.2	none
1036	08.27.98	5	2.48 pm	0.84	40	8.0	–
1037	08.28.98	5	11.49 am	0.77	47	9.4	none
1038	08.28.98	5	11.50 am	0.86	45	9.0	–
1039	08.31.98	4	10.09 am	– 0.96	31	7.8	none
1040	08.31.98	4	3.07 pm	– 0.77	47	11.8	none (10 more during day)
1051	09.01.98	2	9.32 am	0.94	19	9.5	none (21 more during day)
1073	09.02.98	3	9.37 am	0.75	30	10.0	none
1074	09.02.98	5	3.49 pm	– 0.80	58	11.6	
1075	09.03.98	4	9.35 am	– 0.82	37	9.3	none
1076	09.03.98	5	10.30 am	0.93	47	9.4	–
1077	09.03.98	5	10.56 am	– 0.77	45	9.0	–
1078	09.03.98	4	3.42 pm	0.77	37	9.3	–
1079	09.04.98	5	2.50 pm	– 0.76	39	7.8	none
1080	09.04.98	5	3.39 pm	0.76	48	9.6	–
1081	09.04.98	5	3.41 pm	0.85	45	9.0	–
1082	09.10.98	5	3.36 pm	– 0.80	49	9.8	none
1083	09.10.98	3	3.42 pm	0.76	37	12.3	–
1084	09.11.98	4	9.46 am	0.87	49	12.3	none
1085	09.11.98	4	9.48 am	0.77	51	12.8	–
1086	09.11.98	5	9.50 am	0.77	69	13.8	–
1087	09.11.98	3	9.53 am	0.77	36	12.0	–
1088	09.11.98	2	10.03 am	0.92	20	10.0	–
1089	09.11.98	4	2.50 pm	0.82	34	8.5	–
1090	09.11.98	4	2.51 pm	0.82	37	9.3	–
1091	09.16.98	5	3.43 pm	0.81	40	8.0	none
1092	09.29.98	2	2.17 pm	– 0.89	18	9.0	2.17 pm: fed cut funds rate 25 basis points.
1093	09.29.98	2	2.18 pm	– 1.08	18	9.0	–
1094	09.29.98	3	2.20 pm	– 0.84	30	10.0	–
1095	10.01.98	5	11.19 am	– 0.79	51	10.2	none
1096	10.02.98	5	10.18 am	– 0.90	47	9.4	none
1097	10.02.98	4	12.41 pm	0.80	42	10.5	none
1098	10.02.98	5	2.52 pm	0.75	45	9.0	none
1099	10.02.98	4	3.15 pm	0.75	30	7.5	none
1100	10.05.98	4	3.37 pm	0.78	31	7.8	none
1101	10.05.98	4	3.42 pm	0.80	33	8.3	–
1102	10.05.98	4	3.43 pm	0.75	30	7.5	–
1103	10.07.98	5	10.12 am	0.91	41	8.2	none
1104	10.07.98	5	3.06 pm	0.76	41	8.2	none
1105	10.07.98	4	3.26 pm	– 0.75	44	11.0	none
1106	10.07.98	5	3.41 pm	0.91	48	9.6	none
1107	10.08.98	3	9.50 am	– 0.78	32	10.7	none (eight more during day)
1116	10.09.98	4	10.14 am	– 0.77	31	7.8	none
1117	10.15.98	1	3.15 pm	0.89	11	11.0	3.14 pm: Fed cut funds rate and discount rate 25 basis points (not a normal FOMC meeting).
1118	10.15.98	1	3.16 pm	1.00	11	11.0	
1119	10.15.98	1	3.17 pm	1.00	9	9.0	–
1120	10.15.98	1	3.18 pm	1.29	10	10.0	–

k-minute-change							Event
	day	k	end time	size	vol.	ave. vol.	
1121	10.15.98	1	3.19 pm	1.00	8	8.0	–
1122	10.15.98	1	3.21 pm	– 1.19	8	8.0	–
1123	10.15.98	1	3.22 pm	– 1.29	7	7.0	–
1124	10.15.98	5	3.44 pm	0.80	36	7.2	–
1125	10.15.98	5	3.45 pm	0.80	34	6.8	–
1126	10.15.98	5	3.46 pm	0.80	36	7.2	–
1127	11.17.98	2	2.19 pm	0.83	19	9.5	2.15 pm: Fed cut funds rate and discount rate 25 basis points.
1128	11.17.98	2	2.20 pm	1.14	17	8.5	–
1129	01.15.99	5	7.31 am	– 0.78	198	39.6	none
1130	01.15.99	4	8.10 am	0.77	102	25.5	8.10 am: Estado said Brazil central bank won’t
1131	01.15.99	5	8.12 am	0.81	125	25.0	intervene in foreign exchange market.
1132	01.15.99	5	8.16 am	0.82	159	31.8	–
1133	01.31.99	1	5.57 pm	0.85	9	9.0	none
1134	02.23.99	5	10.04 am	– 0.82	68	13.6	10.00 am: Greenspan testimony; economy may be stretched.
1135	03.05.99	2	8.31 am	1.05	152	76.0	8.30 am: payrolls up 275,000.
1136	04.16.99	4	9.48 am	– 0.78	47	11.8	none
1137	04.16.99	5	1.17 pm	– 0.86	54	10.8	none
1138	05.12.99	1	9.47 am	– 0.81	17	17.0	?time: Rubin to announce resignation;
1139	05.12.99	1	9.48 am	– 0.81	12	12.0	Summers is successor.
1140	05.12.99	1	9.50 am	0.88	12	12.0	–
1141	05.18.99	4	2.14 pm	– 0.80	41	10.3	2.11 pm: Fed let rates stand; adopted tightening bias.
1142	06.01.99	5	10.05 am	– 0.85	54	10.8	10.00 am: NAPM 55.2 vs. 52.8 last month.
1143	06.04.99	3	8.33 am	– 0.88	191	63.7	8.30 am: payrolls up 11,000.
1144	06.13.99	1	5.35 pm	– 0.77	2	2.0	none
1145	06.15.99	1	4.19 pm	– 1.01	3	3.0	none
1146	06.16.99	3	8.30 am	0.79	170	56.7	8.30 am: CPI unchanged; core up 0.1%
1147	06.30.99	2	2.17 pm	0.96	24	12.0	2:15 pm: Fed raised funds rate 25 basis points; adapted neutral bias.
1148	06.30.99	3	2.19 pm	0.89	31	10.3	–
1149	08.06.99	2	8.30 am	– 0.83	151	75.5	8.30 am: payrolls up 310,000.
1150	09.03.99	1	8.30 am	0.89	150	150.0	8.30 am: payrolls up 124,000.
1151	09.10.99	1	8.31 am	0.80	93	93.0	8.30 am: PPI up 0.5%; core down 0.1%
1152	09.15.99	1	8.30 am	0.81	106	106.0	8.30 am: CPI up 0.3% core up 0.1%.
1153	10.05.99	2	2.12 pm	– 0.93	32	16.0	2.12 pm: Fed let rates stand; adopted tightening bias.
1154	10.05.99	2	2.13 pm	– 0.80	27	13.5	–
1155	10.08.99	1	8.30 am	0.75	126	126.0	8.30 am: payrolls down 8,000.
1156	10.13.99	2	4.21 pm	– 0.77	5	2.5	none
1157	10.13.99	1	4.22 pm	– 0.84	1	1.0	–
1158	10.15.99	1	8.30 am	– 0.95	150	150.0	8.30 am: PPI up 1.1%; core up 0.8%.
1159	10.20.99	1	4.27 pm	1.21	3	3.0	none

Notes: CPI = consumer price index; core excludes food and energy. PPI = producers price index; core excludes food and energy. Percentage changes are at monthly rates except for the change in the employment cost index, which is at quarterly rate. NAPM = National Association of Purchasing Managers.

Comments on “Events that shook the market” by Ray Fair

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Professor Fair’s findings highlight some of the dilemmas we face in trying to understand price formation and market liquidity, both from a macro and a micro-economic perspective. Firstly, I would like to express my admiration for the amount of work that has gone into collecting data on the millions of 5 minute intervals between 1982 and 1999 and the extensive research of newswires for clues on what might have triggered large intra-day stock price changes. As a practitioner I am appealed by the use of tick data as it reflects the actual behaviour of financial markets, and may shed light on issues such as market dynamics and market structure. The more common use of average daily data is less useful in this regard. My comments on the paper are divided into three parts: general comments, technical issues and a suggestion for further research.

General comments

One important observation is that we still do not understand what moves market prices very well. In the paper only one third of the large price movements could be traced to a single event. Most of the ‘explainable’ events are related to data releases impacting expectations on the future course of monetary policy. This seems to suggest that central banks still matter, which is a comforting thought if you work for one, as I do. The fact that macro-economic data show up as an important determinant of stock prices should not of course be too surprising given the fact that interest rates impact the cost of capital of firms and the net present value of future earnings.

In the search for explanations of price moves it may be useful to have knowledge of the market conditions prevailing at the time of the identified events. In particular, some of the large price changes may have followed the breaching of important technical support and resistance levels that prevailed at the time or may be related a lack of market liquidity. My experience with the monitoring of developments in foreign exchange market suggests that such triggers for large-price moves are frequent. It may also be worthwhile to look into the pattern of order flows during the course of these days. This volume information could shed light on the depth of the market, i.e. the ability of the market to absorb large order flows without substantial price moves. It may also shed light on the behaviour of market participants in the minutes surrounding important data releases. I suspect a bunching of orders and trades at these times. Thus it may be possible to filter out some cases of large intra-day price that may be more related to market functioning than to the incorporation of new information into the market. This may be particularly the case when the large price move does not persist during the course of subsequent five-minute trading intervals.

Technical comments

- With regard to the tick data, I have assumed that these data reflect prices upon which trades (however small or large) have actually taken place. If this is not the case, for instance when such data is based on mid-price of the best bids and offers in the central order limit book, one has to be aware of distortions caused by widening bids and offers during moments of market uncertainty.
- The data shows there are several occasions of successive five-minute periods with large price moves. This suggests that a longer period may be used in such cases to describe a particular event.
- The stock market crash of October 1987 is one of those ‘non explainable’ events during which price developments and a seizing up of market liquidity reflected a widespread loss of confidence in the stability of the financial system. The divergence of valuations in cash and futures markets reflected a fragmentation of price discovery and market liquidity, which was in part triggered by large delays in processing order flows. Analysis of such price divergences on an intra-day basis may allow some judgement on the evolution of market functioning on those tumultuous days.

Suggestion for further research

The main contribution of this paper is to make an extensive database of tick data available and to link large intra-day price moves to certain events. In addition to searching for news events that impacted prices, further research into market conditions prevailing at the time of large price moves seems promising. This information would be useful form a micro-economic perspective, to better understand the functioning and structure of financial markets.

A second suggestion for further research concerns the impact of changes in the shape of the yield curve on stock prices. Here it may be of interest to research if there are different outcomes in the case of Nasdaq and S&P500 stocks. Some analysts have suggested the valuation of technology stocks are less sensitive to the level of interest rates since earnings lie further in the future (in some sense comparable with a high duration bond). Since there is also a very liquid index-futures contract on the Nasdaq-index, it would be of interest to study differences in the behaviour of this contract vis-à-vis the S&P futures for the large intra-day price moves in recent years.

Comments on "Events that shook the market" by Ray Fair

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Discussion

In the empirical literature, the relationship between news and asset prices has typically been analysed by asking two different questions: What is the price impact of a certain type of news? What news has caused a certain large price change? The first, "ex-ante", approach is usually carried out by regressing the changes in asset prices over a variable representing the event or its unexpected component. The second, "ex-post", approach is implemented in two stages: first, "large" price changes are computed and then an event that may explain each such big move is looked for.

Recently intraday data have allowed researchers to gain more profound insights on the relationship between asset prices and information. By using transaction data for the US bond market, Fleming and Remolona (1999a) attempt to identify information that may account for the sharpest price changes, and also estimate the price impact of scheduled macroeconomic announcements. Andersen and Bollerslev (1998) address the same two key questions for the deutsche mark-dollar spot exchange rate.

The paper by Professor Fair extends this line of research to the U.S. stock market, applying the ex-post methodology to a huge amount of transaction data on the S&P 500 futures contract (the sample period runs from 21 April 1982 to 29 October 1999). The paper is a major contribution that provides a highly detailed picture of the price discovery process in the US stock market.

The results of the above-mentioned studies are quite similar for the bond and forex markets. For the on-the-run five-year US Treasury note, Fleming and Remolona (1999a) found that, between September 1993 and August 1994, each of the twenty-five sharpest five-minute price changes could be associated with a just-released macroeconomic announcement. Similarly, Andersen and Bollerslev (1998) were able to associate an event with each of the twenty-five largest five-minute jumps in the deutsche mark-dollar exchange rate recorded between October 1992 and September 1993. In this case, however, the events included 15 releases of macroeconomic data, eight other economic events and three political events.

In Fair's study of the S&P 500 futures contract, 220 of the 4417 trading days in the dataset recorded at least one large price change. The most striking result of the paper is that events were found only for 69 of these days. The latter included 31 macroeconomic announcements, 22 monetary policy events, seven other economic events and nine political events. These findings are consistent with those obtained by Cutler, Poterba and Summers (1989) for the 50 largest daily changes in the S&P 500 index from 1946 through 1987.

My brief comments will discuss three aspects of the work: (1) the definition of a "large" stock price change; (2) the sensitivity of stock prices to news about monetary policy and the business cycle; (3) the main finding that several stock price gaps are apparently unexplained.

Identifying "large" price changes is analogous to a problem of outlier detection. As such, it should take into account the statistical properties of the S&P 500 transactions data, and in particular the persistence in their volatility (Chan, Chan and Karolyi, 1991). As regards the time interval, Fair devises a procedure that finds all the large one to five-minute changes without duplication. This is a convenient solution. Nevertheless, as pointed out by Fair, it is important to consider that five out of the 17 price gaps computed on daily data by Cutler, Poterba and Summers between 1982 and 1987 are not detectable on one to five-minute data, indicating that the results are substantially affected by the sampling frequency. The choice of the minimum size is also very important. Fair sets a level of 0.75 percent in absolute value. The choice of an unconditional threshold implies that during turmoil a huge number of price gaps are selected, while in normal periods large jumps in prices can be missed. The adoption of a time-varying threshold level could significantly increase the share of (conditional) large price changes that are explained by some news event.

Table 1 clearly shows that monetary policy is a factor in stock price gaps. For instance, in the second half of the nineties seven interest rate decisions shook the market. In that period, two Testimonies also caused price gaps. This evidence is not surprising. Since monetary policy decisions in the US are made known in a limited number of scheduled FOMC meetings and speeches per year, market participants tend to revise their expectation about the future stance of monetary policy in the days in which these institutional events take place. It is also interesting to note that in the last three interest rate shocks in 1999 that were accompanied by a statement on the bias of monetary policy it was the latter that determined the correction in stock prices, rather than the change in the policy interest rate. This means that market participants are ready to extract information about the future tightness of monetary policy from any signal provided by the central bank. This fact should be taken into account by empirical studies that attempt to identify monetary policy shocks on high frequency data by focusing exclusively on the unexpected component of the policy interest rate.

It must be emphasised that these events represent purely exogenous monetary policy shocks. As Fair clearly points out with respect to the first monetary surprise detected on 16 July 1982 “for all intents and purposes one can attribute all of the price change to the money supply announcement”. As purely discretionary monetary shocks, these events could be used to identify sectorial responses to monetary policy shocks by carrying out event study analyses of the kind used by Fleming and Remolona (1999b) for the bond market. (To reduce the incidence of non-synchronous trading, the analysis could be limited to the stocks included in the S&P 100.)

While there is no doubt that at least 22 events can be directly ascribed to Fed decisions, the effect of news about payroll, prices, etc. cannot be primarily connected with expectations about the future stance of monetary policy. Business cycle news affects not only expected discount rates, but also expected cash flows and expected excess returns on stocks (notably the perceived “riskiness” of stocks). The relevance of developments in real activity for the stock market is likely to be much more pervasive than the evidence reported in Table 1 seems to suggest. First, as noted by Fair, since many macroeconomic announcements occur at 8.30 am, if we also had data on after-hours trading for the period before 1994 (when the S&P 500 futures contract began to be traded on the GLOBEX market), it is likely that many more effects of macroeconomic data would have been found. Moreover, since information about the business cycle is dispersed over a wide set of indicators (Stock and Watson, 1999), the price impact of each data release can be limited. Finally, if we set a lower minimum level for price gaps, we would certainly find more events linked to macroeconomic announcements. In fact, between September 1993 and August 1994 the impact on the S&P 500 futures price of 17 relevant macroeconomic news events was comparable to that found for bond prices (cf. Table 3).

As mentioned above, the most striking result of the paper is the huge number of large price changes that apparently are not associated with any event, a situation that seems peculiar to the stock market. Cutler, Poterba and Summers (1989) discard the possibility that market analysts can systematically miss news about fundamental values that are instead observed by market participants, and indicate two alternative explanations. First, large market movements may reflect changes in average assessments of fundamental values as investors re-examine existing data or present new arguments. Second, many investors do not trade on the basis of their own assessment of values, but rely on market prices to gauge them-a form of departure from the assumption of rational behaviour. According to Cutler, Poterba and Summers, the latter hypothesis would explain, for instance, why, during the stock market crisis of October 1987, most shares were not traded despite the dramatic drop in their prices.

Recent theoretical research has pointed out how the introduction of fundamental uncertainty in standard equilibrium models can help replicate key aspects of financial data series, and in particular volatility clustering and time-varying risk premia (Veronesi, 1999). If the economy shifts between two unobservable states at random times (say, between high-growth and low-growth states), risk-averse investors will tend to overreact to bad news in good times and to underreact to good news in bad times, since in these cases the incoming information increases the uncertainty of investors about the latent state of the economy. It also implies that the reaction of prices to news will tend to be large in good times and small in bad times.

A different explanation assumes that there are investors who show some kind of cognitive bias and limit the activity of fully rational arbitrageurs. In these cases, prices will tend to overreact or underreact to fundamental news depending on the processes of belief formation assumed for the different types of investors (Schleifer, 2000).

The stylised facts about stock price movements highlighted by Fair can provide a useful yardstick for evaluating the predictive accuracy of these different explanations.

In conclusion, the paper by Professor Fair clearly shows that intraday data are well suited for investigating the relationship between price movements and information arrival for the stock market as well. Full exploitation of the statistical properties of intraday data could substantially improve the outcome of the signal extraction process.

In the US, monetary policy is a factor in large stock price movements. The paper identifies a number of purely exogenous monetary policy shocks that provide suggestions for further research. Such evidence is also relevant to the debate on whether and to what extent central banks should abstain from monetary surprises.

The most striking feature of the US stock market is that many, if not most, large price changes are unexplained. A theoretical explanation of this fact could be provided by recent equilibrium models that assume rational learning or some departure from the hypothesis of rational expectations, along the lines suggested by Cutler, Poterba and Summers (1989) after the stock market crash of 1987.

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Information flows during the Asian crisis: evidence from closed-end funds

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Abstract

A salient feature of the Asian crisis of 1997 was a collapse of stock markets that took place over several months. The dynamics of this collapse raises the question of what information was driving the markets. This paper examines a key aspect of this question: did information flow from the domestic Asian markets to overseas markets, or vice versa? We test for the direction of this information flow by comparing daily returns in several Southeast Asian equity markets with daily returns on US-based closed-end funds that invest in those markets, exploiting the fact that there is no overlap between the trading hours in the two regions. We find that while information flows between local and US markets tended to be roughly evenly balanced *before* the crisis, US market returns assumed a more important role *during* the crisis. This is the case both for the level of daily returns and for the volatility of those returns. We also find that fund returns were more closely tied to broad US market returns during the crisis period. This suggests that the shift in causation between the United States and Asia reflected a greater role for US market sentiment, rather than for the news that became known during US trading hours.

I. Introduction

Did the financial turmoil that affected many emerging economies in the middle and late 1990s stem primarily from developments within those economies or from events in financial markets in the industrial countries? Proponents of the former view have pointed to poor policy choices in the emerging economies, particularly in such areas as exchange-rate policy, banking supervision and corporate governance. Adherents of the latter view emphasise the suddenness and magnitude of the reversal in capital flows to the emerging economies, and the fact that markets seemed to “punish” geographically similar but otherwise sound economies with high credit-risk premia and reduced capital-market access. This debate over events in the recent past is of relevance to a number of current policy issues, including the appropriateness of restrictions on international capital flows, the role of the International Monetary Fund, and the “bailing in” of private sector lenders in sovereign debt workouts.

The two positions are not necessarily mutually exclusive. Some commentators concede that the crisis economies were flawed, but assert that global investors overreacted to their difficulties. Martin Wolf (1998) comments that whatever the policy crimes, these “hardly justify the enormity of the punishment.”[1] Fischer (1998), by contrast, characterises the countries’ problems as “mostly homegrown” and points to a number of common policy faults - specifically macroeconomic overheating, pegged exchange rates, and weak bank supervision - though he also acknowledges significant differences among the countries.[2] Paul Krugman explains the severity and spread of the crisis by likening it to a bank run.[3] The issue then becomes: Who ran? Some analysts have argued that capital outflows represented a self-fulfilling “rush for the exits” by panicked foreign investors, while others claim the outflows were initiated by massive capital flight by “front-running” domestic investors.

This paper does not offer a conclusive resolution to this debate. It does, however, attempt to provide an insight into a key aspect of it, namely: in the period surrounding the crisis, did information about financial market returns in emerging economies flow from the domestic market to overseas markets, or vice versa? In this study, “information” is defined broadly to include anything that might have a material effect on returns, including changes in investor sentiment.

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1 Wolf, M, “Flows and Blows,” *Financial Times*, 3 March 1998.

2 Fischer, S, “The Asian Crisis: A View from the IMF,” speech at the Midwinter Conference of the Bankers’ Association for Foreign Trade, Washington, D.C., 22 January 1998.

3 Krugman, P, “Paradigms of Panic: Asia Goes Back to the Future,” *Slate Magazine*, 12 March 1998.

We test for the flow of information by comparing daily returns in several east Asian equity markets with daily returns on US-based closed-end funds that invest in those markets. Because there is essentially no overlap between the trading hours in the two regions, we can safely assume that all of the information incorporated into a day's trading in Asian markets will be available to those trading the closed-end funds in the United States that same day. Similarly, the information incorporated in a day's closed-end fund trading is fully available for the next day's trade in Asia. Completing the picture of information flows, the populations of investors in the two markets are likely to differ as well. Indeed, a primary function of closed-end funds is to allow investors in mature markets to gain exposure to the corresponding emerging markets, without requiring them to trade directly in those markets. This distinction might be less useful, however, to the extent that foreign investors are active in local Asian markets.

Previous studies of country fund behaviour, such as Frankel and Schmukler (1996) and Bodurtha, Kim and Lee (1995), compare fund prices to their net asset values (NAVs). Bodurtha, Kim and Lee find that that movements in the premia of fund prices over their NAVs are highly correlated and reflect US stock market returns, implying an important role for US investor sentiment in fund returns. Frankel and Schmukler examine prices and NAVs on three closed-end funds investing in Mexico around the December 1994 peso crisis in an attempt to gauge the relative levels of sentiment of local and foreign investors. They find that the funds tended to trade at a discount before the crisis and at a premium afterwards, suggesting that foreign investors were relatively more optimistic than their local counterparts. They further note that NAV returns tended to "cause" price returns in a Granger sense, further supporting the view that the drop in confidence during the crisis had strong local roots.

We examine this relationship below, but we also compare fund returns to local market returns, for two reasons. First, NAVs are available only at a weekly frequency while local returns are available daily, allowing a finer analysis of price behaviour. Second, the time difference between the two markets allows us to study timing issues, rather than simple correlations.

Evidence on the timing of securities returns on essentially identical securities in different markets - ie on whether price movements in one market tend to lead or lag price movements in the other - could be informative either about the timing of the arrival of relevant news, or about the timing of changes in sentiment regarding the level or riskiness of expected returns. In the case of Asian markets and US closed-end funds, it is likely that most, though not all, of the relevant news becomes known during Asian trading hours. Exceptions might be official statements or policy decisions (such as IMF program announcements) by institutions located in the United States and Europe. Significant changes of sentiment, on the other hand, could conceivably occur among either group of investors. A finding that returns in closed-end funds led returns in local markets would thus be evidence for the importance of mature-market investor sentiment in determining emerging-market returns. A finding that returns in local markets led those in closed-end funds, on the other hand, would be less conclusive. While such a finding could indicate an important role for local sentiment, returns could simply be reacting to local news.

We find that information flows between local and US markets tended to be roughly evenly balanced before the "crisis period" beginning in July 1997, but that US market returns assumed a relatively more important role during the crisis. This is the case both for the level of daily returns and for the volatility of those returns. We also find that the funds are more reflective of the broad US market return during the crisis period. This suggests that the shift in causation between the United States and Asia reflected a greater role for US market sentiment, rather than for the news that became known during US trading hours.

Corroborating evidence is provided by the behaviour of US purchasers of Asian equities. Using aggregated data provided by a large securities custodian, we find that the positive influence of local-market returns on fund returns tends to be weaker at times when US investors purchased large amount of Asian equities. In other words, large equity flows to Asia are associated with looser price links between markets, while flows out of Asia are associated with stronger price links. This suggests that US investors tend to be contrarians in their portfolio activities vis-à-vis Asia: they purchase Asian equities just at those times when their opinions differ most strongly from those in the local markets.

To some extent these results can be counted as evidence against the "front-running" hypothesis tested by Kramer and Smith (1995) and Frankel and Shmukler (1996). While we do find, as Frankel and Shmukler do for Mexico, that the funds' prices move from a discount to NAV to a premium after the crisis started, we do not find that sentiment among Asian investors drives changes in sentiment among US investors. Instead, our results are closer to those of Choe, Kho and Stulz (1999), who find that foreign investors in the Korean equity market followed momentum (positive feedback) strategies before the crisis and contrarian (negative feedback) strategies during the crisis.

The next section reviews the debate about the direction of information flows before, during and after the Asian crisis. Section III examines characteristics of closed-end funds, including country funds, and discusses how they might shed light on the information-flow debate. Section IV describes the funds and local returns used in this study and discusses the behaviour of the discount to NAV over the period studied. Section V presents results on spillovers of the level and volatility of returns before, during and after the crisis, while Section VI examines return and volatility spillovers during times of investor inflows and outflows. Section VII concludes.

II. Information flows and the Asian crisis

The second half of 1997 saw the unprecedented collapse of the stock markets and currencies of five Asian countries - Thailand, Indonesia, Malaysia, the Philippines and South Korea. By year's end, the five Asian currencies had shed a third to three quarters of their values. The stock markets of Bangkok, Jakarta, Kuala Lumpur, and Manila had lost USD 370 billion or 63% of the four countries' combined GDP. The Seoul stock market had declined 60%. The debacle effectively ended years of impressive economic performance by these countries.

The first sparks of the Asian crisis may have started in Thailand in March 1997 when loan problems of several finance companies came to light.[4] The Bangkok stock market fell 25% over the next three months and the Thai baht came under increasing pressure. When the Thai authorities devalued the baht in July, the crisis quickly became a regional one, spreading to Indonesia, Malaysia, and the Philippines (Figs. 1 and 2). In October, the Korean won and the Seoul stock market joined the carnage after credit rating agencies downgraded several of the country's banks. In the region, Taiwan stood out as a country that has escaped the crisis virtually unscathed.

The relatively benign macroeconomic conditions of the Asian countries and their somewhat different circumstances make the severity and spread of the crisis a puzzle. Prime Minister Mahathir of Malaysia blames such a run on "highwaymen of the global economy," hedge fund managers in particular.[5] Brown, Goetzmann, and Park (1998), however, estimate the currency positions of ten large hedge funds and find nothing unusual about these funds' net positions or profits during the crash period. If these authors are right, a key issue that remains is whether blame can be placed on other classes of foreign investors.

Certainly a sharp overall decline in net inflows of foreign capital to the crisis countries accompanied the collapse in currency values and stock prices, whether as cause or effect (Table 1). Portfolio investment fell much more sharply than did foreign direct investment in most countries, although the degree and timing of this decline varied. Banking flows, in particular, reversed dramatically from the second half of 1997 onwards (Graph 1).

The behaviour of equity investors is more ambiguous. Data assembled by State Street, a large international securities custodian, indicate that foreign investors often tended to increase their purchases of domestic equities precisely at those times when foreign banks were reducing their exposures (Graph 2). However, in the case of Indonesia, foreign equity sales reinforced cutbacks in bank lending. These data are discussed in more detail in part VII below.

Clearly it is problematic to determine the impact on the crisis countries of "foreigners" as a group, given the divergent response to the crisis shown by different groups of investors. The remainder of this paper is focused on the behaviour of one important group of foreign investors, namely participants in the market for closed-end funds that hold Asian equities. To the extent that these investors are representative of foreign equity investors in Asia as a whole, the analysis can shed light on an important aspect of the questions identified in the Introduction. Foreign equity investors as a group certainly play an important role in local markets in Asia; for example, in the Korean stock market, foreign investors held 12% at 2 December 1996 and 14.73% at 27 December 1997 (Choe, Kho and

4 See Moreno, Pasadilla, and Remolona (1998) for a narrative of the crisis. They also offer an explanation based on Krugman (1998) and McKinnon and Pill (1998).

5 Mahathir, M, "Highwaymen of the Global Economy," *Wall Street Journal*, 23 September 1997.

Stulz, 1999). However, answers to the broader question of the role of foreign investors in the Asian crisis await a detailed study of the role of banks, hedge funds and other investor classes.

III. Closed-end funds

A closed-end fund is an investment vehicle that has a fixed number of shares and invests in a portfolio of stocks, bonds, and other securities, usually with a specialised focus. Closed-end country funds hold portfolios consisting of shares in firms based in a specific country or group of countries.

After its initial offering, new investors can obtain shares in the fund only by purchasing them from other investors. The market price of the fund and the net asset value (NAV) of its holdings tend to differ, because of the difficulty of engaging in arbitrage between fund shares and the shares in the fund's portfolio. In particular, it is difficult to take short positions in most closed-end funds because they are not actively traded. For funds that hold stocks traded in the US domestic market, the price tends to be below the NAV, while US-based funds that specialise in stocks from foreign countries can trade at both large premiums and large discounts (Bodurtha, Kim and Lee, 1995; Bonser-Neal et al, 1990). Funds can be terminated in two ways, either by a change in the fund's structure to that of an open-ended fund or by liquidation, both of which result in the value of the shares equalling that of the fund's NAV. The uncertainty as to the final termination date is another factor hindering arbitrage between the fund's price and its NAV. In other words, funds which trade at a discount to their NAV can be thought of as promising a positive excess return, relative to the underlying assets, over an uncertain horizon.

Explanations for the presence of this implied excess return vary. A traditional view emphasises agency costs: "Because the managers of closed-end funds are perceived to be less responsive to profit opportunities than open-end fund managers, who must attract and retain shareholders, closed-end fund shares often sell at a discount from net asset value" (Downes and Goodman, 1991). More recent analysts, including Lee, Shleifer and Thaler (1991), note that a fund and its component stocks are likely to be held by different clienteles of investors. Specifically, Lee, Shleifer and Thaler find that the funds in their sample tend to be held disproportionately by individual rather than institutional investors. Given the difficulty of arbitrage, the fund price and the NAV can therefore reflect differences in "sentiment" across these clienteles, and indeed the difference between them can act as an index of small-investor sentiment relative to that of the rest of the market. Lee, Shleifer and Thaler further propose that the tendency for closed-end funds to trade at prices below their NAV, that is to offer positive excess returns, compensates for the risk of liquidity-related selling or large swings in sentiment on the part of individual investors.

For country funds, an additional factor influencing the divergence of prices from NAVs is the presence of barriers to the access of foreign investors to local markets. These barriers include legal restrictions, transaction costs, and liquidity premia. They have the effect of enhancing any effects resulting from differences in sentiment, by reinforcing the distinction between the investment clienteles of local markets and those of closed-end funds. Bonser-Neal et al. (1990) find that announcements of reductions in these barriers tend to cause fund prices to decline relative to NAVs, regardless of whether the fund had previously been trading at a premium or a discount. The fact that these announcements reduce premia and increase discounts, rather than reducing the divergence of price from NAV in either direction, would argue against the view that free cross-border portfolio flows drive fund prices and NAVs together while restrictions on flows drive them apart. Instead, it indicates that investors in country funds are willing to accept relatively lower returns when barriers are high, and that removing these barriers reduces one of the attractions of the funds, causing the fund price to fall until investors again are satisfied with the prospective returns.

IV. The behaviour of premia

For the present study, closed-end funds are identified that represent each of the Asian countries considered to have been most affected by the 1997-98 crisis: Korea, Indonesia, Thailand, Malaysia and the Philippines. We include two funds that invested in each of Korea, Indonesia, and Thailand, and one fund investing in each of Malaysia and the Philippines. As controls, we also include two funds for Taiwan, which is considered to have been relatively less affected by the crisis than the other five. This produces a sample of ten funds from six countries (Table 2).

Six of the ten funds sold at prices that were, on average, at a positive premium to their NAVs during 1995-99, of which five were still at a premium on the last day of 1999. For each of the ten funds, there were times during the sample period when it sold at a premium and times when it sold at a discount. Premia tended to be closely correlated for the countries for which two funds are observed, suggesting that investor sentiment specific the country concerned, rather than factors unique to specific funds such as the perceived abilities of the fund managers, tended to be the key factor moving the premia (Table 3).[6] Correlations of premia for funds from different countries are not especially high. The figures in the bottom five lines in each panel in Table 3, however, seem to be consistently higher than those in the top four lines, suggesting that premia for Thailand, Indonesia, Malaysia and Philippines funds were more closely synchronised with one another than they were with those for Korea and Taiwan funds, or than Korea and Taiwan premia were with each other.

While there is no persistent pattern as to whether the funds tended to trade at a premium or a discount, their behaviour before, during and after the crisis illustrates the evolution in investor sentiment towards the region (Graph 3). Premia rose for all of the funds from the crisis countries starting in mid-1997. For some countries, such as Korea and Indonesia, the jump in premia was quite sudden, while for others, such as Thailand, a gradual increase in premia can be detected from late 1996 onwards. By early 1998, all of the funds from the five crisis countries traded at positive premia, while the two Taiwan funds continued to exhibit discounts. Premia declined gradually in the course of 1998 and 1999 in most cases, though for some, especially the two Thailand funds, they remained high and volatile.

It will be useful to define the period from 1 July 1997 to 31 October 1998 as the “crisis era”. This covers the time from the floating of the Thai baht on 2 July 1997, to the stabilisation of markets in the course of October 1998, and thus corresponds roughly to the most acute phase of the crisis in terms of economic developments in the region itself.[7]

Average premia during this crisis period were higher than those for the previous two and a half years (1 January 1995-30 June 1997) for eight of the ten funds studied (Table 4). While during the pre-crisis period, six of these eight sold at a discount to NAV, during the crisis all but one sold at a premium. The two cases where premia declined (that is, discounts were larger) are the two Taiwan funds. In every case, the crisis-period average premium is significantly different from that of the pre-crisis period according to the standard t-test.

Average premia in the fourteen months following October 1998 were lower than during the crisis period for the Korea, Indonesia, and Philippines funds. The premium on the Malaysian Fund fell, but to an insignificant degree, despite the country’s economic recovery. This may reflect Malaysia’s imposition of controls on foreign exchange and portfolio flows from September 1998, thus confirming the results of Bonser-Neal et al. For Thailand, premia continued to rise after the crisis. For Taiwan, the discounts narrowed again after the crisis, while remaining larger than pre-crisis levels.

These observations generally correspond to Frankel and Schmukler’s findings for Mexico, and appear to support their interpretation of those findings, namely that foreign investors in the crisis countries tended to be more optimistic than local investors during the crisis period. After the crisis, these shifts were reversed for Korea, the Philippines, and Indonesia, but not for Thailand or Malaysia. For Taiwan, the discount widened during the crisis and then narrowed afterwards. Taken in isolation, this might seem to be a sign of contagion: US investors turned bearish on Taiwan because of the region’s problems, while local investors remained calm. In conjunction with the results for the other five countries, however, where the shifts in sentiment moved in the opposite direction, the contagion interpretation seems less convincing. Rather than a divergence in sentiment about the valuation of the Taiwanese market, the wider discount for Taiwan may instead have indicated an increase in the risk premium demanded by US investors for Taiwanese assets.

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6 Premia are defined here and elsewhere in the paper as the log of the ratio of price to NAV. The term “premium” will be used generically to refer to the difference between price and net asset value, even when the price is below the NAV (in which case the term “discount” will occasionally be used as well).

7 In terms of disruptions to global financial markets, crisis conditions can be said to have persisted until later in 1998, or even into the early months of 1999. See BIS (1998, 1999a, 1999b).

V. Tracking the direction of influence: daily price changes

Fund premia offer an indication of the relative levels of sentiment of US and Asian investors, but they cannot tell us whether and in what ways the attitudes of these two groups of investors influence one another. In this section, we attempt to answer these questions by examining daily price changes in the two regions, relying on the fact that the two markets are open at different times.

V.1 Impact of Asian local returns on US country funds

Daily prices of closed-end funds are modeled according to the following GARCH(1,1) specification:

$$\begin{aligned}
FR_t^i &= \alpha_0 + \alpha_1 LR_t^i + \alpha_2 RR_t^i + \alpha_3 US_t + \alpha_4 FR_{t-1}^i \\
&\quad + \alpha_5 d_t^{797} + \alpha_6 d_t^{797} LR_t^{\,i} + \alpha_7 d_t^{797} RR_t^i + \alpha_8 d_t^{797} US_t + \alpha_9 d_t^{797} FR_{t-1}^i \\
&\quad + \alpha_{10} d_t^{1198} + \alpha_{11} d_t^{1198} LR_t^{\,i} + \alpha_{12} d_t^{1198} RR_t^i + \alpha_{13} d_t^{1198} US_t + \alpha_{14} d_t^{1198} FR_{t-1}^i \\
&\quad + \alpha_{15} LR_t^{i\,2} + \varepsilon_t^i \\
\varepsilon_t^i &\sim N(0, h_t^i) \\
h_t^i &= \gamma_0 + \gamma_1 \varepsilon_t^{i\,2} + \gamma_2 h_{t-1}^i + \gamma_3 d_t^{797} + \gamma_4 d_t^{1198} + \gamma_5 LR_t^{i\,2} + \gamma_6 d_t^{797} LR_t^{i\,2} + \gamma_7 d_t^{1198} LR_t^{i\,2}
\end{aligned} \tag{1}$$

where *i* indexes funds and countries and the variables are defined as follows:

FR_t^i: The daily log change in the closing price of the fund. Where two funds are available for a given country, the average of the two daily returns is used.

LR_t^i: The daily log change in the closing level of the local stock market index corresponding to the fund, in US dollar terms. For most of the markets studied, a "broad" and a "narrow" market index were available. Where possible, we use the "narrow" indices, in order to match the tendency for the country funds to buy shares of a relatively small number of large-capitalisation stocks in their respective markets.[8] The indices used are listed in Table 5. Asian-close exchange rates were used to translate the local currency returns into dollar returns.

RR_t^i: A regional return index, formed as an equally weighted average of the daily returns on the five local indices *excluding* that of country *i*.

US_t: The daily log change in the S&P 500 index.

d_t^{797}: A dummy variable taking the value one from 1 July 1997 through the end of the sample.

d_t^{1198}: A dummy variable taking the value one from 1 November 1998 through the end of the sample.

This specification has a number of important features.

It permits the fund return to reflect both a US market factor (US_t) and a regional market factor (RR_t). We thus accommodate the findings of Diwan, Errunza and Senbet (1995) and Bodurtha, Kim and Lee (1995), who find that US-based country fund returns reflect both US market returns and home market returns. The regional factor allows for the possibility of contagion effects. US investors might take information from a regional return into account when pricing the country fund, even if the regional return has not yet been incorporated (or not fully incorporated) into the local market.

- We include the lagged fund return to correct for autocorrelation, which was found to be present in many of the fund returns. To the extent that local-market returns reflect the previous day's country-fund returns (as will be discussed below), we want to eliminate the

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[8] For example, the Indonesia Fund invested in 23 local issues as of 30 June 1999, of which seven constituted more than half of the fund's holdings. On the same date, the Thai Fund was invested in 26 local issues, with more than half of its holdings accounted for by ten issues.

impact of autocorrelation in the country-fund returns as much as possible and focus on what "new" is learned from the day's local-market returns.

- Using two dummy variables, d_t^{797} and d_t^{1198}, lets us ask not only whether price behaviour differs during the crisis period, but also whether markets returned to their previous behaviour in its aftermath.
- The estimated returns are characterised by autoregressive conditional heteroskedasticity. This reflects the findings of ARCH-LM tests (results of which are available from the authors). A further discussion of the GARCH specification and the estimated coefficients for the volatility equation can be found in Section VI.
- There is a risk that the volatility equation will pick up a non-linear relationship between local and fund returns, rather than a pure volatility linkage. For this reason, $LR_t^{i\,2}$ is included on the right-hand side of the mean equation. The estimated coefficients (not reported here) are significant in a few cases, and positive in nearly all cases, suggesting that a non-linear, convex relationship probably does exist.

The first panel of Table 6 presents results of the estimation of equation (1) by OLS for the fund returns for the six countries in our sample. The second and third panels of Table 6 respectively report the sums of the coefficients on the local, regional and US return variables during the crisis (7/97-10/98) and afterwards (11/98-12/99) for each equation. The first column of Table 6 reports the results from an estimation of equation (1) using the average of the six fund returns as the dependent variable, the average of the six regional returns as RR_t^i, and dropping the right-hand terms in LR_t^i. Graph 4 shows the coefficients on the local and US market returns in the three periods for the six-country portfolio and the individual countries.

During all three periods studied, both the local return and the US market return are positive and significant (at the 10% significant level) for each fund, confirming the findings of Diwan, Errunza and Senbet (1995) and others. The regional return, reflecting overall investor sentiment in Asia, is significant for all six individual country returns and for the portfolio of funds before July 1997, but has a less consistent impact thereafter for some countries. The adjusted R^2 terms indicate that our four factors (the three shown and the lagged fund return) and two dummies explain between 26 and 49% of fund returns over the sample period.

Before July 1997, a remarkably consistent fraction - between 0.49 and 0.59 - of each country's daily local return is reflected in corresponding closed-end fund returns. US market sentiment is also an important factor for country closed-end funds, with a factor loading ranging from 0.15 to 0.50.

During the Asia crisis, the local return tends to become less important, and the US market return more important. For Korea, Thailand, Indonesia, Malaysia and the Philippines the coefficient on the local return factor falls by an amount between 0.11 and 0.35, while remaining significant. Only for Taiwan, which as noted was relatively unharmed by the crisis, does the effect of the local-return factor rise, though insignificantly. At the same time, the weight of the US market return rises sharply for all six countries, by amounts ranging from 0.38 to 0.72. The weight of the regional factor falls for three of the six countries, becoming insignificant for two of them. The greatest decline in the regional factor is for Taiwan, indicating that during the crisis investors reduced the importance they assigned to regional developments in their day-to-day valuations of the Taiwanese equity market.

For the portfolio of funds, the average local return has a factor weight more than double that of the US market returns before July 1997, with coefficients of 0.76 and 0.34 respectively. During the crisis, both factors are still significant, but their relative weights shift sharply: to 0.51 for the regional average and 0.78 for the S&P 500. This conforms to the picture offered by the individual country returns: during the Asian crisis, the attitudes of US investors towards Asian markets became decoupled from those of local investors, and became more closely tied to patterns of investor sentiment within the US market.

After October 1998, the balance again shifts back to a greater role for the local returns, though the role of the US market factor remains strong. For the six-country portfolio, the local weight rises to 0.75 while the weight on the S&P 500 falls to 0.54. For the six individual-country fund returns, the US factor weight falls in every case, but for five of them (the Philippines is the exception) it stays above its pre-crisis level. The coefficient on the local return rises for all six countries: for four of them, paradoxically, to a level greater than that prevailing before July 1997. It is notable that the two countries, Thailand and Indonesia, where the local return coefficient does not return to its pre-July 1997 level, also witnessed persistently high price/NAV premia after the crisis. This suggests that the high premia correspond to a continuing divergence in sentiment between fund investors in the

US and local investors in Asia, a divergence that diminished sharply in the aftermath of the crisis for the other four countries studied.

V.2 Impact of US sentiment on Asian local returns

Asian local returns can be modeled in a similar way to the US fund returns in the previous section, by modifying equation (1) as follows:

$$
\begin{aligned}
LR_t^i = {} & \beta_0 + \beta_1 FR_{t-1}^i + \beta_2 FRO_{t-1}^i + \beta_3 US_{t-1} + \beta_4 LR_{t-1}^i \\
& + \beta_5 d_{t-1}^{797} + \beta_6 d_{t-1}^{797} FR_{t-1}{}^i + \beta_7 d_{t-1}^{797} FRO_{t-1}^i + \beta_8 d_{t-1}^{797} US_{t-1} + \beta_9 d_{t-1}^{797} LR_{t-1}^i \\
& + \beta_{10} d_{t-1}^{1198} + \beta_{11} d_{t-1}^{1198} FR_{t-1}{}^i + \beta_{12} d_{t-1}^{1198} FRO_{t-1}^i + \beta_{13} d_{t-1}^{1198} US_{t-1} \\
& + \beta_{14} d_{t-1}^{1198} LR_{t-1}^i + \beta_{15} FR_{t-1}^i{}^2 + \eta_t^i \\
\eta_t^i \sim {} & N(0, k_t^i) \\
k_t^i = {} & \delta_0 + \delta_1 \eta_t^i{}^2 + \delta_2 k_{t-1}^i + \delta_3 d_{t-1}^{797} + \delta_4 d_{t-1}^{1198} + \delta_5 FR_{t-1}^i{}^2 + \delta_6 d_{t-1}^{797} FR_{t-1}^i{}^2 + \delta_7 d_{t-1}^{1198} FR_{t-1}^i{}^2
\end{aligned}
\tag{2}
$$

In addition to the terms in equation (1), this equation includes FRO_{t-1}^i, the equally weighted average of the five local returns excluding *i*.

Note that equations (1) and (2), while having several terms in common, do not raise simultaneity issues. The dependent variable in equation (2), LR_t^i, is an independent variable in equation (1), but the dependent variable in equation (1), FR_t^i, is only represented in equation (2) in the form of its lag - which is on the right-hand side of equation (1) as well. Thus, while the inclusion of both the dependent and several of the independent variables from equation (2) on the right-hand side of equation (1) may raise multicollinearity issues, we need not worry about the independence of the disturbance terms in either equation.

Table 7 gives the results of the estimation of equation (2) for each of the six local returns. In the first column, results are presented for the estimation of equation (2) using an equally weighted portfolio of the six local returns as the independent variable, dropping the regional fund-return variable on the right-hand side and using a portfolio of the six fund returns for FR_{t-1}. As before, the lower two panels presents the coefficients for the crisis period (7/97-10/98) and post-crisis period (11/98-12/99), with significance levels derived from F-tests. Graph 5 illustrates the coefficients on the fund returns and the S&P 500.

Before July 1997, the sentiment of fund investors, as represented by the coefficient on the previous day's fund return, is significant at the 10% level for three of the six countries. The magnitude of the effect ranges from 0.08 to 0.13 for the three country returns where it is significant to 0.19 (but not statistically significant) for the portfolio. These effects are consistently smaller than the corresponding effect of local returns on fund returns from Table 6, as one would expect in an environment where most of the news relevant to Asian market returns occurs during Asian trading hours.

During the Asian crisis, this effect increases significantly for two countries (Korea and Indonesia) and for the local return, increases to an insignificant degree for three countries, and falls slightly and insignificantly for Malaysia. It ranges from 0.15 to 0.28 for the four country returns where it is now significant, and reaches 0.45 (and statistically significant) for the six-country portfolio. Thus, at the same time that local returns were becoming less relevant to fund returns, the fund returns tended to become more relevant to the local returns.

In contrast to the increased effect of US closed-fund returns on local markets, the impact of the broader US market return tended to decline during the crisis. Whereas the coefficient on the US_{t-1} variable is statistically significant (at the 10% level) for four of the six country returns, during the crisis it declines for four countries and remains significant for only two. For the regional portfolio, the US factor coefficient falls from 0.17 to 0.12. This suggests that the increased impact of the fund returns on local markets during the crisis reflects the heightened importance of US investors' sentiment *towards those specific markets*, and not merely an increased co-movement of the Asian markets with the US market in general.

After the crisis, the fund-return coefficient tends to fall again. In fact, the fund return is significant and positive for only one local market (the Philippines), compared with three before and four during the crisis; it turns significant and *negative* for Malaysia. For the regional portfolio, the fund-return coefficient remains significant after the crisis, but declines to 0.13.

Meanwhile, the influence of the broad US market return again rises after October 1998 for four of the six countries. It becomes significant at the 5% level for five of them - all except Taiwan, one of the only two where it had been significant *during* the crisis. For the regional portfolio, the coefficient on the S&P 500 return rises to 0.39.

V.3 Summary

A number of stylised facts can be drawn from the results in the previous two sections. During normal times, Asian market returns and closed-end fund returns influence each other, but the effect of the Asian returns on the funds is the greater. During the Asian crisis, the effect of the Asian markets on the funds declines, while that of the funds on the Asian markets increases. Movements in the S&P 500 (a proxy for the US market as a whole) become more important for the fund returns during the Asian crisis, but have a reduced impact on the Asian local markets (even as their indirect impact via the funds increases). After the crisis, the direction of causality again seems to go from the Asian local markets to the country funds, with effects in the reverse direction in most cases even weaker than before. The influence of the broader US market declines, but remains somewhat stronger than it had been before the crisis.

VI. Tracking the direction of influence: daily volatility

The previous section attempted to determine whether and in what ways the flow of “news” between the US and Asian markets changed during the Asian crisis, by looking at changes in market and fund returns. Another form of news, however, is volatility. An increase of price volatility in a given market could indicate a wider divergence of views among investors, increased activity by a previously passive group of investors, or a deterioration in liquidity conditions. When increased volatility in one market return is followed by increased volatility in a related return, after the assumed determinants of the two returns are accounted for, this could indicate that news about participation, liquidity, or changes of opinion in one market is relevant to market values in the other market.

Lagrange multiplier (LM) tests (available from the authors) confirm that, for all of the countries studied, the mean equations in systems (1) and (2) are characterised by autoregressive conditional heteroskedasticity. Using this fact, this section attempts to determine whether volatility is linked across regions - ie whether an unusually volatile day in one market is followed by an unusually volatile day in the other - and whether and how these patterns changed during the crisis period.

The GARCH(1,1) specifications in equations (1) and (2) enable us to ask a number of questions about the volatility of country-fund and local market returns. The coefficients on the time-period dummies in the variance equation of system (1) (γ_3 and γ_4) indicate whether volatility as a whole was higher during those periods compared with January 1995-June 1997. The coefficient on the squared local return (γ_5) in the variance equation indicates whether volatile market returns in Asia are followed by volatile returns on the corresponding country-funds in New York, *irrespective* of its effects on the levels of the returns. The coefficients on the interaction terms (γ_6 and γ_7) indicate whether volatility transfer was accentuated or dampened during these periods. The analogous interpretations hold for the δ coefficients in system (2).

Tables 8 and 9 present the estimated coefficients of the variance equations in (1) and (2). From January 1995 through June 1997, volatility tended to be transmitted strongly from the local markets to the country funds, but not vice versa. The squared local return has a positive, significant effect (at a 10% confidence level) on fund volatility for four of the six countries and for the six-country portfolio (Table 8). This effect changes somewhat for some of the national markets during the crisis and post-crisis periods, but generally not in a strong or consistent way. It rises for three countries (one significantly) after July 1997, then falls for four countries (two significantly, with a significant rise for Malaysia) after November 1998.

In contrast, the transmission of volatility from the funds to the Asian market did seem to change over the course of the sample period (Table 9). The squared fund return has a positive and significant effect on local-return volatility during the pre-crisis period for only one country (Malaysia). During the crisis, this coefficient rises for five of the six countries, and becomes significant for three of them (although, puzzlingly, despite rising it is no longer significant for Malaysia). After the crisis, it declines for four countries, and remains significant and positive for only two of them, while becoming significantly *negative* for Malaysia.

The overall picture that emerges, although clouded by divergent results in certain countries, can be stated as follows. Before July 1997, volatility in local returns was strongly transmitted to country funds, while volatility transmission in the reverse direction was weaker. During the crisis, volatility tended to be transmitted strongly in both directions. After the crisis, volatility transmission in both directions tended to decline, and indeed was lower than before the crisis in most cases. This suggests that, to the extent that country-fund returns reflect the sentiment of US investors towards the Asian markets, this sentiment was much more important for market developments during the crisis than before or after. Local Asian market developments were about equally important for US sentiment regarding Asia before, during and after the crisis period.

There are two notable, and instructive, exceptions to this pattern. For Taiwan, volatility transmission in both directions was insignificant in the pre-crisis period and increased throughout the time under study. The volatility of the Taiwan fund return becomes significant and positive for local-return volatility in the post-crisis period. As already noted, Taiwan escaped the worst effects of the crisis. It is possible that, as western investors sought opportunities in the region in the aftermath of the crisis, Taiwan was seen as an especially promising market, strengthening linkages between western sentiment and local returns.

The other key exception is Malaysia, where the influence of local-return volatility on that of fund returns rose strongly after the crisis, while volatility transmisssion in the opposite direction becomes significantly negative (though small in absolute terms). These findings may reflect the imposition of capital controls in September 1998, as a result of which foreign holders of Malaysian shares were heavily affected by local returns but could do little to influence these returns by reallocating their portfolios. The low coefficient on fund-return volatility for the local return may thus reflect Malaysia's ability to insulate itself from foreign investor sentiment, while the high coefficient on volatility transmission in the opposite direction reflects the continuing interest that US investors had in local developments.

VII. Market sentiment and flows into Asian equities

Changes in sentiment between different classes of investors ought to be associated with portfolio shifts: we would expect investors who have become optimistic about a security's future returns to purchase it from those who are (or have become) pessimistic. Similarly, price changes that are not associated with portfolio shifts might indicate that all classes of investors have experienced a parallel shift in opinion. This section asks whether changes in US investor holdings of Asian equities were associated with differences in sentiment about Asian market prospects, using the local-return/fund-return relationship as an indicator of these differences in sentiment.

US purchases of Asian equities are measured using data from State Street Bank and Trust Co, a large US-based custodian of foreign securities. At August 1998, State Street was estimated to be the custodian for 40% of the securities holdings of US mutual funds (Froot et al, 1998). The data used in this study are net equity purchases and sales settled in the corresponding Asian currency, where the transaction is initiated by non-local investors. Thus, while these figures obviously do not account for all foreign purchases and shares of Asian equities, they are likely to offer a useful indicator of the size and direction of these flows.

Monthly figures were available on net purchases of equities by non-local investors in five of the six countries studied above (all except Taiwan). These were divided by the total capitalisation of the corresponding national stock market to obtain a capital-inflow indicator, summarised in Table 10 and Graph 6. One fact immediately apparent from these data is that they do not correspond neatly to the crisis period - while investors did engage in sustained selling over certain crisis periods from certain countries (for example, sales of Indonesian equities were high in the first months of 1998), one cannot establish a clear link between increased sales of Asian equities by foreign investors and the onset or persistence of crisis conditions in the Asian markets. The volatility of equity flows is generally higher during the crisis period than before or after, though even this is not the case for Indonesia and the Philippines.

To test whether fund inflows affected the pattern of information flows between local and foreign markets, fund returns and local market returns are estimated using the following GARCH specification:

$$
\begin{aligned}
FR_t^i &= \alpha_0 + \alpha_1 LR_t^i + \alpha_2 RR_t^i + \alpha_3 US_t + \alpha_4 FR_{t-1}^i \\
&\quad + \alpha_5 IF_t^i + \alpha_6 IF_t^i LR_t^i + \alpha_7 IF_t^i RR_t^i + \alpha_8 IF_t^i US_t + \alpha_9 IF_t^i FR_{t-1}^i + \alpha_{10} {LR_{t-1}^i}^2 + \varepsilon_t^i \\
\varepsilon_t^i &\sim N(0, h_t^i) \\
h_t^i &= \gamma_0 + \gamma_1 {\varepsilon_t^i}^2 + \gamma_2 h_{t-1}^i + \gamma_3 IF_t^i + \gamma_4 {LR_t^i}^2 + \gamma_5 IF_t^i {LR_t^i}^2
\end{aligned} \tag{3}
$$

and

$$
\begin{aligned}
LR_t^i &= \beta_0 + \beta_1 FR_{t-1}^i + \beta_2 FRO_{t-1}^i + \beta_3 US_{t-1} + \beta_4 LR_{t-1}^i + \beta_5 IF_{t-1}^i \\
&\quad + \beta_6 IF_{t-1}^i FR_{t-1}^i + \beta_7 IF_{t-1}^i FRO_{t-1}^i + \beta_8 IF_{t-1}^i US_{t-1} + \beta_9 IF_{t-1}^i LR_{t-1}^i + \beta_{10} {FR_{t-1}^i}^2 + \eta_t^i \\
\eta_t^i &\sim N(0, k_t^i) \\
k_t^i &= \delta_0 + \delta_1 {\varepsilon_t^i}^2 + \delta_2 h_{t-1}^i + \delta_3 IF_{t-1}^i + \delta_4 {LR_t^i}^2 + \delta_5 IF_{t-1}^i {LR_t^i}^2
\end{aligned} \tag{4}
$$

In both cases, the time dummies from equations (1)-(2) are replaced by the equity inflow indicator, IF_t, for the corresponding month. These specifications allow us to ask, first, whether local returns and fund returns are higher during months when flows into Asian equities are high; second, whether these returns have a greater influence on one another during such months; third, whether the returns are more volatile during such months; and fourth, whether volatility is transmitted more strongly during these months. Since, as noted above, equity inflows were neither notably high nor low during the crisis period, we can be fairly sure that we are not picking up the crisis effects by proxy.

Estimated coefficients for equations (3) and (4) are presented in Tables 11 and 12 respectively, for the five countries for which equity-purchase data were available and, in the first column of each table, for equally-weighted five-country portfolios of the fund returns and local returns.

While the level of the equity inflow variable has no significant effect on the country-fund returns (Table 11), it does have a positive effect on Asian local returns, which is significant at the 10% level or better for three countries and for the five-country portfolio (Table 12). The negative and significant coefficient on the interaction term in Table 11 helps to resolve the puzzle. When local returns are low, a high level of equity inflows corresponds to a high fund return. In other words, fund returns are linked to local returns in normal circumstances, but there are occasions on which local investors are pessimistic (as indicated by a low level of LR) while foreign investors are optimistic (as indicated by a high level of IF), resulting in a relatively higher fund return than one would otherwise have expected (as indicated by a strong negative coefficient on IF*LR).

In local markets, the coefficient on the interaction of equity flows and the fund return tends to be insignificant (Table 12). However, for the five-country portfolio the interaction between equity flows and the S&P 500 return is negative and significant. This suggests that, when foreign investor flows into Asian equities are large, the local markets tend to “decouple” from the US market, while period of low inflows or sales by foreign investors are associated with a strong correlation between local and US market returns.

A similar result holds for the volatilities of the returns. Normally, a volatile local return leads to a volatile fund return, though the reverse effect is less strong. Foreign purchases of Asian equities, by themselves, have little effect on the volatility of either return. However, a high level of foreign investor flows into Asian equities is associated with a lower transmission of volatility from Asia to the United States, for four of the five countries studied and for the five-country portfolio. In other words, excessive foreign investor optimism or pessimism can “override” the normal pattern of linkages between Asian and US markets.

While these effects tend to be significant in statistical terms, the lower two panels of Tables 11 and 12 indicate that they are not very strong in economic terms. When net equity inflows are “shocked” upward by one standard deviation (corresponding to events that, under a normal distribution, should occur about one third of the time), the fraction of the five-country portfolio of local returns incorporated into the five-country portfolio of fund returns falls from 0.63 to 0.58, while the coefficient on local-return volatility in determining fund-return volatility falls from 0.14 to 0.11. The effects in the opposite direction are equally weak: the fund return's impact on local returns falls from 0.19 to 0.16, while the coefficient on fund-return volatility falls from 0.016 to 0.011. In other words, even if our equity-flow indicator reflects genuine shifts in patterns of information flows between local and US equity markets, these shifts tend to be relatively small.

VIII. Conclusion

The results presented in this paper form a more complicated picture than either of the caricatures which have dominated most discussions of the Asian crisis. On the one hand, it is clear that US investors (and, presumably, other investors in developed countries) did not cause the collapse of Asian financial markets by engaging in a massive selloff of Asian securities during the 1997-1998 crisis. Instead, US sentiment as indicated by closed-end fund premia tended to be positive relative to that of Asian investors, both during the period of the crisis and, for some countries, during its aftermath. On a daily level, both the level and volatility of returns on Asia-oriented closed-end country funds tended to be less responsive to local market returns during the crisis. This gap in sentiment was particularly strong during periods when US investors were net purchasers of Asian equities.

At the same time, the gyrations of US sentiment towards Asian markets clearly had an impact on those markets, and this impact was clearly at its strongest during the period of the crisis. Rather than local returns and fund returns influencing one another, during the crisis period the direction of causation clearly ran from the fund returns (both their level and their volatility) towards the local returns. The driving factor here was the sentiment of those US investors oriented towards Asia, rather than the US stock market as a whole: Asian markets, which in non-market times tended to be more or less well-correlated with the S&P 500 index, de-coupled from the broader US market during the crisis period. This decoupling result is supported when equity inflows and outflows, rather than the presence or absence of crisis, are tested for their effects on the relationship between local and fund returns, with a weaker relationship detected in periods when US investor optimism (as proxied by net equity purchases by foreigners) is strong.

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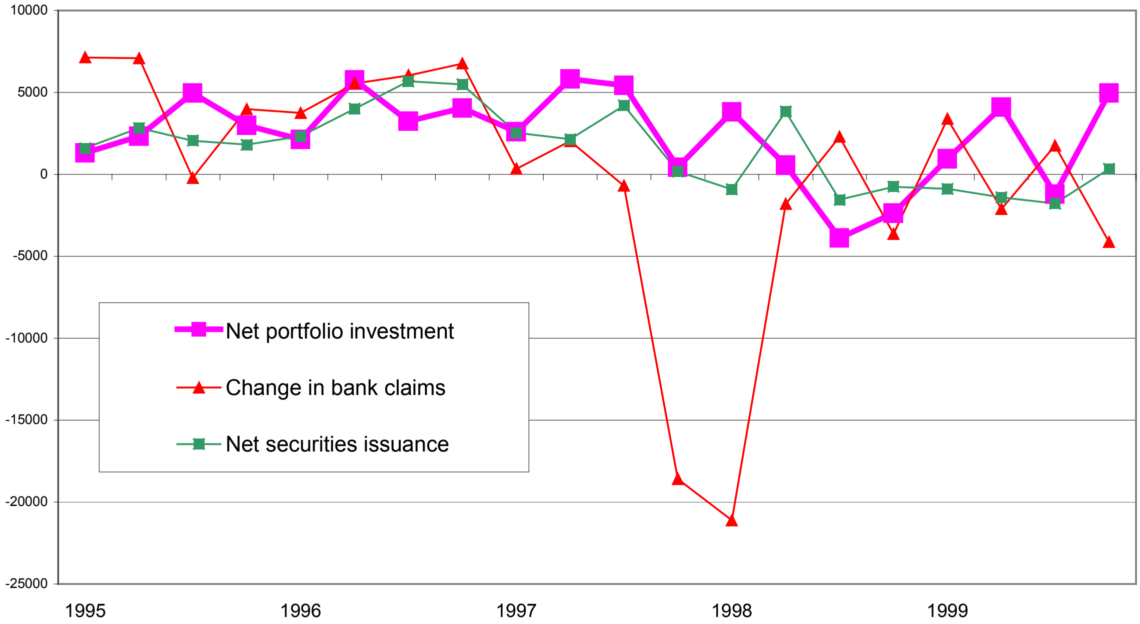
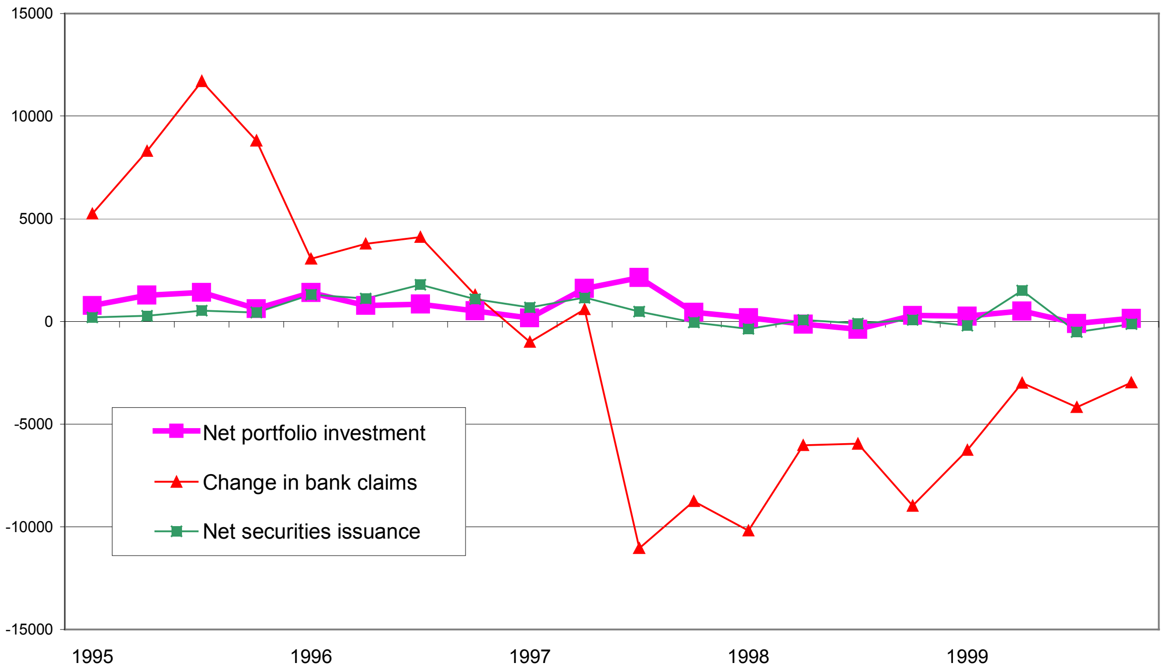
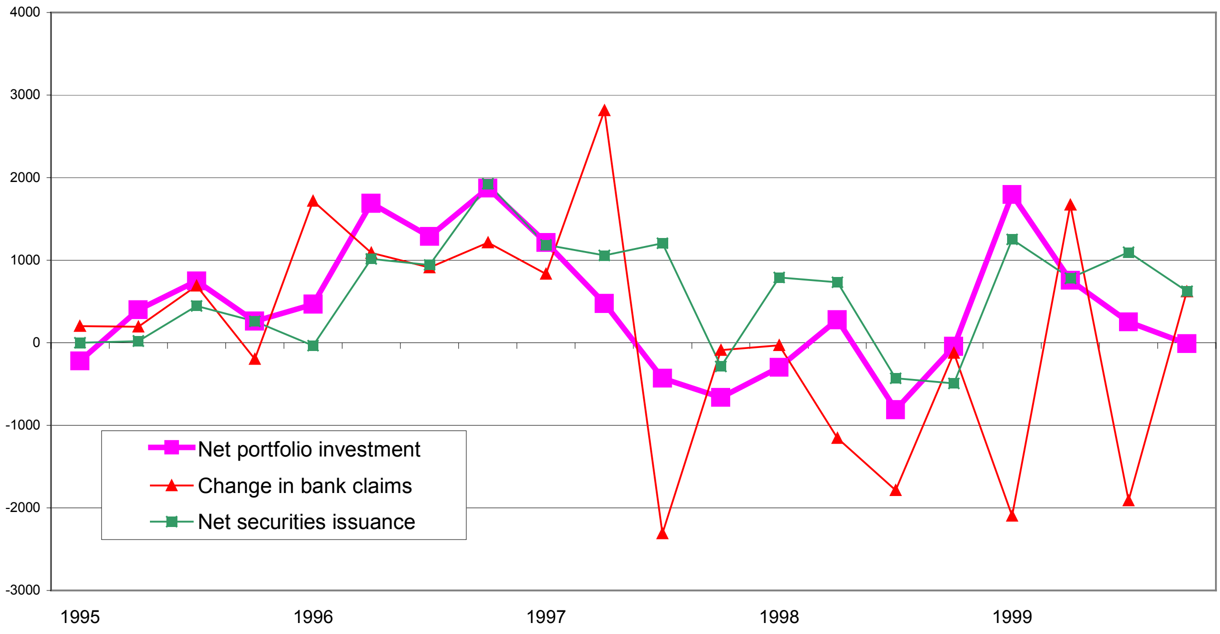
Graph 1a. Components of net portfolio investment (US$M.): Korea



Graph 1b. Components of net portfolio investment (US$M.): Thailand

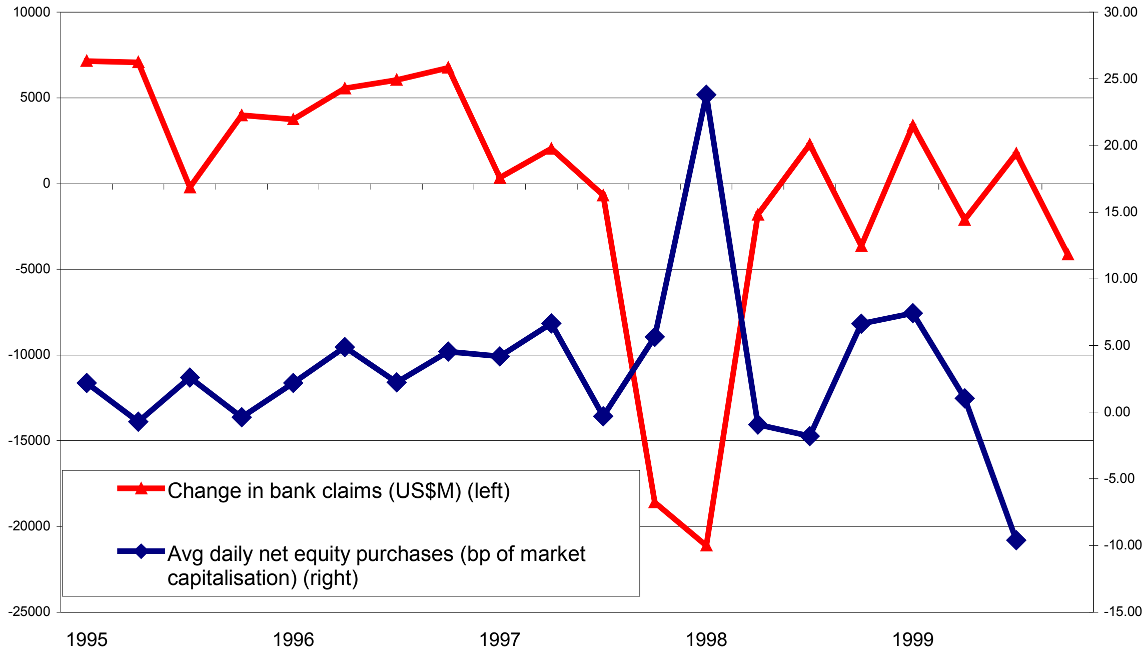


Graph 1c. Components of net portfolio investment (US$M): Philippines

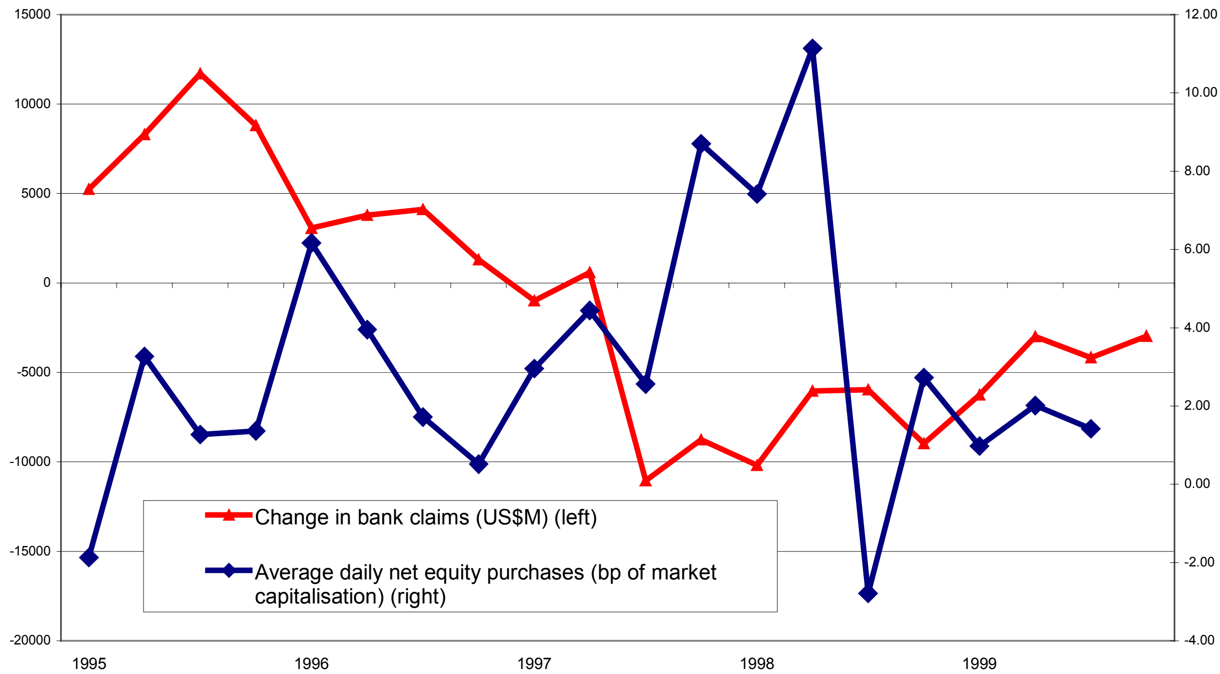


Source (all graphs): IMF, BIS.

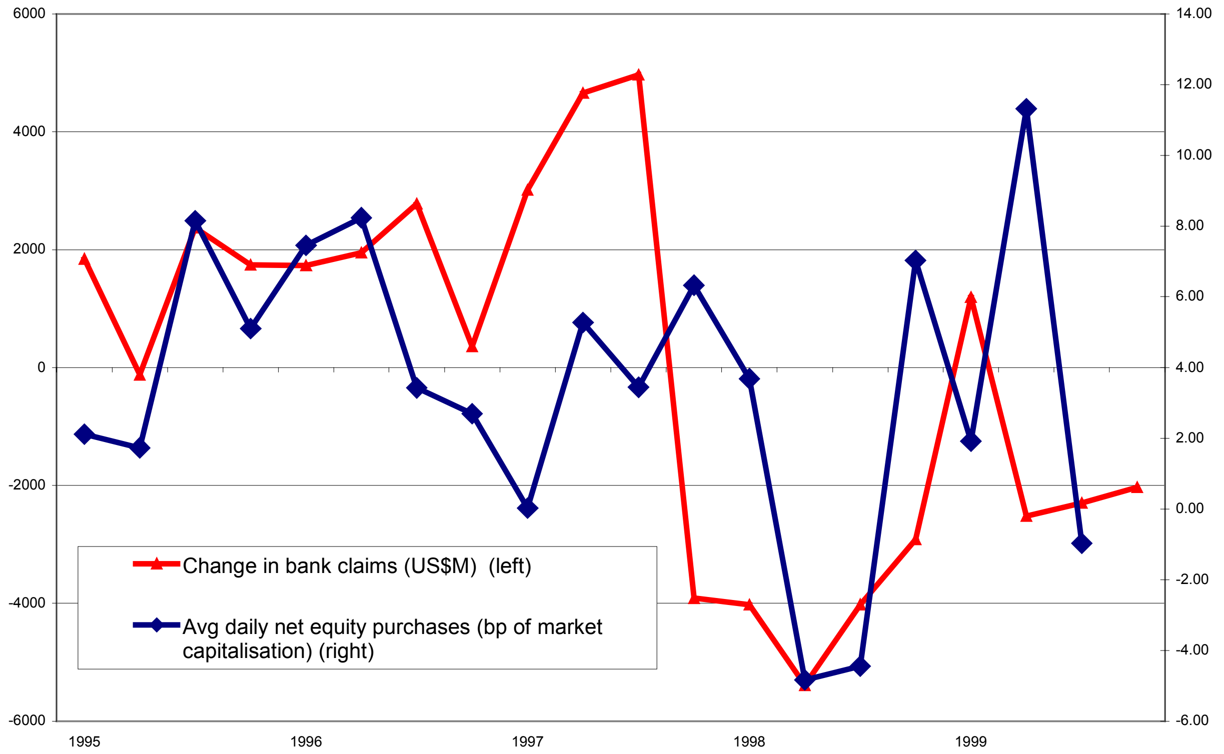
Graph 2a. Banks and equity investors compared: Korea



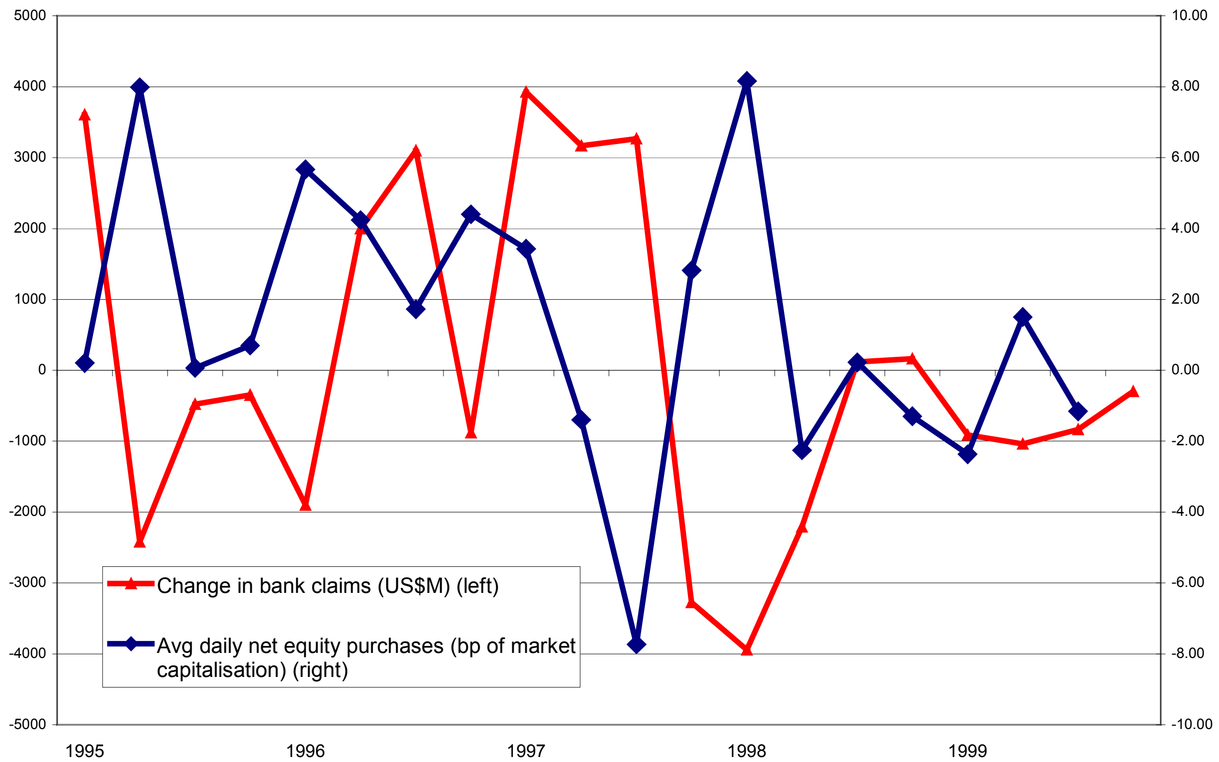
Graph 2b. Banks and equity investors compared: Thailand



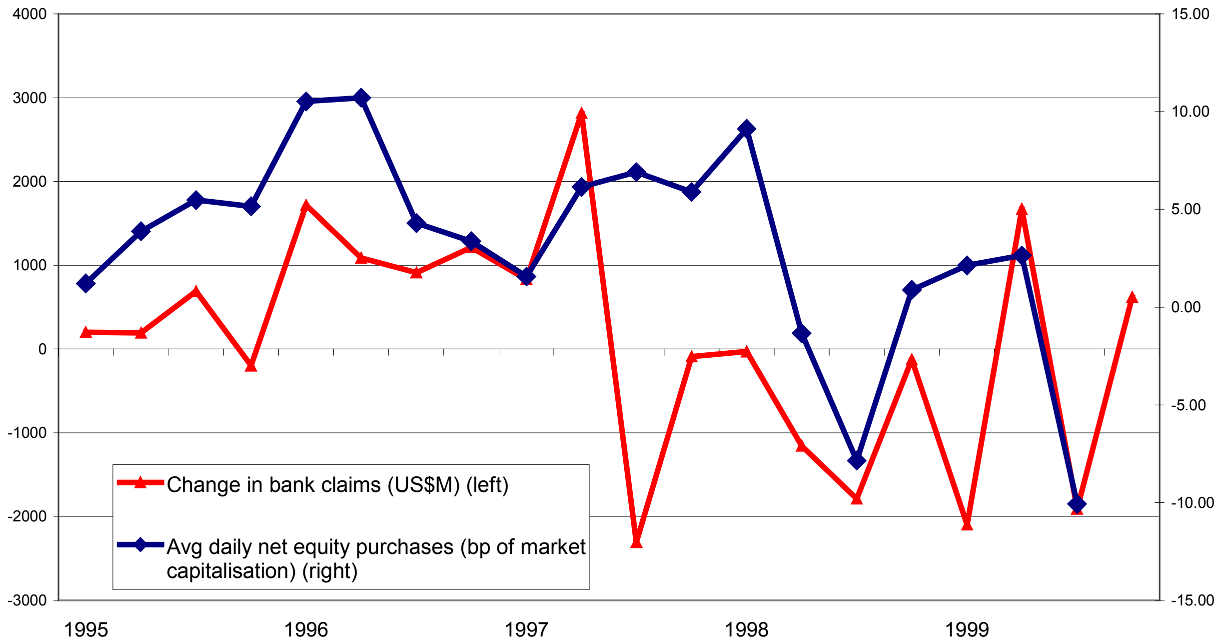
Graph 2c. Banks and equity investors compared: Indonesia



Graph 2d. Banks and equity investors compared: Malaysia

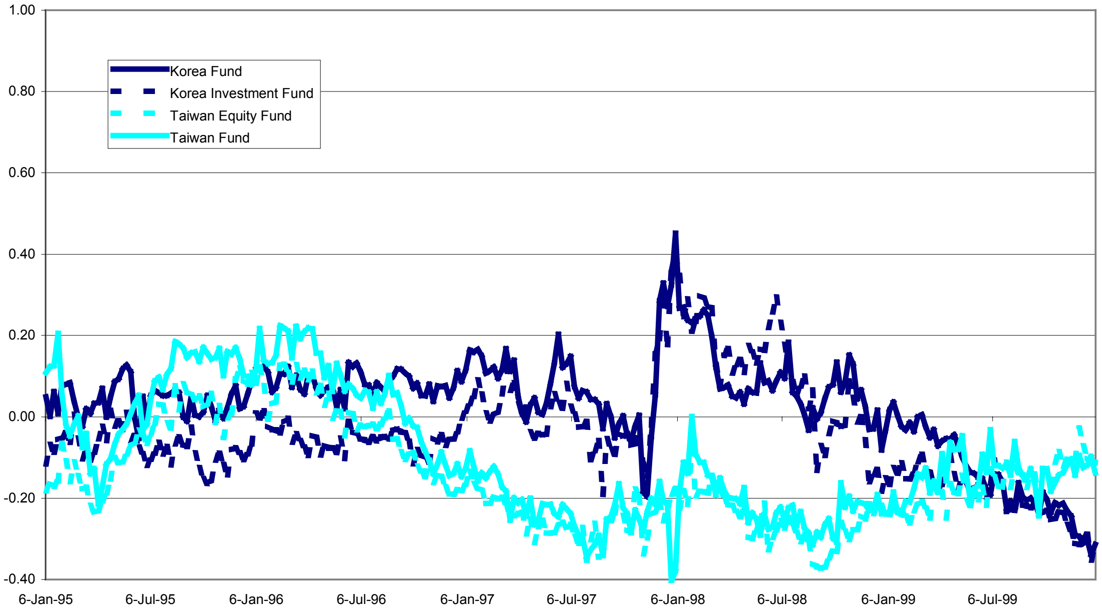


Graph 2e. Banks and equity investors compared: Philippines

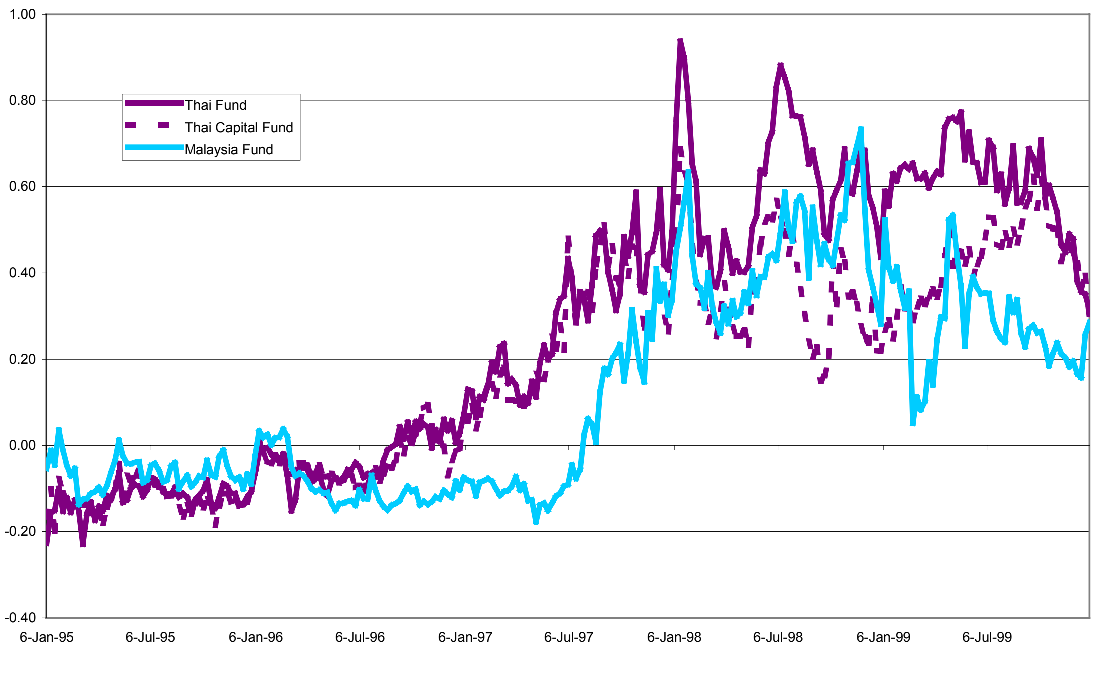


Source (all graphs): BIS, State Street.

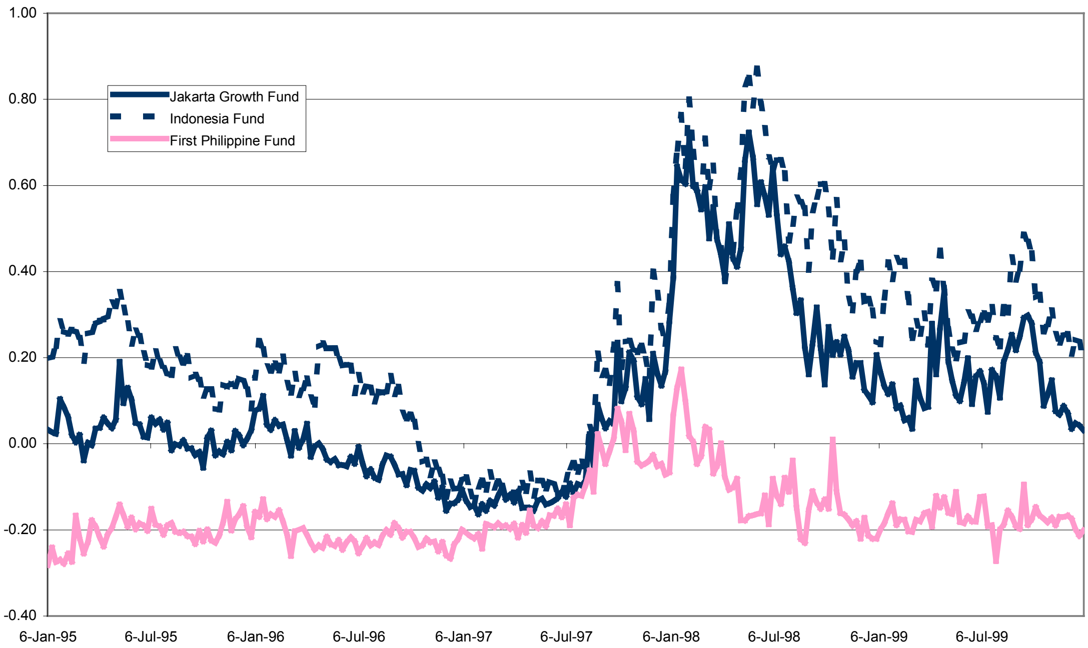
Graph 3a. Premium of price over net asset value



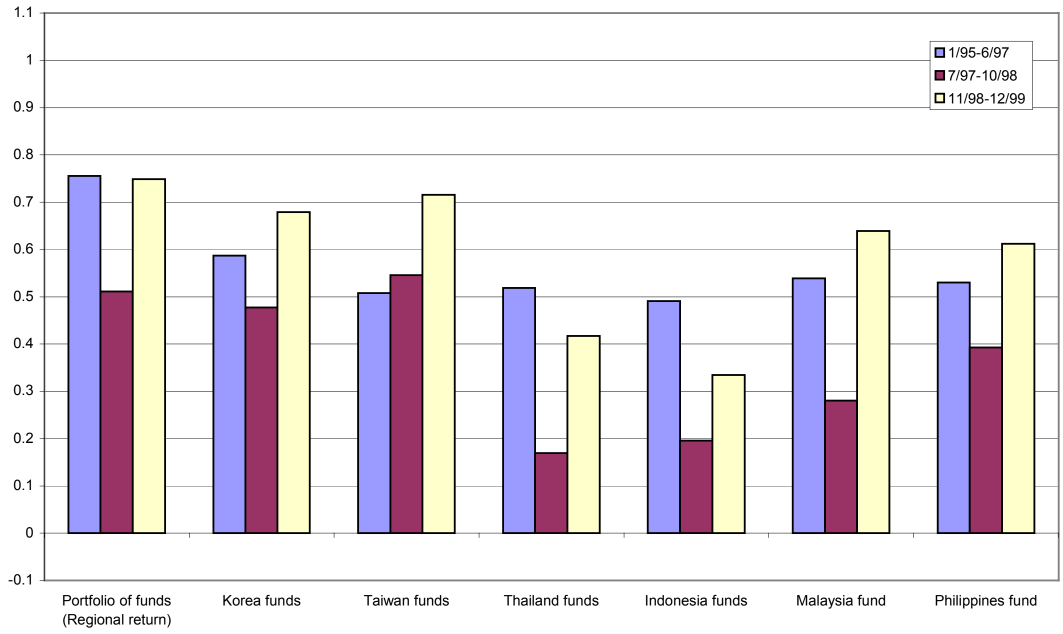
Graph 3b. Premium of price over net asset value (cont.)



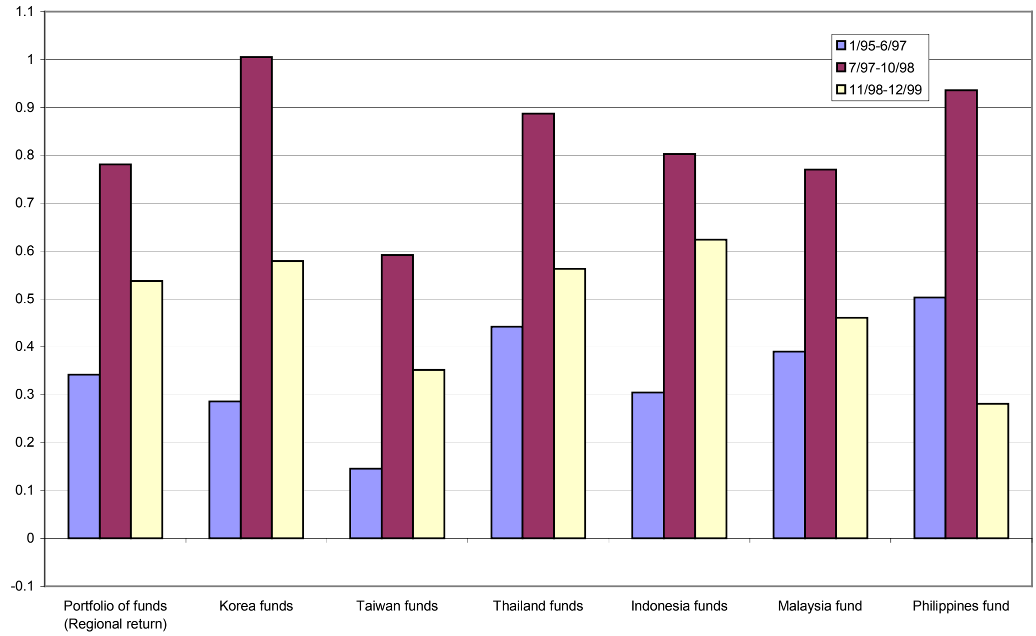
Graph 3c. Premium of price over net asset value (cont.)



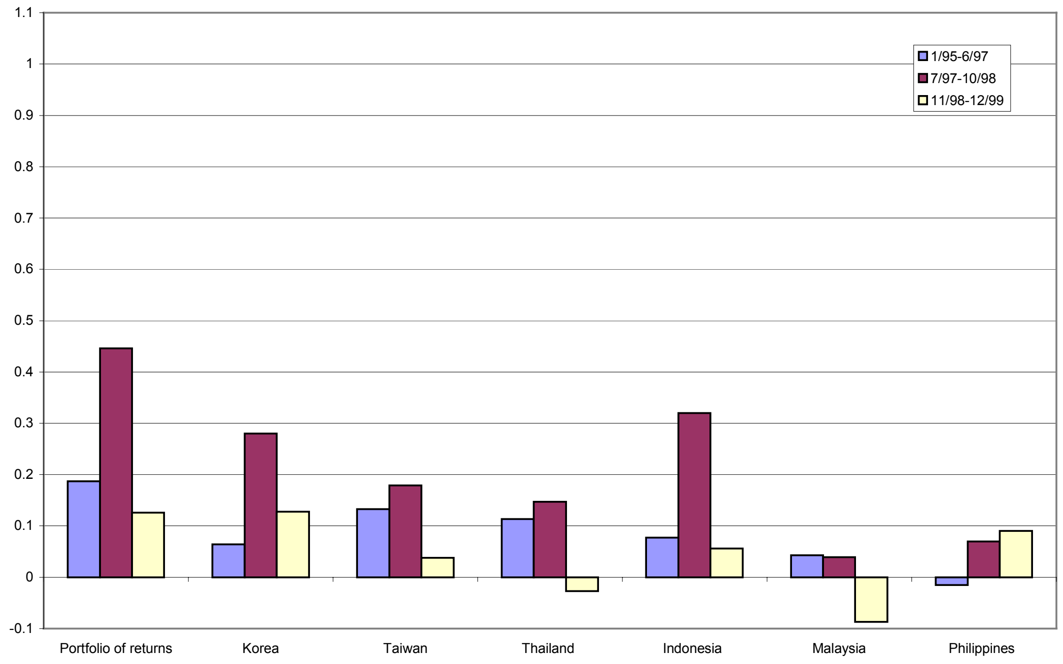
Graph 4a. Effect of local market return on fund return



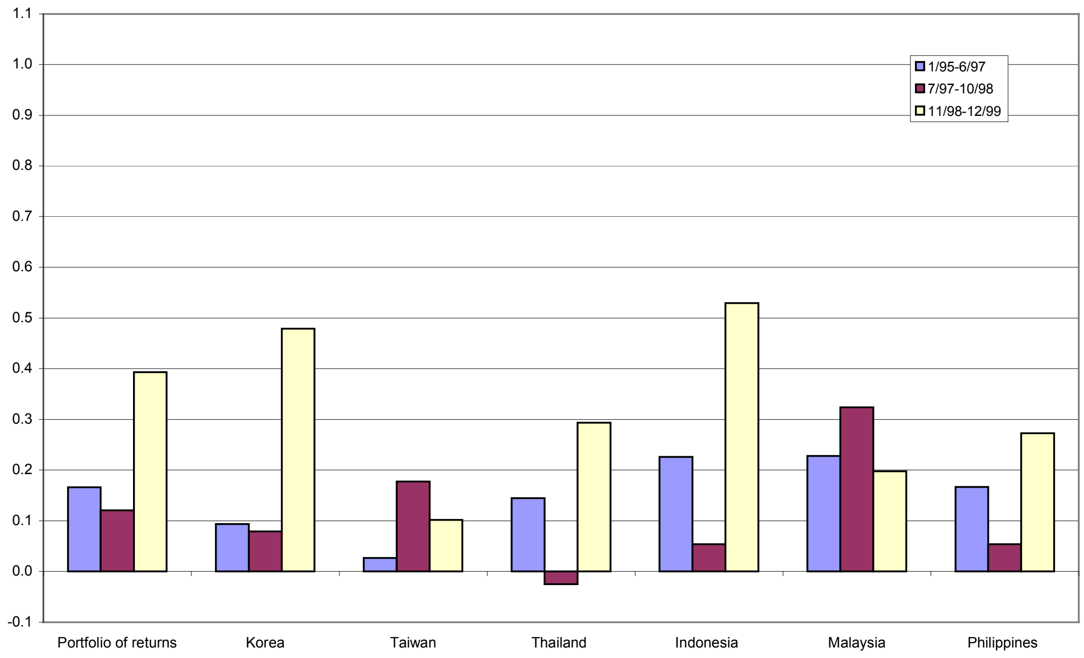
Graph 4b. Effect of S&P 500 return on fund return



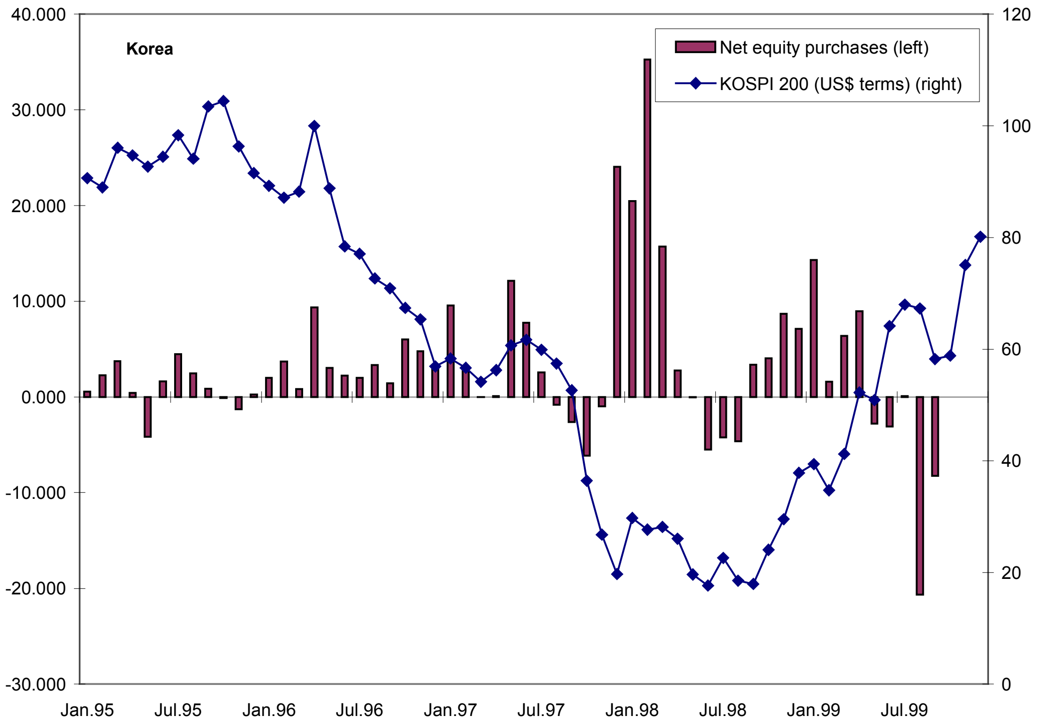
Graph 5a. Effect of previous day's fund return on local market return



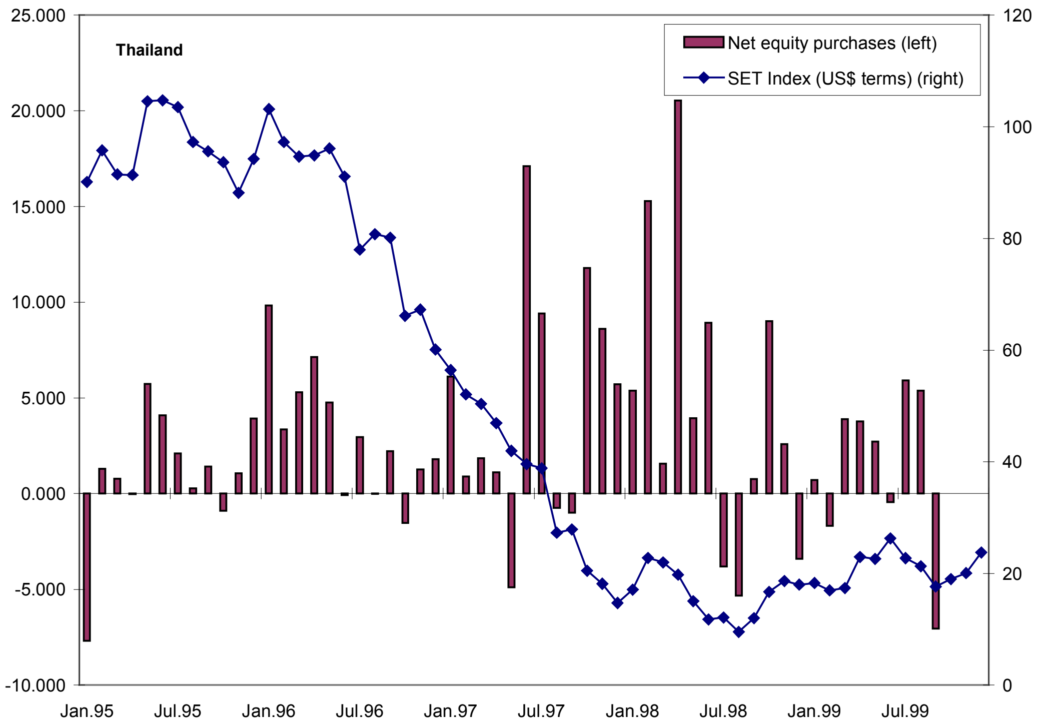
Graph 5b. Effect of previous day's S&P 500 return on local market return



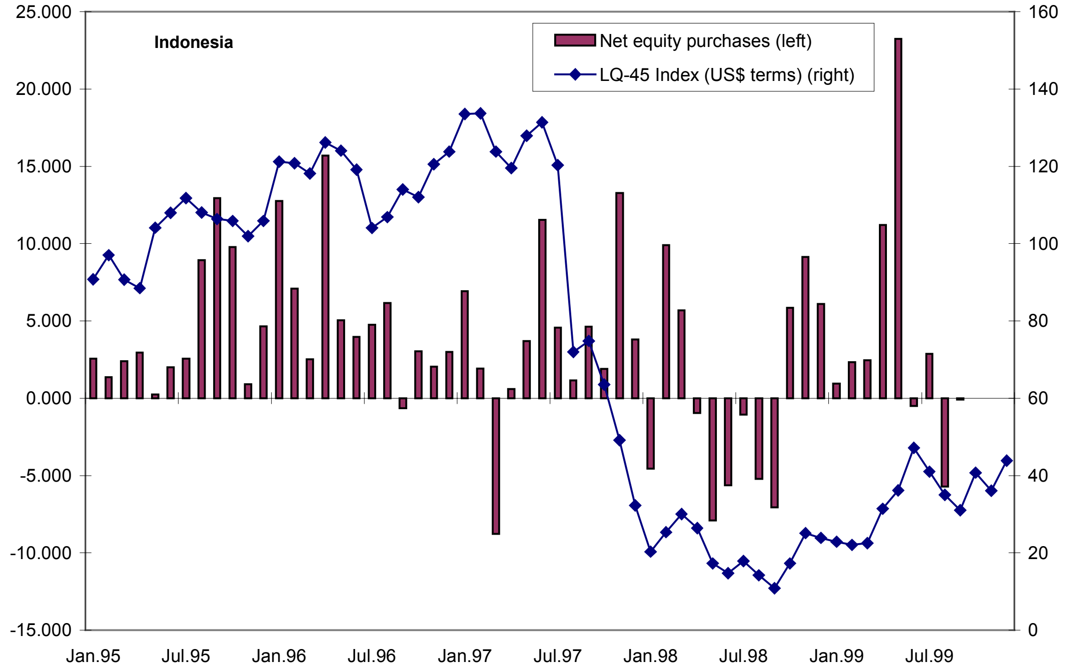
Graph 6. Purchases of equities by foreign investors and local market indices



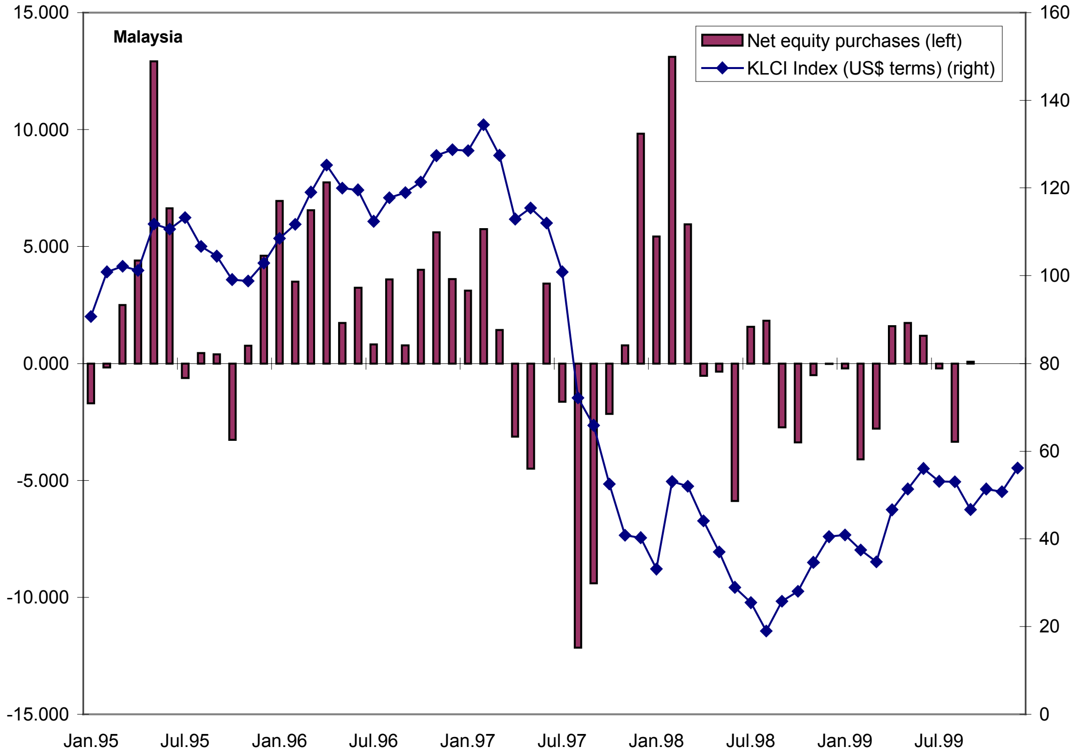
Source: State St., Bloomberg, author calculations.



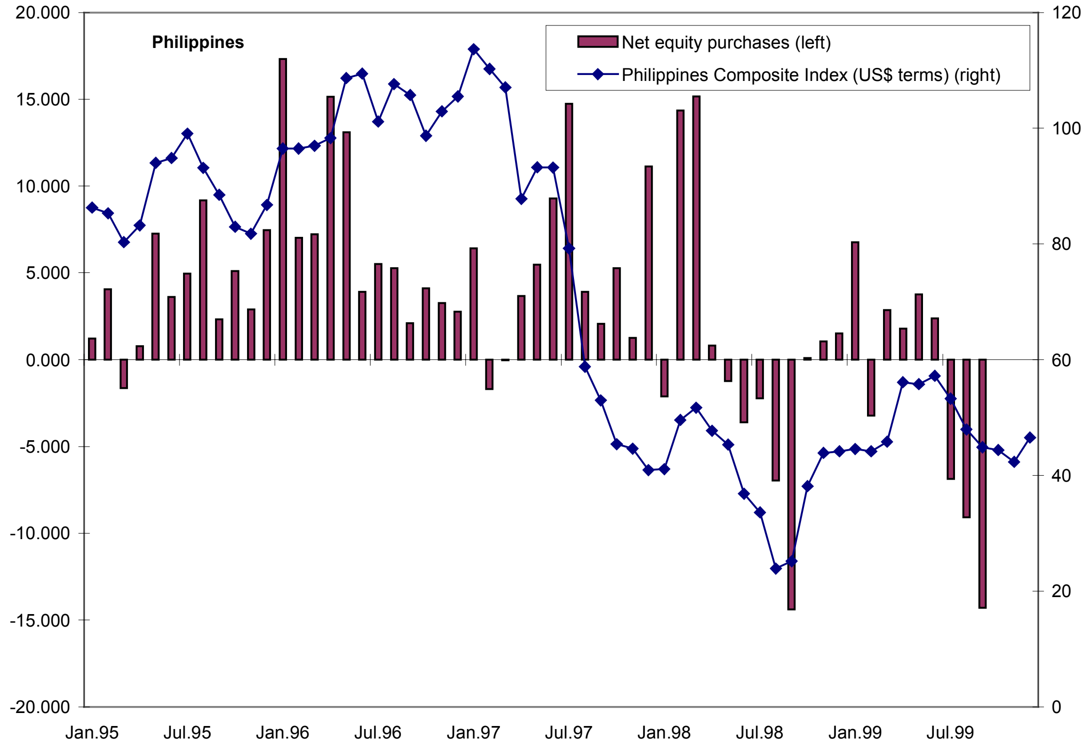
Source: State St., Bloomberg, author calculations.



Source: State St, Bloomberg, author calculations.



Source: State St., Bloomberg, author calculations.



Source: State St., Bloomberg, author calculations.

Table 1
Capital flows to Asian countries
(US$ bn)

	1995	1996	1997	1998	1999
			Indonesia		
Direct investment	3.7	5.6	4.5	– 0.4	– 1.5[1]
Portfolio investment	4.1	5.0	– 2.6	– 1.9	– 1.5[1]
Other investments	2.4	0.2	0.6	– 1.6	0.8[1]
Current account balance	*– 6.4*	*– 7.7*	*– 4.9*	*4.1*	*3.9*[1]
Change in reserves	*– 1.6*	*– 4.5*	*5.1*	*– 2.1*	*– 3.3*[1]
			Korea		
Direct investment	– 1.8	– 2.3	– 1.6	0.6	4.8
Portfolio investment	11.6	15.2	14.3	– 1.9	8.8
Other investments	7.5	11.1	– 10.8	– 2.1	– 12.7
Current account balance	*– 8.5*	*– 23.0*	*– 8.2*	*40.6*	*25.0*
Change in reserves	*– 7.0*	*– 1.4*	*11.9*	*– 31.0*	*– 23.0*
			Malaysia		
Direct investment	4.2	5.1	5.1		
Portfolio investment	– 0.4	– 0.3	– 0.2		
Other investments	3.9	4.7	– 2.1		
Current account balance	*– 8.5*	*– 4.6*	*– 4.8*	*9.4*	*12.5*
Change in reserves	*1.8*	*– 2.5*	*3.9*	*– 10.6*	*– 4.7*
			The Philippines		
Direct investment	1.1	1.3	1.1	1.6	0.7
Portfolio investment	1.2	5.3	0.6	– 0.9	2.8
Other investments	2.7	4.3	5.3	0.9	– 0.4[1]
Current account balance	*– 2.0*	*– 4.0*	*– 4.4*	*1.3*	*7.0*
Change in reserves	*– 0.9*	*– 4.0*	*2.6*	*– 1.9*	*– 4.0*
			Taiwan		
Direct investment	– 1.4	– 2.0	– 3.0	– 3.6	– 1.5
Portfolio investment	0.5	– 1.1	– 8.3	– 2.4	9.1
Other investments	– 7.3	– 5.7	3.1	8.5	5.5
Current account balance	*5.5*	*11.0*	*7.8*	*3.4*	*5.9*
Change in reserves	*3.9*	*– 1.1*	*0.7*	*– 4.8*	*– 18.6*
			Thailand		
Direct investment	1.2	1.4	3.4	6.8	5.3
Portfolio investment	4.1	3.5	4.4	– 0.0	0.8
Other investments	16.6	14.5	– 16.2	– 16.6	– 12.2
Current account balance	*– 13.6*	*– 14.7*	*– 3.0*	*14.0*	*11.0*
Change in reserves	*– 7.2*	*– 2.2*	*9.9*	*– 1.4*	*– 4.6*

[1] Up to third quarter 1999.

Sources: IMF, *Balance of payments*; national data.

Table 2

Asian closed-end country funds

Fund	Date of inception	Market capitalisation ($M.)	Average premium 1995-99	Premium at 31 December 1999
Korea Fund	08/29/84	825	+ 0.03	– 0.31
Korean Investment Fund	02/01/92	71	– 0.04	– 0.31
Taiwan Equity Fund	07/01/94	NA	– 0.15	– 0.11
Taiwan Fund	12/01/86	NA	– 0.09	– 0.14
Thai Fund	02/01/88	97	+ 0.27	+ 0.31
Thai Capital Fund	05/01/90	NA	+ 0.18	+ 0.30
Jakarta Growth Fund	04/01/90	16	+ 0.10	+ 0.03
Indonesia Fund	03/09/90	24	+ 0.24	+ 0.19
Malaysia Fund	05/01/87	70	+ 0.12	+ 0.28
First Philippine Fund	11/01/89	56	– 0.16	– 0.20
Average		**166**	**+ 0.05**	**+ 0.00**

Source: Bloomberg; author's calculations. Market capitalisations are as of year-end 1999 and are not reported by the Taiwan Fund, Taiwan Equity Fund and Thai Capital Fund. The premium, measured weekly, is defined as the natural logarithm of the ratio of price to net asset value.

Table 3
Correlations of fund premia

Correlations of levels

	Korean Inv. Fund	Taiwan Equity Fund	Taiwan Fund	Thai Fund	Thai Capital Fund	Jakarta Growth Fund	Indon. Fund	Malaysia Fund	First Phil. Fund
Korea Fund	0.81	0.02	0.09	– 0.25	– 0.32	0.06	0.00	– 0.15	0.18
Korean Investment Fund		– 0.25	– 0.22	0.11	0.07	0.45	0.32	0.18	0.43
Taiwan Equity Fund			0.94	– 0.66	– 0.62	– 0.30	– 0.24	– 0.51	– 0.36
Taiwan Fund				– 0.70	– 0.69	– 0.31	– 0.23	– 0.53	– 0.40
Thai Fund					0.94	0.63	0.57	0.87	0.52
Thai Capital Fund						0.57	0.47	0.76	0.57
Jakarta Growth Fund							0.93	0.78	0.56
Indonesia Fund								0.77	0.43
Malaysia Fund									0.52

Correlations of weekly changes

	Korean Inv. Fund	Taiwan Equity Fund	Taiwan Fund	Thai Fund	Thai Capital Fund	Jakarta Growth Fund	Indon. Fund	Malaysia Fund	First Phil. Fund
Korea Fund	0.48	0.12	0.02	0.17	0.19	0.13	0.09	0.12	– 0.02
Korean Investment Fund		0.02	– 0.08	0.10	0.04	– 0.04	0.01	– 0.09	– 0.05
Taiwan Equity Fund			0.57	0.08	0.10	0.08	0.02	0.19	0.11
Taiwan Fund				0.14	0.13	0.10	0.06	0.24	0.12
Thai Fund					0.52	0.14	0.08	0.27	0.24
Thai Capital Fund						0.19	0.04	0.21	0.15
Jakarta Growth Fund							0.54	0.23	0.15
Indonesia Fund								0.19	0.08
Malaysia Fund									0.24

Table 4

Average premia over selected time periods

Fund	Jan 1995 - Jun 1997	July 1997 - Oct 1998	Nov 1998 - Dec 1999
Korea Fund	+ 0.069 (0.004)	+ 0.088** (0.014)	– 0.122** (0.014)
Korean Investment Fund	– 0.046 (0.005)	+ 0.085** (0.017)	– 0.179** (0.010)
Taiwan Equity Fund	– 0.074 (0.011)	– 0.258** (0.007)	– 0.178** (0.007)
Taiwan Fund	+ 0.008 (0.012)	– 0.233** (0.008)	– 0.153** (0.007)
Thai Fund	– 0.027 (0.010)	+ 0.539** (0.020)	+ 0.602** (0.013)
Thai Capital Fund	– 0.036 (0.009)	+ 0.379** (0.013)	+ 0.409** (0.013)
Jakarta Growth Fund	– 0.031 (0.007)	+ 0.305** (0.028)	+ 0.149** (0.010)
Indonesia Fund	+ 0.104 (0.012)	+ 0.433** (0.031)	+ 0.317** (0.010)
Malaysia Fund	– 0.083 (0.004)	+ 0.328** (0.021)	+ 0.317 (0.019)
First Philippine Fund	– 0.207 (0.003)	– 0.069** (0.010)	– 0.176** (0.004)
Average	**– 0.03**	**+ 0.16**	**+ 0.10**

Standard errors in parentheses.

** A t-test for the equivalence of means rejects equality between the fund's average premium over the period and its average premium over the immediately preceding period with 95% confidence.

Source: Bloomberg; author's calculations.

Table 5

Local stock markets: capitalisation data and indices used in daily regressions

Country	Market capitalisation (US$bn)		Index	Number of index members
	End-1994	End-1999		
Korea	192	309	KOSPI 200 Index	200
Taiwan	247	376	TWSE Weighted Index	452
Thailand	131	58	Bangkok SET Index	394
Indonesia	47	64	Jakarta LQ-45 Index	45
Malaysia	199	145	KL Composite Index	100
Philippines	56	48	Philippines Composite Index	33

Source: IFC; Bloomberg.

Table 6

Does Asian local market sentiment drive fund prices in New York?

The first panel of this table reports slope coefficients and z-statistics (using Bollerslev-Wooldridge QML standard errors) for the mean equation in the following GARCH system:

$$\begin{aligned} FR_t^i &= \alpha_0 + \alpha_1 LR_t^i + \alpha_2 RR_t^i + \alpha_3 US_t + \alpha_4 FR_{t-1}^i \\ &\quad + \alpha_5 d_t^{797} + \alpha_6 d_t^{797} LR_t^i + \alpha_7 d_t^{797} RR_t^i + \alpha_8 d_t^{797} US_t + \alpha_9 d_t^{797} FR_{t-1}^i \\ &\quad + \alpha_{10} d_t^{1198} + \alpha_{11} d_t^{1198} LR_t^i + \alpha_{12} d_t^{1198} RR_t^i + \alpha_{13} d_t^{1198} US_t + \alpha_{14} d_t^{1198} FR_{t-1}^i \\ &\quad + \alpha_{15} LR_t^{i2} + \varepsilon_t^i \\ \varepsilon_t^i &\sim N(0, h_t^i) \\ h_t^i &= \gamma_0 + \gamma_1 \varepsilon_t^{i2} + \gamma_2 h_{t-1}^i + \gamma_3 d_t^{797} + \gamma_4 d_t^{1198} + \gamma_5 LR_t^{i2} + \gamma_6 d_t^{797} LR_t^{i2} + \gamma_7 d_t^{797} LR_t^{i2} \end{aligned}$$

In the first column, the "portfolio of funds" is an equally weighted average of the country-fund returns; it is regressed on a constant, the US market return, an equally weighted portfolio of local market returns, its own lag, and interactions of these variables with the crisis dummy.

Constant term and coefficients on the lagged dependent variable are not reported. Figures followed by an asterisk (*) are significant at the 10% level; those followed by a double asterisk (**) at the 5% level. The number of observations and the adjusted R^2 of each regression are also reported.

The second panel reports the totals of the coefficients on the local, regional and US returns during the period when d_t^{797} equals 1 (ie $\alpha_1+\alpha_6$, $\alpha_2+\alpha_7$, and $\alpha_3+\alpha_8$). The third panel reports the totals for when d_t^{1198} equals 1 (ie $\alpha_1+\alpha_6+\alpha_{11}$, etc.). Significance levels are based on the F-statistic from a Wald test for the sum of the coefficients equalling zero.

Explanatory variable	Portfolio of funds	Korea funds	Taiwan funds	Thailand funds	Indonesia funds	Malaysia fund	Philippines fund
d_t^{797}	– 0.002* (– 1.81)	– 0.002 (– 1.57)	– 0.001 (– 0.49)	– 0.001 (– 0.85)	– 0.003* (– 1.79)	– 0.002 (– 1.24)	– 0.002 (– 1.34)
d_t^{1198}	0.001 (1.22)	0.003* (1.81)	0.002 (1.29)	0.000 (0.14)	0.002 (0.82)	0.001 (0.29)	0.002 (1.12)
LR_t^i	0.756** (15.72)	0.587** (14.06)	0.508** (12.76)	0.519** (13.46)	0.491** (8.86)	0.539** (9.01)	0.530** (10.76)
$d_t^{797}LR_t^i$	– 0.246** (– 3.57)	– 0.110* (– 1.91)	0.038 (0.55)	– 0.350** (– 3.53)	– 0.295** (– 4.41)	– 0.259** (– 3.19)	– 0.137* (– 1.67)
$d_t^{1198}LR_t^i$	0.238** (2.74)	0.203** (3.96)	0.170** (2.09)	0.248** (2.32)	0.139** (2.27)	0.359** (2.78)	0.219** (2.12)
RR_t^i		0.194** (2.51)	0.322** (2.85)	0.246** (2.84)	0.187** (2.51)	0.230** (2.44)	0.256** (3.41)
$d_t^{797}RR_t^i$		– 0.126 (– 1.33)	– 0.282** (– 2.33)	0.006 (0.05)	0.223** (2.01)	0.026 (0.22)	– 0.148* (– 1.65)
$d_t^{1198}RR_t^i$		0.090 (1.00)	0.066 (0.76)	– 0.113 (– 0.79)	0.069 (0.41)	0.243** (1.97)	0.110 (1.05)
US_t	0.342** (12.33)	0.286** (4.07)	0.146* (1.90)	0.442** (6.86)	0.305** (3.78)	0.390** (5.68)	0.503** (5.42)
$d_t^{797}US_t$	0.439** (6.40)	0.719** (4.93)	0.447** (4.44)	0.445** (2.43)	0.497** (2.99)	0.380** (2.75)	0.433** (2.67)
$d_t^{1198}US_t$	– 0.242** (– 2.80)	– 0.427** (– 2.84)	– 0.240** (– 2.18)	– 0.324 (– 1.59)	– 0.178 (– 0.83)	– 0.309* (– 1.66)	– 0.655** (– 3.99)
N	1302	1302	1302	1302	1302	1302	1302
Adj R^2	0.31	0.49	0.35	0.29	0.30	0.26	0.31

Table 6 (contd)

Effects during 07/97 - 10/98:							
LR_t^i	0.511**	0.477**	0.546**	0.169*	0.196**	0.280**	0.393**
RR_t^i		0.068	0.040	0.252**	0.410**	0.256**	0.108**
US_t	0.781**	1.005**	0.592**	0.887**	0.803**	0.770**	0.936**
Effects during 11/98 - 12/99:							
LR_t^i	0.749**	0.679**	0.716**	0.417**	0.335**	0.639**	0.612**
RR_t^i		0.158**	0.106	0.139	0.479**	0.499**	0.218**
US_t	0.538**	0.579**	0.352**	0.563**	0.624**	0.461**	0.281**

Table 7

Does New York market sentiment drive Asian local returns?

The first panel of this table reports slope coefficients and z-statistics (using Bollerslev-Wooldridge QML standard errors) for the mean equation in the following system:

$$LR_t^i = \beta_0 + \beta_1 FR_{t-1}^i + \beta_2 FRO_{t-1}^i + \beta_3 US_{t-1} + \beta_4 LR_{t-1}^i$$
$$+ \beta_5 d_{t-1}^{797} + \beta_6 d_{t-1}^{797} FR_{t-1}^i + \beta_7 d_{t-1}^{797} FRO_{t-1}^i + \beta_8 d_{t-1}^{797} US_{t-1} + \beta_9 d_{t-1}^{797} LR_{t-1}^i$$
$$+ \beta_{10} d_{t-1}^{1198} + \beta_{11} d_{t-1}^{1198} FR_{t-1}^i + \beta_{12} d_{t-1}^{1198} FRO_{t-1}^i + \beta_{13} d_{t-1}^{1198} US_{t-1} + \beta_{14} d_{t-1}^{1198} LR_{t-1}^i$$
$$+ \beta_{15} {FR_{t-1}^i}^2 + \eta_t^i$$
$$\eta_t^i \sim N(0, k_t^i)$$
$$k_t^i = \delta_0 + \delta_1 {\eta_t^i}^2 + \delta_2 k_{t-1}^i + \delta_3 d_{t-1}^{797} + \delta_4 d_{t-1}^{1198} + \delta_5 {FR_{t-1}^i}^2 + \delta_6 d_{t-1}^{797} {FR_{t-1}^i}^2 + \delta_7 d_{t-1}^{1198} {FR_{t-1}^i}^2$$

In the first column, the "portfolio of local returns" is an equally weighted average of the local dollar returns; it is regressed on a constant, the US market return, an equally weighted portfolio of fund returns, its own lag, and interactions of these variables with the crisis dummy.

Constant term and coefficients on the lagged dependent variable are not reported. Figures followed by an asterisk (*) are significant at the 10% level; those followed by a double asterisk (**) at the 5% level. The number of observations and the adjusted R^2 of each regression are also reported.

The second panel reports the totals of the coefficients on the country-fund nd US returns during the period when d_t^{797} equals 1 (ie $\beta_1+\beta_6$, $\beta_2+\beta_7$, and $\beta_3+\beta_8$). The third panel reports the totals for when d_t^{1198} equals 1 (ie $\beta_1+\beta_6+\beta_{11}$, etc.). Significance levels are based on the F-statistic from a Wald test for the sum of the coefficients equalling zero.

Explanatory variable	Portfolio of returns	Korea local return	Taiwan local return	Thailand local return	Indonesia local return	Malaysia local return	Philippines local return
d_{t-1}^{797}	– 0.002 (– 1.51)	– 0.001 (– 0.55)	– 0.002** (– 2.00)	– 0.002 (– 1.26)	– 0.003 (– 1.33)	– 0.006** (– 3.29)	– 0.003* (– 1.70)
d_{t-1}^{1198}	0.003** (3.40)	0.004** (1.99)	0.002 (1.53)	0.004* (1.65)	0.004 (1.29)	0.006** (3.33)	0.002 (1.21)
FR_{t-1}^i	0.187 (1.59)	0.064 (1.31)	0.133** (3.05)	0.113** (2.30)	0.077** (2.30)	0.043 (1.58)	– 0.015 (– 0.37)
$d_{t-1}^{797}FR_{t-1}^i$	0.259** (2.09)	0.215** (2.57)	0.046 (0.57)	0.033 (0.40)	0.243** (2.45)	– 0.004 (– 0.04)	0.085 (1.23)
$d_{t-1}^{1198}FR_{t-1}^i$	– 0.320** (– 7.28)	– 0.152 (– 1.17)	– 0.141 (– 1.34)	– 0.173* (– 1.71)	– 0.264** (– 2.35)	– 0.126 (– 1.34)	0.020 (0.31)
FRO_{t-1}^i		– 0.066 (– 1.16)	0.024 (0.41)	0.167** (2.50)	0.103** (2.37)	0.091* (1.81)	0.265** (5.54)
$d_{t-1}^{797}FRO_{t-1}^i$		0.177 (1.53)	0.062 (0.82)	0.238** (1.99)	0.335 (1.37)	0.176 (1.53)	0.099 (1.12)
$d_{t-1}^{1198}FRO_{t-1}^i$		0.069 (0.47)	0.019 (0.28)	– 0.092 (– 0.66)	– 0.349 (– 1.24)	– 0.230* (– 1.80)	– 0.166* (– 1.80)
US_{t-1}	0.166 (1.20)	0.094 (1.24)	0.027 (0.39)	0.145* (1.79)	0.226** (4.47)	0.228** (4.50)	0.167** (3.03)
$d_{t-1}^{797}US_{t-1}$	– 0.045 (– 0.28)	– 0.015 (– 0.10)	0.151 (1.35)	– 0.170 (– 0.90)	– 0.172 (– 0.55)	0.097 (0.53)	– 0.113 (– 0.73)
$d_{t-1}^{1198}US_{t-1}$	0.271** (3.34)	0.401* (1.78)	– 0.076 (– 0.67)	0.319 (1.54)	0.476 (1.36)	– 0.126 (– 0.64)	0.219 (1.41)
N	1302	1302	1302	1302	1302	1302	1302
Adj R^2	0.27	0.31	0.07	0.15	0.09	0.13	0.22

Table 7 (contd)

Effects during 7/97 - 10/98:							
FR_{t-1}^{I}	0.446**	0.280**	0.179**	0.147**	0.320**	0.039	0.070
FRO_{t-1}^{I}		0.111	0.086*	0.405**	0.438*	0.267**	0.365**
US_{t-1}	0.121	0.079	0.178**	– 0.025	0.054	0.324*	0.054
Effects during 11/98 - 12/99:							
FR_{t-1}^{I}	0.126**	0.128	0.038	– 0.027	0.056	– 0.087**	0.090**
FRO_{t-1}^{I}		0.180	0.105**	0.313**	0.089	0.037	0.199**
US_{t-1}	0.393**	0.479**	0.102	0.294**	0.530**	0.198**	0.273**

Table 8

Does the volatility of the Asian local market drive the volatility of fund prices in New York?

The first panel of this table reports slope coefficients and z-statistics (using Bollerslev-Wooldridge QML standard errors) for the variance equation from the following system:

$$
\begin{aligned}
FR_t^i &= \alpha_0 + \alpha_1 LR_t^i + \alpha_2 RR_t^i + \alpha_3 US_t + \alpha_4 FR_{t-1}^i \\
&\quad + \alpha_5 d_t^{797} + \alpha_6 d_t^{797} LR_t^i + \alpha_7 d_t^{797} RR_t^i + \alpha_8 d_t^{797} US_t + \alpha_9 d_t^{797} FR_{t-1}^i \\
&\quad + \alpha_{10} d_t^{1198} + \alpha_{11} d_t^{1198} LR_t^i + \alpha_{12} d_t^{1198} RR_t^i + \alpha_{13} d_t^{1198} US_t + \alpha_{14} d_t^{1198} FR_{t-1}^i \\
&\quad + \alpha_{15} LR_t^{i\,2} + \varepsilon_t^i \\
\varepsilon_t^i &\sim N(0, h_t^i) \\
h_t^i &= \gamma_0 + \gamma_1 \varepsilon_t^{i\,2} + \gamma_2 h_{t-1}^i + \gamma_3 d_t^{797} + \gamma_4 d_t^{1198} + \gamma_5 LR_t^{i\,2} + \gamma_6 d_t^{797} LR_t^{i\,2} + \gamma_7 d_t^{797} LR_t^{i\,2}
\end{aligned}
$$

In the first column, the “portfolio of funds” is an equally weighted average of the country-fund returns; it is regressed on a constant, the US market return, an equally weighted portfolio of local market returns, its own lag, and interactions of these variables with the time dummies.

Constant term and coefficients on the lagged dependent variable are not reported. Figures followed by an asterisk (*) are significant at the 10% level; those followed by a double asterisk (**) at the 5% level. The number of observations and the adjusted R^2 of each regression are also reported.

The second and third panels report the totals of the coefficients on the squared local return during the periods when d_t^{797} equals 1 (ie $\gamma_5+\gamma_6$) and when d_t^{1198} equals 1 (ie $\gamma_5+\gamma_6+\gamma_7$.). Significance levels are based on the F-statistic from a Wald test for the sum of the coefficients equalling zero.

Explanatory variable	Portfolio of funds	Korea funds	Taiwan funds	Thailand funds	Indonesia funds	Malaysia fund	Philippines fund
$\hat{\varepsilon}_t^{i\,2}$	0.106** (4.02)	0.043** (2.32)	0.203** (3.73)	0.077** (2.67)	0.131** (4.21)	0.113** (2.34)	0.103** (2.88)
h_{t-1}^i	0.542** (11.77)	0.776** (13.63)	0.502** (7.08)	0.666** (8.01)	0.717** (13.84)	0.396** (3.92)	0.596** (7.56)
d_t^{797}	0.00002** (2.47)	0.00004** (2.18)	0.00002 (0.94)	0.00007 (1.46)	0.00007** (2.18)	0.00019** (2.78)	– 0.00001 (– 0.83)
d_t^{1198}	– 0.00001 (– 0.92)	– 0.00004** (– 2.01)	– 0.00002 (– 0.85)	– 0.00001 (– 0.15)	0.00006 (1.24)	– 0.00011 (– 1.29)	0.00008** (3.34)
LR_t^{i2}	0.121** (2.89)	0.080** (2.59)	0.041 (0.77)	0.094** (2.88)	0.093* (1.68)	0.103 (0.60)	0.066** (2.37)
$d_t^{797}LR_t^{i2}$	– 0.004 (– 0.10)	– 0.037 (– 1.26)	0.001 (0.01)	0.056 (0.71)	– 0.067 (– 1.25)	– 0.028 (– 0.16)	0.218** (3.11)
$d_t^{1198}LR_t^{i2}$	0.054 (1.30)	– 0.014 (– 0.73)	0.016 (0.28)	– 0.102 (– 1.10)	– 0.039* (– 1.93)	0.773* (1.88)	– 0.248** (– 3.07)
N	1302	1302	1302	1302	1302	1302	1302
Adj R^2	0.47	0.48	0.34	0.28	0.29	0.24	0.30
Effects during 7/97 - 10/98:							
LR_t^{i2}	0.116**	0.043**	0.042	0.150	0.026**	0.075**	0.284**
Effects during 11/98 - 12/99:							
LR_t^{i2}	0.170**	0.029**	0.058	0.048	– 0.013	0.848**	0.036

Table 9

Does the volatility of fund prices in New York drive the volatility of the Asian local market?

The first panel of this table reports slope coefficients and z-statistics (using Bollerslev-Wooldridge QML standard errors) for the ***variance*** equation from the following system:

$$LR_t^i = \beta_0 + \beta_1 FR_{t-1}^i + \beta_2 FRO_{t-1}^i + \beta_3 US_{t-1} + \beta_4 LR_{t-1}^i$$
$$+ \beta_5 d_{t-1}^{797} + \beta_6 d_{t-1}^{797} FR_{t-1}{}^i + \beta_7 d_{t-1}^{797} FRO_{t-1}^i + \beta_8 d_{t-1}^{797} US_{t-1} + \beta_9 d_{t-1}^{797} LR_{t-1}^i$$
$$+ \beta_{10} d_{t-1}^{1198} + \beta_{11} d_{t-1}^{1198} FR_{t-1}{}^i + \beta_{12} d_{t-1}^{1198} FRO_{t-1}^i + \beta_{13} d_{t-1}^{1198} US_{t-1} + \beta_{14} d_{t-1}^{1198} LR_{t-1}^i$$
$$+ \beta_{15} FR_{t-1}^i{}^2 + \eta_t^i$$
$$\eta_t^i \sim N(0, k_t^i)$$
$$k_t^i = \delta_0 + \delta_1 \eta_t^{i\,2} + \delta_2 k_{t-1}^i + \delta_3 d_{t-1}^{797} + \delta_4 d_{t-1}^{1198} + \delta_5 FR_{t-1}^i{}^2 + \delta_6 d_{t-1}^{797} FR_{t-1}^i{}^2 + \delta_7 d_{t-1}^{1198} FR_{t-1}^i{}^2$$

In the first column, the "portfolio of local returns" is an equally weighted average of the local dollar returns; it is regressed on a constant, the US market return, an equally weighted portfolio of fund-return residuals, its own lag, and interactions of these variables with the crisis dummy.

Constant term and coefficients on the lagged dependent variable are not reported. Figures followed by an asterisk (*) are significant at the 10% level; those followed by a double asterisk (**) at the 5% level. The number of observations and the adjusted R^2 of each regression are also reported.

The second and third panels report the totals of the coefficients on the squared local return during the periods when d_t^{797} equals 1 (ie $\delta_5+\delta_6$) and when d_t^{1198} equals 1 (ie $\delta_5+\delta_6+\delta_7$.). Significance levels are based on the F-statistic from a Wald test for the sum of the coefficients equalling zero.

Explanatory var.	Portfolio of returns	Korea loc return	Taiwan loc return	Thailand loc return	Indonesia loc return	Malaysia loc return	Philippines local return
$\hat{\varepsilon}_t^{i\,2}$	0.151** (11.58)	0.060** (3.08)	0.106** (3.36)	0.059** (3.09)	0.137** (3.90)	0.084** (4.47)	0.093** (3.07)
h_{t-1}^i	0.599** (36.72)	0.906** (42.05)	0.720** (11.05)	0.913** (39.54)	0.719** (12.96)	0.836** (27.28)	0.853** (18.13)
d_{t-1}^{797}	0.00003** (3.01)	– 0.00001 (– 1.12)	0.00000 (0.33)	0.00000 (0.30)	0.00009 (1.08)	0.00002 (0.95)	0.00004** (2.12)
d_{t-1}^{1198}	– 0.00007** (– 20.22)	0.00002* (1.69)	– 0.00001 (– 0.48)	0.00001 (0.75)	– 0.00002 (– 0.26)	0.00000 (– 0.07)	– 0.00004** (– 1.98)
FR_{t-1}^{i2}	0.002 (0.04)	0.009 (0.76)	0.024 (1.01)	0.008 (0.79)	0.032 (1.47)	0.016** (2.58)	0.012 (1.39)
$d_{t-1}^{797}FR_{t-1}^{i2}$	0.003 (0.04)	0.033* (1.90)	0.021 (0.60)	0.014 (1.12)	0.296** (2.91)	0.073 (1.11)	– 0.014 (– 1.26)
$d_{t-1}^{1198}FR_{t-1}^{i2}$	0.001 (0.08)	– 0.041 (– 1.58)	0.047 (1.14)	– 0.033 (– 1.54)	– 0.263** (– 2.45)	– 0.094 (– 1.39)	0.001 (0.06)
N	1302	1302	1302	1302	1302	1302	1302
Adj R^2	0.25	0.07	0.06	0.13	0.08	0.11	0.21
Effects during 7/97 - 10/98:							
FR_{t-1}^{i2}	0.005	0.042**	0.044	0.022*	0.328**	0.089	– 0.002
Effects during 11/98 - 12/99:							
FR_{t-1}^{i2}	0.005**	0.001	0.091**	– 0.011	0.066*	– 0.005**	– 0.001

Table 10

Summary statistics on monthly purchases of Asian equities

These tables present summary statistics on the monthly net purchases of equities by customers of State Street Bank and Trust Co., a large international securities custodian. Each monthly figure is the ratio of total purchases of equities from the specified country during that month to the country's average stock market capitalisation during the month, expressed in percentage points. Negative values indicate net sales.

	Mean	Std Deviation	Minimum	Maximum
Korea				
Full sample (1/95-9/99)	3.18	8.04	– 20.65	35.26
Pre-crisis (1/95-6/97)	2.84	3.42	– 4.17	12.15
Crisis (7/97-10/98)	7.06	13.42	– 6.13	35.26
Post-crisis (11/98-9/99)	0.74	8.56	– 20.65	14.31
Thailand				
Full sample (1/95-9/99)	3.05	5.36	– 7.68	20.53
Pre-crisis (1/95-6/97)	2.38	4.38	– 7.68	17.10
Crisis (7/97-10/98)	7.45	6.41	– 1.00	20.53
Post-crisis (11/98-9/99)	0.87	4.51	– 7.05	9.00
Indonesia				
Full sample (1/95-9/99)	3.56	5.94	– 8.77	23.24
Pre-crisis (1/95-6/97)	4.42	4.84	– 8.77	15.68
Crisis (7/97-10/98)	2.15	6.24	– 7.91	13.27
Post-crisis (11/98-9/99)	2.97	7.65	– 7.06	23.24
Malaysia				
Full sample (1/95-9/99)	1.31	4.61	– 12.15	13.10
Pre-crisis (1/95-6/97)	2.70	3.67	– 4.49	12.92
Crisis (7/97-10/98)	0.25	7.46	– 12.15	13.10
Post-crisis (11/98-9/99)	– 0.62	2.10	– 4.10	1.83
Philippines				
Full sample (1/95-9/99)	3.19	6.45	– 14.38	17.33
Pre-crisis (1/95-6/97)	5.24	4.39	– 1.69	17.33
Crisis (7/97-10/98)	5.14	6.93	– 3.61	15.16
Post-crisis (11/98-9/99)	– 2.46	6.51	– 14.38	6.76
Five-country average				
Full sample (1/95-9/99)	2.86	4.23	– 6.68	17.58
Pre-crisis (1/95-6/97)	3.51	2.84	– 1.11	11.01
Crisis (7/97-10/98)	4.41	5.95	– 2.34	17.58
Post-crisis (11/98-9/99)	0.30	4.12	– 6.68	5.73

Source: State Street; author's calculations.

Table 11

Does the Asian local market have more influence on fund prices in New York when US investors are buying Asian equities?

The first panel of this table reports slope coefficients and z-statistics (using Bollerslev-Wooldridge QML standard errors) for the system:

$$FR_t^i = \alpha_0 + \alpha_1 LR_t^i + \alpha_2 RR_t^i + \alpha_3 US_t + \alpha_4 FR_{t-1}^i$$
$$+ \alpha_5 IF_t^i + \alpha_6 IF_t^i LR_t^i + \alpha_7 IF_t^i RR_t^i + \alpha_8 IF_t^i US_t + \alpha_9 IF_t^i FR_{t-1}^i + \varepsilon_t^i$$
$$\varepsilon_t^i \sim N(0, h_t^i)$$
$$h_t^i = \gamma_0 + \gamma_1 \varepsilon_t^{i\,2} + \gamma_2 h_{t-1}^i + \gamma_3 IF_t^i + \gamma_4 LR_t^{i\,2} + \gamma_5 IF_t^i LR_t^{i\,2}$$

The inflow variable, IF_t^i, equals the net purchase of equities from country *i* by US investors during the corresponding month, scaled by the country's market capitalisation. Data for this variable for Taiwan were not available. In the first column, the "portfolio of funds" is an equally weighted average of the country-fund returns; it is regressed on a constant, the US market return, an equally weighted portfolio of local market returns, its own lag, and interactions of these variables with an average of the five countries' net-equity-purchase variables.

Constant term and coefficients on the lagged dependent variable are not reported. Figures followed by an asterisk (*) are significant at the 10% level; those followed by a double asterisk (**) at the 5% level. The number of observations and the adjusted R^2 of each regression are also reported.

Explanatory variable	Portfolio of funds	Korea funds	Thailand funds	Indonesia funds	Malaysia fund	Philippines fund
Mean equation:						
IF_t^i	0.004 (0.56)	0.003 (0.35)	– 0.010 (– 0.98)	0.003 (0.38)	0.003 (0.28)	0.008 (1.05)
LR_t^i		0.603** (22.66)	0.440** (11.87)	0.297** (10.04)	0.431** (8.70)	0.510** (11.23)
IF_t^i *LR_t^i		– 0.898** (– 3.82)	– 1.504** (– 2.36)	0.752* (1.71)	– 1.932** (– 2.65)	– 0.194 (– 0.32)
RR_t^i	0.658** (18.90)	0.114** (2.74)	0.135** (2.22)	0.340** (5.09)	0.237** (4.16)	0.150** (2.85)
IF_t^i *RR_t^i	– 1.066* (– 1.93)	– 0.212 (– 0.82)	1.482 (1.64)	– 1.586 (– 1.53)	1.587* (1.90)	0.109
US_t	0.476** (11.28)	0.491** (8.56)	0.517** (6.99)	0.448** (5.50)	0.485** (7.81)	0.565** (7.72)
IF_t^i *US_t	0.108 (0.12)	0.521 (0.59)	1.052 (0.75)	0.341 (0.35)	– 0.358 (– 0.28)	– 1.139 (– 1.12)
Variance equation:						
$\hat{\varepsilon}_t^{i\,2}$	0.126** (3.88)	0.062** (2.56)	0.094** (3.30)	0.057** (3.69)	0.080 (1.59)	0.129** (2.95)
h_{t-1}^i	0.620** (10.42)	0.743** (15.71)	0.774** (16.60)	0.933** (45.62)	0.797** (12.70)	0.625** (9.96)
IF_t^i	0.00003 (0.58)	0.00009 (1.07)	0.00002 (0.30)	0.00002 (1.28)	0.00023 (1.51)	– 0.00012* (– 1.72)
$LR_t^{i\,2}$	0.151** (4.39)	0.076** (4.23)	0.099** (3.17)	0.007 (1.57)	0.116** (2.09)	0.199** (4.31)
IF_t^i *$LR_t^{i\,2}$	– 0.523** (– 2.54)	– 0.162** (– 2.24)	0.150 (0.44)	– 0.086** (– 2.52)	– 0.856** (– 2.04)	– 0.846** (– 2.63)
N	1236	1236	1236	1236	1236	1236
Adj R^2	0.37	0.42	0.25	0.25	0.19	0.27

Table 11 (contd)

Explanatory variable	Portfolio of funds	Korea funds	Thailand funds	Indonesia funds	Malaysia fund	Philippines fund
Effects when net purchases of Asian equities are at their average level:						
LR_t^i	0.628**	0.573**	0.395**	0.322**	0.411**	0.503**
RR_t^i		0.107**	0.180**	0.286**	0.254**	0.154**
US_t	0.479**	0.508**	0.549**	0.460**	0.481**	0.526**
$LR_t^{i\,2}$	0.136**	0.070**	0.104**	0.004	0.107**	0.170**
Effects when net purchases of Asian equities are one S.D. above their average level:						
LR_t^i	0.583**	0.503**	0.314**	0.368**	0.314**	0.491**
RR_t^i		0.091**	0.259**	0.190**	0.333**	0.160**
US_t	0.483**	0.549**	0.605**	0.480**	0.463**	0.457**
$LR_t^{i\,2}$	0.114**	0.058**	0.112**	– 0.001	0.064**	0.119**

Table 12

Does investor sentiment in New York have more influence on local Asian returns when US investors are buying Asian equities?

The first panel of this table reports slope coefficients and z-statistics (using Bollerslev-Wooldridge QML standard errors) for the system:

$$LR_t^i = \beta_0 + \beta_1 FR_{t-1}^i + \beta_2 FRO_{t-1}^i + \beta_3 US_{t-1} + \beta_4 LR_{t-1}^i + \beta_5 IF_{t-1}^i$$
$$+ \beta_6 IF_{t-1}^i FR_{t-1}^i + \beta_7 IF_{t-1}^i FRO_{t-1}^i + \beta_8 IF_{t-1}^i US_{t-1} + \beta_9 IF_{t-1}^i LR_{t-1}^i + \beta_{10} {FR_{t-1}^i}^2 + \eta_t^i$$
$$\eta_t^i \sim N(0, k_t^i)$$
$$k_t^i = \delta_0 + \delta_1 {\varepsilon_t^i}^2 + \delta_2 h_{t-1}^i + \delta_3 IF_{t-1}^i + \delta_4 {LR_t^i}^2 + \delta_5 IF_{t-1}^i {LR_t^i}^2$$

The inflow variable, IF_t^i, equals the net purchase of equities from country *i* by US investors during the corresponding month, scaled by the country's market capitalisation. In the first column, the "portfolio of returns" is an equally weighted average of the local market returns returns; it is regressed on a constant, the US market return, an equally weighted portfolio of fund returns, its own lag, and interactions of these variables with an average of the five countries' net-equity-purchase variables.

Constant term and coefficients on the lagged dependent variable are not reported. Figures followed by an asterisk (*) are significant at the 10% level; those followed by a double asterisk (**) at the 5% level. The number of observations and the adjusted R^2 of each regression are also reported.

Explanatory variable	Portfolio of returns	Korean loc return	Thailand loc return	Indonesia loc return	Malaysia loc return	Philippines loc return
			Mean equation:			
IF_{t-1}^i	0.023** (3.22)	0.024* (1.92)	0.016 (1.32)	0.018 (0.99)	0.035** (3.44)	0.017** (2.56)
FR_{t-1}^i		0.142** (3.23)	0.058 (1.37)	0.241** (4.16)	– 0.002 (– 0.08)	0.037 (1.32)
$IF_{t-1}^i * FR_{t-1}^i$		– 0.145 (– 0.20)	1.435* (1.87)	– 0.435 (– 0.66)	0.745 (1.22)	– 0.388 (– 0.90)
FRO_{t-1}^i	0.213** (5.12)	0.042 (0.88)	0.191** (3.21)	0.442** (2.37)	0.064 (1.42)	0.289** (7.16)
$IF_{t-1}^i * FRO_{t-1}^i$	– 0.744 (– 0.86)	0.428 (0.57)	1.403 (1.46)	– 1.070 (– 0.27)	2.951** (2.20)	– 0.318 (– 0.53)
US_{t-1}	0.293** (6.30)	0.173** (2.53)	0.198** (2.40)	0.099 (0.70)	0.287** (5.51)	0.191** (3.86)
$IF_{t-1}^i * US_{t-1}$	– 2.128** (– 2.34)	– 1.508 (– 1.16)	– 1.904 (– 1.23)	1.973 (0.66)	– 2.086* (– 1.69)	0.304 (0.37)
			Variance equation:			
$\hat{\varepsilon}_t^{i\,2}$	0.061** (3.55)	0.044** (2.76)	0.041** (2.34)	0.270** (2.60)	0.113** (4.35)	0.096** (4.08)
h_{t-1}^i	0.915** (43.51)	0.932** (57.98)	0.951** (61.96)	0.550** (3.69)	0.889** (42.18)	0.900** (38.26)
IF_{t-1}^i	– 0.00001 (– 0.46)	– 0.00002 (– 0.38)	– 0.00002 (– 0.47)	– 0.00096** (– 2.71)	– 0.00005 (– 1.40)	0.00001 (0.64)
$FR_{t-1}^{i\,2}$	0.020** (2.42)	0.025** (3.47)	0.008 (1.42)	0.093 (1.38)	0.000 (0.10)	0.003 (0.89)
$IF_{t-1}^i * FR_{t-1}^{i\,2}$	– 0.128 (– 1.06)	0.017 (0.13)	– 0.013 (– 0.13)	0.005 (0.00)	0.087 (1.07)	– 0.077 (– 1.10)
N	1236	1236	1236	1236	1236	1236
Adj R^2	0.20	0.06	0.13	0.09	0.09	0.19

Table 12 (contd)

Explanatory variable	Portfolio of returns	Korean loc return	Thailand loc return	Indonesia loc return	Malaysia loc return	Philippines loc return
Effects when net purchases of Asian equities are at their average level:						
FR_{t-1}^{i}	0.192**	0.137**	0.101**	0.226**	0.006	0.023
FRO_{t-1}^{i}		0.056	0.233**	0.406**	0.095**	0.278**
US_{t-1}	0.232**	0.123*	0.140**	0.167*	0.265**	0.201**
$FR_{t-1}^{i\,2}$	0.016**	0.026**	0.008	0.093*	0.001	0.001
Effects when net purchases of Asian equities are one S.D. above their average level:						
FR_{t-1}^{i}	0.161**	0.125*	0.178**	0.200**	0.043	0.000
FRO_{t-1}^{i}		0.089	0.308**	0.341	0.243**	0.259**
US_{t-1}	0.144**	0.006	0.038	0.285	0.161**	0.219**
$FR_{t-1}^{i\,2}$	0.011*	0.027*	0.007	0.093	0.005	– 0.004

Comments on "Information flows during the Asian crisis: evidence from closed-end funds" by Benjamin H Cohen and Eli M Remolona

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I am very pleased to be here today and to comment on this very interesting and constructive paper dealing with information flows during the international financial crisis. In my view, the Cohen and Remolona paper is an excellent example of the insights empirical research gives practitioners and policy makers into policy questions regarding market design and regulations. Specifically, the paper examines global flows of information before, during and after the Asian crisis and offers important insights into the debate over the cause of the financial crisis and activities of foreign investors such as hedge funds. I will not go into the content of the paper in detail, but would like to point out some of the most interesting findings.

First, the paper focuses on the direction of information flows with respect to financial market returns in Asian markets; did information flow from the domestic markets to the US market, or vice versa?

As the paper points out, this question is a fundamental one in understanding the Asian financial crisis because, depending on the answer, different lessons could be derived from the experience regarding the cause of the financial crisis.

The finding of the paper in this respect is that while information flows between local and US markets tended to be roughly evenly balanced before the crisis, US market returns assumed a more important role during the crisis. The paper also finds that the shift in causation between the US and Asia reflected US market sentiment playing a greater role than the news that became known during US trading hours.

These findings are very interesting and I am interested to know what exactly they mean. The authors conclude that they reflected changes in investor sentiment, but what transmission mechanisms can be identified and through which changes does US investor sentiment have a material effect in determining emerging-market returns?

In my view, there seem to be several explanations to the mechanisms. One explanation emphasizes information linkage. Connections between US closed-fund returns and local market returns occur through information. If both markets share some common factor regarding risk and return, then fund returns in one market have an impact on those in the other market. In this explanation, the market doesn't need to be connected with other markets in a transactional sense. A shock may be transmitted from the US market to Asian markets, even though investors trading in the former might be completely different from those in the latter. The only thing needed is some shared some common factor regarding risk and return.

Regarding information linkage, the paper provides us with a very interesting insight, namely that the common factor reflected US investors' sentiment toward specific markets, and not merely the news which was known during US trading hours. This seems to imply that global contagion spreads from industrial countries to emerging economies only when investors recognize events or news as important factors to be considered in specific risk-return relationships.

An alternative explanation to the mechanisms producing contagion and amplifying market dynamics could be the effect of portfolio rebalancing. From this point of view, contagion occurs through trading for portfolio rebalancing purposes. A typical example is the repatriation of funds from emerging markets by investors in industrial countries. Actually, it is said that some signs of such behaviour were observed in some Asian countries in the first half of this year, which was affected by US stock market corrections. Such behaviour can perhaps, to some extent, be explained by the decreasing willingness of investors to shoulder credit risk because of losses incurred by US stock market corrections.

In this context, findings in the paper might be interpreted as follows: Investors rebalanced their portfolios because returns in other markets had a direct effect on their risk tolerance. Movements in US markets had such an effect on US-based closed-fund investors, but not so on local market investors. Stronger interconnection between US-based closed-funds and local funds during the crisis might suggest the existence of international investors who invested in both US-based closed-funds and local funds. I think such a supposition is plausible, because such investment in different types of funds, in which the underlying instrument is the same but liquidity is different, could contribute to hedging or arbitraging purposes.

Supposing that a shock was strongly transmitted between the US market and Asian markets through information linkage or through rebalancing effects during the crisis, why was the direction of information flow from US market returns to local markets returns more strongly transmitted and not vice versa? Is it because mature-market investors have better information about emerging markets than local market investors? Or is it because US market returns have a bigger effect on the risk tolerance of international investors than local market returns? I think this is a very interesting question on which future research in this area might focus.

Another interesting finding in the Cohen and Remolona paper concerns market sentiment and capital flows between the US and Asian markets. The paper suggests US investors tended to be contrarians in their portfolio activities vis-à-vis Asia: they purchased Asian equities just when their opinions differed most from investors in the local markets.

As I see it, this finding is actually confirmed by trends in US capital flow data and is very stimulating in understanding the market dynamics during the crisis. According to capital movement statistics by the US Treasury Department as shown in Chart 1, US investors seemed to purchase Asian stocks at low prices even during the Asian crisis and sold them at higher prices to take profits in the latter half of 1999.

In contrast, Japanese investors exhibited different investment behaviour, as shown in Chart 2. It was not until 1999 that net equity investment from Japan to Asian countries turned to a positive figure. These results might suggest that not only local investors but also Japanese investors pursued positive feedback strategies toward Asian equities during the crisis, while US investors followed negative feedback strategies.

What accounts for such a difference between the two types of investors?

Does this mean that Asian investors, including Japanese investors, were so greatly affected by local news that they could not make objective judgement about investment in Asian equities? Or does this difference stem from that of different time horizons? This point is interesting and could be related to the issue regarding information asymmetry between US investors and local investors.

In my view, the debate over investor strategies as to whether positive feedback or negative feedback and the impact on the market is particularly important, in the sense that it could be a key aspect of market dynamics during the crisis. If positive feedback investors become dominant in the market when market strains emerge, they tend to close out their long positions even though there may be strong suspicion that prices have overshot only temporarily. Such herding behaviour exacerbates market pressures. Typically, it is said that such a mechanism occurred in autumn 1998. However, the paper suggests that US-based international investors should perhaps not always be regarded as the culprits of the Asian crisis, whose sharp outflow of portfolio investments brought about declines in asset prices. In fact, they could even be regarded as having helped market prices recover in the healing process after the crisis.

It is said that recently index-based investment style has become more prevalent and having more of an impact on the market. Such investors are typical positive feedback ones and I think the analytical framework in the paper may be applied to analyse such changes in investment style in the market.

In closing, once again I would like to say that many of the findings in the paper contain stimulating points which should be taken into account when considering policy issues regarding market design and regulations. I hope such empirical studies as these will further contribute to such policy issues in the future.

Thank you for your kind attention.

Comments on "Information flows during the Asian crisis: evidence from closed-end funds" by Benjamin H Cohen and Eli M Remolona

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The recent Asian crisis has generated a heated debate regarding the underlying causes and the appropriate policy response. Cohen and Remolona provide a very nice and succinct summary of the literature, outlining opposing positions classifying the crisis as either home grown or driven by developments in international financial markets. In the course of weighing the evidence they review relevant data on the direction of foreign capital flows and the premia on US closed-end Asian equity funds. Although such data is useful, they note it has inherent limitations. In particular, it is at best available at a weekly interval rendering direct analysis of the dynamics of the information flow during the crisis impossible. Being able to study the markets at a daily frequency and investigating whether return relevant information has its origin in Asia and is flowing from Asia to foreign markets or vice versa is of interest for a number of reasons. It permits a more direct examination of the dynamics of the crisis and the role exerted by different factors. For example, it may shed light on whether foreign "speculators" were to blame for the development and what may be achieved by imposing restrictions on international capital flows. It would thus also speak indirectly to the desirability of including private investors in debt workouts and the appropriate role of IMF.

Unfortunately, data that allow us to explore this issue directly is not readily available since the stocks Asian equities typically are not traded outside of the local markets. Cohen and Remolona suggest an interesting way around this problem. They analyse the daily interaction between returns on local Asian equity markets and US closed-end funds investing in these Asian equity markets. Although the net asset value of the stocks underlying the fund is not available daily, the market price of the fund is. Comparing the associated daily fund returns with the daily Asian equity-index returns provides direct evidence on the relative valuation of these assets from a US versus a local perspective.

The above observation motivates the regression-based analysis in the remainder of their paper. The comments below focus on methodological aspect of their approach. My main concern is that the explored regressions are hard to interpret and fail to fully exploit the lack of overlap between the trading in the local Asian and the US market. To illustrate my point it is convenient to consider a stylised framework. Assume first that the portfolio of stocks underlying the US traded closed-end fund is identical to the basket making up the local Asian equity index so that the returns are based on the same underlying assets, and investors may purchase the fund and stock index freely at no transaction costs. Second, assume that the returns are uncorrelated and the expected daily mean returns are small relative to the return innovations. These characteristics are broadly consistent with the hypothesis of an efficient market, although we may anticipate positive autocorrelation in the index returns due to nonsynchroneous trading effects. However, one may readily control for such features in practice. Third, let the calendar day be split into two consecutive segments, the Asian trading day followed by the US trading day. For simplicity, we assume no overlap and no separation between the two trading periods. Consequently, the close-to-close (CC) fund return, $R_t^F(CC)$, is composed of two distinct uncorrelated return innovations, ε_t^F and ε_t^L. Likewise, the close-to-close local return, $R_t^L(CC)$, consists of two uncorrelated components, ε_t^L and ε_{t-1}^F. In contrast, note that the corresponding open-to-close (OC) returns are $R_t^F(OC) = \varepsilon_t^F$ and $R_t^L(OC) = \varepsilon_t^L$. We also stipulate constant return variances, $Var(\varepsilon_t^F) = \Phi_F^2$ and $Var(\varepsilon_t^L) = \Phi_L^2$.

A stylised representation of the Cohen and Remolona regressions take the form,

$$R_t^F(CC) = a_0 + a_1 R_t^L(CC) + u_t^F, \quad (1)$$

and

$$R_t^L(CC) = b_0 + b_1 R_{t-1}^F(CC) + u_t^L, \quad (2)$$

Exploiting the decomposition of the close-to-close returns, we may rewrite (1) as

$$\varepsilon_t^L = a_0 + a_1 (\varepsilon_t^L + \varepsilon_{t-1}^F) + (u_t^F - \varepsilon_t^F), \quad (3)$$

where the "error" term, $u_t^F - \varepsilon_t^F$, is uncorrelated with the "regressor," $\varepsilon_t^L + \varepsilon_{t-1}^F$, while the dependent variable, ε_t^L, is identical to one component of the regressor and orthogonal to the other, ε_{t-1}^F.

Consequently, equation (1) may be interpreted as an "error-in-variables" regression, where the open-to-close local Asian index return is regressed upon a noisy version of itself. If the second component was not included in the regressor, we would trivially have an asymptotic regression coefficient of unity, i.e., plim $a_1 = 1$. Given the "error-in-variables" representation, standard results instead imply,

$$\text{plim } a_1 = 1 - [\Phi_F^2 / (\Phi_L^2 + \Phi_F^2)]. \tag{4}$$

Equation (4) constrains a_1 to the unit interval. Moreover, a_1 increases with the size of the return innovations during local Asian trading relative to US trading. This suggests a simple interpretation of shifts in a_1 across subsamples: A declining a_1 implies that relatively more return relevant information is generated during US trading hours and vice versa. Similar arguments apply to equation (2). We have,

$$\text{plim } b_1 = 1 - [\Phi_L^2 / (\Phi_F^2 + \Phi_L^2)]. \tag{5}$$

Within this stylised setting, we thus have plim $(a_1 + b_1) = 1$, which reinforces the interpretation of the coefficients as representing shares of return relevant information generated in the respective trading hours. These findings suggest the following interpretation of the evidence in the paper. First, the coefficients corresponding to a_1 are almost universally higher than those corresponding to b_1, implying that the local Asian business hours generate the majority of the return relevant information for the Asian stocks. The higher regression R^2 of the specifications corresponding to equation (1) relative to (2) (Table 5 versus Table 6 in the paper) is also consistent with this view. Second, the drop in a_1 and increase in b_1 during the crisis period point to a relative elevation in the role of the US market segment during the crisis itself. These findings support the discussion in the paper.

Unfortunately, the formal justification behind these interpretations break down in the representations where additional variables, correlated with the above return innovations, are introduced into the system. Because the specifications include indicators observed simultaneously with both the local return (regional Asian market returns) and fund returns (US market index), the interpretation is confounded by complicated and likely time-varying correlation effects that are hard to assess within an "error-in-variables" style regression. This is further accentuated if one uses close-to-close returns for the auxiliary indices as they span both the Asian and US trading periods, and hence generally will be correlated with both Asian and fund return innovations. In particular, the direct interpretation of the relative size of the coefficients is lost and the estimates no longer cumulate to a meaningful measure of overall significance. Confirming this observation, the coefficients corresponding to a_1 and b_1 aggregate to less than unity -- slightly so in the pre-crisis period but much so during the crisis itself. Rather than trying to decipher the interactions between the parameter estimates in this setting complicated by the correlations induced by the overlap between close-to-close returns, I encourage the authors to explore an alternative set of regressions that avoid the overlap between return components among the regressors and thus are more readily interpretable. The only additional requirement is access to daily opening prices as well as closing prices. For example, consider

$$R_t^F(CO) = c_0 + c_1 R_t^L(OC) + c_2 RR_t(OC) + v_{1,t}^F, \tag{6a}$$

$$R_t^F(OC) = d_0 + d_1 R_t^L(OC) + d_2 US_t(OC) + v_{2,t}^F, \tag{6b}$$

$$R_t^F(CC) = e_0 + e_1 R_t^L(OC) + e_2 R_{t-1}^F(OC) + v_{3,t}^F. \tag{6c}$$

Notice that equation (6a) relates the fund return to the local Asian return over the period where only the Asian markets are trading. Hence, it checks whether the fund opening price incorporates the return innovation from the Asian market one-for-one ($c_1 = 1$). Some discrepancy may also be found if the assets held by the fund differ systematically from those in the local market index. As such, this serves as a useful control. In addition, one may gauge whether the other regional Asian stock indices impact the pricing of the fund beyond the effect already reflected in the local market ($c_2 > 0$). Equation (6b) studies whether there is a delayed reaction or systematic spill-over in sentiment from Asia to US trading in Asian stocks ($d_1 > 0$). Here, I have also included the contemporaneous US return to check for a direct relation between the US market and the valuation of the local Asian assets. Finally, equation (6c) relates the close-to-close fund return to the local Asian open-to-close return, where we may test $e_1 = 1$ without worrying about an error-in-variables problem. For illustration, I have also included the lagged fund return to control for autocorrelation due to nonsynchroneous trading, $e_2 = 0$. This type of specification is much easier to deal with econometrically, and the associated interpretation of the estimates is more straightforward. Furthermore, modifications of equation (6a-c) that focus on alternative variables or controls are readily constructed.

Cohen and Remolona also study the volatility dynamics of the return series. The spirit of my comments also applies to this analysis. They base the analysis on close-to-close returns that potentially obscure the underlying relationships and certainly render the interpretation of the coefficients much more difficult than necessary. There is a well-established literature on volatility spill-overs that carefully calibrate and standardise non-overlapping return and volatility component in a manner that preserve the interpretation of standard volatility persistence measures from the daily univariate GARCH literature, rendering both marginal coefficients and cumulative effects meaningful. This is simply not attainable with the complex correlation structures that is induced among the various return components generated by overlapping close-to-close returns. Good illustrations of such procedures are provided in Engle, Ito and Lin (1990, Econometrica) and Hamao, Masulis and Ng (1990, Review of Financial Studies).

In conclusion, I find the paper both insightful and stimulating. It provides an excellent introduction to the crisis debate and points to an interesting set of data that may help us address questions directly related to the key issues. They establish that flight by foreign investors was not the driving force in the market collapse. The analysis also suggests that the US traded Asian funds were more important factors in the return dynamics during the crisis than either before or afterwards, and that the US domestic market in turn became more critical for the returns on the US based Asian country funds. This points towards an increased importance of the US market sentiment during the crisis. My more critical comments are motivated by the desire to see the available data exploited to the fullest so that the results are further robustified and the parameter estimates rendered more directly interpretable.

Estimating liquidity premia in the Spanish Government securities market[1]

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Abstract

This paper investigates the presence of liquidity premia in the relative pricing of assets traded on the Spanish government securities market. Firstly we propose a classification of bonds consisting on four categories that identify groups of bonds with a different degree of liquidity. Secondly we estimate liquidity premia, including liquidity parameters in the estimation of the zero-coupon yield curve. The results suggest the existence of a significantly higher yield (adjusted for differences in cash flows) for non-strippable bonds that mostly reflects their lower degree of liquidity. Secondly, post-benchmark bonds display a positive liquidity premium over benchmark issues. Thirdly, the lack of liquidity of pre-benchmark bonds does not seem to be priced. We also show that these pricing discrepancies are robust to the impact of taxes on bonds. The size of relative liquidity premia seems relatively small.

1. Introduction

The liquidity of an asset is generally understood as the ease of its conversion into money.[2] In practice, the conversion of an asset into money involves certain costs: searching costs, delays, broker's commissions, etc. The higher these costs the lower the degree of liquidity of the asset. In most financial markets there is a class of agents known as market makers whose function is to provide liquidity. These agents are ready to buy and sell securities up to a maximum amount and make their profit on the difference between the bid and ask prices. These latter are the prices at which the other market participants can execute surely and immediately sell and buy transactions, respectively. As a consequence, the bid-ask spread reflects the cost incurred by a typical investor to unwind an asset position, which is part of the cost of converting an asset into money. This is why the bid-ask spread is one of the most widely used measures of liquidity.

According to market microstructure models, the bid-ask spread may reflect three different costs faced by market-makers: asymmetric information costs, inventory costs, and order processing costs. However, as Gravelle (BIS, 1999) points out, asymmetric information costs should not be very significant in the case of government securities (GS). This implies that liquidity in the GS markets should be closely linked to the market-makers' inventory risk and order processing costs which ultimately depend on the level of risk of the asset (duration) and the frequency with which a transaction will be executed (turnover). Sometimes, liquidity is measured by some indicators of market activity (turnover, turnover ratio, benchmark status, age).

Because investors value asset liquidity we can expect liquidity and differences in liquidity to be priced. In other words, investors may require a *liquidity premium* for holding illiquid assets in order to compensate them for bearing higher transaction costs. In the literature, there are some papers that test for the existence of liquidity premia in securities markets. The bulk of this literature focuses on equity markets,[3] whereas there are only a few papers that focus on debt markets, most of them using US market data.

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2 See *Market liquidity: Research Findings and Selected Policy Implications* in BIS (1999) for the various dimensions of liquidity.

3 See, for example Amihud and Mendelson (1986,1989).

Papers that test for the existence of a liquidity premium in the GS markets have followed a number of approaches. For example, in Amihud and Mendelson (1991), Kamara (1994) and Garbade (1996) the liquidity premium is only estimated for short-term US Treasuries. They compare the yield of notes with that of bills with the same term to maturity. Since both instruments are identical, except for the fact that notes are more actively traded, yield differences are attributed to differences in liquidity. The liquidity premia in these papers are found to be high and significant. Elton and Green (1998) use volume as a proxy for liquidity and introduce it as an explanatory variable when fitting a zero-coupon curve to the US market. They find that the coefficient of this variable is significant for most of the days considered, although the implied liquidity premium is very small. Warga (1992) proxies liquidity by indicating whether or not an issue is on-the-run (the most recently issued security of a particular maturity). He compares the ex-post monthly excess return over the 30-day Treasury bill rate for two series of constant duration portfolios made up of securities traded on the US market, one containing on-the-run bonds and the other containing the other available bonds. Warga (1992) finds that the portfolios having the on-the-run bonds exhibit a lower return and interprets this as evidence of a liquidity premium. In other papers, like Shen and Starr (1998), liquidity is proxied by the bid-ask spread. In that paper the term premium, proxied by the excess ex-post returns of the six-month US T-bill over the 3-month bill, is estimated and it is found that the bid-ask spread accounts for a substantial portion of this premium.

The estimation of liquidity premia in GS markets is important because among other reasons, it might improve the information content of prices. For example, the existence of liquidity premia may distort the information extracted from the estimated term structure or the estimates of implied inflation expectations obtained by comparing fixed-coupon and inflation linked bonds.

Against this background, the main goal of this paper is to investigate whether there is a liquidity premium in the relative pricing of assets traded on the Spanish GS market. To do this, we first characterise the relative liquidity of bonds. The strippability of the asset and its benchmark status appear to be two relevant determinants of securities' liquidity within each maturity zone. Given this property, we consider four categories of bonds that take into account these elements. The categories are non-strippable bonds and pre-benchmark (bonds that will have the benchmark status in the future), benchmark and post-benchmark (bonds that had the benchmark status in the past) strippable bonds. We find that there are important liquidity differences among these categories of bonds according to different measures of liquidity based on activity measures and bid-ask spread. Benchmark bonds are the most liquid, followed by strippable non-benchmark bonds. And finally non-strippable bonds appear to be very illiquid.

In the second part of the paper we estimate relative liquidity premia. A traditional approach consists on calculating yield spread between non-benchmark and benchmark bonds. However, this estimation of the liquidity premia do not control for risk and tax factors derived from different cash flow structures of bonds. Our methodology follows that of Elton and Green (1998) which is based on the estimation of the term structure of interest rate. This approach allows us a better control of effects related to the cash flow structure of bonds. Concretely, we incorporate liquidity effects in the estimation of zero-coupon yield curve introducing dummy variables for the different categories of bonds. The results suggest the existence of a significantly higher yield, adjusted for differences in cash flows, for non-strippable bonds that seems to mostly reflect their lower degree of liquidity. Second, post-benchmark bonds display a positive liquidity premium over benchmark issues. Third, the lack of liquidity of pre-benchmark bonds does not seem to be priced. Even, in some periods, we find evidence of a negative premium over benchmark issues for pre-benchmark bonds. We also show that these pricing discrepancies are robust to the impact of taxes on bonds. Thus, these results point to the existence of a liquidity premium in the relative pricing of bonds traded on the Spanish GS market, although its size seems relatively small.

The remainder of the paper is organised as follows. Section 2 describes the structure of the Spanish GS market. Section 3 describes the data. Section 4 proposes a classification of bonds which tries to identify bonds with a different degree of liquidity. Section 5 estimates liquidity premia. And, finally Section 6 summarises the main conclusions.

2. Structure of the Spanish government securities market

Two types of instruments are issued by the Spanish Treasury: *Letras del Tesoro* (Treasury bills), which are short-term securities issued at a discount, and *Bonos y Obligaciones del Estado* (State bonds) which are medium and long-term securities with annual coupon payments. Both types of instruments are represented by book entries and issued via regular competitive auctions. Bills are issued at 6-, 12- and 18-month maturities, whereas bonds are issued at 3-, 5-, 10-, 15- and 30-year maturities. Auctions take place on a monthly basis, except for 30-year issues, which are auctioned every two months, and 12- and 18-month issues, which are auctioned fortnightly. In the case of medium and long-term securities issues are reopened over several consecutive auctions until the outstanding amount reaches a minimum level .The securities allocated at such auctions have identical nominal coupon and interest payment and redemption dates.

Since July 1997, the Spanish Treasury has been issuing strippable bonds. They enjoy a more favourable tax treatment for payers of the corporate income tax because the latter are not subject to withholding tax on coupon payments. These securities can be stripped during their life into n+1 zero -coupon assets, where n stands for the number of remaining coupon payments, arising from the cash flow generated by the bond's coupons and principal. The stripping process is conducted by a type of market participant, known as market makers, who assume a number of commitments subject to annual review. They are also allowed to conduct the reverse process (reconstitution). All outstanding bonds issued before July 1997 are non-strippable and do not enjoy the favourable tax treatment of strippable bonds.

Secondary market trades are conducted through three systems, the first two being reserved for market members, while the third is for transactions between market members and their clients. In the first system (known as a blind market) trading is electronically conducted without knowledge of the counterparty's identity. Only those market members who comply with certain requirements and assume a number of commitments can participate in this market segment. For instance, they are obliged to quote during at least 60% of the time of each session the five references of bonds with benchmark status subject to a maximum bid-ask spread. Blind market trades can only be outright transactions, whether spot or forward.

The second trading system (known as second tier) channels all the remaining transactions between market members. Trading is conducted directly between traders or through brokers. Some brokers post indicative bid and ask prices on electronic market information systems such as Reuters and Bloomberg. In this market segment, participants can trade outright (in spot or forward transactions) or enter into repos. Two types of repo transactions are allowed: ordinary repos and blocked repos, the difference being the fact that under the second type the buyer cannot transact freely with the securities purchased, regardless of the buy-back date set.

Clearing and settlement of GS transactions is carried by the *Central de Anotaciones en Cuenta* (Book-Entry System). Each market member holds an account in this system. Individuals or institutions who are not members have to channel transactions through a Managing Institution. The procedure used is one of cash on delivery and transactions are settled three business days later.

In addition to the secondary market, there is a futures and options market (MEFF RF) where the underlying assets are Spanish GS. At present traded contracts include 5- and 10-year futures and options and 30-year futures.

Table 1 provides turnover figures for 1999 for the different market segments and for the different types of transactions on the Spanish GS market. Repos are the type of transaction with the highest activity, which is mainly concentrated in the very short-term (mainly overnight), followed by outright spot purchases. By contrast, trading activity through forward, futures, and options transactions, is more limited.

Table 2 gives a breakdown of the holders of Spanish GS at end-1999. Financial institutions own a very large proportion of the outstanding amount (61.2% for bonds and 91.3% for bills). Non-residents also have a significant share of the market in the case of bonds (32%). By contrast, the share of non-financial companies and households is very low.

3. Data

In this paper we use daily data of prices (quoted and traded), outstanding amounts, and trading activity (trading volume and number of trades) for the Spanish GS market from January 1999 to

April 2000. This information is collected for 34 bonds,[4] 24 of them being outstanding through out the sample period (see Table 3 for a description of the issues). For the short-term, the data refer to Treasury bills and repo transactions.

The price data include the following:

(i) Daily mean prices traded in the blind and second tier markets. This information was provided by the Central de Anotaciones del Banco de España.

(ii) Daily quoted bid and ask prices collected from two sources:

  (a) Reuters-BDE: since August 1998 the Banco de España Research Department has been building a database from prices posted by Reuters at 2 p.m. It includes best quoted bid and ask prices for long and short-term issues; and bid and ask rates for repo transactions. Bid and ask daily prices are calculated using the quotations of four brokers: three[5] of them are pre-determined and the fourth is the one who traded most recently. From these quotations, the highest bid and the lowest ask prices are selected for each issue. If the bid price is higher than the ask,[6] then prices are sequentially removed until a non-negative spread is obtained.

  (b) Bloomberg: this agency provides daily quoted prices calculated as the average bid and ask quotations at the close (determined by Bloomberg to be about 5 p.m.). In this case, by construction, the bid-ask spread is always positive. Bloomberg also provides bid and ask rates for a wide range of repo transactions.

It is worth noting that these databases of quoted prices use different methodologies to calculate daily bid-ask spreads for each instrument. The Reuters-BDE spread is the best proxy for market spread, but it employs non-contemporaneous quotations that may introduce noise and may bias downwards the size of the spread. On the other hand, the Bloomberg spread cannot be interpreted as the market spread, but its main advantage is that it reduces non-contemporaneity problems due to the fact that quotes are averaged. Both the Reuters-BDE and Bloomberg databases suffer from the shortcoming that quotations are indicative, not firm, and therefore may not reflect actual prices. This problem arises from the specific microstructure of this market and means that it is desirable to use traded prices as complementary information.

Trading activity (number of trades and trading volume) is provided by Central de Anotaciones del Banco de España and refers only to transactions between market members, in both, the blind and the second tier market.

4. Classifying bonds by their relative degree of liquidity

4.1. A classification of bonds according to their life cycle and tax treatment

In this section we propose a classification of bonds that tries to identify a number of bonds' groups with a different degree of liquidity. It is important to stress that the approach we follow in this paper of classifying bonds by their relative liquidity is not intended to compare liquidity across different maturity zones but within them. This decision is justified by the methodology we use in the next section to estimate liquidity premia, which consists of fitting a zero-coupon yield curve. With this methodology it would be very difficult to separate term and liquidity effects if we used a classification of liquidity across different maturities.

Our classification takes into account two elements of an issue: its stage of the life cycle and its strippability (whether or not the issue is strippable). We consider the stage of the life because, in general, trading activity and, consequently, liquidity is related to this. It is well-known that in GS markets trading activity tends to be concentrated in a small group of securities known as *benchmark*

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[4] Strips are not included. The percentage of stripped bonds is very low in the Spanish market.

[5] Four in the case of short-term instruments.

[6] In the information provided by Reuters it is not uncommon to find negative bid-ask spreads, possibly due to the lack of contemporaneity in quotations.

issues. The prices of these instruments are used for extracting information for macroeconomic analysis and pricing purposes. In the Spanish GS market there are five benchmark issues corresponding to 3-, 5-, 10-, 15- and 30-year maturities. Since all issued bonds become benchmark issues, the life cycle of bonds consists of three stages. In the first stage, the bond is initially issued and its outstanding volume is relatively small, and henceforth, its trading activity is very low and does not have benchmark status. Henceforth bonds in this stage will be referred to as *pre-benchmark* issues. New fungible auctions increase the outstanding amount, and trading activity also increases until benchmark status is achieved and the issue becomes the most liquid one for a particular maturity. Finally, in the third stage, the bond is replaced by a new benchmark and its activity begins to decrease. Bonds in this stage will be referred to as *post-benchmark* issues.

More specifically, our classification consists on four categories, namely: pre-benchmark, benchmark, post-benchmark and non-strippable issues. The first three categories are only made up of strippable bonds, whereas the last category includes all non-strippable bonds. We have decided to group all non-strippable bonds in one specific category for the following two reasons. Firstly, in our sample, all non-strippable bonds are at a very advanced third stage of their life cycle, therefore they are very illiquid. And, secondly they are subject to a less favourable tax treatment that could be reflected in a tax premium.

Over our sample period there was no official classification of benchmark issues in Spain. As a consequence, the identification of the benchmark issue may be difficult during transition periods. The criterion we use in this paper considers that a new bond becomes the benchmark issue when it has been traded more than the old benchmark for at least three consecutive days. Table 4 gives the dates on which a benchmark bond changes according to this criterion, and compares them with those reported by Reuters and Bloomberg. It can be observed that dates do not coincide for the different criteria, but differences between them rarely exceed one month.

Figure 1 shows the life cycle for all the strippable 10-year bonds of our sample. Over this period two replacements take place, one in February 1999 when the 5.15% bond becomes the new benchmark. The other takes place in October 1999, where the most recently issued bond (4% with maturity at 31/1/10) acquires the benchmark status. This figure displays the evolution of trading activity, outstanding volume and bid-ask spread. It can be seen that the life cycle in trading activity appears also to be incorporated in the evolution of bid-ask spreads, suggesting a very different degree of liquidity of bonds depending on the stage of life they are at.

Table 5 gives a descriptive analysis of previously defined bond categories over our sample period. It can be seen that benchmark issues account, on average, for almost 53% of the daily trading volume (see panel c). Regarding non-strippable bonds, despite the high number of issues (20 issues at the beginning of the period and 16 at the end), they account for just 7% of daily trading volume (see panel c). Finally, the remaining 40% of the trading volume corresponds to strippable non-benchmark bonds - of which pre-benchmark issues account for a 15% and post-benchmark issues for a 25%. There are 4 such bonds at the beginning of the period increasing to 9 by April 2000.

Looking at trading activity by maturity, a high concentration around the 10-year zone is observed, with approximately 33% of the total market activity. The 3- and 5-year maturities account for approximately 25% and 23% respectively. And finally, the zones of 15- and 30-year residual maturity have the lowest activity, with approximately 5% and 3% of total market activity respectively.

Regarding outstanding volume, panel b of Table 5 shows a very different distribution. In particular, a high concentration (more than 50%) is observed in the 0-6 years zone.

4.2. Is this classification useful to identify groups of bonds with a different degree of liquidity?

We now show to what extent the four categories we propose are useful to identify groups of bonds with a different degree of liquidity. To do this we compute some liquidity measures for the different categories of bonds.

Table 6 reports various measures of liquidity, all based on trading activity, with a breakdown by bond category and maturity. These measures are the number of trades, trading volume and turnover ratio. All these indicators are computed per bond and, therefore, can be interpreted as liquidity measures. A feature shared by all the liquidity measures of Table 6 is the remarkable differences in liquidity between the various categories considered.

When comparing liquidity within each maturity zone benchmark issues are the bonds with the greatest liquidity, followed by strippable non-benchmark bonds (for both pre-benchmark[7] and post-benchmark). Finally, non-strippable bonds are notable for their very scarce trading activity.

When comparing liquidity among maturity zones the most liquid bonds are those with a residual maturity between 3 and 10 years. In this regard, it is worth noting that strippable non-benchmark bonds are sometimes even more liquid than 15 and 30-year benchmark issues.

Figure 2 displays the bid-ask spread[8] as a function of the duration for the different categories of bonds. This indicator is used as an additional measure of liquidity. A similar pattern to that reported for trading activity also appears here when comparing the liquidity of bonds within each maturity zone (proxied by duration), although now pre-benchmark issues appear to be more similar, in terms of liquidity, to non-strippable bonds than to post-benchmark strippable issues.

However, note that trading activity and the bid-ask spread would provide a very different ranking of bonds by liquidity if we did not control for time to maturity or duration. As Figure 2 illustrates, the bid-ask spread is positively correlated with time to maturity. This effect clearly dominates the previously reported pattern (benchmark status of the bond). This relationship between spread and time to maturity, which has also been identified in other GS markets,[9] can theoretically be explained by the market-makers' inventory risk and order processing costs which ultimately depend on the level of risk of the asset proxied by time to maturity. Accordingly, a ranking of bonds using the bid-ask spread would classify short-term securities as the most liquid and long-term issues as relatively illiquid. Conversely, a ranking of bonds by the liquidity measures of Table 6 would classify the 10-year bonds as the most liquid issues.

It is worth noting that our classification of bonds can be justified either by the bid-ask spread or by the trading activity measures because, as was mentioned before, we do not intend to compare liquidity across different maturity zones but within them.

Three conclusions may be drawn from the previous analysis. Firstly, the benchmark status and strippability of bonds are two relevant determinants of the liquidity differences within a maturity zone for the bonds in our sample. Secondly, the four categories we consider seem useful to identify groups of bonds with a different degree of liquidity. And, finally there seems to be strong differences in liquidity between bonds, so that it makes sense to test for the existence of a liquidity premium. This is the aim of the next section of the paper.

5. Estimating liquidity premia

5.1 Preliminary analysis

Our aim in this second part of the paper is to find out if differences in liquidity are priced in the Spanish GS market. One simple approach is to approximate the liquidity premium by the yield spread between non-benchmark and benchmark bonds. Applying this measure to the 3,5, 10 and 15-year bonds of our sample always gives a negative figure suggesting the existence of negative liquidity premia. However, it has to be taken into account that some factors may bias this measure. Differences in coupons between bonds imply different risk and tax burdens that may make the investor to demand different yields to maturity for other reasons not related to liquidity effects. The main difficulty when identifying the liquidity premium is the separation of the factors explaining differences in yields to maturity. With this in mind, our proposal is based on the estimation of the term structure of interest rate, because this allows us to control any effect associated with differences in maturity and cash flows. This approach is similar to that followed by Elton and Green (1998).

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7 Although in the case of the turnover ratio no significant differences are observed between pre-benchmark and benchmark issues.

8 For this analysis we use Reuters-BDE database because it provides a spread measure more closely linked to the concept of market spread (see Section 3 for a description of this database and a comparison with the Bloomberg data).

9 BIS (1999) compares bid-ask spreads for "on-the-run" and "off-the-run" issues for Canada, Italy, Japan, UK and US. In all cases they find the same results as ours, except for Japan, where the liquidity of on-the-run 10-year bonds makes their spread narrower than the one for "on-the-run" 5-year bonds. (See Table 3 of the document *Market Liquidity: Research Findings and Selected Policy Implications*)

According to the analysis developed in Section 4, we distinguish four types of bonds in relation to their life cycle and tax-treatment: non-strippable bonds, pre-benchmark, post-benchmark, and benchmark strippable bonds.

As a preliminary analysis of the existence of liquidity premia we estimate a zero-coupon yield curve using the Svensson model[10] and analyse yield errors for bonds. We find a very interesting pattern. Figure 3 shows the estimation for a representative day of the sample, (29/3/99) and Table 7 reports average yield errors for each bond in our sample distinguishing its classification over the sample. The first thing we observe is that most non-strippable bonds appear to have on average positive yield errors, i.e. they are located above the estimated yield curve. The second point to be stressed is the tendency of pre-benchmark bonds to exhibit a lower yield error than benchmark bonds. Finally, the yield error of post-benchmark bonds lies between yield errors of benchmark and non-strippable bonds. These results suggest a preliminary evidence of a positive premium for post-benchmark and non-strippable bonds over benchmark issues. Next section tries to estimate them using a more formal approach.

5.2 Estimating the yield curve incorporating liquidity effects

Previous results indicate that the estimation of the yield curve for the Spanish GS market may be improved if we introduce liquidity and withholding tax affects. Our proposal consists of estimating the term structure of interest rates using the Svensson model, and introducing additional parameters to capture these effects. Elton and Green (1998) follow a similar procedure[11] introducing volume as a measure of liquidity. Given that the relationship between traded volume and liquidity may not be linear, we prefer using dummy variables to classify securities according to their degree of liquidity and tax treatment.

The instantaneous forward rate of term m is modelled in the following way:

$$\varphi_m = \beta_0 + \beta_1 e^{-\frac{m}{\tau_1}} + \beta_2 \frac{m}{\tau_1} e^{-\frac{m}{\tau_1}} + \beta_3 \frac{m}{\tau_2} e^{-\frac{m}{\tau_2}} + \gamma_0 PREBENCH + \gamma_1 POSTBENCH + \gamma_2 NONSTRIP$$

where *PREBENCH*, *POSTBENCH* and *NONSTRIP* are dummy variables which take value 1 respectively for pre-benchmark, post-benchmark and non-strippable bonds. Parameters for dummy variables must be interpreted in relation to the remaining instruments (3, 5, 10, 15 and 30-year benchmark bonds and short-run securities).

Parameters γ_0, γ_1 and γ_2 capture the excess yield demanded on pre-benchmark, post-benchmark and non-strippable bonds respectively. Parameters γ_0 and γ_1 can be directly interpreted as liquidity premia since the tax treatment is the same for all strippable issues. Conversely, parameter γ_2 may incorporate a tax premium in addition to a liquidity premium due to the relatively less favourable tax treatment of non-strippable bonds. The interpretation of remaining parameters is the standard in the Svensson model.

Once we have modelled the instantaneous forward rate including withholding tax and liquidity effects, we estimate the parameter vector $\beta = (\beta_0, \beta_1, \beta_2, \tau_1, \beta_3, \tau_2, \gamma_0, \gamma_1, \gamma_2)$ for each day t as

$$\hat{\beta} = \arg\min \sum_{i=1}^{N} \left[\frac{p_{ti} - p_{ti}(\beta)}{p_{ti}} \right]^2$$

where p_{ti} is the actual price of bond i at time t and N is the number of securities considered in the estimation. And $p_{ti}(\beta)$ is the theoretical price obtained from previous specification of the

[10] See Annex 1 for a description of this model. We use the mid-point between the bid and ask price provided by Bloomberg database. Bonds with a remaining life of under a year are excluded from the estimation. For short-term we use repo and bill rates.

[11] They use non-linear least squares to fit a cubic spline to the after-tax cash flows of bonds.

instantaneous forward rate. We minimise price errors instead of yield errors to obtain a better fit for the maturities we are most interested in, i.e. medium and long-term.

We distinguish four nested models. The basic model, called Model 1, does not consider any liquidity or tax effects, i.e. γ_0, γ_1 and γ_2 are restricted to zero. In Model 2 all these parameters are freely estimated. Model 3 eliminates the dummy variable for pre-benchmark bonds. And finally, Model 4 estimates freely the parameter of the dummy variable for non-strippable bonds, and restricts to zero the parameters of other dummy variables.

We estimate these four models for the whole sample, from January 1999 to April 2000. Over the period 4/1/99 to 30/9/99 the Svensson model does not improve on the Nelson-Siegel model[12] and consequently, we estimate the latter model for this period. Table 8[13] reports a descriptive analysis of the estimated parameters for dummy variables, the root mean square error (RMSE), and the reduction in the RMSE with respect to the basic model.

When all the three dummy parameters are freely estimated (Model 2) the average RMSE is 63% less than that of the basic model. The estimated parameter for the non-strippable bonds dummy is positive and significant for almost the whole sample, indicating that these bonds incorporate a premium over the benchmark issues that ranges from 2 to 10 basis points in our sample. On average, the excess yield demanded on these bonds is 7 basis points.

Regarding post-benchmark bonds, they also appear to incorporate a premium. It takes the value of 5 basis points on average, but in this case there are some periods where the parameter of this variable is not positive and statistically significant. Figure 4 shows the evolution of estimated parameters, $\gamma_0, \gamma_1, \gamma_2$. Shaded areas indicate that γ_2 is not statistically significant. These areas coincide with transition periods in benchmark bonds. During January-February 1999 it occurs a change in 3, 5 and 10-year benchmark bonds. During October-December 1999 there is a change of all benchmark bonds except for 30-year term. Replacements take place gradually in the markets and during these transition periods frontiers between strippable categories are blurred given the small number of bonds in each one. The approach we follow changes the category of a bond in a specific date (see table 4) generating sharp movements in the estimations of premium parameters. The sensitivity of our estimations to benchmark changes does not allow interpreting the evolution of the parameters during transition periods. In any case, results when the categories of bonds are enough differentiated show that post-benchmark bonds include a liquidity premium over benchmark issues.

Regarding the parameter of pre-benchmark bonds, it turns out to be negative on average, suggesting the existence of a negative premium for these bonds over benchmark issues. However, some care must be taken in the interpretation of this parameter. It is negatively significant just for 54% of the sample, and for some periods there is just one bond included in this category. The conclusion drawn from the analysis of this parameter is that pre-benchmark bonds do not seem to include any positive liquidity premium in spite of its low liquidity in terms of trading activity and bid-ask spreads. This result may possibly be explained by the forward-looking behaviour of investors. It is costly to trade with these securities, but investors do not demand a higher yield because, in the future, these costs will be much lower when the bond acquires benchmark status. In addition, the limited supply of these issues may contribute to sustain its price.

Given that pre-benchmark bonds do not appear to have a positive liquidity premium over benchmark issues, and given that there are not many pre-benchmark bonds in our sample we proceed to group pre-benchmark and benchmark bonds into the same category. Model 3 includes only two dummy variables to capture liquidity differences, one for post-benchmark and the other for non-strippable bonds. The reduction of RMSE with respect to the basic model continues to be very high (51%). Non-strippable bonds appear to have a premium of 8 basis points on average, ranging from 5 to 12 basis points. This parameter is statistically significant for the whole sample. For post-benchmark bonds, parameter γ_1, is positive and significant for 76% of the sample, showing on average a liquidity

[12] This model is a particular case of the Svensson model, where the instantaneous forward rate is modelled with just one hump, i.e. β_3 and τ_2 are set to zero and one respectively.

[13] We concentrate on the interpretation of parameters γ_0, γ_1, γ_2, which will be reported in different tables. The remaining parameters of the Svensson and Nelson-Siegel model are presented in Table 12 for all models estimated throughout the paper.

premium of 5 basis points. Therefore, inclusion of pre-benchmark bonds with benchmark bonds increases the size of the premia.

Finally, Model 4 includes just one dummy variable for non-strippable bonds. It allows an average reduction of 37% in the RMSE with respect to the basic model. In this model the average premium for non-strippable bonds over the remaining bonds is 6 basis points. The size of this parameter is lower than in Model 3 because in this case the group of bonds excluded from dummy variables includes post-benchmark issues.

Summing up, the results of this section suggest the existence of a positive liquidity premium for post-benchmark issues over benchmark bonds, although its size is very small - similar to that found by Elton and Green (1998) in the US market. Conversely, the lack of liquidity of pre-benchmark issues does not seem to be priced and, even a negative premium over benchmark issues is found in some periods. Finally, there is clear evidence of a positive premium for non-strippable bonds in the whole sample. However, as noted before, this premium could not be interpreted directly as a liquidity premium since the excess yield demanded on non-strippable bonds may also include a withholding tax premium. In the next section we try to separate these two components.

5.3 Separation of withholding tax and liquidity effects for non-strippable bonds

Previous results show strong evidence of a premium for non-strippable bonds. This premium may be due to liquidity or withholding tax effects. In order to separate these two elements we re-estimate Model 3 taking into account the withholding tax (Model 5). To do this we modify the cash flows of non-strippable bonds in the following way. We reduce the amount of coupon payments by 18% (withholding tax rate during the sample period) and include a new cash flow stream corresponding to the compensation for the withholding tax. The dates we assign to these cash flows are selected to coincide with the quarterly tax returns that Spanish corporations are obliged to make. It has to be noted that this procedure considers the maximum cost faced by non-strippable bond holders since they can reduce it through *coupon washing transactions*.[14] Consequently, the estimated premium for non-strippable bonds can be interpreted as a lower bound for the true liquidity premium.

Table 9 reports the estimated parameters, which are comparable with those for Model 3. This model reduces the premium for non-strippable bonds by only 1.2 basis points, in comparison with Model 3, whereas the premium for post-benchmark bonds does not change significantly. These results suggest that most of the estimated premia for non-strippable bonds in Models 2 to 4 correspond to liquidity premia. According to the results of Table 9, the average liquidity premium for non-strippable bonds is at least 6.8 basis points. This minimum liquidity premium for non-strippable bonds is still higher, on average terms, than the one found for post-benchmark bonds. This fact is logical given their strong differences of liquidity.

5.4 Robustness to tax effects

The above results may be sensitive to including general tax effects in the estimation of the zero-coupon yield curve. That is, they may be affected by the estimation of a pre-tax instead of a post-tax yield curve. It is very difficult to estimate a post-tax yield curve, because tax treatment is very different across investors. In this section we estimate a post-tax yield curve considering the tax treatment for Spanish corporations, which hold two thirds of the outstanding volume of bonds (see Table 2).

The theoretical price is expressed as the present value of after-tax cash flows imposing a tax rate of 35% for both interest payment and capital gains. Capital gains or losses at redemption date are distributed over the life of the bond using a constant yield method, and assuming that the bond is held until maturity.

Table 10 shows the main results of the re-estimation of Model 3 using after-tax flows (Model 6). It can be seen that in this new model the sign and significance of estimated premia remain unchanged, but their absolute size is reduced by around 35%. The parameters that capture the asymptotic value of the

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[14] These transactions consist of selling non-strippable bonds to non-resident investors, who are not subject to withholding tax, before the date of the coupon and buying them again after that date.

instantaneous forward rate ($\beta_0 = \varphi_\infty$) and the instantaneous forward rate for an infinitely small term to maturity ($\beta_0 + \beta_1 = \varphi_0$) are also reduced by around 35%. Consequently, the liquidity premia expressed as a percentage of these interest rates remain almost unchanged (see Table 11). These results indicate that previously estimated relative liquidity premia are robust to tax effects.

Finally, it is interesting to see that the model with taxes further reduces the RMSE. In average the reduction over the basic model is 77%, which compares with the 51% reduction of the model with two dummies without taxes (model 3). This result suggests the relevance of introducing tax effects when fitting zero-coupon yield curves.

5.5 Impact of liquidity premia on the estimated yield curve

In this subsection we study the impact of the introduction of liquidity effects on the estimated zero-coupon yield curve. To do this we compare the estimated parameters for the different models we have considered in subsections 5.2 to 5.4.

Table 12 reports the mean and the standard deviation of the basic parameters of the term structure.[15] The mean of the parameters' estimates for the models that introduce liquidity effects without taxes (Models 2 to 5) are very close to those of the basic model (Model 1). This is not a surprising result given the relatively small liquidity premia we have found. Conversely, when taxes are introduced (Model 6) some parameters change significantly. β_0 and β_1 are now around 35% (the tax rate) below the estimated level for Model 3, whereas the changes in the other parameters are less dramatic. These changes imply that after-tax zero-coupon rates are around 35% below before-tax rates.

Figure 5 depicts the zero-coupon yield curve at 29/3/99 for two of the models we have estimated, using the Nelson-Siegel approach. The models we consider in this figure are: the basic model (Model 1), and the model that introduces dummy variables for both post-benchmark and non-strippable bonds (Model 3). The yield curve of the model that accounts for liquidity (Model 3) stands slightly above (with a maximum difference of 9 basis points) that of the basic model for short-term maturities and slightly below (with a maximum difference of 12 basis points) for medium-term horizons. For long-term maturities both models display very similar rates.

Figure 6 shows the estimated 1-year forward curves implied by the yield curves depicted in Figure 5. Estimated forward rates for horizons shorter than 6 years are slightly lower for the model that accounts for liquidity effects, whereas the opposite is the case for longer horizons.

Figures 7 and 8 are the same as Figures 5 and 6, except for the fact that they are estimated at 31/3/00 using the Svensson model. In this case, the zero-coupon rates are lower for the model that introduces liquidity effects in comparison with the basic model, for all horizons. It is worth noting that for the medium and long-term horizons the difference between the two curves is smaller than that found with Figure 5. This is not surprising given the reduction in the estimated liquidity premia in the second half of the sample.

The introduction of liquidity effects has allowed us to identify liquidity premia and to reduce the error in the estimation of the zero-coupon yield curve. However, the implications for information content of the estimated yield curve are not very important. This is an expected result given the small size of the estimated liquidity premium.

5.6 Biases in quoted prices

The results of the previous analysis have been derived from the quoted prices posted by brokers in Bloomberg. We know that these prices are not firm, but merely informative and, consequently, we do not know how accurate they are. The existence of a bias in quoted prices may have an impact on the liquidity premia we have estimated. In this subsection we study the existence of biases in quoted prices in comparison with traded prices and derive conclusions on their impact on the estimated liquidity premia. To do this, for each day and security we compute the difference in yield between the mean traded price and the mid-quoted price. Table 13 shows the average of these yield differences for

[15] From 4/01/99 to 30/9/99 we estimate Nelson-Siegel model, and from 1/10/99 to 14/4/00 Svensson model.

each security, distinguishing its classification along the sample in the four categories we have considered along this paper (pre-benchmark, benchmark, post-benchmark and non-strippable).

Three important features emerge from Table 13. First, average yield differences are very small for all securities (between 0 and 2.3 basis points), most of them being not significantly different from zero. Second, yield differences tend to be positive for most bonds (those figures statistically different from zero are all positive). Third, yield differences tend to be higher for pre-benchmark and non-strippable bonds.

These results suggest the existence of a small negative bias in quoted prices in comparison with traded prices, which seems to be more important for pre-benchmark and non-strippable issues. This would mean that there is a negative bias in the estimated liquidity premia for non-strippable and pre-benchmark bonds. That is, the actual liquidity premium for non-strippable bonds is higher than the estimated, whereas the absolute value of the liquidity premium for pre-benchmark issues is lower. However, these biases in the liquidity premia should be relatively small given the very small size of the bias in quoted prices. Consequently, the reported results of sections 5.1 to 5.5 seem to be robust to the data set we have used.

6. Conclusions

The analysis developed in the first part of the paper has showed that there are important liquidity differences among securities traded in the Spanish GS market. The strippability of the asset –strippable bonds are the most liquid- and the benchmark status –benchmark bonds are more liquid- appear to be two relevant determinants of the securities' liquidity within each maturity zone.

The second part of the paper analyses the presence of liquidity premia in the relative pricing of assets traded in the Spanish GS market. The estimation is carried out introducing liquidity parameters in the Svensson model of the zero-coupon yield curve. These liquidity parameters allow us to estimate the excess yield of pre-benchmark, post-benchmark and non-strippable bonds over benchmark issues. This methodology improves the traditional approach of the estimation of liquidity premium – which consists on computing the yield spread between non-benchmark and benchmark bonds - because it allows an appropriate control of effects associated with differences in maturity and cash flows.

The main conclusions that can be drawn from the estimations made are the following. Firstly, results suggest the existence of a positive and significant premium for post-benchmark bonds (both strippable and non-strippable). However, these premia are very small and similar to those found by Elton and Green (1998) in the US market. Secondly, the lack of liquidity of pre-benchmark bonds does not seem to be priced in the market. Even, in some periods, a negative premium over benchmark issues arises for these bonds. This is a somewhat surprising result that can possibly be explained by the forward-looking behaviour of market participants. Thirdly, estimated liquidity premia seem robust to tax effects and to the data set we have used (quoted prices instead of actual traded prices). Finally, regarding the implications for the estimated zero-coupon yield curve, the introduction of liquidity effects do not have significant effects due to small size of the estimated liquidity premia.

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Annex

The Svensson model

Svensson (1994) specifies a smooth parametric function for the yield curve. The functional form for the instantaneous forward rate of term *m*, φ_m, is the following:

$$\varphi_m(\beta) = \beta_0 + \beta_1 e^{-\frac{m}{\tau_1}} + \beta_2 \frac{m}{\tau_1} e^{-\frac{m}{\tau_1}} + \beta_3 \frac{m}{\tau_2} e^{-\frac{m}{\tau_2}} \tag{A1}$$

where

$\beta = (\beta_0, \beta_1, \beta_2, \tau_1, \beta_3, \tau_2)$ is the vector of parameters.

m is the term to maturity.

$\beta_0 = \varphi_\infty$ is the asymptotic value of the instantaneous forward rate.

$\beta_0 + \beta_1 = \varphi_0$ is the instantaneous forward rate for an infinitely small term to maturity.

τ_1 indicates the position of the first internal maximum or minimum.

β_2 determines the magnitude and direction of the first hump. If $\beta_2 > 0$, there is a maximum at τ_1, and a minimum if $\beta_2 < 0$.

τ_2 indicates the position of the second internal maximum or minimum.

β_3 determines the magnitude and direction of the second hump. If $\beta_3 > 0$, there is a maximum at τ_2, and a minimum if $\beta_3 < 0$.

The zero-coupon rate is the mean of integration of the instantaneous forward rate between period 0 and the term to maturity, $r_m = \frac{1}{m}\int_0^m \varphi_\theta d\theta$:

$$r_m(\beta) = \beta_0 + (\beta_1 + \beta_2)\frac{\tau_1}{m}(1 - e^{-\frac{m}{\tau_1}}) - \beta_2 e^{-\frac{m}{\tau_1}} + \beta_3 \frac{\tau_2}{m}(1 - e^{-\frac{m}{\tau_2}}) - \beta_3 e^{-\frac{m}{\tau_2}} \tag{A2}$$

The continuously compounded discount factor is $d(m) = e^{-mr_m}$:

$$d(m,\beta) = e^{-\beta_0 m - (\beta_1+\beta_2)\tau_1(1-e^{-\frac{m}{\tau_1}}) + \beta_2 m e^{-\frac{m}{\tau_1}} - \beta_3\tau_2(1-e^{-\frac{m}{\tau_2}}) + \beta_3 m e^{-\frac{m}{\tau_2}}} \tag{A3}$$

The theoretical price of bond *i* at time *t* is the present value of its cash flows:

$$p_{it}(\beta) = \sum_{j=t}^{M} d(j-t,\beta) f_{ij} \tag{A4}$$

where f_{ij} represents the coupon and principal payments of bond *i* at date *j*, and *M* is the redemption date.

The parameter vector is estimated by the minimisation of some distance between the theoretical and actual prices of the bonds. Sometimes the minimisation problem is specified in terms of yield errors instead of price errors, or some combination of both. Another possibility is estimation by maximum likelihood.

Table 1
Turnover in the Spanish government securities market
1999, EUR millions

	Total	Between market members	Through registered dealers
Outright spot purchases	1,817,465	498,958	1,318,507
Bills	79,839	31,609	48,230
Bonds	1,737,626	467,349	1,270,277
Repo transactions	11,254,309	2,785,567	8,468,742
Forward transactions	37,901	–	–
Futures and options transactions	356,146	–	–

Source: Banco de España

Table 2
Holders of Spanish government securities
End-1999, percentages

	Bills	Bonds
Banking sector	49.5	31.1
Other financial institutions	41.8	30.1
Mutual funds	35.8	16.8
Pension funds	0.5	2.9
Insurance companies	0.3	10.1
Other	5.2	0.3
Non-financial companies	1.7	3.8
Households	4.6	1.2
Non-residents	2.3	32.0
Other	0.1	1.8

Source: Banco de España.

Table 3

Bonds included in the sample

Coupon (%)	Issue Date	Maturity Date	Remaining life at 4/1/99 (years)
10.65	30/07/86	30/07/01	2.57
12.25	25/03/90	25/03/00	1.22
11.3	15/11/91	15/01/02	3.03
10.3	15/04/92	15/06/02	3.45
10.9	15/02/93	30/08/03	4.65
10.5	17/05/93	30/10/03	4.82
8.2	15/12/93	28/02/09	10.16
8.0	17/01/94	30/05/04	5.41
7.4	15/02/94	30/07/99	0.57
10.0	15/11/94	28/02/05	6.16
10.1	15/09/95	28/02/01	2.15
10.15	15/09/95	31/01/06	7.08
9.4	15/12/95	30/04/99	0.32
8.4	15/03/96	30/04/01	2.32
8.8	15/03/96	30/04/06	7.32
7.8	17/06/96	31/10/99	0.82
8.7	15/07/96	28/02/12	13.16
7.9	15/10/96	28/02/02	3.15
6.75	15/11/96	15/04/00	1.28
7.35	16/12/96	31/03/07	8.24
5.0	15/07/97	31/01/01	2.08
5.25	15/07/97	31/01/03	4.08
6.0	15/07/97	31/01/08	9.08
6.15	15/07/97	31/01/13	14.08
6.0	15/01/98	31/01/29	30.10
5.15	10/07/98	30/07/09	10.58
4.25	7/08/98	30/07/02	3.57
4.5	10/08/98	30/07/04	5.57
4.75	7/12/98	30/07/14	15.58
4.0	11/05/99	31/01/10	11.08
3.0	13/07/99	31/01/03	4.08
3.25	12/07/99	31/01/05	6.08
4.6	15/02/00	30/07/03	4.57
4.95	14/02/00	30/07/05	6.57

Table 4

Dates of change in benchmark bonds according to different criteria

Term	Benchmark		Criteria		
	Coupon (%)	Maturity date	Ours	Bloomberg	Reuters
3	5.00	31/1/01			
	4.25	30/7/02	28/1/99	12/1/99	2/2/99
	3.00	31/1/03	2/12/99	4/1/00	19/11/99
5	5.25	31/1/03			
	4.50	30/7/04	13/1/99	12/1/99	2/2/99
	3.25	31/1/05	2/12/99	4/1/00	19/11/99
10	6.00	31/1/08			
	5.15	30/7/09	26/2/99	12/1/99	2/2/99
	4.00	31/1/10	6/10/99	4/1/00	19/11/99
15[1]	6.15	31/1/13			
	4.75	30/7/14	2/12/99		
30[1]	6.00	31/1/29			

[1] Bloomberg and Reuters do not provide benchmark information on these terms.

Table 5

Main figures for Spanish GS bonds by categories and term to maturity

a) Number of issues

residual life (years)	pre-benchmark	post-benchmark	non-strippable	benchmark	total
0-2		0.9	5.9		6.9
2-4	1.0	1.5	4.8	1	8.3
4-6	0.5	0.3	3.1	1	4.9
6-9		0.9	3.1		4.0
9-11	0.4	0.4	0.8	1	2.6
11-16	0.7	0.3	1.0	1	3.0
16-30				1	1.0
Total	**2.6**	**4.4**	**18.7**	**5**	**30.7**

b) Total outstanding amount

residual life (years)	pre-benchmark	post-benchmark	non-strippable	benchmark	total
0-2		4.7%	13.4%		18.1%
2-4	1.2%	6.5%	8.6%	4.0%	20.4%
4-6	1.3%	4.6%	7.2%	4.1%	17.2%
6-9		7.9%	7.1%		15.0%
9-11	1.6%	5.6%	2.7%	4.7%	14.5%
11-16	1.5%	5.0%	0.8%	4.2%	11.6%
16-30				3.3%	3.3%
Total	**5.6%**	**34.3%**	**39.7%**	**20.3%**	**100.0%**

c) Aggregate trading volume

residual life (years)	pre-benchmark	post-benchmark	non-strippable	benchmark	total
0-2		4.5%	2.1%		6.6%
2-4	2.9%	6.8%	2.2%	12.7%	24.7%
4-6	3.5%	4.2%	0.7%	14.6%	23.0%
6-9		3.5%	1.7%		5.3%
9-11	7.6%	4.4%	0.1%	20.4%	32.6%
11-16	1.0%	1.4%	0.3%	2.6%	5.4%
16-30				2.5%	2.5%
Total	**15.0%**	**24.9%**	**7.3%**	**52.9%**	**100.0%**

d) Total number of trades

residual life (years)	pre-benchmark	post-benchmark	non-strippable	benchmark	total
0-2		3.4%	1.4%		4.8%
2-4	1.8%	6.5%	1.9%	12.5%	22.6%
4-6	3.5%	5.1%	0.8%	15.4%	24.8%
6-9		3.7%	1.7%		5.4%
9-11	6.2%	4.8%	0.1%	21.9%	33.0%
11-16	1.3%	1.4%	0.3%	3.2%	6.1%
16-30				3.2%	3.2%
Total	**12.8%**	**24.9%**	**6.2%**	**56.2%**	**100.0%**

For each day in the sample bonds are classified according to their residual life and category. Panel (a) reports the average number of bonds in each class. In the remaining panels, activity measures are added for the bonds included in each class. Values are expressed in percentage terms with respect to the whole market. Figures in this table are averages of the daily data from January 1999 to April 2000.

Table 6

Liquidity measures for the Spanish Government Bonds

a) Number of transactions per bond

residual life (years)	pre-benchmark	post-benchmark	non-strippable	benchmark
0-2		11.9	0.8	
2-4	4.5	17.4	1.4	43.9
4-6	6.3	9.5	0.9	54.3
6-9		13.2	1.9	
9-11	28.0	17.0	0.5	77.2
11-16	4.6	4.9	0.9	11.3
16-30				11.3
Total	**16.5**	**16.9**	**0.3**	**39.6**

b) Trading volume per bond (EUR millions)

residual life (years)	pre-benchmark	post-benchmark	non-strippable	benchmark
0-2		90.5	7.0	
2-4	50.0	106.4	9.5	258.2
4-6	49.5	55.0	5.3	295.9
6-9		72.0	11.3	
9-11	199.0	89.8	2.5	414.3
11-16	20.2	28.5	6.8	53.6
16-30				51.1
Total	**45.5**	**63.2**	**6.1**	**153.3**

c) Turnover ratio

residual life (years)	pre-benchmark	post-benchmark	non-strippable	benchmark
0-2		0.9%	0.1%	
2-4	2.1%	1.0%	0.2%	2.9%
4-6	2.5%	0.8%	0.1%	3.2%
6-9		0.4%	0.2%	
9-11	4.3%	0.7%	0.0%	4.0%
11-16	0.6%	0.3%	0.4%	0.6%
16-30				0.7%
Total	**2.4%**	**0.7%**	**0.2%**	**2.3%**

For each day in the sample bonds are classified according to their residual life and category. For each class we compute the average of the activity measures for the bonds included. Figures in this table are averages of the daily data from January 1999 to April 2000.

Table 7

Errors in the estimated zero-coupon yield curve

Issue			Average estimated yield errors (basis points)			
Coupon (%)	**Maturity Date**	**Average duration**	**Pre-bench**	**Benchmark**	**Post-bench**	**Non-strip**
12.25	25/3/00	1.02				22.9
6.75	15/4/00	1.09				20.5
10.1	28/2/01	1.47				3.1
5.0	31/1/01	1.49		0.6	– 1.0	
8.4	30/4/01	1.58				– 2.5
11.3	15/1/02	2.16				2.7
7.9	28/2/02	2.31				1.3
10.3	15/6/02	2.49				1.6
4.25	30/7/02	2.81	– 3.7	– 7.1		– 3.7
3.0	31/1/03	3.09	– 5.6	0.0		
5.25	31/1/03	3.17		– 7.8	– 3.8	
4.6	30/7/03	3.26	– 1.4			
10.9	30/8/03	3.39				0.9
10.5	30/10/03	3.53				0.8
8.0	30/5/04	4.05				– 2.2
10.0	28/2/05	4.46				2.5
4.5	30/7/04	4.52	–14.0	– 9.5	– 4.6	
3.25	31/1/05	4.86	– 4.9	– 3.4		
4.95	30/7/05	4.94	– 4.6			
10.15	31/1/06	5.05				5.2
8.8	30/4/06	5.28				2.7
7.35	31/3/07	6.01				3.2
6.0	31/1/08	6.77		3.0	3.9	
8.2	28/2/09	7.04				1.3
5.15	30/7/09	8.02	– 7.1	– 4.2		– 4.1
4.0	31/1/10	8.43	– 7.1	– 6.5		
8.7	28/2/12	8.44				5.4
6.15	31/1/13	9.48		2.1	2.2	
4.75	30/7/14	10.87	– 2.0	– 1.3		
6.0	31/1/29	14.57		0.0		
Average			**– 5.6**	**– 2.8**	**– 1.6**	**4.3**

Figures are average of daily errors. Zero coupon-yield curves are estimated with the Svensson model for the whole sample.

Table 8

Liquidity premia estimates

(basis points)

	Model 1 Basic model	Model 2			Model 3		Model 4
		γ_0	γ_1	γ_2	γ_1	γ_2	γ_2
Mean		– 3.1	4.6	6.9	5.3	8.0	6.3
Max		4.5	10.7	10.4	12.4	11.5	9.2
Min		–11.1	–6.0	2.5	– 3.7	5.1	3.4
Std Dev		4.0	3.5	1.7	3.9	1.6	1.5
% Signif (-)		54%	0%	0%	0%	0%	0%
2% Signif (+)		0%	72%	99%	76%	100%	100%
Mean RMSE	*1.61*	*0.59*			*0.79*		*1.02*
Mean error reduction		*63.35%*			*50.90%*		*36.60%*

Zero-coupon yield curves are estimated using Nelson-Siegel model until September 1999 and Svensson model for the remaining sample. Dummy variables are introduced to capture liquidity differences. The instantaneous forward rate of term *m* is modelled in the following way for Svensson model:

$$\varphi_m = \beta_0 + \beta_1 e^{-\frac{m}{\tau_1}} + \beta_2 \frac{m}{\tau_1} e^{-\frac{m}{\tau_1}} + \beta_3 \frac{m}{\tau_2} e^{-\frac{m}{\tau_2}} + \gamma_0 PREBENCH + \gamma_1 POSTBENCH + \gamma_2 NONSTRIP$$

Model 1 (basic model) does not consider any liquidity effect i.e. γ_0, γ_1 and γ_2 are restricted to zero. Model 2 estimates freely all parameters. Model 3 eliminates the dummy variable for pre-benchmark bonds, γ_0. Model 4 estimates the parameter for non-strippable bonds, and restricts to zero the parameters of the other dummy variables. Premium figures are expressed in basis points. Mean error reduction refers to the reduction of RMSE with respect the basic model.

Table 9

Impact of withholding tax on liquidity premia estimates

(basis points)

	Model 5	
	γ_1	**γ_2**
Mean	5.4	6.8
Max	12.6	10.5
Min	– 3.4	3.8
Std Dev	4.0	1.7
% Signif (-)	0%	0%
% Signif (+)	78%	100%
Mean RMSE	0.82	
Mean error reduction	49.06%	

Zero-coupon yield curves are estimated using Nelson-Siegel model until September 1999 and Svensson model for the remaining sample. Dummy variables are introduced to capture liquidity differences. The instantaneous forward rate of term *m* is modelled in the following way for Svensson model:

$$\varphi_m = \beta_0 + \beta_1 e^{-\frac{m}{\tau_1}} + \beta_2 \frac{m}{\tau_1} e^{-\frac{m}{\tau_1}} + \beta_3 \frac{m}{\tau_2} e^{-\frac{m}{\tau_2}} + \gamma_1 POSTBENCH + \gamma_2 NONSTRIP$$

Model 5 is the same as Model 3 but taking into account withholding tax effects for non-strippable bonds. Liquidity premium figures are expressed in basis points. Mean error reduction refers to the reduction of RMSE with respect the basic model (Model 1).

Table 10

Impact of taxes on liquidity premia estimates

(basis points)

	Model 6	
	γ_1	**γ_2**
Mean	3.3	5.0
Max	8.1	7.4
Min	– 2.3	3.3
Std Dev	2.6	1.0
% Signif (-)	0%	0%
% Signif (+)	76%	100%
Mean RMSE	0.37	
Mean error reduction	77.01%	

Zero-coupon yield curves are estimated using Nelson-Siegel model until September 1999 and Svensson model for the remaining sample. Dummy variables are introduced to capture liquidity differences. The instantaneous forward rate of term *m* is modelled in the following way for Svensson model:

$$\varphi_m = \beta_0 + \beta_1 e^{-\frac{m}{\tau_1}} + \beta_2 \frac{m}{\tau_1} e^{-\frac{m}{\tau_1}} + \beta_3 \frac{m}{\tau_2} e^{-\frac{m}{\tau_2}} + \gamma_1 POSTBENCH + \gamma_2 NONSTRIP$$

Model 6 is the same as Model 3 but using after-tax cash flows. Liquidity premium figures are expressed in basis points. Mean error reduction refers to the reduction of RMSE with respect the basic model (Model 1).

Table 11

Relative liquidity premia with and without taxes

(percentage points)

	γ_1/β_0	**γ_2/β_0**	**$\gamma_1/(\beta_0+\beta_1)$**	**$\gamma_2/(\beta_0+\beta_1)$**
Model 3, before-tax yield curve	0.85	1.30	2.07	3.05
Model 6, after-tax yield curve	0.83	1.26	1.98	2.93

Table 12

Estimated parameters in different term structure models

		β_0		β_0		β_0		τ_0		β_0		τ_0	
		mean	**sd**	**mean**	**sd**	**mean**	**sd**	**mean**	**sd**	**mean**	**sd**	**mean**	**sd**
Model 1	4/1/99-30/9/99	6.23	0.18	–3.74	0.15	–0.66	1.16	4.89	2.29				
	1/10/99 - 14/4/00	6.42	0.23	–3.28	0.59	40.64	0.37	2.96	1.22	–40.42	0.36	3.03	1.27
Model 2	4/1/99-30/9/99	6.18	0.27	–3.60	0.33	–1.89	1.57	3.34	0.72				
	1/10/99 - 14/4/00	6.43	0.24	–3.42	0.64	40.72	1.15	2.33	0.43	–40.89	1.30	2.39	0.43
Model 3	4/1/99-30/9/99	6.29	0.20	–3.80	0.18	–1.22	1.46	4.42	1.97				
	1/10/99 - 14/4/00	6.42	0.24	–3.42	0.64	40.89	1.50	2.31	0.45	–41.06	1.64	2.36	44.32
Model 4	4/1/99-30/9/99	6.27	0.20	–3.79	0.18	–1.01	1.39	4.57	2.02				
	1/10/99 - 14/4/00	6.40	0.24	–3.36	0.65	40.62	1.02	2.31	0.74	–40.82	1.07	2.36	0.75
Model 5	4/1/99-30/9/99	6.29	0.20	–3.82	0.18	–1.18	1.47	4.47	1.99				
	1/10/99 - 14/4/00	6.42	0.24	–3.44	0.63	40.42	0.27	2.23	0.43	–40.63	0.27	2.28	0.43
Model 6	4/1/99-30/9/99	3.92	0.15	–2.28	0.14	–0.70	1.19	4.01	1.52				
	1/10/99 - 14/4/00	4.11	0.14	–2.16	0.40	40.46	0.16	2.06	0.40	–40.58	0.16	2.09	0.40

Zero-coupon yield curves are estimated using Nelson-Siegel model until September 1999 and Svensson model for the remaining sample. Dummy variables are introduced to capture liquidity differences. The instantaneous forward rate of term *m* is modelled in the following way for Svensson model:

$$\varphi_m = \beta_0 + \beta_1 e^{-\frac{m}{\tau_1}} + \beta_2 \frac{m}{\tau_1} e^{-\frac{m}{\tau_1}} + \beta_3 \frac{m}{\tau_2} e^{-\frac{m}{\tau_2}} + \gamma_0 PREBENCH + \gamma_1 POSTBENCH + \gamma_2 NONSTRIP$$

Model 1 (basic model) does not consider any liquidity effect i.e. γ_0, γ_1 and γ_2 are restricted to zero. Model 2 estimates freely all parameters. Model 3 eliminates the dummy variable for pre-benchmark bonds, γ_0. Model 4 estimates the parameter for non-strippable bonds, and restricts to zero the parameters of the other dummy variables. Model 5 is the same as Model 3 but taking into account withholding tax effects for non-strippable bonds. Model 6 is the same as Model 3 for after-tax cash flows. Interest rates are expressed in percentage points.

Table 13

Biases in quoted prices

Issue			Average differences in yield (basis points)			
Coupon (%)	**Maturity Date**	**Average duration**	**Pre-bench**	**Benchmark**	**Post-bench**	**Non-strip**
12.25	25/3/00	1.02				1.0
6.75	15/4/00	1.09				1.2[1]
10.1	28/2/01	1.47				1.0[1]
5.0	31/1/01	1.49		0.2	0.1	
8.4	30/4/01	1.58				1.2[1]
11.3	15/1/02	2.16				0.2
7.9	28/2/02	2.31				0.2
10.3	15/6/02	2.49				0.4
4.25	30/7/02	2.81	1.0[1]	0.0	0.1	
3.0	31/1/03	3.09	1.9[1]	0.8[1]		
5.25	31/1/03	3.17		– 1.1	0.2	
4.6	30/7/03	3.26	0.0			
10.9	30/8/03	3.39				0.6[1]
10.5	30/10/03	3.53				0.4
8.0	30/5/04	4.05				0.2
10.0	28/2/05	4.46				0.1
4.5	30/7/04	4.52	– 0.1	0.0	0.3	
3.25	31/1/05	4.86	2.3[1]	0.5		
4.95	30/7/05	4.94	1.2[1]			
10.15	31/1/06	5.05				– 0.2
8.8	30/4/06	5.28				0.0
7.35	31/3/07	6.01				0.2
6.0	31/1/08	6.77		– 0.4	0.1	
8.2	28/2/09	7.04				1.1[1]
5.15	30/7/09	8.02	– 0.3	– 0.2	0.4	
4.0	31/1/10	8.43	0.3	0.6[1]		
8.7	28/2/12	8.44				0.9[2]
6.15	31/1/13	9.48		0.0	0.4	
4.75	30/7/14	10.87	– 0.2	0.0		
6.0	31/1/29	14.57		0.0		
Average			0.3	– 0.1	0.2	0.3

[1] Significant at 5%. [2] Significant at 10%

For each day and security we compute the difference in yield between the mean traded price and the mid quoted price. Figures are daily averages for the whole sample (January 1999 to April 2000).

Figure 1
Life cycle for strippable 10-year bonds included in the sample

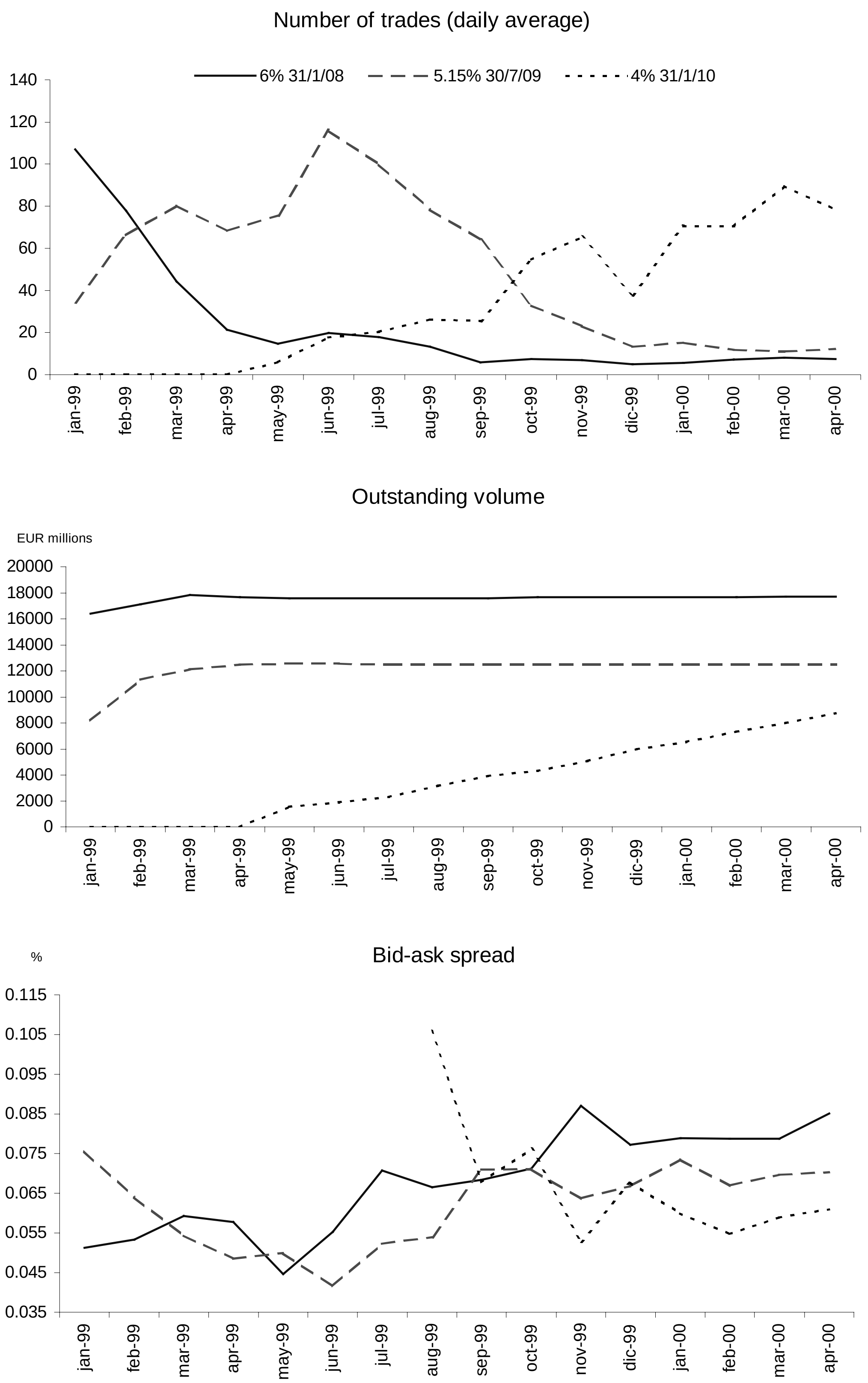


Figure 2
Average of daily bid-ask spread
(percentage points)

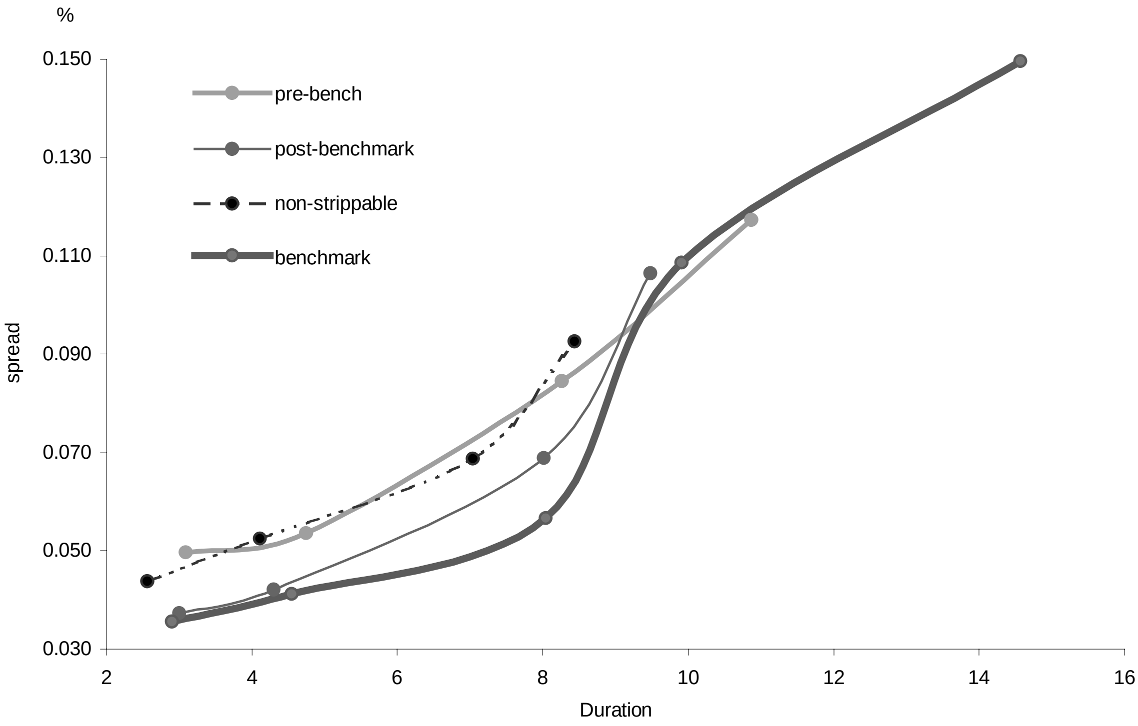


Figure3
Zero-coupon yield curve
(Nelson-Siegel model, 29 March 1999)

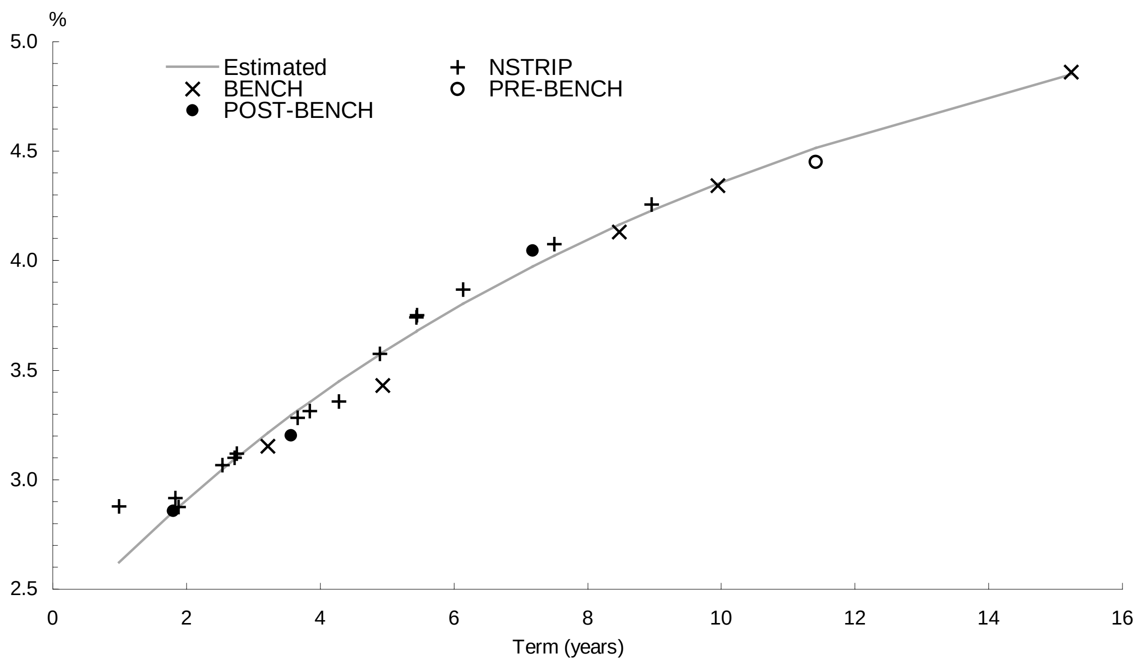
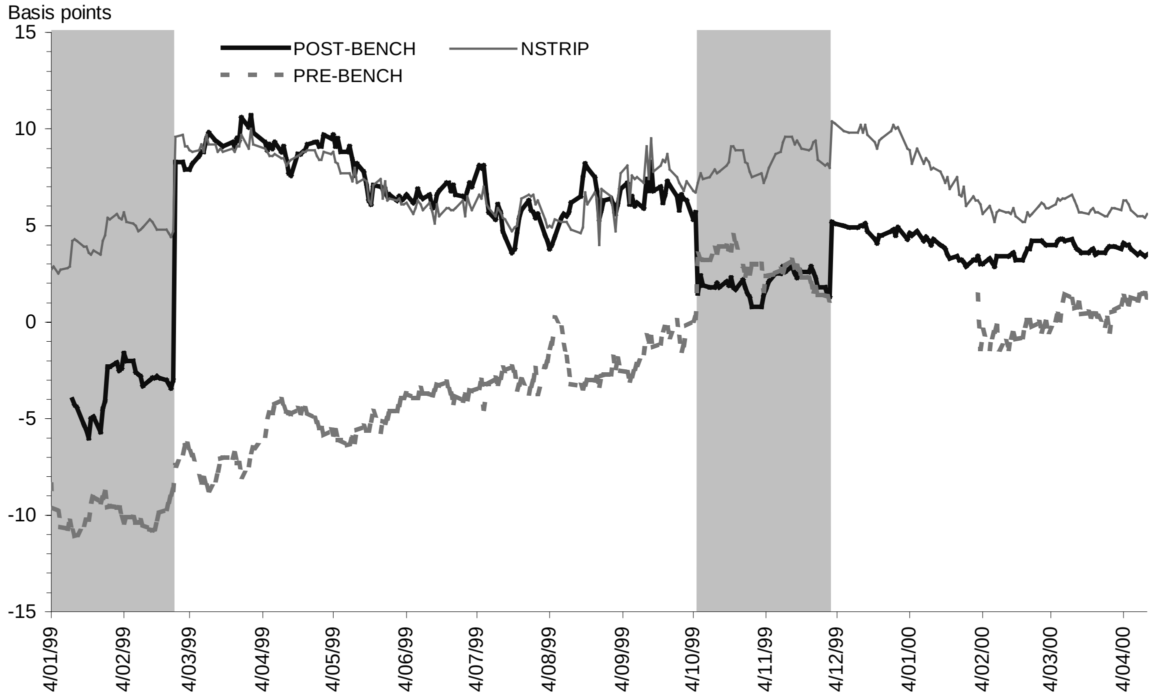


Figure 4
Evolution of estimated parameters



Zero-coupon yield curves are estimated using Nelson-Siegel model until September 1999 and Svensson model for the remaining sample. Dummy variables are introduced to capture liquidity differences. The instantaneous forward rate of term *m* is modelled in the following way for Svensson model:

$$\varphi_m = \beta_0 + \beta_1 e^{-\frac{m}{\tau_1}} + \beta_2 \frac{m}{\tau_1} e^{-\frac{m}{\tau_1}} + \beta_3 \frac{m}{\tau_2} e^{-\frac{m}{\tau_2}} + \gamma_0 PREBENCH + \gamma_1 POSTBENCH + \gamma_2 NONSTRIP$$

This estimation corresponds with Model 2. Shaded areas indicate that γ_2 is not statistically significant.

Figure 5

Zero-coupon yield curve

(Nelson-Siegel model, 29 March 1999)

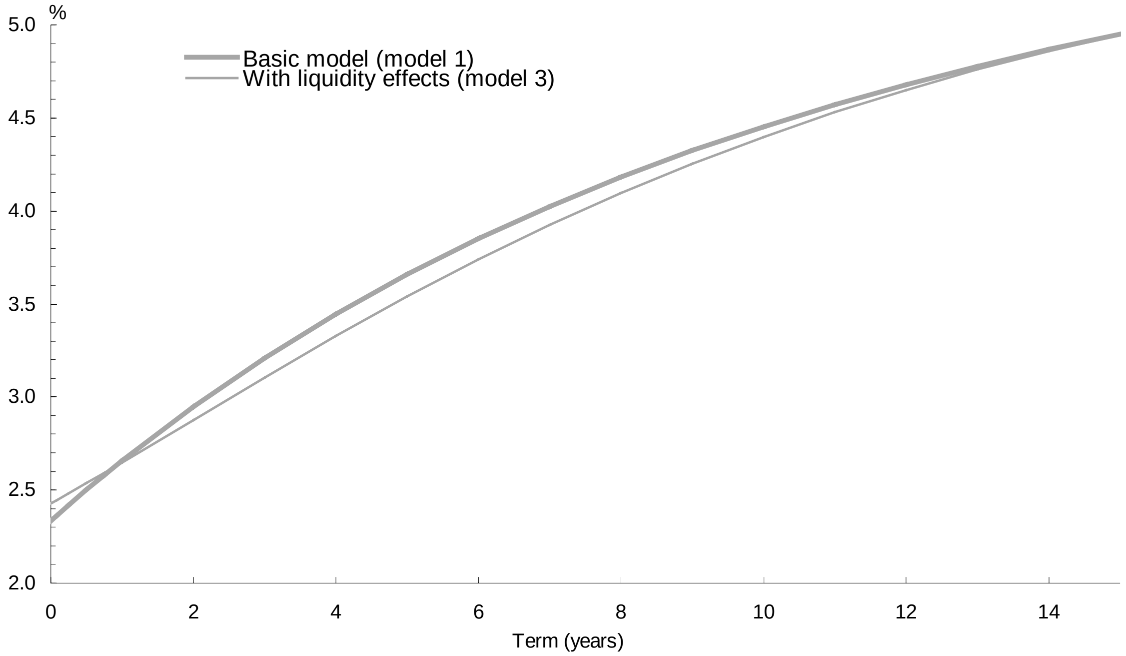


Figure 6

One-year forward curve

(Nelson-Siegel model, 29 March 1999)

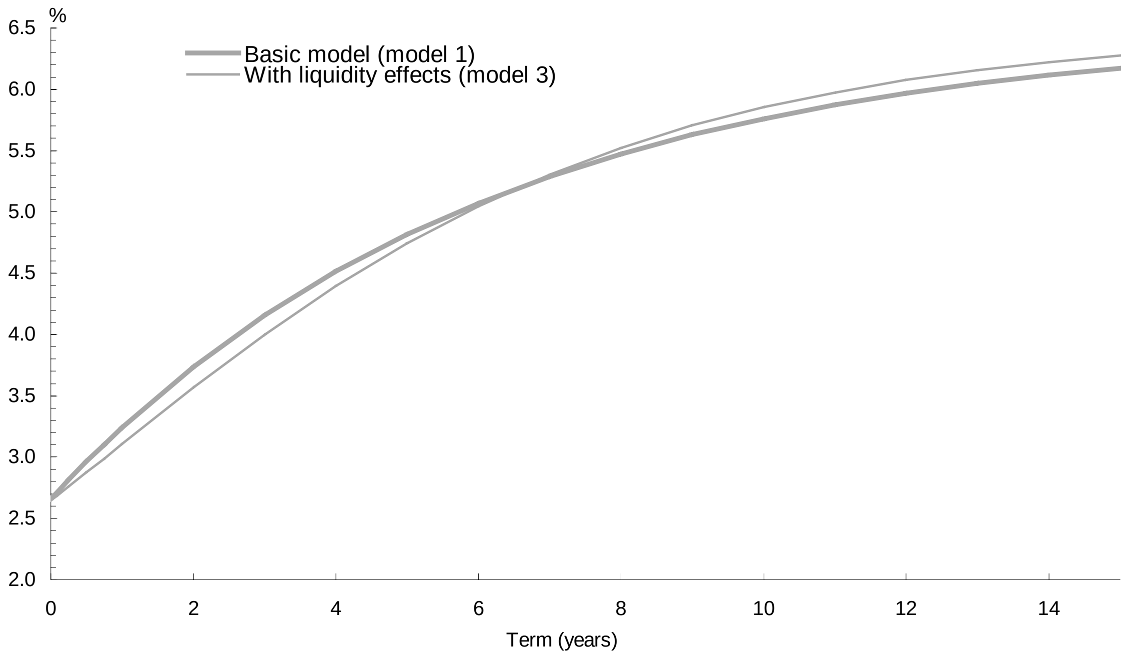


Figure 7
Zero-coupon yield curve
(Svensson model, 31 March 1999)

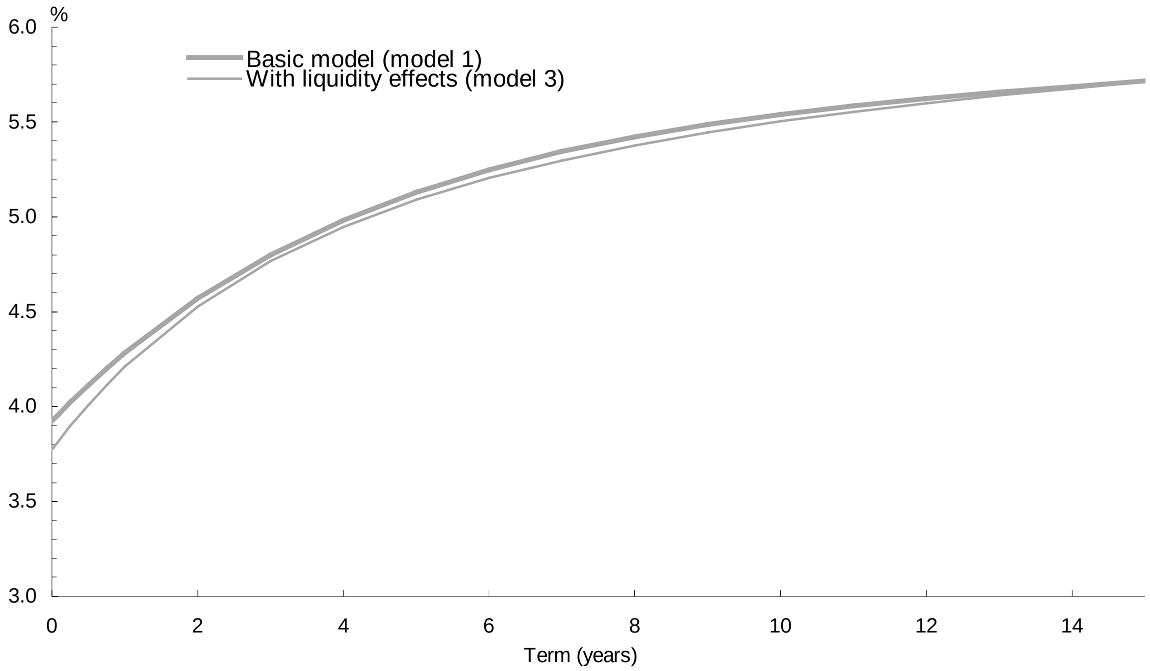
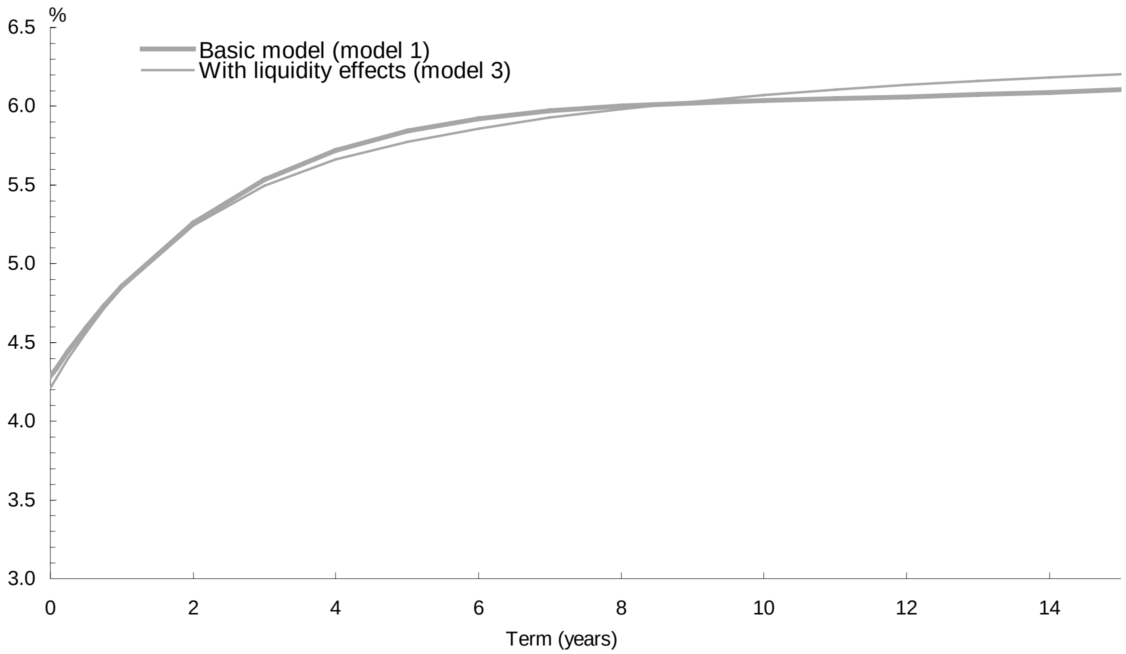


Figure 8
One-year forward curve
(Svensson model, 29 March 1999)



Comments on "Estimating liquidity premia in the Spanish Government securities market" by F Alonso, R Blanco, A del Río and A Sanchís

Christian Upper, Deutsche Bundesbank

Liquidity premia for fixed income securities have commanded increasing attention since their sharp increase during the summer of 1998. They normally obtained by comparing the yields of two bonds with similar maturity and default risk but different liquidity. Blanco et al follow a more sophisticated course. Instead of pairwise comparisons of two securities, they estimate a zero-coupon yield curve for Spanish government securities and include dummy variables for three categories of bonds which they deem to be less liquid. The categories are non-strippable post-benchmark bonds, strippable post benchmark bonds and strippable pre-benchmark securities. The coefficients of the dummies give an estimate for the premium commanded by the corresponding category over a hypothetical benchmark security of the same maturity.

Yield spreads may arise due to a variety of factors, and liquidity differences is only one of them. The authors follow a two-step strategy in order to make sure that what they estimate does indeed reflect liquidity premia and not something else. They begin by showing that market liquidity does vary across the four categories of bonds. For my taste, they rely too much on activity indicators such as trading volume or turnover ratios. Such variables show that the market was liquid enough in order to handle the amount of activity observed, but not how easily securities could be traded. However, given the scarcity of readily available measures for market liquidity, the use of activity indicators may be unavoidable. Nonetheless, perhaps the fact that bid-ask spreads increase with maturity does indicate that shorter term paper is more liquid than longer term paper, although they will not be able to identify this with their methodology. As a second step, the authors rule out the main alternative factor – differences in taxation - by repeating the estimation with post tax returns. They find that the estimates for the coefficients barely change. Hence the interpretation of the yield spread as a liquidity premium seems acceptable.

The authors find that non-strippable bonds (all bonds issued before July 1997) and post-benchmark bonds command a yield premium in the range of 5 to 10 basis points over the benchmark, while pre-benchmark issues (on-the-run issues which eventually will become benchmarks) show a negative yield spread over benchmark paper, although this is significant only about one half of the time.

Let me first make some remarks on the methodology before discussing the results. The procedure used requires considerably more effort than simple pairwise comparisons. So what is the value-added and is it worth the effort?

Firstly, by using a yield curve as a baseline, the authors reduce the importance of pricing errors for individual bonds, e.g. due to non-synchronised trading. Such ‚noise' should be present mainly in the prices of the less liquid bonds, so that the gain is not so much by using a better benchmark but by aggregating over several bonds in the individual categories. Secondly, using a yield curve reduces the bias that arises because of differing maturities that are sometimes unavoidable in pairwise comparisons. This should improve the accuracy of the estimates in particular for shorter maturities, where the yield curve tends to be steeper. Thirdly, their approach allows them to test whether the estimates for the yield curve are sensitive to the presence of liquidity premia. Thankfully, they find that they are not. This should be of great relieve to all those users of yield curves who do not adjust for liquidity differences.

A final advantage is that their approach allows them to distinguish between the relative importance of different attributes of a bond in determining the liquidity premium. This would enable us to break down the liquidity premium into different components. Unfortunately, the authors do not make much out of this possibility since the categories into which they group the bonds are mutually exclusive. For example, they do not included the non-strippable bonds in the post-benchmark category. As a consequence, we don't know whether the liquidity premium of non-strippable bonds over strippable post-benchmark securities is statistically significant. Using non-exclusive categories would permit them to identify the marginal effects of each attribute.

As in any empirical paper, the question arises whether the variables (in their case the dummies) included are the correct ones. Now I don't know much about the market for Spanish government securities beyond what I learned in the paper, so I'm not qualified to make judgements on this issue. However, I wonder why they omitted the volume outstanding of the securities. In Germany, the on-the-run/off-the-run spread is quite sensitive to the size of the issues. Is this not the case in Spain? Another possibility would be to include measures of liquidity in a more direct manner. The authors are right to point out that the relationship between liquidity premia and liquidity itself is likely to be non-linear, but this could be dealt with by using dummies e.g. for low/average/high liquidity. Of course, one always has to bear in mind that with only about 30 bonds at any point in time, the number of variables that one may include is rather limited.

Let me now briefly comment on the results. Figure 4 of the paper shows the evolution of the estimated liquidity premia over time. Two rather distinct periods can be identified, separated by two months in which the coefficient of POST-BENCH turned out to be insignificant. Between March and September 1999, the coefficients of NSTRIP and POST-BENCH are virtually identical. As I mentioned before, using non-exclusive categories could have told us whether this was statistically significant. In the second period, from December 1999 to April 2000, the two coefficient differ by a much greater extend. That of POST-BENCH declines from 5 to 3 basis points, while that of NSTRIP halves from 10 to 5 basis points. Unfortunately, the paper does not discuss the evolution of liquidity premia over time, although this may be due to its more methodological focus. All in all, the movement over time of the estimated premia is rather different from the pattern followed by a (admittedly very simple) measure of liquidity spreads in the German market, which has seen a sharp rise in liquidity premia during the summer of 1999. It would be interesting to know why.

Let me conclude by saying that there is a lot of value added in their approach. I found it a very inspiring paper and I'm looking forward to follow ups. There are two extensions in which I would be particularly interested. The first would be to go beyond the Spanish market and consider yield differences between bonds of different nationalities. At the beginning of August, Spanish 10 year benchmark bonds yielded 18 basis points and Portuguese and Italian bonds 26 basis points more than German paper. Why is this the case? Which role does the presence of a liquid futures market for the German Bundesanleihe play? These are questions which wait to be answered, although admittedly the task of putting together a EMU-wide dataset seems aweinspiring. The second extension would be to include credit spreads into the analysis. So far, it has been difficult to disentangle the effect of creditworthiness and liquidity. For example, at the Bundesbank, we have argued that the yield spread of German bank bonds over Bunds mainly reflect liquidity rather than credit premia, but we haven't been able to test this in a rigorous manner. With your methodology, this could be done.

Comments on "Estimating liquidity premia in the Spanish government securities market" by F Alonso, R Blanco, A del Río and A Sanchís

Oreste Tristani, European Central Bank[1]

This is an interesting paper that presents quantitative estimates of the premium implicit in the price of Spanish government securities to compensate investors for liquidity risk and tax factors. The main results are obtained fitting a yield curve model through bond prices, measured daily, over the maturity spectrum. The estimation is repeated for the number of available days, from January 1999 to April 2000. Liquidity premia are identified through a number of dummy variables attached to less liquid bonds, i.e. issues that are not on-the-run.

The results show that the latter category of bonds appears to command a relatively small premium, of the order of a few basis points, broadly comparable to the estimates obtained by Elton and Green (1998) for US government securities. This is reassuring, since it confirms that in developed economies, excluding times of market stress, liquidity premia on government securities can be considered as negligible in terms of their macroeconomic implications. Nevertheless, their precise measurement can be useful in the management of government debt, for example allowing a more informed assessment of the relative desirability of alternative issuing practices.

The authors (henceforth ABRS) also mention that knowing the size of liquidity premia on government bonds is important for central banks, when they try to extract macroeconomic information (mainly on inflation expectations) from the yield curve. This activity crucially involves filtering out from the curve any kind of risk-premia, amongst which liquidity premia, that are unrelated to macroeconomic factors. The more precise the technique adopted to implement the filtering, the more precise the estimation of macroeconomic information reflected in the yield curve.

To filter out liquidity premia, central banks typically concentrate on benchmark issues. However, this methodology is relatively crude, since liquidity premia (compensation for the lack of immediacy, depth, tightness and resiliency of the market) are likely to characterise all bonds, not just the *relatively* more illiquid ones. The assumption that benchmark bonds are not just *more* liquid, but completely (liquidity) risk-free is not necessarily warranted. To improve the precision of the measurement of inflationary expectations implicit in the yield curve, one would really need an estimate of the liquidity premium implicit in the return on benchmark issues.

Unfortunately, ABRS' methodology does not provide indications in this respect, but their exercise remains interesting for the other reasons emphasised above.

Is the Svensson model precise enough?

As already mentioned, the magnitude of the premia measured by ABRS is "moderately" small, of the order of a few basis points. In this respect, ABRS' paper is a commendable effort to look for refinement and precision in the measurement and estimation of yield curves. A fundamental choice in the analysis is therefore the selection of the yield curve model. As is often the case in economics, we have many available models to fit one phenomenon, and no model appears to unequivocally dominate the others under all circumstances.

ABRS dismiss altogether dynamic models, presumably because, although theoretically more appealing, their simpler versions do not provide an accurate description of empirical yield curves. The choice falls instead on static "estimated" models, such as simple polynomial splines (e.g. the cubic one suggested by McCulloch, 1971, 1975), which have the advantages of being both manageable and to only require estimation of a small number of parameters to fit the data. Among the available static models, ABRS favour Svensson's (1994) modification of the Nelson and Siegel (1987) model. The main advantage of the Nelson and Siegel model with respect to simple polynomial splines is that the latter, as noted by Shea (1984), tend to produce unstable forward rates, especially at longer

[1] The opinions expressed are personal and should not be attributed to the European Central Bank.

maturities. Nelson and Siegel's formulation, on the contrary, imposes an asymptote on the forward rates curve, so that the latter always stabilises as the maturity increases. Svensson's modification of this model essentially represents a way to increase the degrees of freedom available in the estimation, so that a larger number of shapes for the yield curve can be accommodated.

Though more stable at long maturities Nelson and Siegel's methodology is not, however, "perfect". Like spline models, it essentially amounts to fitting a certain functional form through bond prices. As in any estimation, there will typically be a discrepancy between actual and fitted values. As a result of Svensson's modification, the functional form used by ARBS is very flexible and it can accommodate a large number of shapes for the yield curve. Nevertheless, it will not, excluding exceptional circumstances, be able to replicate *exactly* observed prices. In fact, it is not uncommon for it to produce fitted yields to maturity that are a few basis points different from observed yields.

Whether this imperfect fit is acceptable will ultimately depend on the scope of the analysis. When zero coupon rates and long-term forward rates are used to gauge average future inflationary expectations, a discrepancy between actual and fitted values of up to 10 basis points can probably be considered negligible. Indeed, Svensson takes the view that "for financial analysis, the estimation of forward rates is done with a number of different methods, some rather complex to achieve sufficient precision. For monetary policy analysis, the demand for precision is arguably less, which can be traded for increased robustness and simplicity of the estimation method".

Is the extended Nelson and Siegel model sufficiently precise for the estimation of liquidity premia? I believe that the answer is not obvious, since the magnitude of pricing errors generated by the Svensson model is comparable to that of the liquidity premia resulting from ABRS' estimation. Indeed, liquidity premia are defined as the discrepancy between the observed price of an illiquid bond and the estimated price of a perfectly liquid theoretical bond with the same maturity and coupons. An insufficiently precise estimate of the latter would obviously bias the calculation of the premia.

The authors might therefore want to spend some time convincing the reader that it is indeed liquidity premia that they are measuring, and not (at least in part) a mispricing error due to an insufficiently precise model. Since they repeat their yield curve estimation for a large number of days, thus a large variety of shapes for the yield curve, it should be possible to argue that any potential mispricing error is likely to have zero mean over the days contained in the sample. Hence, even if daily premia could occasionally be affected by a pricing error, the average of these premia over the sample period is likely to genuinely represent the compensation for liquidity risk requested by investors.

An alternative model

There are, however, alternative methodologies, such as that presented in Brousseau and Sahel (1999), that allow for a more precise replication of the yield curve.

The fundamental difference between the Nelson and Siegel methodology and that adopted by Brousseau and Sahel is that the latter provides an exact representation of the yield curve, and not a statistical approximation of it. Rather than postulating a certain functional form for the curve, and then fitting it to the data, Brousseau and Sahel use a non-parametric procedure that allows the shape of the curve to be (almost) completely data-determined.

This *bootstrapped* yield curve is defined as one in which the observed prices of all the available securities are replicated without any pricing error. Since only a finite number of bonds is available for the construction of a continuous curve, the maturities for which no traded bonds are available must, like in the other methodologies, be interpolated. In this case, however, the interpolation is such as to leave the estimates of prices of existing securities equal to their observed values.

A bootstrapped curve obviously bears a more complex mathematical representation than one obtained through, for example, the Svensson methodology. Specifically, it cannot be represented through a small number of parameters and it can typically only be obtained numerically. Brousseau and Sahel construct the curve in an iterative fashion, starting from the available security with the shortest maturity and then moving on by increasing maturity. For illustrative purposes, they concentrate on the US and German securities markets and report the results of some bootstrapped curves obtained assuming different sorts of interpolation methods (see Brousseau and Sahel, for details). They then compare these results with those estimated through a simple spline model.

Interestingly, the difference between the yield curves obtained through the two methodologies turns out to be, on average, of the order of a few decimals of a percentage point, i.e. potentially large enough to produce significant differences in the estimation of liquidity premia. This seems to confirm that some caution is necessary in the interpretation of ABRS's results. It must be taken into account, however, that the comparison presented in Brousseau and Sahel (1999) is with respect to a simple spline model. Compared to the latter, Svensson's model is likely to improve significantly the fit of the observed price data, so that it remains unclear whether its pricing errors can be considered negligible for the purpose of estimating liquidity premia.

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Does market transparency matter? A case study

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Abstract

We analyse a change in the degree of transparency of MTS, the electronic inter-dealer market for Italian Government bonds, namely the July 1997 move to the anonymity of quotes. Our evidence supports the hypothesis that a decrease in transparency makes liquidity traders worse-off, whereas large/informed traders find it less costly to execute block trades. The evidence is also consistent with the "waiting game" hypothesis of Foster and Viswanathan (1996): under anonymity, traders tend to delay their trades in an attempt to acquire information through the order flow.

From a public welfare perspective, our results indicate that the move to anonymity has been accompanied by an increase in market liquidity and by a reduction in volatility, a phenomenon that is also partly explained by the growth in Italy's prospects for early participation in the EMU. The speed of information aggregation on MTS increases, as shown by an improvement of the MTS lead over the futures market. In a European perspective, the current organisation and performance of MTS place the market in a competitive position with respect to other sovereign bond markets and may contribute to their integration under the single currency.

1. Introduction[2]

The electronic inter-dealer market for Italian Government bonds, known as MTS (from ***Mercato Telematico dei Titoli di Stato***), is characterised in international comparison by a high degree of transparency (Inoue 1999). In July 1997, 10 years after its inception, MTS switched to a new operation regime in which the names of market-makers who post bid and ask quotes for each security are not revealed.

The switch seems worth investigating because it prompts a number of interesting questions for financial economists and regulatory authorities. What was the reason for the switch? Who benefited from it? How did it affect market performance, in terms of liquidity, efficiency and price volatility? Is market microstructure theory consistent with the evidence? Has the switch altered the way Italian T-bonds are traded on MTS as compared to the over-the-counter market? Can we derive any regulatory policy lessons from the experience of MTS? What are the implications for the development of an integrated sovereign bond market in the European single-currency area?

Market microstructure theory shows that the existence of information asymmetries among participants is a key element in understanding how a financial market is organised and works. If the market is populated by two types of agents with different information endowments and objective functions, the better-informed and the liquidity-motivated, then a given market set-up may be optional for one group but, generally, not for the other. Similarly, a change in the set-up may benefit one group at the expense of the other. The dynamic relationship between the two groups has implications for the consolidation or fragmentation of trading in different markets and for asset price volatility. It can be argued that in a bond market, like ours, the absence of "inside" information on an asset's fundamental value reduces the scope for heterogeneity of beliefs with respect to a stock market. However, we observe that the concept of private information must also include knowledge that dealers may acquire

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on the order flow and on the trading intentions of large customers, an argument which applies to the bond and foreign exchange markets as well as to the stock market. This knowledge causes an update of beliefs and may be profitably exploited at the expense of other market participants, according to a notion that is also at the basis of the literature on block trading and dual-trading. From an empirical viewpoint, some studies support this hypothesis in the forex market and in the bond market (see Lyons 1995 and Scalia 1998a), although there is also evidence to the contrary (Proudman 1995).

The models of information asymmetry point to one conclusion: liquidity traders in general prefer more transparency, informed traders prefer less transparency. This notion was pioneered by the Grossman (1998) model of sunshine trading. Sunshine trading, ie disclosing pre-trade information on the direction of price-contingent orders, removes the possibility that those orders are information-based and thus eases the inference problem of market-makers. This should lower execution costs for sunshine traders and possibly increase overall trading volume. Forster and George (1992) explore the effects of various degrees of traders' anonymity on the distribution of wealth within the two groups of market participants. They conclude that if monitoring who is trading in a centralised system gives information on liquidity trades, then disclosing the identity of current participants lowers execution costs for liquidity traders, provided that there is sufficient competition among privately informed agents. This clearly reduces the expected profit of informed traders. The analysis of Pagano and Röell (1996) provides similar results: in a comparison of alternative trading systems, it is shown that greater transparency, such as that provided by a centralised order execution system with full disclosure, reduces the average trading cost for liquidity traders. Madhavan (1995) examines the issue of post-trade information disclosure and market fragmentation vs consolidation in a two-period dynamic model. The model provides an unambiguous prediction on the implication of different disclosure rules for informed traders and "large" (or strategic) liquidity traders: they should prefer non-transparency because it facilitates dynamic trading strategies, like "working" a large order over time. Without mandatory disclosure, dealers also prefer not to disclose trades voluntarily because they profit from the reduction in price competition. Naik, Neuberger and Viswanathan (1994) investigate the relationship between delayed trade disclosure rules and execution costs in a dynamic market setting with risk averse dealers. If there are two stages of trading, first a public investor who trades with market-maker A, and then A who trades with other competing market-makers, a delayed disclosure rule of the first-round trade by A grants him the possibility in the second round to exploit the information conveyed by the trade itself. In turn, A passes on part of the associated profit to the public investor. The authors stress the fact that their conclusion has a more general bearing: any time less-than-full disclosure of large and informative orders occurs, the dealers who intermediate the order and their customers should be better-off.

Our summary of models that explore the effects of information asymmetry and market transparency is far from exhaustive. However, theory provides an unambiguous prediction in our context: under assumption that significant informational asymmetries exist, the switch that took place on MTS in July 1997 should have shifted the balance between liquidity traders' and informed/large traders' profits in favour of the latter. For the purpose of the tests to be conducted in the following sections, we set forth two hypotheses:

- Hypothesis I: liquidity traders have been made worse-off by the move to anonymity (we shall call this hypothesis the "liquidity trader's curse").
- Hypothesis II: (the "large trader's blessing"): large/informed traders have been made better off.

Our brief survey also suggests a third implication, related to the previous ones. By not disclosing the names of market-makers, the 1997 switch has made the structure of MTS more similar to that of the over-the-counter inter-dealer broker market, where dealers negotiate trades without revealing their identity. We should consider the possibility that dealers in the opaque over-the-counter market (trading either through a broker or vis-à-vis) benefit from the price discovery function of the highly transparent MTS, thus free-riding on the information disseminated by the latter (see also Madriagal 1996). Hence, before the 1997 shift there would have been two types of free-riding. The first would have been among MTS members, and it is captured by Hypothesis I. The second would have been by the OTC market at the expense of MTS. If the switch to anonymity has reduced the second type of free-riding, making MTS more similar to the OTC market, then the incentives for the informed/large dealers to trade over-the-counter rather than on the regulated market have declined. We have the following hypothesis:

- Hypothesis III ("decline of OTC free-riding"): trading volume on the OTC market has fallen since the MTS shift.

The events that preceded the market move seem broadly consistent with Hypothesis I-III. At the end of 1996 the proposal of anonymity was put forward by a group of MTS specialists (which we may assimilate to the informed/large traders of theory), led by one with foreign affiliation. The main argument advanced by the proponents was that the shift would increase the welfare of the most skilled market players, thus enhancing competition and market efficiency. In fact, the proponents' complaint about the regime of full transparency was that it allowed small dealers to mirror the moves of the big players. Understandably, some small traders had reservations. The treasury and the Banca d'Italia raised no objections. In the end the management board (in which small dealers are lowly represented) approved the proposal, which became effective on 14 July 1997.

The MTS switch of 1997 is also interesting for another reason. Foster and Viswanathan (1996) have explored the possibility that informed traders' signals are different, giving an incentive to informed traders to forecast the price forecasts of others. This may induce each informed trader to delay his transactions and wait for the other traders' moves to reveal more information. The Foster-Viswanathan model has the following prediction for intraday trading activity on MTS.

- Hypothesis IV (the "waiting game"): after the switch to anonymity, the increasing dispersion in traders' opinions reduced market turnover in the early stages of trading and increased it in the later stages.

The previous discussion explains the first objective of this paper. By analysing various market indicators before and after the MTS switch to anonymity, we wish to conduct a test of the four above mentioned hypotheses: the liquidity trader's curse, the larger trader's blessing, the decline of free-riding and the waiting game hypothesis. Another contribution of our test is that we use an original and extensive data set as compared to that of the existing literature.

Should the hypothesised worsening of terms for liquidity traders be the unique, or even the main concern for market regulators? This question raises the more general problem of which market design maximises public welfare, which brings us to the subject of normative economics and regulatory policy. O'Hara (1995) tries to qualify the three goals of a market set forth by Domowitz (1990). They are (i) reliable price discovery, (ii) broad-based price dissemination, and (iii) effective hedging against price risk.

Concerning the first goal, O'Hara argues that the ability to find a market-clearing price is enhanced by scale and possibly by the existence of multiple settings which suit the needs of different types of traders.

The achievement of broad-based price dissemination is a more contentious issue because the free-riding problem discussed earlier sets up a trade-off between market transparency and the price discovery process in the same market. Broadly speaking, market transparency certainly improves public welfare. However, in a dynamic setting, if the same asset is traded in market A and market B, and market A becomes more transparent, then it is possible that large/informed traders will move from A to B. In this case, it is conceivable that the price discovery process in market A will diminish, to the advantage of market B.[3] Therefore, it is not clear where on the ideal market transparency scale the costs from a reduction in informed trading will outweigh the benefits from greater information dissemination (see eg Franks and Schaefer 1995). This issue is at the heart of our investigation.

The third goal, namely hedging of price risk, reflects the concern for the market's ability to provide insurance to liquidity traders. The empirical counterpart of this goal is the minimisation of execution costs for liquidity trades and the improvement in general of market liquidity, defined as the property whereby the price impact of an order is small. O'Hara introduces a fourth goal of optimal market design:

> "(…) another, perhaps greater, function of the market that is not recognised in the working definitions given above (…) is the role of market efficiency. How well and how quickly a market aggregates and impounds information into the price must surely be a fundamental goal of market design."

[3] The above statement has no precise implications on the change in the degree of price discovery that derives from a change in the transparency of A for the whole market, given by the sum of A and B.

However, she also notes that the search for market efficiency presents two main problems. First, raising the speed of information aggregation may in principle increase price volatility, which is not desirable. Second, since market efficiency is positively related to the extent of information-based trading, which in turn generates losses for liquidity traders, the goal of efficiency may conflict with that of minimising execution costs for the uniformed. To summarise, although the issue of the optimal design of a financial market remains in O'Hara's words an open question, it seems safe to claim that, provided "sufficient" scale and transparency, the contribution of a market to public welfare should be measured along three dimensions: liquidity, volatility, efficiency (where the second variable clearly exerts a negative effect).

We observe that in the case of the government bond market the pursuit of public welfare along these lines is consistent with the objective of minimising the cost of public debt servicing and with the operating objectives of the monetary and regulatory authorities (Santini 1997): to carry out liquidity management operations that do not affect the smooth functioning of the market, to obtain information about market expectations, to improve monetary policy implementation in general, to conduct micro-prudential policy.[4]

The second objective of our empirical investigation is therefore of a regulatory nature. Because normative economics in this area does not show unambiguously what is the welfare-maximising degree of market transparency, we wish to develop a case study based on the previously noted event. We observe a change of regime in the arrangement of our market. By estimating the three performance variables defined earlier, both before and after the shift, we try to empirically assess whether it afforded a higher or lower level of welfare. We shall also try to keep into account an important macroeconomic factor that may have influenced the performance of MTS during our sample period, namely the fiscal consolidation process which brought about a sharp improvement in Italy's prospects for early participation in the European Monetary Union. To this extent, we shall provide estimates of the relative weight of the microeconomic effect on our market performance measures, as distinct from the microeconomic effect related to the shift to anonymity.

The paper proceeds as follows. Section 2 describes the main features of the market. Section 3 presents evidence and tests on Hypotheses I-IV. Sections 4-6 provide estimates and tests on market liquidity, volatility and efficiency, respectively. Section 7 discusses the empirical evidence against the background of theory, the regulatory implications and the prospects for the development of an integrated securities market in the single currency area. Section 8 summarises and concludes. An Appendix table provides a brief history of the Italian government bond market during the last decade.

2. The market

The securities listed on MTS include all recent Italian Treasury issues: the three-, six- and 12-month bills known as BOTs, the 18- and 24-month bills known as CTZs, the floating-rate notes with initial life of seven years known as CCTs, and the fixed-coupon BTPs with initial life of three, five, 10 and 30 years. The minimum order size is five billion lire, which is by far the modal trade size. Market members are of three types: specialists, primary dealers and ordinary members.[5] Specialists and primary dealers are committed to quoting firm two-way quotes on a wide range of securities, to being competitive in terms of tightness of spreads, and to maintaining a share on the primary and secondary market above a certain threshold, with stricter requirements applying to specialists.[6] Both categories may apply for bond and cash lending with the Banca d'Italia. Ordinary members can trade at the quoted prices. Specialists and primary dealers can also trade at somebody else's quotes.[7] In practice over 60% of transactions take place between two market-makers (specialists and primary dealers). In

[4] A detailed survey of market structure and regulation in government securities markets is provided by Dattels (1995).

[5] Strictly speaking, specialists are included in the class of primary dealers. Upon demand and subject to the selection criteria set by the Treasury and the Banca d'Italia, a primary dealer may be upgraded to the status of specialist. Downgraded specialists maintain the status of primary dealers.

[6] The requisites for specialists are market share above 3% on the primary market and above 1.5% on MTS. Primary dealers must maintain a minimum share of 0.5% on MTS.

[7] The July 1997 shift to anonymity was accompanied by a further innovation: all quotes at the same price made by different market-makers are aggregated, leading to an aggregate volume figure associated with each outstanding quote.

what follows we shall refer to the players who initiate a trade as "traders", without distinguishing whether they are market-makers or ordinary members.

Trading hours are from 09:00 to 17:10. The market trading mechanism is fully integrated. Each member's video-terminal serves three functions: (1) publication of pre- and post-trade information, including the five best bid and ask quotes for each security,[8] (2) trade execution at a key-press, and (3) automatic clearing and settlement onto the centralised systems for bank reserves and government bonds managed by the central bank.[9]

In the spring of 1997 the run-up to the annual review of the specialists' requisites, including a check of their market share, contributed to the growth in overall trading volume observed on MTS. Some specialists may have inflated their transactions on an intraday basis, without affecting their open positions at the end of the day, in an attempt to improve their turnover score. After the June 1997 review, the Treasury and the Banca d'Italia decided to lengthen the observation period to two years and to hold the next review in January 2000.[10] Partly as a consequence of this process, daily trading volume changed from an average 36 trillion lire in the second half of 1996 to 45 trillion in the first half of 1997; since then, it has stabilised at around 33 trillion lire.

The data-set employed in the empirical analysis of the following sections includes all MTS transactions, and the identity of the traders, in the period from 1 September 1996 to 31 May 1998. The old regime data sample runs from 1 September 1996 to 13 July 1997 (period 1). The new regime sample goes from 14 July 1997 to 31 May 1998 (period 2). The two samples are approximately equal in length, about 10½ months each. To be precise, there are 213 working days in period 1 and 221 working days in period 2.

3. Evidence on theoretical predictions

3.1 Hypothesis I - The liquidity trader's curse

The first type of evidence we should like to gather is that concerning the change, if any, in the degree of market participation by the informed/large dealers and the liquidity/small traders. To this extent, Table 1 provides summary statistics on the average number of active traders on a daily basis, ranked according to their market share, before and after the switch to anonymity. If we consider the smallest traders (below 0.1% of trading volume) we note that they decrease in number from 15 in period 1 to three in period 2. The second smallest class of traders (between 0.1% and 0.25%) decreases from 110 to 84. The third class (between 0.25% and 0.5%) increases slightly from 59 to 65 traders. The fourth class (0.5 to 1%) increases from 34 to 41 traders. Overall, if we set a threshold for "small" traders at 1%, we note that their average number decreases from 218 to 193. The two classes of the largest traders, from 1 to 2.5% and above 2.5%, both increase, with their sum going from 42 to 48 traders. The reduction in the number of small traders is also evidenced by the data on market concentration, provided in the lower part of the table. The Herfindahl concentration index of traders increases from 3.2 to 3.8%. The degree of concentration measured on the market-makers' side increases from 5.0 to 5.8% on average. The null hypothesis of equal means before and after anonymity is rejected. These results are consistent with Hypothesis I.[11]

3.2 Hypothesis II - The large trader's blessing

The greater concentration among market-makers seems consistent with the hypothesis that large traders have been made better-off. The category of informed and/or large traders can also be detected ex post based on the occurrence of large trades. On MTS a "block trade" as such hardly ever occurs. Due to the prudence of market-makers who post firm quotes, also in terms of size, 99% of transactions

[8] Prices are quoted clean, as a percentage of par value.

[9] Further details on the functioning of MTS can be found in Banca d'Italia (1994).

[10] The switch to anonymity was also viewed as a measure to avoid the inflation of trading volume.

[11] The daily behaviour of the concentration indices has a mixed pattern. It is relatively stable in the first half of each year, but it tends to increase slightly towards year-end.

occur at or below five times the minimum size of five billion lire. Traders wishing to exchange a large amount of bonds respond to this behaviour by working the order over time. This would suggest to proxy large trades by tracking down the continuations of trades made by the same trader on the same bond on each working day. Things are complicated, however, by the fact that during our sample period a "race for volume" took place (see previous section), and many trades were inflated, ie offset by trades of opposite sign within the same day. To control for this phenomenon, we proxy large trades as follows. Within each working day, we compute the net daily change in each trader's holdings of each listed bond. When the net change in absolute terms is larger than a given threshold, we count one "large trade" for each continuation of trades above the same threshold. Our working variable is then defined as the ratio of large trades so defined to total daily volume.

The evidence is presented in Figure 1, Panel A, for a threshold of 50 billion lire, and Panel B, for a threshold of 100 billion lire. Each panel reports the daily series of the large trade ratio and an interpolating function. In the case of a threshold of 50 billion lire, the ratio generally lies between 10% and 20%. Panel A shows that the series increases from period 1 to 2, and the tests of equal mean and of equal distribution across periods are rejected. The series obtained with a threshold of 100 billion lire generally lies between 0% and 10%. The evidence across periods in analogous: the ratio increases from period 1 to period 2, indicating that, under anonymity, it has become easier to build/unwind large positions on MTS, and the tests of equal mean and distribution are rejected. These findings support Hypothesis II.

3.3 Hypothesis III - The decline of OTC free-riding

In order to gather evidence on the hypothesised shift from OTC to MTS transactions, we used the information contained in the monthly statistical reports of MTS market-makers to the Banca d'Italia. These reports include the OTC trading volume in government bonds of each dealer, with a breakdown for trades carried out with residents and non-residents. We corrected the residents' figures for the effect of double counting by scaling them down by the share of MTS turnover involving trades between two market-makers. We thus obtained an estimate of the OTC volume that is comparable with the MTS exact figures that we possess. We then calculated the ratio of OTC volume over total inter-dealer volume (OTC plus MTS). The resulting figures are given in Table 2. It shows that the OTC share tends to increase from the end of 1996 onward. The highest OTC share figures are observed in July 1997 (37.3%) and in May 1998 (39.6%). The subdued OTC share in the spring of 1997 may partly be explained by the race for volume that took place on MTS and that no longer occurred under anonymity. The evidence of Table 2 is at odds with Hypothesis III.

3.4 Hypothesis IV - The waiting game

If the waiting game hypothesis holds, dealers should try to delay their trades on an intraday basis in the attempt to acquire more information through trade flow, and we would expect a shift of trading volume for the early stages of trading to the later stages. In order to analyse intraday turnover on MTS, we chose the benchmark 10-year BTP issue, which is generally the most heavily traded security. The evidence is given in Figure 2, which shows the intraday relative volume on the benchmark BTP, ie the share of trading volume observed in each half-hour interval of the day over the total daily volume of the bond. The key findings that emerge from Figure 2 are as follows. First, trading volume increases from the first half-hour of trading (09:00-09:30) to the second half-hour. Second, like in most financial markets, there is a decline in trading activity after 12:30 for about 1½ hours. Third, trading activity remains steady after 14:30 (we recall that the closing interval after 17:00 lasts only 10 minutes, ie one third duration of the other intervals). Finally, we note that from period 1 to period 2 volumes decline slightly in the morning intervals and increase correspondingly in the intervals after 14:30. In fact, 3.2% of total daily volume shifts from trading before 14:30 in period 1 to after 14:30 in period 2. The hypothesis of identical distribution of volumes is rejected in nine out of 17 intraday intervals by the Kolmogorov-Smirnov test. The hypothesis of identical means is rejected in six out of 17 intervals by the *t*-test. The last finding seems consistent with Hypothesis IV.

4. Liquidity

Various definitions have been provided in the literature for the concept of market liquidity. Perhaps the most popular one is "a market is liquid if the impact of a trade on price is small". However, the liquidity concept has several other dimensions (see eg O'Hara 1995; Muranga and Shimizu 1997). The richness of our data-set allows us to conduct an empirical study of market liquidity along different definitions. The first and simplest indicator of market liquidity is turnover. For the reason explained in Section two, namely that trading volume should have been biased by the dealers' effort to maintain their status before the June 1997 review, we do not think that it is useful to compare total MTS trading volume before and after the market move to anonymity. Instead, we prefer to focus our attention on the number of bonds that were actively traded on each day. The second indicator of liquidity is the bid-ask spread. The (half-) spread is the reward paid by traders to market-makers for their services, which provides immediacy to those wishing to buy or sell a security. The third indicator of liquidity is the market impact of a trade, which is related to the adverse selection problem faced by market-makers and which varies directly with the perceived arrival of orders from informed traders. We present the evidence on each of the above mentioned indicators respectively in the three following subsections.

4.1 Active bonds

We choose two statistics to describe turnover on the active bonds. We first rank the bonds traded on each day by their volume of transactions. We then consider those bonds below the median and take (1) their number (ie one half of the total number of traded bonds) and (2) their share over total daily trading volume. These statistics are plotted on a daily basis in Figure 3. It shows that the number of the 50% least-traded bonds tends to increase in period 1, and thereafter it declines slightly. On average, this number changes from 63.6 in period 1 to 65.1 in period 2. On the other hand, the volume share of the least traded bond shows an increasing trend, and it doubles on average from 6.9% before anonymity to 14.3% after anonymity. The tests of the hypotheses that the mean and distribution of market share by class do not change are rejected.

We note that in addition to the review of the specialists' status (see Section 2), there were also reviews of the primary dealers' status at the end of 1996 and 1997. One of the requisites was related to each dealer's ability to make a market in the illiquid bonds. We attribute the observed increase in the share of the least traded bonds at year-end to the dealers' attempt to qualify in the annual review. This phenomenon seems to have been particularly significant at the end of 1997.

4.2 Bid-ask spread

Our intraday data-set does not include data on the bid-ask spread. In order to obtain estimates of the fixed-cost of trading associated with the existence of the spread, we use our intraday transactions data to fit the two-equation empirical model of trade and quote revision proposed by Foster and Viswanathan (1993) (see Hasbrouck 1991 for a thorough discussion). This model is as follows:

(1) $$q_t = \alpha + \sum_{i=2}^{N} \delta_i 1_{d_t=i} + \sum_{j=1}^{3} \beta_j dp_{t-j} + \sum_{k=1}^{3} \iota_k q_{t-k} + \tau_t$$

(2) $$dp_t = 2c\left[1_{q_t>0} - 1_{q_{t-1}>0}\right] + \sum_{i=2}^{N} 2c_i \left[1_{q_t>0} - 1_{q_{t-1}>0}\right] 1_{d_t=i} + \lambda\tau_t + \sum_{j=2}^{N} \lambda_j 1_{d_t=j} \tau_t + v_t$$

Where q_t is the signed trade size (eg –5 indicates a public scale of five billion lire at the current bid price) and dp_t is the price change that occurred between the previous trade and the current trade. $1_{d_t=i}$ is an indicator variable equal to one if trade *t* occurs in the *i*-th half-hour interval of the day and 0 otherwise. $1_{q_t=0}$ is an indicator variable equal to 1 if trade *t* is a public buy and 0 otherwise. Equation (1) tries to model the expected value of the incoming order conditional on the past record of orders and prices; the residual τ_t is the unexpected component, or the innovation brought about by the order and potentially related to informed trading.[12] The residual in turn becomes one of the explanatory variables of the price change caused by the order, given by the equation (2). In it, the coefficient *c* is an estimate of the "fixed" component of transaction costs. Assuming that the "true" (and unobservable) value of the bond does not change, *c* measures the difference between the transaction price and the true price, corresponding to one half of the spread, ie to the compensation for the market marking

[12] Equation (1) is run using the logit method.

services provided by the dealer who posted the quote. In practice, since the true bond price ***does*** change over time, if we take 2*c* we do not obtain the actual spread but an unbiased (and noisy) estimate of it.[13] In equation (2) we allow for the possibility that 2c changes during the day, by introducing dummy variables for the half hour intervals $i = 2,\dots,N$, where N is the last interval of the day (from 17:00 to the market close at 17:10). In the same equation, the λ coefficient measures the adverse selection component of trading cost, or market impact of a trade, which enters the total cost of trading when the trade itself is not expected by the market-makers on the basis of the past order flow. Again, we allow for the possibility that λ changes during the day, by introducing (N-1) interval dummies. This estimation approach, which recognises the dynamic nature of trading costs, is similar to those employed in a number of previous studies.[14]

The evidence on the intraday spread estimates for the benchmark 10-year BTP is plotted in Figure 4. The first fact that we note is that 2c is roughly W-shaped during the day. The spread has three peaks: at the open, before 14:30 and at the close. The peak between 14:00 and 14:30 (08:00-08:30 US Eastern Standard Time) is related to the market's uncertainty concerning the opening prices of the United States financial markets. The peak may also be related on some days to the upcoming release of United States' macroeconomic indicators. This finding is analogous to previous evidence for MTS (Scalia 1998a) and to the behaviour of the United States' T-bond market (Fleming and Remolona 1997). The second fact that we note is that the spread in period 2 is uniformly lower than in period 1. In particular, the spread in the initial and final intervals of the day declines from 2 to 1.4 basis points of price.

It may be argued that the estimated reduction of the spread, which is positively related to the asset's expected volatility may have been caused by the general improvement in the Italian Treasury bond market, brought about by the increase in Italy's prospects for early participation in the EMU. This poses the problem of distinguishing the effects that MTS anonymity and the macroeconomic change have had on our market performance variables. As a control variable for macroeconomic improvement, we chose the 10-year BTP-Bund-yield differential.[15] Figure 5, Panel A shows the series of the estimated bid-ask spread and the BTP-Bund yield differential on a daily basis. This yield differential fell from around 3% in September 1996 to 1% in July 1997 and fell again to 0.25% in May 1998. The bid-ask spread series shows a declining trend in period 2. In that period the differential and the spread are clearly associated.[16]

What are the relative weights of the micro- and macroeconomic effects on the spread? In order to provide an answer, we run a regression of the spread estimate over a constant, the differential a dummy equal to one in the second period, and the product of the previous two variables. The weights are obtained as the product of the estimated coefficients by the average value of each variable, as a percentage of total. These weights are reported in Figure 5, Panel B. The weight of the microeconomic effect, related to the dummy variable, is equal to 56%. The macroeconomic variable, ie the differential, accounts for 10%, and the third variable (the differential times the dummy) accounts for 34%. Adopting a cautious stance, and attributing the last estimate entirely to the macroeconomic effect, we observe that the microeconomic effect accounts for over one half of the total improvement in the bid-ask spread from period 1 to period 2.

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13 In order to control for residual heteroskedasticity caused the different length of time between subsequent trades, we weight each observation in equation (2) by the inverse square root of the time elapsed since the previous trade. We thus run equation (2) with the weighted least squares method.

14 Equation (2) is instantaneous, ie there are no lagged effects of prices or quantities. According to Hasbrouck (1991), the inclusion of lagged terms in the price equation would be justified under the following circumstances: (a) inventory effects are in place, such that dealers seek to smooth the holdings of bonds over time; (b) there is "price-discreteness", due to a large tick-size; (c) prices adjust slowly to new information. In our setting, we think that the case for hypotheses (a)-(c) is weak, and the inclusion of lagged terms would only affect the efficiency of the estimates. Therefore, we see no compelling reason for departing form the Foster-Viswanathan instantaneous-modelling approach. The average adjusted R-square of our daily equations is equal to 0.36.

15 Another plausible proxy might be the market perceived probability of Italy's early participation in the EMU. This estimated probability measure and the BTP-Bund spread are strongly correlated.

16 This is confirmed by a simple regression of the spread over a constant and the differential (not reported for simplicity). We also perform a Chow stability test that the regression coefficients are identical between period 1 and period 2. The results show that the differential is directly related to the spread; however, this effect is limited to period 2, and the stability test is rejected.

4.3 Market impact

The intraday evidence on the estimates of the market impact λ is plotted in Figure 6. The first finding is that in period 1 there are minor variations of λ during the day, whereas in period 2 there is a tendency for market impact to increase in the early afternoon intervals. The second finding is that λ is uniformly lower in period 2 than in period 1.

Has λ been influenced by the general macroeconomic improvement of the market? Figure 7, Panel A shows the market impact series and the yield differential series. The evidence, again, is that the differential is positively related to the spread in period 2, but unrelated to it in the earlier period.[17]

The results on the weights of the micro- and macroeconomic effects are given in Panel B, obtained with the same methodology of the previous subsection. The microeconomic effect turns out to be extremely large, equal to around 69% of the total price impact. The macroeconomic effect accounts for the remaining 31%.

5. Volatility

We estimate volatility on an intraday basis as the squared log-difference of the bench mark 10-year BTP prices, taken at half-hourly intervals. The resulting evidence is presented in Figure 8. Intraday volatility displays a U-Shape. Although its estimate declines in the last interval of the day, we recall that the different length of the interval itself does not make the corresponding value comparable to estimates for earlier intervals.[18] Volatility is largest in the initial interval of period 1, when it is equal to 0.03%. Throughout the rest of the day it is much lower, generally below 0.015%, and it rises after 14:30. The second fact that we note is that volatility in period 2 is uniformly lower than in period 1. In particular, volatility in the first half hour of trading declines from 0.030 to 0.011. Moreover, after 14:30 the increase in volatility is less pronounced.

Figure 9, Panel A provides evidence on the relationship between the BTP-Bund yield differential and volatility on a daily basis. The picture is slightly different from the case of the spread and market impact. A direct relationship between yield differential and volatility is found; this is significant in period 2 only; however, the Chow stability test between periods can not be rejected. The evidence of panel B is that the microeconomic effect has a weight of 37% on volatility, ie much smaller than in the case of the cost measures, whereas the macroeconomic effect accounts for the remaining 63%.

6. Efficiency

The notion of financial market efficiency implies that prices fully reflect all available information. As is well known, Fama (1970) distinguishes three types of efficiency: ***weak form efficiency***, which requires that no investor can earn excess returns based on historical price information; ***semi-strong-form efficiency***, which implies that no investor can earn excess returns by applying trading rules based on any publicly available information; and ***strong-form efficiency***, which implies that no investor can earn excess returns using any type of information, whether public or private. While strong-form efficiency is unachievable if one accepts the view that information asymmetries are a relevant factor in explaining dealers' behaviour, weak-form and semi-strong-form efficiency are in principle attainable by a financial market. In particular, the hypothesis of weak-form efficiency has been tested by empirical studies on leads and lags between cash and futures markets for the same security, in which prices are strictly correlated due to a no-arbitrage argument. The evidence in the case of bond markets is available for Japan and Italy. In Japan the JGS inter-dealer cash market is driven by the futures market, with cash prices lagging behind the price of the 10-year JGS contract traded on the Tokyo Stock Exchange by two minutes on average (Miyanoya, Inoue and Higo 1997). In the case of the Italian BTPs there is evidence of reciprocal causality between the futures contract traded on LIFFE and the benchmark 10-year BTP traded on MTS in the years 1992-1993; furthermore,

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[17] The Chow stability test between period 1 and 2 is rejected.

[18] Under the hypothesis that bond prices follow a Brownian motion, our (squared) volatility proxy in the last interval should be multiplied by 30'/10'=3 in order to express it in half-hourly terms.

the futures lead can not be exploited to earn excess returns on MTS, consistent with weak-form efficiency of MTS with respect to LIFFE (Scalia 1998b; see also Angeloni et al 1996).

Has MTS changed its record of efficiency with respect to LIFFE following its switch to anonymity? This question is relevant because traders in the two marketplaces are not fully integrated, particularly concerning their access to information on monetary policy implementation, the Treasury's issuing decisions and the order-flow. The empirical analysis that follows seeks to update previous evidence, while improving the type of data and the power of the causality test.

Our data sample includes all MTS transactions on the benchmark 10-year BTP and all BTP futures transactions at LIFFE in the period from September 1996 to May 1998.[19] We also employ an intraday data-set, obtained from the Reuters service, that contains market prices and quotes at five-minute intervals on the following financial instruments: the three-month eurolira futures contract at LIFFE (last trade price), the Deutsche Mark/US Dollar exchange rate (last bid), and the 10-year Bund futures contract traded at LIFFE (last trade price). The general motive for the inclusion of these variables in a VAR analysis of causality is to take into account the behaviour of the world financial markets that potentially may explain the behaviour of BTP prices, ie we should like to include in a parsimonious way all the relevant information set. We observe that, compared with previous studies, we take a step from the notion of weak-form efficiency to that of semi-strong-form efficiency, which involves the predictability of prices based on all publicly available information. The specific reasons for this set of variables are as follows. The short-term rate futures captures the attitude of domestic monetary policy. The DM/USD exchange rate is the reference exchange rate for Europe, reflecting the relative degree of monetary tightness between the United States and Germany. The Bund futures prices incorporate the attitude of investors towards the European fixed-income market.

After taking the log-differences of our intraday time series at five-minute intervals (simple differences for the eurolira rate), for each day in our sample we ran a VAR system of equations in order to check if any pattern of causality emerges among the prices of our financial instruments, and in particular between BTP cash and futures prices.[20] The evidence on absolute contemporaneous correlation among variables is given in Table 3, Panel A. The evidence on the VAR estimates is contained in Panel B, which gives summary statistics (frequency and mean) on the coefficients that turned out to be significantly different from zero across all days. The maximum lag length with significant statistical power in both samples is 10 minutes (two lags). However, since the second lag of variables turns out to be significant in a negligible number of cases, for ease of presentation the table reports only the evidence for the first lag of variables.

The key facts that emerge from our estimates are as follows. First, as with previous evidence from many financial markets worldwide, all our series display substantial mean-reversion at five-minute intervals. In particular, the average mean-reversion coefficient for the BTP cash price is –0.41 in period 1 and –0.35 in period 2; the averages for the BTP futures are –0.33 and –0.41, the averages for the eurolira rate are –0.33 and –0.31. Second, contemporaneous correlation of price changes between cash and futures BTP is extremely high (0.72 and 0.64 on average in periods 1 and 2,[21] as one would expect based on the no-arbitrage principle. Third, causality between cash and futures BTP runs in both directions. In particular, the five-minute average lead of LIFFE declines from 0.39 to 0.34, while the average lead of MTS is almost unchanged, from 0.33 to 0.32. Furthermore, while the number of days in which LIFFE displays a significant lead on MTS declines from 30 in period 1 to 18 in period 2, the corresponding frequency for the MTS leads increases from 17 days in period 1 to 25 days in period 2. Finally, there is evidence of positive two-way causality between price changes of the Bund futures, on one side, and of the BTP cash and futures, on the other side. Interestingly, we observe that contemporaneous correlation increases over time (from 0.47 to 0.51 for the benchmark BTP, from 0.49 to 0.55 for the BTP futures) and that causality from the Bund to the BTP becomes positive in a number of cases in period 2. These phenomena are consistent with the hypothesis that, thanks to the

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19 The futures data-set was kindly made available by LIFFE.

20 The VAR model is estimated in the interval 09:00-17:10 (opening hours of MTS) on a daily basis. The number of lags is selected by minimising the Akaike information criterion.

21 The fact that we use the benchmark BTP, which is not necessarily the cheapest-to-deliver bond for the futures contract, diminishes the power of the no-arbitrage principle in our case, thus reducing the correlation between cash and futures.

improvement in the prospects of first-round participation of the lira in the EMU, in period 2 the Italian and German bond markets have become more integrated.

Compared with the evidence on causality for the years 1992-1993, when the LIFFE lead over MTS was of 15 to 30 minutes with an intensity of 0.25-0.30, in recent years the lead has become much shorter, and the frequency of cases in which it is longer than five minutes is just four days out of 309. The MTS lead has increased compared to 1992-1993.

7. Discussion and regulatory policy implications

We summarise the main empirical findings of the previous sections.

A. Small traders' participation on MTS decreases from period 1 to period 2.

B. Large traders' participation increases.

C. Large trades on MTS become more frequent in period 2.

D. The share of OTC transactions over total inter-dealer trading increases slightly from period 1 to period 2.

E. The shape of intraday trading volume on the benchmark bond is slightly displaced towards the late trading intervals of the day, from period 1 to period 2.

F. The share of trading volume of the 50% least traded bonds on MTS doubles from period 1 to period 2.

G. The intraday bid-ask spread is W-shaped, and the spread in period 2 is uniformly lower than in period 1.

H. The market impact λ is uniformly lower in period 2 than in period 1.

I. Volatility is U-shaped and uniformly lower in period 2.[22]

J. The increase in Italy's prospects for early participation in the EMU is correlated to the improvement in spread, market impact and volatility in period 2, but virtually uncorrelated to them in period 1. The macroeconomic effect explains between 31% and 63% of the improvement in market performance.

K. Causality between BTP cash prices on MTS and futures prices at LIFFE runs in both directions.

L. From period 1 to period 2 the intensity of causality from either market becomes similar, the frequency of the LIFFE lead declines, the frequency of the MTS lead increases.

The first theoretical hypothesis that we made was that the smaller MTS traders, who are most likely to be liquidity motivated and uninformed, have been made worse-off by the market move to anonymity. Finding A is clearly consistent with the "liquidity trader's curse". Some small traders, although formally MTS members, may have withdrawn from active market participation because under anonymity they have less control on the "real game" played by the large traders, thus being unable to mirror their moves. It seems likely that either or both of the following phenomena may have occurred in period 2: (1) small players deal more frequently on an OTC basis through large dealers, and are prepared to pay a commission for the superior information possessed by the latter; (2) small players participate more actively in the uniform-price auctions of Treasury securities. The counterpart to this are findings B and C, suggesting that the "large trader's blessing" has indeed occurred. Under anonymity large traders are better able to carry out big inventory adjustments, which in period 1 were presumable executed on the OTC market.

The estimated increase in the share of OTC trading volume (finding D), although a small amount, is somewhat puzzling. It is the opposite of Hypothesis III. The "decline of OTC free-riding" hypothesis is actually related to two considerations. First, anonymity makes MTS more similar to the OTC inter-dealer-broker market. This increases ceteris paribus the incentives to trade on MTS. Second, under

[22] Concerning the findings G, H and I, for control purposes, we also ran the empirical tests on market liquidity and volatility using a different set of securities, namely the just-off-the-run five-year BTPs. In a ranking of daily trading volume these bonds generally lie between the fifth and the 15th most traded issues. The evidence (available from the authors) confirms the findings for the 10-year benchmark bonds. However, finding E is no longer observed on five-year BTPs.

anonymity it becomes more difficult for the OTC market to free-ride on price and order-flow information provided by MTS. This phenomenon may cause an increase in trading cost on the OTC market, and may induce some dealers to trade directly on MTS. Are there any other reasons for the observed increase in OTC turnover? It is possible that the OTC inter-dealer brokers have reacted to the 1997 MTS shift, which increased competition between the two markets, by reducing spreads. We note an interesting example concerning competition between the OTC market and MTS. Cantor Fitzgerald, one of the major brokers trading Italian bonds from London, had often used MTS through an intermediary in the past; at the end of 1997, it applied for membership with MTS and in May 1998 it started trading large volumes directly on the Italian market. Since MTS prices reflect a more homogeneous market set-up, it is now conceivable that they can be straightforwardly applied to OTC transactions. If this is so, then OTC free-riding on MTS might have even increased after the MTS switch.

The fourth theoretical prediction that we investigated is the waiting game hypothesis of Foster and Viswanathan. Finding E is consistent with this hypothesis: under anonymity the order flow information of MTS has become less useful to dealers, and they tend to wait longer in order to extract more information. There is a further reason for the slight displacement of the intraday profile of trading volume in period 2. During our sample period the Italian market has become increasingly integrated with the other major financial markets. Among these, the US market is an important source of information and it has been a growing source of investment into the Italian market. Hence, the information and orders that start arriving on MTS from 14:30 onward, ie from the opening of the US market and the release time of most US macroeconomic indicators, have increased over time, and this clearly contributes to the observed shift in intraday trading volume on MTS.

Findings F to L represent in our opinion impressive evidence on the improvement in the performance of MTS in recent years. In interpreting these results, we face an attribution problem. As we argued earlier on, two distinct factors may have played a role, namely the switch to anonymity, a one-time event that took place in the middle of 1997, and the steady progress of public finance of 1996-1997. We tried to distinguish between these two factors, and obtained results that show that the microeconomic effect amounts for 31 to 63% of the variation in the market performance variables. In the case of the two cost measures, the microeconomic effect is more important. In the case of volatility, the macroeconomic effect takes first place. This is not surprising, since market volatility may be expected to be more sensitive to macroeconomic conditions than trading costs.

An additional factor that may have played a role is the listing to repo contracts on MTS starting in December 1997. Repo contracts on Treasury bonds have been traded among dealers on the OTC market for long before that date. However, cash traders greatly benefited from the inception of repo trading directly on MTS, through a reduction in the cost of setting up short positions. This may help explain why the speed of price discovery on MTS has increased with respect to the futures market (findings K and L).

From a regulatory point of view, the evidence presented in this paper has several implications. The first implication is domestic. The move to anonymity has furthered the reform process of the Treasury bond market that the Italian regulatory authorities initiated in 1994 (see the Appendix table). This reform was aimed at restoring the competitive role of MTS with respect to the OTC market, by opening up the former to foreign investors, lowering transaction costs and promoting competition among dealers. Since 1994 MTS has greatly increased efficiency and turnover relative to the OTC market. As we have shown, the 1997 shift helped to enhance this competition, affording higher levels of welfare for those who invest in Italian Treasury bonds. The improvement of the secondary market should also have benefited the issuer, through a reduction in the cost of debt servicing. We conclude that the 1997 innovation on MTS has proved successful.

The second regulatory implication follows from the first one. Looking at the Italian Treasury bond market from a more general perspective, we note that the market has made a remarkable progress in just one decade, from an opaque, lowly liquid market with negligible foreign participation to a highly transparent and liquid market with a large participation of international investors. This progress has been similar in nature to developments in other industrialised countries, but in the Italian case it has been more intense. To this extent, MTS has played a key role. The ideas that have underlain the MTS inception and development have proved successful in the medium term. These ideas are: (1) full automation of the trading mechanism; (2) transparency; (3) large participation; (4) inside and outside competition. We believe that the experience of MTS may be useful for those emerging countries wishing to establish a liquid and efficient financial market in a relatively short time horizon.

In 1998 the market was fully privatised. A major development took place in September 1998, namely the listing of a eurolira 10-year bond issued by the European Investment Bank and of a large group of German government bonds.[23] The listing of sovereign bonds from other countries is also planned. In the perspective of EMU, it has been argued that the likely integration of the European bond markets might imply either a strong cooperation among sovereign issuers, or a "race to benchmark status" (McCauley and White 1997). In both cases, the role of each country's government bonds within the European market will be positively affected by the liquidity conditions of the domestic market and by the availability of the securities in the portfolios of international investors, even more than by the creditworthiness of the issuer. In this view, the improvement in the liquidity of MTS, along with the decision by the Italian Treasury to convert all outstanding debt in euros on 1 January 1999, places the Italian issues in a strong position among the partner countries' issues.

8. Conclusion

We analysed a change in the organisation of the electronic inter-dealer market for Italian Treasury bonds known as MTS, namely the shift to the anonymity of quotes in July 1997. The implications of this event were investigated in the light of market microstructure theory and from a public welfare perspective. We employed an extensive data-set which includes all transactions carried out on MTS with the identity of the traders, in the period from September 1996 to May 1998. In addition, we used intraday prices for the BTP futures contract traded at LIFFE and for a set of financial instruments that may be viewed as explanatory variables for the dynamics of BTP prices. Our evidence supports the hypothesis that the decrease in transparency makes liquidity traders worse-off, whereas large/informed traders find it less costly to execute block trades. The evidence is also consistent with the "waiting game" hypothesis of Foster and Viswanathan (1996) on intraday trading: under anonymity, traders tend to delay their trades in an attempt to acquire information through the order flow. From a public welfare perspective, our results indicate that the move to anonymity has been accompanied by an increase in market liquidity and by a reduction in volatility, a phenomenon that is also partly explained by the growth in Italy's prospects for early participation in the EMU. The speed of information aggregation on MTS increases, as shown by an improvement of the MTS lead over the futures market. From a regulatory policy perspective our evidence suggests that, despite the welfare loss suffered by small traders, the move to anonymity has afforded an overall improvement in market performance. In this respect, the experience of MTS may be useful for the development of market mechanisms in emerging countries. Finally, in a European perspective, the current organisation and performance of MTS place the market in a competitive position compared to other cash markets for government bonds, and may contribute to a closer integration of these markets under the EMU.

[23] Contracts on these bonds are cleared and settled through international depository entities (Euroclear and Cedel).

Table 1

Dealers' participation on MTS

	Period 1	Period 2	p-value>t[1]
Number of traders with a market share of:[2]			
Less than 0.1%	15	3	
0.1 - 0.25%	110	84	
0.25 - 0.5%	59	65	
0.5 - 1%	34	41	
1 - 2.5%	27	30	
2.5% or more	15	18	
Total	**260**	**241**	
Herfidahl concentration index			
among all traders:			
daily average (%)	3.2		0.00
(standard deviation)	(0.7)		
among market-makers:			
daily average (%)	5.0	5.8	0.00
(standard deviation)	(0.9)	(1.0)	

[1] A p-value at or below 0.05 implies rejection of the null hypothesis of identical means by the *t*-test. [2] The traders' shares are daily averages (213 days for period 1, 221 days for period 2).

Table 2

Monthly trading volume on OTC market and MTS

(trillion lire and percentage values)

	OTC[1]	MTS	OTC share on total %
1996			
September	310	812	27.6
October	393	915	30.1
November	326	892	26.7
December	284	717	28.3
1997			
January	343	1136	23.3
February	360	834	30.2
March	356	735	32.6
April	396	898	30.6
May	463	1048	30.6
June	528	946	35.8
July	508	854	37.3
August	322	562	36.4
September	455	898	33.6
October	457	928	33.0
November	363	730	33.2
December	355	611	36.8
1998			
January	297	658	31.1
February	295	621	32.2
March	379	726	34.3
April	296	566	34.4
May	315	481	39.6

[1] Estimated values.

Table 3

Intraday evidence on price causality

(averages of daily estimates)

Panel A: contemporaneous correlations

	10-year benchmark BTP	10-year BTP future	3-month eurolira	10-year Bund future	D-Mark/ US dollar
	Period 1				
10-year benchmark BTP	1.00	0.72	– 0.37	0.47	0.08
10-year BTP future		1.00	– 0.42	0.49	0.08
3-month eurolira			1.00	– 0.29	– 0.05
10-year Bund future				1.00	0.05
D-Mark/US Dollar					1.00
	Period 2				
10-year benchmark BTP	1.00	0.64	0.08	0.51	0.01
10-year BTP future		1.00	0.08	0.55	0.01
3-month eurolira			1.00	0.09	– 0.01
10-year Bund future				1.00	0.01
D-Mark/US Dollar					1.00

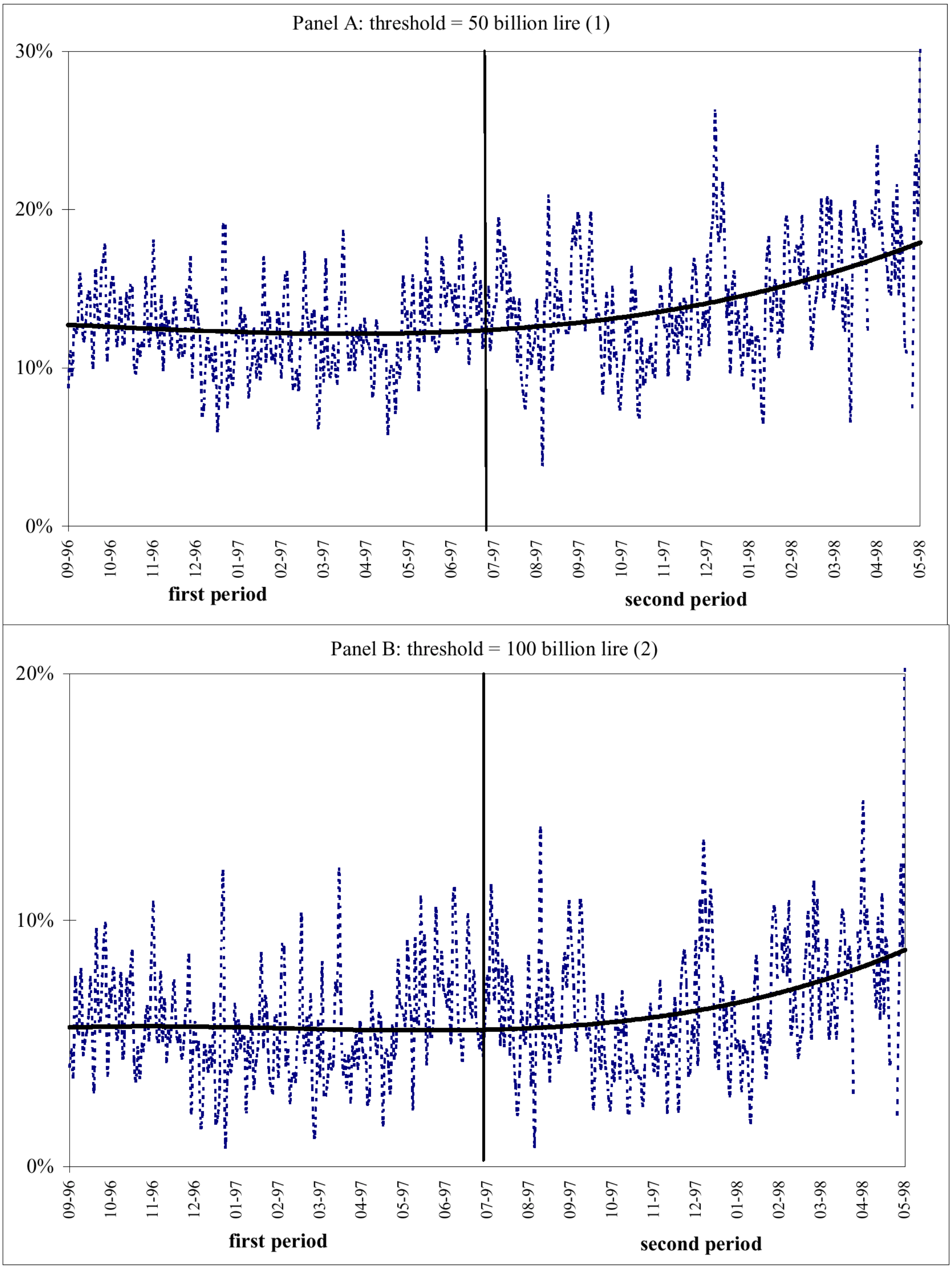
Panel B: lead-lag estimates[1]

	10-year benchmark BTP		10-year BTP future		3-month eurolira		10-year Bund future		D-Mark/ US Dollar	
	Average (2)	No days	Average (2)	No days	Average (2)	No days	Average (2)	No days	Average (2)	No days
	Period 1									
10-year benchmark BTP	– 0.41	38	0.39	30	– 0.81	19	– 0.10	20	0.18	12
10-year BTP future	0.33	17	– 0.33	28	– 1.37	11	– 0.02	20	0.19	8
3-month eurolira	– 0.03	18	– 0.08	34	– 0.33	85	– 0.01	15	– 0.03	10
10-year Bund future	0.17	11	0.27	22	– 0.41	15	– 0.32	36	0.09	10
D-Mark/US Dollar	0.08	13	0.00	11	– 0.50	8	– 0.17	13	– 0.27	37
	Period 2									
10-year benchmark BTP	– 0.35	35	0.34	18	0.12	9	0.29	17	0.11	9
10-year BTP future	0.32	25	– 0.41	39	0.10	8	0.32	21	0.13	7
3-month eurolira	0.04	12	0.08	14	– 0.31	56	0.01	8	0.00	4
10-year Bund future	0.15	18	0.20	20	0.34	9	– 0.34	27	– 0.01	9
D-Mark/US Dollar	– 0.07	9	0.31	6	0.11	5	0.30	21	– 0.25	18

[1] Causality at five-minute level runs from the variables along the top row to the variables along the first column on the left. Due to gaps in the intraday series, 159 days and 150 days were employed for the estimates respectively in period 1 and period 2. [2] Average estimated causality over the days where estimated causality is non-zero with 95% confidence. [3] Number of days in which the estimated causality is non-zero with 95% confidence.

Figure 1

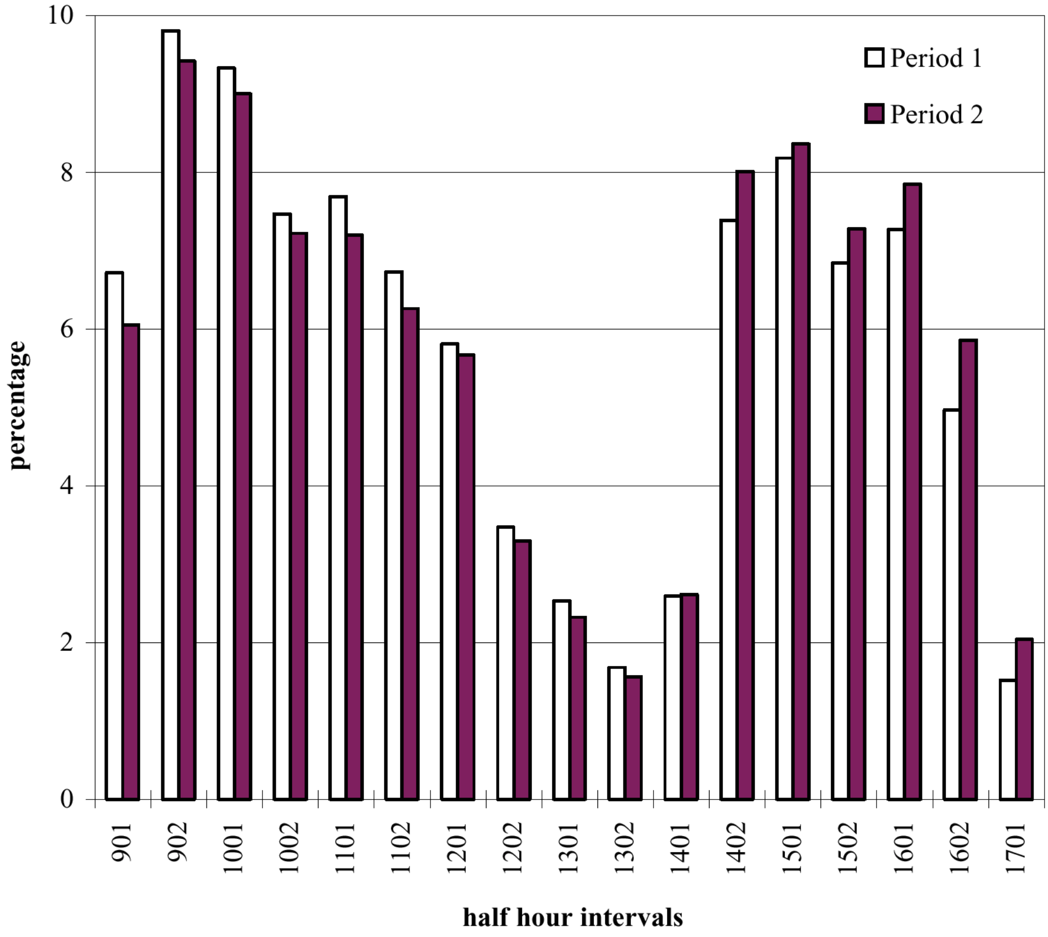
BLOCK TRADES ON MTS AS A SHARE OF TOTAL TRADING VOLUME



(1) A block trade is assumed whenever the net daily change in a trader's holding of an issue is worth at least 50 billion lire. - (2) A block trade is assumed whenever the net daily change in a trader's holding of an issue is worth at least 100 billion lire.

Figure 2

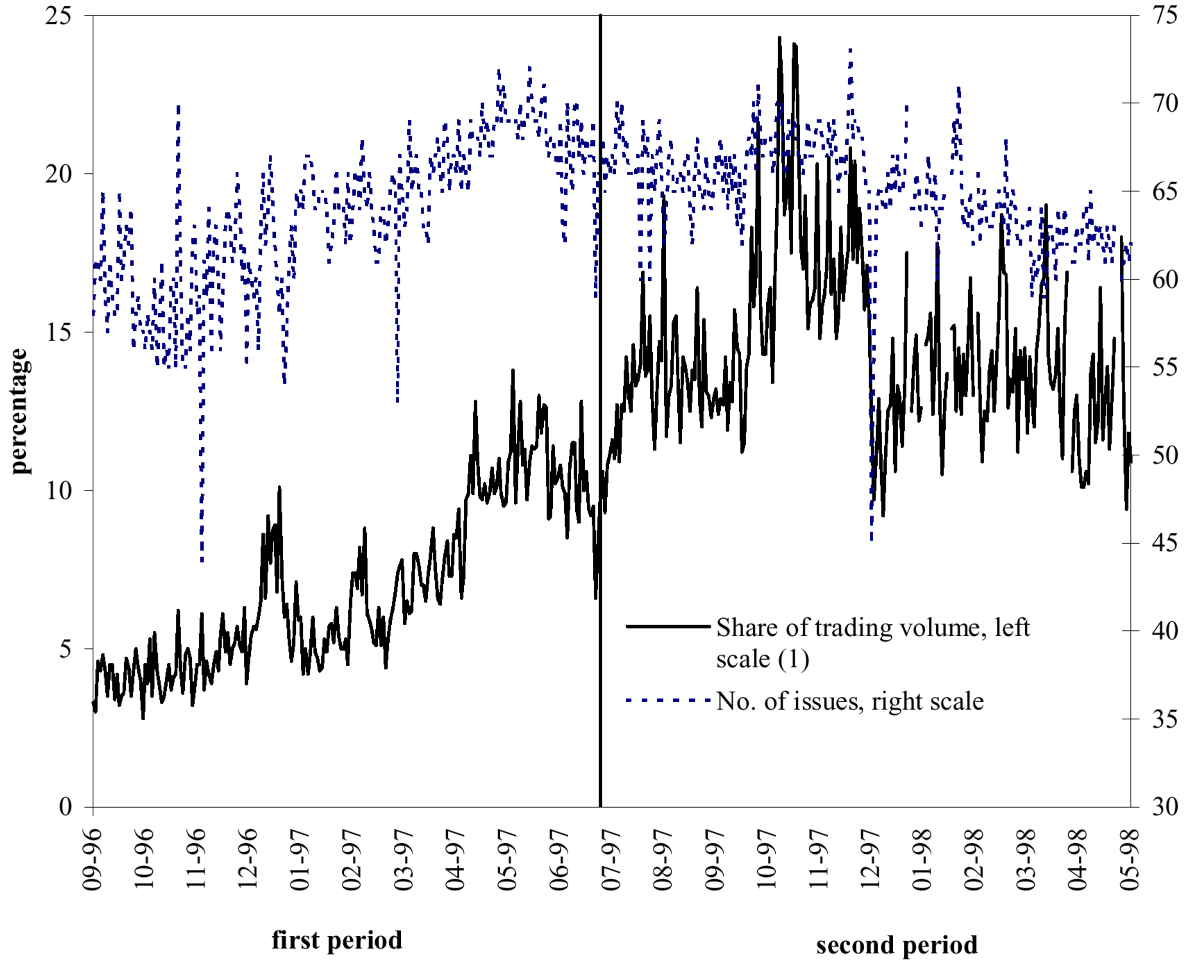
MTS: INTRADAY TRADING VOLUME ON 10-YEAR BENCHMARK BTP (1)



(1) Data are averages of shares of daily volume. The Kolmogorov-Smirnov test rejects the null hypothesis of identical distributions in period 1 and 2 for 9 out of 17 intervals, at 5% level. The *t*-test rejects the null hypothesis of identical means for 6 out of 17 intervals, at 5% level.

Figure 3

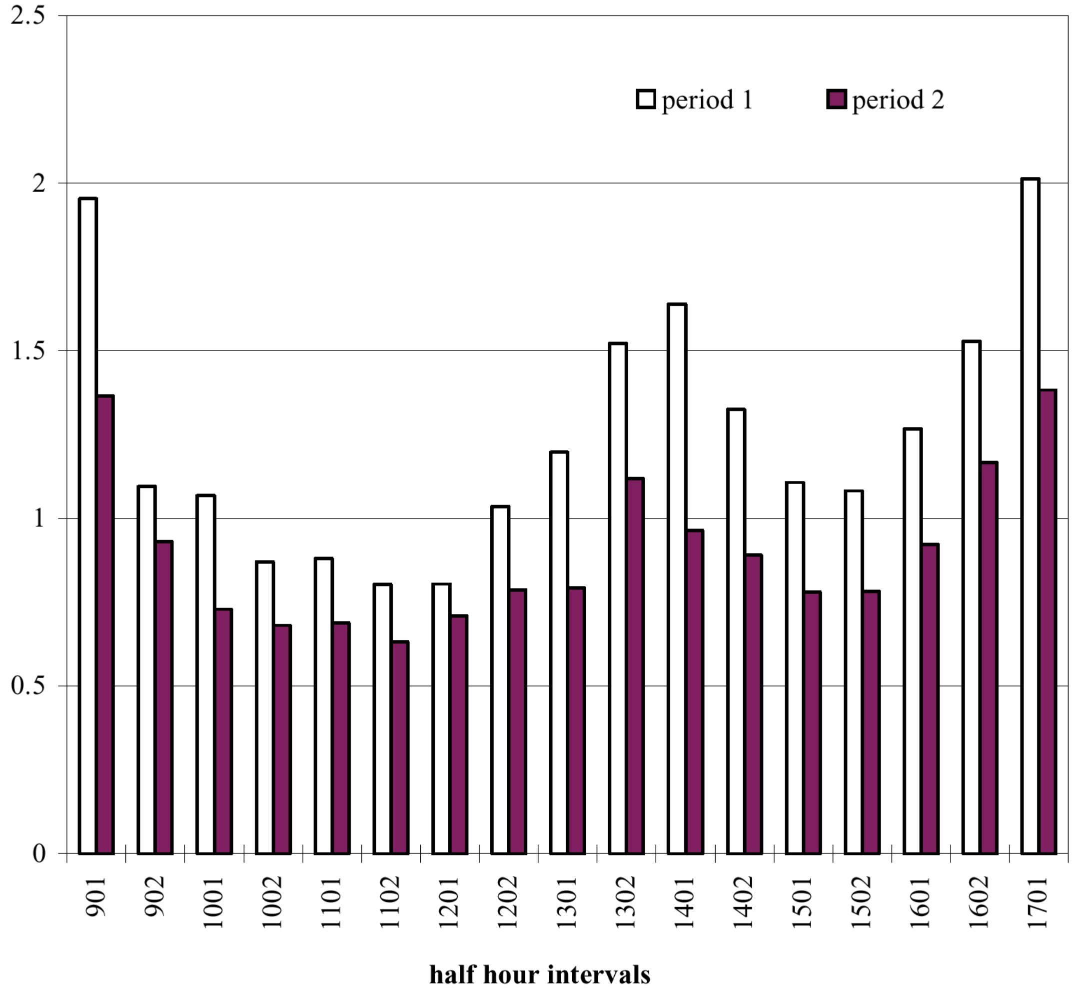
LEAST LIQUID BONDS ON MTS: NUMBER OF ISSUES AND SHARE OF TRADING VOLUME



(1) The Kolmogorov-Smirnov test rejects the null hypothesis of identical distributions in period 1 and 2 at 5% level. The *t*-test rejects the null hypothesis of identical means at 5% level.

Figure 4

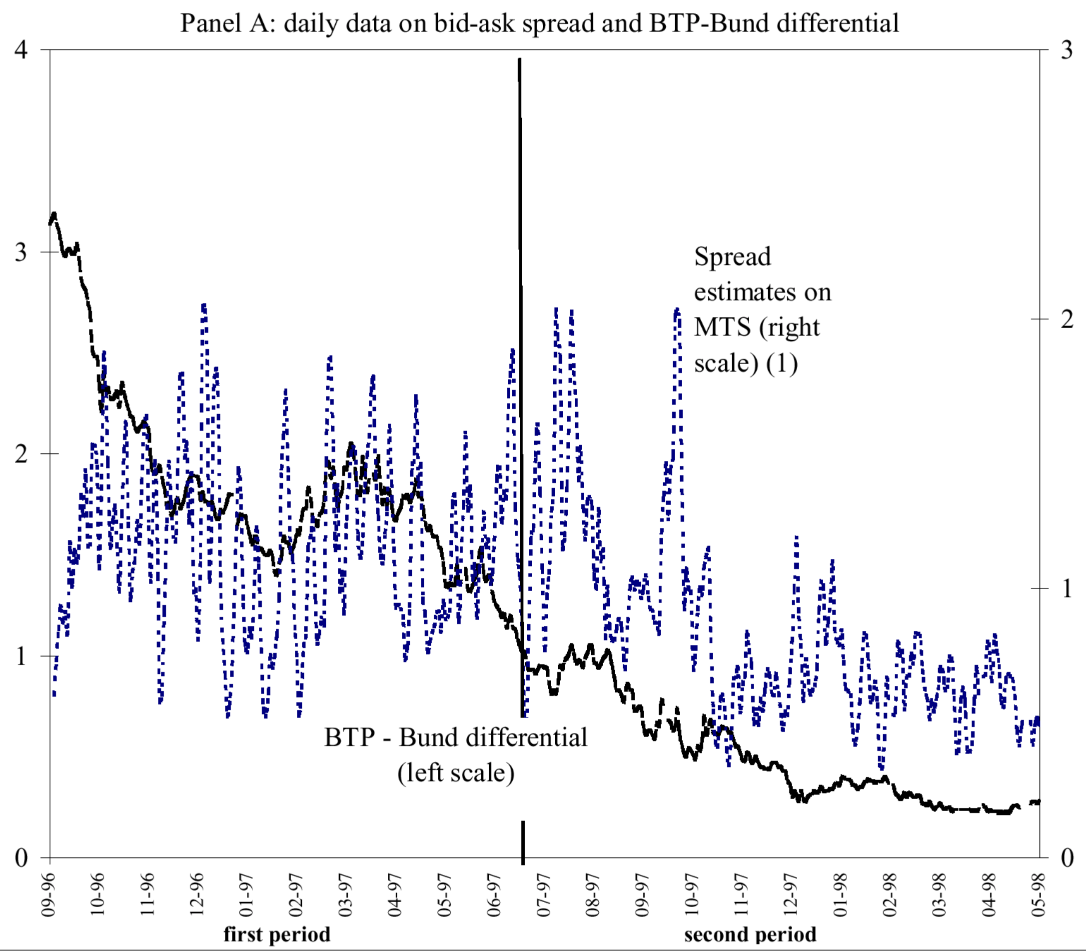
INTRADAY SPREAD ESTIMATES ON 10-YEAR BENCHMARK BTP (1)

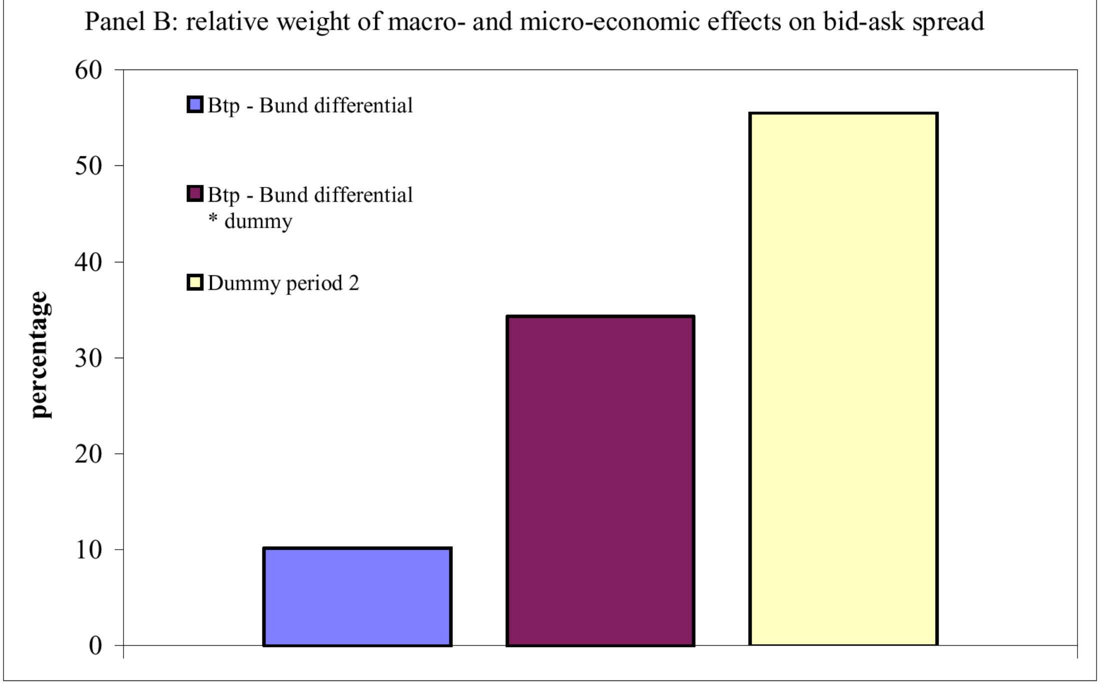


(1) Data are in basis points. See equations *(1)* and *(2)* in the text. The Kolmogorov-Smirnov test rejects the null hypothesis of identical distributions in period 1 and 2 for 16 out of 17 intervals, at 5% level. The *t* -test rejects the null hypothesis of identical means for 16 out of 17 intervals, at 5% level.

Figure 5

MACRO- AND MICRO-ECONOMIC EFFECTS ON BID-ASK SPREAD

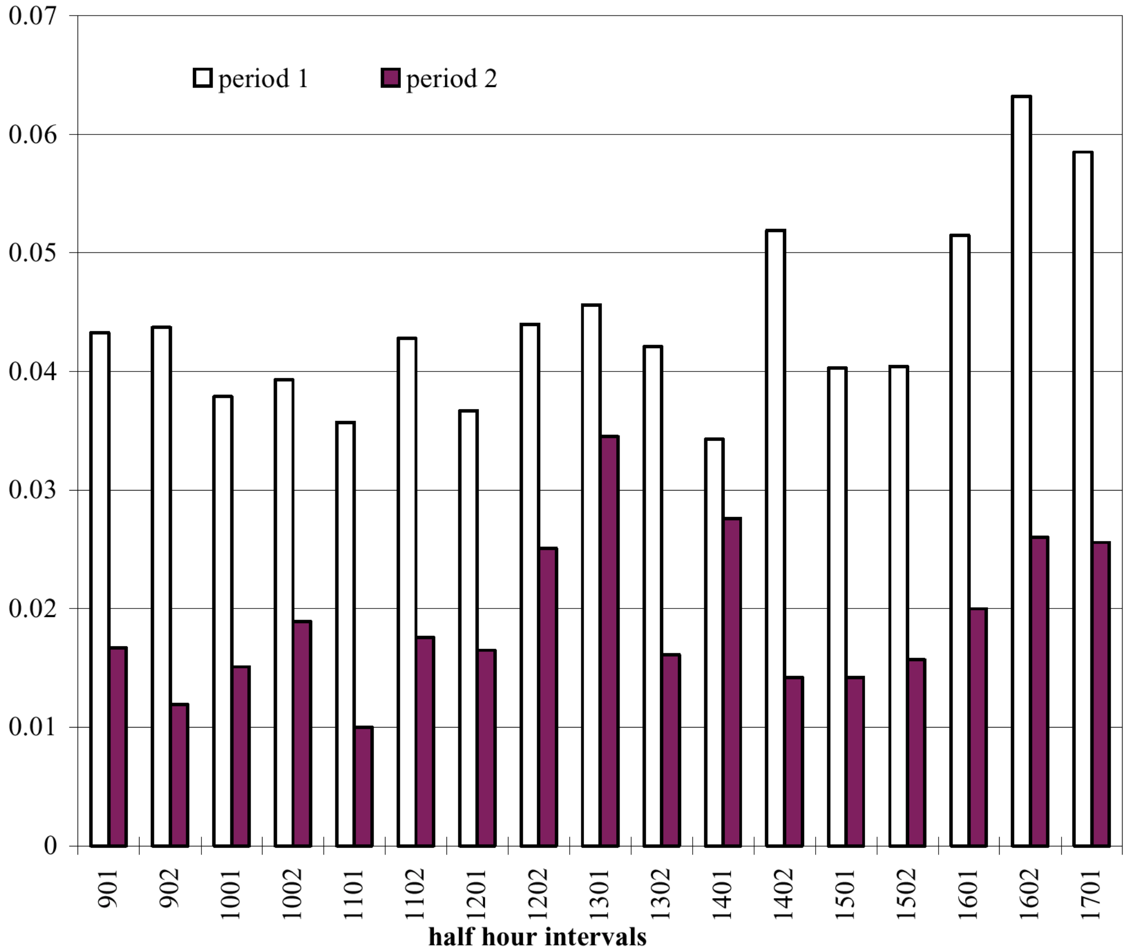




(1) Spread estimates on MTS, in basis points, are moving averages over three days.

Figure 6

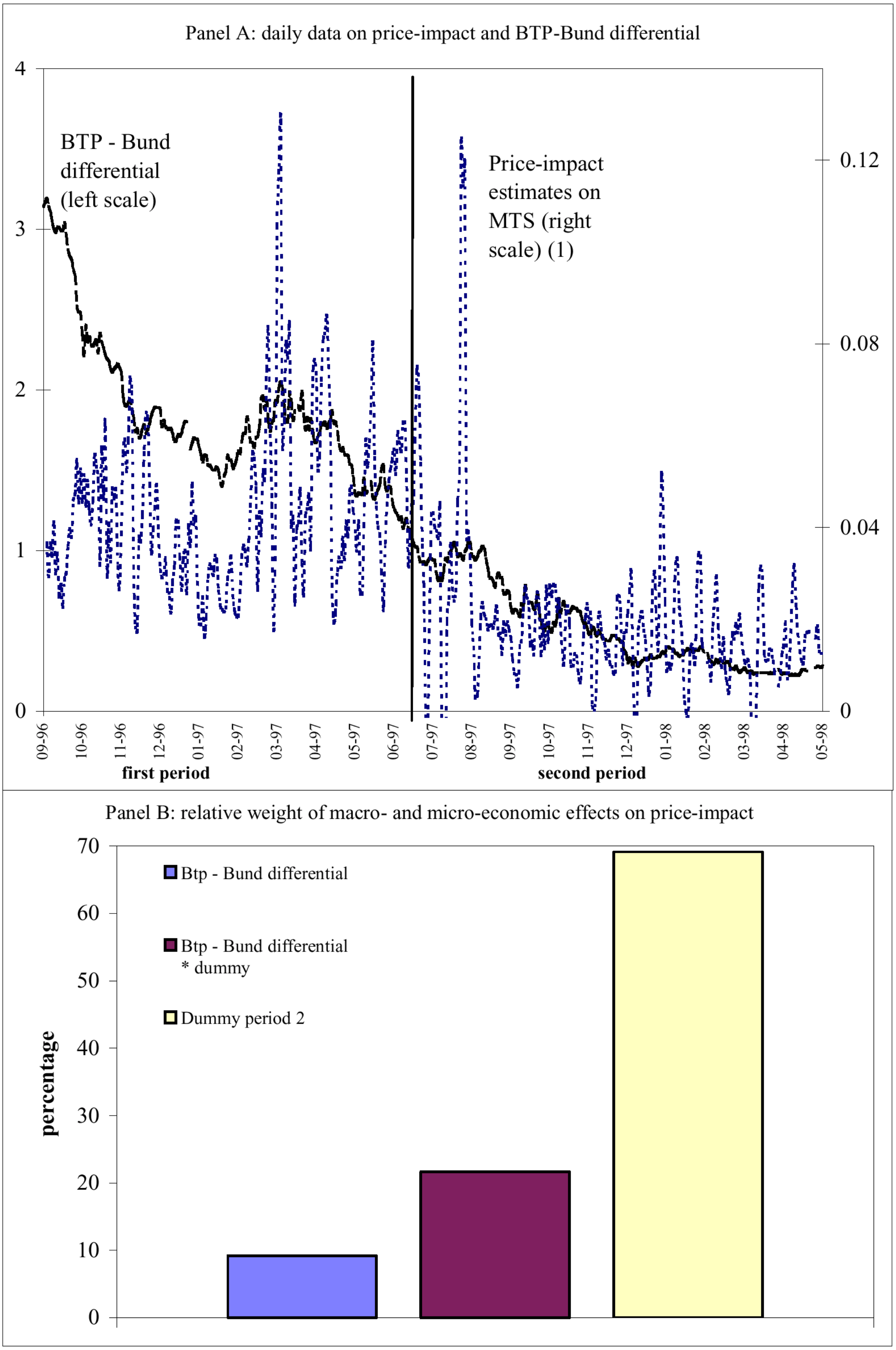
INTRADAY PRICE IMPACT ESTIMATES
ON 10-YEAR BENCHMARK BTP (1)



(1) Basis points per 1 billion lire order size. See equations *(1)* and *(2)* in the text. The Kolmogorov-Smirnov test rejects the null hypothesis of identical distributions in period 1 and 2 for 17 out of 17 intervals, at 5% level. The *t* -test rejects the null hypothesis of identical means for 15 out of 17 intervals, at 5% level.

Figure 7

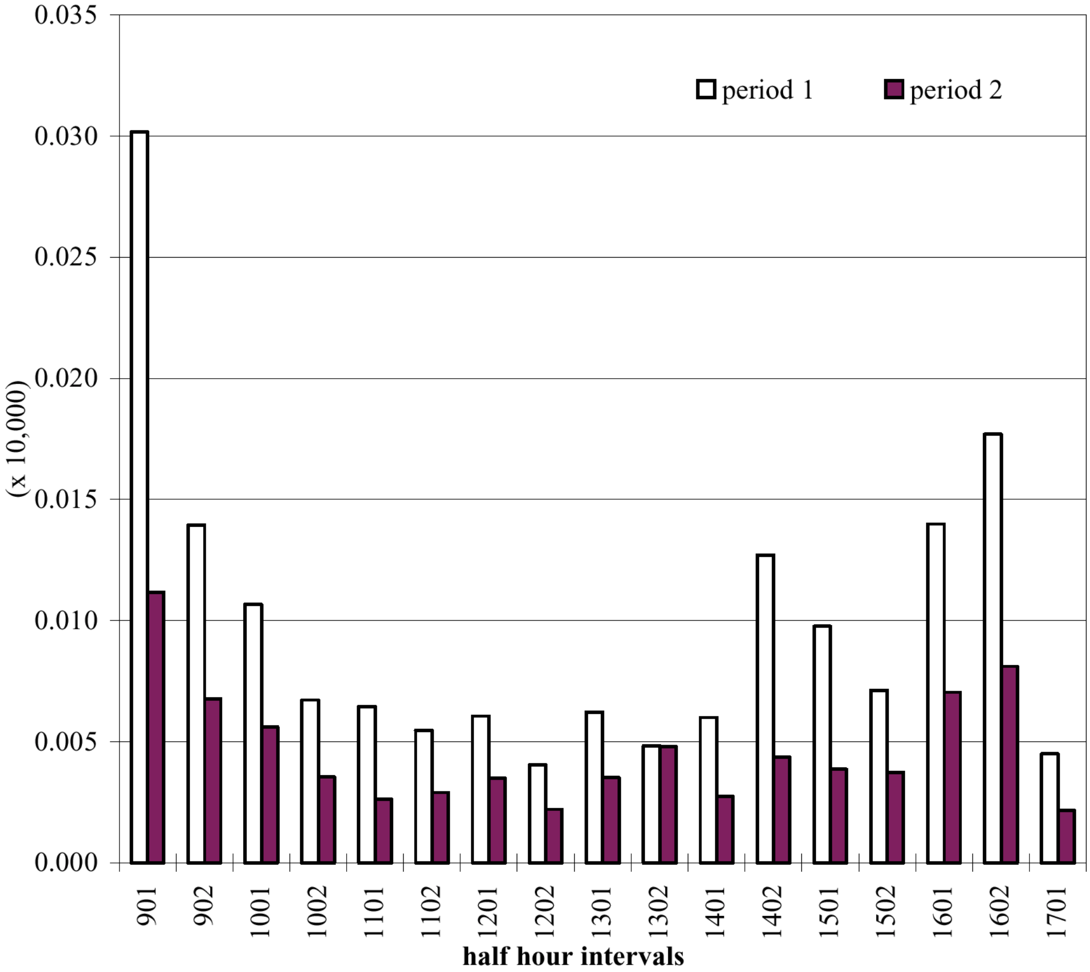
MACRO- AND MICRO-ECONOMIC EFFECTS ON PRICE IMPACT



(1) Price-impact estimates on MTS, in basis points, are moving averages over three days.

Figure 8

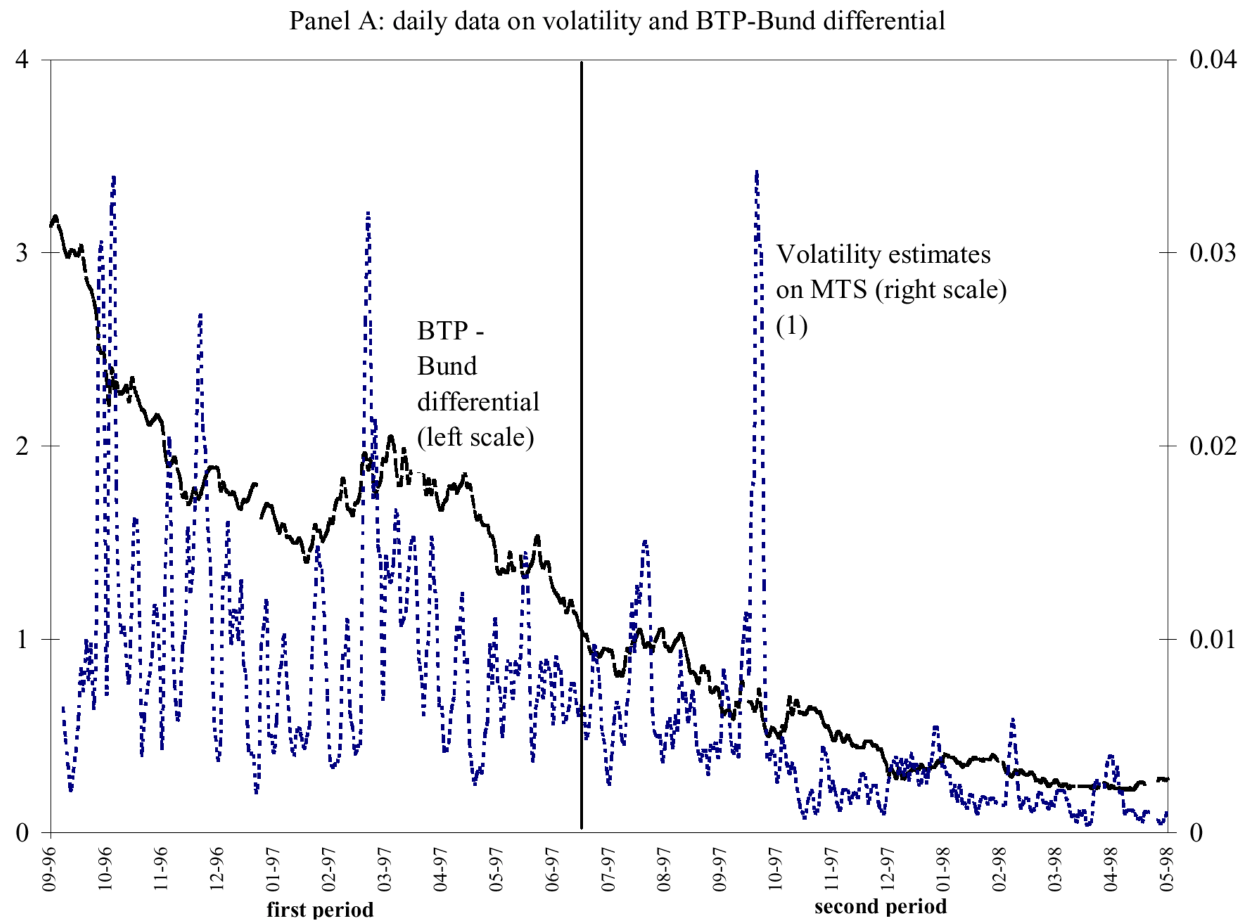
INTRADAY PRICE VOLATILITY ON 10-YEAR BENCHMARK BTP (1)

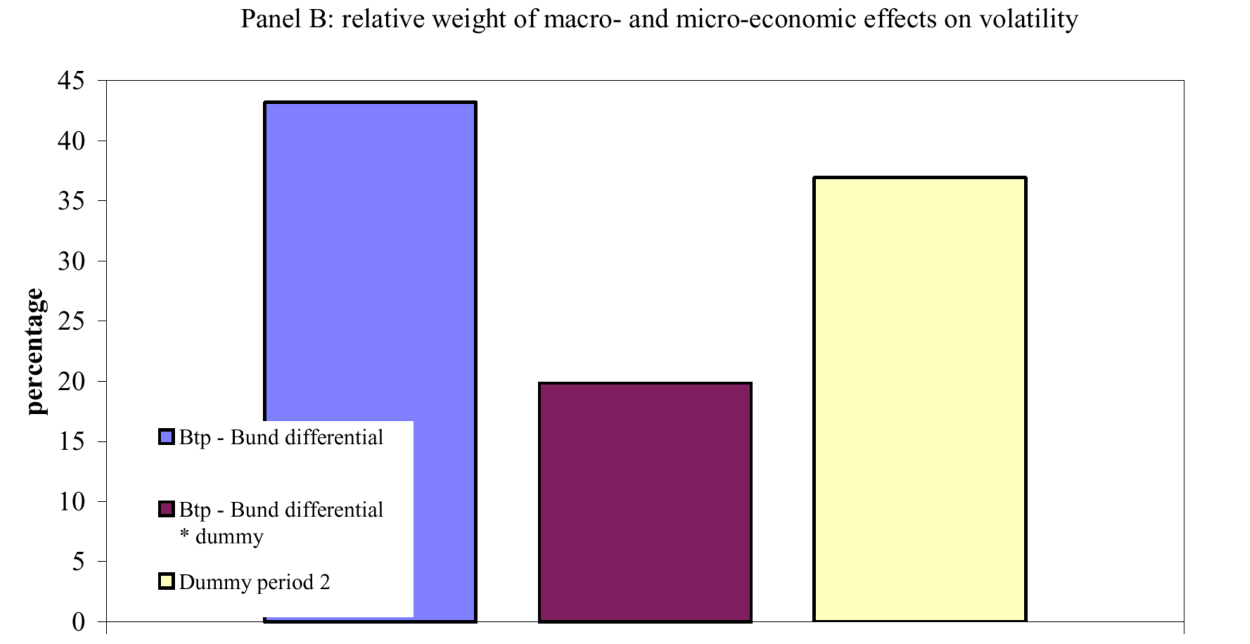


(1) The Kolmogorov-Smirnov test rejects the null hypothesis of identical distributions in period 1 and 2 for 17 out of 17 intervals, at 5% level. The *t*-test rejects the null hypothesis of identical means for 16 out of 17 intervals, at 5% level.

Figure 9

MACRO- AND MICRO-ECONOMIC EFFECTS ON VOLATILITY





(1) Volatility estimates on MTS, in basis points, are moving averages over three days.

Appendix

The development of the Italian Government Bond Market in the last decade

Year	International integration	Changes in market microstructure	New instruments
1998	Liberalisation of capital flows (partial)	Inception of MTS Start of regular reopenings of Treasury auctions Floor to bid prices abolished for T-bills, uniform price auction introduced for other bonds.	
1990	Liberalisation of capital flows (full)		Real-time securities transferral at the central depository Banca d'Italia
1991	10-year BTP futures at LIFFE (London)		
1992			Inception of the Italian futures market (MIF)
1993	First US$ global bond issue by the Republic of Italy. Prohibition of direct financing of the Treasury by the Banca d'Italia		First insurance of 30-year BTPs
1994	Reform of MTS	Treasury starts publishing timetable of auctions Electronic bid submission at auctions Reserved reopenings for "specialists in government activities" Continuous trading on MOT, the electronic retail market	
1995			First issuance of CTZs (two-year zero coupon bonds) CCT indexation fully matched with contemporaneous six-month bills
1996	EU investment Service Directive made effective		
1997	Withholding tax abolished for foreign investors Remote access to MTS for foreign primary dealers	Monitoring functions to the MTS management board	Treasury bond repo trading starts on MTS
1998		First ad hoc reopenings of Treasury auctions	Book-entry system for all new treasury issues Coupon-strips traded on MTS

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Comments on "Does market transparency matter? A case study" by A Scalia and V Vacca

Agnes Van den Berge, Banque Nationale de Belgique

The paper "Does market transparency matter? A case study" discusses the influence of a decrease in transparency ***resulting from anonymous trading*** on the Italian MTS electronic trading system, a dealer system. The study supports the theoretical evidence that a decrease in transparency makes liquidity traders (those traders who know only the price process) worse off whereas large informed traders (those who know more about fundamental asset values than others) are better off because they can better exploit their private information.

The study indicates that the decrease in transparency was associated with an increase in market liquidity, a reduction in trading costs and in price volatility and with an increase in market efficiency, defined as the degree to which prices fully reflect all available information.

Market transparency is usually defined as "the ability of market participants to observe the information on the trading process". However, transparency has many dimensions because a market has many kinds of participants and many types of information. The information can either be public (available to all market participants eg publicly announced statistics) or private (not available to all market participants and including both inside information about fundamentals and information on order flows or customer behaviour).

The study discusses the impact of a change in transparency of the Italian MTS resulting from a move to anonymity where the names of the market makers who post bid and ask quotes were no more revealed. In the study the change of transparency only treated one type of information in a specific market microstructure namely a quote-driven dealer market.

Academic findings are far from conclusive regarding the relationship between the level of transparency and the liquidity of bond markets. But in a dealer market, such as MTS, which is yet highly transparent, decreasing certain kinds of transparency can sometimes be beneficial. However, this may not lead to general conclusions about the relationship between transparency and liquidity which the study does not but I would like to stress the importance of this for the audience.

Transparency of market information has two aspects namely pre-trade quotes and post-trade information on prices/quantities actually transacted.

An early disclosure of information on specific orders, including the names of the dealers posting the orders linked to the size of these orders, does indeed appear counterproductive for the liquidity of dealer markets because of the risk of disclosing the movements in the market-maker's books. A too immediate (eg real-time) dissemination of this information to the market may reduce the incentive for dealers to make markets. Other elements of the pre-trade price transparency, such as the publication of aggregated volumes by limits, are beneficial to the liquidity of the market.

Where the study finds that the reduction of pre-trade transparency (move to anonymity) has had positive effects on the liquidity of the market, it does not give any indication of the effects of a disclosure of more detailed information after trade execution (post-trade transparency). Post-trade transparency makes markets fairer but it becomes harder for market makers to unwind positions quietly as prices would be more responsive to trades. Therefore it also may reduce liquidity. In a dealer market, the right balance has to be found between the level of transparency and the interests of the involved market-makers.

After these general observations on the interaction between transparency and liquidity, the reading of the paper leads us to three, more specific remarks and comments.

Firstly, the data used in the empirical analysis are the transactions in the period from September 1996 to end May 1998. It is mentioned in the study that macroeconomic effects have had a very important impact on the market performance. It has to be stressed that during the second part of the period under review, financial markets in Europe and thus also government bond markets were largely influenced by the impending introduction of the euro and the gradual emergence of pan-European financial markets. This was certainly the case in Italy.

Secondly, in accordance with the theoretical models, the decrease in transparency caused a reduction in the number of small traders. Even if the number of the larger traders increased, the overall number of market participants declined. In a context of bond markets where a limited number of global market-makers captures a growing size of the order flow from institutional and retail investors, the increasing market concentration should be a matter of concern because of its negative impact on the liquidity of the market. This applies even more to smaller markets where the number of market participants often is limited.

Finally, the empirical evidence shows that the reduction in transparency was accompanied by a decline in volatility but that the microeconomic effect (introduction of pre-trade anonymity) explained only 37 per cent of this evolution. Should the anonymity move have been taken today, then it seems doubtful that there would be a significant impact on volatility. At present, the most important dealers in the eurozone automatically derive their posted prices from a set of (exogenous) parameters like the corresponding yield of the Bund-future.

Let me now give a few remarks on the Belgian experience with electronic trading platforms for government bonds.

MTS Belgium has been operational since 5 May 2000. It introduced the Italian model and started directly with anonymous trading. At present, the dealer quotes and the order book are only available to market participants. Discussions are under way with information vendors to allow them to disseminate market information on their screens (with a certain delay) which will improve the market transparency.

All fixed rate OLO bonds with a remaining life to maturity of over 1.25 years are currently traded on MTS Belgium (16 bonds), representing a total outstanding amount of 138 billion euro.

For the time being, 16 primary dealers in Belgian government bonds and 1 market maker have access to the system. Market access will further be extended to domestic and foreign financial intermediaries in the capacity of price taker.

Five Belgian bonds are currently traded on ***EuroMTS*** representing a total outstanding amount of 48 billion euro. The first introduction took place on 9 September 1999.

Since 3 July 2000, twenty Belgian bonds are traded on ***Broker-Tec***.

The market share of electronic trading can be estimated at roughly one third of the total turnover of purchase/sale transactions in OLO bonds. Turnover in OLO bonds on Broker Tec has been marginal since its launch.

The electronic trading of Belgian government bonds, especially their introduction on EuroMTS followed by the launch of a domestic MTS, has resulted in lower transaction costs in terms of fees and bid/ask-spreads compared to the OTC market. This is also due to the straight through processing facilities MTS provides. As such, it has improved the liquidity of the secondary market. Further improvements in market transparency and an increasing number of market participants should give an additional boost to market liquidity in the future.

Conclusions

If conceptually a totally transparent market should be favourable for the liquidity of bond markets, practically in a dealer market, a compromise has to be found between the level of transparency and the involvement of the market-makers who expect a return on the capital invested in the market making activity. In Italy, where a reduction in market transparency (move to anonymity) has led to an improvement of market liquidity, the MTS market seemed initially “too transparent”. By contrast, the introduction of electronic trading in Belgium increased market transparency which, together with lower transaction costs, improved market liquidity.

Comments on “Does market transparency matter? A case study” by Antonio Scalia and Valeri Vacca

Peter Rappoport, JP Morgan

The asymmetric information view of markets points out that traders’ behaviour will be driven in specific ways by the environment in which they trade. It predicts, for example, that bid-offer spreads should be wider, the greater is the chance that a market maker has to trade with informed individuals, and that liquidity may be lowered by transparent trading, because transparency limits the return to market-making. Any theory that can successfully predict liquidity conditions would very quickly find a place as a market staple.

To test such a theory, a controlled experiment would be the best, and one actually appears to have been provided by the move of the MTS system to anonymous trading in mid-1997. The authors argue cogently that this change in rules would have four principal observable implications. Some are related to the shift in the “balance of power” towards large market specialists, to the detriment of MTS “liquidity traders”, and those in the OTC market who faced a reduced flow of information. Others follow from the change in the optimal trading strategy under the new rules.

The authors provide a clear and erudite exposition of an impressive battery of tests of these hypotheses. With one exception, they find that things moved in the predicted direction following the switch to anonymous trading. As liquidity improved following the introduction of anonymous trading, they suggest that anonymity may be a desirable feature to incorporate into the design of new markets.

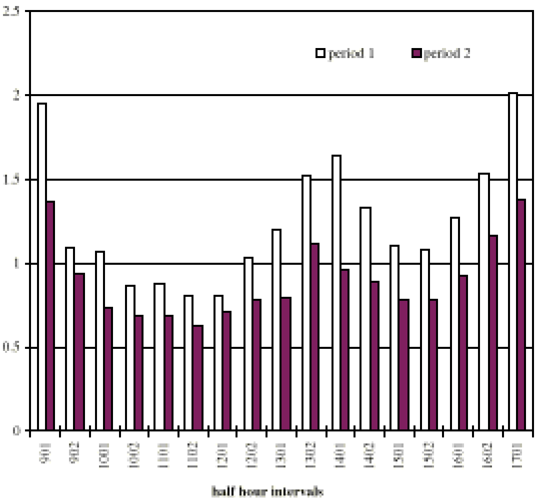
My interest in this paper is in where it leaves one on the broader questions related to liquidity I outlined at the start. Should the paper’s evidence give one a new respect for the asymmetric information dimension of market microstructure, or is there something else going on? Essentially, the question is not so much one of whether the predicted directions of the responses to lower transparency are confirmed. It is more a matter of whether the effects of the change are large, relative to the other influences on the way markets trade. Here, things are less clear.

For example, the only hypothesis contradicted by the authors’ evidence is that the OTC market should have suffered, because its ability to “free-ride” on information about sources of MTS flows was curtailed. However, the authors cogently argue that there were other extenuating circumstances at the time that could have led to the continued growth in the OTC share of the market. But then, how big are these extenuating circumstances in other instances?

I have little doubt that asymmetric information considerations are present where prices are set by market makers. But at least in the bond world, it seems hard to believe they are dominant. The only information advantage that seems to be around in government bonds concerns not so much an inside track on fundamentals, such as interest rate policy or macroeconomic news, but knowledge of flows, ie that a big liquidity trade is imminent, from which profits can be made by “positioning ahead”. Here, there is no winner’s curse in having traded with the informed: the fundamental value of the securities bought (sold) has not necessarily fallen (risen). The only thing that has been missed is the opportunity to make life difficult for the trader who knows of the liquidity flows, and, thereby, to increase the chance of gaining some of the returns from the flows for oneself. Asymmetry of information may be more important in equity markets, where information on individual stocks’ fundamentals can plausibly flow slowly enough among market participants for market makers to worry about the winner’s curse. However, in corporate bond markets, which presumably dance to the same fundamentals tune as equity markets, inventory management appears, to me at least, to be a more pressing concern, and a more proximate determinant of bid-offer spreads.

The evidence presented by the paper does have something to offer on the importance of asymmetric information, but it is not very encouraging. As predicted by the theory, a move to lower transparency should lower the bid-offer spread. Figure 4 in the paper, reproduced below shows the bid-offer spread during half-hour periods in the trading day.

INTRADAY SPREAD ESTIMATES ON 10-YEAR BENCHMARK BTP (1)



Indeed, the curve shifted down once trading was made anonymous. However, the magnitude of the shift is small in comparison with the fluctuation in bid offer throughout the day, both before and after June 1997. This cycle is also evident in BTP prices and price volatility. Bid-offer seems to be widest at the times when, perhaps, traders are least closely focussed on their screens. Is wide bid-offer a simple way of trading on autopilot? Why doesn't someone in this highly competitive market quote a narrower bid-offer at these times? Or are they times when there is a higher density of informed traders? Probably not: Figure 2 shows that the volume of trading is lowest at the times when bid-offer is widest.

One can engage in the obvious drole speculations about what drives the daily cycle. However, the simple fact is that when you trade on MTS has more effect on the liquidity you will experience than the rules under which you trade. So it seems like the first order of business is to understand why these fluctuations can take place. And at first blush, here as in other instances mentioned above, asymmetric information does not appear to be the most promising answer.

Short introduction on the work of the Johnson-group

Eloy Lindeijer, De Nederlandsche Bank

Following the liquidity crisis of autumn 1998 the Committee on the Global Financial System formed a working group to examine the events surrounding the market stresses that evolved during that period. Special focus was given to the suddenness of liquidity deterioration and mechanisms that contributed towards the widespread withdrawal of risk taking. The group, which was chaired by Karen Johnson of the Federal Reserve Board, included representatives of central banks and monetary authorities from G-10 countries and emerging markets. Market participants were interviewed in financial centres in Europe, Asia and the Americas to get a global perspective and the group assembled a large data set on key financial indicators to compare the events of end-1998 with previous episodes of market turbulence.

With regard to the run up to the liquidity crisis the Johnson-group noted that risk spreads and volatilities in G10 markets were on the low side of experience, particularly given the market turbulence in Asia during 1997 and 1998. One explanation put forward was the scale of resources devoted to risk arbitrage, driven by widespread emulation of relative value strategies pioneered by firms such as LTCM and the perception of risk free bets. This had helped reinforce linkages across markets, narrowing spreads and reducing volatilities in a way that reinforced macro-economic trends, such as declining capital market interest rates. This 'overextention' led to a severe market reaction, for which Russia's de facto default served as a trigger.

Market dislocation reached a peak during the recapitalisation of LTCM as worries about mass sales, risk concentration and above all systemic risk led to widespread risk withdrawal. Indicative of the severe hedging pressures related to the unwinding of so-called carry trades was the 9% move in usd-japanese yen on October 7-8. In some niche products (such as Danish mortgage backed securities) arbitrageurs had built positions that amounted to multiples of typical daily market turnover, thus making it difficult to unwind positions without substantial market impact. In fixed-come markets this phenomenon led to an explosion of swap and credit spreads. 1998 also produced some of the largest stock market declines for the decade, in some case just falling short of the 1987 crash. Volatility of bond and credit spreads, however, generally remained below the peaks experienced in 1991 and 1994.

Compared with earlier episodes of market turbulence the Johnson-group concluded the autumn of 1998 was an unusual, but not an extremely rare event. It demonstrated that the 'abnormal' returns earned by relative value arbitrage players were in fact not risk free when viewed over a longer time span. Policy makers concluded that some of the mechanisms that produce contagion and amplify market turbulence can be addressed by better risk management, capital and disclosure. Interestingly, some commentators argue the opposite case: that the response of policy makers may actually contribute towards greater herd behaviour in the future (triggered by disclosure of market sensitive data) and has already resulted in more permanent reduction of market liquidity in the main markets, as market makers and speculators have withdrawn risk bearing capital. My personal inclination is to argue that liquidity has not deteriorated in global financial markets, but has returned to more sustainable levels subsequent to processes of deleveraging. With regard to herding it is important that policy makers create an environment that makes it less likely that such flows create systemic concerns. I would argue that disclosure and transparency increases the resilience of the financial system through greater market discipline.

A second important but not unrelated issue concerns the impact of technology on the functioning of markets. The emergence of the e-commerce is impacting market liquidity today in ways that are not very well understood. Competing electronic trading platforms may be contributing to a fragmentation of liquidity in previously centralised markets. Alternatively new smart order routing systems may be providing a single portal to all available pools of liquidity, effectively re-creating one central market place. The current state of liquidity of global financial markets may be an interesting issue for further discussion today, time permitting.

Let me conclude with one of the remaining imprints of the Johnson-study for my work as a central banker. Understanding the structure and functioning of markets is important, both form a monetary policy and systemic risk perspective. End-1998 some of the most visible manifestations of market stress occurred in markets not always directly followed by central banks. As more risks become priced in a world that is marked to market, central banks have realised the need to build up expertise outside the traditional realm of money and forex markets. Developments in swap, credit and equity markets and the impact of e-commerce have become important and are therefore monitored on an on-going basis. The work programme of the CGFS is a good example of this.

Market liquidity under stress: observations from the FX market

Francis Breedon, Lehman Brothers

Given the thorough exposition of the theoretical aspects of liquidity given elsewhere in the conference, I will focus my remarks on a market perspective to these issues. First, outlining in general how periods of market stress appear from an FX dealing floor, and second looking in detail at whether the euro could be subject to such an event.

FX markets under stress

Market makers tend to identify two types of stress events in FX markets, the first involving high volatility and high turnover the second high volatility and low turnover. In general, high volatility and high turnover events are good times for market makers, since they can achieve the seemingly contradictory combination of keeping quoted spreads relatively narrow while substantially increasing their profit per trade. This is achieved through higher effective spreads whereby market makers find it easier to "read" incoming trades and so quote accordingly. Examples of this type event include the periodic 10 yen moves of dollar/yen, and the fallout for a developed country currency crisis. Although large trades are difficult to execute in such periods, liquidity in terms of the ability to execute a trade remains, the cost is that effective spreads are wide.

High volatility low turnover events are rarer and represent bad times for market makers. These usually occur after a high volatility high turnover event but change in character as the market has lost two-way interest. Despite being able to read future trades, market makers are unable to make money as it is impossible to lay off trades elsewhere. Therefore, these any trades have a disproportionate impact on prices and market makers find themselves on the wrong side of these moves. The reaction is to widen quoted spreads considerably or not quote at all. Although these event are rare, the aftermath of the Asian currency crisis supplied a few examples. In particular, the Indonesian Rupiah went through an extended period of low turnover and high volatility. Interestingly, this occurred once international investors had almost totally withdrawn from trading, while domestic investors still had a strong demand for dollars.

From a policy-maker point of view it could be argued that high turnover high volatility events are unfortunate, but preferable to the low turnover variety. Certainly, the majority of currency moves in the Asian crisis occurred after the initial crisis, and the Rupiah's near 80% fall has had real and lasting effects on the economy.

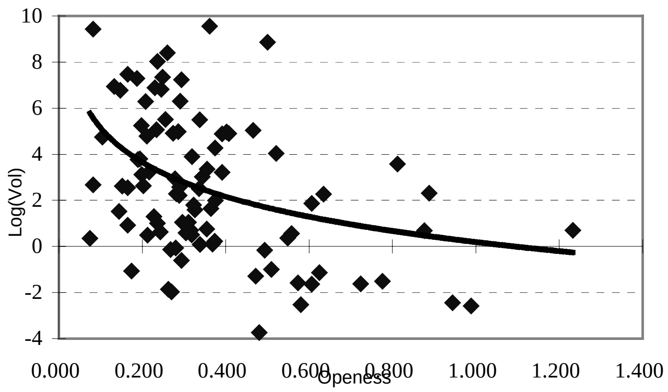
Liquidity prospects for the euro

Although the fall in value of euro/dollar since inception is its best known feature, that fall has been remarkably orderly, with very few stress events along the way. Certainly, in comparison to dollar/yen, which has made a number of one day 10% moves in recent years, the euro's moves seem orderly. Is this simply a coincidence or is the euro naturally more resilient to stress events?

The first point to note about the euro is that euro/dollar volatility has changed relative to dollar/DM. The key reason for this is the phenomenon of 'benign neglect' whereby central banks of large closed economies put less weight on the exchange rate than those of more open ones. The reason for this is simple, with imported goods making up a smaller proportion of CPI, the inflation impact of exchange rates movements is less in more closed economies. The chart below shows the impact of this benign neglect for a sample of nearly 100 countries, more closed economies do indeed seem to have more volatile exchange rates.

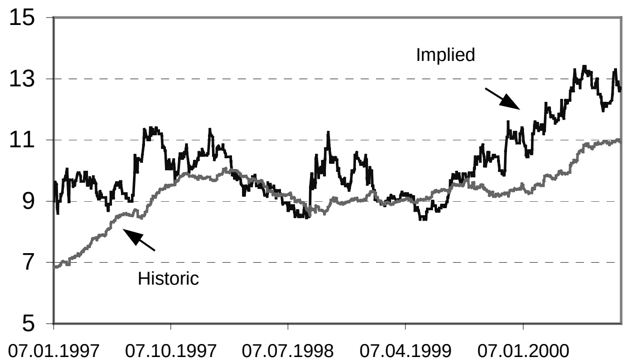
Exchange Rate volatility and openness

(Imports as % of GDP vs. log real effective exchange rate volatility)



Volatility of the DM and Euro

(1 Year Historic and Implied $/DM volatility)

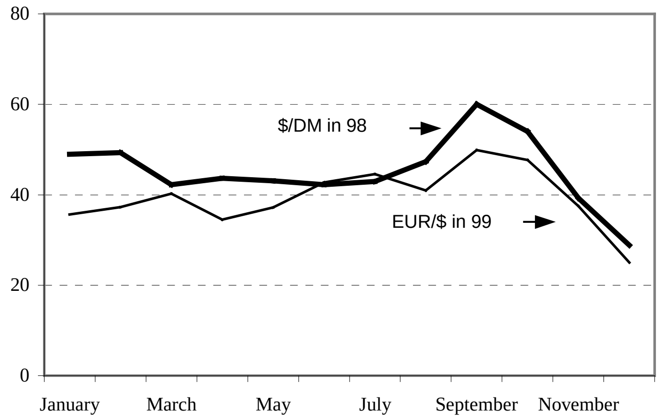


In the case of euro/dollar, the impact of benign neglect can be seen in the chart above as the volatility of euro/dollar has risen well above that of dollar/DM before it. This is true in both actual historic volatility, and expected volatility (from FX options)

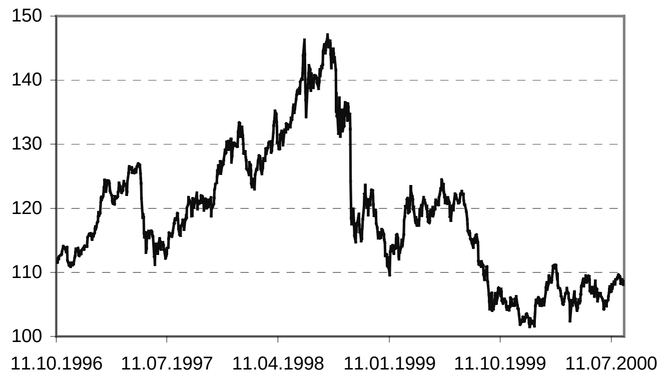
Overall, benign neglect suggests that euro/dollar will be volatile, but will this volatility escalate into stress events? One piece of evidence that suggests it may is shown in the chart below. Turnover in euro/dollar has fallen markedly relative to dollar/DM. This lower turnover may indicate less depth in the market and so a higher propensity to stress events.

EBS turnover in $/DM and EUR/$ - 1998-99

(Monthly average of daily turnover in $bn, source EBS)



$/JPY



However, a closer look at stress events in dollar/yen suggest that although euro/dollar is more susceptible to these events than was dollar/DM, we are unlikely to have one soon. As the chart above shows, dollar/yen has had three main stress events (daily movements of around 10%) in the last 4 years (May 1997, September 1999 and October 1999). A notable feature of these events is that the have all been sharp appreciations of the yen. ("dollar/yen goes up by the stairs and down by the lift" as some traders put it). The reason given for this is that the large interest rate differential between the US and Japan (currently around 6%) makes being long dollar/yen a standard carry trade (i.e. even if the exchange rate does not move, the interest rate return of borrowing yen and depositing in dollars is substantial). For a number of reasons, it is carry trades like these that are the most susceptible to stress events, and tend to have skewed returns such that they generate small positive returns for long periods and then one off large negative returns.

Since the US euro-area interest rate differential is still quite small (about 2%) euro/dollar does not yet have the characteristics of a standard carry trade. However, carry trade analysis would suggest that if euro/dollar were to have a sharp move, it is more likely to be an appreciation than depreciation.

The puzzling decline in financial market liquidity[1]

Avinash Persaud, State Street

Market participants frequently complain about lack of market liquidity - the ability to get in and out of financial markets without driving the price against you. The worrying thing is that complaints of illiquidity are even levelled against the largest markets, the smooth functioning of which have economic importance. Consequently, illiquidity and the resulting dislocation of markets, also worries central bankers. The foreign exchange market, for example, is the largest market by turnover - daily turnover is an estimated $1.8 trillion - or 20 times that of the New York stock exchange. But on 28 February 2000, the euro fell 5% against the yen in five hours as the market absorbed a modest sale of euros. The difference between hedging foreign exchange exposure that morning or afternoon could have wiped out a year's profit margin for a typical exporter. Previously, as turnover was the only observable measure of liquidity and markets have grown rapidly, complaints about liquidity have generally been put down to a few freakish events and investors being unwilling to give up hard-earned returns to market makers. However, new data analysed below suggests investors' tight hold has been well founded and that, in many markets, liquidity never recovered from the body blows of 1998. It would appear that the forces that are sapping market liquidity are growing inadvertently aided unsuspectingly, by the activities of unsuspecting regulators.

Liquidity - measured by the price-impact of trading

There are few measures, but many different meanings, of liquidity. In a macroeconomic sense, liquidity often refers to interest rate or money supply conditions influenced by a central bank. In a microeconomic sense, liquidity refers to the ability or cost of transacting in markets. The two concepts are related but separate and it is the latter meaning of liquidity, the cost of transacting, that I wish to focus on. Poor liquidity equates to large transaction costs which limit the size and growth of markets and is likely to result in a loss of economic welfare. In most financial markets, transaction costs comprise commissions and the price impact of trading - how sensitive prices are to an additional purchase or sale. As commissions are typically fixed or move with market conditions (which are proportional to price-impact of trading), a good measure of changing liquidity conditions is simply a measure of the price impact of trading. What exactly does this mean?

In liquid conditions, all else being equal, the decision by one section of the market to sell the euro on one day should not drive the euro lower. The co-movement of flows and prices - or, put another way, the price impact of selling - should be small. On another day, when liquidity conditions have worsened, the same sale might drive the euro down and have a bigger price impact. To reveal this shift in liquidity conditions, one needs to observe not just the development of prices, but underlying gross transactions. This has not been done before because, while price data is available, reliable and representative, flow data was not - until recently.

Using State Street's $6.2 trillion of portfolio holdings data - about 12% of the world's tradeable securities - Froot, O'Connell & Seasholes (1999) have calculated a daily measure of the price impact of buying or selling of foreign securities faced by cross-border investors in 42 countries. Correlations between State street's portfolio flow data and, where available, official data range from 0.70 to 0.80 in major markets, suggesting that the data is representative. In essence, their methodology involves regressing returns on contemporaneous cross-border investor buys and sells. The coefficients on buys and sells are published as the Liquidity Index. The regression equation takes into account the timing of trades, the "anticipation effect" and outliers (see FDO Partners & State Street Bank, 2000). The results challenge many assumptions. In terms of the price impact of trading foreign equities, the liquidity faced by cross-border investors is substantially lower today than in 1997. In emerging markets, liquidity conditions are even lower than during the turmoil of 1998.

1 This was previously published in the June 2000 edition of Risk Magazine.

The declining path of liquidity

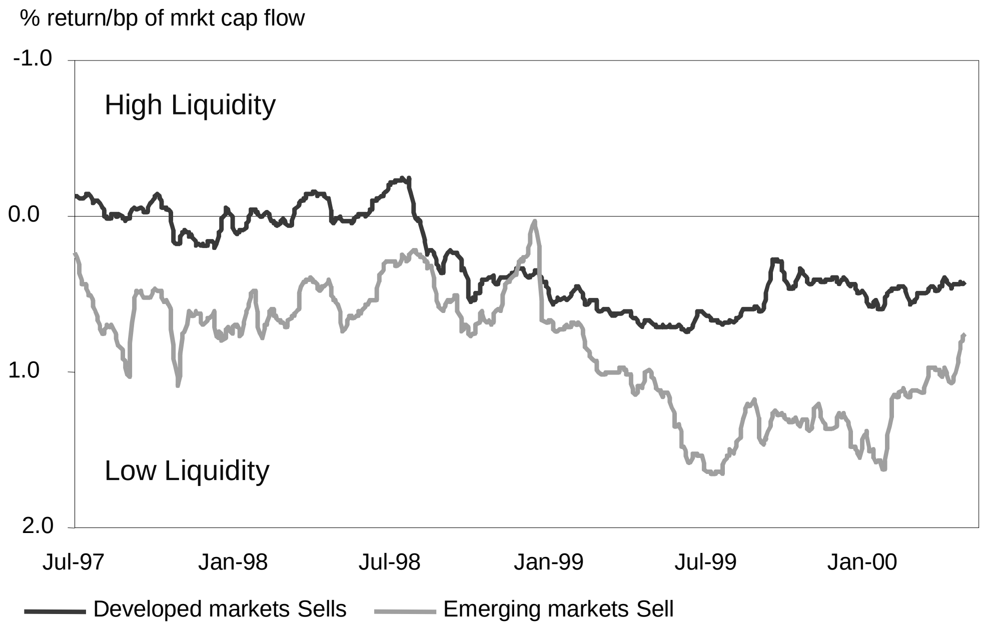
Up to June 1997, liquidity was high in both developed and emerging markets for cross-border investors. From that point, emerging market liquidity started a long decline. This decline accelerated after the South Korean "devaluation" in November 1997, received a temporary boost at the time of the three interest rate cuts by the US Federal Reserve and talk of new lending facilities for the International Monetary Fund, and then slumped back ahead of year 2000 computer concerns in 1999. The crisis in Korea appears to have had a more systematic impact on flows and liquidity than the previous devaluations in Thailand, Indonesia and Malaysia.

It was only just before the Russian crisis in August 1998 that liquidity conditions started to turn down in developed markets. Developments in South Korea and Russia proved systemic. Year 2000 concerns appear to have been concentrated on emerging markets, having little liquidity impact on developed markets. While the US Federal Reserve Bank's interest rate cuts at the end of 1998 offered some temporary respite for liquidity in emerging markets and arguably offset the impact of the collapse of the hedge fund, Long-Term Capital Management, it appears to have been less successful in supporting liquidity in developed markets, see Chart 1.

Chart 1

Liquidity index for emerging and developed markets
- a measure of the price-impact of trading

(based on State Street's $6 trillion of portfolio holdings data)



Markets are bigger and thinner

The decline of liquidity across developed and emerging markets contrasts strongly with the sharp rebound in foreign security prices in 1999, for both developed and emerging markets. It also contrasts strongly with recovering portfolio flows since October 1998 and growing turnover in the foreign exchange market. The net decline in cross-border portfolio flows in developed and emerging markets between late 1997 to October 1998 was reversed by June 1999, in the case of developed markets. Flows returned, but liquidity did not.

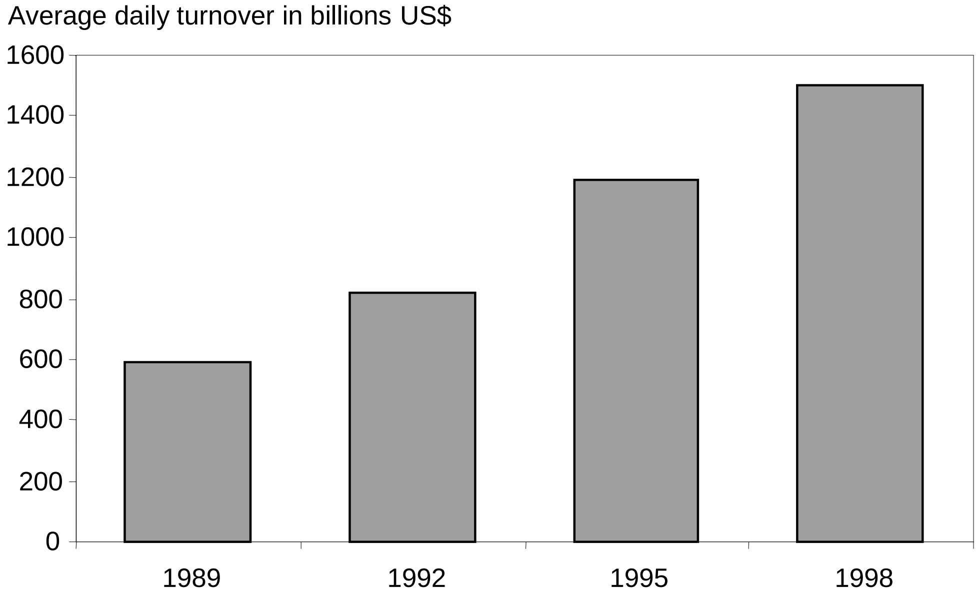
Explaining the liquidity puzzle

There are three factors that may explain this liquidity puzzle. First some have argued that liquidity may have been hindered by the uncertainty attached to two one-off factors: the arrival of European economic and Monetary Union (EMU) in January 1999 and the anticipation of the year 2000 "bug" in January 2000. We have seen that the latter may have impacted liquidity in emerging markets. But

EMU should have been a force for increased liquidity in the securities markets and, although foreign exchange turnover may have been 5-10% higher without EMU, foreign exchange turnover has not actually fallen. It is hard to explain this decline in liquidity by one-off factors alone.

Chart 2

Daily turnover in the foreign exchange market



ECNs and electronic broking

A second factor drawing liquidity has been the rise in the number of exchanges, or perhaps more accurately, the rise in the number of different ways of transacting, such as electronic alternative trading systems (ATS) and electronic communication networks (ECNs). ECNs such as Instinet and Island have recently taken a large part of the market share of trading on the Nasdaq (the world's second largest equity market) and Electronic Broking System has established a significant share of the foreign exchange markets. These systems operate well and reduce costs when markets are large and participants have different views. As a result they draw liquidity away from outside these systems. However, when markets are small or participants adopt the same view, no one is obliged to make a market on the crossing network or broking systems so liquidity vanishes and little is left outside those systems to help. Given the economic importance of liquidity, regulators and central bankers ought to pay more attention to supporting the development of liquid exchanges without harming innovation. There is a growing demand for liquidity. Pension fund managers initially rejoiced at the recent announcement of the merger of the London Stock Exchange with the Deutsche Bourse principally because of the liquidity benefits.

More herding, less liquidity

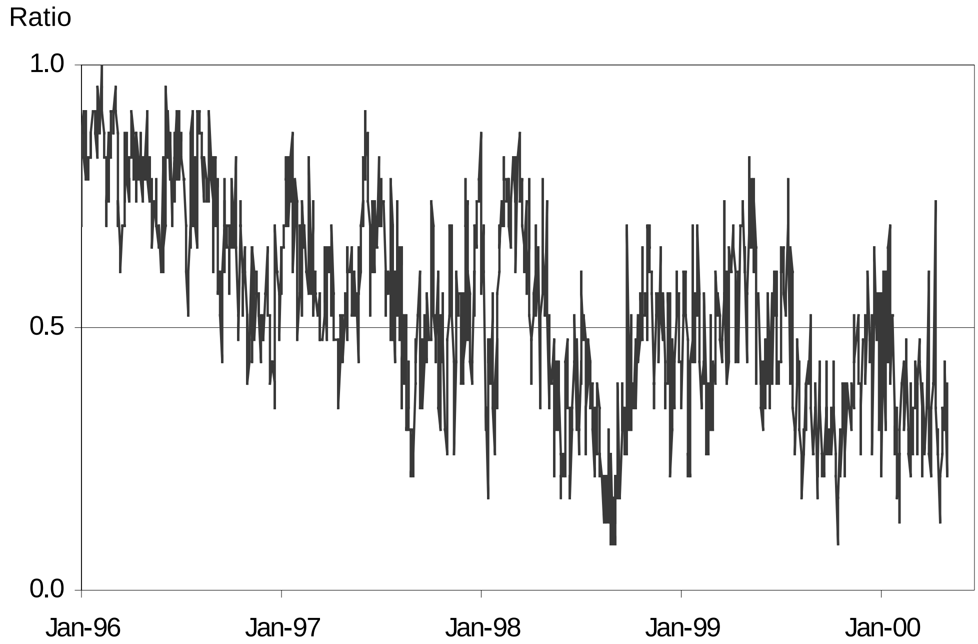
The third, related factor is that markets are increasingly driven by herding investors. Two decades ago, international capital flows were strongly related to international trade and banking flows that helped to finance the shortfalls. One decade ago, absolute return investors taking bets on currencies and bond markets played an important role in boosting liquidity. Today, international capital flows are driven by institutional investors seeking better returns than their competitors.

There is plenty of theoretical support and empirical evidence that relative-return investors behave like a herd, rushing in and out of markets together. Increasing the size of the herd does little to support liquidity, indeed it could reduce it. If everyone is trying to enter or exit the door at the same time, having more people do so doesn't make it easier to squeeze past. One measure of the herd is portfolio flows across 23 emerging markets. Although these 23 markets exhibit very different economic fundamentals, politics and markets, there are times when all 23 markets are receiving net inflows as in most of 1996 when buying emerging markets was all the rage, or in late 1998, when almost all 23 were receiving net outflows. Chart 3 shows the ratio of the number of emerging market countries receiving net cross-border inflows over the previous month (or more precisely, rolling 20 business days). This ratio can be used as a measure of investors' preference or aversion to risk more generally.

Chart 3

Ratio of emerging markets receiving net cross-border inflows

(based on State Street's $6 trillion portfolio holdings data)



Risk management may make matters worse

To some extent, the combination of herding investors and tighter risk management systems has aggravated liquidity conditions. Herding investors hit their risk management limits together, causing them to sell the same markets together, reducing liquidity and causing those markets to exhibit even greater volatility than in the past, which in turn feeds back into risk management systems and leads to further sales. Thus, the cycle becomes self-feeding. In these environments, markets need contrarian investors able to run sizeable losses and buy near the bottom in the hope that they will be the first to benefit from the turnaround. Interestingly, unregulated, hedge funds are more likely to do this than regulated investors and banks. Indeed, given that such hedge funds are putting the risk capital of wealthy professionals at risk (and not the life savings of the average person in the street), they are eminently suitable for this task.

Hedge funds may support liquidity

The irony is that, by pressuring counterparties to insist on greater disclosure and tougher collateral requirements, regulators have been discouraging the kind of hedge funds that make big macro bets. In the process, they may have damaged market liquidity which in turn has made it more difficult for these hedge funds to operate. It is noteworthy that the pattern of liquidity faced by cross-border institutional investors in emerging markets follows very closely the dwindling size of macro hedge funds. This process has recently enveloped the twin icons of the hedge fund industry, Julian Robertson and George Soros.

Regulators need to examine how to achieve the right balance of regulation and provide support to liquidity. Given that herding by institutional investors continues, that ECNs and ATSs are growing in number and that big macro hedge funds are on the defensive, one can expect liquidity to remain in scarce supply. Consequently, the managing of liquidity will become increasingly important for every participant in financial markets. With this in mind, and armed with this new set of daily data on liquidity in 41 different markets (Froot, O'Connell & Seasholes, 1999), State Street intend to initiate a regular series of research on liquidity, its path, its drivers and key changes.

Avinash Persaud, Global Head of Research

Summary of methodology

The Liquidity Index (see main text) was produced by FDO partners, LLC and State Street Bank. State Street Bank is the world's largest custodian with more than $6.2 trillion of assets under custody. The Liquidity Index is based on daily cross-border transactions data, collected by State Street Bank & Trust. From the universe of total daily transactions, State Street separates out only those transactions that originate in one country but settle in another currency - cross-border portfolio flows.

The methodology involves regressing returns on contemporaneous buys and sells. The coefficients on buys and sells from the regression are published as the "Liquidity Index - a measure of the price impact of trading". The regression specification takes into account the timing of trades, the "anticipation effect" and outliers.

This measure is both a measure of liquidity faced by cross-border investors and a measure of whether cross-border flows are "driving" the current direction of a market or not. Either interpretation can be justified.

Adjustment to the data

Timing of trades, anticipation effects, outliers

In order to take into account the fact that a trade can take place in the same day but across several time zones, each day's data is actually the average of returns on that day and the preceding day.

Earlier work by Professor Ken Froot et al has shown that State Street's portfolio flows have predictive power for future equity returns. To try to isolate the contemporaneous price impact of trades, this correlation is controlled for by including lagged terms of buys and sells in the regression specification.

A single outlier can cause coefficients to shift wildly, as the outlier enters the rolling window used to estimate the regression and when it leaves. In order to reduce these distortions outliers are removed in the following way. Each day, for each country, the amount of bought and sold (as a fraction of market capitalisation) is compared to the mean buy or sell trades over the past trading days. Any point that is more than four standard deviations away from the mean is replaced with the mean.

Regression equation

To estimate the price impact cost, we run a regression of returns on contemporaneous buys, contemporaneous sells, lagged buys, lagged sells and a constant.

$$R_{i,t} = \alpha_i + \theta_{B,i}\,Buy_{i,t} + \theta_{S,i}\,Sell_{i,t} + \Lambda_{B,i}\,Buy_{i,t-p} + \Lambda_{S,i}\,Sell_{i,t-p} + \varepsilon_{i,t}$$

where:

$R_{i,t}$ = the average daily return of market "i" over the week from "t–(p+1)" to "t"

$Buy_{i,t}$ = the average fraction of market capitalization of market "i" bought by clients of State Street Bank over the week from "t–(p+1)" to "t"

$Sell_{i,t}$ = the average fraction of market capitalization of market "i" sold by clients of State Street Bank over the week from "t–(p+1)" to "t"

The two parameters of interest are:

$\theta_{B,i}$ parameter related to the price impact of foreign buying

$\theta_{S,i}$ parameter related to the price impact of foreign selling

We expect the parameter $\theta_{B,i}$ to be positive most of the time, indicating that foreign purchases happen contemporaneously with price increases. During periods when the price impact costs of buying are high, the $\theta_{B,i}$ parameter will be more positive than during times when the price impact costs are low.

Likewise, we expect the $\theta_{S,i}$ parameter to be negative, indicating that foreign sales happen contemporaneously with price decreases. During periods when the price impact costs of selling are high, the $\theta_{S,i}$ parameter will be more negative than during times when the price impact costs are low. Here because the $\theta_{S,i}$ parameter is multiplied by –1, it will be reflected into the positive y-axis.

The magnitude of the parameters depends on a number of factors including State Street's share of foreign trading in a particular market and the units that the flows and returns are recorded in (eg decimal form, percentages, basis points, etc). We express returns (in USD) in percentages (multiply them by 100) and flows in basis points of market capitalization (multiply them by 10,000).

Contemporaneous correlation

Returns are highly correlated in markets around the world. To account for this correlation, we estimate a system of regression equations and allow for correlation between the residuals. For one region or country, we have:

$R_{i,t} = \alpha_i + \theta_{B,i} Buy_{i,t} + \theta_{S,i} Sell_{i,t} + \Lambda_{B,i} Buy_{i,t-p} + \Lambda_{S,i} Sell_{i,t-p} + \varepsilon_{i,t}$

which can be written more succinctly: $Y_i = X_i \beta_i + \varepsilon_i$

We then stack a series of regions or countries:

$$\begin{bmatrix} Y_1 \\ Y_2 \\ \vdots \\ Y_N \end{bmatrix} = \begin{bmatrix} X_1 & 0 & \cdots & 0 \\ 0 & X_1 & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & X_1 \end{bmatrix} \begin{bmatrix} \beta_1 \\ \beta_2 \\ \vdots \\ \beta_N \end{bmatrix} + \begin{bmatrix} \varepsilon_1 \\ \varepsilon_2 \\ \vdots \\ \varepsilon_N \end{bmatrix}$$

$$\Sigma = \begin{bmatrix} \sigma_{11} & \sigma_{12} & \cdots & \sigma_{1N} \\ \sigma_{21} & \sigma_{22} & \cdots & \sigma_{2N} \\ \vdots & \vdots & \ddots & \vdots \\ \sigma_{N1} & \sigma_{N2} & \cdots & \sigma_{NN} \end{bmatrix}$$

The GLS estimator is:

$$V = \Sigma \otimes I_N$$

$$\hat{\beta} = \left[X'V^{-1}X\right]^{-1} X'V^{-1}Y$$

The GLS estimator $\hat{\beta}$ is actually a vector of region or country parameters $\hat{\beta} = \left[\hat{\beta}_1, \hat{\beta}_2, \ldots, \hat{\beta}_N\right]'$ where each region or country parameter is $\hat{\beta}_i = \left[\hat{\alpha}, \hat{\theta}_{B,i}, \hat{\theta}_{S,i}, \hat{\Lambda}_{B,i}, \hat{\Lambda}_{S,i}\right]'$. Again, we are interested only in the parameter estimates of $\left\{\hat{\theta}_{B,i}, \hat{\theta}_{S,i}\right\}$ which represent the contemporaneous co-movement of prices and flows.

All regions and countries are unable to be estimated together because regions are linear combinations of countries

For further detail see, Liquidity Index - Technical Document, FDO Partners LLC & State Street Bank, April 2000.

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Measuring liquidity under stress

Christian Upper, Deutsche Bundesbank

In my contribution to the panel, I want to focus on how to measure liquidity under stress. Since financial markets seem to behave quite differently in periods of stress than under normal conditions, measures of liquidity that tend to work well most of the time may no longer be meaningful in turbulent markets.

In line with much of the literature, I define market liquidity as the "ease" with which a security can be traded. The word "ease" can be substituted by the word "cost". The cost of trading includes commissions, fees and taxes as well as the bid-ask spread, the price impact of a trade, and the cost caused by prices movements if a trade cannot be executed at once or when it is split up into a sequence of smaller transactions. In normal times, the cost of trading is particular relevant for those investors who try to make a profit by trading on small movements in prices and, consequently, turn over positions relatively quickly. For such agents, even small differences in the cost of a transaction can make or break a trading strategy. Investors with longer time horizons shouldn't really worry about market liquidity.

In periods of stress, the relative importance of the individual factors determining market liquidity changes. Prices become much more volatile than they are normally, and a strong desire on part of the traders to close undesired positions seems to dominate all other motivations for trading. From conversations with practitioners, I have the impression that costs arising from the price impact of trades or the inability to close positions quickly when prices are moving against them tend to eclipse the bid-ask spread and administrative costs. As a consequence, any measure for liquidity under stress should meet two requirements:

1. it should reflect the costs of closing large positions quickly; and
2. it should be available at a relatively high frequency as even relatively brief bouts of illiquidity may impose heavy costs on market participants if prices are volatile.

In order to illustrate the issues involved, I consider the liquidity of 10 year German government bonds market during two episodes of stress. The first took place in the aftermath of the Russian debt moratorium on 17 August, and the second during the sharp revaluation of the yen vis-à-vis the US dollar on 8 and 9 October of the same year.[1] As you will find out, the microstructure of the German bond market is broadly similar to that of foreign exchange markets, so the discussion should have some relevance beyond the particular market segment considered.

Although German government bonds are listed, the role of exchanges is negligible, with the bulk of trading taking place over the counter. This is especially true for the inter-dealer market, where trading takes place either directly between banks or through inter-dealer brokers (Freimakler). These act as intermediaries, but have no obligation to provide firm quotes. More recently, a sizeable proportion of inter-dealer trading has shifted to the electronic trading systems Euro-MTS, but electronig trading did not exist in 1998. The OTC market seems to be altogether rather opaque. Participants observe quotes but do not receive timely information on other traders' past transactions beyond anecdotal evidence. The only comprehensive data set is maintained by the German securities regulator, which receives information on every single transaction in Germany, regardless of whether or not it takes place on an exchange. However, this data is not published on a regular basis - not even in aggregated form - although it has been available for research purposes.

After this brief excursion into the microstructure of the German government bond market, let me turn to our main topic. A measure for liquidity that is often used by practitioners is trading volume. As a macroeconomist raised on representative agent models, where markets can be liquid without any trades actually taking place, I initially found this surprising. Market liquidity is only one, and not necessarily the most important, factor affecting trading activity. For example, if traders have sufficiently large hedging needs, then they may trade even if market liquidity is relatively low. What volume does tell us, however, is whether trading was possible at all.

[1] The analysis is based on Upper (2000).

Turnover in the German government bond market was particularly high just after the Russian debt moratorium, with a high average size of trade. Both turnover and the average trade size declined throughout September, and remained relatively - but not unusually - low on 8 and 9 October. Should we conclude that the market was very liquid during the first episode but not during the second? Not necessarily. Perhaps market participants had much stronger hedging needs after the Russian moratorium than during the events in the forex market.

A measure for liquidity that is easily available at any frequency and is directly linked to the cost of trading is the *quoted* bid-ask spread. However, since there is no commitment to market making in the OTC market for Bundesanleihen, quoted spreads are merely indicative, ie agents do not have to honour the prices they promise. While this should not matter in normal times, it is a problem in periods of stress. Traders told me that in the aftermath of the Russian moratorium, some traders simply did not bother to answer the phone. In addition, quoted spreads apply to transactions up to a specified amount only, which, furthermore, may vary over time. Eg if market participants cut this amount in periods of stress, then this would leave the quoted spread unchanged, even though market depth would fall.

One way out is to compute *effective* spreads from transactions data.[2] Unlike quoted spreads, effective spreads refer to the *average* transaction and therefore capture at least some of the cost of closing large positions. Unfortunately, this comes at a cost. Effective spreads have to be estimated, which introduces noise and reduces the frequency at which they are available. For the four German government bonds of our sample, the highest frequency which could be computed was one day.

Figure 2 shows the effective bid-ask spreads computed with Roll's (1984) measure. While this measure has important shortcomings[3], it is easy to estimate and does require a direction of trade indicator (which I have not been able to construct). I find that effective bid-ask spreads rose from values of less than one basis point during the first, tranquil, half of the year to 2-2½ basis points during the second half of August. This clearly indicates a *reduction* in market liquidity, in contrast to the *increase* suggested by the turnover series. Effective spreads gradually declined until the beginning of October, but soared to to around 4 basis points on 8 and 9 October.

How can we interpret the apparently contradictory findings of the activity series and the bid-ask spreads? First of all, the two measures should be viewed in conjunction. The increase in spreads after the Russian moratorium looks less menacing if we bear in mind that the market has been able to handle a considerably higher volume as well as larger trade sizes than under normal conditions. In early October, in contrast, there was no surge in volume which could explain at least part of the widening of the bid-ask spreads, which therefore looks much more threatening.

Let me conclude by restating the two requirements that a measure for liquidity under stress should meet. Firstly, it should reflect the cost of closing large positions within a short period of time. Secondly, it should be available at high frequencies. Neither trading volume nor the effective bid-ask spread considered completely satisfy both requirements on their own. Taken together, however, we do get a fair picture of a market that seems to have been rather effective in handling a high degree of activity in August but became rather illiquid in October.

The ascent of electronic trading should - in principle - open up new sources of high quality data, which would permit us to compute more elaborate measures of liquidity. It is crucial, though, that central banks actually get access to this data. The experience in this respect has so far not been very encouraging.

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[2] For a survey of different ways to compute effective spreads refer to Huang & Stoll (1997).

[3] In particular, the Roll-measure is biased downwards in the presence of asymmetric information or inventory effects. Since such effects are more likely to occur in times of stress when price volatility adds to inventory risk and there may be more scope for private information, the Roll measure probably underestimates any stress-related reduction of market liquidity.

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Order flow and exchange rate dynamics

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Abstract

Macroeconomic models of nominal exchange rates perform poorly. In sample, R^2 statistics as high as 10% are rare. Out of sample, these models are typically out-forecast by a naïve random walk. This paper presents a model of a new kind. Instead of relying exclusively on macroeconomic determinants, the model includes a determinant from the field of microstructure- order flow. Order flow is the proximate determinant of price in all microstructure models. This is a radically different approach to exchange rate determination. It is also strikingly successful in accounting for realised rates. Our model of daily changes in log exchange rates produces R^2 statistics above 50%. Out of sample, our model produces significantly better short-horizon forecasts than a random walk. For the DM/$ spot market as a whole, we find that $1 billion of net dollar purchases increases the DM price of a dollar by about 0.5%.

> *Omitted variables is another possible explanation for the lack of explanatory power in asset market models. However, empirical researchers have shown considerable imagination in their specification searches, so it is not easy to think of variables that have escaped consideration in an exchange rate equation.*
>
> Richard Meese (1990)

1. Motivation: microstructure meets exchange rate economics

Since the landmark papers of Meese and Rogoff (1983a, 1983b), exchange rate economics has been in crisis. It is in crisis in the sense that current macroeconomic approaches to exchange rates are empirical failures: the proportion of monthly exchange rate changes that current models can explain is essentially zero. In their survey, Frankel and Rose (1995) write "the Meese and Rogoff analysis at short horizons has never been convincingly overturned or explained. It continues to exert a pessimistic effect on the field of empirical exchange rate modelling in particular and international finance in general."[2]

Which direction to turn is not obvious. Flood and Rose (1995), for example, are "driven to the conclusion that the most critical determinants of exchange rate volatility are not macroeconomic." If determinants are not macro fundamentals like interest rates, money supplies, and trade balances, then what are they? Two alternatives have attracted attention. The first is that exchange rate determinants include extraneous variables. These extraneous variables are typically modeled as rational speculative bubbles (Blanchard 1979, Dornbusch 1982, Meese 1986, and Evans 1986, among others). Though the jury is still out, Flood and Hodrick (1990) conclude that the bubble alternative remains unconvincing. A second alternative to macro fundamentals is irrationality. For example, exchange rates may be determined in part from avoidable expectational errors (Dominguez 1986, Frankel and Froot 1987, and Hau 1998, among others). On *a priori* grounds, many economists find this second alternative unappealing. Even if one is sympathetic to the presence of irrationality, there is a wide gulf between its presence and accounting for exchange rates empirically. Until it can produce an empirical account, this too will remain an unconvincing alternative.

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[2] The relevant literature is vast. Recent surveys include Frankel and Rose (1995), Isard (1995), and Taylor (1995).

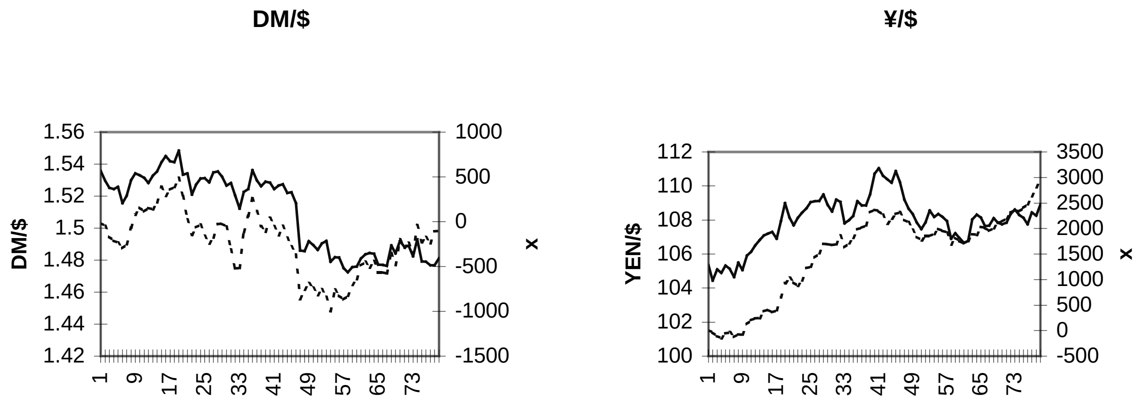
Our paper moves in a new direction: the microeconomics of asset pricing. This direction makes available a rich set of models from the field of microstructure finance. These models are largely new to exchange rate economics, and in this sense they provide a fresh approach. For example, microstructure models direct attention to new variables, variables that have “escaped the consideration” of macroeconomists (borrowing from the opening quote). The most important of these variables is order flow.[3] Order flow is the proximate determinant of price in all microstructure models. (That order flow determines price is therefore robust to differences in market structure, which makes this property more general than it might seem.) Our analysis draws heavily on this causal link from order flow to price. One level deeper, microstructure models also provide discipline for thinking about how order flow itself is determined. Information is key here - in particular, information that currency markets need to aggregate. This can include traditional macro fundamentals, but is not limited to them. In sum, our microeconomic approach provides a new type of alternative to the traditional macro approach, one that does not rely on extraneous information or irrationality.[4]

Turning to the data, we find that order flow does indeed matter for exchange-rate determination. By “matter” we mean that order flow explains most of the variation in nominal exchange rates over periods as long as four months. The graphs below provide a convenient summary of this explanatory power. The solid lines are the spot rates of the DM and Yen against the Dollar over our four-month sample (1 May to 31 August 1996). The dashed lines are marketwide order-flow for the respective currencies. Order flow, denoted by x, is the sum over time of signed trades between foreign exchange dealers worldwide.[5]

Figure 1

Four months of exchange rates (solid) and order flow (dashed)

1 May - 31 August 1996



Order flow and nominal exchange rates are strongly positively correlated (price increases with buying pressure). Macroeconomic exchange rate models, in contrast, produce virtually no correlation over periods as short as four months.

3 Order flow is a measure of buying/selling pressure. It is the net of buyer-initiated orders and seller-initiated orders. In a dealer market such as spot foreign exchange, it is the dealers who absorb this order flow, and they are compensated for doing so. (In an auction market, limit orders absorb the flow of market orders.)

4 Another alternative to traditional macro modelling is the recent “new open-economy macro” approach (eg Obstfeld and Rogoff 1995). We do not address this approach in this paragraph because, as yet, it has not produced an empirical literature.

5 For example, if a dealer initiates a trade against another dealer's DM/$ quote, and that trade is a $ purchase (sale), then order flow is +1 (–1). These are cumulated across dealers over each 24-hour trading day (weekend trading - which is minimal - is included in Monday). In spot foreign exchange, roughly 75% of total volume is between dealers (25% is between dealers and non-dealer customers).

To address this more formally, we develop and estimate a model that includes both macroeconomic determinants (eg interest rates) and a microstructure determinant (order flow). Our estimates verify the significance of the above correlation. The model accounts for about 60% of daily changes in the DM/$ exchange rate. For comparison, macro models rarely account for even 10% of monthly changes. Our daily frequency is noteworthy: though our model draws from microstructure, it is not estimated at the transaction frequency. Daily analysis is in the missing middle between past microstructure work (tick-by-tick data) and past macro work (monthly data). Bridging the two helps clarify how lower-frequency exchange rates emerge from the market's operation in real time.

To complement these in-sample results, we also examine the model's out-of-sample forecasting ability. Work by Meese and Rogoff (1983a) examines short-horizon forecasts (1 to 12 months). They find that a random walk model out-forecasts the leading macro models, even when macro-model "forecasts" are based on realised future fundamentals. Subsequent work lengthens the horizon beyond 12 months and finds that macro models begin to dominate the random walk (Meese and Rogoff 1983b, Chinn 1991, Chinn and Meese 1994, and Mark 1995). But results at shorter horizons remain a puzzle. Here we examine horizons of less than one month. (Transaction data sets that are currently available are too short to generate statistical power at monthly horizons.) We find that at horizons from one-day to two-weeks, our model produces better forecasts than the random-walk model (over 30% lower root mean squared error).

The relation we find between exchange rates and order flow is not inconsistent with the macro approach, but it does raise several concerns. Under the macro approach, order flow should not matter for exchange rate determination: macroeconomic information is publicly available - it is impounded in exchange rates without the need for order flow. More precisely, the macro approach typically assumes that: (1) all information relevant for exchange rate determination is common knowledge; and (2) the mapping from that information to equilibrium prices is also common knowledge. If either of these two assumptions is relaxed, however, then order flow will convey information about market-clearing prices. Relaxing the second assumption should not be controversial, given the failure of current exchange-rate models. Direct evidence, too, corroborates that order flow conveys relevant information (Lyons 1995, Yao 1997, Covrig and Melvin 1998, Ito, Lyons and Melvin 1998, Cheung and Wong 1998, Bjonnes and Rime 1998, Evans 1999, Naranjo and Nimalendran 1999, and Payne 1999.)[6]

Note that order flow being a proximate determinant of exchange rates does not preclude macro fundamentals from being the underlying determinant. Macro fundamentals in exchange rate equations may be so imprecisely measured that order-flow provides a better "proxy" of their variation. This interpretation of order flow as a proxy for macro fundamentals is particularly plausible with respect to expectations: standard empirical measures of expected future fundamentals are obviously imprecise.[7] Orders, on the other hand, reflect a willingness to back one's beliefs with real money (unlike survey-based measures of expectations). Measuring order flow under this interpretation is akin to counting the backed-by-money expectational votes.

This paper has six remaining sections. Section two contrasts the micro and macro approaches to exchange rates. Section three develops a model that includes both micro and macro determinants. Section four describes our data. Section five presents our results. Section six provides perspective on our results. Section seven concludes.

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[6] The standard example of order flow that conveys non-public information is orders from central bank intervention. (Within our four-month sample, however, the Fed never intervened.) Probably more important on an ongoing basis is order flow that conveys information about "portfolio shifts" that are not common knowledge. A recent event provides a sharp example. Major banks attribute the yen/dollar rate's drop from 145 to 115 in Fall 1998 to "the unwinding of positions by hedge funds that had borrowed in cheap yen to finance purchases of higher-yielding dollar assets" (The Economist, 10.10.98). This unwinding - and the selling of dollars that came with it - was forced by the scaling back of speculative leverage in the months following the Long Term Capital Management crisis. These trades were not common knowledge as they were occurring. (See also section 6 below, and Cai et al. 1999.)

[7] One might argue that expectations measurement cannot be driving the negative results of Meese and Rogoff because they use the driving variables' *realised* values. However, if the underlying macro model is incomplete, then realised values still produce an incorrect expectations measure.

2. Models: spanning the micro-macro divide

A core distinction between a microstructure approach to exchange rates and the traditional macro approach is the role of trades in price determination. In macro models, trades have no distinct role in determining price. In microstructure models, trades have a leading role - they are the proximate cause of price adjustment. It is instructive to frame this distinction by contrasting the structural models that emerge from these two approaches.

Structural models: macro approach

Exchange-rate models within the macro approach are typically estimated at the monthly frequency. When estimated in changes they take the form:

(1) $$\Delta p_t = f(\Delta i, \Delta m, \ldots) + \varepsilon_t.$$

where Δp_t is the change in the log nominal exchange rate over the month (DM/$). The driving variables in the function $f(\Delta i,\Delta m,\ldots)$ include changes in home and foreign nominal interest rates i, money supply m, and other macro determinants, denoted here by the ellipsis.[8] Changes in these public-information variables drive price - there is no role for order flow. Any incidental price effects from order flow that might arise are subsumed in the residual ε_t. These models are logically coherent and intuitively appealing. Unfortunately, they account for almost none of the monthly variation in floating exchange rates.

Structural models: microstructure approach

Equations of exchange-rate determination within the microstructure approach are derived from the optimisation problem faced by price setters in the market - the dealers.[9] These models are all variations on the following specification:

(2) $$\Delta p_t = g(\Delta x, \Delta I, \ldots) + \nu_t$$

Now Δp_t is the DM/$ rate change over two transactions, rather than over a month as in the macro models. The driving variables in the function $g(\Delta x,\Delta I,\ldots)$ include order flow Δx, the change in net dealer positions (or inventory) ΔI, and other micro determinants, denoted by the ellipsis. Order flow can take both positive and negative values because the counterparty either purchases (+) at the dealer's offer or sells at the dealer's bid (–). Here we use the convention that a positive Δx is net dollar purchases, making the theoretical relation positive: net dollar purchases drive up the DM price of dollars. It is interesting to note that the residual in this case is the mirror image of the residual in equation 1: it subsumes any price changes due to determinants in the macro model $f(\Delta i,\Delta m,\ldots)$, whereas the residual in equation 1 subsumes price changes due to determinants in the micro model $g(\Delta x,\Delta I,\ldots)$.

Microstructure models predict a positive relation between Δp and Δx because order flow communicates non-public information, and once communicated, it is reflected in price. For example, if there is an agent who has superior information about the value of an asset, and that information advantage induces the agent to trade, then a dealer can learn from those trades (purchases indicate good news about the asset's value, and vice versa). Empirically, estimates of a relation between Δp and Δx at the transaction frequency are uniformly positive and significant. This is true for many different markets, including stocks, bonds, and foreign exchange.

The relation in microstructure models between Δp and ΔI is not our focus in this paper, but let us clarify nonetheless. This relation is referred to as the inventory-control effect on price. The inventory-control effect arises when a dealer adjusts his price to control fluctuation in his inventory. For example, if a

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8 The precise list of determinants depends on the model. Meese and Rogoff (1983a) focus on three models in particular: the flexible-price monetary model, the sticky-price monetary model, and the sticky-price asset model. Here our interest is simply a broad-brush contrast between the macro and microstructure approaches. For specific models see Frenkel (1976), Dornbusch (1976), and Mussa (1976), among many others.

9 Empirical work using structural micro models includes Glosten and Harris (1988), Madhavan and Smidt (1991), and Foster and Viswanathan (1993), all of which address the NYSE. Structural models in a multiple-dealer setting include Snell and Tonks (1995) for stocks, Lyons (1995) for currencies, and Vitale (1998) for bonds.

dealer has a larger long position than is desired, he may shade his bid and offer downward to induce a customer purchase, thereby reducing his position. This affects realised transaction prices, which accounts for the relation. (These idiosyncratic inventory effects on individual dealer prices do not arise in the model developed in the next section.)

Spanning the micro-macro divide

To span the divide between the micro and macro approaches, we develop a model with components from both:[10]

$$\Delta p_t = f(\Delta i, \ldots) + g(\Delta x, \ldots) + \eta_t. \qquad (3)$$

The challenge is the frequency mismatch: transaction frequency for the micro models versus monthly frequency for the macro models. In the next section we develop a model in the spirit of equation (3). We estimate the model at the daily frequency by using micro determinants that are time-aggregated. We focus in particular on order flow Δx. Our time-aggregated measure spans a much longer period than is addressed elsewhere within empirical microstructure.

3. Portfolio shifts model

Overview

One source of exchange rate variation in the model is portfolio shifts on the part of the public. These portfolio shifts have two important features. First, they are not common knowledge as they occur. Second, they are large enough that clearing the market requires adjustment of the spot exchange rate.

The first feature - that portfolio shifts are not common knowledge - provides a role for order flow. At the beginning of each day, public portfolio shifts are manifested in orders in the foreign exchange market. These orders are not publicly observable. Dealers take the other side of these orders, and then trade among themselves during the day to share the resulting inventory risk. The market learns about the initial portfolio shifts by observing this interdealer trading activity. By the end of the day, the dealers' inventory risk is shared with the public.

The second important feature is that the initial portfolio shifts, once absorbed by the public at the end of the day, are large enough to move price. This requires that the public's demand for foreign-currency assets is less than perfectly elastic. If the public's demand is less than perfectly elastic, different-currency assets are imperfect substitutes, and price adjustment is required to clear the market.[11] In this sense, our model is in the spirit of the portfolio balance approach to exchange rates. In another sense, however, our model is very different from that earlier approach. Portfolio balance models are driven by changes in asset supply. Asset supply is constant in our model. Rather, our model identifies two distinct components on the demand side. The first is driven by innovations in public information (standard macro fundamentals). The second is driven by non-public information. This non-public information takes the form of portfolio shifts. The model does not take a stand on the underlying determinants of these portfolio shifts (though we do address this issue in section 6).

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[10] Goldberg and Tenorio (1997) develop a model for the Russian ruble market that includes both macro and microstructure components. Osler's (1998) trading model includes macroeconomic "current account traders" who affect the exchange rate in flow equilibrium.

[11] For evidence of imperfect substitutability across U.S. stocks, see Scholes (1972), Shleifer (1986) and Bagwell (1992), among others. Substitutability across currencies is likely to be lower than across same-currency stocks. Though direct evidence of this lower substitutability is lacking, some point to home bias in international portfolios as indirect evidence. Note, too, that the size of the order flows the DM/$ spot market needs to absorb are on average more than 10,000 times those absorbed in a representative U.S. stock (eg the average daily volume on NYSE stocks in 1998 was $9.3 million, whereas the average daily volume in DM/$ spot was about $300 billion).

Specifics

Consider a pure exchange economy with T trading periods and two assets, one riskless, and one with a stochastic payoff representing foreign exchange. The T+1 payoff on foreign exchange, denoted F, is composed of a series of increments, so that $F=\sum_{t=1}^{T+1} r_t$. The increments r_t are i.i.d. Normal$(0,\Sigma_r)$ and are observed before trading in each period. These realised increments represent the flow of publicly available macroeconomic information over time (eg changes in interest rates).

The foreign exchange market is organised as a decentralised dealership market with N dealers, indexed by i, and a continuum of non-dealer customers (the public), indexed by $z\in[0,1]$. Within each period (day) there are three rounds of trading. In the first round dealers trade with the public. In the second round dealers trade among themselves to share the resulting inventory risk. In the third round dealers trade again with the public to share inventory risk more broadly. The timing within each period is:

Daily timing

Round 1			Round 2			Round 3	
r_t	Dealers Quote	Public Trades	Dealers Quote	Interdealer Trade	Order Flow	Dealers Quote	Public Trades

The dealers and customers all have identical negative exponential utility defined over time-T wealth.

Trading round 1

At the beginning of each period t, all market participants observe r_t, the period's increment to the payoff F. On the basis of this increment and other available information, each dealer simultaneously and independently quotes a scalar price to his customers at which he agrees to buy and sell any amount.[12] We denote this round-one price of dealer i as P_{i1}. (To ease the notational burden, we suppress the period subscript t when clarity permits.) Each dealer then receives a net customer-order realisation c_{i1} that is executed at his quoted price P_{i1}, where $c_{i1}<0$ denotes a net customer sale (dealer i purchase). Each of these N customer-order realisations is distributed Normal$(0,\Sigma_{c1})$, and they are independent across dealers. (Think of these initial customer trades as assigned - or preferenced - to a single dealer, resulting from bilateral customer relationships for example.) Customer orders are also distributed independently of the public-information increment r_t.[13] These orders represent portfolio shifts on the part of the non-dealer public. Their realisations are not publicly observable.

Trading round 2

Round 2 is the interdealer trading round. Each dealer simultaneously and independently quotes a scalar price to other dealers at which he agrees to buy and sell any amount. These interdealer quotes are observable and available to all dealers in the market. Each dealer then simultaneously and independently trades on other dealers' quotes. Orders at a given price are split evenly across any dealers quoting that price. Let T_{i2} denote the (net) interdealer trade initiated by dealer i in round two. At the close of round 2, all dealers observe the net interdealer order flow from that period:

12 The sizes tradable at quoted prices in major FX markets are very large relative to other markets. At the time of our sample, the standard quote in DM/$ was good for up to $10 million, with a tiny bid-offer spread, typically less than four basis points. Introducing a bid-offer spread (or price schedule) in round one to endogenise the number of dealers is a straightforward - but distracting - extension of our model.

13 A natural extension of this specification is that customer orders reflect changing expectations of future r_t.

$$\Delta x = \sum_{i=1}^{N} T_{i2} \tag{4}$$

Note that interdealer order flow is observed without noise, which maximises the difference in transparency across trade types: customer-dealer trades are not publicly observed but interdealer trades are observed. In reality, FX trades between customers and dealers are not publicly observed. Though signals of interdealer order flow are publicly observed, it is not the case that these trades are observed without noise. Adding noise to Eq. (4), however, has no qualitative impact on our estimating equation, so we stick to this simpler specification.

Trading round 3

In round three, dealers share overnight risk with the non-dealer public. Unlike round one, the public's motive for trading in round three is non-stochastic and purely speculative. Initially, each dealer simultaneously and independently quotes a scalar price P_{i3} at which he agrees to buy and sell any amount. These quotes are observable and available to the public at large.

The mass of customers on the interval [0,1] is large (in a convergence sense) relative to the N dealers. This implies that the dealers' capacity for bearing overnight risk is small relative to the public's capacity. Dealers therefore set prices so that the public willingly absorbs dealer inventory imbalances, and each dealer ends the day with no net position. These round-3 prices are conditioned on the round-2 interdealer order flow. The interdealer order flow informs dealers of the size of the total inventory that the public needs to absorb to achieve stock equilibrium.

Knowing the size of the total inventory the public needs to absorb is not sufficient for determining round-3 prices. Dealers also need to know the risk-bearing capacity of the public. We assume it is less than infinite. Specifically, given negative exponential utility, the public's total demand for the risky asset in round-3, denoted c_3, is a linear function of the its expected return conditional on public information:

$$c_3 = \gamma\left(E[P_{3,t+1}|\Omega_3]-P_{3,t}\right)$$

where the positive coefficient γ captures the aggregate risk-bearing capacity of the public, and Ω_3 is the public information available at the time of trading in round three.

Equilibrium

The dealer's problem is defined over four choice variables, the three scalar quotes P_{i1}, P_{i2}, and P_{i3}, and the dealer's interdealer trade T_{i2} (the latter being a component of Δx, the interdealer order flow). The appendix provides details of the model's solution. Here we provide some intuition. Consider the three scalar quotes. No arbitrage ensures that, within a given round, all dealers quote a common price. Given that all dealers quote a common price, this price is necessarily conditioned on common information only. Though r_t is common information at the beginning of round 1, order flow Δx_t is not observed until the end of round 2. The price for round-3 trading, P_3, therefore reflects the information in both r_t and Δx_t.

Whether Δx does influence price depends on whether it communicates any price-relevant information. The answer is yes. Understanding why requires a few steps. First, the appendix shows that it is optimal for each dealer to trade in round 2 according to the trading rule:

$$T_{i2} = \alpha\, c_{i1}$$

with a constant coefficient α. Thus, each dealer's trade in round 2 is proportional to the customer order he receives in round 1. This implies that when dealers observe the interdealer order flow $\Delta x = \Sigma_i T_{i2}$ at the end of round 2, they can infer the aggregate portfolio shift on the part of the public in round 1 (the sum of the N realisations of c_{i1}). Dealers also know that the public needs to be induced to re-absorb this portfolio shift in round 3. This inducement requires a price adjustment. Hence the relation between the interdealer order flow and the subsequent price adjustment.

The pricing relation

The appendix establishes that the change in price from the end of period t-1 to the end of period t is:

(5) $$\Delta P_t = r_t + \lambda \Delta x_t$$

where λ is a positive constant. That this price change includes the innovation in payoffs r_t one-for-one is unsurprising. The $\lambda\Delta x_t$ term is the portfolio shift term. This term reflects the price adjustment required to induce re-absorption of the public's portfolio shift from round 1. For intuition, note that $\lambda\Delta x = \lambda\Sigma_i T_{i2} = \lambda\alpha\Sigma_i c_{i1}$. The sum $\Sigma_i c_{i1}$ is this total portfolio shift from round 1. The public's total demand in round 3, c_3, is not perfectly elastic, and λ insures that at the round-3 price $c_3 + \Sigma_i c_{i1} = 0$.

Empirical implementation

Getting from equation (5) to an estimable model requires that we specialise the macro component of the model—the public-information increment r_t. We choose to specialise this component to capture changes in the nominal interest differential. That is, we define $r_t \equiv \Delta(i_t - i_t^*)$, where i_t is the nominal dollar interest rate and i_t^* is the nominal non-dollar interest rate (DM or Yen). This yields the following regression model:

(6) $$\Delta P_t = \beta_1 \Delta(i_t - i_t^*) + \beta_2 \Delta x_t + \eta_t$$

Our choice of specialisation has some advantages. First, this specification is consistent with monetary macro models in the sense that these models call for estimating ΔP using the interest differential's change, not its level. (As a diagnostic, though, we also estimate the model using the level of the differential, a la Uncovered Interest Parity; see footnote 17.) Second, in asset-approach macro models like the Dornbusch (1976) overshooting model, innovations in the interest differential are the main engine of exchange rate variation.[14] Third, from a purely practical perspective, data on the interest differential are readily available at the daily frequency, which is certainly not the case for the other standard macro fundamentals (eg real output, nominal money supplies, etc).

Naturally, this specification of our macro component of the model has some drawbacks. It is certainly true that, as a measure of variation in macro fundamentals, the interest differential is obviously incomplete. One can view it as an attempt to control for this key macro determinant in order to examine the importance of micro determinants. One should not view it as establishing a fair horse race between the micro and macro approaches.

4. Data

Our data set contains time-stamped, tick-by-tick data on actual transactions for the two largest spot markets - DM/$ and ¥/$ - over a four-month period, 1 May to 31 August 1996. (For more detail than we provide here, see Evans 1997.) These data were collected from the Reuters Dealing 2000-1 system via an electronic feed customised for the purpose. Dealing 2000-1 is the most widely used electronic dealing system. According to Reuters, over 90% of the world's direct interdealer transactions take place through the system.[15] All trades on this system take the form of bilateral electronic conversations. The conversation is initiated when a dealer uses the system to call another dealer to request a quote. Users are expected to provide a fast two-way quote with a tight spread, which is in turn dealt or declined quickly (ie within seconds). To settle disputes, Reuters keeps a temporary record of all bilateral conversations. This record is the source of our data. (Reuters was unable to provide the identity of the trading partners for confidentiality reasons.)

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[14] Cheung and Chinn (1998) corroborate this empirically. Their surveys of foreign exchange traders show that the importance of individual macroeconomic variables shifts over time, but "interest rates always appear to be important."

[15] Direct trading accounts for about 60% of interdealer trade and brokered trading accounts for the remaining 40%. As noted in footnote 3, interdealer transactions account for about 75% of total trading in major spot markets (at the time of our sample). Therefore, relative to the total market, our data set represents about 90% of 60% of 75%, or about 40%. For more detail on the Reuters Dealing 2000-1 System see Lyons (1995) and Evans (1997).

For every trade executed on D2000-1, our data set includes a time-stamped record of the transaction price and a bought/sold indicator. The bought/sold indicator allows us to sign trades for measuring order flow. This is a major advantage: we do not have to use the noisy algorithms used elsewhere in the literature for signing trades. A drawback is that it is not possible to identify the size of individual transactions.[16] For model estimation, order flow Δx is therefore measured as the difference between the number of buyer-initiated trades and the number of seller-initiated trades. This shortcoming of our data - as well as others - must be kept in perspective, however. If our data were noisy and our results negative, then concern about data quality would be serious indeed - the negative results could easily be due to poor data. We show below, however, that our results are quite positive. This cannot be the result of noise in our data.

Three features of the data are especially noteworthy. First, they provide transaction information for the whole interbank market over the full 24-hour trading day. This contrasts with earlier transaction data sets covering single dealers over some fraction of the trading day (Lyons 1995, Yao 1998, and Bjonnes and Rime 1998). Our comprehensive data set makes it possible, for the first time, to analyse order flow's role in price determination at the level of "the market." Though other data sets exist that cover multiple dealers, they include only brokered interdealer transactions (see Goodhart, Ito and Payne 1996, and Payne 1999). More important, these other data sets come from a particular brokered-trading system, one that accounts for a much smaller fraction of daily trading volume than the D2000-1 system covered by our data set. (There is also evidence that dealers attach more informational importance to direct interdealer order flow than to brokered interdealer order flow. See Bjonnes and Rime 1998.)

Second, our market-wide transactions data are not observed by individual FX dealers as they trade. Though dealers have access to their own transaction records, they cannot observe others' transactions on the system. Our data therefore represent activity that, at the time, participants could only infer indirectly. This is one of those rare situations where the researcher has more information than market participants themselves (at least in this dimension).

Third, our data cover a relatively long time span (four months) in comparison with other micro data sets. This is important because the longer time span allows us to address exchange-rate determination from more of an asset-pricing perspective than was possible with previous micro data spanning only days or weeks.

The three variables in our Portfolio Shifts model are measured as follows. The change in the spot rate (DM/$ or ¥/$), Δp_t, is the log change in the purchase transaction price between 4 pm (GMT) on day t and 4 pm on day t-1. When a purchase transaction does not occur precisely at 4 pm, we use the subsequent purchase transaction (with roughly one million trades per day, the subsequent transaction is generally within a few seconds of 4 pm). When day t is a Monday, the day t-1 price is the previous Friday's price. (Our dependent variable therefore spans the full four months of our sample, with no overnight or weekend breaks.) The daily order flow, Δx_t, is the difference between the number of buyer-initiated trades and the number of seller-initiated trades (in thousands), also measured from 4 pm (GMT) on day t-1 to 4 pm on day t (negative sign denotes net dollar sales). The change in interest differential, $\Delta(i_t – i_t^*)$, is calculated from the daily overnight interest rates for the dollar, the deutschemark, and the yen (annual basis); the source is Datastream (typically measured at approximately 4 pm GMT).

5. Empirical results

Our empirical results are grouped in four sets. The first set addresses the in-sample fit of the portfolio shifts model. The second set addresses robustness issues. The third set addresses the direction of causality. The fourth set of results addresses the model's out-of-sample forecasting ability (in the spirit of Meese and Rogoff 1983a).

[16] This drawback may not be acute. There is evidence that the size of trades has no information content beyond that contained in the number of transactions. See Jones, Kaul, and Lipson (1994).

5.1 In-sample fit

Table 1 presents our estimates of the portfolio shifts model (equation 6) using daily data for the DM/$ and ¥/$ exchange rates. Specifically, we estimate the following regression:

(7) $\Delta p_t = \beta_1 \Delta(i_t - i_t^*) + \beta_2 \Delta x_t + \eta_t$

where Δp_t is the change in the log spot rate (DM/$ or ¥/$) from the end of day t-1 to the end of day t, $\Delta(i_t - i_t^*)$ is the change in the overnight interest differential from day t-1 to day t (* denotes DM or ¥), and Δx_t is the order flow from the end of day t-1 to the end of day t (negative denotes net dollar sales).[17]

The coefficient β_2 on our portfolio shift variable Δx_t is correctly signed and significant, with t-statistics above 5 in both equations. To see that the sign is correct, recall from the model that net purchases of dollars - a positive Δx_t - should lead to a higher DM price of dollars. The traditional macro-fundamental - the interest differential - is correctly signed, but is only significant in the yen equation. (The sign should be positive because, in the sticky-price monetary model for example, an increase in the dollar interest rate i_t requires an immediate dollar appreciation - increase in DM/$ - to make room for the expected dollar depreciation required by uncovered interest parity.) The overall fit of the model is striking relative to traditional macro models, with R^2 statistics of 64% and 45% for the DM and yen equations, respectively. In terms of diagnostics, the DM equation shows some evidence of heteroskedasticity, so we correct the standard errors in that equation using a heteroskedasticity-consistent covariance matrix (White correction).[18]

Table 1

In-sample fit of portfolio shifts model

$\Delta p_t = \beta_1 \Delta(i_t - i_t^*) + \beta_2 \Delta x_t + \eta_t$

	$\Delta(i_t - i_t^*)$	Δx_t	Diagnostics		
			R^2	Serial	Hetero
DM	0.52 (0.35)	2.10 (0.20)	0.64	0.78 0.41	0.08 0.02
Yen	2.48 (0.92)	2.90 (0.46)	0.45	0.50 0.37	0.96 0.71

The dependent variable Δp_t is the change in the log spot exchange rate from 4 pm GMT on day t-1 to 4 pm GMT on day t (DM/$ or ¥/$). The regressor $\Delta(i_t - i_t^*)$ is the change in the one-day interest differential from day t-1 to day t (* denotes DM or ¥, annual basis). The regressor Δx_t is interdealer order flow between 4 pm GMT on day t-1 and 4 pm GMT on day t (negative for net dollar sales, in thousands). Estimated using OLS. Standard errors are shown in parentheses (corrected for heteroskedasticity in the case of the DM). The sample spans four months (May 1 to August 31, 1996), which is 89 trading days. The Serial column presents the p-value of a chi-squared test for residual serial correlation, first-order in the top row and fifth-order (one week) in the bottom row. The Hetero column presents the p-value of a chi-squared test for ARCH in the residuals, first-order in the top row and fifth-order in the bottom row.

[17] Though the dependent variable in standard macro models is the change in the log spot rate, the dependent variable in the Portfolio Shifts model in equation (6) is the change in the spot rate without taking logs. These two measures for the dependent variable produce nearly identical results in all our tables (R^2s, coefficient significance, lack of autocorrelation, etc). Here we present the log change results - equation 7 - to make them directly comparable to previous macro specifications.

[18] To check robustness, we examine several obvious variations on the model. For example, in the spirit of Uncovered Interest Parity, we include the level of the interest differential in lieu of its change. The level of the differential is insignificant in both cases. We also include a constant in the regression, even though the model does not call for one. The constant is insignificant for both currencies. Estimating the whole model in levels rather than changes produces a pattern similar to that in Table 1: order flow is highly significant, the interest differential is insignificant, and R^2 is 0.75 for the DM equation and 0.61 for the Yen equation. With this levels regressions, however, beyond the usual concerns about non-stationarity, there is also strong evidence of serial correlation and heteroskedasticity (both tests are significant at the 1% level for both currencies). Finally, recall that our price series is measured from purchase transactions. Results using 4 pm sale prices are identical. We address additional robustness issues in the next subsection.

The size of our order flow coefficient is consistent with past estimates based on single-dealer data. The coefficient of 2.1 in the DM equation implies that a day with 1000 more dollar purchases than sales induces an increase in the DM price by 2.1%. Given the average trade size in our sample of $3.9 million, $1 billion of net dollar purchases increases the DM price of a dollar by 0.54% (= 2.1/3.9). At a spot rate of 1.5 DM/$, this implies that *$1 billion of net dollar purchases increases the DM price of a dollar by 0.8 pfennig*. At the single-dealer level, Lyons (1995) finds that information asymmetry induces the dealer he tracks to increase price by 1/100th of a pfennig (0.0001 DM) for every incoming buy order of $10 million. That translates to 1 pfennig per $1 billion. Though linearly extrapolating this estimate is certainly not an accurate description of single-dealer behaviour, with multiple dealers it may be a good description of the market's aggregate elasticity.

The striking explanatory power of these regressions is almost wholly due to order flow Δx_t. Regressing Δp_t on $\Delta(i_t – i_t^*)$ alone, plus a constant, produces an R^2 statistic less than 1% in both equations, and coefficients on $\Delta(i_t – i_t^*)$ that are insignificant at the 5% level.[19] That the interest differential regains significance once order flow is included, at least in the Yen equation, is consistent with omitted variable bias in the interest-rates-only specification. (The correlation between the two regressors Δx_t and $\Delta(i_t – i_t^*)$ is 0.02 for the DM and –0.27 for the Yen, though both are insignificant at the 5% level.)

Order flow's ability to account for the full four months of exchange rate variation is surprising, not only from the perspective of macro exchange rate economics, but also from the perspective of microstructure finance. Recall from section 2 that structural models within microstructure finance are typically estimated at the transaction frequency - they make no attempt to account for prices over the full 24-hour day. Our regression is at the daily frequency. One might have conjectured that the net impact of order flow over the day would be zero (each day accounts for about one million transactions). This conjecture would be consistent with a belief that cumulative order flow mean-reverts rapidly (eg within a day). But rapid mean reversion is clearly not the behaviour displayed by cumulative order flow in Figure 1. This lack of mean reversion provides some room for the lower frequency relation we find here.

The lack of strong mean reversion in our measured order flow deserves further attention, particularly considering that half-lives of individual dealer positions can be as short as 10 minutes (Lyons 1998). The key lies in recognising that our measure of order flow reflects interdealer trading, not customer-dealer trading. Consider a scenario that illustrates why our measure in Figure 1 can be so persistent. (Recall that Figure 1 displays cumulative order flow, defined as the sum of interdealer order flow, Δx_t, from 0 to t.) Starting the scenario from $x_t=0$, an initial customer sale does not move x_t from zero because x_t measures interdealer order flow only. After the customer sale, then when dealer i unloads the position by selling to another dealer j, x_t drops to –1. A subsequent sale by dealer j to another dealer, dealer k, reduces x_t further to –2.[20] If a customer happens to buy dealer k's position from him, then x_t remains at –2. In this simple scenario, order flow measured only from trades between customers and dealers would have reverted to zero - the concluding customer trade offsets the initiating customer trade. The interdealer order flow, however, does not revert to zero. Note, too, that this difference in the persistence of the two order-flow measures - customer-dealer versus interdealer - is also a property of the Portfolio Shifts model. In the Portfolio Shifts model, customer order flow in round three always offsets the customer order flow in round one. But the interdealer order flow, which only arises in round two, does not net to zero. This non-zero Δx_t serves as a carrier of value in our estimating equation.

5.2 Robustness

In this section we address three robustness issues beyond those examined in the previous section. They correspond to the following three questions: (1) Might the order-flow/price relation be non-linear?

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[19] There is a vast empirical literature that attempts to increase the explanatory power of interest rates in exchange rate equations by introducing interest rates as separate regressors, introducing non-linear specifications, etc. This literature has not been successful, so we do not pursue this line here. Note that the lack of explanatory power from traditional fundamentals is not unique to exchange rate economics: Roll (1988) produces R^2s of only 20% using traditional equity fundamentals to account for daily stock returns, a result he describes as a "significant challenge to our science".

[20] This repeated passing of dealer positions in the foreign exchange market is referred to as the "hot potato" phenomenon. See Burnham (1991) and Lyons (1997).

(2) Does the relation depend on the gross level of activity? and (3) Does the relation depend on day of the week?

Might the order-flow/price relation be non-linear?

The linearity of our Portfolio Shifts specification depends crucially on several simplifying assumptions, some of which are rather strong on empirical grounds. It is therefore natural to investigate whether non-linearities or asymmetries might be present.[21] A simple first test is to add a squared order-flow term to the baseline specification. The squared order-flow term is insignificant in both equations. We also test whether the coefficient on order flow is piece-wise linear, with a kink at $\Delta x_t=0$. If true, this means that buying pressure and selling pressure are not symmetric. A Wald test that the two slope coefficients are equal cannot be rejected for the DM equation. There is some evidence of different slopes in the Yen equation however: the test is rejected at the 4% marginal significance level. In that case, the point estimates show a greater sensitivity of price to order flow in the downward direction, though both estimates remain positive and significant.

Does the order-flow/price relation depend on the gross level of activity?

Another natural concern is whether the order-flow/price relation in Table 1 is state contingent in some way, perhaps depending on the market's overall activity level. Our data set provides a convenient measure of overall activity, namely the total number of transactions. As a simple test, we partition our sample of trading days into quartiles, from days with the fewest transactions to days with the most transactions. We then estimate separate order-flow coefficients for each of these four sample partitions. In both the DM and Yen equations, all four of the order-flow coefficients are positive. In the DM equation, the coefficients are slightly U-shaped (from fewest transactions to most, the point estimates for β_2 are 2.7, 2.0, 1.9, and 3.3). In the Yen equation, the coefficients are monotonically increasing (from fewest transactions to most, the point estimates for β_2 are 1.0, 1.1, 3.5, and 4.1).

In terms of theory, this result for the Yen is consistent with the "event-uncertainty" model of Easley and O'Hara (1992), but the DM result is not. The event-uncertainty model predicts that trades are more informative when trading intensity is higher. Key to understanding their result is that in their model, new information may not exist. If there is trading at time t, then a rational dealer raises her conditional probability that an information event has occurred, and lowers the probability of the "no-information" event. The upshot is that trades occurring when trading intensity is high induce a larger update in beliefs, and therefore a larger adjustment in price.

Does the order-flow/price relation depend on day of the week?

Another state-contingency that warrants attention is day-of-the-week effects.[22] To test whether day-of-the-week matters, we partition our sample into five sub-samples, one for each weekday (recall that weekends are subsumed in our Friday-to-Monday observations). In both the DM and Yen equations, all five of the resulting order-flow coefficients are positive. In the DM equation, the Tuesday coefficient is the largest, and the Wednesday coefficient is the smallest. The Yen equation also shows that Tuesday's coefficient is the largest, but in this case the Monday coefficient is smallest. More important, a Wald test that the coefficients are equal across the five days cannot be rejected at the 5% level in either equation (though in the case of the Yen, it can be rejected at the 10% level).

5.3 Causality

Under our model's null hypothesis, causality runs strictly from order flow to price. Accordingly, under the null, our estimation is not subject to simultaneity bias. (We are not simply "regressing price on quantity," as in the classic supply-demand identification problem. Quantity - ie volume - and order flow are fundamentally different concepts.) Within microstructure theory more broadly, this direction of

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[21] We pursue these (simple) non-linear specifications with the comfort that outliers are not driving our results - a fact that is manifest from Figure 1.

[22] In terms of theory, the model of Foster and Viswanathan (1990) is a workhorse for specifying day-of-the-week effects. In their model, there is periodic variation in the information advantage of the informed trader. This advantage is assumed to grow over periods of market closure, in particular, over weekends, making order flow on Monday particularly potent.

causality is the norm: it holds in all the canonical models (Glosten and Milgrom 1985, Kyle 1985, Stoll 1978, Amihud and Mendelson 1980), despite the fact that price and order flow are determined simultaneously. The important point in these models is that price innovations are a function of order flow innovations, not the other way around.[23] That said, alternative hypotheses do exist under which causality is reversed. The following taxonomy frames the causality issue, and identifies specific alternatives under which causality is reversed, so that the merits of these alternatives can be judged in a disciplined way.

Theoretical overview

The timing of the order-flow/price relation admits three possibilities, depending on whether order flow precedes, is concurrent with, or lags price adjustment. We shall refer to these three timing hypotheses as the *Anticipation* hypothesis, the *Pressure* hypothesis, and the *Feedback* hypothesis, respectively.

Within each of the three hypotheses - Anticipation, Pressure, and Feedback - there are also variations. Under the Anticipation hypothesis, for example, order flow can precede price adjustment because prices adjust fully only after order flow is commonly observed - in low-transparency markets like foreign exchange, order flow is not commonly observed when it occurs (Lyons 1996). Order flow might also precede price because price adjusts only after some piece of news anticipated by order flow is commonly observed (eg the short-lived private information in Foster and Viswanathan 1990). Under the Pressure hypothesis the two main variations correspond to microstructure theory's canonical model types - information models and inventory models. In information models, observing order flow provides information about payoffs (Glosten and Milgrom 1985, Kyle 1985). In inventory models, order flow alters equilibrium risk premia (Stoll 1978, Ho and Stoll 1981).[24] Under the Feedback hypothesis, order flow lags price because of feedback trading. Negative-feedback trading is systematic selling in response to price increases, and buying in response to price decreases (eg Friedman's celebrated "stabilising speculators"). Positive-feedback trading is the reverse. Variations on the Feedback hypothesis are distinguished by whether this feedback trading is rational (an optimal response to return autocorrelation) or behavioural, meaning that it arises from systematic decision bias (DeLong et al. 1990, Jegadeesh and Titman 1993, Grinblatt et al. 1995).

Under the Pressure hypothesis, causality runs from order flow to price, despite their concurrent realisation.[25] For the Anticipation hypothesis, the second variation noted above - where price adjusts only after some piece of news anticipated by order flow is observed - is probably not relevant to foreign exchange (in contrast to equity markets, where insider order flow can anticipate a firm's earnings announcement, for example). The other variation of the Anticipation hypothesis - where order flow affects price with a delay because it is not commonly observed - is relevant to foreign exchange. In this case, causality still runs from order flow to price, but the effects are delayed. As noted in the Data section, order flow in this market is not common knowledge when realised. Consequently, lags in price adjustment do not violate market efficiency (conditional on public information). One way to test this variation of the Anticipation hypothesis is by introducing lagged order flow to our Portfolio Shifts model. Rows one and three of Table 2 present the results of this regression: lagged order flow is insignificant. At the daily frequency, lagged order flow is already embedded in price.[26]

[23] Put differently, order flow in these models is a proximate cause. The underlying driver of order flow is non-public information (information about uncertain demands, information about payoffs, etc). Order flow is the channel through which this type of information is impounded in price.

[24] Within this inventory-model category, there is an additional distinction between price effects that arise at the marketmaker level (canonical inventory models) and price effects that arise at the marketwide level, due to imperfect substitutability (eg our Portfolio Shifts model). In the case of price effects at the marketmaker level, these effects are often modeled as changing risk premia. But sometimes, largely for technical convenience, models are specified with risk-neutral marketmakers who face some generic "inventory holding cost."

[25] This does not imply that price cannot influence order flow. Price does influence order flow in microstructure models (both for the usual downward sloping demand reason, and because agents learn from price). It is still the case that - in equilibrium - price innovations are functions of order flow innovations, not vice versa. Our Portfolio Shifts model is a case in point.

[26] As another check along these lines, we also decompose contemporaneous order flow into expected and unexpected components (by projecting it on past order flow). In our model, all order flow Δx is unexpected, but this need not be the case in the data. We find, as the model predicts, that order flow's explanatory power comes from its unexpected component.

Under the Feedback hypothesis, causality can go in reverse, that is, from price to order flow.[27] Within exchange-rate economics, a natural first association is Friedman's stabilising speculators, which is negative-feedback trading (rational). Though the direction of causality in this case is reversed, one would expect to find an order-flow/price relation that is negative. We find a positive relation. If instead positive-feedback trading were present and significant, then one would expect order flow in period t to be positively related to the price change in period t-1. In daily data, this corresponds to Δx_t being explained, at least in part, by Δp_{t-1}. If our order-flow coefficient in Table 1 is picking up this daily-frequency positive feedback, then including lagged price change Δp_{t-1} in the Portfolio-Shifts regression should weaken, if not eliminate, the significance of order flow. Rows two and four of Table 2 present the results of this regression. Past price change does not reduce the significance of order flow, and is itself insignificant. These results run counter to the positive-feedback hypothesis at the daily frequency.

Table 2

Portfolio shifts model: Alternative specifications

$$\Delta p_t = \beta_1 \Delta(i_t - i_t^*) + \beta_2 \Delta x_t + \beta_3 \Delta x_{t-1} + \eta_t$$
$$\Delta p_t = \beta_1 \Delta(i_t - i_t^*) + \beta_2 \Delta x_t + \beta_3 \Delta p_{t-1} + \eta_t$$

	$\Delta(i_t - i_t^*)$	Δx_t	Δx_{t-1}	Δp_{t-1}	Diagnostics		
					R^2	Serial	Hetero
DM	0.40 (0.36)	2.16 (0.18)	0.29 (0.19)		0.65	0.76 0.48	0.39 0.03
	0.42 (0.35)	2.17 (0.18)		0.11 0.07	0.66	0.60 0.44	0.38 0.01
Yen	2.48 (0.91)	2.90 (0.36)	-0.20 (0.35)		0.47	0.07 0.41	0.55 0.84
	2.64 (0.91)	2.98 (0.36)		-0.13 (0.09)	0.48	0.21 0.63	0.52 0.81

The dependent variable Δp_t is the change in the log spot exchange rate from 4 pm GMT on day t-1 to 4 pm GMT on day t (DM/$ or ¥/$). The regressor $\Delta(i_t - i_t^*)$ is the change in the one-day interest differential from day t-1 to day t (* denotes DM or ¥, annual basis). The regressor Δx_t is interdealer order flow between 4 pm GMT on day t-1 and 4 pm GMT on day t (negative for net dollar sales, in thousands). Estimated using OLS. Standard errors are shown in parentheses (corrected for heteroskedasticity in the case of the DM). The sample spans four months (1 May to 31 August 1996), which is 89 trading days. The Serial column presents the p-value of a chi-squared test for residual serial correlation, first-order in the top row and fifth-order (one week) in the bottom row. The Hetero column presents the p-value of a chi-squared test for ARCH in the residuals, first-order in the top row and fifth-order in the bottom row.

Empirical reality

The theoretical overview above cannot resolve the fact that, in daily data, all three hypotheses - Anticipation, Pressure, and Feedback - may produce a relationship that appears contemporaneous. A concern therefore remains that the positive coefficient on order flow in Table 1 might be the result of positive-feedback trading that occurs intraday. We offer two additional types of evidence against this alternative interpretation of our results. The first is a set of three arguments why intraday positive feedback is an unappealing hypothesis in this context. The second is an explicit analysis of bias, designed to calibrate how extreme the positive feedback would have to be to account for the key moments of our data. (These moments include, but are not limited to, the moments that produce our order-flow coefficient in Table 1.)

There are three reasons, a priori, why the hypothesis of intraday positive-feedback trading is unappealing. First, direct empirical evidence does not support it: there is no evidence in the current literature of positive-feedback trading in the foreign exchange market. Second, if systematic positive-feedback trading were present, it would be irrational: intraday studies using transactions data find no evidence of the positive autocorrelation in price that would make positive-feedback an optimal

[27] Note that the Feedback hypothesis does not imply that causality runs wholly in reverse. For example, the Feedback hypothesis does not rule out that feedback trading can affect prices.

response (Goodhart, Ito, and Payne 1996). Third, the fallback possibility of irrational positive-feedback trading is difficult to defend. Recall that the order flow we measure is interdealer order flow. Though systematic feedback trading of a behavioural nature (ie not fully rational) might be a good description of some market participants, dealers are among the most sophisticated participants in this market.

Bias analysis

To close this section on causality, let us consider what it would take for positive-feedback trading to account for our results. Specifically, suppose intraday positive-feedback trading is present - Under what conditions could it account for the key moments of our data? These moments include, but are not limited to, the moments that produce our positive order-flow coefficient in Table 1. We show below that these conditions are rather extreme. In fact, through a broad range of underlying parameter values, feedback trading would have to be *negative* to account for the key moments of our data.

We start by decomposing measured order flow Δx_t into two components:

(8) $$\Delta x_t = \Delta x_{t1} + \Delta x_{t2}$$

where Δx_{t1} denotes exogenous order flow from portfolio shifts (a la our model), with variance equal to Σ_{x1}, and Δx_{t2} denotes order flow due to feedback trading, where

(9) $$\Delta x_{t2} = \gamma \Delta p_t$$

Suppose the true structural model can be written as:

(10) $$\Delta p_t = \alpha \Delta x_{t1} + \varepsilon_t$$

where ε_t represents common-knowledge (CK) news, and ε_t is iid with variance Σ_ε. By CK news we mean that both the information *and* its implication for equilibrium price is common knowledge. If both conditions are not met, then order flow will convey information about market-clearing prices (recall the discussion in the introduction). If feedback trading is present ($\gamma \neq 0$), then α will be a reduced form coefficient that depends on γ. Note that under these circumstances, equation (10) is a valid reduced-from equation that could be estimated by OLS if one had data on Δx_{t1}.

With data on Δx_t and Δp_t only, suppose we estimate

(11) $$\Delta p_t = \beta \Delta x_t + \varepsilon_t$$

If $\gamma \neq 0$, our estimates of β will suffer from simultaneity bias. To evaluate the size of this bias, consider the implications of equations (8) through (10) for the moments:

$$\beta = \mathrm{Cov}(\Delta p_t, \Delta x_t) \,/\, \mathrm{Var}(\Delta x_t)$$

$$\delta = \mathrm{Var}(\Delta p_t) \,/\, \mathrm{Var}(\Delta x_t)$$

From equations (8) through (10) we know that:

$$\Delta x_t = (1+\gamma\alpha)(\Delta x_{t1}) + \gamma \varepsilon_t$$

Solving for expressions for $\mathrm{Cov}(\Delta p_t, \Delta x_t)$, $\mathrm{Var}(\Delta p_t)$, and $\mathrm{Var}(\Delta x_t)$, we can write:

$$\beta = \mathrm{Cov}(\Delta p_t, \Delta x_t) \,/\, \mathrm{Var}(\Delta x_t) = \left(\alpha(1+\gamma\alpha)\Sigma_{x1} + \gamma\Sigma_\varepsilon\right) / \left((1+\gamma\alpha)^2\Sigma_{x1} + \gamma^2\Sigma_\varepsilon\right)$$

$$\delta = \mathrm{Var}(\Delta p_t) \,/\, \mathrm{Var}(\Delta x_t) = \left(\alpha^2\Sigma_{x1} + \Sigma_\varepsilon\right) / \left((1+\gamma\alpha)^2\Sigma_{x1} + \gamma^2\Sigma_\varepsilon\right)$$

Now, define an additional parameter:

$$\phi = \Sigma_\varepsilon / \Sigma_{x1}$$

This parameter represents the ratio of CK news to order-flow news. With this parameter ϕ we can rewrite the key coefficients as:

$$\beta = \left(\alpha(1+\gamma\alpha) + \gamma\phi\right) / \left((1+\gamma\alpha)^2 + \gamma^2\phi\right)$$

$$\delta = \left(\alpha^2 + \phi\right) / \left((1+\gamma\alpha)^2 + \gamma^2\phi\right)$$

Using the sample moments for Cov($\Delta p_t, \Delta x_t$), Var(Δp_t), and Var(Δx_t), we can solve for the implied values of the α and γ for given values of ϕ. The following table presents these implied values of α and γ.

Note that even for values of ϕ above 2, the feedback trading needed to generate our results is actually *negative*. Note too that the parameter α - the order-flow-causes-price parameter - is not driven to zero until ϕ reaches values well above 10. To invalidate our causality interpretation, then, CK news would have to be one to two orders of magnitude more important that order-flow news. In our judgement this is too extreme to be compelling.

To close this section on causality, it is not enough for the sceptical reader to assert simply that order flow and price are both "endogenous," or that we are merely observing a "simultaneous relationship". These points are true. But they are also true within the body of microstructure theory reviewed above. And within that body of theory, price innovations are still driven by order flow innovations. This section is our effort to bring some disciplined thinking to an otherwise superficial debate.

Table 3

Bias analysis

$$\Delta x_{t2} = \gamma \Delta p_t$$
$$\Delta p_t = \alpha \Delta x_{t1} + \varepsilon_t$$

	$\phi=\Sigma_\varepsilon/\Sigma_{x1}$	α	γ
DM	0	2.4	–0.05
	0.1	1.2	–0.51
	1	1.9	–0.12
	2	2.1	–0.03
	10	2.0	0.16
	100	0.0	0.36
Yen	0	2.4	–0.18
	0.1	1.3	–0.58
	1	2.2	–0.23
	2	2.4	–0.15
	10	2.8	–0.02
	100	0.0	0.21

The table shows the values for the parameters α (order-flow-causes-price) and γ (price-causes-order-flow) implied by the sample moments and given values for the parameter ϕ. The parameter ϕ is the ratio of common-knowledge news to order-flow news.

5.4 Out-of-sample forecasts

To control for the myriad specification searches conducted by empiricists, a tradition within exchange rate economics has been to augment in-sample model estimates with estimates of models' out-of-sample forecasting ability. Accordingly, we present results along these lines as well. The original work by Meese and Rogoff (1983a) examines forecasts from 1 to 12 months. Our four-month sample does not provide sufficient power to forecast at these horizons. Our horizons range instead from one day to two weeks. The Meese-Rogoff puzzle is why short-horizon forecasts do so poorly, and our focus is definitely on the short end (though not so short as to render the horizon irrelevant from a macro perspective).

Table four shows that the portfolio shifts model produces better forecasts than the random-walk (RW) model. The forecasts from our model are derived from recursive estimates that begin with the first 39 days of the sample. Like the Meese-Rogoff forecasts, our forecasts are based on realised values of the future forcing variables - in our case, realised values of order flow and changes in the interest differential. (Thus, they are not truly "out-of-sample forecasts". We chose to stick with the Meese-Rogoff terminology.) The resulting root mean squared error (RMSE) is 30 to 40% lower than that for the random walk.

Note that our 89-day sample has very low power at the one- and two-week horizons. Even though our model's RMSE estimates are roughly 35% lower at these horizons, their out-performance is not statistically significant. With a sample this size, the one-week forecast would need to cut the RW model's RMSE by about 50% to reach the 5% significance level. (To see this, note that for the DM a two-standard-error difference at the one-week horizon is about 0.49, which is roughly half of the RW model's RMSE of 0.98). The two-week forecast would have to cut the RW model's forecast error by some 54%. More powerful tests at these longer horizons will have to wait for longer spans of transaction data.

Table 4

Out-of-sample forecasts errors

Root mean squared errors (×100)

	Horizon	RW	Portfolio shifts	Difference
DM	1 day	0.44	0.29	0.15 (0.033)
	1 week	0.98	0.63	0.35 (0.245)
	2 weeks	1.56	0.96	0.60 (0.419)
Yen	1 day	0.40	0.32	0.08 (0.040)
	1 week	0.98	0.64	0.33 (0.239)
	2 weeks	1.34	0.90	0.45 (0.389)

The RW column reports the RMSE for the random walk model (approximately in %age terms). The Portfolio Shifts column reports the RMSE for the model in equation (6). The Portfolio Shifts forecasts are based on realised values of the forcing variables. The forecasts are derived from recursive model estimates starting with the first 39 days of the sample. The Difference column reports the difference in the two RMSE estimates, and, in parentheses, the standard errors for the difference, calculated as in Meese and Rogoff (1988).

6. Discussion

The relation in our model between exchange rates and order flow is not easy to reconcile with the traditional macro approach. Under the traditional approach, information is common knowledge and is therefore impounded in exchange rates without the need for order flow. This apparent contradiction can be resolved if either: (1) some information relevant for exchange rate determination is not common knowledge; or (2) some aspect of the mapping from information to equilibrium prices is not common knowledge. If either is relaxed then order flow conveys information about market-clearing prices.

Our portfolio shifts model resolves the contradiction by introducing information that is not common knowledge - information about shifts in public demand for foreign-currency assets. At a microeconomic level, dealers learn about these shifts in real time by observing order flow. As the dealers learn, they quote prices that reflect this information. At a macroeconomic level, these shifts are difficult to observe empirically. Indeed, the concept of order flow is not recognised within the international macro literature. (Transactions, if they occur at all, are strictly symmetric, and therefore cannot be signed to reflect net buying/selling pressure.)

If order flow drives exchange rates, then what drives order flow? From a valuation perspective, there are two distinct views. The first view is that order flow reflects new information about *valuation numerators* (ie future dividends in a dividend-discount model, which in foreign exchange take the form of future interest differentials). The second view is that order flow reflects new information about *valuation denominators* (ie anything that affects discount rates). Our portfolio shifts model is an example of the latter: order flow is unrelated to valuation numerators - the future r_t. This type of order flow can be rationalised with, for example, time-varying risk tolerance, time-varying hedging demands, or time-vary transactions demands. (In presenting the model, we did not take a stand on a specific rationalisation.) An example consistent with the valuation-numerators view is the "proxy-for-expectations" idea introduced in the introduction. That is, an important source of innovations in exchange rates is innovations in expected future fundamentals, and in real time these may be well proxied by order flow.

Note that separating valuation numerators from valuation denominators has implications for the concept of "fundamentals." Order flow that reflects information about valuation numerators - like expectations of future interest rates - is in keeping with traditional definitions of exchange-rate fundamentals. But order flow that reflects valuation denominators encompasses nontraditional exchange-rate determinants, calling, perhaps for a broader definition. In any event, exploring these links to deeper determinants is a natural topic for future research. This will surely require a retreat back into intraday data.[28]

The practitioner view versus the academic view

Another perspective on order flow emerges from the difference between academic and practitioner views on price determination. Practitioners often explain price increases with the familiar reasoning that "there were more buyers than sellers." To most economists, this reasoning is tantamount to "price had to rise to balance demand and supply." But these phrases may not be equivalent. For economists, the phrase "price had to rise to balance demand and supply" calls to mind the Walrasian auctioneer. The Walrasian auctioneer collects "preliminary" orders and uses them to find the market-clearing price. Importantly, the auctioneer's price adjustment is immediate - no trading occurs in the transition. (In a rational-expectations model of trading, for example, this is manifested in all orders being conditioned on the market-clearing price.)

Many practitioners have a different model in mind. In the practitioner model there is a dealer instead of an abstract auctioneer. The dealer acts as a buffer between buyers and sellers. The orders the dealer collects are actual orders, rather than preliminary orders, so trading does occur in the transition to the new price. The dealer determines new prices from the new information about demand and supply that becomes available.

Can the practitioner model be rationalised? At first blush, it appears that trades are taking place out of equilibrium, implying irrational behaviour. But this misses an important piece of the puzzle. Whether these trades are out-of-equilibrium depends on the information available to the dealer. If the dealer knows at the outset that there are more buyers than sellers (eventually pushing price up), then it may not be optimal to sell at a low interim price. If the buyer/seller imbalance is not known, however, then rational trades can occur through the transition. In this case, the dealer cannot set price conditional on all the information available to the Walrasian auctioneer. This is precisely the story developed in canonical microstructure models (Glosten and Milgrom 1985). Trading that would be irrational if the dealer could condition on the auctioneer's information can be rationalised in models with more limited (and realistic) conditioning information.

Relation between our model and the flow approach to exchange rates

Consider the relation between our model, with its emphasis on order flow, and the traditional "flow approach" to exchange rates. Is our approach just a return to the earlier flow approach? Despite their apparent similarity, the two approaches are distinct and, in fact, fundamentally different.

A key feature of our model is that order flow plays two roles. First, holding beliefs constant, order flow affects price through the traditional process of market clearing. Second, order flow also alters beliefs because it conveys information that is not yet common knowledge. That is:

$$\text{Price} = P\big(\Delta x, B(\Delta x, \ldots), \ \ldots\big)$$

Price P thus depends both directly and indirectly on order flow, Δx, where the indirect effect is via beliefs B. Early attempts to analyse equilibrium with differentially informed individuals ignored the information role - the effect of order flow on beliefs. Since the advent of rational expectations, models that ignore this information effect from order flow are viewed as less compelling.

[28] The role of macro announcements in determining order flow warrants exploring. This, too, requires the use of intraday data. A second possible use of macro announcements is to introduce them directly into our Portfolio Shifts specification, even at the daily frequency. This tack is not likely to be fruitful: there is a long literature showing that macro announcements are unable to account for exchange rate first moments (as opposed to second moments; see Andersen and Bollerslev 1998).

This is the essential difference between the flow approach to exchange rates and the microstructure approach. Under the flow approach, order flow communicates no information back to individuals regarding others' views/information. All information is common knowledge, so there is no information that needs aggregating. Under the microstructure approach, order flow does communicate information that is not common knowledge. This information needs to be aggregated by the market, and microstructure theory describes how that aggregation is achieved, depending on the underlying information type.

7. Conclusion

This paper presents a model of exchange rate determination of a new kind. Instead of relying exclusively on macroeconomic determinants, we draw on determinants from the field of microstructure. In particular, we focus on order flow, the variable within microstructure that is - both theoretically and empirically - the driver of price.[29] This is a radical departure from traditional approaches to exchange rate determination. Traditional approaches, with their common-knowledge environments, admit no role for information aggregation. Our findings suggest instead that the problem this market solves is indeed one of information aggregation.

Our Portfolio Shifts model provides an explicit characterisation of this information aggregation problem. The model is also strikingly successful in accounting for realised rates. It accounts for more than 60% of daily changes in the DM/$ rate, and more than 40% of daily changes in the Yen/$ rate. Out of sample, our model produces better short-horizon forecasts than a random walk. Our estimates of the sensitivity of the spot rate to order flow are sensible as well, and square with past estimates at the individual-dealer level. We find that for the DM/$ market as a whole, $1 billion of net dollar purchases increases the DM price of a dollar by about 0.5%. This relation should be of particular interest to people working on central bank intervention (though care should be exercised in mapping central bank orders to subsequent interdealer trades).

Two issues raised by our measure of order flow deserve some remarks. First, though our measure captures a substantial share of total trading, it remains incomplete. As data sets covering customer-dealer trading and brokered interdealer trading become available, the order-flow picture can be completed (see, eg Payne 1999). A second interesting issue raised by our order-flow measure is whether its relation to price would change if order flow were observable to dealers in real time (ie if the market were more transparent). From a policy perspective, the effects of increasing order-flow transparency may be important: unlike most other financial markets, the FX market is unregulated in this respect. The welfare consequences are not yet well understood.

So where do the results of this paper lead us? In our judgement they point toward a research agenda that borrows from both the macro and microstructure approaches. It is not necessary to decouple exchange rates from macroeconomic fundamentals, as is common within microstructure finance. In this way, the approach is more firmly anchored in the broader context of asset pricing. (Though we freely admit that longer time series will be necessary to implement the macro dimension fully.) Nor is it necessary to treat exchange rates as driven wholly by public information, as is common within the macro approach. The information aggregation that arises when one reduces reliance on public information is well suited to microstructure: there are ample tools within the microstructure approach for addressing this aggregation. In the end, this two-pronged approach may help locate the missing middle in exchange rate economics - that disturbing space between our successful modelling of very short and very long horizons.

We close by addressing the obvious challenge for this agenda: What drives order flow? Here are two strategies, among many, for shedding light on this question. The first strategy involves disaggregating order flow. For example, interdealer order flow can be split into large banks versus small banks, or investment banks versus commercial banks. Data sets on customer order flow can be split into non-financial corporations, leveraged financial institutions (eg hedge funds), and unleveraged financial institutions (eg mutual and pension funds). Do all these trade types have the same price impact? If not, whose trades are most informative? This will clarify the underlying sources of non-CK information,

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29 This primacy of order flow within microstructure should mitigate standard concerns about data snooping. (In the words of Richard Meese 1990, "At this point exchange rate modelers can be justly accused of in-sample data mining.") The variable we introduce - order flow - is the obvious a priori driving variable from microstructure theory.

which brings us closer to a specification of this market's information structure. The second strategy involves focusing on periods where we expect CK information to be most important, for example, periods encompassing scheduled macro announcements. Does order flow account for a smaller share of the price variation within these periods? Or is order flow an important channel for resolving uncertainty about the mapping from public information to price? Whatever the result, the important point is that the what-drives-order-flow question is not beyond our grasp.

Appendix: model solution

Each dealer determines quotes and speculative demand by maximising a negative exponential utility function defined over terminal wealth. Because returns are independent across periods, with an unchanging stochastic structure, the dealers' problem collapses to a series of independent trading problems, one for each period. Within a given period t, let $W_{i\tau}$ denote the end-of-round τ wealth of dealer i, where we use the convention that W_{i0} denotes wealth at the end of period t-1. (To ease the notational burden, we suppress the period subscript t when clarity permits.) With this notation, and normalising the gross return on the riskless asset to one, we can write the dealers' problem as:

(A1) $$\underset{\{P_{i1},P_{i2},P_{i3},T_{i2}\}}{Max} \quad E\left[-\exp(-\theta W_{i3} \mid \Omega_i)\right]$$

s.t.

$$W_{i3} = W_{i0} + c_{i1}(P_{i1} - P'_{i2}) + \left(D_{i2} + E[T'_{i2} \mid \Omega_{i2}]\right)\left(P_{i3} - P'_{i2}\right) - T'_{i2}\left(P_{i3} - P_{i2}\right)$$

$P_{i\tau}$ is dealer i's round-τ quote and a ′ denotes an interdealer quote or trade received by dealer i. The dealers' problem is defined over four choice variables: the three scalar quotes P_{i1}, P_{i2}, and P_{i3}, and the dealer's outgoing interdealer trade in round 2, T_{i2}. This outgoing interdealer trade in round 2 has three components:

(A2) $$T_{i2} = c_{i1} + D_{i2} + E[T'_{i2} \mid \Omega_{Ti2}]$$

where D_{i2} is dealer i's speculative demand in round 2, and $E[T'_{i2}|\Omega_{Ti2}]$ is the dealer's attempt to hedge against incoming orders from other dealers (this term is zero in equilibrium). The last three terms in W_{i3} capture capital gains/losses from round-1 customer orders c_{i1}, round-2 speculative demand D_{i2}, and the round-2 position disturbance from incoming interdealer orders T'_{i2}. The conditioning information Ω_i at each decision node (3 quotes and 1 outgoing order) is summarised below.

$$\Omega_{Pi1} \equiv \left\{ {}_{k=1}^{t}\{r_k\}, {}_{k=1}^{t-1}\{\Delta x_k\} \right\}$$

$$\Omega_{Pi2} \equiv \left\{ \Omega_{Pi1}, c_{i1} \right\}$$

$$\Omega_{Ti2} \equiv \left\{ \Omega_{Pi2} \right\}$$

$$\Omega_{Pi3} \equiv \left\{ \Omega_{Pi2}, \Delta x_t \right\}$$

Conditional variances

This appendix repeatedly uses several conditional return variances. These variances do not depend on conditioning variables' realisations (eg they do not depend on dealer i's realisation of c_{i1}. These conditional variances are therefore common to all dealers and known in period one. (It is a convenient property of the normal distribution that realisations of conditioning variables affect the conditional mean but not the precision of the condition mean.) This predetermination of conditional variances is key to the derivation of optimal quoting and trading rules.

Equilibrium

The equilibrium concept we use is Bayesian-Nash Equilibrium, or BNE. Under BNE, Bayes rule is used to update beliefs and strategies are sequentially rational given those beliefs.

Solving for the symmetric BNE, first we consider properties of optimal quoting strategies.

PROPOSITION 1: A quoting strategy is consistent with symmetric BNE only if the round-one and round-two quotes are common across dealers and equal to:

$$P_{1,t} = P_{2,t} = P_{3,t-1} + r_t$$

where $P_{3,t-1}$ is the round-three quote from the previous period, and r_t is the public-information innovation at the beginning of period t.

PROPOSITION 2: A quoting strategy is consistent with symmetric BNE only if the common round-three quote is:

$$P_{3,t} = P_{2,t} + \lambda\Delta x_t$$

The constant λ is strictly positive.

Proof of propositions 1 and 2

No arbitrage requires that all dealers post a common quote in all periods. (Recall from section three that all quotes are scalar prices at which the dealer agrees to buy/sell any amount, and trading with multiple partners is feasible.) Common prices require that quotes be conditioned on commonly observed information only. In rounds one and two, this includes the previous period's round-three price, plus the public-information innovation at the beginning of period t, r_t. (Dealer i's round-two quote therefore cannot be conditioned on his realisation of c_{i1}.)

The equations that pin down the levels of these three prices embed the dealer and customer trading rules. When conditioned on public information, these trading rules must be consistent with equilibrium price. This implies the following key relations:

(A3) $$E[c_{i1}|\Omega_{Pi1}] + E[D_{i2}(P_{1,t})|\Omega_{Pi1}] = 0$$

(A4) $$E[c_{i1}|\Omega_{Pi1}] + E[D_{i2}(P_{2,t})|\Omega_{Pi1}] = 0$$

(A5) $$E[\Sigma_i c_{i1}|\Omega_{Pi3}] + E[c_3(P_{3,t})|\Omega_{Pi3}] = 0$$

The first two equations simply state that, in expectation, dealers must be willing to absorb the demand from customers. The third equation states that, in expectation, the public must be willing at the round-3 price to absorb the period's aggregate portfolio shift. These equations pin down equilibrium price because any price except that which satisfies each would generate net excess demand in round-2 interdealer trading, which cannot be reconciled since dealers trade among themselves.

That $P_{1,t}=P_{2,t}=P_{3,t-1}+r_t$ follows directly from the fact that expected value of c_{i1} conditional on public information Ω_{Pi1} is zero, and expected speculative dealer demand D_{i2} is also zero at this public-information-unbiased price. To be more precise, this statement *postulates* that the dealer's demand D_{i2} has this property; we show below in the derivation of the optimal trading rule that this is the case.

That $P_{3,t}=P_{2,t}+\lambda\Delta x_t$ follows from the fact that Δx_t is a sufficient statistic for the period's aggregate portfolio shift $\Sigma_i c_{i1}$. Given the aggregate portfolio shift must be absorbed by the public in round 3, $P_{3,t}$ must adjust to induce the necessary public demand. Specifically, the round-3 price must satisfy:

$$c_3(P_{3,t}) = -\Sigma_i c_{i1}$$

Given the optimal rule for determining T_{i2} (which we establish below), we can write $\Sigma_i c_{i1}$ in terms of interdealer order flow Δx_t as:

$$\Sigma_i c_{i1} = (1/\alpha)\Delta x_t$$

and since the specification of c_3 in the text is:

$$c_3 = \gamma\Big(E[P_{3,t+1}|\Omega_3] - P_{3,t}\Big)$$

this implies a market-clearing round-3 price of:

$$P_{3,t} = E[P_{3,t+1}|\Omega_3] + (\alpha\gamma)^{-1}\Delta x_t$$
$$= \sum_{i=1}^{t}\left(r_i + \lambda\Delta x_i\right)$$

with $\lambda=(\alpha\gamma)^{-1}$, which is unambiguously positive. This sum is the expected payoff on the risky asset (the r_i terms), adjusted for a risk premium, which is determined by cumulative portfolio shifts (the Δx_i terms). This yields equation (5) in the text:

(5) $$\Delta P_t = r_t + \lambda\Delta x_t$$

where ΔP_t denotes the change in price from the end of round three in period t-1 to the end of round three in period t.

Equilibrium trading strategies

An implication of common interdealer quotes $P_{2,t}$ is that in round two each dealer receives a share 1/(N-1) of every other dealer's interdealer trade. This order corresponds to the position disturbance T'_{i2} in the dealer's problem in equation (A1). Given the quoting strategy described in propositions one and two, the following trading strategy is optimal and corresponds to symmetric linear equilibrium:

PROPOSITION 3: The trading strategy profile:

$$T_{i2} = \alpha c_{i1}$$

$\forall i \in \{1,...,N\}$, with $\alpha>0$, conforms to a Bayesian-Nash equilibrium.

Proof of proposition 3: optimal trading strategies

As noted above, because returns are independent across periods, with an unchanging stochastic structure, the dealers' problem collapses to a series of independent trading problems, one for each period. Because there are only N dealers, however, each dealer acts strategically in the sense that his speculative demand depends on the impact his trade will have on subsequent prices.

It is well known that if a random variable W is distributed $N(\mu,\sigma^2)$ and the utility function $U(W)= -\exp(-\theta W)$, then:

(A6) $$E[U(W)] = -\exp\left[-\theta(\mu - \theta\sigma^2/2)\right]$$

Maximising E[U(W)] is therefore equivalent to maximising $(\mu - \theta\sigma^2/2)$. This result allows us to write the dealers speculative-demand problem as:

(A7) $$\underset{D_{i2}}{\text{Max}} \quad D_{i2}\left(E[P_3|\Omega_{Ti2}]-P_2\right) - D_{i2}^2(\theta/2)\sigma^2$$

where the information set Ω_{Ti2} is defined above, and σ^2 denotes the conditional variance of $E[P_3|\Omega_{Ti2}]-P_2$. Now, from Proposition two, we can write:

(A8) $$E[P_3|\Omega_{Ti2}] - P_2 = E[\lambda\Delta x|\Omega_{Ti2}]$$

And from the definitions of Ω_{Ti2} and Δx we know that:

(A9) $$E[\lambda\Delta x|\Omega_{Ti2}] = \lambda T_{i2}$$

The expected value of the other dealers' trades in Δx is 0 under our specification because (i) customer trades are mean-zero and independent across dealers and (ii) there is no information in the model other than customer trades to motivate speculative demand. This fact also implies that dealer i's trade in round 2, T_{i2} from equation (A2), is equal to:

$$T_{i2} = D_{i2} + c_{i1}$$

Therefore, we can write the dealer's problem as:

(A10) $$\underset{D_{i2}}{\text{Max}} \quad D_{i2}\lambda(D_{i2} + c_{i1}) - D_{i2}^2(\theta/2)\sigma^2$$

The first-order condition of this problem is:

(A11) $$2\lambda D_{i2} + c_{i1} - \theta\sigma^2 D_{i2} = 0$$

which implies a speculative demand of:

(A12) $$D_{i2} = \left(\frac{1}{\theta\sigma^2 - 2\lambda}\right)c_{i1}$$

This demand function and the fact that $T_{i2}=D_{i2}+c_{i1}$ imply:

(A13)

$$T_{i2} = \left(\frac{1}{\theta\sigma^2 - 2\lambda} + 1 \right) c_{i1} \equiv \alpha c_{i1}$$

The second-order condition for a maximum, $(2\lambda - \theta\sigma^2)<0$, insures that $\alpha>0$.

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Comments on "Order flow and exchange rate dynamics" by Martin Evans and Richard Lyons

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At first glance, explaining the exchange-rate dynamics appears to be a challenge. Since the seminal work by Meese and Rogoff (1983), we know that macro determinants are helpless to explain changes in exchange rate. Moreover, microstructure models are considered to be adequate to predict exchange rate, but only at a very high frequency.

This paper is very ambitious, since it aims at using a theoretical determinant of the exchange rate from the microstructure field to explain the exchange rate dynamics at a daily frequency. The authors argue that this new determinant, the order flow, is helpful to forecast exchange rate, because it reveals to investors information on portfolio shifts, which is not common knowledge.

They develop a theoretical model based on the microstructure approach. In this model, the initial portfolio shifts by investors are not publicly observable. Market dealers take the other side of these orders and then trade among themselves to share risk. Order flow is then defined as the sum of net intradealer trades.

Order flow is one of these theoretical concept, which measure is truly challenging. Here, the measurement problem is overcome by the use of a perfectly adequate database. First, the database contains tick-by-tick data, which allows all transactions between dealers to be recorded. But, more importantly, for each transaction, a buyer/solder indicator is available, allowing trades to be signed and therefore order flow to be measured.

The order flow is computed as the difference between the number of buyer-initiated trades and the number of solder-initiated trades. Therefore, the size of trades cannot be used to measure order flow more accurately. In particular, during turbulent periods, there may be asymmetries between the size of buyer-initiated trades and the size of solder-initiated trades.

It is also worth noting that the period covered by the database is pretty long, since the data covers four months for the DM-US and Yen-US exchange rates. Of course, a longer period of time would be desirable to assess more precisely the ability of order flow to forecast exchange rate. But empirical results obtained for the two more traded currencies are rather convincing. A longer sample would allow several specification of the portfolio shift model to be tested. In particular, the effect of order flow may well be a non-linear one.

At this stage, two issues arise: First, can the new explanatory variable be considered as exogenous with respect to the exchange rate? Second, can the empirical evidence obtain by Evans and Lyons been seen as solving the puzzle highlighted by Meese and Rogoff?

1. Exogeneity of order flow

I begin with the exogeneity of order flow. Empirical estimates of the portfolio-shift model are quite impressive: the authors obtain a strong and significant effect of the order flow in explaining change in exchange rate. The interest-rate spread has a correctly signed effect, but it is weakly significant. R^2s are very large, as compared to usual standards. The portfolio-shift model appears to have a much larger explanatory power than standard macro models.

But this raises the issue of order flow exogeneity with respect to the exchange rate. In other words, can order flow be used in forecast exercises? The authors give several convincing arguments why causality goes from order flow to price, and not from price to order flow. However, order flow may well depend, to some extent, on prices. In this case, using order flow to forecast exchange rates would appear to be less promising. In particular, out-of-sample forecasts computed by the authors reveal that the portfolio shift model outperforms significantly, but rather weakly, the usual random walk. But this exercise is performed using realized order flow and interest rates. A real-time forecast would require a preliminary forecast of order flow and interest rates. This would further reduce the forecast ability of order flow. Yet this reduction would be even larger if order flow has to be forecast using past exchange rates.

2. The Meese and Rogoff puzzle

Concerning the Meese and Rogoff puzzle, I would like to make two remarks.

First, Meese and Rogoff stated that macro models fail to predict exchange rate. The failure of macro models seems to come from their inability to represent exchange rate expectations accurately. This paper proposes a new proxy for exchange-rate expectations, the order flow, which seems to be much more accurate than previous proxy variables. From this point of view, this paper succeeds in solving the Meese and Rogoff puzzle, since it exhibits a theoretically and empirically pertinent determinant of exchange rate.

On the other hand, this study has to be extended to longer periods and, more importantly, to other currencies. Indeed, it is usual in the exchange-rate analysis to obtain that a theoretical model is valid for one exchange rate but not for other ones. This is the sense of the result of Meese and Rogoff (1983).

Conclusion

To sum up, the approach developed in this paper is very appealing, since it shows that a natural micro determinant of exchange rate is able to explain a large part of the exchange-rate dynamics.

Two avenues can be explored to answer the question “What drives order flow?” On one hand, as suggested by the authors, further disaggregating order flow would provide information on the kind of trades which is more informative on future exchange rate. On the other hand, one may try to identify macro determinants of order flow. The next step may be to derive the central tendency of exchange rate, which results from all transactions between individual customers and dealers interacting in the market place. This central tendency may be a major determinant of order flow. Since the central tendency has to be modeled in the macro area, this may help to fill the gap between micro and macro approaches.

To conclude, I would like to ask a few questions. First, since the portfolio shift model appears to have a strong forecast ability, how could it be used in practice to forecast exchange rate? More precisely, are the data used to compute order flow available in real time, in order to take advantage of the forecast ability at the one-day horizon, for instance?

Comments on "Order flow and exchange rate dynamics" by Martin Evans and Richard Lyons

Robert N McCauley

This is a rare empirical study of exchange rate economics that leaves the reader with the feeling that he has actually learned something about the subject. It also passes the Leontief test of working to analyse new data rather than just re-squeezing a well-squeezed orange. Let me comment on the theory, data and findings as presented and then make some suggestions as to how the analysis might evolve to address the central concerns of this workshop.

Theory

Evans and Lyons are ambivalent about the theoretical position of order flow. In places it is the expression of other determinants of exchange rates. In other places, it is a competing explanation. Conceiving of order flow as an explanation of exchange rate movements strikes me as similar to conceiving of the number of automobiles on the road in a city as an explanation of variations in the speed of traffic. True, there are ways of thinking about exchange rate movements or traffic jams that ignore order flow or vehicular density. But a traffic engineer wants to know why drivers are out there on the road. It is useful to know that order flow is a proximate cause of exchange rate changes, but one hopes that we can use data on order flow to point to the more remote causes, which may or may not be fundamentals. Conceptually, order flow can be seen as tapping into portfolio shifts - and I find odd the assertion that "Portfolio balance models are driven by asset supply" (p8). Empirical tests of the portfolio theory have sought to find the effect of varying asset supplies, it is true, but the theory contains predictions about the effect of demand shifts, which this paper can be read as verifying.

Data

After applauding the effort entailed in obtaining the Reuters 2000 data, let us put them into context. We are told in the text that the share of Reuters 2000 in direct interdealer trading is 90%--but what share of the spot market is captured? The data are drawn from a period, May-August 1996, when the spot foreign exchange market was undergoing rapid structural change. In particular, electronic broking was growing, apparently not only at the expense of voice broking but also at the expense of direct dealing (BIS (1997), p91).

Lets look at the figures. In 1995 and 1998, 65% and 68% of dollar-DM and dollar-yen spot trading was interdealer: therefore the Reuters 2000 data on trading among dealers do not cover one-third of the spot market (BIS (1996, 1999), Table 1-B, Table E-2). Then, within the interdealer market, brokers account for a third of trades in 1995, rising to about a half in 1998 (Table). So the Reuters data cover well less than half of the spot market overall and one-half to two-thirds of the interdealer market (accepting Reuters estimate of its share of the direct interdealer market).

The authors recognise the drawback of having data only on buy/sell hits (p14). The authors' own theorising says that dollars, not hits, matter, and I can only agree. Especially in light of the Reuters 2000 system capturing only a part, albeit a very significant part, of all transactions, we have an errors in variable problem here at the very least.

One small but significant suggestion regarding interest rates. Overnight rates do not well capture changes in the expected path of policy rates. Something like the three-month rate, three months forward, better captures the swings in policy expectations that move interest rates.

Results

Whatever the slippage between what the authors want to measure and what they do measure, their results suggest the weight of order flow moves the exchange rate. Certainly market participants spend a lot of time and analysts devote a lot of words to assessing how other market participants are positioning and repositioning.

The Study Group on Market Dynamics of the Working Group on Highly Leveraged Institutions (HLIs) learned that HLIs can release their private information on their own positions and intentions to their own benefit. Furthermore, to my mind, it was evident that the sheer weight of HLI positions, particularly in middle-sized markets, could absorb large fractions of the institutionally determined capacity of day-to-day market participants, including exporters and portfolio managers, to take the other side. Under these circumstances, price formation could be very sensitive to news, rumours, or further, noisy positioning. The authors refer to the dollar-yen's sharp move in Fall 1998, but the charts covering that period for the Australian, New Zealand, Singaporean dollars and the South African rand also deserve a look (Market Dynamics (2000)). Order flow and cumulated order flow matter.

What is missing from these results, however, is an active exploitation of the insight that liquidity in a market is variable over time. Let me close on the interaction of order flow and liquidity.

The interaction of order flow and liquidity

The paper makes its most salient contribution to today's discussion when it separates the data into quartiles based on daily transactions volumes. The authors find a U-shaped pattern of coefficients on order flow for the DM, indicating the largest impacts of order flow on days with the lightest and heaviest trading, and an upward pattern for the yen, indicating the strongest impact amid heaviest trading. This is a rough cut, however, particularly if Reuters 2000 gains (or loses) market share during periods of high volatility.

It is known, however, that trading activity varies more by the hour of the day than by day. Consider the claim of the Reserve Bank of Australia that highly leveraged institutions in mid-1998 deliberately traded during Sydney lunch time, or in the slow period between Sydney's wind-down and London's wind-up, in order to have maximum effect on the Australian dollar's exchange rate (Market Dynamics (2000), pp127-8). To address these claims, one would want to approach this question of the variable impact of order flow by time of day.

A finer approach might also seek to separate out the response of order flow to major releases of news building ono the findings of Fleming and Remolona. Liquidity and order flow dynamics might differ quite a bit between normal periods and immediate post-news periods.

In short, the authors should arm themselves with curiosity regarding the waxing and waning of liquidity and see what their data set can tell them. While this paper has made an important contribution to our understanding of exchange rate determination, much work on foreign exchange market liquidity remains to be done.

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Table
Spot interdealer foreign exchange trading by mode, 1995 and 1998

	1995			1998		
	UK	**US**	**JP**	**UK**	**US**	**JP**
Direct			63	64	53	54
Brokered			37	36	47	46
Of which: electronic		*10*	*12*	*25*	*33*	*36*

Sources: Bank of England (1998), p 352; Federal Reserve Bank of New York (1998), pp 6, 8; Bank of Japan (1998), p 224; BIS (1999), p 15.

Note: UK and US data are for spot trading as a whole; US voice broker share for spot trading assumed to be same as for all trading.

Trading volumes, volatility and spreads in FX markets: evidence from emerging market countries

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Abstract

This paper provides empirical evidence on the relationship between trading volumes, volatility and bid-ask spreads in foreign exchange markets. It uses a new data set that includes daily data on trading volumes for the dollar exchange rates of seven currencies from emerging market countries. The sample period is 1 January 1998 to 30 June 1999. The results are broadly consistent with the findings of the literature that used futures volumes as proxies for total foreign exchange trading. I find that in most cases unexpected trading volumes and volatility are positively correlated, suggesting that both are driven by the arrival of public information, as predicted by the mixture of distributions hypothesis. I also find evidence of a positive correlation between volatility and spreads, as suggested by inventory cost models. However, contrary to the prediction of these models, I do not find evidence of a significant impact of unexpected trading volumes on spreads.

1. Introduction

This paper looks at the relationship between trading volumes, volatility and bid-ask spreads in foreign exchange markets. A number of studies on the microstructure of foreign exchange markets have looked at this issue from both a theoretical and an empirical point of view. From a policy perspective, the issue is important for its implications for the analysis of market liquidity and its relationship with risk. Broadly speaking, a market can be considered to be liquid when large transactions can be executed with a small impact on prices (BIS (1999a)). In practice, however, no data are available that allow to measure this definition of foreign exchange market liquidity directly. Instead, trading volumes or bid-ask spreads are often used as indirect measures. Volatility is often considered as a measure of risk.

The empirical microstructure literature has typically found a positive correlation between volumes and volatility. A theoretical explanation of this finding is that volume and volatility are both driven by a common, unobservable factor, which is determined by the arrival of new information. This theory, also known as the mixture of distributions hypothesis, predicts that volatility moves together with unexpected trading volumes. A further common finding of the literature is that volume and spreads are positively correlated. The explanation provided by microstructure theory is that bid-ask spreads are determined inter alia by inventory costs, which widen when exchange rate volatility increases. Through the mixture of distributions hypothesis, this also establishes a positive link between unexpected volumes and spreads.

An important drawback of empirical studies in this area is that good data on foreign exchange trading volumes are generally not available at high frequencies. The most comprehensive source of information on trading in foreign exchange markets, the *Triennial Survey of Foreign Exchange and Derivatives Market Activity* published by the BIS, for example, does not provide much information on the time series behaviour of trading volumes. Researchers have therefore looked at alternative data sources to find proxies for foreign exchange market turnover.

This paper uses a new data set that for the first time matches daily data on trading volumes, volatility and spreads. The data set covers the dollar exchange rates of seven currencies from emerging market countries, the Colombian peso, the Mexican peso, the Brazilian real, the Indian rupee, the Indonesian rupiah, the Israeli shekel and the South African rand. The data cover the period from January 1998 to June 1999. Since there is not much offshore trading in these currencies, local transaction volumes are fairly representative of total trading. In order to allow a comparison with foreign exchange markets in industrial countries and with the results from previous studies, the paper also looks at trading volumes from the Tokyo interdealer yen/dollar market. Finally, for the Mexican peso, data from a fairly active currency futures market on the Chicago Mercantile Exchange were also obtained. These data allow a direct comparison of volumes on foreign exchange and futures markets.

A main finding using this data set is that in most cases unexpected trading volume and volatility are positively correlated, suggesting that they both react to the arrival of new information, as the mixture of distributions hypothesis predicts. This result is in line with the findings of the literature that relies on futures data. The markets for the Mexican peso and the real, however, provide important exceptions. In these two cases, the relationship between unexpected volumes and volatility is not statistically significant. I also find evidence of a positive correlation between volatility and spreads, as suggested by inventory cost models. However, the results do not show a significant impact of trading volumes on spreads.

The remainder of the paper is organised as follows. Section 2 reviews the main contributions to the literature on the relationship between trading volumes, volatility and spreads. Section 3 describes the data set that is used. In Section 4, I present some descriptive evidence on the relationship between volumes, volatility and spreads. I then use regression analysis to test whether the mixture of distribution hypothesis holds in foreign exchange markets. Section 5 concludes.

2. Literature review

There is an extensive literature on the relationship between trading volumes and volatility in financial markets. Karpoff (1987) provides a good overview of the early literature. Most of the research has focused on stock markets and futures markets, for which data on volumes are more easily available. An important finding is that trading volume and price variability are positively correlated at different frequencies. The coefficient is highest for contemporaneous correlations. However, it does not always appear to be very sizeable. Evidence was found by Harris (1986) and Richardson et al. (1987) for stock markets and by Cornell (1981) for commodity futures.

The empirical work on foreign exchange markets has suffered from the problem that good data on trading volumes are not easily available for foreign exchange markets, since unlike equity markets, they are for a large part decentralised. Different data sources were used to describe the time series behaviour of trading volumes.[1] Many studies used data on futures contracts, which can be easily obtained, to proxy for interbank trading volumes. Studies that have found a positive correlation between volumes and volatility in these markets include Grammatikos and Saunders (1986), Batten and Bhar (1993) and Jorion (1996). An obvious drawback of these data sets is that trading in futures is very small compared to OTC volumes (Dumas (1996)). In the first quarter of 1998, for example, total turnover of currency futures traded on organised exchanges amounted to roughly $70 billion (BIS (1998)), compared to total OTC turnover in spot, forward and swap markets of about $1500 billion (BIS (1999b). While these two markets may still be closely linked through arbitrage (Lyons (2000)), little evidence is available on this link.

A widely used source of information on foreign exchange trading, the *Triennial Survey of Foreign Exchange and Derivatives Market Activity* published by the BIS, provides extensive cross-sectional information but very little information on the time series behaviour of turnover. Hartmann (1998b) makes efficient use of these data by combining a large cross-section of exchange rates taken from the BIS Triennial Survey with two time series observation into a panel but still faces the problem of having limited time series information.

A number of papers analysed data on indicative quotes provided by Reuters through its FXFX page.[2] However, these quotes do not represent actual trades and it is not possible to infer from a quote for which volume it is given. Spreads that are quoted on the Reuters screen are generally far from actual, traded spreads.[3] Moreover, it is common for banks that act as data providers to program an automated data input, e.g. by having a particular quote entered at regular time intervals. This is especially true for smaller banks that may have an interest in quoting prices in order to advertise their presence in a particular market segment. Finally, when an important event occurs, traders are likely to act and trade rather than entering data for Reuters. Hence, Reuters tick frequency may be low at times

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1 A good overview of the characteristics of data sets used in the literature can be found in Lyons (2000).

2 See, eg Goodhart and Figliuoli (1991), Bollerslev and Domowitz (1993). And Melvin and Yi (2000).

3 See also Hartmann (1998a) , p.142.

of high trading activity and high when markets are calm. The relationship between quote frequency and actual trading activity is therefore likely to be quite noisy.

An alternative source of information is the Bank of Japan's data set on brokered transactions in the Tokyo yen/dollar market, which has been used by Wei (1994) and Hartmann (1999). Again, a problem with these data is that they represent only a fraction (about one-sixth) of total turnover in the Tokyo yen/dollar market and not more than 5% of the global yen/dollar market.

A fruitful approach has been to look at high-frequency data on actual transactions in the OTC market. One such data set, used by Lyons (1995), covers all transactions that a foreign exchange dealer in New York entered with other dealers in one week in 1992. Goodhart, Ito and Payne (1996) analyse data on electronically brokered interdealer transactions that occurred on one day in 1993. While these data provide a wealth of information, including information on the direction of order flows, they necessarily cover only a limited segment of foreign exchange markets and span a relatively short time period.

Different theoretical explanations have been offered for the co-movement of trading volumes and volatility. Early work was based on models of "sequential information arrival" (Copeland (1976,1977)), according to which information reaches one market participant at a time. As that agent reacts to the arrival of news, his demand curve will shift, thereby leading to a positive correlation between volume and volatility. An alternative explanation of the volume-volatility correlation is based on the "mixture of distribution hypothesis" first proposed by Clark (1973). According to this hypothesis, volume and volatility are both driven by a common, unobservable factor. This factor reflects the arrival of new public information, and determines a positive correlation between unexpected turnover and unexpected volatility. Tauchen and Pitts (1983) show that volume and volatility can co-move for two reasons. First, as the number of traders grows, market prices become less volatile. Second, given the number of traders, an increase in volume reflects a higher disagreement among traders and hence leads to higher volatility. This link is stronger when new information arrives at a higher rate.

A number of studies looked at the relationship between volumes and volatility by testing the common observation that in both bond and equity markets, an upward movement of the market is generally associated with higher transaction volumes. The empirical evidence for this relationship is weak. While some studies (e.g. Epps (1975, 1977), Rogalski (1978) and Richardson et al. (1987)) have found weakly supportive evidence, other studies (e.g. James and Edminster (1983) and Wood et al. (1985)) have failed to detect any significant correlation. Moreover, no favourable evidence has been found for futures markets.

Models that explain bid-ask spreads in terms of inventory costs establish a link between bid-ask spreads, volatility and trading volumes. One determinant of inventory costs is the cost of maintaining open positions, which is positively related to price risk.[4] According to this view, exchange rate volatility increases price risk and thereby pushes up spreads. Supportive evidence is provided by Bessembinder (1994), Bollerslev and Melvin (1994) and Hartmann (1999), who found a positive correlation between spreads and expected volatility measured by GARCH forecasts.

A second determinant of inventory costs is trading activity. Trading volumes can have a different impact on spreads depending on whether they are expected or unexpected. Expected trading volumes should be negatively correlated with spreads to the extent that they reflect economies of scale and are associated with higher competition among market makers (Cornell (1978)). By contrast, unexpected trading volumes should have a positive impact on spreads to the extent that they are associated with higher volatility through the mixture of distribution hypothesis.

4 The microstructure literature analyses two other types of costs, order processing costs (i.e. costs of providing liquidity services) and asymmetric-information costs (Bessembinder (1994), Jorion (1996) Hartmann (1999), Lyons (2000)). While order processing costs are arguably small in foreign exchange market (Jorion (1996)), there is no consensus on the importance of asymmetric-information costs.

3. The data

Exchange rates

This paper looks at the dollar exchange rates of seven currencies from emerging market countries: the Indonesian rupiah, the Indian rupee, the Mexican peso, the Brazilian real, the Colombian peso, the South African rand and the Israeli shekel.[5] The sample period is January 1998 to June 1999.[6] According to the IMF classification, during this period these exchange rates either floated independently or were managed in some form (Table 1). Figure 1 and Table 2 provide information on their behaviour. As a benchmark for comparison, information on the yen/dollar rate is also shown. Overall, all seven emerging market currencies depreciated against the dollar. The depreciation was particularly sharp following the Russian crisis in August 1998, except in the case of the real.

The real depreciated at a constant rate against the dollar in 1998. Following intense speculative pressures, Brazilian authorities had to float it freely on 15 January 1999 and it plunged from 1.1 to almost 2.2 real/dollar. The real recovered somewhat in the following months, and in June 1998 traded around a level of 1.75 to the dollar. The Mexican peso depreciated through most of 1998 and early 1999. However, it moved within a much smaller range than the real. The Mexican peso's weakness was exacerbated following the Russian crisis in August 1998, when it depreciated by around 20% against the dollar. It also fell, albeit not as much, in January 1999, during the period of speculative attacks on the real. The Colombian peso's downward trend and fluctuation range against the US dollar were similar to those of the Mexican peso. In August 1998, pressure on the peso stepped up, inducing the authorities to widen the intervention band by 9%, effectively devaluing the currency. The peso came under renewed pressure in March 1999 and June 1999, when it devalued by about 20%.

The Indian rupee's behaviour was characterised by periods of stability followed by sharp downward movements. The Indonesian rupiah fell sharply in January and in July 1998. Its volatility declined in 1999 but remained still very high.

The South African rand depreciated markedly against the dollar around the time of the Russian crisis in the summer of 1998. In the following months it recouped part of its losses but in 1999 trended down again. In 1998, the shekel followed a slightly depreciating trend against the dollar. It fell by 20% after the Russian crisis but stabilised in the following weeks.

Figure 3 and Table 3 report information on the percentage exchange rate changes for the seven currencies. For the real, statistics are presented also separately for the period January-June 1999, during which it floated. Between January 1998 and June 1999, the average daily percentage change of most of the currencies was significantly positive, with the exceptions of the Colombian peso and the rupee. Their standard deviation ranged between 0.29 for the rupee and 4.24 for the rupiah. Over the same period, the standard deviation of the yen/dollar rate was close to 1. Most exchange rate changes exhibited positive skewness, consistently with their downward trend, and leptokurtosis. Table 4 suggests that the exchange rate changes exhibited very little persistence.

Volatility

Figure 3 shows the historical volatility of the seven exchange rates computed over moving windows of one month. Summary statistics are reported in Table 5. In terms of their volatility, the seven exchange rates can be divided into two groups. The real and the rupiah experienced sharp volatility spikes and were characterised by very high average volatility. By contrast, the historical volatilities of the rupee, the shekel and the Colombian peso remained quite low, averaging less than 8%. The rand's volatility was relatively low on average, but it spiked at about 50%.

Table 5 shows that the volatility of the yen/dollar exchange rate during the same period averaged 15%. This is much less than the volatility of the exchange rate of the real or the rupiah vis-à-vis the dollar, but more than the volatility of the other five exchange rates in this data set.

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[5] See, eg Goodhart and Figliuoli (1991), Bollerslev and Domowitz (1993). And Melvin and Yi (2000).

[6] The choice of the sample period is dictated by the availability of data on trading volumes.

A finding that is common for exchange rates and other asset prices is the existence of volatility clustering, i.e. the fact that periods of persistent turbulence are followed by periods of relative calm. In the finance literature, this phenomenon has typically been described by some ARCH-type models. Tests for ARCH-type effects presented in Table 6 suggest that the exchange rates under investigation - possibly with the exception of the Colombian peso - exhibit some ARCH-type behaviour. This behaviour seems to be fairly well captured by a GARCH (1,1) model.

Finally, Table 7 present evidence that for half of the exchange rates, volatility follows a time trend. The trend is positive for the Colombian peso and the real and negative for the rupee and rupiah. There is no evidence of weekend effects.

Trading volumes

A large literature has documented the behaviour of trading volumes in stock markets and bond markets. By contrast, there is relatively little work on foreign exchange markets because of the difficulty of obtaining data with a sufficiently high frequency in this decentralised market. In this paper I use daily data on local turnover on seven exchange rates from emerging market countries. The sample period is January 1998 to June 1999, except for the rupiah, for which data are available only from 1 January to 30 June 1999. Since there is not much offshore trading in these currencies, local transaction volumes are fairly representative of total trading.[7] As a benchmark, I also included trading volumes from the Tokyo interbank market, which were used in previous studies. The time series of the trading volumes are shown in Figure 1 and summary statistics are reported in Table 8.

A comparison with data from the 1998 Triennial Survey suggests that taken together, trading in the seven currency pairs accounts for roughly 1-2% of total foreign exchange market turnover. The individual markets are characterised by very different levels of activity. The most active markets were those for the Mexican peso, the real and the rand. To get an idea of the size of these markets, trading of dollars against one of these currencies averaged about one-third of local trading of Canadian or Australian dollars against the US dollar in April 1998. Taking the yen/dollar market as a benchmark, the size of these markets is slightly smaller than the interbank market in Tokyo. The market for the rupee also appeared to be quite active ($3.5 billion a day), much more than the market for the Indonesian rupiah ($1 billion per day), which in 1995 and 1996 was the most active foreign exchange market in emerging market countries with an average daily turnover of more than $8 billion (BIS (1997)). Transaction volumes for the Colombian peso and the shekel instead are relatively small.

Figure 1 reveals some interesting facts about the behaviour of trading volumes. First, it shows also that the foreign exchange turnover involving the real, the Colombian and the Mexican peso and the rand fell substantially around the end of August 1998, at time when the Russian crisis sparked a global reduction of liquidity in financial markets. The decline in trading volumes was particularly sharp in Mexico and Brazil.

Second, while trading volumes behaved differently around periods of speculative pressures, they always fell substantially once the pressures abated.[8] This fact can be illustrated with the examples of the real and the Colombian peso. At the time of the speculative attack on the real in January 1999, trading volumes for the real rose for a short time and became more volatile. After Brazilian authorities decided to float the real on 12 January 1999, trading of reais against dollars remained very volatile for several weeks and before shrinking substantially around mid-February. In the following months, turnover built up again steadily but remained below levels observed before the crisis.

The Colombian peso was subject to two rounds of pressure in 1999. When pressure mounted first in March, there appears not to have been a significant increase in foreign exchange turnover. Following this period, daily turnover dried up from about $200 mn to less than $50 mn. At the time of the second attack on the peso in June 1999, turnover increased sharply to $150 mn or more. Since our sample period ends on 30 June 1999, it is not possible to verify how trading volumes behaved after the second attack.

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[7] Until 1997 there was a quite active rupiah/dollar market in Singapore. However, most of this activity disappeared following the Asian crisis.

[8] This is true even when volumes are measured at constant exchange rates.

Regressions of trading volumes on a constant, a linear time trend and a dummy variable capturing weekend effects show a negative trend for trading in the Colombian and Mexican peso, the real (until January 1999) and the rand. By contrast, there is evidence that turnover of the rupiah followed a positive time trend. The volume of yen/dollar transactions brokered in Tokyo instead did not exhibit any time trend. Graph 2, which plots turnover in selected emerging markets over a longer horizon (and at a lower frequency), puts these findings in perspective. It suggests that trading in emerging market currencies rose in the mid-1990s but in most cases fell after the Asian crisis and again at the time of the global reduction of liquidity in financial markets in 1998. Finally, Table 7 shows also that weekend effects appear to be present in about half of the cases.

For the Colombian peso, data on the transaction frequency is also available. It is sometimes argued that the combination of these data and trading volumes provides a good indicator on market liquidity. Table 8 suggest that for the Colombian peso/dollar exchange rate the behaviour and the statistical properties of turnover and transaction frequency are very similar.

Table 10 presents some summary statistics on the changes of daily trading volumes. On average, trading volumes for the Colombian peso, the rand, the real and the rupiah appear to be most stable. The statistic for the rupiah is particularly noteworthy and indicative of the low variability of turnover associated with the low activity after the Asian crisis. For the Colombian peso, the statistics for turnover and transaction frequency appear to be very similar. The Mexican peso and the shekel exhibit the highest average variability of trading volumes.

Futures contracts for three exchange rates, the Mexican peso/dollar, the real/dollar and the rand/dollar rate are traded on the Chicago Mercantile Exchange.[9] Table 9 provides summary statistics for these series. Trading is most active for the Mexican peso/dollar contracts, but in any case much less - no more than 3% - than OTC turnover. Figure 3 shows that on many days there is actually no trading on the exchanges at all involving the rand or the real. Exchange traded and OTC volumes appear not to have co-moved closely in 1998 and 1999. The correlation coefficient of daily percentage changes is less than 1% for the Mexican peso, 4% for the rand and 18% for the real.

Spreads

Table 12 reports summary statistics on bid-ask spreads, expressed as a fraction of the exchange rate. As a benchmark for comparison, it also provides information on the spreads on brokered yen/dollar transactions in Tokyo. A caveat about these data is that they refer to indicative quotes rather than to actual transactions. Spreads on actual trades may be much smaller (and possibly less volatile), especially when one looks at electronically brokered transactions.[10] Subject to these caveats, Table 12 shows that spreads on exchange rates from emerging market countries are much higher and much more volatile than spreads on the yen/dollar exchange rate. Their average in 1998 and 1999 ranged from 0.12% of the exchange rate for the Mexican peso to more than 2% for the rupiah, compared to spreads around 0.05% for the yen/dollar market. Table 7 provides some additional information on the variations of the bid-ask spreads.

4. Volumes, volatility and spreads

Trading volumes and volatility

A well-established fact for stock markets is that trading volumes are positively correlated with price variability.[11] A similar result has also been found for foreign currency futures (Grammatikos and Saunders (1986)). The middle panels in Graph 2 suggest that when turmoil hits foreign exchange markets, both volatility and trading volumes increase. Once the currency has depreciated markedly, trading volumes tend to fall substantially.

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[9] There is also active trading in real/dollar futures contracts on the BM&F in São Paolo and on the rand in London.

[10] See Lyons (1995) and Hartmann (1999) and Cheung and Chinn (1999). Most of yen/dollar spot transactions between banks are nowadays conducted through electronic brokers and have spreads about two to three hundredths of a US cent.

[11] See Karpoff (1987, pp.112-3) for an overview of this literature.

In terms of the contemporaneous correlation between daily foreign exchange turnover and exchange rate volatility, Table 14 shows positive coefficients for five out of seven emerging market exchange rates (the dollar exchange rates of the Colombian peso, rand, rupee, rupiah and shekel). The correlation is also positive for the yen/dollar rate traded in the Tokyo interbank market. This holds irrespectively of whether exchange rate volatility is measured by absolute values of percentage changes, squared returns or the standard deviation of daily returns computed over rolling windows of one-month. By contrast, I find a negative correlation for the real and the Mexican peso.

Regressions of volatility on a constant, a time trend, a day-of-the-week dummy and trading volumes gives positive and statistically significant coefficients in all cases except the real and the Mexican peso (Table 16). For the Colombian peso, a regression of volatility on the number of deals per day also gives a positive and statistically significant coefficient. With the exception of the real and the Mexican peso, these results are consistent with the finding of a positive correlation of volatility and volume found for currency futures (Grammatikos and Saunders (1986), Jorion (1996)). Table 15 shows that for the Mexican peso, exchange traded data give a positive correlation between volatility and volumes. This result suggests that exchange traded data may not always be an appropriate proxy for total interbank trading.

Another interesting issue is whether trading volumes are “directional” in the sense that they increase (decrease) when asset prices increase (decrease). Testing this hypothesis involves testing for a positive correlation of volumes and price changes. There is some, albeit weak, evidence in favour of this hypothesis for stock and bond markets (Karpoff (1987)). By contrast, there is no supporting evidence for futures markets. Table 14 does not provide clear evidence of a positive relationship between volumes and price changes in foreign exchange markets. For some exchange rates - the Colombian and the Mexican peso, rand, rupee and shekel - changes are positively correlated with trading volumes but the correlation coefficients are small. The correlation coefficient is negative for the real and rupiah. The same is true for yen/dollar rate traded in Tokyo.

A related question is whether trading volume and exchange rate changes are related in an asymmetric fashion, i.e. whether volumes are higher when a currency appreciates or when it depreciates against the dollar. Karpoff (1987) argued that most empirical findings for stock markets support what he calls an “asymmetric volume-price change hypothesis”, i.e. a fundamentally different relationship between volume and price changes for positive and negative price changes. The evidence presented in Table 14 appears only weakly consistent with this hypothesis. For the Mexican peso, rupee, shekel and the yen, the relationship between volumes and exchange rate changes is almost identical in the two cases. For the other currencies, volumes change more in the case of a depreciation than in the case of an appreciation.

The positive correlation between volumes and volatility found for most of the exchange rates is unlikely to be a reflection of changes in the number of traders active in these markets. These changes appear rather to have occurred in the mid-1990s, when banks increasingly moved into emerging markets, and after the Asian crisis, when the sharp fall in turnover was accompanied by a significant decline in the number of traders. A more plausible explanation for the positive correlation between turnover and volatility is that both variables are driven by a common, unobservable factor. According to the mixture of distribution hypothesis, this factor reflects the arrival of new information, and determines a positive correlation between unexpected turnover and unexpected volatility. Moreover, as Tauchen and Pitts (1983) show, volume and volatility can co-move because, given the number of traders, an increase in volume may reflect a higher disagreement among traders, which also leads to higher volatility. The co-movement is closer when new information arrives at a higher rate.

To test this hypothesis, I split volatility and trading volumes into expected and unexpected components. I use estimates from a GARCH(1,1) model to describe expected volatility. This model appears to fit the time series well.[12] Ideally, volatility implied in option prices could be used, since there is evidence that it outperforms GARCH models in providing forecasts of future volatility.[13] However, option contracts for currencies of emerging market countries are not very liquid, particularly after the Asian crisis. The GARCH(1,1) model can be written as:

[12] Following a common practice of the literature, the GARCH-model is fitted on the entire time series, therefore yielding in-sample forecasts.

[13] Jorion (1996); Galati and Tsatsaronis (1996).

$$(1) \quad R_t = \mu + r_t, \quad r \sim N(0,h_t), \quad h_t = \alpha_0 + \alpha_1 r_{t-1}^2 + \beta h_{t-1}$$

where R_t is the return, μ its mean and h_t its conditional variance at time t.

In order to measure expected trading volumes, I used the Box-Jenkins analysis to select a parsimonious time series representation for the volume series, which are taken in logs. Time series models were fitted on the levels of trading volumes since Augmented Dickey-Fuller test suggest that they are stationary. AR models, in most cases of first order, seemed appropriate to represent the turnover series. These models allow to split trading volumes into an expected and an unexpected component.

The regression equation that is estimated takes the following form:

$$(2) \quad R_{t+1}^2 = \alpha + \beta_1 h_{t+1} + \beta_2 E_t(v) + \beta_3 [v_{t+1} - E(v)] + \beta_4 t + \beta_5 w_t + \varepsilon_{t+1}$$

where total volatility R^2_{t+1} is defined as squared returns, expected volatility h_{t+1} is the one-step-ahead conditional return variance from a GARCH(1,1) specification, and log-volumes are decomposed into an expected component $E_t(v)$ and an unexpected component $[v_{t+1}-E_t(v)]$ by using a fitted AR series and its residuals. A linear time trend and a dummy capturing weekend effects are also included.

Table 17 reports the regression results. The coefficient on unexpected turnover is positive and statistically significant at 1% or 5% in all the regressions for exchange rates from emerging market countries, except those for the Mexican peso and the real. For these two currencies, the coefficients are negative but not statistically significant. A positive, significant coefficient is also found for the yen/dollar rate traded in Tokyo. Except for the Mexican peso and the real, the results support the idea that information flow drives volatility and volumes, as implied by the mixture of distributions hypothesis. This result is consistent with the conclusion of the literature that used data on currency futures (Jorion (1996)). These results are independent of market size: they hold for both for the smallest market (Colombian peso/dollar) as for the biggest market (ran/dollar) in emerging market countries, as well as for the even bigger yen/dollar interbank market in Tokyo. This is consistent with the finding presented in Batten and Bhar (1993) for futures markets. Differently from Jorion's results, however, expected volumes also have a positive, significant effect on volatility in three cases (rupee, shekel and yen).

Table 18 shows the results for regressions that also include expected volatility, measured by the GARCH forecast, among the explanatory variables. The coefficient on unexpected turnover remains positive and statistically significant in most cases. This is in line with the results presented in Jorion (1996). Again, the coefficient on unexpected trading volume it is negative but not statistically significant for the Mexican peso and the real. In these cases only, the GARCH volatility forecast is also significant.

Overall, the results support the idea that the arrival of new public information drives the positive correlation between volumes and volatility, as postulated by the mixture of distribution hypothesis. Favourable evidence is found for four out of six exchange rates from emerging market countries and for the Tokyo interbank yen/dollar market. These findings appear to be independent of market size. By contrast, the mixture of distribution hypothesis appears not to hold for the Mexican peso/dollar and real/dollar markets.

Trading volumes and bid-ask spreads

Figure 2 highlights that in foreign exchange markets in emerging market countries, bid-ask spreads spiked during times in which volatility sharply increased and turnover fell. While spreads tended to narrow shortly after these episodes, in some cases they remained wide for some time. Table 14 shows that in foreign exchange market in emerging market countries, spreads and volatility are positively correlated. In most cases spreads and trading volumes are negatively correlated, a result that contrasts with findings of the early literature.[14] By contrast, the behaviour of spreads appears totally unrelated to changes in volumes and volatility in the Tokyo yen/dollar interbank market, as indicated by correlation coefficients close to zero.

[14] See eg Glassman (1987).

Models that explain bid-ask spreads in terms of inventory costs establish a link between bid-ask spreads, volatility and trading volumes. One determinant of inventory costs is the cost of maintaining open positions, which is positively related to price risk (Jorion (1996)). According to this view, exchange rate volatility increases price risk and thereby pushes up spreads. Supportive evidence is provided by Bessembinder (1994), Bollerslev and Melvin (1994) and Hartmann (1999), who found a positive correlation between spreads and expected volatility measured by GARCH forecasts.

A second determinant of inventory costs is trading activity. Trading volumes can have a different impact on spreads depending on whether they are expected or unexpected. Expected trading volumes should be negatively correlated with spreads if they reflect economies of scale and are associated with higher competition among market makers (Cornell (1978)). By contrast, unexpected trading volumes should have a positive impact on spreads to the extent that they are associated with higher volatility through the mixture of distribution hypothesis.[15]

To test these assertions, I regressed bid-ask spreads on the GARCH variance forecasts and measures of expected and unexpected trading volumes:

$$(3) \quad S_t = \alpha + \beta_1 h_{t+1} + \beta_2 E_{t-1}(v) + \beta_3 [v_t - E_{t-1(v)}] + \beta_4 w_t + \varepsilon_t$$

The results are presented in Table 19. Consistently with the findings of the literature, the coefficient on the GARCH variance forecast is positive and statistically significant, suggesting that volatility influences bid-ask spreads through its effect on inventory costs. However, in contrast to the predictions of the theory, I do not find positive and significant coefficients on unexpected volumes. The coefficients on expected volumes are also not statistically significant. This latter result is not surprising, since the sample period is likely to be too short to allow for changes in these foreign exchange markets that lead to more efficient trade processing and higher competition among market makers.

5. Conclusions

This paper tried to provide a contribution to the literature on the microstructure of foreign exchange markets by investigating the empirical relationship between trading volumes, volatility and bid-ask spreads. Until now most of the research in this area has relied on data on futures markets, since good data on turnover in foreign exchange markets were not easily available. One important critique of this approach is that volumes in futures markets are not representative of total foreign exchange market activity. This paper uses a new data set that includes daily data on trading volumes for the dollar exchange rates of seven currencies from emerging market countries, the Indonesian rupiah, the Indian rupee, the Mexican peso, the Brazilian real, the Colombian peso, the South African rand and the Israeli shekel. To allow a comparison with other studies, it also looks at trading volumes from the Tokyo interdealer yen/dollar market. The data set covers the sample period from 1 January 1998 to 30 June 1999.

An important result is that unexpected trading volumes and volatility are positively correlated, suggesting that they both respond to the arrival of new information, as the mixture of distributions hypothesis predicts. This is consistent with the findings of the literature that relies on futures data. It suggests that the observation that futures markets are not representative is not “damaging” (Dumas (1996)). The markets for the Mexican peso and the real, however, provide important exceptions. In these two cases, the relationship between unexpected volumes and volatility is negative but not statistically significant. Moreover, for the Mexican peso data from foreign exchange market and from futures market give opposite results, as unexpected futures volumes and volatility are positively correlated.

I also find evidence of a positive correlation between volatility and spreads, as suggested by inventory cost models. This result is also consistent with the findings of the literature. However, in contrast to previous studies I do not find evidence of a significant impact of unexpected trading volumes on spreads.

[15] See Jorion (1996).

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6. Tables and Graphs

Table 1

Exchange rate regimes, 1998–99

Currency	Exchange rate regime
Colombian peso	Managed floating within an intervention band
Mexican peso	Independent float
Rand	Independent float
Real	Managed floating within an adjustable band against the US dollar (mid-1995 to January 1999)
Rupee	Independent float
Rupiah	Free float (since the crisis in July 1997)
Shekel	Managed with respect to a basket of currencies, with margins of fluctuations of +/– 15%.[1]

[1] The shekel is fully convertible since May 1998.

Source: IMF Exchange Rate Arrangements and Restrictions, 1999.

Table 2

Summary statistics for the exchange rates levels

	Mean	Std. Dev.	Min	Max
Colombian peso	1486.19	116.28	1305.83	1752.18
Mexican peso	9.33	0.66	8.03	10.65
Rand	5.73	0.51	4.88	6.62
Real	1.36	0.3	1.12	2.19
Rupee	41.75	1.41	38.3	43.52
Rupiah	9614.66	2193.46	6000	16745
Shekel	3.89	0.24	3.55	4.37
Memo item: Yen/dollar	126.80	9.38	108.80	147.25

Notes: The sample period is 1 Jan. 1998 - 1 July 1999.

Table 3

Summary statistics for the percentage changes of exchange rates

	Mean	Std. Dev.	Test mean=0	Skewness	Kurtosis	Min	Max
Colombian peso	0.09	0.59	2.77 (0.01)	2.42 (0)	21.34 (0)	– 1.9	5.55
Mexican peso	0.04	0.76	1.07 (0.28)	1.1 (0)	6.15 (0)	– 2.94	4.49
Rand	0.06	1.05	1.14 (0.25)	– 1.33 (0)	17.14 (0)	– 8.88	4.58
Real	0.13	1.4	1.78 (0.08)	1.39 (0)	28.15 (0)	– 10.77	11.41
Real (1.1-1.7.1999)	0.32	2.4	1.49 (0.1)	0.58	7.59	– 10.77	11.41
Rupee	0.03	0.29	1.81 (0.07)	– 0.24 (0)	20.3 (0)	– 2.13	2.01
Rupiah	0.14	4.24	0.67 (0.05)	1.05 (0.06)	7.84 (0)	– 18.47	22.6
Shekel	0.04	0.5	1.5 (0.14)	1.08 (0)	7.4 (0)	– 2.25	2.87
Memo item: Yen/dollar	– 0.02	1.02	– 0.29 (0.8)	– 1.00	5.57	– 6.6	3.3

Notes: The sample period is 1 Jan 1998 - 1 July 1999.

Table 4

Sample autocorrelation coefficients

	1	**2**	**5**	**10**	**20**	**60**
Colombian peso						
exchange rate	0.18	0.02	– 0.04	0.04	– 0.01	– 0.05
volume	– 0.25	– 0.06	– 0.1	0.0	0.03	0.08
bid-ask spread	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.
number of deals	– 0.24	– 0.02	0.03	0.0	0.05	0.06
Mexican peso						
exchange rate	– 0.03	0.02	0.02	0.09	0.06	0
volume	– 0.11	– 0.03	0.01	0.01	– 0.04	– 0.01
bid-ask spread	– 0.21	– 0.02	0.01	0.02	– 0.02	– 0.1
Rand						
exchange rate	– 0.003	– 0.04	0.21	0.09	– 0.08	0.009
volume	– 0.18	– 0.24	0.13	0.15	0.09	0.1
bid-ask spread	– 0.03	– 0.004	– 0.01	0.05	– 0.01	– 0.005
Real						
exchange rate	0.15	0.16	– 0.11	0.11	– 0.01	– 0.03
volume	– 0.18	0.01	– 0.09	0.11	0.23	0.04
bid-ask spread	– 0.08	0.39	– 0.05	0.005	– 0.02	– 0.02
Rupee						
exchange rate	0.03	0.00	– 0.02	0.023	0.01	– 0.07
volume	– 0.00	– 0.00	– 0.00	– 0.00	– 0.00	– 0.00
bid-ask spread	– 0.15	0.20	0.05	0.03	– 0.02	– 0.02
Rupiah						
exchange rate	0.16	0.04	0	0.06	0.04	0.01
volume	– 0.4	0.12	0.09	– 0.15	0.04	– 0.01
bid-ask spread	– 0.01	– 0.004	– 0.004	– 0.005	– 0.004	– 0.002
Shekel						
exchange rate	0.12	0.2	– 0.02	0.13	– 0.05	0.03
volume	– 0.15	– 0.01	0.22	0.08	0.02	0.18
bid-ask spread	– 0.38	0	– 0.08	0.09	– 0.04	0.04
Memo item: Yen/dollar						
exchange rate	0.14	0.005	– 0.02	0.03	0	0.05
volume	– 0.29	– 0.06	0.21	0.15	0.06	0.02
bid-ask spread	– 0.37	– 0.08	0.09	0.05	0.07	0.02

Notes: The sample period is 1 Jan 1998 - 1 July 1999. All variables are expressed in percentage changes. Historical volatilities are computed with daily data over rolling windows of 20 business days. Bid-ask spreads are expressed as percentage of the exchange rate.

Table 5

Summary statistics for the 1-month historical volatility

	Mean	Std. Dev.	Min	Max
Colombian peso	7.94	4.66	1.64	21.97
Mexican peso	10.53	5.83	3.28	27.3
Rand	13.11	10.2	2.09	49.2
Real	10.98	18.24	0.56	76.15
Real (1.1.-1.7.1999)	30.54	20.69	6.61	76.15
Rupee	3.73	2.75	0.69	10.66
Rupiah	55.47	38.85	8.35	175.26
Shekel	6.55	4.08	2.09	21.2
Memo item: Yen/dollar	15.33	5.51	7.76	32.46

Notes: Sample period: 1 Jan. 1998 - 1 July 1999. Historical volatilities are computed with daily data over rolling windows of 20 business days.

Table 6

Test for ARCH effects for the percentage change of the exchange rate

Lags	1	2	5	10
Colombian peso	0.00	0.03	0.44	11.77**
Mexican peso	16.25**	21.18**	83.95**	83.8**
Rand	8.1**	9.73**	59.16**	75.56**
Real	15.69**	15.35**	24.33**	40.3**
Rupee				
Rupiah	16.82**	23.89**	67.72**	46.27**
Shekel	25.3**	71.2**	74.13**	75.6**
Memo item: Yen/dollar				

Notes: The table reports results for Engle's (1982) Lagrange multiplier-type test of time-varying heteroschedasticity regression residuals. The test statistics are reported for different lags (1 to 30). ** = the test statistic is significant at the 1% level, suggesting the rejection of the null hypothesis of no ARCH effects.

Table 7
Regressions of volatility on a linear trend

	Constant	Time	Dummy	R^2
Colombian peso	– 0.63* (– 1.94)	0.001** (3.69)	– 0.09 (0.72)	0.01
Mexican peso	0.45 (0.98)	0.00 (0.30)	– 0.06 (– 0.39)	0.01
Rand	2.74* (1.94)	– 0.001 (– 1.34)	– 0.54 (– 1.55)	0.01
Real (1)	– 11.95** (– 4.14)	0.01** (4.02)	– 1.02 (– 1.19)	0.02
Rupee	0.60** (3.00)	– 0.00** (– 2.76)	0.06 (0.89)	0.03
Rupiah	177.22** (4.94)	– 0.16** (– 4.81)	– 4.38 (– 0.86)	0.11
Shekel	0.13 (1.01)	0.0001 (0.90)	0.03 (0.41)	0.01
Memo item: Yen/dollar	1.18* (1.93)	– 0.00 (– 0.15)	– 0.31 (– 1.42)	0.01

Notes: Regressions of volatility, computed as squared returns, on a linear time trend over the sample period 1 Jan 1998 - 1 July 1999. T-statistics are in parentheses. Trading volumes are expressed in US$ millions. (1) Sample period 1.1.1999-1.7.1999.

*,** = significantly different from zero at the 5% and 1% level, respectively.

Table 8

Summary statistics for FX trading volumes

	Mean	Std. Dev.	Min	Max
Colombian peso: Trading volume	165	72	16	362
Colombian peso: Number of trades per day	231	102.7	29	452
Mexican peso	8827	2281	635	15812
Rand	9535	2410	3432	21568
Real	10849	4090	2671	31219
Real (1)	8153	2549	3051	16965
Rupee	3478	1351	1	8211
Rupiah	1072	250	611	1871
Shekel	772	244	5	1698
Memo item: Yen/dollar (3)	12944	4453	1737	41341

Notes: Average daily turnover, in US$ millions. The sample period is Jan 1998 - June 1999. (1) Sample period 1.1.1999-1.7.1999. (2) Number of deals per day. (3) Traded on the Tokyo interbank market. In April 1998, average daily *global* turnover for yen/dollar transactions amounted to US$ 267 billion. Total FX turnover in April 1998 was US$ 1500 billion.

Table 9

Summary statistics for exchange traded volumes

Series	Mean	Std Dev.	Min	Max
Real	16	35	0	277
Real (1)	224	484.27	0	4164
Mexican peso	247	160	0	934
Mexican peso (1)	4632	2968.35	0	17076
Rand	12	22	0	270
Rand (1)	142	275.35	0	3361

Notes: Notional values, in US$ millions. The sample period is January 1998 - June 1999. (1) Number of deals per day.

Table 10

Summary statistics for percentage changes of OTC trading volumes

	Mean	Std. Dev.	Test mean=0	Skewness	Kurtosis	Min	Max
Colombian peso	7.5	45.16	3.17 (0)	2.95	20.17	– 84.02	411.72
Colombian peso (2)	6.97	47.6	2.8 (0)	4.54	38.93	– 81.77	505.71
Mexican peso	12.78	109.09	2.3 (0.02)	10.34	126.58	– 93.04	1542.37
Rand	2.65	26.24	1.6 (0.11)	1.24	4.4	– 64.2	141.19
Real (1)	5.48	38.24	1.6 (0.12)	4.9	38.32	– 73.49	318.59
Rupee							
Rupiah	3.20	27.24	1.2 (0.25)	0.52	-0.13	– 50.27	68.44
Shekel	4.05	30.70	4.6 (0)	7.73	85.99	– 100.00	104.77
Memo item: Yen/dollar (3)	7.98	46.6	3.6 (0)	2.77	13.78	– 78.06	346.88

Notes: The sample period is Jan 1998 - June 1999. (1) Sample period 1.1.1999-1.7.1999. (2) Number of deals per day. Traded on the Tokyo interbank market.

Table 11

Regressions of trading volumes on a linear trend

	Constant	Time	Dummy	R^2
Colombian peso: Trading volume	315.64** (19.43)	– 0.33** (– 9.35)	– 12.70 (– 1.70)	0.27
Colombian peso: Number of trades per day	506.72** (23.41)	– 0.59** (– 13.80)	– 15.93 (– 1.64)	0.44
Mexican peso	11345** (20.94)	– 5.35** (– 4.46)	– 563.28** (– 2.15)	0.07
Rand	12992** (9.25)	– 6.54** (– 2.45)	– 996.28** (– 3.10)	0.07
Real (1)	28454** (16.15)	– 33.83** (– 10.66)	– 223.45 (– 0.47)	0.39
Rupee	6786** (11.79)	– 3.41** (– 6.01)	126.94 (0.90)	0.08
Rupiah	– 296.27 (– 0.54)	2.37** (2.46)	– 3.91 (– 0.07)	0.09
Shekel	127.48** (7.71)	– 0.04 (– 1.25)	17.73* (2.04)	0.02
Memo item: Yen/dollar (3)	18141** (17.21)	– 9.74** (– 4.59)	– 1993** (– 3.46)	0.09

Notes: Regressions of trading volumes on a linear time trend, over the sample period 1 Jan 1998 - 1 July 1999. T-statistics are in parentheses. Trading volumes are expressed in US$ millions. (1) Sample period 1.1.1999-1.7.1999. (2) Number of deals per day. (3) Traded on the Tokyo interbank market. In April 1998, average daily *global* turnover for yen/dollar transactions amounted to US$ 267 billion.

*,** = significantly different from zero at the 5% and 1% level, respectively.

Table 12

Summary statistics for the bid-ask spreads

	Mean	Std. Dev.	Min	Max
Colombian peso	n.a.	n.a.	n.a.	n.a.
Mexican peso	0.12	0.08	0.02	0.94
Rand	0.20	0.63	0.01	12.27
Real	0.16	0.32	0.01	2.68
Real (1.1.-1.7.1999)	0.44	0.44	0.01	2.70
Rupee	0.37	0.23	0.00	2.36
Rupiah	2.11	1.49	0.14	8.70
Shekel	0.29	0.16	0.12	1.18
Memo item: Yen/dollar	0.05	0.02	0.02	0.09

Notes: Sample period: 1 Jan 1998 - 1 July 1999. Bid-ask spreads are expressed as percentage of the exchange rate. The data source is DRI.

Table 13

Summary statistics for the percentage changes of bid-ask spreads

	Mean	Std. Dev.	Test mean=0	Skewness	Kurtosis	Min	Max
Colombian peso	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.
Mexican peso	18.44	87.47	4.15 (0)	4.15	4.15	– 84.9	853
Rand	34.4	401.6	1.68 (0.1)	18.55	357.1	– 97.3	7763
Real	40.4	258.7				– 92.3	3491
Real (1.1.–1.7.1999)	69.57	417.68	1.89 (0.06)	6.71	46.76	– 84.8	3443.98
Rupee	0.03	0.29	1.81 (0.07)	– 0.24 (0)	20.3 (0)	– 2.13	2.01
Rupiah	521.54	7999	1.28 (0.2)	17.5	322.17	– 1056	150148
Shekel	4.07	28.8	2.78 (0)	1	2.31	– 57.3	135.12
Memo item: Yen/dollar	12.9	58.4	4.3 (0)	1.39	2.38	– 70.3	237.1

Notes: Sample period: 1 Jan 1998 - 1 July 1999. Bid-ask spreads are expressed as percentage of the exchange rate.

Table 14

Correlations

	Δxr, volume	**Δ^{+}xr, volume**	**Δ^{-}xr, volume**	**|ΔXR|, volume**	**Volatility, volume**	**Hist. vol., volume**	**Δxr, spread**	**Spread, volume**	**Spread, volatility**
Colombian peso: Trading volume	0.08	0.14	0.05	0.08	0.09	0.01	n.a.	n.a.	n.a.
Colombian peso: Number of trades per day	0.13	0.15	0.03			0.13	n.a.	n.a.	n.a.
Mexican peso	0.06	– 0.10	0.17	– 0.11	– 0.10	– 0.28	0.26	– 0.15	0.64
Rand	0.16	0.27	– 0.11	0.20	0.14	– 0.08	0.15	– 0.05	0.14
Real	– 0.17	– 0.42	0.26	– 0.37	– 0.21	– 0.57	0.31	– 0.47	0.56
Real (1.1.-1.7.1999)	– 0.25	– 0.31	– 0.02	– 0.05	– 0.01	– 0.28	0.32	– 0.30	0.35
Rupee	0.10	0.37	– 0.35	0.38	0.28	0.40	0.13	0.38	0.38
Rupiah	– 0.09	0.17	– 0.30	0.21	0.10	– 0.02	– 0.05	– 0.20	0.30
Shekel	0.04	0.30	– 0.27	0.28	0.25	0.09	0.17	0.16	0.43
Memo item: Yen/dollar	– 0.16	0.47	– 0.45	0.45	0.40	0.15	– 0.09	0.08	0.08

Notes: Correlation coefficients are computed for the period 1 Jan 1998 - 1 July 1999. Exchange rate changes are percentage changes; volumes are expressed in logarithms; bid-ask spreads are expressed as a fraction of the spot rate; volatility is computed as squared returns; historical volatility is computed as standard deviations of daily percentage exchange rate changes over rolling windows of 20 business days.

Table 15

Correlations for exchange traded volumes

	Δxr, volume	**Δ^{+}xr, volume**	**Δ^{-}xr, volume**	**|ΔXR|, volume**	**Volatility, volume**	**Hist. vol., volume**	**Spread, volume**
Mexican peso	0.09	0.28	-0.11	0.22	0.16	0.04	0.07

Notes: Correlation coefficients are computed for the period 1 Jan 1998 - 1 July 1999. Volumes refer to notional amounts in US$ millions. The other variables are defined as in Table 14.

Table 16

Volatility and trading volume: unconditional regressions

	Constant	Time trend	Volume	R^2
Colombian peso	– 3.86** (– 2.88)	0.0022** (4.16)	0.41* (2.14)	0.02
Colombian peso (2)	– 1.62 (– 0.83)	0.0008 (0.42)	0.23* (2.46)	0.04
Mexican peso	2.89 (1.54)	0.0002 (0.44)	– 0.28 (– 1.58)	0.00
Rand	– 2.11 (– 0.36)	– 0.02** (– 2.75)	2.17* (2.30)	0.06
Real(1)	121.19 (1.75)	– 0.02 (– 1.05)	– 10.80* (– 1.92)	0.07
Rupee	0.21 (1.24)	– 0.001** (– 2.65)	0.05** (2.87)	0.04
Rupiah	20.10 (0.76)	– 0.05 (– 1.80)	5.01** (3.49)	0.07
Shekel	– 1.06* (– 2.68)	0.0002* (2.06)	0.23** (2.75)	0.04
Memo item: Yen/dollar (3)	– 32.18** (– 2.98)	0.0023* (2.45)	3.26** (3.03)	0.12

Notes: Regressions of exchange rate volatility, defined as squared returns, on a constant, a linear time trend and trading volumes (in logarithms) over the sample period 1 Jan 1998 - 1 July 1999. T-statistics based on White's asymptotically consistent standard errors are in parentheses. Coefficients on trading volumes are multiplied by 100. (1) Sample period 1.1.1999-1.7.1999. (2) Number of deals per day. (3) Traded on the Tokyo interbank market.

*,** = significantly different from zero at the 5% and 1% level, respectively.

Table 17

Trading volumes and volatility

	Constant	E(v)	v-E(v)	R^2
Colombian peso	– 4.04 – (2.47)	0.44 (1.91)	0.34* (2.27)	0.02
Mexican peso	3.09 (0.82)	– 0.28 (– 0.78)	– 0.30 (– 1.50)	0.00
Rand	7.50 (0.66)	1.31 (0.98)	3.08* (2.47)	0.07
Real (1)	69.93 (1.04)	14.28 (1.26)	– 2.83 (– 0.62)	0.15
Rupee	– 1.70** (– 2.74)	0.29** (3.16)	0.04** (2.91)	0.05
Rupiah	17.41 (0.41)	5.32 (1.19)	4.97** (2.96)	0.07
Shekel	– 2.63** (– 2.87)	0.56** (2.91)	0.18** (2.65)	0.05
Memo item: Yen/dollar (2)	– 27.27* (– 2.04)	2.78* (2.03)	3.20** (3.06)	0.11

Notes: Regressions of total volatility on a constant, a linear time trend, expected volumes and unexpected volumes. The sample period is 1 Jan 1998 - 1 July 1999. Total volatility is defined as squared returns; log-volumes are decomposed into an expected and unexpected component by using a fitted AR series and its residuals. T-statistics based on White's asymptotically consistent standard errors are in parentheses. (1) Sample period 1.1.1999-1.7.1999. (2) Traded on the Tokyo interbank market.

*,** = significantly different from zero at the 5% and 1% level, respectively.

Table 18

Trading volumes and volatility

	Constant	GARCH	E(v)	v-E(v)	R^2
Colombian peso	– 3.53 (– 1.85)	0.01 (1.03)	0.39 (1.50)	0.34* (2.22)	0.02
Mexican peso	– 0.03 (– 0.01)	0.08** (2.89)	– 0.64 (– 0.18)	– 0.28 (– 1.50)	0.09
Rand	7.19 (0.62)	0.09 (1.66)	0.14 (0.10)	3.04* (2.27)	0.10
Real (1)	– 83.43 (– 0.69)	0.23 (1.73)	20.27 (1.57)	– 1.69 (– 0.36)	0.20
Rupee	– 1.47* (– 2.22)	0.01 (1.06)	0.24* (2.47)	0.04* (2.53)	0.05
Rupiah	17.09 (0.41)	– 0.00 (– 0.08)	5.16 (1.02)	5.00** (2.76)	0.07
Shekel	– 0.90 (– 1.44)	0.08** (3.52)	0.14 (1.02)	0.08 (1.51)	0.19
Memo item: Yen/dollar (2)	– 14.50 (– 1.47)	0.11 (1.90)	1.31 (1.31)	2.93** (3.23)	0.13

Notes: Regressions of total volatility on a constant, a linear time trend, expected volatility, expected volumes and unexpected volumes. The sample period is 1 Jan 1998 - 1 July 1999. Total volatility is defined as squared returns; expected volatility is the one-step-ahead conditional return variance from a GARCH(1,1) specification; log-volumes are decomposed into an expected and unexpected component by using a fitted AR series and its residuals. T-statistics based on White's asymptotically consistent standard errors are in parentheses. (1) Sample period 1.1.1999-1.7.1999. (2) Traded on the Tokyo interbank market.

*,** = significantly different from zero at the 5% and 1% level, respectively.

Table 19

Spreads, trading volumes and volatility

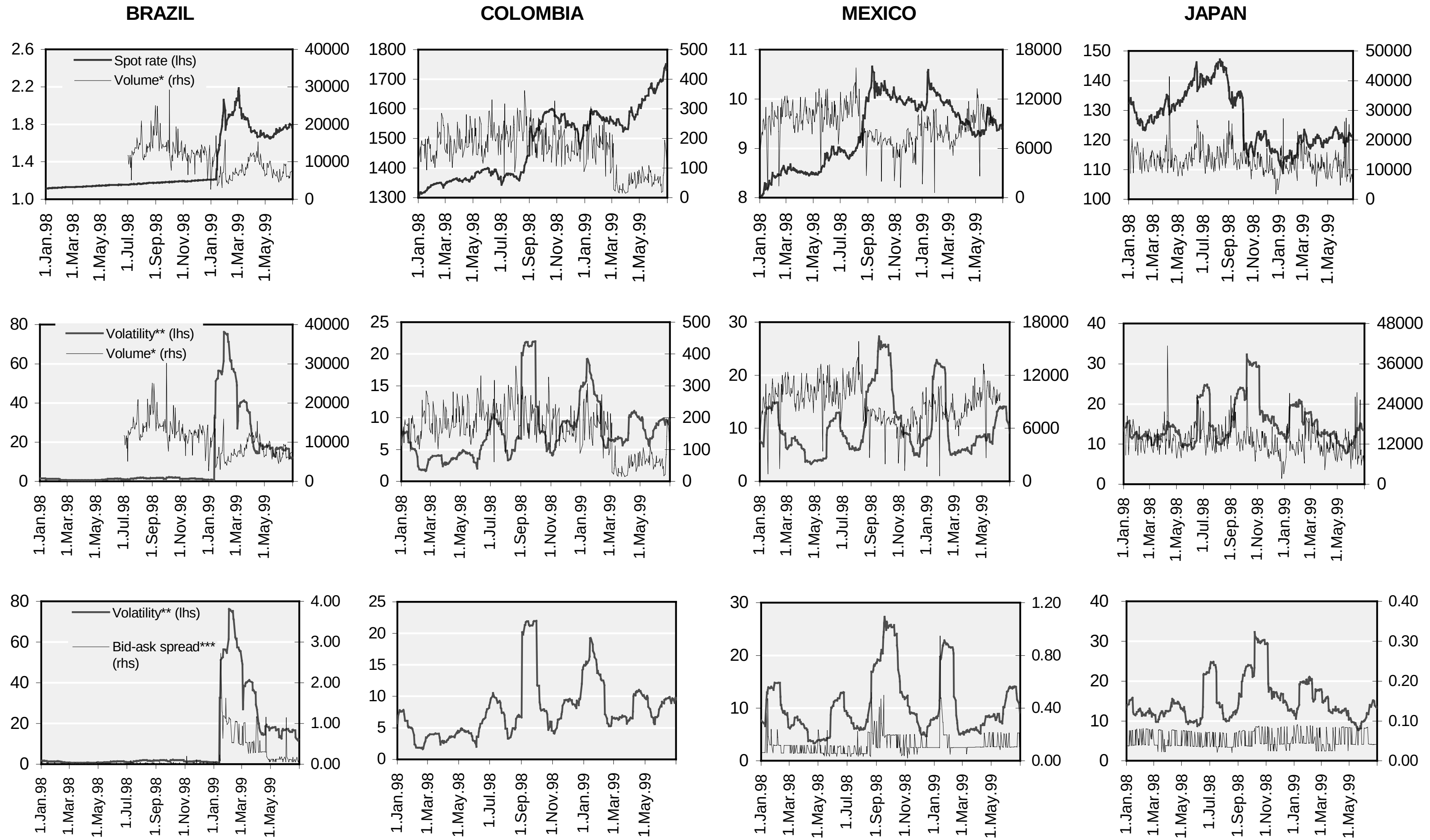
	Constant	GARCH	E(v)	v-E(v)	R^2
Colombian peso	n.a.	n.a.	n.a.	n.a.	n.a.
Mexican peso	0.47** (2.83)	0.01** (5.32)	– 0.05 (– 2.65)	– 0.01 (– 1.21)	0.20
Rand	0.96 (1.42)	0.01 (1.50)	– 0.10 (– 1.45)	– 0.16 (– 1.09)	0.05
Real(1)	– 0.10 (– 0.06)	0.01** (7.31)	0.02 (0.13)	– 0.26 (– 0.99)	0.47
Rupee	0.15 (0.22)	0.03** (6.81)	0.01 (0.16)	0.02 (1.61)	0.22
Rupiah	7.77 (1.73)	0.01* (2.28)	– 1.01 (– 1.55)	– 0.44 (– 1.74)	0.10
Shekel	0.23* (2.13)	0.03** (14.40)	– 0.03 (– 1.42)	– 0.01 (– 0.95)	0.62
Memo item: yen/dollar (2)	0.08 (1.18)	0.00 (1.25)	– 0.00 (– 0.48)	0.00 (0.54)	0.02

Notes: Regressions of bid-ask spreads on expected volatility, expected volumes and unexpected volumes estimated over the sample period 1 Jan 1998 - 1 July 1999. Total volatility is defined as squared returns; expected volatility is the one-step-ahead conditional return variance from a GARCH(1,1) specification; volumes are decomposed into an expected and unexpected component by using a fitted AR(1) series and its residuals. T-statistics based on White's asymptotically consistent standard errors are in parentheses. (1) Sample period 1.1.1999-1.7.1999. (2) Traded on the Tokyo interbank market.

*,** = significantly different from zero at the 5% and 1% level, respectively.

Graph 1

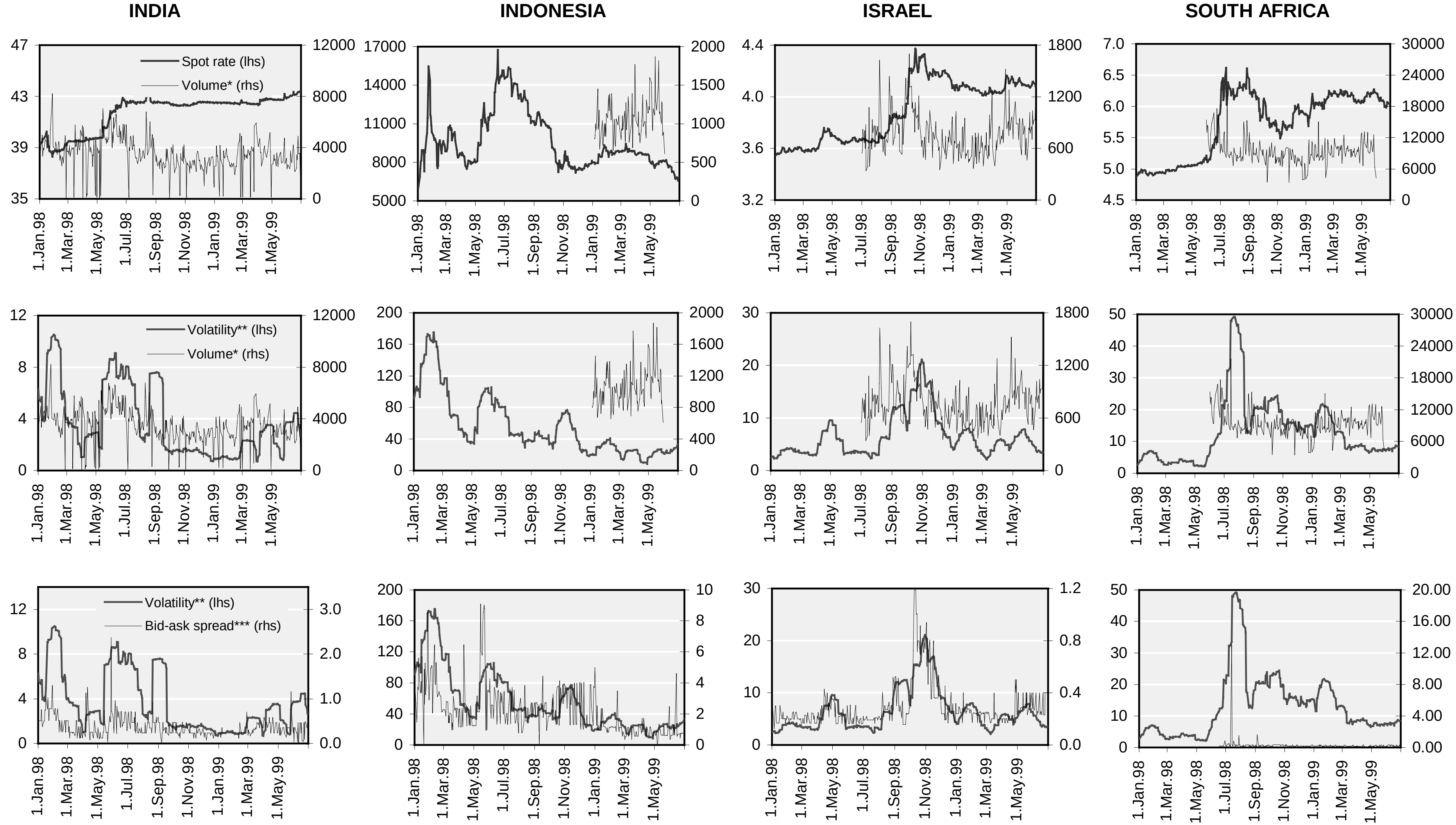
Trading volumes, volatility and spreads



* Expressed in million US$. ** 1-month historical volatility. *** As a percentage of the mid quote.

Graph 1 (continued)

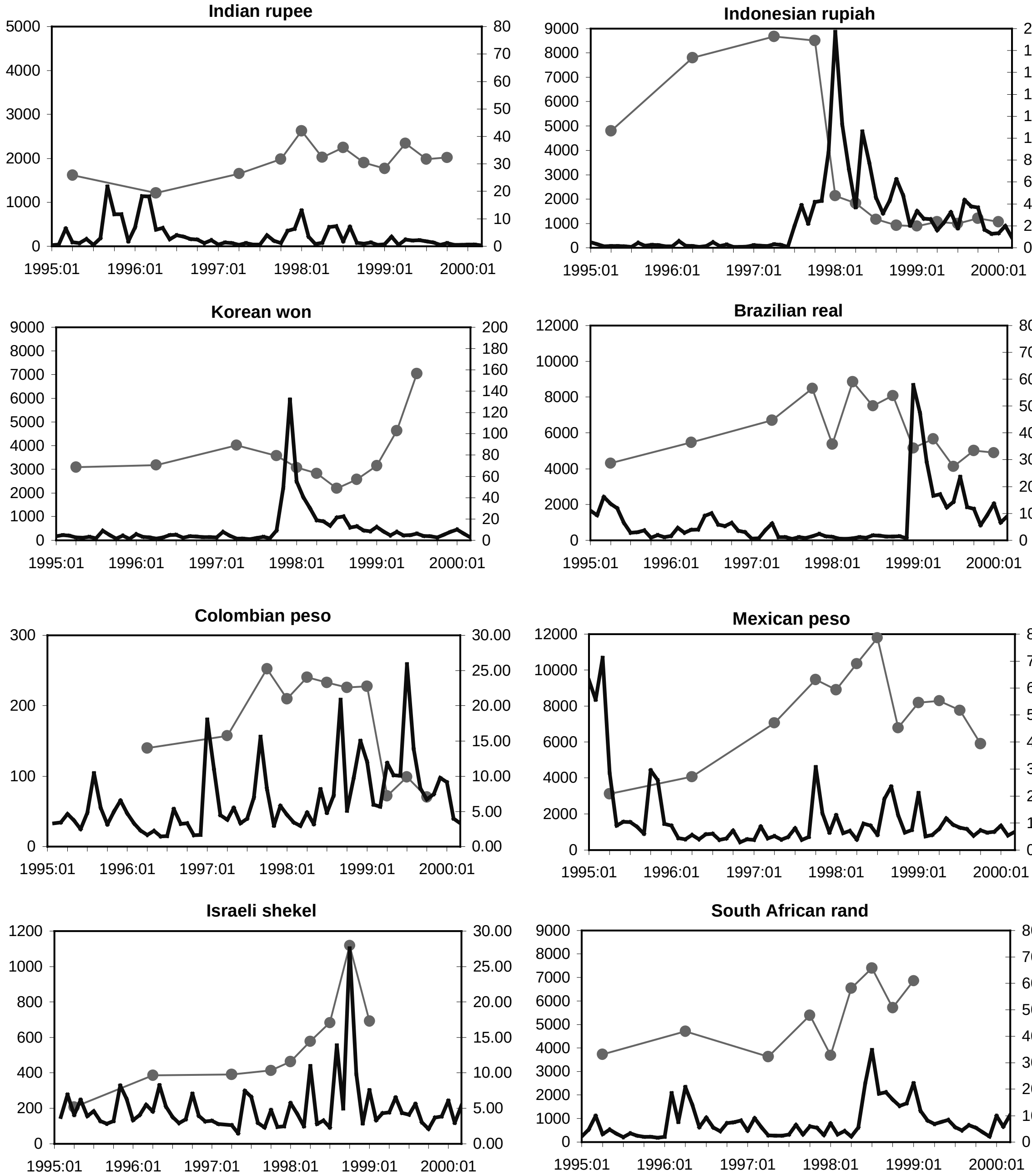
Trading volumes, volatility and spreads



* Expressed in million US$. ** 1-month historical volatility. *** As a percentage of the mid quote.

Graph 2

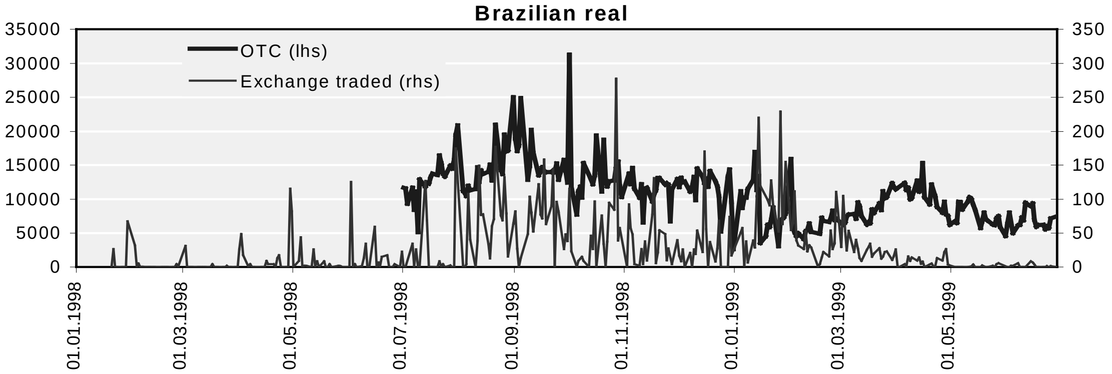
Trading volumes and volatility in selected emerging markets

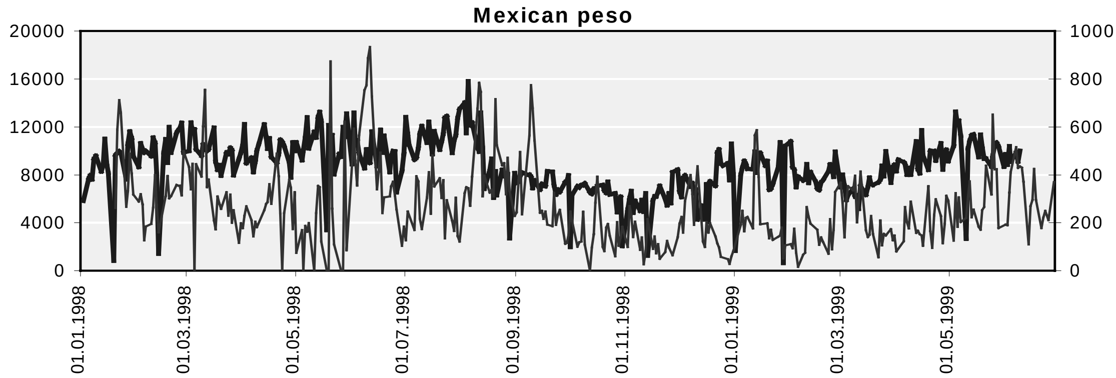


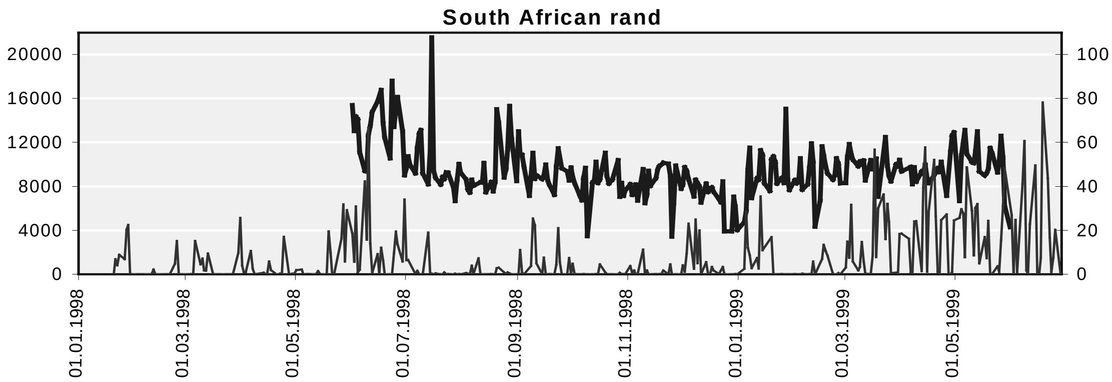
Notes: Trading volumes refer to local turnover in the domestic currency, per trading day in the month shown (in millions of US$). In the case of Mexican and Brazil, turnover includes other currencies. Volatility is computed as one-month annualised standard deviation of percentage changes in the exchange rate against the US dollar.

Graph 3

OTC and exchange traded foreign exchange market turnover







Notes: OTC turnover is defined as in Graph 1. Exchange traded turnover refers to notional values of futures contracts transacted on the Chicago Mercantile Exchange.

Comments on: "Volumes, volatility and spreads in FX markets: evidence from emerging market countries" by Gabriele Galati

Javiera Ragnartz, Sveriges Riksbank

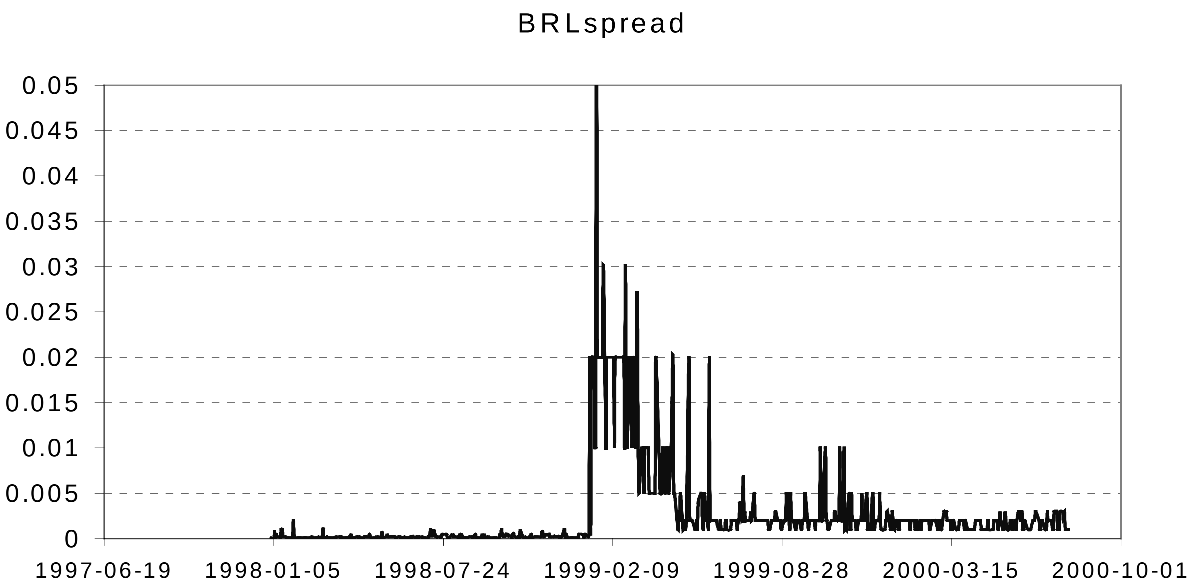
First of all I would like to thank the BIS for arranging this conference and for giving me the opportunity of discussing this great paper. I think the paper covers an area that is interesting for central bankers that try to understand the behaviour and development of exchange rates. The paper contributes to a better understanding of the structure of the currency market.

The paper investigates the relationship between trading volumes, volatility and bid-ask spreads in the foreign exchange markets of seven emerging market exchange rates against the dollar. The yen/dollar exchange rate is also used for comparison purposes. The sample covers June 1998 and the entire year of 1999. Unexpected trading volumes and volatility are found to be positively correlated, with the Mexican peso and the Brazilian real being exceptions. There is also evidence that volatility and bid-ask spreads are positive correlated. However, there is no evidence of a significant impact of trading volumes on spreads. My comments are mainly questions and suggestions to future developments.

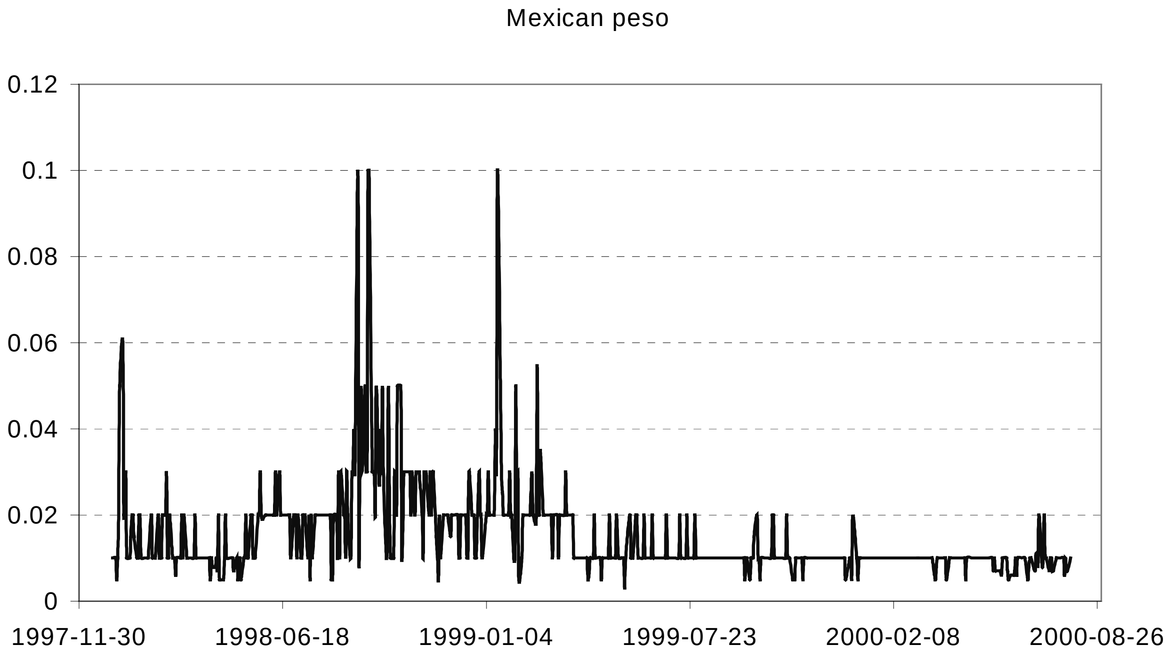
First of all I would also like to ask the author why he thinks the Brazilian real and the Mexican Peso are exemptions. Could it be the case that these exchange rates are more effective than the other markets since these are larger and more heavily traded?

Secondly, I think that the sample period might be too short and more importantly it covers a period characterised by financial stress. The graphs 1a and 1b show that the bid-ask spreads of the Brazilian real and the Mexican peso increased significantly during the fall of 98, a time of high volatility in the currency markets in general following the Russian crisis. However, these spreads decreased significantly in late 99 and have since remained quite stable at a lower level. Thus, there is a possibility that the results might be altered if a longer sample including year 2000 is used.

1a. Bid-ask spreads for the Brazilian real, 1998-01-01—2000-07-30



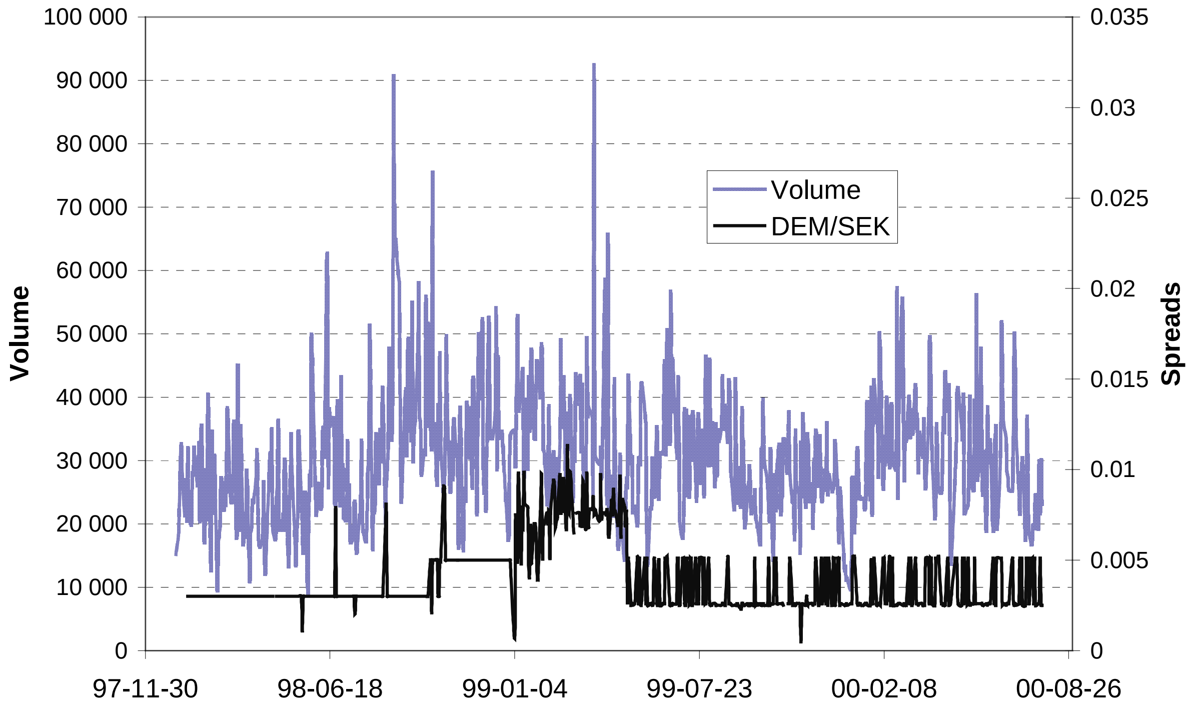
1b. Bid-ask spreads for the Mexican peso, 1998-01-01—2000-07-30



Gabriele Galati uses the Yen/dollar exchange rate in order to compare the results with those of an effective and more liquid market. But the results for the Yen/dollar tend to differ somewhat form the other currency pairs. For instance table 14 in the paper shows that the coefficients tend to be higher for this exchange rate than those for the emerging markets exchange rates. Table 16 shows that the Yen/dollar exchange rate regression tends to have a significantly higher R^2- value. What does the author think is the reason to this? Could the answer be as simple as the Yen/dollar market being more effective?

Finally, I would like to take the opportunity and promote research on the Swedish krona. The Swedish krona behaved quite similar to other emerging market currency during the Asian crisis of 1997 and the following Russian crisis during 1998. It appears that market participants tend to leave the Swedish market during times of financial stress independent if Sweden is affected by the crisis or not. Since the Swedish krona behaves this way it would be nice to see if the same results would be found for the Swedish currency. Graph 2 shows volume and bid-ask spreads on the SEK/DEM exchange rate. The turnover data is not completely clean in the sense that it could include other exchange rates, but should be approximately the SEK/DEM turnover. The graphs show that even though the volume decreased in late 1999 it has once again increased lately. However spreads appears to have been quite constant since the end of 1999.

2. DEM/SEK spreads and volume, 1998-01-01—00-07-30



Comments on "Trading volume, volatility and spreads in FX markets: evidence from emerging market countries" by Gabriele Galati

Alain Chaboud, Federal Reserve System

The nature of the interaction between volume, volatility and the bid-ask spread in foreign exchange markets is still poorly understood, and Gabriele Galati's paper makes excellent use of a new set of transaction volume data to shed light on the subject. The paper provides a test of what has become known as the "mixture of distributions hypothesis," which predicts that trading volume and volatility, simultaneously driven by the arrival of information, should be positively correlated. Using data for seven currencies of large emerging market economies, the paper offers support for that prediction, with the important exceptions of the Brazilian real and the Mexican peso. It also finds a positive correlation between volatility and bid-offer spreads, in line with inventory cost models. I offer a few remarks on the analysis of the link between volume and volatility, and briefly touch on the subject of foreign exchange turnover data.

The study of the correlation between volatility and volume in this paper is rendered more interesting, but perhaps also more difficult, by the sample period of the study, 1998 and 1999, during which the financial markets of many emerging economies experienced turmoil. In that light, the failure of the Brazilian and Mexican currency markets to show a positive correlation between volume and volatility is not altogether a surprise. It may not be much of a stretch to think that volatility and volume could exhibit a positive relation only up to a certain level of volatility, with the correlation becoming negative as volatility grows so large that many investors withdraw from the particular market. The predictions of the mixture of distributions hypothesis model would then be realised in "normal" market conditions, but not in turbulent market conditions. This (unproven) scenario may be easier to swallow for the case of emerging market currencies than for major currencies, but the sharp movement of the dollar/yen exchange rate in October 1998, reportedly accompanied by (likely in part caused by) thin market conditions, is one recent example of this type of phenomenon occurring in a major market. One could perhaps then interpret positive correlation between volume and volatility as a sign of a "healthy" or liquid markets, while the presence of negative correlation between volume and volatility could be taken as a symptom of inadequate liquidity. If such a non-linear relationship exists between volatility and volume, simple linear regression techniques may not be fully adequate for proper inference, and, given the interesting results in the present paper, this may constitute a good opportunity for further research.

As is commonly done, the volatility measure used in this paper is based on net daily changes in exchange rates, matching the daily frequency of the turnover series. While this may appear to be a virtue, it would perhaps be preferable to use a volatility measure that better reflects the cumulative amount of price movement over the trading day and not just the net movement. This would be more in the spirit of the mixture of distributions hypothesis, where the arrival of each new piece of information results in both trading activity and movement in the price. With the daily turnover data representing the sum of all trading activity, matching them with a cumulative measure of price movement would avoid a case where, for instance, two salient pieces of news on the same day moved the exchange rate sharply but in opposite directions, yielding both a high trading volume and a misleadingly low measure of volatility. As an alternative measure of volatility, one could use the daily sum of absolute five - minute exchange rate movements, or some other estimate of integrated volatility. If high-frequency price data are unavailable, using the daily high-low range instead of the daily return as a measure of volatility has been shown to be a step in the right direction.

In a very interesting aside, the paper reports extremely low correlation coefficients between turnover in three of the domestic spot markets in the study and corresponding futures markets turnover from the Chicago Mercantile Exchange. Though the correlation coefficients may be affected by the very low liquidity of those particular futures contracts, this finding still raises a (large red) flag, as currency futures turnover data have been widely used as a proxy for spot turnover data. For spot currency markets, over-the-counter and highly decentralised, transaction volume data for most currency pairs have generally been very difficult if not impossible to obtain. Several researchers have displayed great ingenuity and patience in assembling data sets such as the one in this study, but all these data have been limited in time and scope. The widespread use of electronic trading for most interdealer transactions in recent years carries with it hope that comprehensive high-frequency transaction data for most major currency pairs may soon be available for research.

Sending the herd off the cliff edge: the disturbing interaction between herding and market-sensitive risk management practices[1]

Avinash Persaud, State Street

Summary

In the international financial arena, G7 policymakers chant three things: more market-sensitive risk management, stronger prudential standards and improved transparency. The message is that we do not need a new world order, but we can improve the workings of the existing one. While many believe this is an inadequate response to the financial crises of the last two decades, few argue against risk management, prudence and transparency. Perhaps we should, especially with regards to market-sensitive risk management and transparency. The underlying idea behind this holy trinity is that it better equips markets to reward good behaviour and penalise the bad, across governments and market players. However, while the market is discerning in the long-run, there is now compelling evidence that in the short run, market participants find it hard to distinguish between the good and the unsustainable; they herd; and contagion is common.

Critically, in a herding environment, tighter market-sensitive risk management systems and more transparency actually makes markets less stable and more prone to crisis. This perverse response may help to explain the growing instability of the financial system. Over the past ten years the system has been in crisis in almost four. In response, the emphasis on market sensitive risk management practices and transparency must be tempered rather than extended vis-a-vis credit ratings agencies and the risk management of investors. The incentives of regulators should be realigned to include the impact of herding behaviour on foreign markets as well as local institutions. Financial market institutions should be encouraged to set aside collateral for systemic risk or to buy liquidity options from central banks during good times. Countries should facilitate the inflow of capital from investors who herd least or leave slowly: foreign direct investors, equity portfolio investors and, surprisingly, hedge funds.

A cyclical debate

The debate on reform of the international financial system follows a cycle. In the middle of each crisis - and there have been at least six since the Debt Crisis which started in Mexico in 1982 - there are deafening demands for the whole-scale reform of the entire international financial system. A few months on from the end of each crisis and those demands fade. There were clear and present parallels between calls made in previous crises and those made in the thick of the last crisis, for the IMF to become a lender of last resort, injecting substantial liquidity in times of crisis, and for hedge funds to be regulated. Every crisis inspires plans for a new financial architecture and as the crisis ends, most of these plans are tidied away.

1 This essay won first prize in the Institute of International Finance's Jacques de Larosiere Awards in Global Finance 2000.

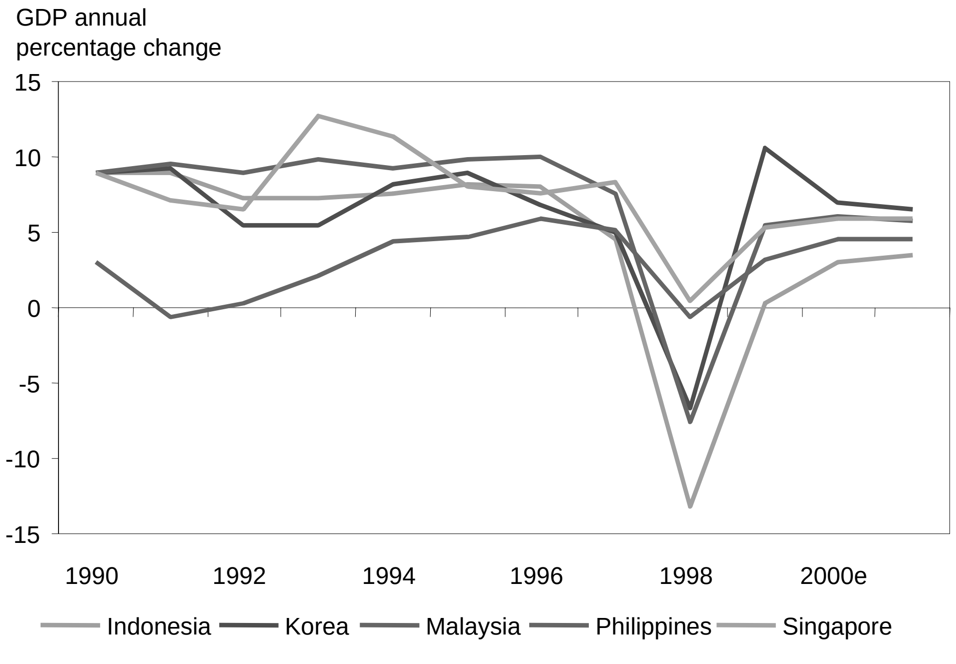
Table 1
Global financial crises in the 1990s

Date	Crisis	Countries where the real exchange rate fell by more than 10% over one month
1992-1993	"EMS"	UK, Italy, Spain, Portugal, Sweden, Finland, Denmark, Norway, Belgium, France, Ireland, India, Venezuela
1994-1995	"Tequila"	Columbia, Venezuela, Mexico, Turkey, Japan
1997-1999	"Asia"	Thailand, Philippines, Indonesia, Malaysia, Taiwan, Korea, Brazil Columbia, Israel, Peru, South Africa, Zimbabwe, Russia, Sweden, Switzerland, Spain

Source: State Street Bank.

Underlying this cycle of debate is that while the demand to make systemic changes is naturally strong in the middle of a crisis, the consensus on what is wrong and what to do is generally weak. Moreover, while recent crises have appeared sharper and more global than before, they have been shorter-lived. Before a consensus on what to do to avoid crises can grow, they are over, and countries previously in crisis begin to enjoy economic rebound and the return of international capital flows. This was not the case during the Latin-American Debt Crisis of the mid-1980s or after the EMS crisis in 1992-93 when economic recovery was held back by a cheap dollar and European governments exerting fiscal restraint. But it was the case in the last two crises in Mexico and Asia, see Chart 1. We also live in an age where ambitions are limited. We no longer walk on the moon. In this environment, the view that often gains ground a few months after the crisis is that there are risks to meddling with a financial system that works most of the time and there are things that can be safely done to improve the workings of the market the rest of the time.

Chart 1
The rapid rebound in Asian GDP



Source: State Street Bank.

The proposals that emerge post-crisis, therefore, tend to focus on making it easier for the market to reward good behaviour and penalise the bad. The emphasis is not on changing the rules of the game, but on strengthening the players: stronger risk management, more prudential standards and improved transparency. One of the key responses of the Interim Committee of the IMF to the latest crisis and the desire to avoid a next one was the adoption on 26 September 1999 of a new Code of Good Practices on Transparency in Monetary and Financial Policies. Incidentally, these measures are all relatively inexpensive to implement. There is declining political support for large packages of tax-payers money to bail out foreign countries in trouble.

How more market-sensitive risk management can create risk

While many believe that market sensitive risk management, prudential standards and transparency are probably not enough to avoid future crises, they believe these measures will probably help to provide the right discipline for governments and can surely do no harm. These measures are likely to be a positive force in the long run when markets are better at discerning between the good and bad. But in the short-run, there is growing evidence that market participants find it hard to discern between the good and the unsustainable, they often herd and contagion from one crisis to another is common. The problem is that in a world of "herding", tighter market-sensitive risk management regulations and improved transparency can, perversely, turn events from bad to worse, creating volatility, reducing diversification and triggering contagion. How can this happen?

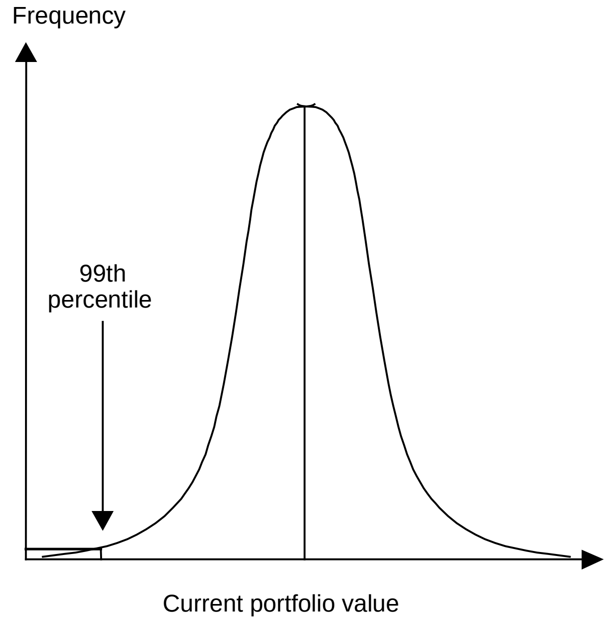
Let us explore the interaction between herding, market-sensitive risk management and transparency in bank lending. It is important to note that bank lending remains a powerful feature of modern-day crises. For example, the five Asian crisis countries - Thailand, Malaysia, South Korea, Indonesia and the Philippines - received USD 47.8 billion in foreign bank loans in 1996. In 1997, banks withdrew USD 29.9 billion, a net turnaround of almost USD 80 billion in one year. By contrast, equity portfolio flows remained positive throughout 1997, see "Portfolio Flow Indicator - Technical Document, State Street Bank and FDO Partners, 1999".

The growing fashion in risk management, supported by the Basel Committee on Banking Supervision, is a move away from discretionary judgements about risk and a move to more quantitative and market-sensitive approaches (for an early reference see, the Supervisory Treatment of Market Risks, Basel Committee on Banking Supervision, 1993). This is well illustrated by how banks now tend to manage market risks by setting a DEAR limit - daily earnings at risk. DEAR answers the question: how much can I lose with, say, a 1% probability over the next day. It is calculated by taking the bank's portfolio of positions and estimating the future distribution of daily returns based on past measures of market correlation and volatility. Both rising volatility and rising correlation will increase the potential loss of the portfolio, increasing DEAR. Falling volatility and correlation will do the opposite. Banks set a DEAR limit: the maximum dollar amount they are prepared to put at risk of losing with a 1% probability. When DEAR exceeds the limit, the bank reduces exposure, often by switching into less volatile and less correlated assets such as US dollar cash. (See RiskMetrics Technical Manual, RiskMetrics Group, London, 1999.)

By herding behaviour I mean that banks or investors like to buy what others are buying, sell what others are selling and own what others own. There are three main explanations for why bankers and investors herd. First, in a world of uncertainty, the best way of exploiting the information of others is by copying what they are doing. Second, bankers and investors are often measured and rewarded by relative performance so it literally does not pay a risk-averse player to stray too far from the pack. Third, investors and bankers are more likely to be sacked for being wrong and alone than being wrong and in company. (For further explanations of herding see Investor Behaviour in the October 1987 Stock Market Crash: Survey Evidence by R Shiller, NBER discussion paper 2446, 1990.)

Figure 1

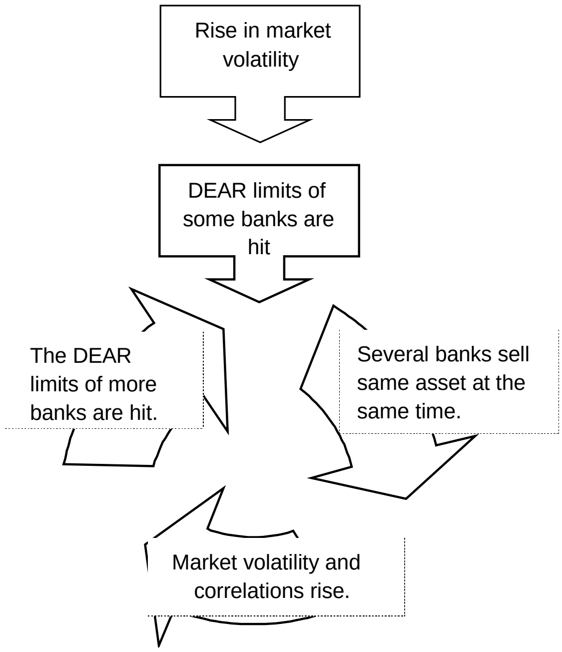
Representation of VAR: histogram of portfolio values



Imagine that over time a herd of banks have acquired stocks in two risky assets that have few fundamental connections, say, Korean property and UK technology stocks. Imagine too that some bad news causes volatility in UK technology stocks and the banks most heavily invested there find that their DEAR limits are hit. As these banks try and reduce their DEAR by selling the same stocks (Korean property and UK technology) at the very same time, there are dramatic declines in prices, rises in volatility in both markets and in the correlation between Korean and UK markets. Rising volatility and correlation triggers the DEAR limits of banks less heavily invested in these markets but invested in other markets. As they join the selling milieu, volatility, correlation and contagion rises.

The key to this environment is that market participants behave strategically in relation to one another, but DEAR measures risk "statically" - without strategic considerations. Previous volatility and correlations were measured over a period of time when the herd gradually built up and are therefore almost certain to underestimate the impact on prices, volatility and correlations when many investors sell the same asset at the same time. This strategic behaviour can be modeled more formally using game theory. (Some attempts to do so can be found in "Risk management with interdependent choice" by Stephen Morris and Hyun Song Shin, Oxford Review of Economic Policy, Autumn 1999.)

Figure 2
A vicious cycle of herding and DEAR limits



Stress-testing

It has been suggested that stress-testing could avoid this contagion - testing how a portfolio of positions perform under made-up scenarios. In practice it does not do so for two reasons. First, the most popular stress test is to see what would happen to a portfolio of positions if a past crisis was repeated. But this is not very meaningful. As we have observed above the spread and focus of crises relate to where positions are and unless positions are identical - which is unlikely given the memory of the past crisis - the next crisis will be different from the past. The best stress test is to assume that everybody has the same positions as you do and you cannot get out of any or them without large losses. This test is seldom attempted. It is hard to estimate the spread and depth of positions or the impact on liquidity and hence potential losses. In this age of quant, risk managers mistakenly prefer to worry about quantifiable risks than unquantifiable ones. Even if the risks could be estimated, banks would treat the result with suspicion. It would be like telling a lending institution that when assessing risk, assume none of the loans are repaid and the historical volatilities and correlation suggest it is a far-fetched scenario, though it is exactly the scenario that Long-Term Capital, the failed hedge fund, and others found themselves in during September and October of 1998.

Several financial institutions suffered serious losses in 1998, but few of these were life threatening. An additional critical point is that even if stress testing worked to save banks from trouble, it may not save a country. One of the interesting aspects of the Asian crisis was that while short-term external debt exposures for Asian countries were large enough to "bring down" countries under the dynamic of herding we have discussed above, the exposures were not big enough to bring down foreign banks. One of the key challenges is to realign the incentives of regulators to worry about the concentration of exposures in foreign countries as well as in local banks.

Let us add another dimension to our nightmare scenario. Further assume that a country has recently signed up to the Special Data Dissemination Standard (SDDS) - one of the lasting responses of the 1995 Tequila crisis - and the 1999 Code of Good Practice, and as a result, has started publishing its foreign exchange reserves daily. In this case bankers and investors with more modest exposures would observe that as risks grow - prices are falling and volatility rising - other bankers and investors are leaving the country rapidly. In this heightened environment they will view the country's loss of reserves as doubly increasing the risk that they are left wrong and alone. This will trigger a further rush for the exit.

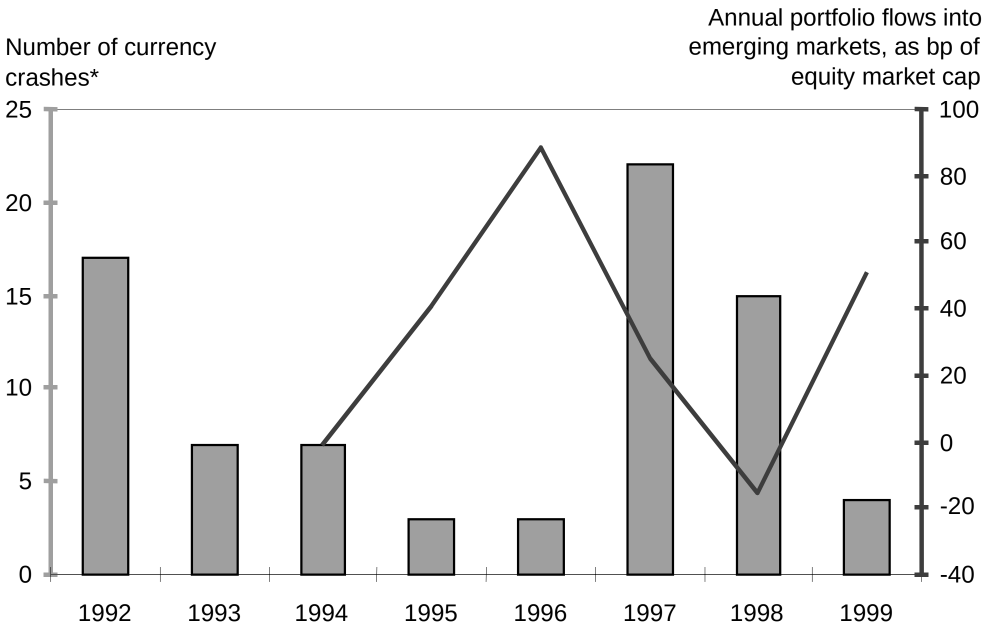
The reason why this is a major challenge to the current regulatory framework is that herding is frequent and that even short-lived financial crises have real economic impact. While herding behaviour is hard to prove directly given the paucity of reliable data on the positions of financial institutions, there is a now a growing body of evidence that markets behave as if market participants herd.

In the foreign exchange markets for example, if we define a crash as a 10% fall in the real exchange rate over three months, there have been 78 crashes across 72 countries since the EMS crisis began in September 1992. These are not distributed evenly over time or distributed with deteriorating fundamentals, but they cluster. Contagion is rife with 70% of crashes occurring in just three years. This contagion does not move predictably along the lines of trade, but along the lines of shared investors. The stepping-stones of the most recent crisis, for example, were from Thailand and Indonesia to Korea, on to Russia and then to Brazil. These countries share very little trade in common. Furthermore, crashes are invariably preceded by booms as the herd moves into place. Chart 2 shows the number of foreign exchange crashes per year across 72 countries as bars and the annual cross-border portfolio flows into emerging markets as a line. Note how investors rushed into emerging markets in 1995 and 1996, prior to the crashes in 1997 and 1998.

Further evidence of herding and the problems of a static value-at-risk analysis can be found by looking at the distribution of daily market returns. In Chart 3, we imagine we are a risk manager in January 1997 looking at the distribution of daily returns of a portfolio of OECD currencies versus the dollar over the previous five years. The distribution is well behaved and fairly symmetrical - though not around zero. According to this actual distribution she would expect a more than 1% decline in this portfolio's value in a day around 5% of the time. Three years later and if she survived, she would have found that her portfolio fell by more than 1% in a day more than 10% of the time and the distribution of returns looked very different (Chart 4). (It can be shown that the difference between these two distributions follows a beta distribution consistent with herding behaviour.)

Chart 2

"Crashes" and "booms" in the foreign exchange market

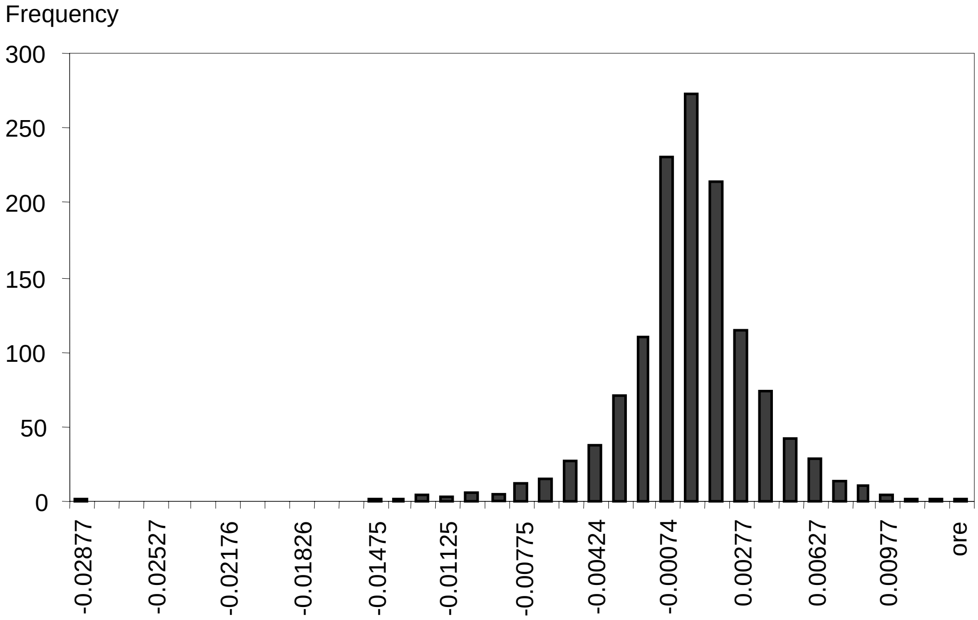


Source: State Street Bank.

The predominance of herding behaviour and its lethal combination with the practice of DEAR limits may explain why the 1990s have been such a decade of financial dislocation: the financial system has been in crisis for 40 out of the 120 months or 33% of the time. This instability has real economic impact. Although international portfolio flows have recovered from dips in 1998, they remain highly concentrated in just five markets: Hong Kong, Korea, Singapore and Taiwan - hardly the most capital needy countries either given their high domestic savings and big current account surpluses. Many other markets have found it hard to raise foreign capital.

Chart 3

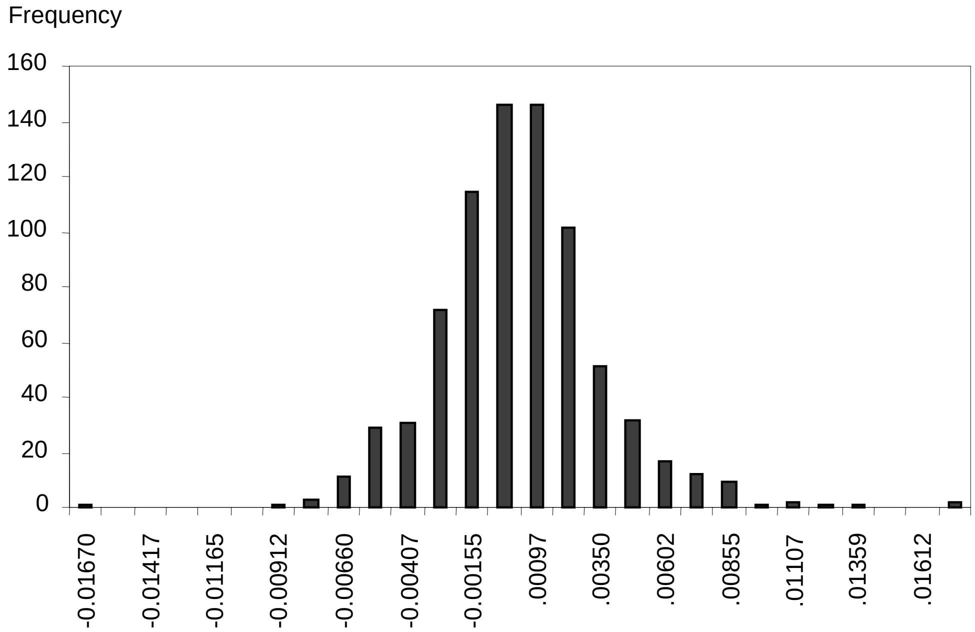
Distribution of average daily dollar returns of an OECD less US portfolio of currencies, 1992-96



These financial crises also have a direct impact on GDP. For example, while there has been a strong rebound in GDP in 1999 in Asia in general and in South Korea in particular, the rebound has not offset the loss of GDP during the crisis period. One way of estimating the lost GDP of the Asian crisis is to estimate where GDP would be today if Asian economies had continued the more modest but sustainable growth rates experienced in the five years before their current account deficits began to widen in 1993-94. Were it not for the crisis and its preceding boom, GDP would be an aggregate of USD 130 billion higher in South Korea, Thailand, Malaysia and Indonesia. Another measure of this lasting impact is the elevation of poverty levels in Asia today compared with 1997.

Chart 4

Distribution of average daily dollar returns of an OECD less US portfolio of currencies, 1997-99



The paradox is that if one or two banks followed a DEAR limit and others did not, those banks would have an effective risk management system that at the margin would support the financial system. But if every bank were to follow the same approach, given that these banks follow each other into and out of markets, the DEAR limit would contribute to systemic risk. It is ironic therefore that the Basel Committee on Banking Supervision is supporting the rapid adoption of these systems across all banks and encouraging investors to follow suit, for an early reference see "An internal model-based approach to market risk capital requirements", Bank for International Settlements, Basel, 1995. There is a further paradox with transparency. The more herding investors and banks know about what each other are up to the more unstable markets may become. In the long run transparency and DEAR limits are a good development, but they are harmful in the short-run in the context of herding behaviour.

What should policymakers and regulators do?

Herding presents a classic example of the need for intervention. The individual incentives of herding investors create systemic risks. Moreover, if regulators were so coordinated that they behaved like one global regulator, they would be best placed to make an intervention. Through the privileged formation they have as a regulator of individual bank balance sheets they know when banks are herding. This requires a different focus. Today regulators are warned by other regulators when banks in their jurisdiction have exposures that threaten themselves, not whether banks around the world have exposures which together threaten a foreign market and could become contagious.

However, if this information were made public, in the context of herding investors, random shocks could quickly evolve into financial crises. But how should regulators respond if they notice herding in a particular market? They should require the bank to put aside an extra amount of capital for "strategic risk" without specifying which markets carry that risk. Applying tighter risk management requirements for those specific markets in which the herd has appeared will only make the stampede more vicious when negative news strikes. Collateral requirements are like a tax on banks and are very unpopular. However, banks could be given the choice of either putting up strategic collateral from time to time, or by buying liquidity options from the central bank during good times.

Whether these solutions would work or not, the whole concept of market-sensitive risk management practices needs to be seriously reconsidered in the context of herding and the authorities should rethink their extension to the use of credit rating agencies and the risk management of long-term investors.

It is arguable that regulators have actually promoted herding through risk management systems. They may also have done so in their zeal for disclosure of bank positions and central bank reserves. Indeed, there is a role for one unregulated investor who is encouraged to buy near the bottom of markets through the absence of risk, capital disclosure and credit concerns. Such investors would make the system safer but would be high risk and so should be restricted to those who can afford to lose. If this investor had to be invented she would look something like a hedge fund. Interestingly, as the big-betting hedge funds have been undermined by the disclosure and credit policies of banks, market liquidity has fallen and volatility has risen. Just as the big macro hedge funds fade away we may find that they supported the market as much as they exploited it.

Those of you unable to stomach regulators promoting hedge funds will be relived to note that there are other kinds of flows that do not herd so much, foreign direct investment for example. Further, during the Mexican and Asian crises, equity portfolio flows also revealed less herding than bond flows. It would appear that bond investors are keen to get out before they are held in by a debt moratorium or orderly work out. This raises some interesting questions for those trying to build in burden-sharing and orderly work out provisions into bond constitutions.

Transparency in data and governance is clearly a good thing in the long-run and promotes the right behaviour of governments. Governments should be encouraged to disclose more information every month and quarter, but not on a daily basis. In an environment of herding investors, there is not a good case for insisting that countries release central bank reserve data with such high frequency. It is telling that during the EMS crisis, many of the developed countries that have just adopted the Code of Good Practice on Transparency found it helpful to delay the monthly publication of their official reserves or to camouflage its information for several months. Small vulnerable emerging markets will find it even more helpful not to publish their reserves every day or every week and should not be forced to do so.

How safe was the "Safe Haven"?
Financial market liquidity during the 1998 turbulences

Christian Upper, Deutsche Bundesbank [1]

Summary

The turbulences in the international financial markets during the summer and autumn of 1998 put the price formation and liquidity provision mechanism in many markets under severe strain. As part of the large-scale portfolio rebalancing that took place, investors shifted a large part of their holdings into cash and into instruments that were perceived as having a low risk and being highly liquid. One of these "safe havens" was the market for German government securities. The paper examines the liquidity of the secondary market for four German benchmark government bonds during this period. The analysis is based on a unique dataset provided by the German securities' regulator, which covers every single transaction of the four bonds in Germany. This feature is particularly attractive for the bond market, where OTC transactions account for most trading.

The volatility of yields of the four bonds more than doubled in the wake of the Russian devaluation on 17 August 1998, and experienced a further peak in early October. It was accompanied by a widening in the yield spread between the individual bonds, which soared to more than twenty basis points from less than five basis points during the first half of the year. The cost of trading, as measured by the effective spread, increased even in a "safe" haven like the market for ten year German government bonds, indicating a reduction in liquidity. Nevertheless, the market was able to handle a statistically significantly higher than usual number of transactions and turnover. In this sense, liquidity provision has been remarkably effective in dealing with the turbulences.

Effective bid-ask spreads are positively related to unexpected trading volume, which should reflect the amount of private information in the market. Nevertheless, surprises volume cannot explain the surge in spreads that occurred during the turbulences.

1. Introduction

The turbulences in the international financial markets during the summer and autumn of 1998 put the price formation and liquidity provision mechanism in many markets under severe strain. Both the International Monetary Fund and the Bank for International Settlements[2] cite anecdotal evidence that "market liquidity dried up temporarily even in the deepest and most liquid markets as risks were repriced and positions deleveraged"(IMF [1998]). As part of the large-scale portfolio rebalancing that took place, investors shifted a large part of their holdings into cash and into instruments that were perceived as having a low risk and being highly liquid. One of these "safe havens" was the market for German government securities. The purchases of German debt securities by non-residents, which had averaged some of 11.4 billion DM per month during the first half of 1998, soared to 34.5 billion DM in July and remained close to this level in August. Most of this increase went into bonds issued by the Federal government, in particular those with maturities around ten years, whose yields fell accordingly.[3] The safe-haven effect disappears from the data as the turbulences reached their climax

[1] Discussion paper 1/00, Economic Research Group of the Deutsche Bundesbank. Parts of the paper have been written during a visit at the Bank of England's Financial Intermediaries Division, which I thank for its hospitality. I am particularly grateful to Charles Goodhart for helping me to arrange the visit and taking an active interest in its success. My work has benefited greatly from discussions with Torsten Arnswald, John Board, Dietrich Domanski, Heinz Herrmann, Manfred Kremer, Narayan Naik, Anthony Neuberger, Richard Payne, Victoria Saporta, Anne Vila and several staff at both the Bundesbank and the Bank of England. It would not have been possible without the generous help of Patrick Umlauf and Michael Kollak of the Bundesamt für den Wertpapierhandel, who made the data available. I would also like to thank Gerhard Klopf and Uwe Ullrich for valuable information on the institutional structure of the German bond market. All remaining errors are mine. The discussion papers published in this series represent the authors' personal opinions and do not necessarily reflect the views of the Deutsche Bundesbank.

[2] IMF (1998) and BIS (1999a).

[3] As a consequence, the spread between the benchmark government bonds and bonds issued by other institutions widened significantly. See Deutsche Bundesbank (1998).

in early October. Purchases of German bonds by non-residents fell back to 8.6 billion DM in September and were actually negative in October, when foreign investors reduced their holdings of German bonds by 10 billion DM.

This paper analyses the evolution of prices and indicators for market liquidity during 1998 for four benchmark 10 year Bundesanleihen. These securities represent the most liquid and actively traded segment of the German fixed income market, and were the main beneficiary of the flight into quality. In particular, I focus on the question of how the turbulences affected market liquidity and price formation. In other words, how did the safe haven weather the storm? Beyond this immediate question, the paper intends to further our understanding of the mechanisms for the provision of liquidity in securities markets more generally, especially in periods of stress.

The market for Government debt is of particular concern to central banks for a number of reasons.[4] Firstly, the prices of government securities provide important indicators for monetary policy, eg through the yield curve, and serve as benchmarks for the pricing of other securities. Such indicators will only be accurate if prices are efficient, ie if they accurately reflect all available information. If markets are not liquid enough, then this may not be the case.

Secondly, government bonds are important hedging vehicles for market participants. A breakdown in market liquidity turns tradable assets turn into non-marketable loans or at least induces large changes in prices. This can lead to substantial losses for those market participants who rely on the ability to turn over positions quickly and at a low price. Upper bounds for the cost of a temporary interruption of tradability have been derived using an options pricing approach by Longstaff (1995). Such costs arise from investors not being able to optimally time their trades. The upper bounds are concave in the time during which the asset cannot be sold, indicating that the cost of illiquidity may be very high even if the period of non-tradability lasts only a short period of time.

In addition, continuous liquidity is the key assumption of standard option pricing and risk management models, which are based on dynamic hedging strategies. If the underlying assets cannot easily be traded, then options can no longer be priced correctly and the riskiness of a portfolio cannot be assessed. While this does not directly impose a cost on investors, it contributes to a rise in uncertainty about their financial health and may trigger off massive sales or runs that could possibly call into question the stability of the financial system. The prevention of such periods of illiquidity should therefore be of serious concern to policymakers and central banks in their quest to safeguard financial stability.

A third reason for why central banks should be interested in the market for government bonds lies in their role as fiscal agents. If secondary markets are not very liquid, then this may add to the borrowing costs of the government. This cost of illiquidity can be substantial. For example, Amihud & Mendelson (1991) find that, after accounting for brokerage fees, the yield of US-Treasury Bills was on average 39 basis points lower than that of Treasury Notes of the same residual maturity but with a bid-ask spread about four times as high.[5]

Market microstructure theory has identified three factors which affect liquidity in financial markets: order processing costs, inventory control considerations and adverse selection problems.[6] Order processing costs include exchange fees and taxes as well as the more immediate costs of handling transactions. They should be fairly constant over time unless changes in the microstructure of the market or in technology take place, and are therefore an unlikely candidate when it comes to explaining variations in market liquidity.

Inventory and adverse selection effects are much more likely to vary even in the short run. The former arise from uncertainty about the order flow and thus the size of the inventory as well as from uncertainty about future prices and hence the valuation of the portfolio.[7] Adverse selection problems arise when a group of investors has private information on the value of the asset and will want to trade

[4] This has been pointed out by several papers in BIS (1999b).

[5] They argue that the larger spread of the T-Notes is the result of a large proportion of the securities being buried in investors' portfolios as a consequence of buy-and-hold strategies.

[6] A survey of the theoretical literature on market microstructure is O'Hara (1995).

[7] Models of inventory control include, among others, Ho & Stoll (1981, 1983) and O'Hara & Oldfield (1986).

only if the current ask price is below, or the bid price above the fundamental value of the asset. In either case, the market maker would make a loss, which he has to recoup from other investors by charging a positive bid-ask spread. This idea, which goes back to Bagehot (1971), has been formalised in a dynamic setting by Glosten & Milgrom (1985) and Kyle (1985), who analyse the pricing strategy of a risk-neutral market maker facing a sequence of informed traders (Glosten & Milgrom) and a strategically acting information monopolist (Kyle). In their models, market liquidity is a decreasing function of the precision of private information relative to the information available to the market maker. If the degree of asymmetric information increases beyond a threshold, then market makers would have to widen their bid-ask spreads by an extent that would drive away liquidity traders (Glosten & Milgrom) or would make prices infinitely sensitive to volume (Kyle). In either case, an equilibrium of the continuous trading involving a market maker would not exist. But even if continuous trading may not be feasible in equilibrium, Madhavan (1992) shows that it may still be possible to trade in auctions that take place at discrete time intervals over which all trades are aggregated and executed at a single price. They may even work if market makers withdraw from trading altogether.

There is by now a large literature on the importance of inventory control and adverse selection effects in the price formation process.[8] However, with the exception of the work of Proudman (1995) and Vitale (1998) on the UK gilt market, so far mainly equity markets have been looked at. This may be due to the fact that data is more easily available for equities, which tend to be traded on organised exchanges, than for bonds, which are mostly traded over the counter. Also, existing models of adverse selection are based on the assumption that some investors have superior information on the payoff of the asset. This is unlikely to be the case for government bonds, where cash flows are perfectly known. Private information may nevertheless play a role in these markets since some investors may know more about short term supply and demand conditions, eg on trading strategies of institutional investors.

Proudman (1995) estimated the importance of asymmetric information in the market for UK government bonds, using the VAR approach of Hasbrouck (1991). He finds no indication for either asymmetric information or inventory control. This is partly in contrast with the results obtained by Vitale (1998), who finds that while customer orders do not seem to provide any information to market makers, this is not the case for inter-dealer transactions, which do seem to carry information. This is exactly what we would expect from the modified adverse selection story informally set out in the previous paragraph. Vitale finds no indication for inventory effects, and comes to the conclusion that the market for gilts is very deep.

Both the inventory and the adverse selection models give hints at why market liquidity may dry up, but they don’t actually model this explicitely. Instead, they assume that the factors driving the degree of liquidity (volatility of the underlying price process, behaviour of noisy traders, precision of private information) are exogenous and constant over time. If we allow these factors to vary over time, possibly even endogenously, then the equilibria of the models may be different or may not even exist. There simply is, at least to my knowledge, no satisfactory theory of liquidity breakdowns. The same applies to the empirical literature, which assumes that spreads, and hence liquidity, is constant over time.

The present paper loosens this assumption and allows liquidity to vary across days, although not within a day. I find that while liquidity is relatively constant over time when markets operate in normal conditions, this is not the case during periods of stress. During the financial market turbulences of August to October, 1998, the effective bid-ask spreads in the market for German government bonds more than doubled, indicating a significant worsening of liquidity.

The paper is structured as follows: Section 2 reviews the literature on the 1998 turbulences in the international financial markets and presents a dating scheme that is used in the remainder of the paper. In the following section, I discuss the concept of market liquidity before turning to the issue of how to measure it. Section 4 gives an overview of the institutional structure of the market for Bunds and section 5 describes the data. Empirical results are presented in sections 6 and 7, which focus on yields and indicators for liquidity, respectively. A final section concludes.

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8 A good example is Huang & Stoll (1997), which also contains a comprehensive review of the literature.

2. The 1998 turbulences in international financial markets

In the summer and autumn of 1998, many international financial markets experienced a period of severe stress which provoked fear of a worldwide recession and deflation. Asset prices became very volatile, the bid-ask spreads in many, but not all, markets increased dramatically, and hitherto stable pricing relationships between different assets broke down.[9] Episodes of collapsing liquidity in securities markets are nothing new, but few if any have affected markets as important as the US bond market or the dollar-yen market.[10]

This section gives a brief outline of the events associated with the turbulences. It is largely based on the report of the Committee on the Global Financial System (BIS [1999c]), whose narrative-historic dating scheme, which informally draws on a large number of statistics, is presented in Table 1.

While it is tempting to date back the start of turbulences in the international financial markets to the Thai devaluation in July 1997, it was not until after the Russian default on 17 August 1998 that the turbulences spread to the markets of the developing economies. But even before, in July and the first half of August, there had been signs that strategies of relative value arbitrage, which involve buying an asset that is perceived to be undervalued and simultaneously selling a similar asset that is expected to fall in price, had led to losses at several large investors. Consequently, the BIS dating begins with a period of mounting tensions, ranging from 6 July to 14 August. Nevertheless, asset price and exchange rate volatility as well as bid-ask spreads were relatively low during this subperiod.

The Russian devaluation and effective default on its short term debt on 17 August caused sizeable losses to some investors and led to a general deleveraging of positions. This affected primarily the yield spreads of non-benchmark securities relative to the benchmarks and, to a lesser extend, price volatility and bid-ask spreads. The recapitalization of LTCM, a large hedge fund, on 23 September, marked the beginning of the third, most turbulent, subperiod, where yield spreads, volatility and bid-ask spreads soared to record levels. The turbulences began to subside after the inter-meeting rate cut by the Federal Reserve on 15 October, which marks the beginning of the fourth and final period, although yield spreads, volatility and bid-ask spreads remained high until the end of the year.

Table 1

The 1998 Financial Market Turbulences

Dates		Events
I.	6 July - 14 Aug	6 July: Salomon Brothers arbitrage desk disbanded 20 July: First Wall Street Journal on LTCM losses 23 July: Japanese sovereign debt placed under review
II.	17 Aug - 22 Sept	17 Aug: Russian effective default and rouble devaluation 2 Sept: LTCM shareholder letter issued
III.	23 Sept - 15 Oct	23 Sept: LTCM recapitalisation 29 Sept: Federal Reserve interest rate cut early Oct: Interest rate cuts in Spain, UK, Portugal and Ireland 7/8 Oct: Large appreciation of Yen relative to US dollar related to closing of “yen carry trades” 12 Oct: Japanese Diet approves bank reform legislation 14 Oct: BankAmerica reports 78% fall in earnings 15 Oct: Federal Reserve cuts rate between meetings
IV.	16 Oct - 31 Dec	5 Nov: Bank of England cuts rates 13 Nov: Brazil formally requests IMF programme 17 Nov: Federal Reserve cuts rates, Japanese sovereign debt downgraded 2 Dec: IMF Board approves programme for Brazil 3 Dec: Coordinated rate cut by European central banks 10 Dec: Bank of England cuts rates

Source: BIS (1999c).

9 See data appendix of BIS (1999c) for details.

10 Davies (1999) compares the 1998 turbulences to previous episodes.

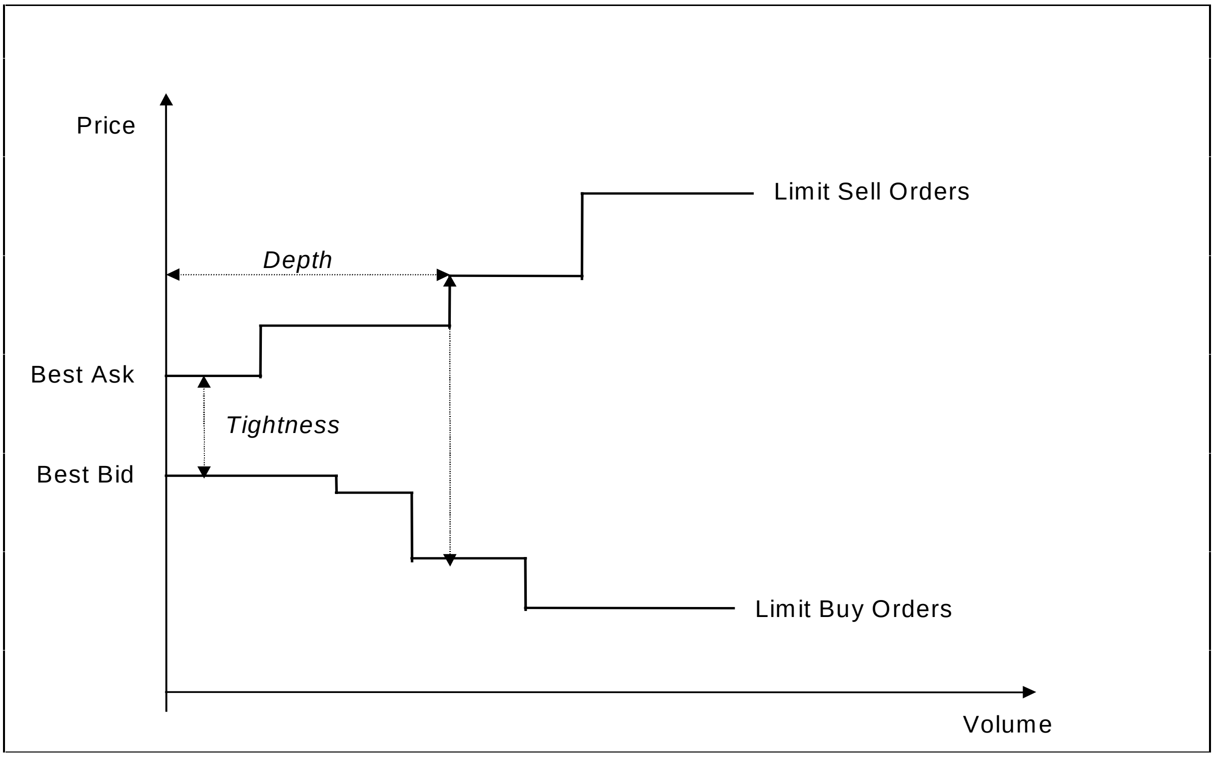
3. Liquidity measurement

3.1 The four dimensions of market liquidity

The liquidity of a secondary market refers to the "ease" with which securities can be traded. As is so often the case, liquidity is in general easy to recognise but difficult to define. Four dimensions of liquidity have been mentioned in the literature: tightness, depth, resiliency and immediacy.

A market is *tight* if there are enough limit orders or quotes in the vicinity of the last trading price such that new buy and sell orders can be executed without great discontinuities in prices. Tightness is directly measured by the bid-ask spread, which in Figure 1 corresponds to the vertical distance between the lowest ask and the highest bid prices. A market is *deep* if large orders can be executed without much effect on prices. In Figure 1, market depth is given by order size, measured on the horizontal axis, and the relevant spread for that amount, measured on the vertical axis.

Figure 1
Dimensions of market liquidity



Both tightness and depth are essentially static concepts, which describe a market at a given point in time. This is in contrast with the concept of *resiliency*, which describes the speed at which price movements due to temporary imbalances between buy and sell orders attract new orders which drive the price back to the fundamentally justified level. This is important because large orders are often split into several small orders which are executed sequentially. In a resilient market, an investor doing this would nevertheless obtain a price which, in the absence of any news, would on average be close to the current market price. In Figure 1, resiliency refers to the speed with which the bid and the ask schedules move back to their initial positions after an order has been executed.

The final dimension of liquidity is *immediacy*, which refers to the time that passes between the placing of a market order and its execution. This seems to have been a problem on the New York Stock Exchange during the stock market crash of October 1987, when orders had to wait up to 75 minutes for execution, apparently due to long printing queues.[11] Unfortunately, immediacy cannot be tested with transactions or quote data, since the time when an order was submitted is not known.

3.2 Measures for market liquidity

Market liquidity has traditionally been proxied by indicators for market activity like the number of trades, trading volume or the spacing between trades.[12] These measures have the advantage that they are immediately available or at least easy to compute, but unfortunately are not directly related to any of the four dimensions of liquidity presented in the previous subsection. Market liquidity is only one, and not necessarily the most important, factor affecting trading activity. For example, if traders have sufficiently large hedging needs, then they may trade even if market liquidity, in terms of tightness, depth, resiliency and immediacy, is relatively low.

Nevertheless, even though there is not a one to one relationship between trading activity and liquidity, indicators such as turnover or the intervals between trades do show that markets are sufficiently liquid to sustain a given amount of trading. The converse is not the case, however. The absence of transactions does not necessarily imply that markets are not liquid but may be due to the agents not wanting to trade at the market clearing prices. Such no-trade equilibria are common in theoretical models with homogenous agents. So, while we should not discard activity indicators altogether, it would be preferable to have measures that are linked to at least one of the four dimensions of market liquidity.

A measure for liquidity that is both easily available and in line with one of the four dimensions mentioned before is the *quoted* bid-ask spread, which indicates the tightness of a market. In quote-driven markets such as the over-the-counter market for German government bonds quoted spreads can be observed by traders at any point in time.[13] Unfortunately, quoted spreads may not represent the true cost of trading since many transactions take place at prices inside the quoted bid-ask bracket. In addition, they apply to transactions up to a specified amount only, which, furthermore, varies over time. E.g., if market participants cut this amount in periods of stress, then this would leave the quoted spread unchanged.

We can overcome the problems associated with *quoted* spreads by computing *effective* spreads. In contrast to the former, effective spreads refer to the average transaction and thus captures not only the tightness but also the depth of a market. They are computed from transactions data and thus reflect the price actually *paid* by traders and not the price *announced* by dealers.

The computation of the effective spread goes back to the seminal work of Roll (1984), who uses the negative first order autocorrelation in transaction prices that results from transactions being undertaken alternately at bid and ask prices. Assuming that buy and sell orders arrive with equal probability and that the bid-ask spread effectively paid by the traders remains constant over the estimation period, he shows that the spread is given by

$$s = 2\sqrt{-Cov\left(\Delta P_t, \Delta P_{t-1}\right)}$$

Unfortunately, there are a number of problems associated with the Roll measure. First of all, as is shown in Stoll (1989), in the presence of adverse selection or inventory effects it may underestimate the actual spread. Secondly, a bias may be induced by the fact that the number of bid and offer transactions may not be balanced.[14] Nevertheless, its simplicity and the ease with which it is

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[11] Brady (1989).

[12] For examples, see BIS (1999b).

[13] Since there is no commitment to market making in the OTC market for Bundesanleihen, quoted spreads are merely indicative, ie agents do not have to honour the prices they promise. While this should not matter in normal times, it may be a problem in periods of stress.

[14] See Choi et al (1988) on how to adjust Roll's estimator for this.

implemented make the Roll measure a good starting point when measuring liquidity, although its limitations have to be borne in mind.[15]

More recently, several other ways for computing spreads from transactions data have been developed.[16] Their computation is more involved than that of the Roll measure, which is why they have not been included in this paper.

4. The market for German government bonds[17]

This section describes the trading arrangements in the market for German government bonds. As is generally the case in the microstructure literature, the institutional structure of a market is of great importance since it conditions the availability of data and is important for the interpretation of the results. A description of the structure of the market for German government securities is also of interest in its own right, since it differs from that of other securities markets.

The market for German 10 year Bundesanleihen (Bunds) represents the most liquid segment of the European bond market and provides a benchmark for the pricing of long maturity bonds throughout Europe. German government bonds are traded both on organised exchanges and over the counter (OTC), with the latter accounting for the bulk of trading. This is especially true for the inter-dealer market, where trading takes place either directly between banks or through inter-dealer brokers (Freimakler). These act as intermediaries, but have no obligation to provide firm quotes. While the academic literature would suggest that dealers use brokers mainly in order to obtain anonymity,[18] conversations with market participants have hinted at a more mundane reason: Brokers seem to be more efficient in matching counterparties because of their more active participation in the market.

While there is a trend towards electronic trading, this stage had not yet been reached during 1998, when brokers posted their quotes on screen but the actual transactions took place over the phone. Electronic trading is more important in the client market. Large banks run proprietary trading systems, where they provide quotes for their customers. Recently, electronic trading systems as the Euro-MTS have become more important also in the inter-dealer market, but in 1998 their role was negligible.

The OTC market seems to be rather opaque. Participants observe quotes but do not receive timely information on other traders' past transactions beyond anecdotal evidence. In principle, it would be possible to obtain this information from clearing data, but it would arrive with a delay of up to three days and would be expensive to collect.

The role of exchanges is minor in comparison to the OTC market, even though all issues are listed. Trading takes place almost exclusively on the Frankfurt Stock Exchange, the role of regional exchanges is negligible. In 1998, German government bonds traded more or less entirely by open outcry since the computerised trading system XETRA did not permit traders to link their quotes to the price of the Bund future.[19] Prices are formed by a specialist broker (Kursmakler), who has to ensure price continuity. If the potential market clearing price lies outside a pre-specified corridor, he interrupts trading for five minutes and calls for new quotes. He may undertake some trading on his own account,

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15 Two practical difficulties may arise when implementing Roll's method. Firstly, the estimate for $Cov\ (\Delta P_t, \Delta P_{t-1})$ may be negative. In this case, I drop the observation from the sample. Secondly, the fact that the autocovariance of prices is not known but has to be estimated from the data induces a small sample bias. This can become a problem when estimating the effective spread over short time intervals like individual days. However, Roll (1984) argues that the bias is relatively small, eg for a sample of 60 observations it would amount to about two percent of the spread, which would seem acceptable. I deal with the problem dropping all intervals with less than 50 observations.

16 They are surveyed in Huang & Stoll (1997), who also present their own measure for estimating and decomposing the spread into the components due to order-processing, inventory control and adverse selection costs.

17 The description of the market for Bunds draws from conversations with Gerhard Klopf and Uwe Ullrich, of the trading desks of the Bundesbank and Deutsche Bank, respectively. I focus the market for Bunds as it was in 1998 and ignore many recent developments.

18 At least temporarily, as the banks receive details on their counterparty for clearing purposes. Only very recently did brokers act as a counterparty, thus preserving anonymity even after clearing.

19 This possibility has been provided in the summer of 1999, but it remains to be seen whether this will shift more activity onto XETRA.

but this occurs mainly to offset imbalances in order size and does not seem to play a big role in providing market liquidity.

5. Data

5.1 Description

The data covers all transactions in Germany during 1998 of four government bonds (Bunds) with a residual maturity of 8 to 10 years. It includes all *Bundesanleihen* (long term bonds issued by the federal government) with an original maturity of 10 years that were issued between January 1997 and July 1998. Details on the individual bonds are given in Table 2. The securities have been selected because they were delivered for the Eurex Bund future in 1998.

Bundesanleihen are usually first issued in January and July, and the size of the issue is increased during the following months. The autumn of 1998 marked an exception to this rule, as in the wake of falling interest rates a new issue with a lower coupon was brought to the market outside the normal issuing schedule in November. As a consequence, the 4.75% bond first issued in July 1998 has a volume outstanding of only 17 billion DM compared to 30 billion DM for the other three bonds in our sample.

The data has been provided by the German securities' regulator *Bundesaufsichtsamt für den Wertpapierhandel (BAWe)*, which receives notice of *every single* transaction in Germany, irrespective of whether or not it takes place on an exchange. This feature is particularly attractive for work on the bond market, where OTC transactions account for most trading.

Bunds are also traded abroad, in particular in London. Unfortunately, we do not have any information on transactions outside Germany, although there is anecdotal evidence that the market share of London has fallen in recent years, albeit by a lesser extent than the one in the market for Bund futures.

In principle, we should have two observations for each trade: one concerning the final buyer and another for the final seller of the security. An exception are trades with intermediaries that reside outside Germany and are therefore not subject to the reporting requirement. For each observation, we have a date and time stamp, price, volume, as well as indicators for whether it is the sale or the purchase leg of a transaction, whether the reporting bank's inventory is affected, whether the reporting bank acts as an agent for a customer or on its own account, and whether trade takes place over the counter or on an exchange.

Some observations on the quality of the data are in order. Firstly, the precision of the time stamps (up to the nearest second) may be more apparent than real. It is probably high for trades that take place on an exchange, but may be low for OTC-transactions. An indication for this is that many trades are reported as taking place outside normal office hours, often at midnight sharp or at prices that seem out of date. Inaccurate time stamps represent a severe problem since the reliability of most indicators for liquidity crucially depend on the correct sequence, although not so much the timing, of trades. Secondly, other variables may have been reported incorrectly. While most of such errors have been eliminated by the BAWe using supplementary information, some mistakes especially in prices or volume may remain.

Matching the buy and sell legs of transactions and filtering the data can reduce the severeness of these problems. Unfortunately, matching the two legs is not straightforward, in part because of errors in the data and in part because of missing notifications. I have matched all purchases and sales that occurred on the same day at the same price and with the same volume. If the time stamps of these pairs of "matched" observations do not coincide, I have used the earlier one. Of a total of 145 thousand observations in our raw data, only about half could be matched in this way.[20] Since the remaining "unmatched" observations account for a large proportion of market activity, I have decided to keep them in the data. After matching, we are left with 109 thousand observations, which further

[20] The BAWe successfully matches the various notices for more than three quarters of the trades but uses supplemental data to achieve this. They use the matches in order to hunt down missing notifications and to eliminate contra-trades. Since the matches are stored in a different database, using this information in the present study would have incurred computational costs that seemed out of proportion to the limited benefits of doing so.

reduce to 99 thousand after filtering.[21] Of these, 66 thousand, 2/3 of the total, were not due to matches (see Table 2).

Table 2
Summary statistics[1]

Coupon	6%	6%	5,25%	4.75%
Maturity	04.01.07	04.07.07	04.01.08	04.07.08
Volume outstanding (in billion DM)	30	30	30	17[2]
Number of trades				
- unmatched	21,366	33,616	64,594	25,586
- matched	16,608	22,832	48,191	19,551
- matched and filtered	14,799	22,407	44,216	17,679
of which not due to matches	10,490	14,247	29,018	12,198
Client trades[3]	2,186	2,966	5,678	2,283
Sales[3]	548	806	1,663	827
Purchases[3]	534	794	2,010	618
Could not be attributed[3]	1,104	1,366	2,005	838
Inter-dealer trades[3]	2,123	5,194	9,520	3,198
Total volume (DM billion)	273.2	486.2	780.4	281.1
Client trades[3]	64.7	113.9	159.7	54.1
Inter-dealer trades[3]	25.9	57.0	99.6	41.9
Average trade size (DM million)	18.5	21.7	17.6	15.9
Client trades[3]	29.6	38.4	28.1	23.7
Purchases[3]	17.7	22.6	15.5	16.8
Sales[3]	19.6	21.4	15.1	14.1
Not attributable[3]	43.5	58.9	54.5	38.3
Inter-dealer trades[3]	12.2	11.0	10.5	13.1

[1] 2 January 1998 - 30 December 1998 except for 4.75% due in 2008, which covers 7 July - 30 December 1998 only. [2] 12 billion on 8 July, 2 billion on 21 August and 3 billion on 31 August. [3] Matches only.

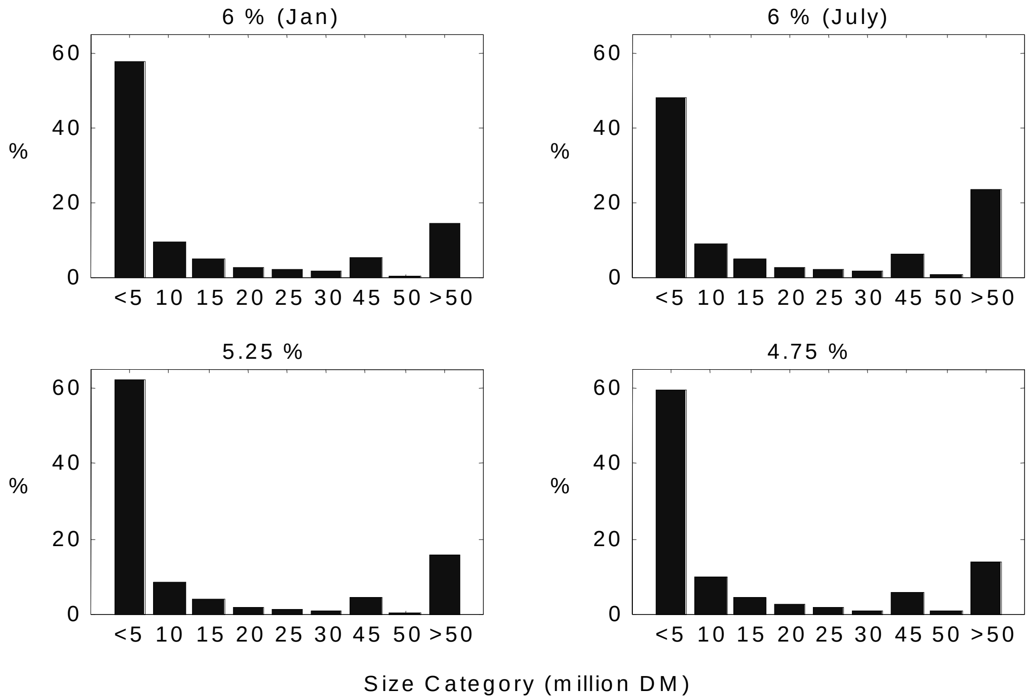
Of the 99 thousand observations of our matched and filtered sample, 44 thousand, or just under half, are for the 5.25% bond due in January 2008, which was on-the-run (ie the most recent issue) during the first half of 1998. There are far less transactions for the two 6% bonds, with 15 and 22 thousand trades, respectively. The 4.75% bond issued in July 1998, which was on the run between July and November, when a new Bundesanleihe not included in our sample was issued, was traded 18 thousand times.

As was mentioned before, our dataset contains information on whether the sale or purchase leg was an agency trade or on the intermediary's own account. In the case of "matched" observations, this allows us to distinguish inter-dealer trades, where banks appear on both sides, from client trades, which cover the remainder. A transaction is classified as a client sale (purchase) if the sale (purchase) involved a client and the counterparty was a bank. If both the purchase and the sale are agency trades, then the transaction is counted as non-attributable client trades. Unfortunately, this classification is not possible for "unmatched" trades.

Client trades account for less than 40% of the "matched" transactions, and inter-dealer trades for the remainder. Of the former, about 30% were identified as purchases and approximately the same number as sales, the remaining client transactions could not be attributed.

[21] I have deleted all observations that are reported to take place either before 8:00 or after 18:00 and those with prices outside a bracket of three standard deviations around a second order polynomial fitted to the prices of all transactions that occur during the same day. While the deletion of observations necessarily leads to a loss in information, I believe that this is much less severe than the problems associated with wrong time stamps or outliers in prices.

Figure 2(a)
Distribution of trade size
(all trades)

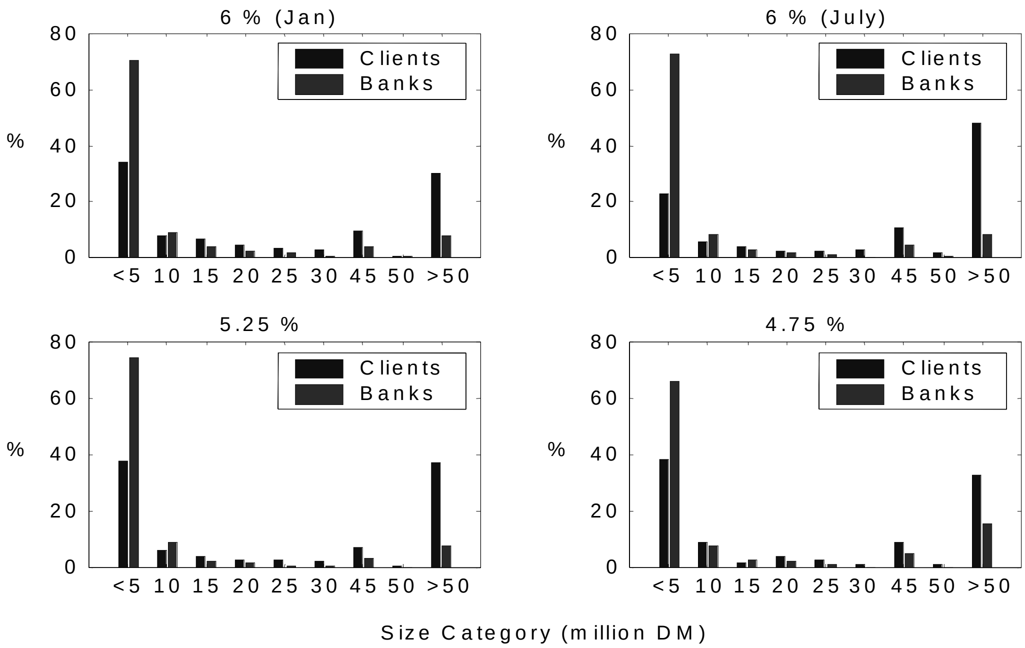


The distribution of trade sizes of all transactions is shown in Figure 2(a). Small trades of up to 5 million DM accounted for around one half of all transactions and large trades above 50 million DM for a further tenth to one fifth. These aggregate figures mask important differences between the different types of transactions, as can be seen from Figure 2(b). Inter-dealer Trades are much more likely to be small than client transactions, where about one third of all trades tops 50 million DM. As a consequence, the average trade size was considerably larger for client trades than for inter-dealer transactions for all four bonds. This gap exists for all three classes of client transactions (sales, purchases and not attributable), although it is by far largest for trades which involve clients on both sides.

Figure 2(b)

Distribution of trade size

(client and inter-dealer trades)



Our data also contains information on whether or not a trade affected the reporting bank's inventory, which permits us to construct a series for the aggregate inventory of the banking sector at any given point in time by simply adding up all transactions which affect inventories, and assuming that the inventory at the beginning of the year was zero. In the case of the two 6% bonds, this latter assumption is arbitrary, the inventory series for these bonds can therefore only be interpreted in terms of movement over time but not in terms of the actual level of inventory.

Figure 3 shows the inventories at the end of each trading day for the four bonds.[22] They tend to fluctuate within a relatively narrow band around zero. On more than 50% of all days, they are between –10 and +10 million DM, which is considerably less than the average transaction size, and exceed 100 million or are below –100 million on only 5% of trading days.

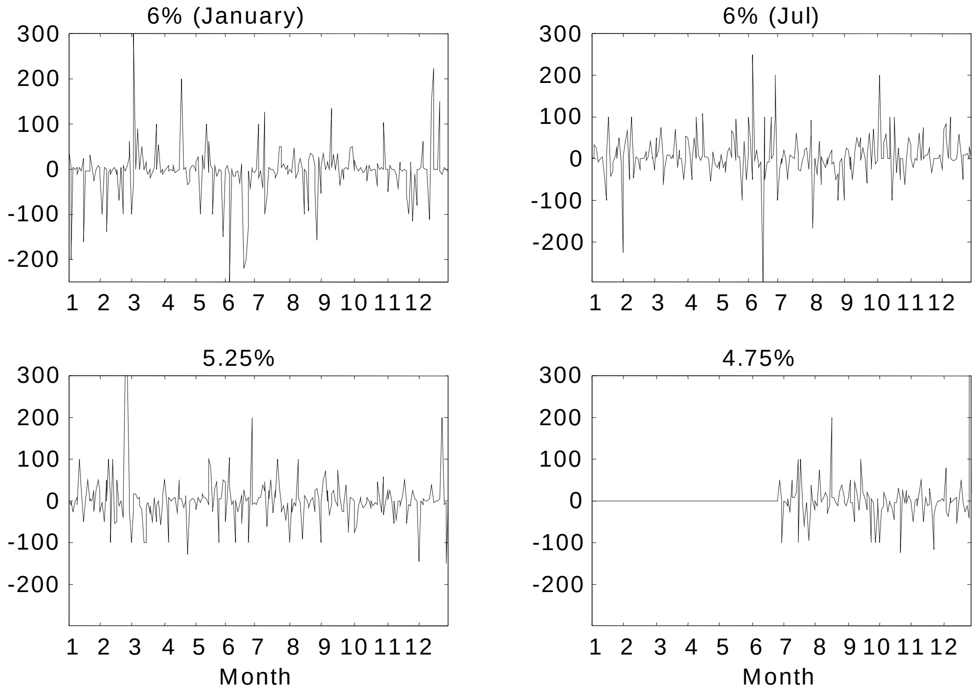
According to our figures, between 78% and 86% of all trades took place over the counter,[23] although this may understate the true figure as many OTC-transactions after execution are entered into an exchange's trading system in order to take advantage of the automated clearing facilities, and reporting banks may "forget" to enter the variable correctly.

[22] End-of-day inventories have been computed from the unfiltered data, since dropping any transaction would reduce the reliability of the data. In addition, misreported time stamps or prices do not have any effects on the quality of the series.

[23] Trades are counted as OTC if identified by the *börslich*-flag or if recorded before 8:00 or after 18:00 hours.

Figure 3

End-of-day inventories



In million DM

5.2 Intra-day patterns of trading activity

Trading activity is very unevenly distributed over the trading day, as can be seen from figures 4(a) – (c), which plot the distribution of the number of trades, turnover and average trade sizes within half hourly intervals. Trading activity shows two peaks during the day. The first peak occurs in the morning between 9:30 to 10:00, when around 10% of all trades take place, which however account for a much higher share of daily turnover. The average trade size in this half hour period is consequently more than twice as high as during the remainder of the day. The second peak in activity occurs between 11:30 and 12:00, when between 12% to more than 20% of all trades take place. These tend to be small trades, however, with an average trade size below that of any other interval. Trading is least heavy in the early morning, before 8:30, at lunchtime between 13:00 to 14:00, and in the evening after 17:00.

Figure 4 (a)
Intra-day distribution of number of trades

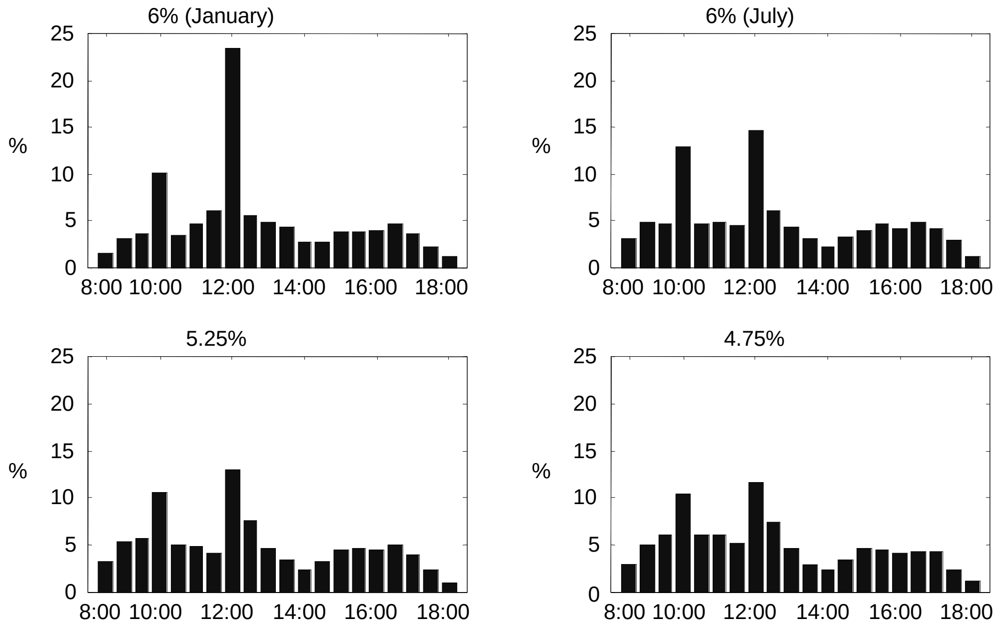


Figure 4 (b)
Intra-day distribution of trading volume

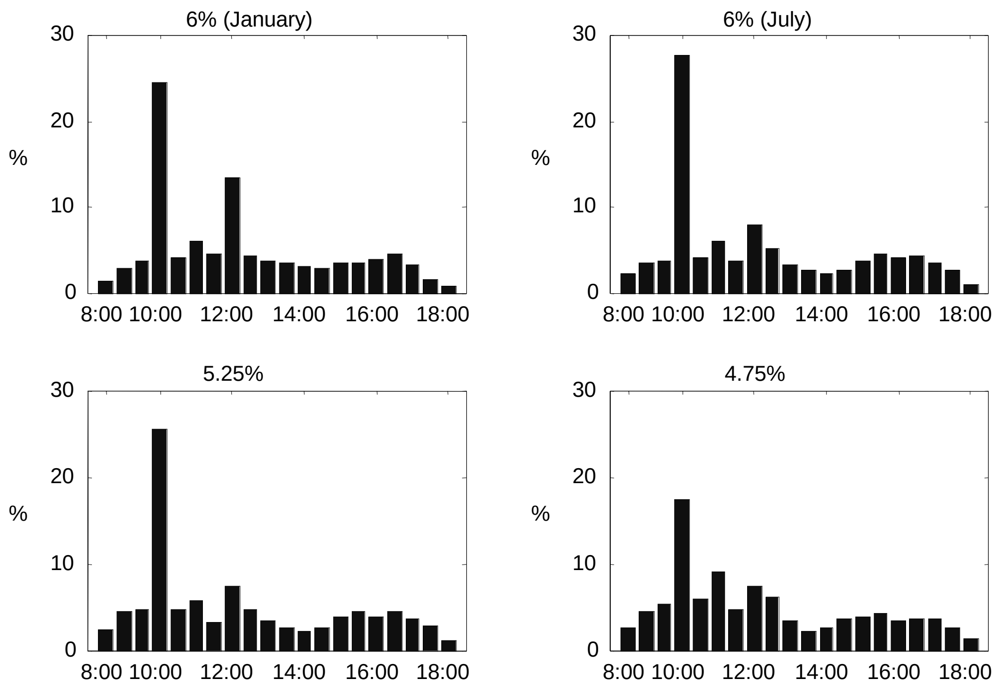
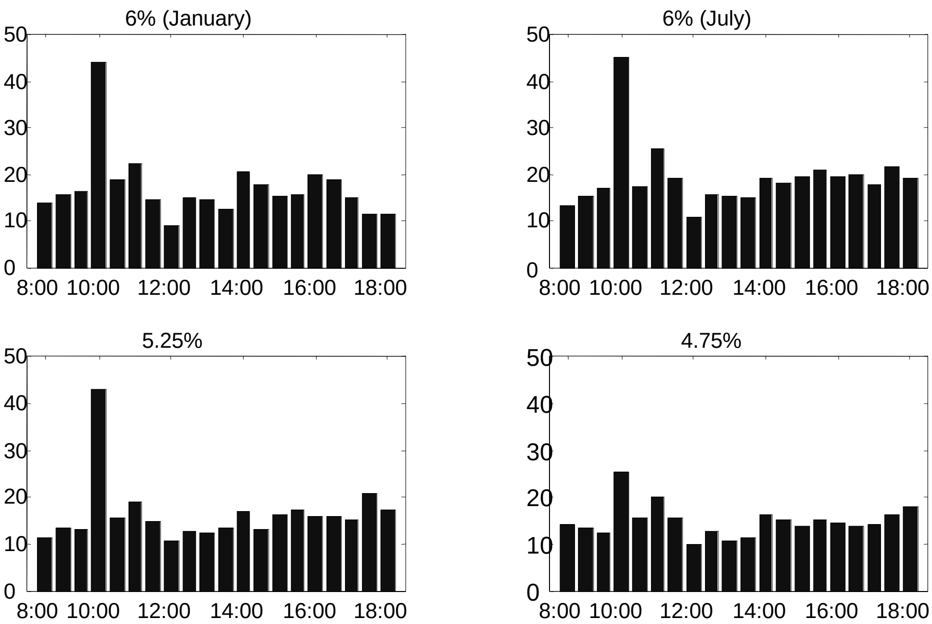


Figure 4 (c)

Intra-day distribution of average trade size



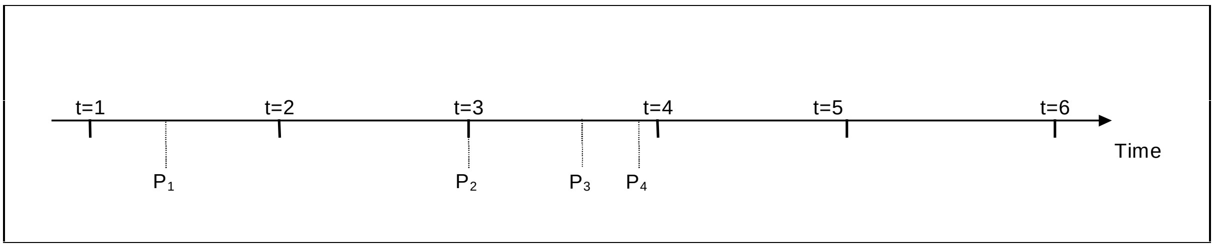
In million DM

5.3 Construction of equally-spaced time series

The presence of intra-day patterns of trading activity can induce biases in the measurement of volatility and correlations if we do not ensure that only trades that take place at the same time are compared. For some purposes, I convert the irregularly-spaced transaction series into an equally-spaced time series of prices. How this was done is illustrated in Figure 5. Transactions occur at irregular time intervals at prices P_1, P_2 etc. We lay a grid of fixed time intervals t=1,2...T over the trading day, and assign to each point of the grid the price of the last available transaction. For example, the price at t=2, $P_{t=2}$ becomes P_1, $P_{t=3} = P_2$ and $P_{t=4} = P_4$. If no transaction occurs during an interval, we still take the last available transaction, thus $P_5 = P_{t=5} = P_4 = P_{t=4}$. If there has not been any transaction on the same trading day, eg at t=1, then the observation is dropped. This seems preferable to using the closing transaction of the previous day since much more news tends to arrive over night than in any brief trading interval. Observations followed by other transactions within the same time interval, like P_3, are discarded.

Figure 5

Construction of an equally spaced time series



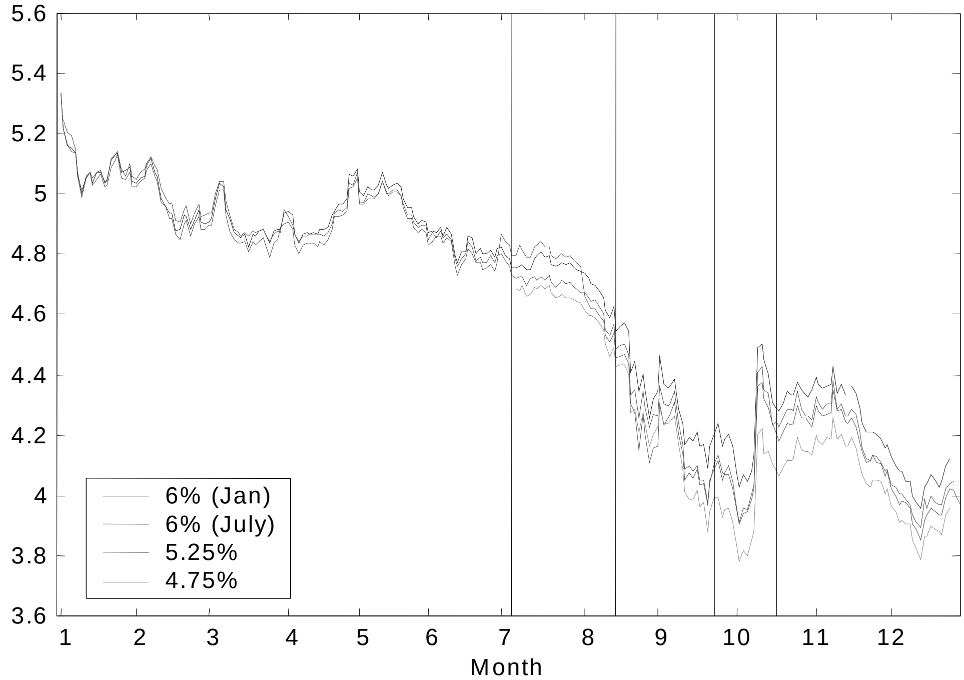
In deciding the length of the fixed intervals, one faces a trade-off between the number of periods without any observations and the amount of observations that are discarded. For the data of this paper, five minute intervals seem a reasonable compromise between the two, although there is some indication that this frequency is too high for days with low trading activity. However, since the focus of the analysis are the turbulences during the summer and autumn, when, as we shall see in section 7, trading activity was particularly large, this seemed less of a problem than loosing information during these periods.

6. Yields of bunds in 1998

Figure 6 shows the yields of the four bonds of our sample during 1998.[24] During the first half of the year, they fell from around 5.3% at the beginning of January to about 4.8%. The decline accelerated sharply during the second half of July and August. Yields reached their nadir at around 4% in early October. After a sharp rebounce, they again declined to values of close to 4% in December. It is interesting to note that the sharp fall in yields began about a month *before* the Russian default, ie during the first period according to the BIS (1999c) classification presented in section 2.

Figure 6

Yields to maturity



The month of July marked a break not only in the overall movement of yields but also in the pricing relationship between the individual bonds. During the first half of the year, the yields of the three bonds then available moved very closely together, rarely diverging by more than five basis points. In July, the yield spread between the individual securities widened sharply to 12 to 14 basis points between the most recent issue, the 4.75%, and the 6% (July) issue.[25] Such a rise in the spread between on-the-run and previous issues seems to have been a quite general phenomenon that also occurred in several other financial markets (BIS [1999c]), and was probably due to the inflow of foreign capital which was invested mainly in the most recent issues.

The yield spread quickly narrowed in August, and even changed sign towards the end of that month. This was probably due to an unusually large open interest in the Bund Future, which aroused fears of a shortage of paper for physical delivery for the September contract. The specification of the Bund

24 In order to prevent distortions due to intraday movements in prices or trading activity, the price at 12:00 or last available transaction before was taken.

25 The 6% (January) bond dropped out of the basket of deliverable securities and is therefore ignored.

Future tends to make it cheaper to deliver a bond with a coupon close to 6% rather than bonds with lower coupons. If the demand for physical delivery exceeds the amount on the market of the bond that is cheapest to deliver, then some investors have to settle with more expensive paper. This would drive up the price of the cheapest to deliver by an amount determined by the difference in delivery costs. This is precisely what seems to have happened in August of 1998, when investors expected a shortage of the 6% (July) and expected to have to settle with the 5.25% or even the 4.75%. When it became apparent in early September that there would be no shortage of the cheapest to deliver because many market participants had closed their open interest, the relationship between the yields quickly reverted to the previous pattern with a positive spread between the most recent and preceding issues. The yield spread between the 6% (July) and the 4.75% reached its maximum of 21 basis points on 13 October and then declined to around 10 basis points in November and December. At that time a new Bund had been issued, such that our figures do no longer represent the on-the-run/of-the-run spread.

The change in the pricing relationship between the bonds does not extend to high frequencies, though. As can be seen from Table 3, the correlation between the yields of the individual bonds, measured at both the daily and hourly level, did not differ in any economically significant way between the two halves of 1998, although their first differences (ie the daily or hourly changes in the yields) do show a higher correlation during the second half of the year than during the first half.

Table 3
Correlations of yields

		Levels		First differences	
		5.25%	4.75%	5.25%	4.75%
1.	Daily				
(a)	01.01.99 – 30.12.99				
	6% (July)	0.994	n/a	0.932	n/a
	5.25%		n/a		n/a
(b)	01.01.99 – 30.06.99				
	6% (July)	0.981	n/a	0.884	n/a
	5.25%		n/a		n/a
(c)	07.07.99 – 30.12.99				
	6% (July)	0.986	0.980	0.951	0.947
	5.25%		0.994		0.970
2.	Hourly				
(a)	01.01.99 – 30.12.99				
	6% (July)	0.993	n/a	0.313	n/a
	5.25%		n/a		n/a
(b)	01.01.99 – 30.06.99				
	6% (July)	0.973	n/a	0.234	n/a
	5.25%		n/a		n/a
(c)	07.07.99 – 30.12.99				
	6% (July)	0.984	0.976	0.363	0.447
	5.25%		0.993		0.436

Let us now turn to the volatility of bond yields. Since the expressions “turbulent” and “stress” are almost synonymous for “high volatility”, this is what we would expect during the second half of 1998, in particular during the time between the Russian default of 17 August and the US rate cut on 15 October.

We can compute a measure for the price volatility of each stock on each trading day from the equally spaced price series by summing intra-day squared holding returns

$$\hat{\sigma}_t^2 = \sum_i \left(\Delta \log P_i\right)^2 .$$

Andersen et al (1999) show that this approaches the price volatility of a continuous process as the intervals between the observations goes to zero. The resulting series of volatility intra-day volatility is shown in Figure 7.

Figure 7
Intra-day volatility of yields in 1998

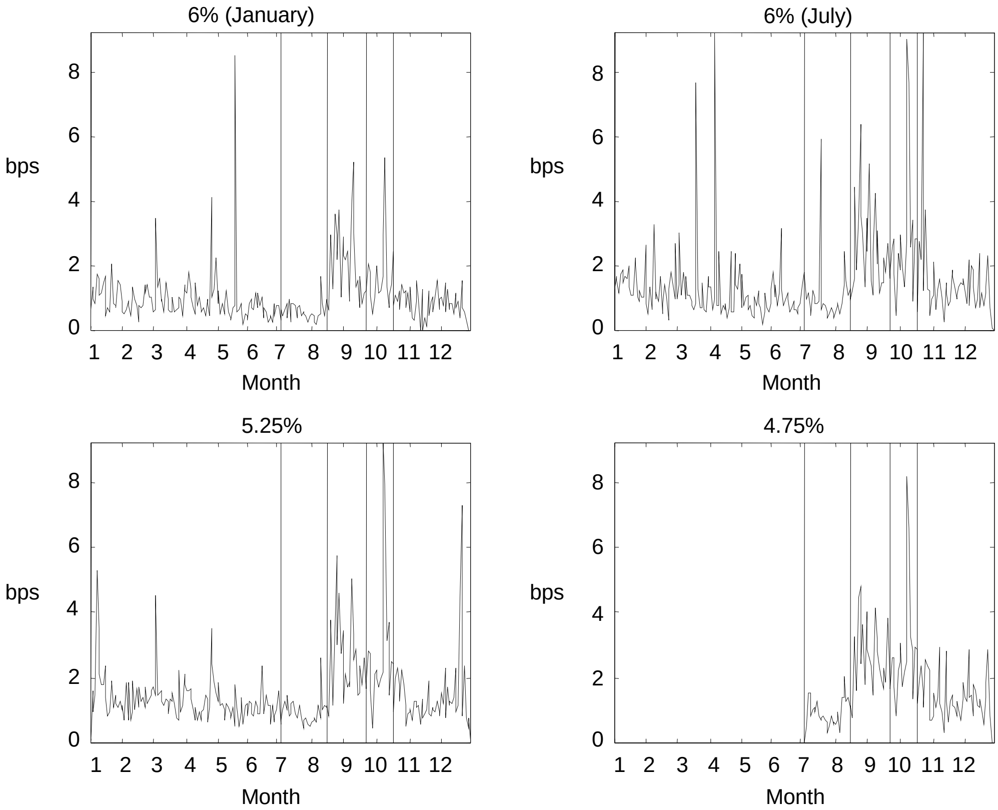


Figure 7 is roughly consistent with what we know from other markets. Volatility remains low during the first period of the turbulences, but soars after the Russian default. Our measure remains high throughout periods II and III, but gradually declines after the US intermeeting rate cut on 15 October, although it remained at around two basis points well above its initial level during most of November and December.

7. Market liquidity in 1998

What were the effects of the capital market turbulences of the summer and autumn of 1998 on the liquidity of the market for German government bonds? Let us begin by looking at activity indicators like the number trades as well as market turnover, before turning to the effective spread as a direct measure for market tightness and depth.

7.1 Activity indicators

The daily number of trades and daily turnover, depicted in figures 8 and 9, decreased somewhat during the first subperiod covering July and part of August relative to the tranquil period, but soared after the Russian default. This coincided with massive inflows of foreign capital into the German bond market of over 30 billion DM per month in July and August, compared to an average of just over 10 billion DM during the first half of 1998. After peaking in late August, activity fell back to normal. Another brief surge in trading took place in the beginning of October, although for some reason this did not affect the 6% (July) bond. In early November, both the number of trades as well as turnover of the 5.25% and the 4.75% dropped due to the introduction of a new bond.

Figure 8

Daily number of trades in 1998

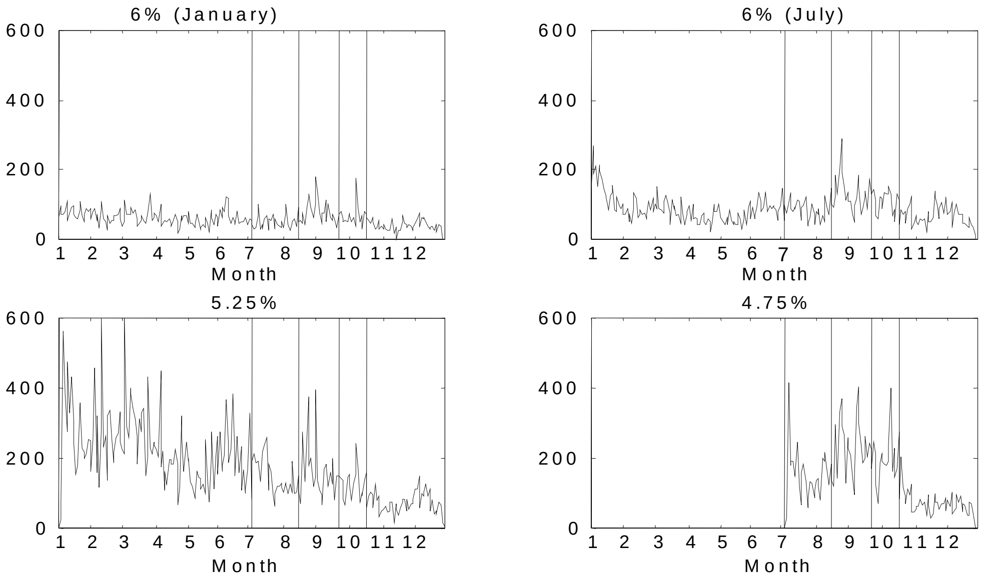
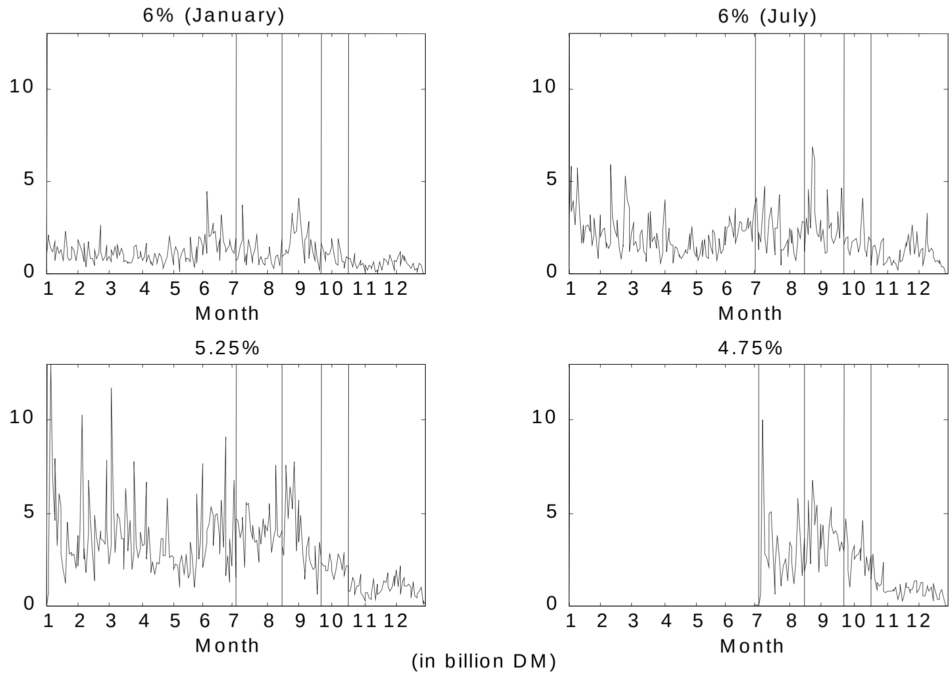


Figure 9

Daily turnover in 1998



Are the changes in trading activity during the period of turbulence statistically significant? This can be tested by regressing the activity measure, in our case the daily number of trades or daily turnover, on a set of dummy variables that are equal to one in the respective subperiod and zero otherwise. In order to isolate the effects of the turbulences from the effects of a change in the on-the-run issue, I included an further dummy D_{recent} indicating whether or not the bond was the most recent issue. In order to reduce multicolinearity, D_{recent} and the dummies for the first and last period of the turbulences have not been included in the equation for the 4.75%.

Table 4
Trading activity during the 1998 turbulences

	6% (Jan.)		6% (July)		5.25%		4.75%	
Dep. variable	(i)	(ii)	(i)	(ii)	(i)	(ii)	(i)	(ii)
Expl. variables								
$D_{6.7.-14.8.}$	-0.00 (-0.01)	**-0.29 (-1.85)	***0.38 (3.47)	**0.29 (2.04)	0.16 (0.33)	0.52 (0.94)	-	-
$D_{17.8.-22-9.}$	***0.61 (4.84)	**0.35 (1.87)	***1.09 (8.37)	***0.75 (4.48)	0.46 (0.91)	0.48 (0.83)	***0.93 (8.45)	***0.94 (6.65)
$D_{23.9.-15.10.}$	***0.46 (3.05)	*-0.333 (-1.46)	***0.96 (6.08)	*0.29 (1.38)	0.37 (0.73)	0.05 (0.09)	***0.68 (5.12)	***0.50 (2.94)
$D_{16.10.-31.12.}$	0.07 (0.42)	***-0.92 (-3.83)	***0.69 (4.13)	0.07 (0.30)	0.28 (0.54)	-0.22 (-0.37)	-	-
D_{recent}	-	-	0.25 (0.63)	0.15 (0.28)	0.16 (0.33)	0.11 (0.20)	-	-
# observations	250	250	250	250	249	249	124	124

Dependent variable: (i) daily number of trades, (ii) daily turnover, normalised by their mean. T-values in brackets. Dummies equal to 1 during indicated period and zero otherwise, dummy *recent* one if on-the-run and zero otherwise. Explanatory variables also include a constant and time trend. Newey-West correction for autocorrelation and heteroscedasticity. Significant on *90%, **95%, ***99% confidence level

The results of the regressions are displayed in Table 4. They indicate that the surge in both the number of trades and in turnover during the second period of the turbulences is statistically highly significant. The only exception to this is the 5.25% bond, where none of the dummies, including , D_{recent}, proved to be significant. There is also some indication that the number of trades, although not in all cases turnover, was higher during the third period than on average. The individual bonds behaved quite differently during the initial and final periods of the turbulences, when some of them became more actively traded and the others less so.

The results indicate that markets have been sufficiently liquid to sustain a larger than normal number of trades and turnover, but this does not necessarily imply that market liquidity remained unaffected. It is therefore necessary to look at direct indicator for market liquidity and not solely rely on activity indicators like the number of trades of turnover.

7.2 Effective spreads

Let us now turn to a measure of liquidity that is directly related to both the tightness and the depth of the market, the effective spread measured by Roll's statistic. Before proceeding, a word of caution is needed. The Roll measure is likely to underestimates the "true" spread in the presence of inventory control or adverse selection effects. Since these are more likely to occur in times of market stress (when prices are more volatile, adding to inventory costs, and there is more scope for private information), our results are liable to underestimate any worsening in market liquidity.

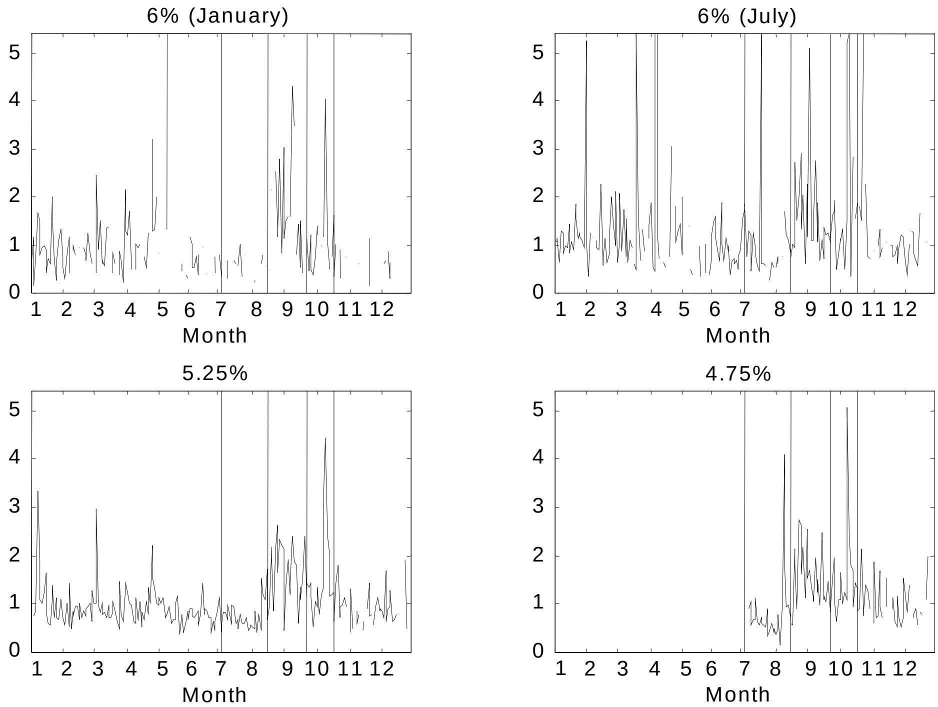
Table 5 shows the Roll measures for the effective spread estimated from transactions prices. The spreads appear to have been very low, and consequently the market very liquid, during the first subperiod of the turbulences ranging from the beginning of July to the eve of the Russian default a month and half later. After the Russian default, effective spreads increase to 1.73 to 1.9 basis points, although they remain below their level of the first half of the year for the two 6% bonds. They rise further after the LTCM rescue, to more than 2 basis points. The only exception is the 6% (January) bond, whose effective spread declines to 1.3 basis points. The spreads ease during the fourth and final subperiod to levels near or even below that of the first half of the year. It is possible that the odd result of a very high spread for this bond during the period of tranquillity is due to data problems.

Table 5
Effective bid-ask spreads (Roll measure, in basis points)

Period	6% (January)	6% (July)	5.25%	4.75%
2 Jan – 3 July	2.60	1.95	1.13	n/a
6 July – 14 Aug	0.58	1.54	0.78	1.18
17 Aug – 22 Sept	1.90	1.88	1.80	1.73
23 Sept – 15 Oct	1.28	2.51	2.09	2.02
16 Oct – 31 Dec	0.82	1.48	1.21	1.11

The estimation of effective spreads over prolonged periods of time has the advantage that one does not have to worry about having too few observations, but their results may mask important developments during these periods. I therefore complement the analysis from Table 5 by estimating the Roll measure for each trading day of the year. In order to reduce the small sample bias associated with this measure, I only consider days with more than 50 trades, which explains the gaps in the graphs, mainly for the two 6% bonds, which also show a lot of noise.

Figure 10
Effective spreads (Roll measure) in 1998



(in basis points of yields)

The daily spreads reproduced in Figure 10 broadly confirm the picture obtained from Table 5. Effective spreads for the 5.25% bond fluctuated around one basis point during the first half of the year and fell somewhat during July and early August. During the second half of August, they soared to values of more than two basis points but eased somewhat during September. A dramatic increase occurred around 8 and 9 October, when the spread for the 5.25% reached 4.4 and those of the 4.75% and the 6% July bonds broke through the 5 basis point barrier. Afterwards, effective spreads quickly declined to around 1.5 basis points in November and December.

We can repeat the exercise of the previous section and test whether the changes in spreads are statistically significant. Since changes in the effective spread may be due to variations in order size, I include the daily average trade size as additional explanatory variable. The results are presented in Table 6.

Table 6

Daily effective spreads during the 1998 turbulences

	6% (Jan.)	**6% (July)**	**5.25%**	**4.75%**
Expl. variables				
$D_{6.7.-14.8.}$	**-1.294 (-1.77)	0.114 (0.31)	-0.206 (-0.44)	-
$D_{17.8.-22-9.}$	-0.664 (-0.88)	*0.584 (1.35)	*0.754 (1.54)	***0.461 (3.51)
$D_{23.9.-15.10.}$	-1.069 (-1.17)	**0.935 (1.73)	*0.800 (1.59)	***0.530 (3.39)
$D_{16.10.-31.12.}$	*-1.531 (-1.44)	0.635 (1.08)	0.249 (0.49)	-
D_{recent}	-	-0.293 (-0.24)	-0.165 (-0.35)	-
Av. trade size	*0.522 (1.33)	**0.447 (1.73)	0.065 (0.70)	-0.010 (-0.05)
# observations	250	250	249	124

Dependent variable: Roll measure for the effective spread, normalised by their mean. T-values in brackets. Dummies equal to 1 during indicated period and zero otherwise, dummy *recent* one if on-the-run and zero otherwise. Explanatory variables also include a constant and time trend. Newey-West correction for autocorrelation and heteroscedasticity. Significant on *90%, **95%, ***99% confidence level.

The coefficients of the dummies for the second and third periods of the turbulences are significantly positive for three of the four bonds. This indicates that liquidity had indeed worsened after the Russian default and only recovered with the inter-meeting rate cut of the Fed. We find no evidence for higher effective spreads either in the run-up period between 6 July and 14 August or the cooling-down period after 16 October. The results obtained for the 6% (January) bond do not contradict this picture but are consequence of a relatively high spread during the first half of the year.

We can test for the importance of the adverse selection component of the spread by running a regression of the effective spreads on expected and unexpected trading volume (Bessembinder [1994], Hartmann [1999]). Unpredictable volume should mainly be due to the arrival of information, which increases the scope for adverse selection problems, and should therefore be positively related to the spread. A surge in predictable volume, in contrast, permits market makers to take advantage of scale economies and hence reduce spreads. In addition, a rise in expected volume that goes hand in hand with a larger number of orders (rather than an increase in order size) reduces the risk of inventory imbalances due to the law of large numbers.

I estimate the following equation:

$$S_{it} = \beta_{i0} + \beta_{i1} evol_{it} + \beta_{i2} uvol_{it} + \beta_{i3} evola_{it} + \varepsilon_{it},$$

where expected volume $evol_{it}$ of bond i is proxied with the (in-sample) fitted values from a standard ARMA model,[26] and unexpected volume $uvol_{it}$ with the corresponding residuals. In order to control for variations in the inventory holding costs, I control for the expected price volatility $evola_{it}$, also obtained from an ARMA model. Since unexpected volume may be determined by the bid-ask spread, we face a problem of endogeneity that could induce a bias in our estimation results (Hartmann [1999]). Unfortunately, our dataset does not contain any variable that is exogenous and could be used as an

26 The prefered specification was an ARMA(1,1) model. Dummies for weekday effects turned out not to be significant.

instrument for $uvol_{it}$.[27] We therefore have to live with the endogeneity problem, although we have to interpret the results in light of the possible bias. The results of the estimation are displayed in Table 7.

Table 7

Effective spreads and trading volume

	6% (July)	**5.25%**	**4.75%**
Expl. Variables			
Exp. volume	0.202	*0.164	0.103
	(0.87)	(1.89)	(0.92)
Unexp. volume	0.201	***0.524	***0.659
	(0.52)	(3.72)	(4.12)
Exp. volatility	**0.349	***0.555	***0.337
	(2.44)	(7.78)	(4.36)
R^2	0.06	0.46	0.25
# observations	124	137	109
F-Test: H_0: $\beta_1=\beta_2$	-	**5.4	***11.4

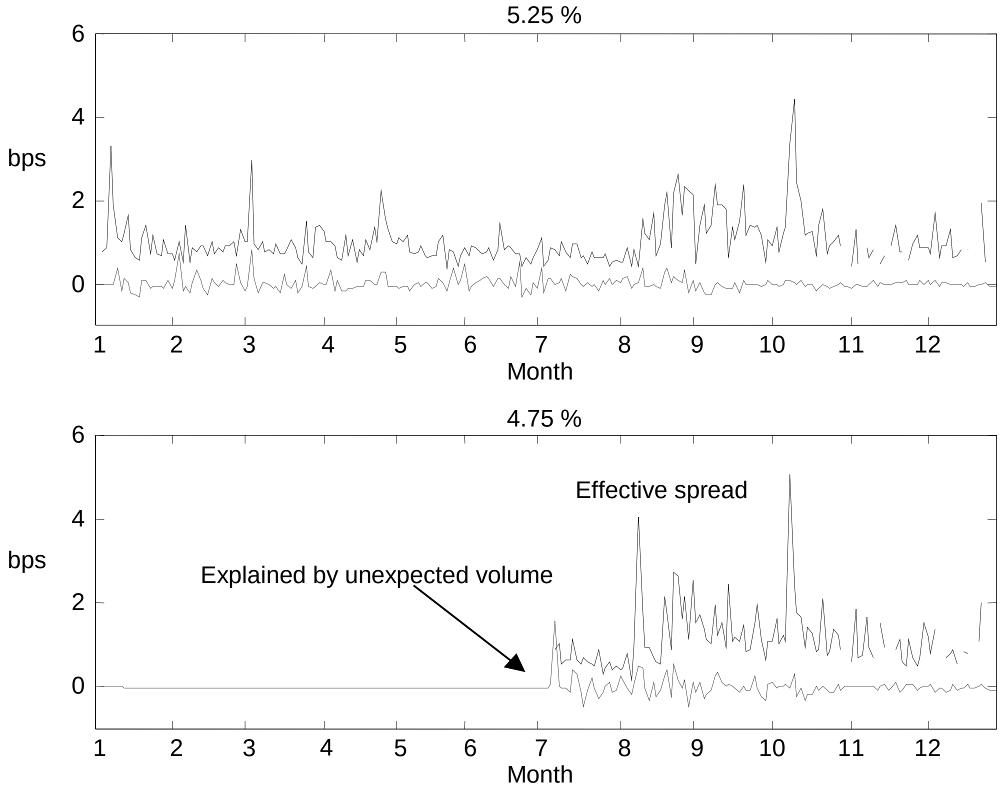
Sample: 27 May - 30 December, 1998, except for the 4.75%, which started on 6 July. Dependent variable: Roll measure for the effective spread, normalised by their mean. T-values in brackets. Volume: Total volume of all deliverable bonds. Expected Volume: 1-step out-of-sample forecasts from ARMA-model. Unexpected Volume=Volume-Expected Volume. Explanatory variables also include a constant. Newey-West correction for autocorrelation and heteroscedasticity. Significant on *90%, **95%, ***99% confidence level.

The relationship between unexpected volume and effective spreads is positive and highly significant for the 5.25% and the 4.75% bonds, which is in line with our hypothesis. The fact that the relevant coefficient is not significant for the 6% (July) bond may be due to the inordinate amount of noise in the series for the spread. The coefficients on expected volume are positive, but either insignificant or only weakly significant, which indicates that there do not seem to be much economies to scale in market making. The importance of inventory holding costs is shown by the highly significant and positive coefficient for expected volatility. We can reject the hypothesis that the coefficients of our two volume measures are equal for two of the bonds, which is in accordance with the results obtained for the forex market (see Hartmann [1999] and references therein).

27 Hartmann (1999) uses the quoting (tick) frequency on the Reuters FXFX page. Such quote information is not available to us, since our data only contains information on transactions.

Figure 11

Unexpected volume and effective spreads



The strong statistical relationship between unexpected volume and bid-ask spreads cannot explain the rise in the latter that occured from August to October. Figure 11 plots the spreads and the proportion that can be explained by unexpected volume ($\beta_{i2}uvol_{it}$). Spreads seem to be related to surprises in volume only in the very short run (a day) but not over prolonged periods of time. When the effective bid-ask spreads surged after the Russian devaluation and later during the uncertainty about the yen-carry-trades, unexpected volume remained essentially flat. This is not consistent with the hypothesis that the fear of adverse selection caused market makers to widen their spreads.

7.3 Did banks destabilise the market?

More generally, did market makers on average play a stabilising or a destabilising role during the turbulences? In particular, did they buy when prices were falling and sell when they were rising, or was it the other way around? We explore this question by regressing the changes in inventory associated with each transaction on the first difference in yields as well as on dummies equal to one during each of the turbulent subperiods and zero otherwise, multiplied by the yield change. A negative coefficient would indicate that banks were selling into the falling market or buying when prices were increasing. Since inter-dealer trades do not have any effect on the inventory of the aggregate banking sector we omit them and confine our analysis to client trades only.

Table 8
Changes in inventories and prices

	6% (January)	6% (July)	5.25%	4.75%
Δ yields	-0.008 (-0.11)	0.036 (0.47)	0.042 (0.77)	0.098 (0.69)
$D_{6.7.-14.8.} \times \Delta$ yields	0.058 (0.03)	0.017 (0.05)	0.002 (0.01)	-
$D_{17.8.-22-9.} \times \Delta$ yields	0.026 (0.11)	-0.107 (-0.70)	-0.004 (-0.04)	-0.095 (-0.53)
$D_{23.9.-15.10.} \times \Delta$ yields	0.001 (0.00)	0.021 (0.096)	-0.29 (-0.27)	-0.038 (-0.20)
$D_{16.10.-31.12.} \times \Delta$ yields	-0.096 (-0.12)	0.016 (0.08)	-0.054 (-0.34)	-
# observations	7,853	9,266	20,111	8,035

Client transactions only. Regression of changes in inventories on a constant and changes in yields. Variables standardised by dividing through standard deviation. T-values in brackets. Newey-West correction for autocorrelation and heteroscedasticity. Significant on * 90%, ** 95% confidence level

The results reproduced in Table 8 give no evidence on a correlation between changes in inventory and the first difference in prices. This indicates that market makers did not exacerbate price fluctuations by selling into falling markets or buying when prices were rising. On the other hand, they did not, on average, stabilise by behaving anticyclically. This is not really surprising given the small size of the inventory at most times.

8. Conclusions

The paper examines the liquidity of the secondary market for four German benchmark government bonds during the financial market turbulences in the summer and autumn of 1998. The analysis is based on a unique dataset provided by the German securities regulator, which covers every single transaction in these bonds in Germany.

The empirical findings of the paper can be summarised as follows:

The intra-day volatility of yields more than doubled in the wake of the Russian devaluation on 17 August and experienced a further peak on 8 and 9 October.

The yield spread between the individual bonds of our sample soared from less than five basis points during the first half of the year to more than twenty basis points in early October. No equivalent breakdown in the pricing relationship can be found at daily or hourly level, however. This indicates that the bonds remained close substitutes for hedging purposes even during periods of stress.

The surge in volatility and the building up of the yield spread occurred at a time of heavy trading activity. For three of the four bonds, both the number of trades and daily turnover were statistically significantly higher between 17 August and 15 October than during the first half of the year.

The effective bid-ask spreads (measured by Roll's [1984] approach) on the two most heavily traded bonds increased from just over one basis point to 1.8 basis points after the Russian de-facto default on 17 August, and rose above two basis points after the LTCM rescue operation. They peaked around 8 and 9 October at around five basis points. This rise is statistically significant even after controlling for changes in the average transaction size.

Effective bid-ask spreads are positively related to unexpected trading volume, which should reflect the amount of private information in the market. Nevertheless, surprises volume cannot explain the surge in spreads that occurred during the turbulences.

How does this answer our initial question on "how did the "safe-haven" weather the storm"? The finding that the effective spread increased considerably during the turbulences clearly indicates that the market suffered from a reduction in liquidity. This result is even more clear cut if we consider that our estimates for the effective spread based on the Roll [1984] measure are downward biased in the presence of inventory control and adverse selection effects, which are likely to be important particularly in times of great volatility.

Nevertheless, while agents may have faced a higher cost of trading, they were still able to trade. Our data does not tell us whether individual agents were able to turn over their position as quickly as they would have liked, but the fact that the market has been able to sustain a significantly higher than normal number of trades and volume suggests that such problems were not widespread, if they occurred at all. In this sense, liquidity provision has been remarkably effective in dealing with the turbulences, especially when seen in contrast to what happened in the US bond market.

Apart from the issue of financial stability, I mentioned two other reasons for why a central bank should monitor the liquidity in the secondary market for government bonds: the quality of indicators and the borrowing costs of government. Indicator quality clearly suffers if it becomes more costly to trade. Hence our findings provide further evidence for the already widely known fact that indicators are the least likely to be accurate at times when they are most needed.

The present paper considers only a small, albeit important, segment of the market. It would be interesting to extend the analysis to a) the futures market, which is far more active than the spot market, and b) bonds from other issuers. It would also be of interest to compute a measure of liquidity that is more robust to the presence of inventory control and adverse selection effects than the Roll measure. Both extensions are left for future work.

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FX impact of cross-border M&A

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- Globalisation has led to a steady increase in the number of cross-border M&A activities, and so these deals are increasingly discussed in the FX market. However, since M&A activity is still measured in billions of dollars per month, as compared with the trillion-plus a day traded in the FX market, it is not clear what impact these deals actually have in practice.
- Using a large sample of past M&A deals, we find that these deals do indeed have an impact on exchange rates, causing, on average, a 1% appreciation of the target nation currency relative to the acquirer. We also find that such currency moves can be predicted using publicly available data.
- From our results, we construct an M&A currency pressure indicator that can be used to help interpret and predict currency moves due to M&A.

Introduction

With globalisation leading to a steady increase in cross-border M&A activity, the latest deals have become a common topic of discussion in the FX market. However, cross-border M&A flows are still only in the billions of dollars per month, compared with FX turnover estimated at $1.5tr per day and so, despite the interest they generate, it is not clear whether these flows are indeed significant in practice.

Number of cross-border M&A, 1982-1994

	1982	1988	1994
UK–US	58	177	180
Euro-area[1]–US	26	67	256

[1] France, Italy, Germany. Source: Vasconcellos (1998).

This research note aims to answer three key questions:

1. Does M&A have a significant impact on currencies?
2. Can this impact be predicted using public information available at the time?
3. If there is an impact, how can information on M&A activity best be summarised?

Despite the interest these deals create in the FX market, there are no previous studies to draw on to analyse this question. Economists seem more interested in the economic motivation for M&A (see for example Kang [1993] and Vasconcellos [1998]), rather than their impact.

Estimating the impact of M&A

Although the aim is simple, studying the impact of cross-border M&A activity on currencies is often complicated by the intricate financing structure of such deals. We use a event study method (see, for example, Campbell, Lo and MacKinlay [1997]) with which we test whether the returns to holding the target nation's currency relative to that of the acquirer where cross-border M&A deals have been announced is significantly different from zero. By averaging over a large number of similar deals, it can be determined if there is a consistent excess return.

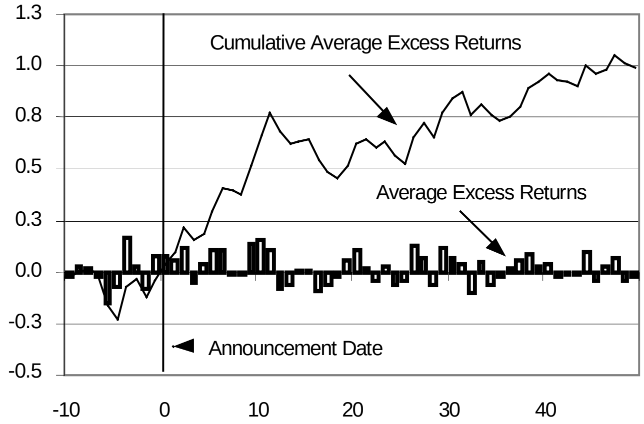
The first step to estimating excess returns is to construct a sample of deals to analyse. Our sample is based on information published in the IFR DataBase and includes all transactions that took place in 1997-99 and that are greater than US$400m. For simplicity, we concentrate on a subset of possible country pairs: US-euro area, US-UK, US-Canada and US-Japan. The table illustrates some sample statistics.

Cross-border M&A sample (Transactions 1997-99, >US$400m)		
	Avg size in US$ m	**Number**
US-Euro area	4,076	38
US-UK	9,855	33
US-Canada	3,140	19
US-Japan	3,128	7

The graph below shows the key result from our event study. Taking an equally weighted average of all the deals in our sample, we construct average daily excess returns, and the cumulate of those excess returns over a period starting 10 days before an M&A deal is announced until 50 days after.

As the chart shows, our results seem to support the hypothesis that M&A does have a significant impact. In particular, the period shortly after a deal is announced shows a strong upward movement in the target nation's currency (relative to the acquirer's currency). This strengthening seems to persist, such that, 50 days after the announcement, the target nation's currency, on average, has appreciated by 1% relative to the acquirers (ie, is 1% stronger than the forward rate 10 days before the announcement would have predicted).

Returns around cross-border M&A



The next task is to measure whether such a rise is statistically significant over the time period. For event studies, the J-statistic is used to assess the significance of the returns:

$$J_1 = \left(\frac{\overline{CAR}}{\sigma^2} \right)^{1/2} \sim N(0.1)$$

$$\overline{CAR} = \frac{1}{N} \sum_{i=1}^{N} ExCumReturns_i$$

The J_1 statistic follows a standard normal distribution. N is the total number of cross-country transactions being analysed, while σ is variance of the cumulated average returns. Furthermore, *ExCumReturns*$_i$ represent the cumulative average excess return at each moment in the 60-day window. Therefore, the J_1 statistic aggregates across countries, but not across time.[1]

1 The results are based on simple averages of returns without taking into account the size of the deal. One reason for using the simple average is that the J_1 statistic is not appropriate for weighted-averages.

Significance tests for M&A impact

Days from announcement	Cumulative returns	J1 Stat	Prob not 0
-10	-0.033	-0.06	0.48
0	0.025	0.05	0.52
5	0.308	0.76	0.78
10	0.671	1.28	0.90
20	0.653	1.37	0.91
30	0.862	1.81	0.97
50	1.006	2.46	0.99

The table above indicates that our result is statistically significant. With this key result, we go on to analyse four detailed aspects of M&A:

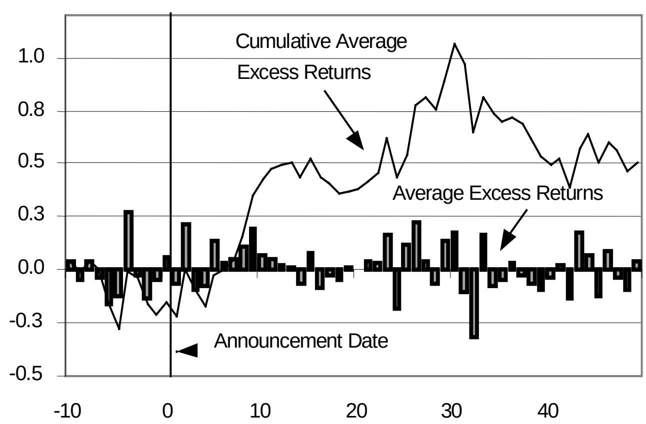
1. *Individual country experience*. Does the M&A impact on currencies differ across countries?
2. *M&A-financed through stock swaps*. Do M&A deals that do not involve cash influence currencies?
3. *M&A from a third currency perspective*. Is the rise in the target currency relative to the acquirer currency also a rise relative to a third currency, or does the acquirer's currency fall too?
4. *Deal Size effects*. Do large deals have a larger currency impact?

Individual country experiences

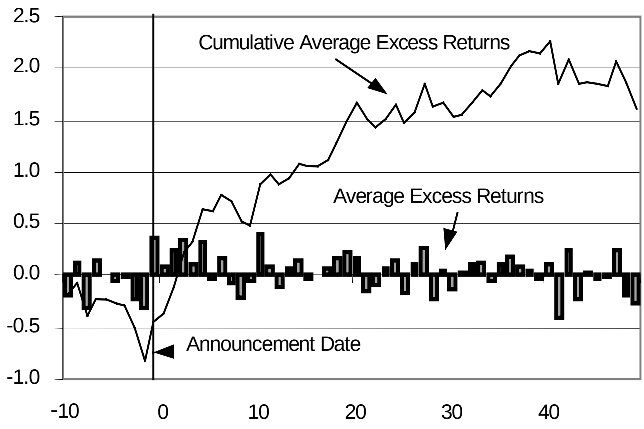
The graphs below show individual country profiles. As can be seen, for the following country couplings, US–euro area, US–UK and US–Canada, the pattern of cumulated excess returns is similar to the average. Only in Japan's case is the profile of returns different. Instead of showing an initial rise followed by a levelling in cumulative returns after the announcement of an M&A deal, there is much more volatility in returns. However, this results probably simply reflect our small sample of US-Japan deals (only seven were recorded).

Another interpretation, however, is that M&A deals have less impact the larger the countries involved. Certainly, our results for the euro area display lower returns than in the cases of the UK and Canada. Unfortunately, our sample is too small to test this proposition rigorously, but it does seem intuitively plausible.

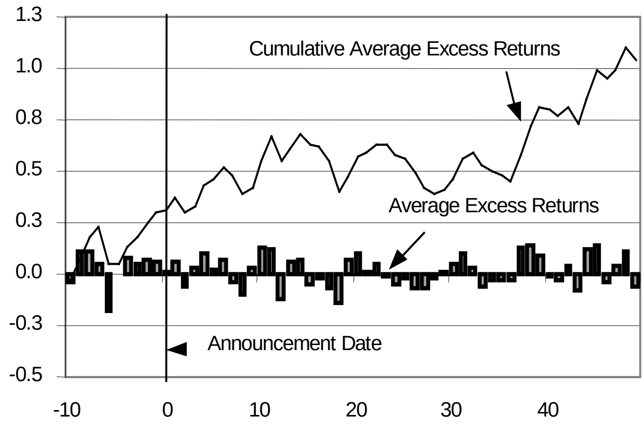
Currency returns: US–Europe



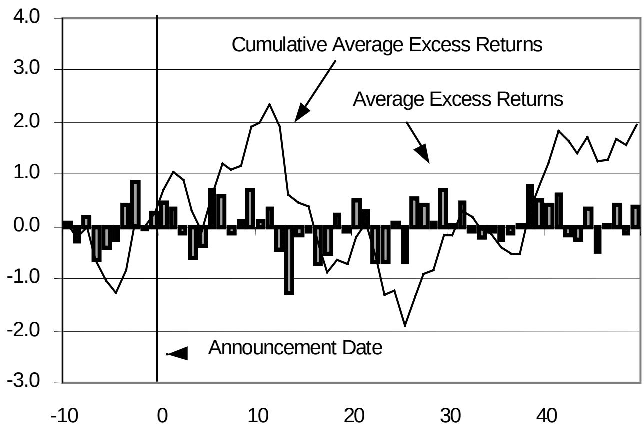
Currency returns: US–Canada



Currency returns: US-UK



Currency returns: US-Japan

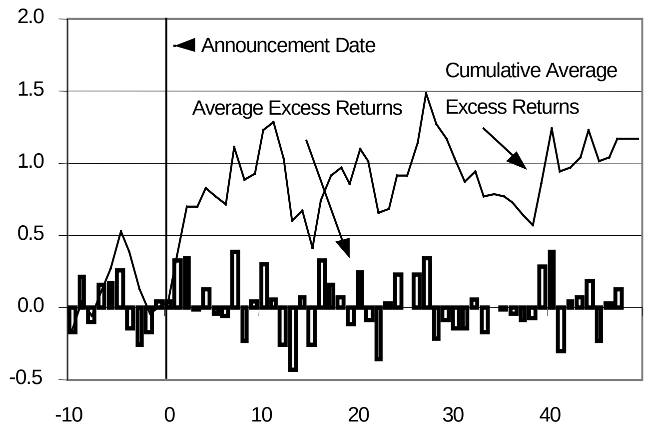


Stock swaps and M&A activity

Although the analysis so far has included M&A deals financed through stock swaps, many argue that it is the cash element of an M&A deal (ie, the acquirer’s need to buy FX to pay for the target) that influences the currency. What sets stock swaps apart from other ways of financing M&A activity is that there should be no foreign exchange transaction between the countries involved in the M&A.

Therefore, *a priori*, we expected no consistent pattern of either average or cumulative average excess returns. However, by looking at the evidence of the seven M&A deals financed via stock swaps in our sample, we were surprised to find that indeed cumulative positive excess returns are present (although not statistically significant).

Currency returns: stock swaps

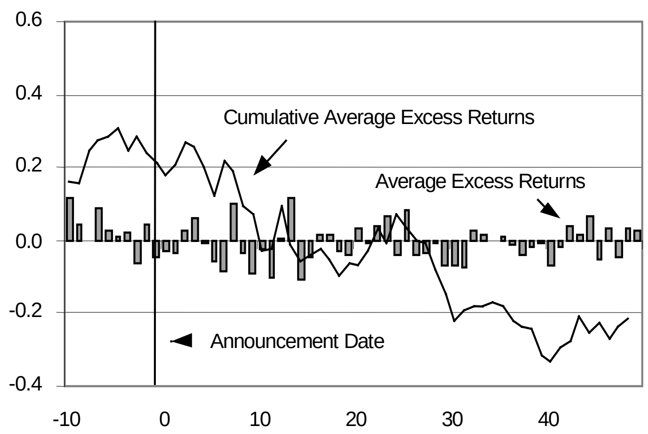


M&A from a third-currency perspective

While our results show how the target nation currency rises relative to the acquirer, it is not clear which of those currencies is actually causing that excess return. Is the target currency going up or the acquirer's going down?

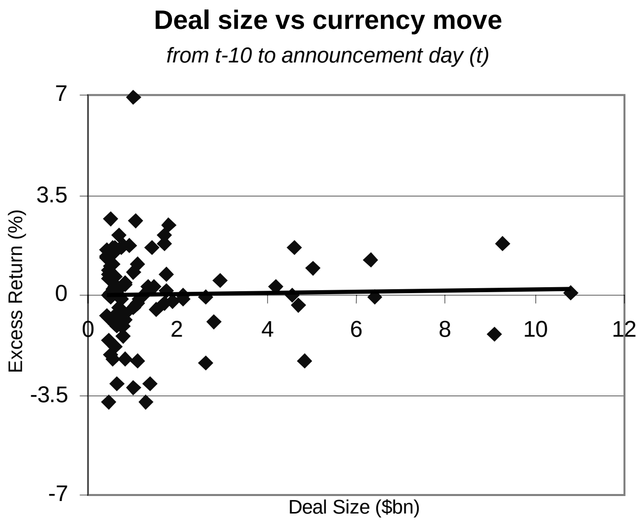
The chart below shows the acquirer's currency relative to SDRs. It indicates negative returns of about 0.5% from announcement day to 50 days post-announcement. Comparing this result with our 1% return on target relative to acquirer indicates that, from a third-currency perspective, an M&A deal results in equal and opposite moves in the target and acquirer currency.

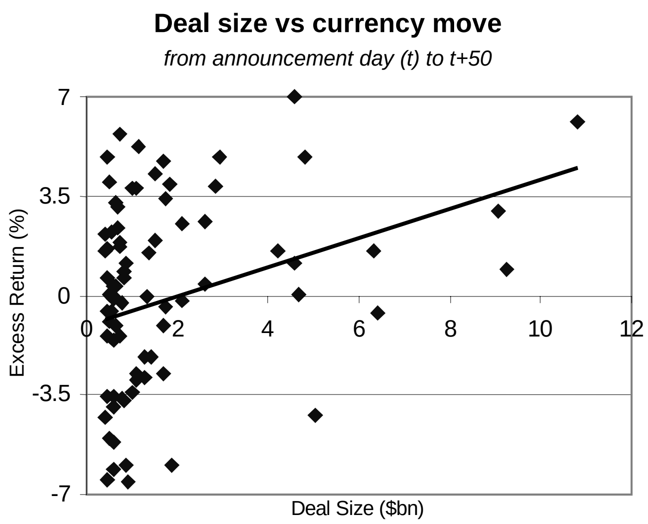
Currency returns: SDRs - acquirer nation



Deal size effects

While the results so far indicate a significant impact of M&A deals on currencies, the event study framework does not give an indication of how this effect changes with deal size. The two scatter plots below show the relationship between deal size and currency impact for two distinct periods. The first looks at the period from 10 days before announcement day to announcement day to see if there is a significant pre-announcement day effect for larger deals. The second chart compares deal size with currency impact for the full post-announcement period (announcement day to 50 days after).





The results of this comparison are fairly encouraging. They show little pre-announcement effect even for big deals, but a fairly clear relationship between deal size and currency impact over the post-announcement period. Roughly speaking, in our sample, every $1 billion has a 0.5% impact. However, separate results for three mega-deals (>$25 billion) show that the currency impact does seem to peak at about 5%

An indicator of the impact of M&A activity

Using our results, we have constructed a simple indicator of M&A activity that can be used to help interpret and predict currency moves. The indicator keeps track of recent deals and gives an estimate of currency pressure in percentage basis points that those deals put on the currency. Following our third country results, we give a negative value to an acquirer nation and positive to a target. The table below shows recent coefficients for this indicator.

Currency pressure coefficients for $6bn deal

Days from announcement	<20	<50
Coefficients for target	0.03	0.01
Coefficients for acquirer	-0.01	-0.01

These coefficients represent the expected daily return of being long of the target nation currency when a deal averaging US$5.8bn is announced. Following the results on deal size, when constructing this indicator we give a larger weight to larger deals (though with a cutoff for "mega-deals").

Given that our results for individual countries implied – but did not prove – that an M&A deal of a given size has a smaller impact the larger the countries involved, we also present a GDP-weighted version of our indicator.

M&A currency pressure indicator – 6 July 1999

(% bp per day)

Country	Unweighted	GDP-weighted
Australia	0.34	4.94
Canada	0.48	4.55
Chile	0.16	12.22
Czech Republic	0.10	10.47
Denmark	-0.17	-5.60
Euro	-1.25	-1.08
Greece	-0.75	-33.02
Norway	0.41	14.60
Poland	0.07	2.72
Singapore	0.08	4.20
South Korea	0.14	1.66
Sweden	0.18	4.22
Switzerland	-0.52	-9.34
United Kingdom	1.15	5.25
United States	-0.41	-0.30

Conclusion

As well as confirming the widely held market belief that cross-border M&A activity does influence exchange rates, our analysis shows:

1. The appreciation tends to occur after the public announcement of the deal and so does not rely on inside information.
2. Statistical tests confirm that this excess return is statistically significant (at the 99% level) and so is probably not just a quirk of our data.
3. Separate tests on deals that do not involve cash (ie, stock swaps) indicate that these too have a currency impact.
4. From a third-currency perspective, currency moves around M&A tend to lead to a fall in the acquirer currency and a rise in the target currency of equal size.

Using our results, we are now in a position to monitor M&A effects on currencies which are clearly an important supplement to standard economic analysis.

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