BANK FOR INTERNATIONAL SETTLEMENTS

BIS Quarterly Review

March 2007

International banking and financial market developments

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This publication is available on the BIS website (www.bis.org).

ISSN 1683-0121 (print)

ISSN 1683-013X (online)

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Notations used in this Review

e	estimated
lhs, rhs	left-hand scale, right-hand scale
billion	thousand million
…	not available
.	not applicable
–	nil
0	negligible
$	US dollar unless specified otherwise

Differences in totals are due to rounding.

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Overview: markets rally until late February[1]

Prices of risky assets rallied between end-November and late February as the outlook for economic growth appeared to improve, while implied volatilities remained near record lows. In this environment, yields rose in major government bond markets, and perceptions grew among investors that monetary policy might turn out to be tighter in the foreseeable future than previously expected. In the United States, data releases indicated surprising strength in the economy, in particular during the first half of the period under review, leading to subsiding expectations among investors of a near-term lowering of policy rates by the Federal Reserve.

Corporate profitability and the ongoing strength of merger and acquisition (M&A) activity contributed to rallies in global equity markets. At the same time, spreads on risky corporate debt fell to all-time lows during the period, reflecting strong investor risk appetite, sound corporate balance sheets and surprisingly low default rates, particularly for higher-yielding credits. Spreads in some collateralised debt obligation (CDO) markets, mainly those centred on the housing sector in the United States, widened significantly over the past two months, possibly foreshadowing a broader turn in the credit cycle in the months to come.

As in high-yield credit markets in advanced economies, spreads on emerging market debt hit historical lows in the first two months of 2007, while equity prices continued to increase. Local events affected some individual countries negatively, but seemed to have little overall effect on investors' perceptions of emerging asset markets up until the last week of February. Instead, investors were largely anticipating continued strength in emerging economies in general, as well as an improving outlook for the US economy. The strong appetite for risk among investors is likely to have been another important factor behind asset price developments in emerging markets during the period under review.

Bond yields rise on strong economic data

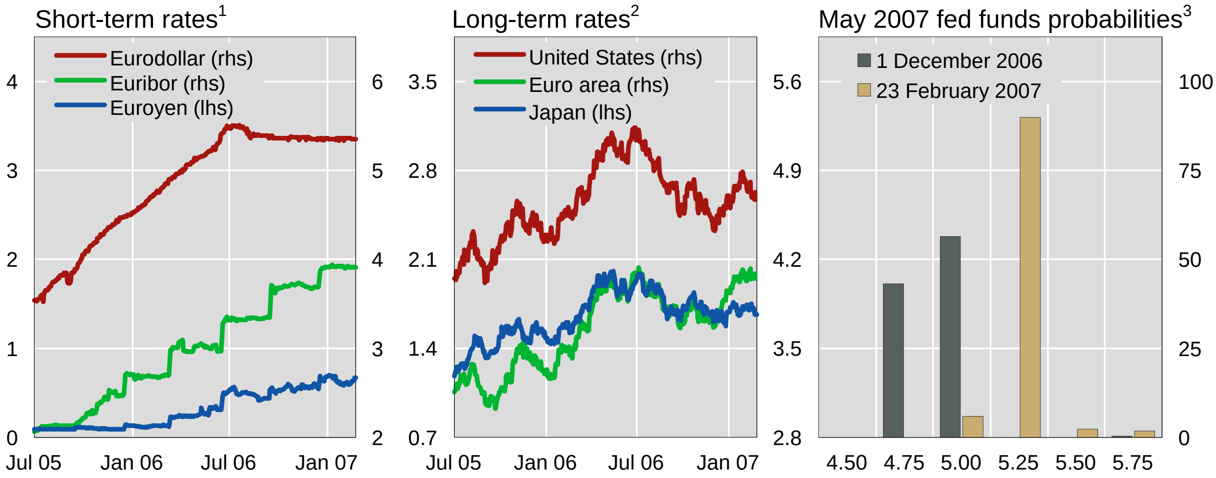
Yields on government bonds rise …

Government bond yields in major industrialised economies rose towards the end of 2006 and in early 2007. After having reached its lowest level in almost a

[1] The period covered in this Overview is from end-November 2006 to 23 February 2007.

Interest rates

In per cent



[1] Three-month rates. [2] Ten-year government bond yields; for the euro area, German bund. [3] Implied risk neutral probabilities for the fed funds rate following the 9 May 2007 FOMC meeting, as implied by prices of options contracts on fed funds futures on the dates indicated in the graph. The probabilities are backed out along the lines suggested by J Carlson, B Craig and W Melick, "Recovering market expectations of FOMC rate changes with options on federal funds futures", *Federal Reserve Bank of Cleveland Working Paper*, 05-07, July 2005.

Sources: Bloomberg; BIS calculations. Graph 1

year at the start of December 2006, the 10-year US Treasury bond yield subsequently increased by almost 50 basis points to 4.90% at the end of January, before retreating to around 4.70% by late February (Graph 1, centre panel). In the euro area, long-term bond yields followed suit, with the 10-year German bund yield rising almost 40 basis points to 4.05%. Yields also edged up in Japan, but by substantially less than in Europe and the United States. By late February, the Japanese 10-year bond yield stood at just below 1.70%, less than 10 basis points higher than at the start of December last year. Short-term money market interest rates remained steady in the United States, reflecting the unchanged monetary policy stance of the Federal Reserve, while they rose in the euro area and Japan, where policy rates were tightened (Graph 1, left-hand panel).

A factor contributing to the rise in bond yields, particularly in the United States and to some extent in Europe, was a growing perception among investors that monetary policy might turn out to be tighter than previously expected. At the beginning of December, prices on federal funds futures options suggested that markets considered a 25 basis point rate cut by the Federal Reserve a near certainty in the first five months of 2007, and that the probability of two 25 basis point cuts during that period was high (Graph 1, right-hand panel). However, by late February, the options market suggested that an easing of policy rates by May 2007 was almost fully ruled out, with the probability of no change standing at 90%. Moreover, a pronounced upward shift in the entire fed funds futures curve indicated that the odds for a rate cut in the second half of the year had fallen considerably (Graph 2, left-hand panel). The market's assessment that policy rates would remain steady for some time was reinforced by the FOMC statement on 31 January and the

… as expectations of a "certain" Fed rate cut evaporate

Federal Reserve Chairman's testimony to Congress two weeks later, which acknowledged that economic growth was firmer, but which also pointed to improving core inflation figures. In the euro area, the ECB delivered an expected 25 basis point rate hike on 7 December, and left policy rates unchanged in January and February. Developments in the EONIA forward rate curve suggested that by late February markets expected monetary policy in the euro area to be tightened more in the course of 2007 than had been anticipated in November 2006 (Graph 2, centre panel).

The Bank of England surprises markets …

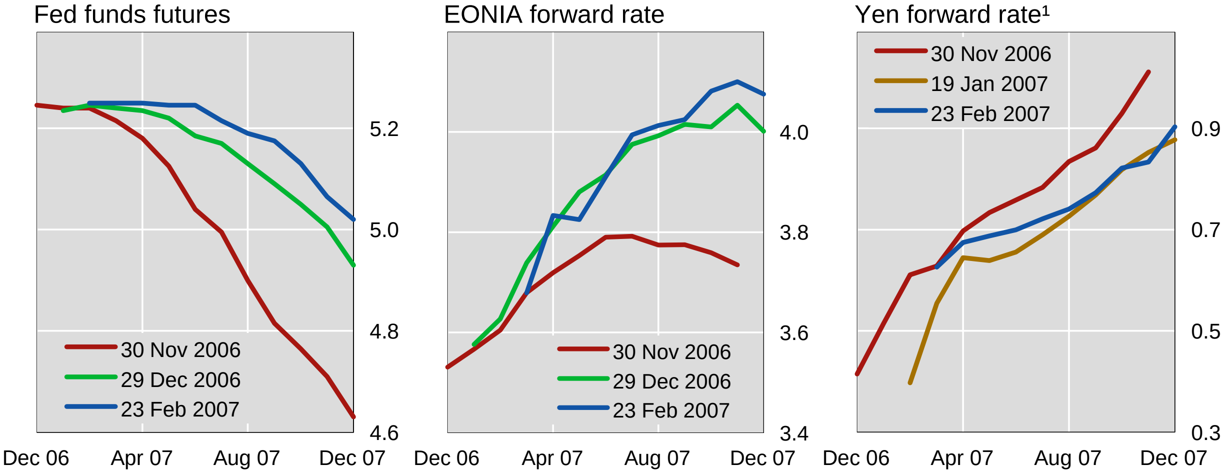
The period under review saw a couple of monetary policy decisions that surprised markets. The Bank of England tightened monetary policy on 11 January in a move that was almost entirely unanticipated by investors, and which resulted in a significant rise in UK bond yields, and even some spillover effects beyond the United Kingdom. The decision reflected data that had been made known to the Bank of England, but not yet to markets, indicating that UK CPI inflation had reached 3%, ie fully 1 percentage point above the Bank's inflation target.

… as does the Bank of Japan

In Japan, in the days before the Bank of Japan's monetary policy meeting on 17–18 January, market expectations indicated a probability of a rate hike around 70–80%, as measured by money market rates as well as by analyst surveys. However, on the day before the announcement of the decision, news reports began to circulate that a policy tightening was unlikely, causing bond yields to fall considerably. The subsequent official announcement that interest rates would remain unchanged thus had little impact on bond yields. Implied volatilities on short-term yen interest rates rose after the monetary policy decision became known, while interest rates fell and implied forward rates shifted downwards. The segment of the implied forward curve corresponding to the second half of 2007 was little changed following the Bank of Japan's decision to raise its benchmark interest rate by 25 basis points to 0.5% at its

Forward curves

In per cent



[1] Implied one-month rates; calculated from Libor fixing rates.

Sources: Bloomberg; BIS calculations. Graph 2 subsequent monetary policy meeting on 21 February (Graph 2, right-hand panel).

Upbeat economic news …

The link between perceptions about the strength of economic activity and market expectations about monetary policy was evident both in the United States and in the euro area. A number of data releases, in particular in December and January, indicated surprising resilience of the US economy, which, in combination with falling oil prices, underpinned a shift in investors' expectations towards a more optimistic outlook. In the euro area too, data releases continued to paint an upbeat picture with respect to current and future economic activity. As the positive news accumulated, US and euro area nominal bond yields rose, largely reflecting rising real yields (Graph 3, left-hand panel). Towards the end of the period under review, however, a partial recovery of oil prices as well as some less favourable data releases seemed to dampen investors' optimism about the US economy, which in turn prompted a partial retreat of US bond yields.

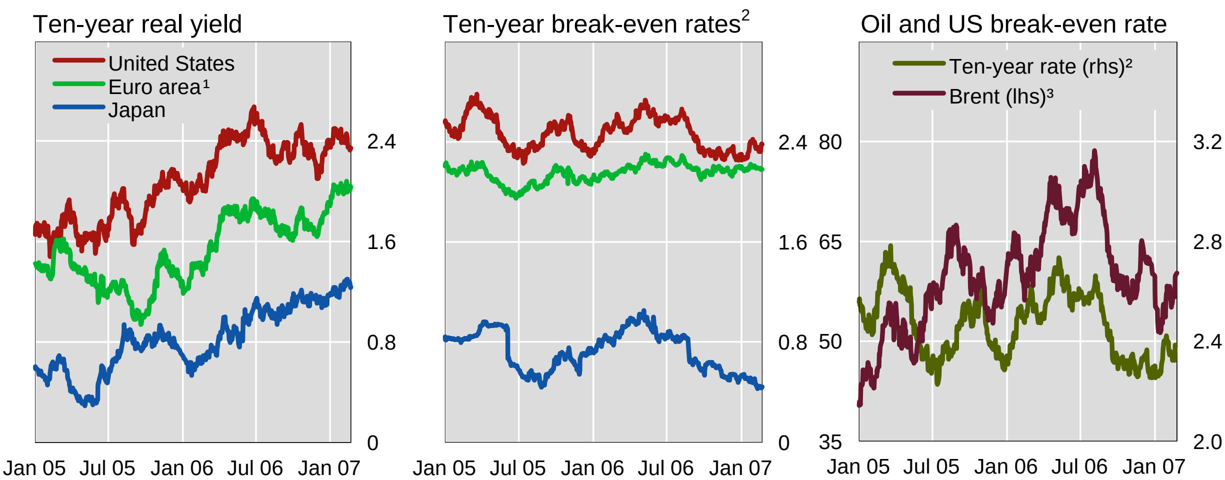
… lifts real and nominal yields …

The rise in nominal bond yields in major industrialised economies since end-November was to some extent restrained by incoming inflation data, which mostly indicated stable or reduced price pressures. Break-even inflation rates remained fairly stable in Europe and the United States (Graph 3, centre panel). Falling oil prices might have contributed in this regard, at least until they started to recover in mid-January (Graph 3, right-hand panel). However, break-even inflation rates in the United States rose somewhat following the release of higher than expected US inflation figures in the second half of February.

… while leaving break-even rates little changed

Estimates of nominal term premia in the US term structure of interest rates suggest that part of the increase in bond yields was due to premia increasing from the extraordinarily low levels reached at the beginning of December (Graph 4, left-hand panel). The rise in the 10-year premium was almost fully mirrored by an increase in the two-year segment, suggesting that the premium investors demanded to bear interest rate risk had increased in parallel across

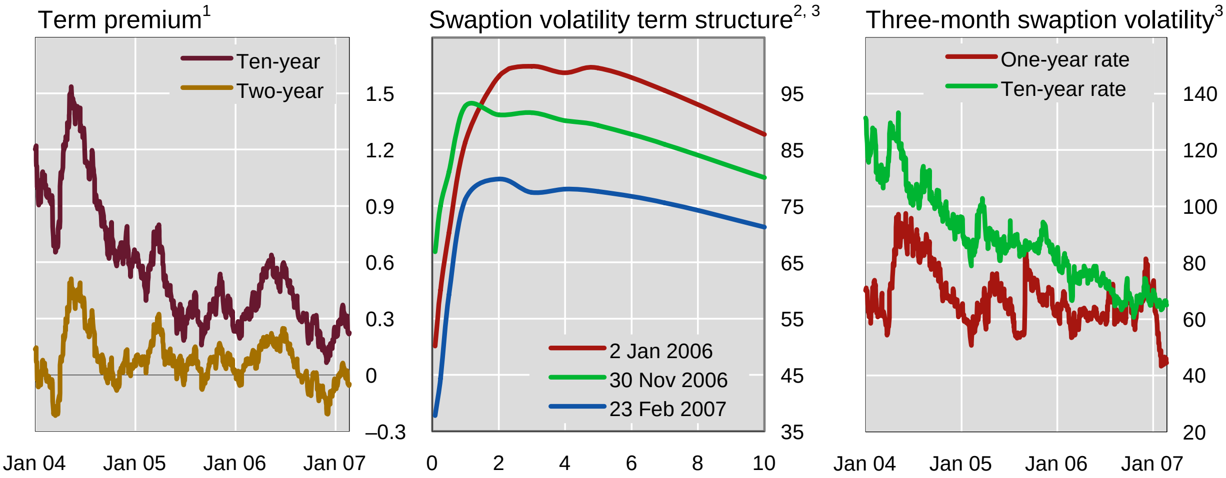
Real yields, oil price and break-even inflation



[1] Based on French government bonds linked to euro area inflation, in per cent. [2] Nominal minus real government bond yields, in per cent. [3] In US dollars.

Sources: Bloomberg; national data. Graph 3

US term premia and swaption volatility



[1] US term premium in per cent, based on D Kim and J Wright, "An arbitrage-free three-factor term structure model and the recent behavior of long-term yields and distant-horizon forward rates", *International Finance Discussion Paper*, 2005-33, Federal Reserve Board, August 2005. [2] Term structure of implied volatilities on one-year swap rates. The horizon on the x-axis is measured in years. [3] In basis points per year.

Sources: Bloomberg; BIS calculations. Graph 4

the entire term structure. However, rising premia accounted for only around half of the overall increase in nominal yields, implying that a large part of this rise was due to upward revisions of the expected trajectory of future interest rates by investors.

Implied interest rate volatilities drop to new lows …

Uncertainty about the outlook for short-term interest rates fell to new lows in January and February, as measured by the three-month implied volatility on swaptions on one-year swap rates (Graph 4, right-hand panel). By contrast, the implied volatility on 10-year swap rates fell considerably less than that on one-year rates, suggesting that a particularly pronounced reduction in uncertainty may have taken place at the short end of the maturity spectrum. Possibly, this was a result of growing perceptions among investors that the Federal Reserve would be likely to keep interest rates on hold for some time. This was also reflected in a substantial narrowing of implied fed funds probability densities over possible outcomes of future rates during the period under review (Graph 1, right-hand panel).

… across all horizons

The fall in implied volatilities was not limited to near-horizon swaptions only. The entire term structure of implied swaption volatilities on one-year swap rates shifted downwards between end-November 2006 and late February 2007 (Graph 4, centre panel). While for horizons beyond two years this represented a continuation of the decline in implied short-term interest rate volatility that had been ongoing for much of 2006, for shorter horizons the decline since end-November was a reversal of earlier increases in the second half of 2006. The downward shift in the term structure of volatilities since November was essentially parallel across the entire spectrum up to 10 years ahead, suggesting either that reduced uncertainty about short-term interest rate developments was perceived as a structural and hence persistent phenomenon, or that the decline in implied volatilities in part reflected a reduction in markets' required compensation for volatility risk across all horizons.

M&A activity buoys equity markets

Global equity markets continued to rally during the period under review, with markets in Japan and Europe outperforming those in the United States (Graph 5, left-hand panel). The TOPIX index rose by 13% between end-November 2006 and late February, reaching a 15-year high during the period. Similarly, the European DJ STOXX index rose by 9%, and the S&P 500 by 4%, both reaching six-year highs.

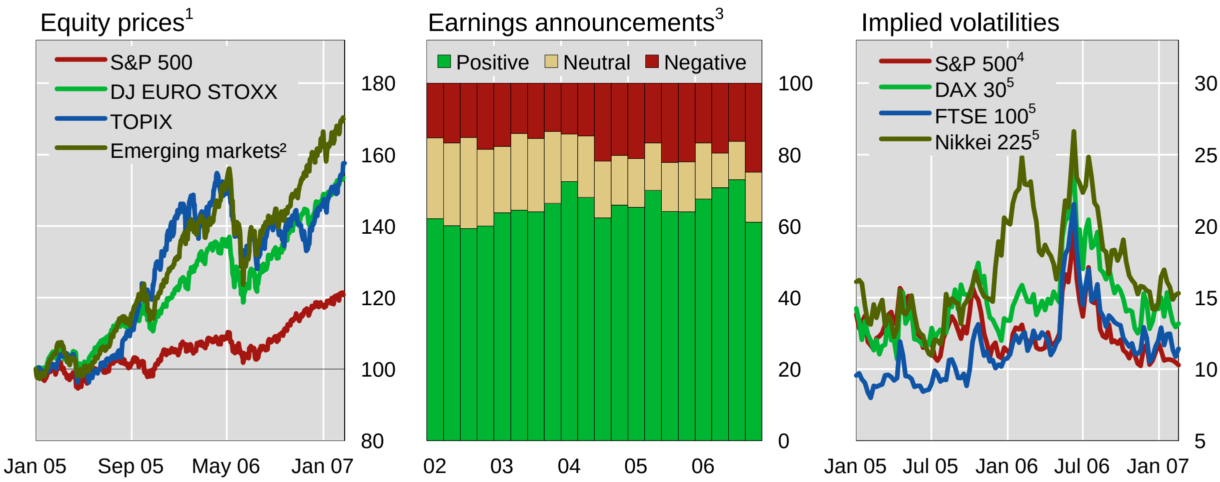
S&P 500 reaches a six-year high

Equity markets in the United States took their cue from incoming information on whether fourth quarter corporate earnings growth would be as robust as in previous quarters. But market participants also reacted to signals of a possible slowdown in the US housing sector. The S&P 500 briefly touched a six-year high on 24 January, after Yahoo! and Sun Microsystems reported better than expected earnings, only to reverse these gains the next day following a worse than expected existing home sales figure and disappointing earnings from other bellwether companies. The index had recovered by early February, boosted in part by positive earnings news, but temporarily retreated following announcements on 8 February by HSBC Holdings and New Century Financial Corporation which pointed to a deterioration in the subprime mortgage loan market in the United States (see below).

Earnings growth slows in the United States

Overall, US corporate earnings for the fourth quarter of 2006 exceeded analysts' expectations, although growth seemed to slow relative to previous quarters. With the reporting season for companies in the S&P 500 near completion by late February, positive surprises again outnumbered negative, but by a smaller margin than in previous quarters (Graph 5, centre panel). The

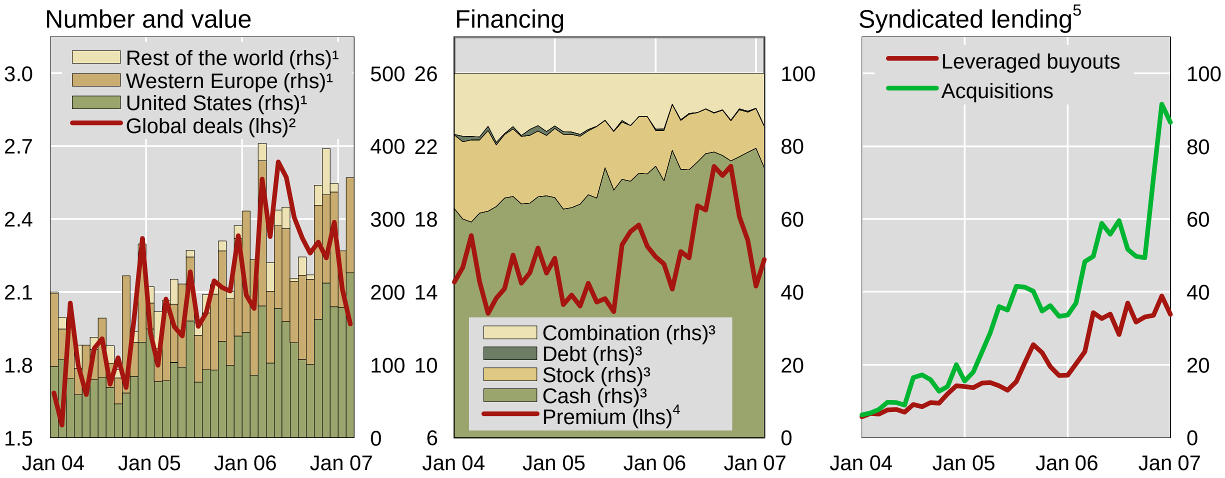
Equity markets and corporate earnings



[1] In local currency; 3 January 2005 = 100. [2] Morgan Stanley Capital International index. [3] Based on 12-month trailing earnings per share; companies in the S&P 500 Index, in per cent. [4] Chicago Board Options Exchange VIX volatility index; weekly averages, in per cent. [5] Volatility implied by the price of at-the-money call options contracts on stock market indices; weekly averages, in per cent.

Source: Bloomberg. Graph 5

M&A activity



[1] Value of announced mergers and acquisitions, in billions of US dollars. [2] Number of global deals, in thousands. [3] Share of global M&A activity financed by instrument, in per cent. [4] Average premium; three-month moving average, in per cent. [5] In billions of US dollars; three-month moving averages.

Sources: Bloomberg; Dealogic Loanware; Standard & Poor's; national data. Graph 6

share of companies which reported positive surprises (61%) was the lowest since the fourth quarter of 2002, while the share reporting negative surprises (25%) was the highest since the third quarter of 1998. Moreover, aggregate earnings growth for the fourth quarter, at just under 10% (on a share-weighted basis), was considerably less than the 15% or more seen at a similar stage in the previous three earnings seasons.

Generally strong corporate earnings also buoyed equity markets in Europe, as did positive data about euro area and US macro developments. Positive earnings news and rumours of corporate takeovers led to 10 consecutive daily advances for the DJ STOXX index in the first half of December, the longest such rally since 1997. Although incoming earnings news at times dampened investors' enthusiasm, equities continued to trend upwards through mid-February, with the index hitting a six-year high on 24 January.

Yen depreciation supports Japanese equities

After a lack of direction in the third quarter of 2006, Japanese equity markets finally found their footing in late November. While strong corporate earnings were key, the rally in part also reflected the weakening of the yen over the period under review, allegedly caused by carry trade positions which put downward pressure on funding currencies (see the box on page 8). The rally started after 22 November, when the yen hit a record low against the euro, and ultimately propelled the TOPIX index to a 15-year high by 21 February. Firms in the iron and steel and maritime transportation sectors saw the largest gains, although automobile and consumer electronics firms were also amongst the top performers during the period.

News of mergers and acquisitions also helped to support equity markets in both Europe and the United States. The announced purchase of Mellon Financial by the Bank of New York on 4 December, for example, helped to boost the S&P 500 Index by 1%. On 7 February, a private equity group, Blackstone, won the bidding war for Equity Office Properties Trust with its

Detecting FX carry trades

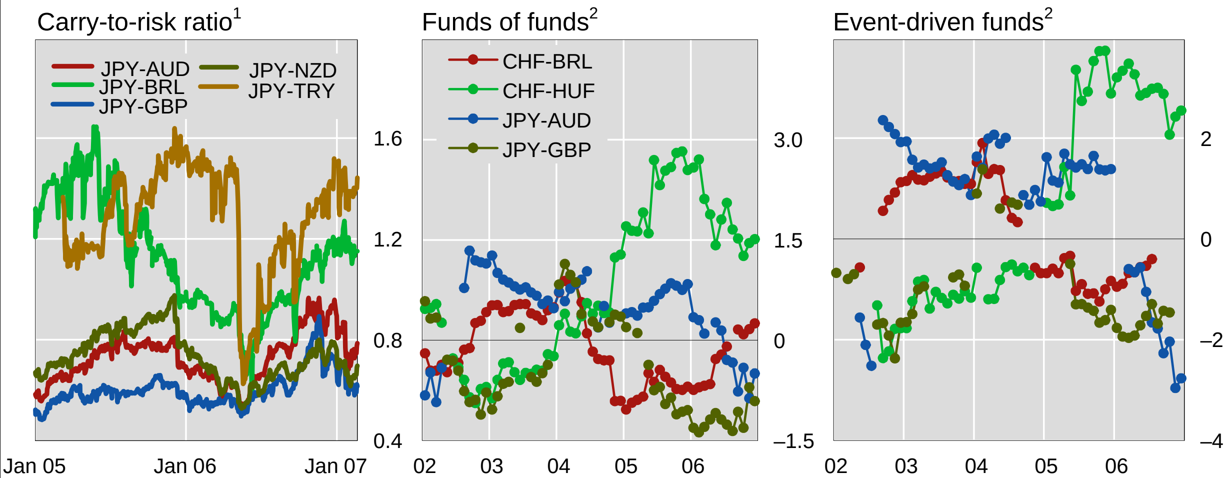
Patrick McGuire and Christian Upper

Many market participants have attributed the recent weakness of the Japanese yen and the Swiss franc to a pickup in carry trades funded in these currencies. However, measuring the volume of carry trades on the basis of publicly available information is problematic, both because of incomplete data, which makes it difficult to distinguish carry positions from other trades, and because of a lack of consensus on what exactly constitutes a carry trade.[①] This box draws on various sources of data in an attempt to measure whether carry trade activity has been high in recent months. Overall, the evidence is mixed. Data on positions in the derivatives market are broadly consistent with growth in activity, in particular for trades funded in yen, while data on cross-border bank lending are more difficult to interpret. Similarly, hedge fund returns appear to be sensitive to carry trade payoffs, but the results are far from conclusive.

The expected payoff of a carry trade depends on the interest rate differential and the likelihood of adverse exchange rate movements. The carry-to-risk ratio[②] – an ex ante measure of the attractiveness of specific currency pairs – indicates that yen-funded carry trades targeting emerging market currencies (eg the Brazilian real or the Turkish lira) have become increasingly attractive since mid-2006, whereas those targeting the currencies of more advanced economies have lost some of their lustre in recent months (Graph A, left-hand panel). Carry trades funded in francs show similar patterns, although they generally have lower carry-to-risk ratios than trades funded in yen.

Hedge funds are reportedly heavily involved in carry trades. Indeed, style analysis regressions suggest that hedge fund returns are partly driven by carry trade payoffs.[③] For many hedge fund families, including funds of funds and event-driven funds, proxies for the ex post payoff of various carry trade positions turn out to be statistically significant, although their overall net effect on hedge fund returns does not seem to have increased recently (Graph A, centre and right-hand panels). In some cases, the estimated coefficients are negative, perhaps reflecting positioning in expectation of an unwinding of the carry trade, or on the relative performance of different currency pairs.

Hedge fund returns and carry trade payoffs



[1] Three-month interest rate differential divided by the implied volatility of the currency option; the funding currency is the Japanese yen. [2] Coefficients from 24-month rolling stepwise regressions of hedge fund returns on market indices and proxies for carry trade payoffs; the first currency in the pair is the funding currency.

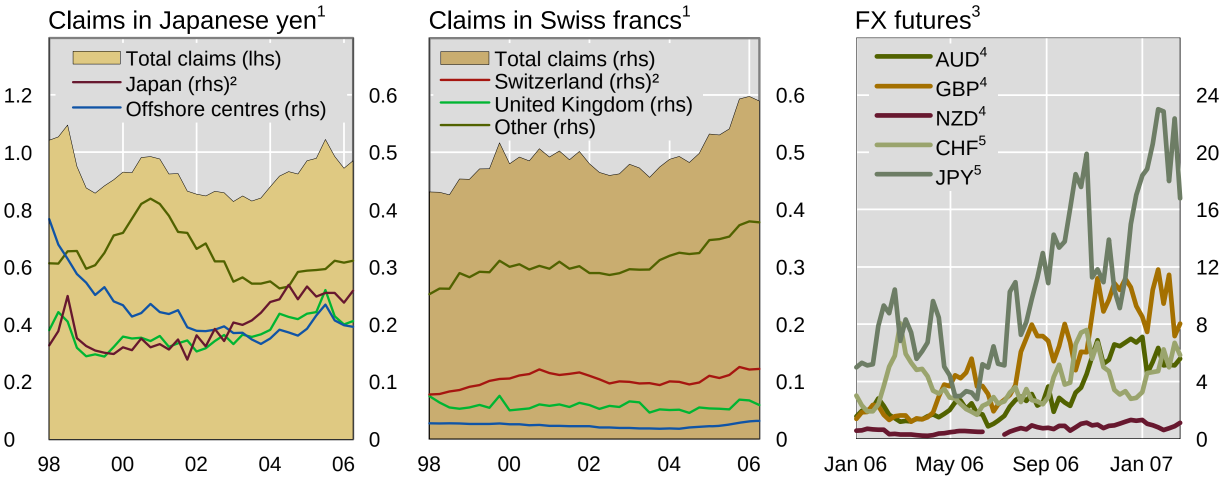
Sources: Datastream; HFR; JPMorgan Chase; BIS calculations. Graph A

Other sources of data might also throw light on this issue. Carry trades can be done through outright borrowing and lending or through derivatives, which are often hedged in the cash market, thus potentially leaving footprints in the BIS international banking statistics. However, the extent to which movements in these data reflect carry trade activity is difficult to quantify since global claims flows reflect many types of economic activity. A pickup in claims on residents in financial centres, where many hedge funds or proprietary trading desks are located, might arguably be more likely to reflect carry trade activity, at least in the narrow sense of the term, than a similar rise elsewhere in the reporting area. With this in mind, the evidence for a recent pickup in carry trade activity in the

BIS international banking statistics is mixed (Graph B, left-hand and centre panels).④ The rise in the stock of outstanding yen-denominated claims in the second half of 2005 did in part reflect greater credit to residents of the United Kingdom and offshore centres; however, claims have since fallen. Swiss franc claims grew in the first half of 2006, although claims on borrowers in these financial centres have remained relatively small. The rise in claims in the first half of 2006 was primarily the result of greater lending to residents in the euro area.

Data on open positions in exchange-traded FX futures in potential funding and target currencies provide the strongest evidence for a growth in carry trade activity in recent months. Non-commercial ("speculative") short positions in yen futures traded in the United States rose between mid-2006 and late February 2007, particularly during periods of yen depreciation (Graph B, right-hand panel). By contrast, speculative short positions in the franc yield little evidence of an increase in futures-based carry trades over this period. Data on speculative long positions in FX contracts on the main developed-country target currencies increased considerably in the second half of 2006, but declined somewhat in early 2007 (Graph B, right-hand panel), consistent with the rise and subsequent fall in the carry-to-risk ratio over this period. However, the weekly movements in this ratio appear to explain little of the changes in speculative positions, although the relationship is statistically significant for some currencies.⑤

Tracking the carry trade



[1] BIS reporting banks' (in currency reporting countries) cross-border claims and claims on residents, by residency of counterparty; in trillions of US dollars, at constant end-2006 Q3 exchange rates. [2] Excluding claims of banks located in the country. [3] In billions of US dollars; derived using exchange rates at the beginning of 2006. [4] Non-commercial short positions. [5] Non-commercial long positions.

Sources: CFTC; BIS. Graph B

Evidence from the OTC derivatives market is sketchier still than that from exchange-traded contracts. The BIS semiannual OTC derivatives survey indicates that, up to end-June 2006, positions in FX contracts denominated in the main funding and target currencies grew faster than the market as a whole. More up-to-date settlement data from CLS Bank show some increase in the volumes of FX swaps denominated in yen, francs and sterling in late 2006. However, turnover in these contracts has been relatively stable, suggesting that activity may be mainly driven by factors other than carry trades.

① Some observers classify as carry trades all foreign currency lending, including, for example, foreign currency bond purchases by Japanese households or Swiss franc-denominated mortgage borrowing by residents of central European countries. This box deals primarily with speculative trades with offsetting long and short positions, since these are arguably more likely to be unwound quickly should market disruptions occur. ② Defined as the three-month interest rate differential weighted by the implied volatility of exchange rates. ③ Style analysis consists of panel regressions of hedge funds' returns on explanatory variables which track the returns of broad market indices, as well as proxies for carry trade payoffs. The analysis is based on hedge fund return data from HFR and 24-month rolling panel regressions on individual hedge fund families. See P McGuire et al, "Time-varying exposures and leverage in hedge funds", *BIS Quarterly Review*, March 2005, for a more detailed discussion. ④ The figures do not include claims of banks in the United States on US residents. ⑤ Data from other exchanges also provide evidence. Open interest in futures traded in Japan increased sharply, in particular in contracts on the Australian and the New Zealand dollar. Positions in the US dollar contract traded in Brazil also grew in late 2006. Futures on the Turkish lira are traded on the Turkish Derivatives Exchange, but volumes are low by international standards.

$39 billion bid, the largest leveraged buyout on record. More broadly, indicators of M&A financing for the most part pointed to continued high levels of activity, although the number of announced deals has declined somewhat in recent months (Graph 6, left-hand panel).

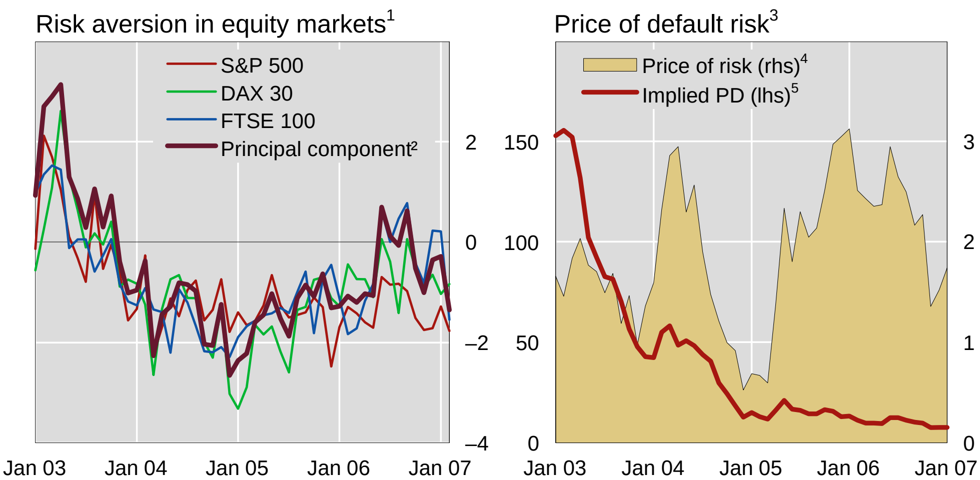
In addition to earnings and M&A activity, greater investor risk appetite also seemed to contribute to the rallies in the major equity markets. Implied volatilities remained low relative to mid-2006, in particular for the S&P 500 (Graph 5, right-hand panel). Implied volatility is influenced by both perceptions of future market volatility and investors' aversion to such volatility. These can be disentangled by comparing the distributions of expected returns implied by option prices with that of historical returns. Measures of risk aversion derived in this way indicate a decrease during the period under review, with the common component of the individual measures for various equity markets reaching its lowest level since July 2005 (Graph 7, left-hand panel).

High-yield credit spreads touch historical lows

Record low credit spreads reflect …

Corporate credit markets continued to rally between end-November and late February. High-yield credit was especially strong, with spreads falling to record lows in some markets (Graph 8). US dollar high-yield asset swap spreads tightened by 59 basis points, and ended the period near the lowest level on record (198 basis points). Similarly, spreads in euro and sterling high-yield credit markets touched record lows during the period. Investment grade credit markets also rallied, with spreads in the US dollar market tightening by 9 basis points, and spreads on euro and sterling debt by 5 and 7 basis points respectively.

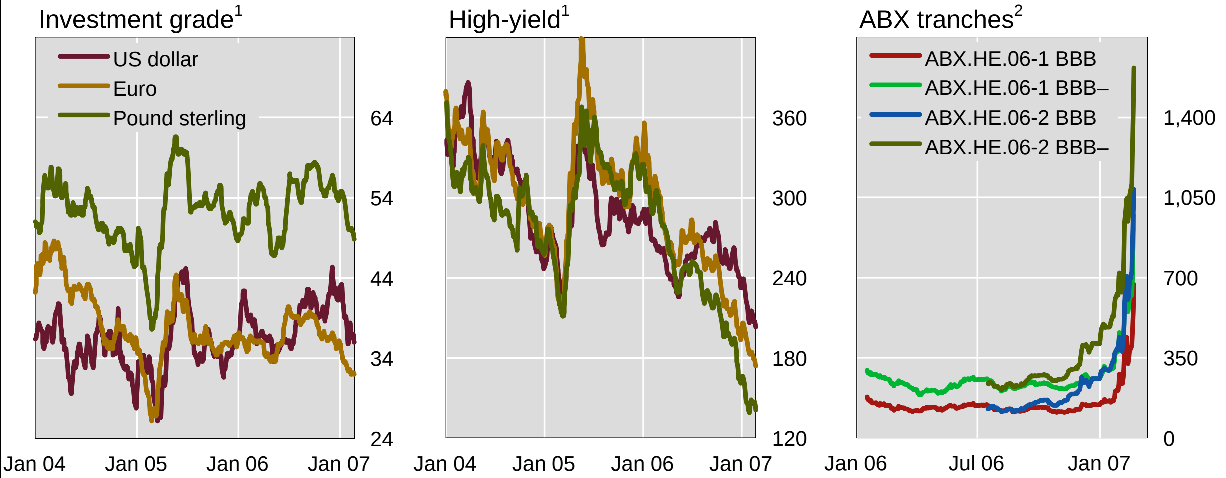
Risk appetite in equity and credit markets



[1] Derived from the differences between two distributions of returns, one implied by option prices, the other based on actual returns estimated from historical data. [2] First principal component of risk appetite indicators estimated for the S&P 500, DAX 30 and FTSE 100. [3] Based on one-year spreads and default probabilities for the constituents of the DJ CDX.NA.IG.4 index. [4] Ratio of risk neutral to empirical probabilities of default. [5] Probability of default (PD) implied by one-year CDS spreads, assuming a recovery rate of 40%, in basis points.

Sources: Bloomberg; Chicago Mercantile Exchange; Eurex; London International Financial Futures and Options Exchange; Markit; BIS calculations. Graph 7

Credit spreads



[1] Merrill Lynch corporate bond indices; asset swap spreads over Libor; five-day moving average, in basis points. [2] Home equity floating closing on-the-run spreads, in basis points.

Sources: Bloomberg; JPMorgan Chase; Merrill Lynch; Moody's Investors Service; BIS calculations. Graph 8

As in equity markets, greater investor risk appetite seemed to contribute to the rally in credit markets. A simple estimate of risk appetite in credit markets can be constructed as the ratio of risk neutral default probabilities derived from credit spreads to those derived from underlying balance sheet information (Graph 7, right-hand panel). This ratio has trended downwards since the summer of 2006, hitting its lowest level in November since March 2005.

… strong demand for structured products …

Investor demand for structured credit products remained strong in the fourth quarter of 2006, also helping to keep credit spreads low. Global issuance of funded CDOs in 2006, at $489 billion, was the highest on record, with particularly robust activity in the fourth quarter. Issuance of synthetic CDOs, which package and securitise credit default swaps on a range of companies, also soared in 2006 to an estimated $450 billion, double the amount issued in 2005. Arrangers of such products often hedge their positions in the cash markets, possibly putting downward pressure on credit spreads.

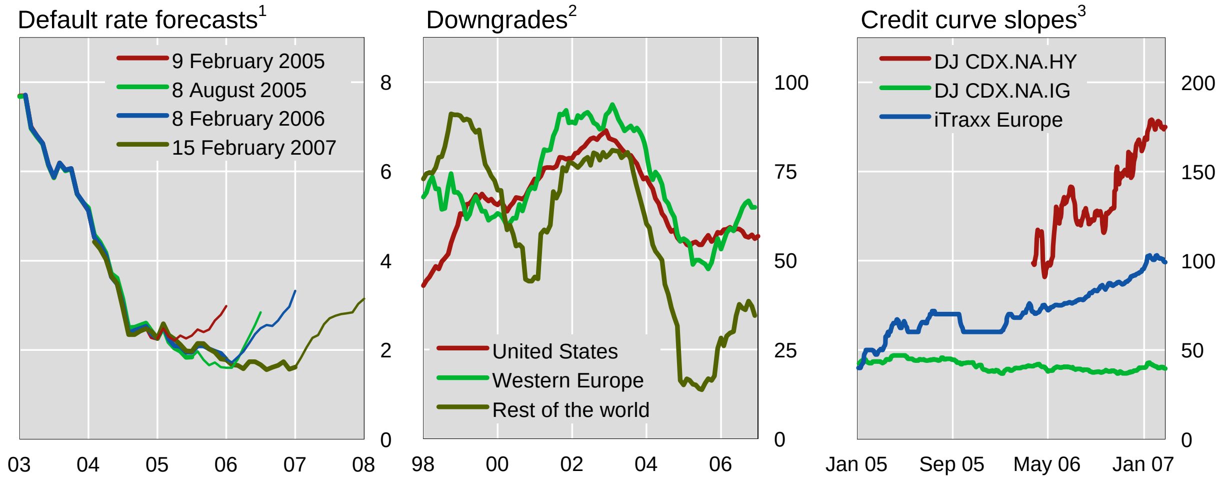
… sound corporate fundamentals…

The rally in credit markets has been underpinned by earnings growth and generally strong corporate balance sheets. On an aggregate basis, corporate profits as a share of GDP have trended upwards in all major markets since mid-2001, and liquid assets relative to debt remain at elevated levels. Moreover, in the United States at least, corporate debt as a share of cash flow fell in 2006 for the fifth year in a row. The apparent health of the corporate sector has led to the surprisingly low realised default rates for speculative grade credit, which have hovered near 2% since 2005, consistently below forecasts (Graph 9, left-hand panel).

… and unusually low corporate default rates

Market participants generally expect default rates to rise in 2007, although there does not appear to be much concern over a sudden and widespread deterioration in credit quality. That said, the difference in spreads between long- and short-maturity high-yield CDS indices has been rising since at least May 2006 in both US and European markets (Graph 9, right-hand panel). Spreads at all maturities have tightened, but those on short- and medium-

Default rates and credit quality



[1] Speculative grade default rate forecasts by Moody's at the time of the legend date, in per cent; the thin lines refer to forecasts one year ahead of the legend date; the thick lines refer to historical default rates. [2] Non-financial corporate downgrades as a percentage of all rating changes by Moody's; 12-month trailing average. [3] Ten-year minus three-year (10- minus five-year for iTraxx Europe) slope for on-the-run CDS spreads; five-day moving average.

Sources: JPMorgan Chase; Moody's Investors Service; BIS calculations. Graph 9

maturity instruments have narrowed the most. While strong investor demand for shorter-term instruments may have played a role, the steepening of the term structure of these contracts may indicate that market participants' sanguine view of default risk over the near term has not entirely spilled over into their longer-term expectations.

Credit investors also seem unconcerned about the ongoing global M&A boom and its possible implications for credit quality. Debt financing of these deals, tracked by syndicated loans earmarked for acquisitions and leveraged buyouts, has risen sharply in recent months, possibly signalling a rise in corporate leverage levels. Equity financing of deals has become less common, accounting for around 12% of announced deals in 2006, compared to 17% in 2005 and 19% in 2004 (Graph 6, centre and right-hand panels). On average, the premiums paid in recent deals have not been particularly high.

Home equity CDO spreads widen significantly

Problems in the subprime sector of the US mortgage market have become more visible, although it is not yet clear how these might spill over into the broader credit markets.[2] Spreads on non-investment grade tranches of home equity CDOs had widened considerably in December, reflecting rising delinquency rates and news of the bankruptcy of several subprime lenders in the United States (Graph 8, right-hand panel). But following HSBC's announcement on 8 February that more funds would have to be set aside to cover bad debts in its subprime lending portfolio, and New Century Financial's downward revision of its 2007 loan production forecast, spreads widened by an additional 200+ basis points in the space of two days. Spreads widened further following an announcement on 20 February of large net losses for the fourth

[2] See A Frankel, "Prime or not so prime? An exploration of US housing finance in the new century", *BIS Quarterly Review*, March 2006, for a discussion of risks in subprime mortgage markets.

quarter of 2006 by Novastar, another large subprime lender in the United States.

Emerging market spreads decline further

In emerging markets, spreads continued to tighten while equity prices rose further during the period under review. Between end-November and late February, the EMBI Global spread index fell from 200 to 170 basis points, hitting all-time lows along the way (Graph 10, right-hand panel). Emerging market CDS spreads also continued on their downward path, although not reaching the lows seen prior to the May–June sell-off. Between the end of November and late February, the MSCI Emerging Market equity index rose by 7%, on top of the 21% increase seen in the first 11 months of 2006.

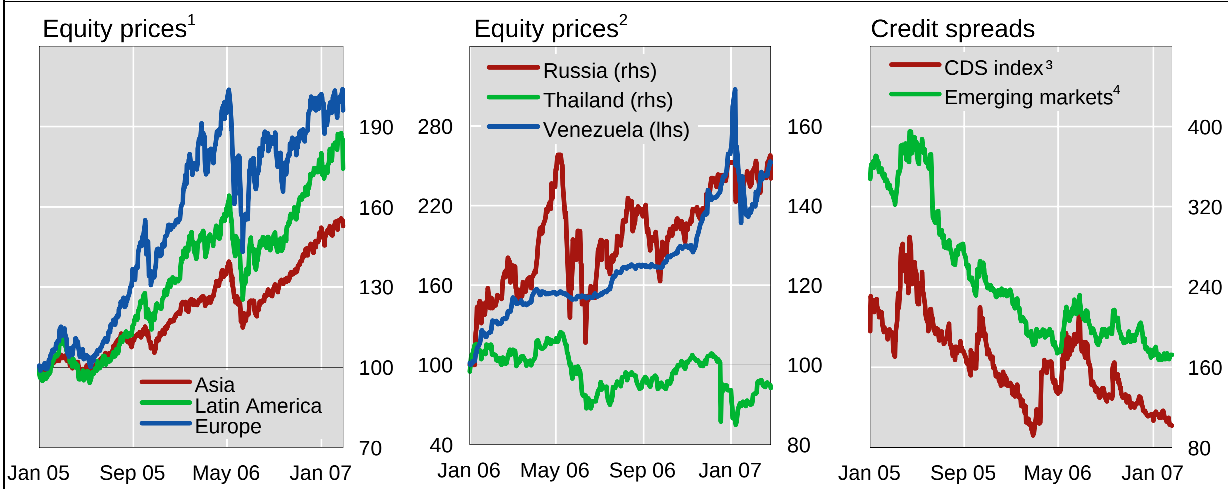
Temporary volatility in emerging asset markets …

In this generally positive market environment, some emerging equity markets suffered a temporary setback and a bout of volatility at the very beginning of 2007, as stock prices suddenly fell sharply. Among the hardest hit was the Russian equity market, which lost almost 7% between end-2006 and 9 January (Graph 10, centre panel). While few immediate triggers for these abrupt price movements were apparent, the prolonged fall in both oil prices and a number of other commodities is likely to have contributed to the decline. Another factor could have been a sense among investors that the Russian market, in particular, was overdue for a correction, following a 51% increase in 2006. Although the episode led some market participants to question whether the risk appetite of investors for emerging market assets was declining, within a couple of weeks most of the losses suffered in January had been recouped.

… triggered by local events …

High-volatility market flare-ups occurred in some countries, triggered by local events, but these remained largely localised. On 18 December, the Bank of Thailand announced the introduction of capital controls aimed at stemming the inflow of speculative capital. The measure stipulated that financial

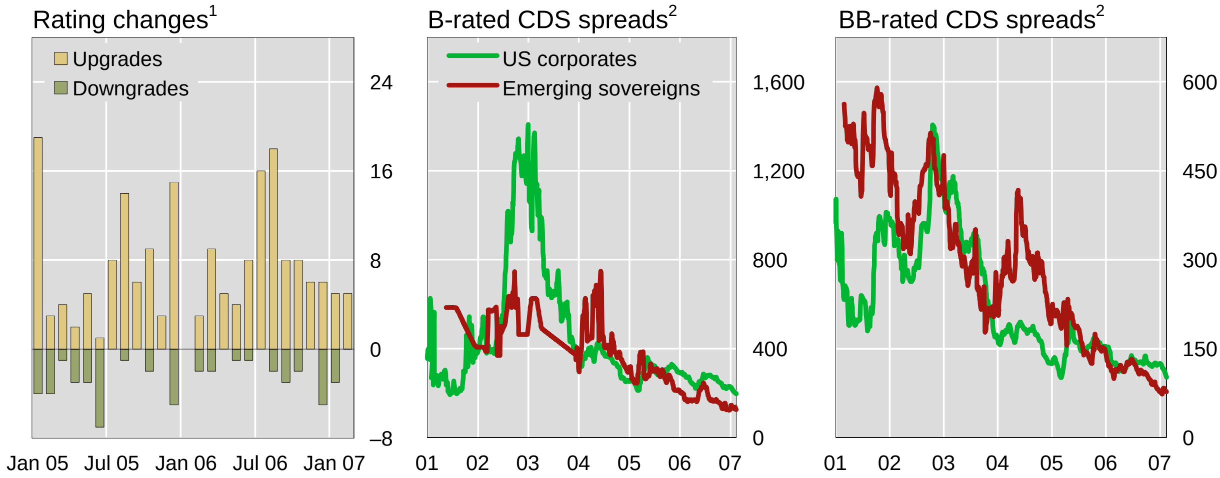
Emerging market equity prices and sovereign spreads



[1] In local currency; 3 January 2005 = 100. [2] In local currency; 3 January 2006 = 100. [3] Emerging markets CDX five-year on-the-run CDS mid spread, in basis points. [4] EMBI Global index; sovereign spread over government bond yields, in basis points.

Sources: Datastream; JPMorgan Chase; BIS. Graph 10

Rating changes and credit spreads



[1] Monthly long-term foreign and local currency sovereign rating changes from Fitch, Moody's and Standard & Poor's. [2] Daily median CDS spreads; five-day moving averages, in basis points.

Sources: Bloomberg; Dealogic; Markit; BIS calculations. Graph 11

institutions would be required to withhold 30% of foreign currencies exchanged for Thai baht in unremunerated accounts. Moreover, investors would be able to recoup their funds only after keeping their investment in Thailand for at least one year, or face losing a third of the deposit. The measures were motivated by concerns about a sharp appreciation of the Thai currency, which prior to the announcement had strengthened by 4% against the US dollar in one month and by 16% since the beginning of 2006. The increase in the value of the baht, which had gathered pace after the military coup in September, was beginning to hurt Thai exporters progressively more, prompting the authorities to take action. It was thought that the capital inflows were going into the domestic bond market, and the measures were designed to discourage such inflows specifically. In the days following the announced capital controls, the baht depreciated by up to 4%, but the largest effect was seen in the Thai stock market, which lost around 15% on the day after the announcement, suggesting that a large part of the flows had been going into the equity market rather than the bond market (Graph 10, centre panel). On that day, net sales of equities by non-residents reached a record $700 million. As a consequence, the government decided to introduce a number of exemptions to the controls, including for investments in equities. By late February, about two thirds of the stock market losses had been recouped, while the baht had resumed its upward path and had strengthened a further 5% compared to the level seen before the introduction of capital controls. Overall, the effect of the Thai turbulence on markets in Asia and elsewhere was limited and temporary.

Two Latin American countries also experienced some turbulence as a result of political factors. Spreads on Ecuador's external debt soared to more than 1,000 basis points as markets priced in the increased likelihood of losses following a presidential election in which the winning candidate had publicly discussed the possibility of an Argentine-style default. In Venezuela, meanwhile, credit spreads rose and equity prices fell by almost 20% following the announcement that a number of private companies would be nationalised. Falling oil prices also weighed on prices of Venezuelan assets at times.

… has little overall impact on spreads

Apart from such bouts of temporary volatility, generally favourable economic conditions and improvements in the outlook for the US economy continued to support emerging market asset prices up until late February. Strong domestic growth prospects and a generally positive fiscal outlook played a role as well. Nonetheless, investors' strong appetite for risk also appeared to be an important factor behind the continued positive developments in emerging asset markets during the period under review. While positive rating changes continued to outnumber negative ones in the past few months – Standard & Poor's, for example, raised India's debt rating to investment grade in January – the ratio of positive to negative changes was lower in the three-month period between December and February than it had been in any three-month period since mid-2005 (Graph 11, left-hand panel). This, however, had little impact on the pace of narrowing of emerging market credit spreads.

The attractiveness of emerging market debt for investors was also evident in the continued decoupling of such spreads when compared to US corporate spreads within the same rating category (Graph 11, centre and right-hand panels). Seen over a relatively long time horizon, emerging market B and BB-rated CDS spreads have generally tended to exceed those on comparable corporate CDS spreads, with the notable exception of a period in 2002–03 following the collapse of Enron and the revelation of a string of corporate governance improprieties in the United States. However, as of around mid-2005, a significant shift in the relative pricing of emerging market and corporate credit seems to have taken place, with spreads on the former now tending to be lower than those on the latter. This tendency has persisted through recent months. For example, since the sell-off in May–June 2006, the median value of five-year CDS spreads on BB-rated emerging market debt has fallen by 35 basis points, while the corresponding CDS spread on US corporates has declined by only 10 basis points. This could suggest a significant reassessment of the riskiness of emerging market credit vis-à-vis corporate credit within the same rating category, and/or a substantial reduction in the price of emerging market credit risk.

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Highlights of international banking and financial market activity[1]

The BIS, in cooperation with central banks and monetary authorities worldwide, compiles and disseminates several datasets on activity in international banking and financial markets. The latest available data on the international banking market refer to the third quarter of 2006. The discussion of the international debt securities market and exchange-traded derivatives markets draws on data for the fourth quarter of 2006.

The international banking market

Locational banking statistics

Growth in cross-border claims is steady in the third quarter

In the third quarter of 2006, total cross-border claims grew by 16% on a year-on-year basis, a growth rate which was up slightly from the previous quarter and in line with historical averages. Total claims of BIS reporting banks expanded by $808 billion, bringing the stock of claims to $25 trillion.

Growth continued to be particularly robust for sterling- and dollar-denominated claims (Graph 1). For the second consecutive quarter, claims denominated in sterling increased at a more rapid rate than those in all other major currencies, at 23% year on year. The 18% expansion of dollar-denominated claims was well above their seven-year average growth rate of 11%, while growth in euro claims was somewhat more subdued at 14%.

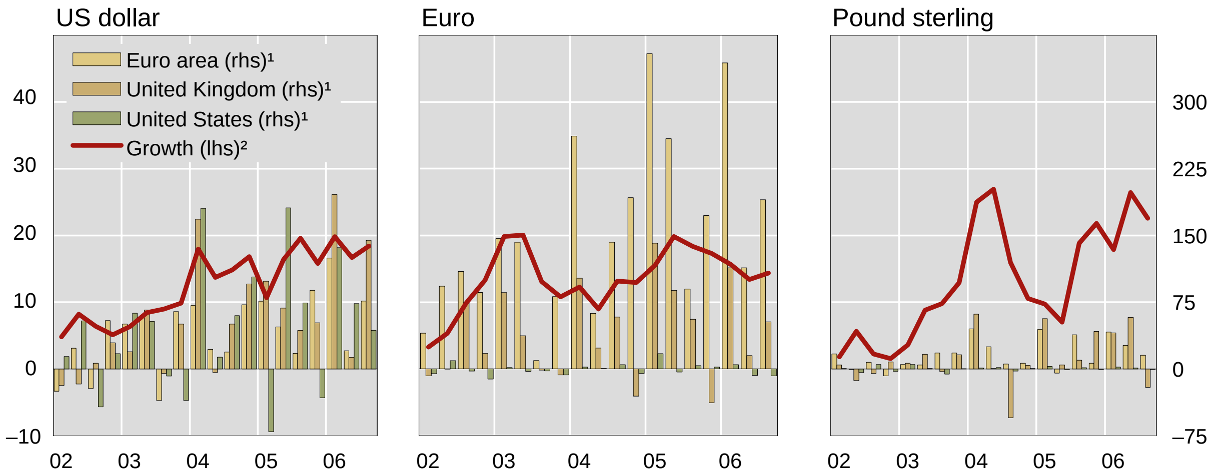
Yen-denominated claims on non-banks increased for the first time in three quarters, pushing growth to 2.5% on a year-on-year basis. Yen claims of banks in Germany and Japan fell while those of banks in France, Switzerland and the United Kingdom increased.

Growth in Swiss franc-denominated lending cools

Growth in Swiss franc-denominated claims on non-banks was more modest than in previous quarters. After gains of $6 billion and $2 billion the previous quarter, banks in the United Kingdom and Austria reported increases in these claims of only $300 million and $500 million, respectively. The cooling

[1] Queries concerning the locational banking statistics and international debt securities should be addressed to Ryan Stever, those regarding the consolidated banking statistics to Goetz von Peter, and those concerning the derivatives statistics to Christian Upper.

Cross-border claims by reporting country and currency



[1] Exchange rate adjusted quarterly changes in amounts outstanding, in billions of US dollars. [2] Annual changes, in per cent, calculated as the sum of exchange rate adjusted changes in amounts outstanding between periods *t*–3 and *t*, divided by the amount outstanding in period *t*–4.

Graph 1

was most evident for Austrian banks, where the expansion in Swiss franc-denominated claims in the latest quarter was the lowest in three years.

The growth in claims of banks in the developed world was steady at 16%. Banks in the United Kingdom were responsible for about a third ($208 billion) of this growth (Graph 2, centre panel), and a large share of those claims ($155 billion) were on non-banks (Graph 2, left-hand panel). Banks in both Japan and the United Kingdom have channelled an ever greater proportion of their total claims to non-bank borrowers (Graph 2, centre panel).

Banks in Canada are particularly active

Banks in Canada were particularly active in the third quarter compared with past experience, as their Canadian dollar-denominated cross-border claims on non-banks surged by $17 billion. Most of this increase was vis-à-vis residents of the United States. This high growth occurred in a quarter of unusually high differentials between US and Canadian short-term interest rates accompanied by below average volatility of the Canadian/US dollar exchange rate.

Funding from banks in Switzerland to offshore centres grew markedly in the third quarter, on the heels of impressive gains in the second (Graph 2, right-hand panel). Although the largest previous quarterly increase in net claims on offshore centres was $6 billion, in the second and third quarters of 2006 these claims rose by $51 billion and $66 billion, respectively. The increase stemmed largely from growth in dollar-denominated bank loans and dollar-denominated inter-office activity.

There was also unusually strong activity between banks in offshore centres and the United States, with dollar-denominated net claims on the United States increasing by a record $147 billion. This was primarily driven by a reduction in dollar deposits from both banks and non-banks.

Emerging markets are strong net depositors

Emerging markets were strong net depositors in the third quarter (Graph 3, left-hand panel). While growth in deposits placed by emerging market residents accelerated sharply, the growth of claims on these residents remained steady. As a result, reporting banks' net claims on emerging market countries decreased by $64 billion, 23% of total net claims.

Residents of Colombia and Mexico contributed heavily to the bulk of new deposits from Latin America. The new deposits from Mexican residents ($10 billion) into reporting banks were twice as large as those in any previous quarter. The majority of those flows were into banks in the Caribbean offshore centres. The new deposits from residents of Colombia, however, were mostly placed with banks in the United Kingdom.

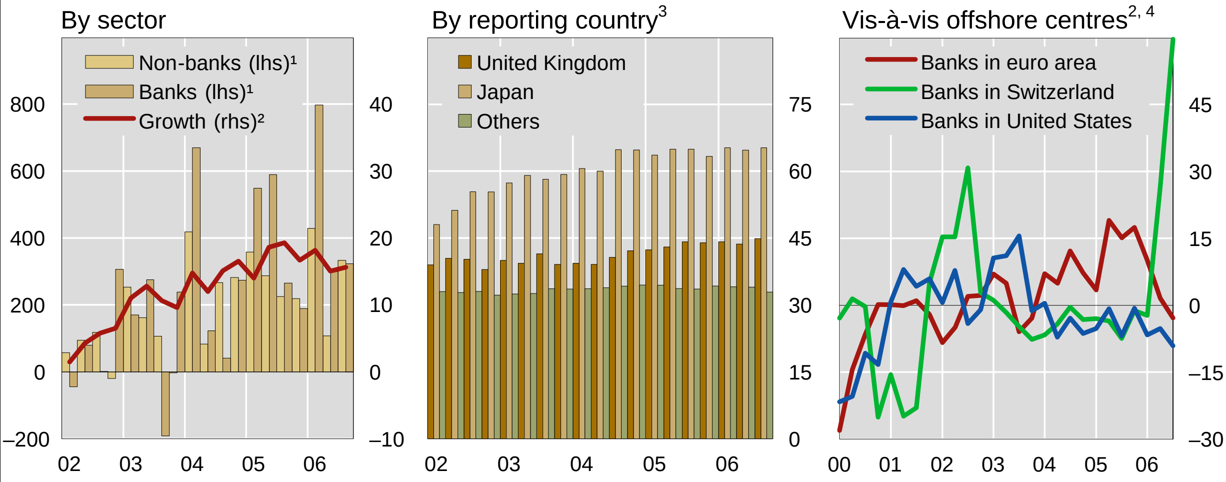
Reporting banks' claims on Korea, Taiwan (China)[2] and Thailand were the principal drivers of the growth of claims on developing Asia (Graph 3, centre panel). New claims on Thailand ($2 billion) were mostly in yen, while the $5 billion in new claims on Taiwan were primarily in dollars. Claims on Korean borrowers expanded by $27 billion, accounting for well over half of the growth of claims on the developing Asia-Pacific region. Over the second and third quarters of 2006, reporting banks' stock of claims on Korea rose by more than $50 billion (50% year on year), or more than all the growth in claims on Korea over the last five years combined.

Foreign currency borrowing in Korea is strong

Much of the increase in loans to Korea was in foreign currency, perhaps due to lower interest rates than locally available accompanied by borrowers' expectations of local currency appreciation. While authorities implemented measures in August 2006 to restrain such inflows, an immediate dampening impact was not apparent in the third quarter, at least based on quarterly data.[3]

The supply of funds from oil-exporting countries grew rapidly in the third quarter, with deposits from Russia in reporting banks expanding by more than $13 billion. While sterling deposits placed by Russian residents fell by

Total claims of BIS reporting banks in the developed world



[1] Exchange rate adjusted quarterly changes in amounts outstanding, in billions of US dollars. [2] Annual changes, in per cent, calculated as the sum of exchange rate adjusted changes in amounts outstanding between periods *t*–3 and *t*, divided by the amount outstanding in period *t*–4. [3] Share of claims on non-banks, in per cent. [4] Net claims.

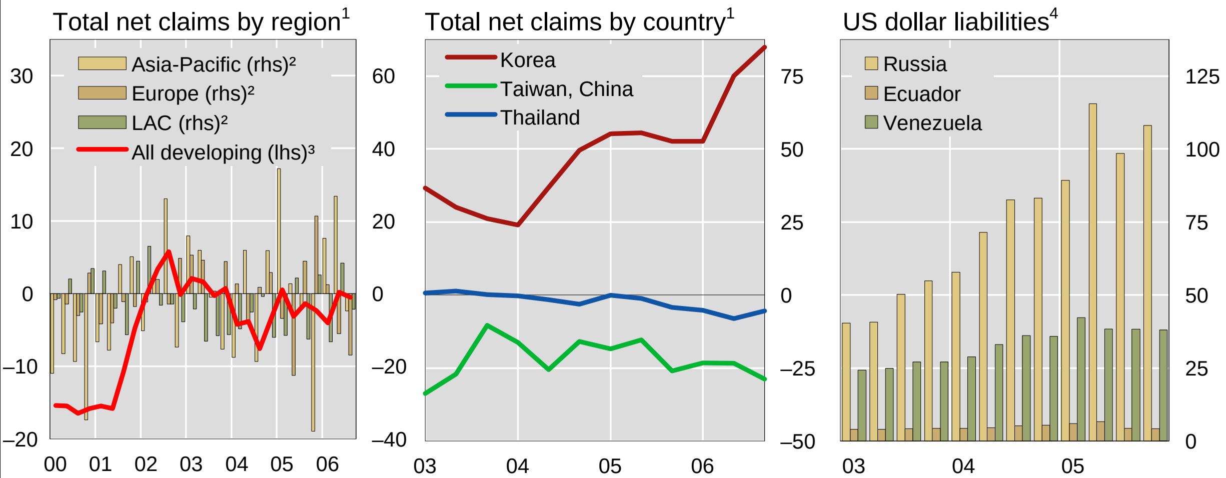
Graph 2

[2] Hereinafter Taiwan.

[3] Bank of Korea, *Financial Stability Report*, October 2006.

BIS reporting banks' positions vis-à-vis emerging economies

By vis-à-vis countries



[1] "Total net claims" is defined as total assets minus total liabilities, in billions of US dollars. [2] Exchange rate adjusted quarterly changes in amounts outstanding, in billions of US dollars; LAC = Latin America and Caribbean. [3] Annual changes, in per cent, calculated as the sum of exchange rate adjusted changes in amounts outstanding between periods *t*–3 and *t*, divided by the amount outstanding in period *t*–4. [4] Total liabilities, in billions of US dollars.

Graph 3

$4 billion, dollar and euro deposits increased by $6 billion and $10 billion, respectively (Graph 3, right-hand panel). Nearly a half of the new deposits were placed in banks in the euro area and another $5 billion in banks located in the United Kingdom.

Deposits from OPEC member states rebounded from a slight dip in the second quarter. This was due principally to growth in US dollar deposits from OPEC countries in reporting banks, which rebounded to $44 billion following a $4 billion decline. Residents of Ecuador and Venezuela continued to reduce their dollar deposits (Graph 3, right-hand panel), while residents of Libya, Saudi Arabia and the United Arab Emirates were among the largest contributors to the growth in OPEC's dollar deposits.[4]

Consolidated banking statistics on an immediate borrower basis

Expansion in consolidated banking claims driven mostly by French, UK and Swiss banks

The consolidated banking statistics show that the expansion during the third quarter of 2006 was driven mostly by French, UK and Swiss banks. Foreign claims[5] were extended primarily to borrowers in the United States, emerging markets and the euro area, even as claims on banks in Germany and the Netherlands declined noticeably. French banks alone accounted for over a quarter of new foreign claims, and for over 40% of new local claims in local currencies, mainly the result of acquisitions of banks in Italy and Greece.

[4] These data should be interpreted with caution since the United States does not provide a complete breakdown of positions vis-à-vis individual oil-exporting countries in the Middle East but only for the Middle East region as a whole (which includes non-OPEC members). Thus, data for many individual countries as well as OPEC do not include data from banks in the United States.

[5] Foreign claims comprise international claims and local claims in local currencies. Local claims are those booked by foreign offices on residents of the country where the foreign office is located.

Australian and Canadian banks also raised their local currency claims, mostly on the main neighbouring countries, while Swiss banks built up local claims on the United States.

Emerging economies remained active in the international banking market. As lenders, Brazilian, Indian and Turkish banks substantially raised their international claims, albeit from a small base. As borrowers, emerging market residents received 20% of new international claims, twice their share in the stock outstanding. Some 80% of new credit went to emerging Asia and emerging Europe, while growth in claims on Latin American residents slowed. The doubling of BIS reporting banks' claims on the public sectors of Korea and Taiwan, mostly in the short-term segment, stood out. In reporting countries' emerging market portfolios, the share of public sector claims gained 1 percentage point, to stand at 17%, and the share of claims maturing within a year edged up to 48%. This is in contrast to the continued decline of these shares in reporting banks' overall portfolios.

Consolidated banking statistics on an ultimate risk basis

Exposures on an ultimate risk basis increase vis-à-vis almost all countries

The consolidated banking statistics on an ultimate risk basis indicate that reporting banks increased their exposures to almost all countries in the third quarter of 2006, in spite of political and military tensions in a number of countries. Claims on borrowers in several smaller emerging markets, such as Egypt, Israel and Ukraine, jumped by a quarter or more. From the lenders' perspective, Greek and Norwegian banks raised their portfolio share allocated to emerging markets by more than 10 percentage points in the course of a year, while US banks shifted their emerging market portfolios significantly towards public sector borrowers.

The value of derivatives positions declines

The relative sizes of contingent liabilities and derivatives positions in the portfolios of reporting banks have moved over time. Since reporting began in early 2005, the value of derivatives positions has decreased more often than increased. For example, the latest quarter's decline of 9% offset the previous increase in the market value of derivatives, largely reflecting the decline in long-term rates that reduced the market value of interest rate swaps.[6] By contrast, contingent facilities have expanded slightly faster than banks' total foreign claims. Within this category, guarantees have increased, while credit commitments have declined, as a share of foreign claims. From the perspective of the ultimate borrowers, the United States continued to secure the largest share of credit commitments by far, while the euro area displaced the United Kingdom as the region attracting the largest share of guarantees.

Risk transfers remain a small part of emerging market exposures

Risk transfers remain a relatively small part of portfolio exposures on emerging markets (Graph 4). Banks can use risk transfers to reduce their portfolio exposure when measured on an ultimate risk basis.[7] A gross outward risk transfer takes place when a US bank's loan to a Mexican corporate, for

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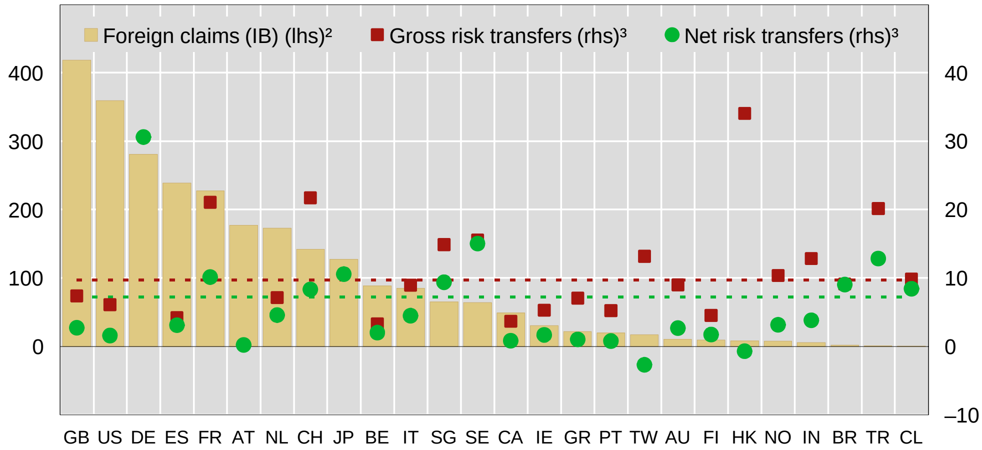
[6] Only positive market values of derivatives positions are reported.

[7] See P McGuire and P Wooldridge, "The BIS consolidated banking statistics: structure, uses and recent enhancements", *BIS Quarterly Review*, September 2005 for a discussion.

Risk transfers out of emerging markets, by reporting country[1]

Positions at end-September 2006



AT = Austria; AU = Australia; BE = Belgium; BR = Brazil; CA = Canada; CH = Switzerland; CL = Chile; DE = Germany; ES = Spain; FI = Finland; FR = France; GB = United Kingdom; GR = Greece; HK = Hong Kong SAR; IE = Ireland; IN = India; IT = Italy; JP = Japan; NL = Netherlands; NO = Norway; PT = Portugal; SE = Sweden; SG = Singapore; TR = Turkey; TW = Taiwan, China; US = United States.

Note: The dotted lines represent the overall shares of the corresponding colours as weighted averages. Austria, Germany and Japan report only net risk transfers and are excluded from average gross risk (outward).

[1] Of the 30 countries reporting consolidated banking statistics (IB), four do not report risk transfers and are excluded. Risk transfers are calculated vis-à-vis all emerging markets combined; net risk transfers exclude transfers within emerging markets. [2] Cross-border claims plus local claims in all currencies booked by foreign offices of domestically owned banks of reporting countries; amounts outstanding; in billions of US dollars. [3] As a percentage of foreign claims (IB).

Source: BIS international consolidated banking statistics. Graph 4

instance, is guaranteed by a third party outside Mexico. But if the guarantor resides in another emerging market, this also gives rise to an inward risk transfer into that emerging market. In the statistics, such inward transfers offset more than half of outward transfers, resulting in *net* risk transfers out of all emerging markets combined of only 7.2% of foreign claims.[8] Net risk transfers were, in fact, smaller than the additional contingent exposure taken on through guarantees extended to emerging market borrowers (10.5% of foreign claims). This was the case for the reporting population as a whole, as well as for most individual reporting countries listed in Graph 4.

The use of risk transfers varies by bank nationality

Banks of different nationalities vary in the way they use risk transfers to alter their portfolio exposures. German banks transferred more than 30% of their exposures out of emerging markets, while US banks, with even larger exposures to emerging markets, did so by less than 2%. In the case of US and Spanish banks, their strong local presence in Latin America may substitute for risk transfers.[9] Since their risk transfers measured in gross and net terms were equal, Brazilian banks appeared to concentrate on guarantors outside

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[8] This percentage has remained fairly stable and has never exceeded 10% since reporting began.

[9] See P McGuire and N Tarashev, "The international banking market", *BIS Quarterly Review*, March 2006 for a host country perspective.

emerging markets By contrast, Hong Kong and Taiwanese banks' larger outward risk transfers were completely offset by inward risk transfers from emerging markets.

The international debt securities market

Issuance robust in the fourth quarter

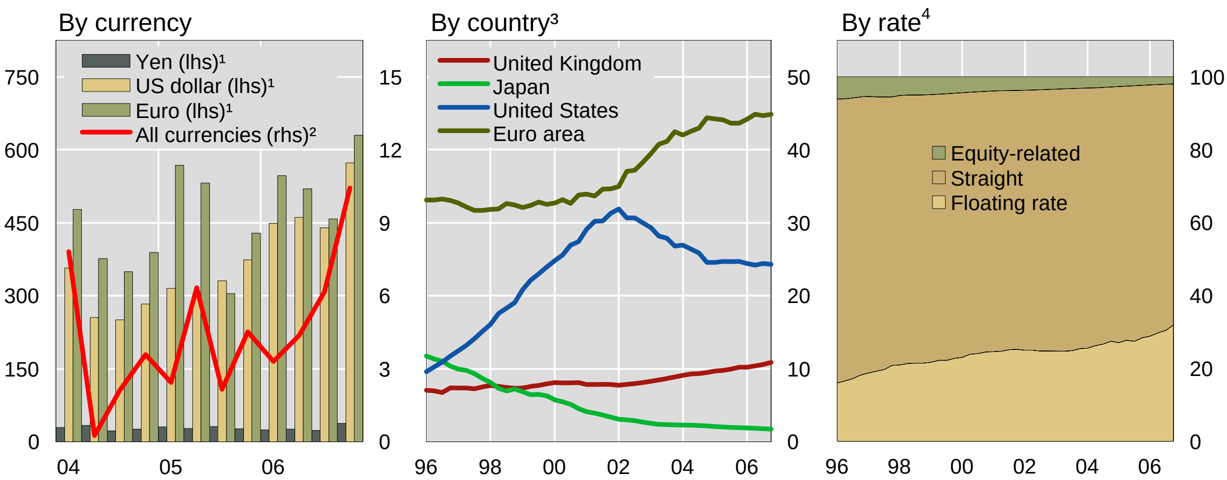
International debt securities issuance was robust in the fourth quarter of 2006. Gross issuance of international bonds and notes increased by over 10% year on year, well above the average growth of 4% since the first quarter of 2000 and the highest rate of growth since the third quarter of 1999. Net issuance surged by 14% year on year to $883 billion, despite a seasonal tendency for it to be sluggish in the fourth quarter. As a result, the amount outstanding of all international bonds and notes increased by 8% from the previous quarter.

The expansion of bond and note issuance in 2006 was strongest in dollar-denominated securities, although the euro and yen also saw significant growth in debt outstanding (Graph 5, left-hand panel). Gross dollar issuance expanded by 12% year on year, the largest increase in five years. Gross euro and yen issuance expanded by $630 billion and $37 billion, respectively. The level of euro-denominated gross issuance remains larger than that of any other currency in both gross and net terms.

In the developed world, issuance was firm in most countries but growth was particularly concentrated in Europe. The euro area was responsible for a full 41% of the $2.1 trillion in total gross issuance and 38% of net issuance. By contrast, the United States' share, which had reached a peak in the first quarter of 2002 at 31%, fell further to 24% (Graph 5, centre panel).

Growth in issuance from the United Kingdom continued to accelerate, increasing by 15% year on year, compared to 9% the previous quarter, and this nudged the United Kingdom's growing share of debt outstanding up to 11%

International bond and note issuance



[1] Announced issues, in billions of US dollars. [2] Annual changes, in per cent, calculated as the sum of exchange rate adjusted announcements between periods *t*–3 and *t*, divided by the sum of exchange rate adjusted announcements between periods *t*–4 and *t*–1. [3] Share in total of all countries, by issuer nationality, in per cent. [4] Share in amounts outstanding, in per cent.

Sources: Dealogic; Euroclear; ICMA; Thomson Financial Securities Data; BIS. Graph 5

(Graph 5, centre panel).

Canada and Japan both experienced high levels of activity in the international debt securities market in the fourth quarter. The $25 billion in gross issuance from Canada marks a 17% year-on-year increase following the previous quarter's growth of 6%. Growth in Japan's gross issuance also surged, but the real sign of growth in Japan was net issuance of $9 billion in bonds and notes. From the beginning of 2000, Japan had made average quarterly net repayments of $111 million. Since the beginning of 2006, however, there has been a run of four consecutive quarters of positive net issuance.

Floating rate debt accounted for a good deal of issuance in international debt securities markets (Graph 5, right-hand panel). Over the last five years, around one quarter of gross issuance of bonds and notes has been floating rate. In the last quarter, that share rose to a record high of 32%. The move away from fixed rate debt is even more pronounced in net terms – with the corresponding share rising a full 5 percentage points from the previous quarter and 20 percentage points more than the average over the last four years.

The share of floating rate debt rises to a record high of 32%

Mortgage-backed securities once again accounted for a large proportion of issuance. For the second consecutive quarter, the largest single issue was a $7 billion offering by a special purpose securitisation trust (Canada Housing Trust) advised by Canada Mortgage and Housing Corporation. In addition, the United States' Federal Home Loan Bank, Fannie Mae and, to a lesser extent, Freddie Mac all had several large issues.

In emerging markets, securities borrowings rose sharply in Europe, Latin America and Asia-Pacific. In emerging Europe, net issuance was nearly 50% greater than the previous quarterly high since 2001. Latin America's growth in net issuance was the largest since 2001. Asia-Pacific experienced strong growth too, over 10% year on year.

Among the countries in the developing world with particularly high net issuance were Russia, the United Arab Emirates, Brazil, Korea, Mexico and Indonesia. Russian net issuance of $16 billion was up by 78% year on year, well above the previous quarter's 24% and about four times the quarterly average over the last five years. Gross issuance in Brazil and Mexico was $7.3 billion and $3.6 billion, respectively, while Indonesia had positive net issuance for the first time in three quarters at $2.6 billion.

Derivatives markets

Trading on the international derivatives exchanges slowed in the fourth quarter of 2006. Combined turnover of interest rate, currency and stock index derivatives fell by 7% to $431 trillion between October and December 2006.[10] As in the previous quarter, the deceleration in activity was primarily the consequence of seasonal factors, which tend to dampen trading in the interest rate segment towards the end of the year. By contrast, activity in derivatives on

10 All growth rates in the section on exchange-traded derivatives refer to quarter-on-quarter increases.

stock indices increased by 5%, although at $45 trillion it remained just below the peak of the second quarter. Turnover in exchange-traded currency contracts rose by 19% to just under $5 trillion, the highest level on record. Finally, activity in futures and options on commodities, which are not included in the above total since notional amounts are not available, increased by 12% in terms of the number of contracts traded, mainly due to a sharp rise in the trading of agricultural commodities in China.

Seasonal factors weigh on turnover of money market derivatives

In the absence of significant monetary policy surprises,[11] turnover in derivatives on short-term interest rates fell by 10% in the fourth quarter of 2006, close to the estimates of seasonal factors presented in the March 2006 *BIS Quarterly Review* (pp 45–6). A particularly strong decline in activity took place in the market for federal funds futures, where turnover more than halved.[12] Trading in options on federal funds initially remained strong, but dropped considerably in November. However, this was not reflected in a reduction in the open positions in the market. Quite on the contrary, open interest in federal funds options almost doubled during the course of the quarter. This suggests that the decline in activity was primarily the result of less short-term trading rather than lower positions. Turnover in derivatives on three-month interest rates declined by a much smaller amount than trading in contracts on overnight rates. For example, trading volumes in three-month eurodollar contracts dropped by 6%, while trading in the equivalent euro and sterling contracts fell by 8% and 17%, respectively. Activity in euroyen futures and options was stable.

Active trading in stock index contracts

Increasing equity prices lifted activity in futures and options on stock indices in most regions. Growth was particularly rapid in some of the smaller markets, such as those for contracts on Hungarian (210%), Indian (78%), Swedish (41%) and Brazilian (37%) equities. Turnover in contracts on euro area and US indices expanded by 17% and 9%, respectively. The only major market posting a decline in activity was Korea, where turnover fell by 10%. The Korean market for stock index derivatives briefly overtook the US market as the world's busiest in the third quarter of 2005, but its size has stagnated since.

Soaring turnover in contracts on central European currencies

Trading in exchange rate linked contracts on the international derivatives exchanges picked up markedly in the final quarter of 2006 and surpassed the peak reached in the second quarter. Among the major currencies, particularly strong growth took place in futures and options on the pound sterling, where turnover increased by 36% to $618 billion. Even higher rates of growth were recorded in contracts on central European currencies such as the Hungarian forint (151%), the Czech koruna (71%) and the Polish zloty (57%). Combined open interest in these currencies almost tripled between end-September and

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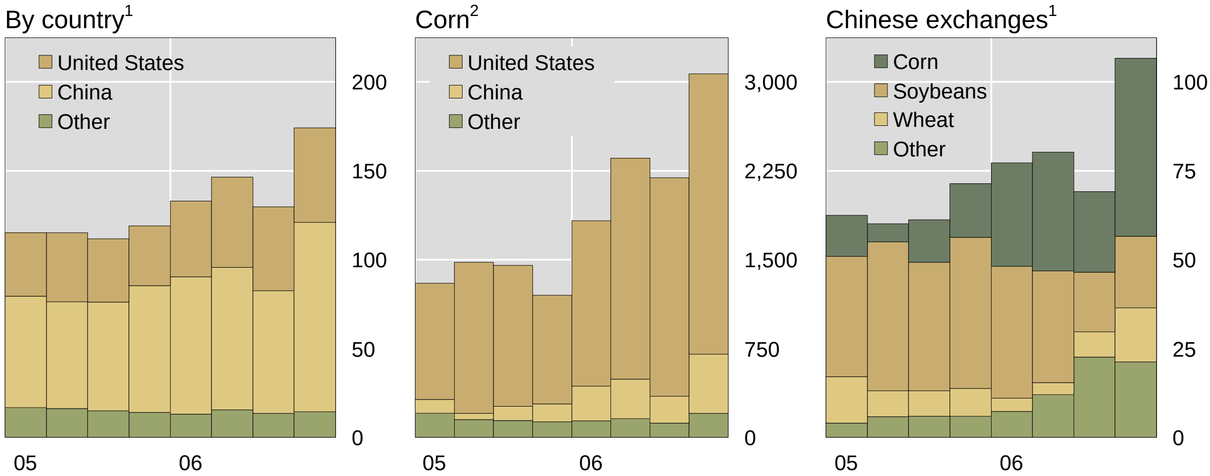
[11] In early January, both the Bank of England and the Bank of Japan surprised markets, the former by increasing policy rates and the latter by leaving them unchanged (Overview, p 3). However, neither event could have affected turnover in the final quarter of 2006.

[12] The payoffs of federal funds derivatives are directly related to policy rates, which makes these contracts particularly suitable for taking positions on the future course of monetary policy. See C Upper, "Derivatives activity and monetary policy", *BIS Quarterly Review*, September 2006, pp 65–76.

Turnover of derivatives on agricultural commodities



[1] In millions of contracts. [2] In millions of tons.

Sources: FOW TRADEdata; BIS calculations. Graph 6

end-December, after declining in the wake of the sell-off in May and June. In the case of Hungary, the rise in positions could be explained by the attractiveness of the forint for carry trades (see the box on p 8), although this explanation is much less likely to hold for positions in the zloty or koruna, where interest rates are much more in line with those in the euro area.

High turnover in agricultural commodities ...

Heavy activity in derivatives on agricultural commodities (34%) offset weaker trading in contracts on energy (–9%) and precious metals (–14%) in the fourth quarter of 2006. Turnover in derivatives on base metals remained stable.

Turnover in agricultural commodities was primarily driven by a surge in the number of contracts on corn (84%), whose prices had increased from approximately \$200 per bushel[13] to almost \$400 in the period under review. Although the rise in corn prices has been widely attributed to the expansion of ethanol production related to fuel conversion programmes in the United States, this factor is unlikely to explain the surge in turnover. Trading in derivatives on corn grew much more rapidly in China (121%) than in the United States (30%), bringing China's share in global turnover measured in terms of the number of contracts traded in agricultural derivatives to 60% (Graph 6). This compares to a share of just over 30% for the United States. However, this measure ignores the much smaller size of Chinese contracts. For example, the futures contract on corn traded on the Dalian Commodity Exchange refers to 10 tons, compared to 5,000 bushels (or approximately 125 tons) for regular-sized and 1,000 bushels (about 25 tons) for mini-sized contracts at the Chicago Board of Trade. Adjusted for the size of contracts, turnover in China accounts for only 16% of worldwide corn turnover, compared to the United States' 77%. In the case of wheat, the share of China is even smaller.

... reflects surge in activity on Chinese exchanges

[13] A bushel of corn weighs approximately 25.5 kg.

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Interpreting sovereign spreads[1]

Sovereign spreads can be broken up into two components: the expected loss from default and the risk premium, with the latter reflecting how investors price the risk of unexpected losses. We show that the risk premium is often the larger part of the spread.

JEL classification: G15, F34.

Recent years have seen a substantial and steady narrowing of sovereign spreads in emerging debt markets. These spreads are the differentials between yields on emerging market debt and those on what might be considered risk-free government bonds of the corresponding duration. The average spread on the EMBI+ index, a widely watched index of emerging market debt prices, for example, fell from about 1,020 basis points in October 2002 to 170 basis points in December 2006.

Does this mean that the borrowers in these markets have become less risky? Much of the recent literature on sovereign spreads has not been very helpful in answering this question. In principle, sovereign spreads reflect both expected losses from default and risk premia. The latter would depend on both the risk of unexpected losses and on how investors price this risk. The literature, however, has not paid enough attention to this distinction, often implicitly assuming that in some way such spreads primarily measure the risk of default.

In this article, in line with the asset pricing literature, we propose an analytical framework for interpreting sovereign spreads. We estimate expected losses from default and risk premia by using data on credit default swap (CDS) spreads and default histories of rated bonds, considering both sovereign and corporate bonds. We find that the expected loss component of the spread is small, while the risk premium plays a bigger role even in periods of relatively low credit spreads.

1 We would like to thank participants at seminars at the BIS, Hong Kong Institute for Monetary Research, Asian Development Bank Institute, Bank of Japan, Bangko Sentral ng Pilipinas, Bank of Thailand, Chinese University of Hong Kong and Hong Kong University for Science and Technology, Claudio Borio and Frank Packer for helpful comments. This paper was written while Eliza Wu was visiting the BIS. All errors remain our own and the views expressed here are solely ours and do not reflect those of the BIS.

The remainder of this article is structured as follows. The first section reviews the literature on default risk and risk premia for both sovereign bonds and corporate bonds. The second section proposes default probabilities as a measure of sovereign risk and illustrates the concept by providing estimates based on historical data on defaults of rated bonds. The third section shows how to decompose sovereign spreads into expected losses from default and risk premia. The final section summarises the results and suggests topics for further research.

Default risk, risk premia and sovereign spreads

One component is the expected loss

A sovereign spread, like any other credit spread, is supposed to compensate investors for default risk.[2] An obvious component of this compensation is the expected loss from sovereign default. For investors who hold the sovereign bond to maturity, this loss is simply the product of the probability of default and the loss-given-default. The probability of default is itself a simple measure of default risk. For investors who plan to sell before maturity, the expected loss would also include the probability of a deterioration in credit quality, short of default.

A less obvious component of the spread is the risk premium. Such a premium compensates investors for the fact that the realised loss from default may exceed the expected loss. Such a default risk is asymmetric because the possible losses from default are large relative to the possible gains from an absence of default. Jarrow et al (2005) have laid down the conditions for the absence of a default risk premium in a world of risk-averse investors. First, defaults on different bonds must be independent. Second, investors must be able to diversify away any idiosyncratic risks by holding a sufficiently large portfolio of bonds. Whether these conditions hold is an empirical question. Can we tell from the data whether there is a sovereign risk premium and, if so, how significant it is?

The risk premium on corporates tends to be rather large

In the case of corporate bonds, the empirical evidence points to a rather large risk premium. Indeed, this risk premium is estimated to be such a large part of credit spreads that Driessen (2005) has dubbed the phenomenon the "credit spread puzzle". Driessen estimates an average premium of 189 basis points after accounting for tax and liquidity effects. Berndt et al (2005) estimate an average premium of a similar magnitude, and moreover find that the risk premium varies greatly over time. For BBB/Baa-rated corporate bonds, Amato and Remolona (2003) suggest that default correlations account for about three quarters of the risk premium and undiversifiable idiosyncratic risk for one quarter. While it is not clear whether sovereign defaults are more highly correlated than corporate defaults, it could be argued that idiosyncratic risk is

2 For less liquid instruments, the spread may also contain a liquidity premium. In the case of US corporate bonds, US local government taxes (which apply to income on corporate bonds but not on US Treasury securities) may also explain part of the spread.

harder to diversify for sovereign bonds because there are fewer available issues.[3]

Nonetheless, the presumption that credit spreads measure just default risk and not risk premia is common among recent papers that propose structural models to measure probabilities of sovereign defaults. Gapen et al (2005) and Oshiro and Saruwatari (2005), for example, apply the standard structural Merton model for corporate credit risk by defining for countries concepts of balance sheet leverage and option volatility. They then judge their approaches to be good ones because they find their risk indicators to be highly correlated with market spreads over time. Diaz Weigel and Gemmill (2006) fit a similar structural model to par Brady bond prices to derive a "distance-to-default" measure of sovereign risk. They then express surprise that country-specific variables account for only 8% of the explained variance of the distance-to-default measure. However, a possible reason for their result is that their distance-to-default measure largely reflects risk premia that are driven by investors' time-varying risk aversion.

Models of sovereign default are often estimated with market spreads

Measuring sovereign risk

In this section, we provide estimates of probabilities of sovereign default as a measure of risk for sovereigns. For present purposes, we rely largely on information from credit ratings, deriving default probabilities from the historical performance of rated bonds. We then examine the power of this measure of risk for explaining the cross-sectional variation of sovereign spreads.

The use of credit ratings

To develop a measure of sovereign risk, we rely on information from credit ratings. In the country risk literature, however, this contrasts with another preferred source of information about sovereign risk, the *Institutional Investor* country ratings.[4] Nonetheless, there are good reasons to rely on credit ratings instead. As explained by Borio and Packer (2004), such ratings have the following advantages: (a) rating agencies explain their criteria and rating methodologies while respondents to the *Institutional Investor* survey do not; (b) rating agencies regularly review and report the correspondence of their ratings with historical default rates; and (c) rating agencies stake their business on the accuracy of their ratings, while respondents to the *Institutional Investor* survey are anonymous and do not have to account for their ratings. Moreover, Micu et al (2006) find that corporate credit default swap spreads react significantly to announcements by credit rating agencies. Since we wish to estimate sovereign risk as judged by market participants, it is important to use information on which they evidently rely.

Credit ratings do contain information

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3 As of April 2006, for example, Moody's rated the bonds of no more than 92 sovereign issuers. Given their skewed return distributions, these bonds are not nearly enough for a diversified portfolio (see Amato and Remolona (2003)).

4 These ratings are featured in Baek et al (2005), Reinhart et al (2003) and Ul-Haque et al (1996).

An important disadvantage of ratings in this regard is that, as Altman and Rijken (2004) among others point out, rating agencies focus on a long-term horizon, using a "through-the-cycle" rating methodology. As a result, ratings respond only to the component of credit quality changes that the agencies perceive to be permanent. Sovereign spreads, however, may reflect risk assessments by investors who do care about credit quality in the short term. Hence, ratings are not likely to provide precise point-in-time measures of risk. To abstract from possible short-term variations in market risk assessments that may be reflected in spreads, we will derive only cross-sectional risk premia and we will do so by comparing assessments implied by ratings only with averages of such credit spreads over time.

Ratings do not provide point-in-time estimates

We use ratings performance information from the three leading international credit rating agencies, namely Moody's, Standard & Poor's and Fitch. We do so for several reasons. First, in spite of differences in agency methodologies, market participants have established a clear correspondence between the ratings scales of the three agencies. For instance, a Aa rating from Moody's implies the same risk as a AA rating from Standard & Poor's. Second, Micu et al (2006) find that two ratings are better than one: credit spreads react to a rating change by one agency even when it is preceded by a similar rating change by another agency. Moreover, it is fairly common at any given time for rating agencies to disagree on a given credit, resulting in "split ratings". In these situations, Cantor et al (1997) find that bond spreads tend to be priced at the average of the ratings. In this article, we estimate a default probability for each sovereign rating based on the average of the frequencies of default for that rating as observed by the three agencies.

Two ratings better than one

We focus only on foreign currency ratings of sovereign debt and ignore ratings on local currency debt. This allows us to isolate sovereign default risk from confounding factors like inflation expectations and foreign exchange and liquidity risks that non-resident investors are likely to face in the case of local currency denominated debt (for a discussion on domestic versus foreign currency sovereign ratings, see Packer (2003)).

Calculating sovereign default probabilities

Our sample consists of 26 emerging market countries. There are 10 Latin American, seven European, six Asian and three Middle East and African (MEA) countries. Table 1 reports the number of countries in each rating grade for sovereign ratings by Moody's, Standard & Poor's and Fitch together with the number of cases for which "split ratings" occur. Most of these emerging market sovereigns tend to be rated single-A at best, and in nearly 70% of the cases the ratings are split.

Split ratings are common

To calculate sovereign default probabilities, we map sovereign ratings onto cumulative default rates for each given rating. Moody's, Standard & Poor's and Fitch publish average cumulative default rates by rating for various investment horizons and they do so separately for corporate debt and sovereign debt. We take the five-year cumulative default rate for each rating and annualise it by assuming a constant default probability during the five-year horizon. This horizon is chosen consistently with the predominant five-year

We use both sovereign and corporate defaults

Sample description				
Moody's	S&P	Fitch	Number of countries	Number of split ratings
Aaa–AA	AAA–AA	AAA–AA	0	0
A	A	A	7	6
Baa	BBB	BBB	5	2
Ba	BB	BB	11	6
B	B	B	2	3
Caa down	CCC down	CCC down	1	1

Sources: FitchRatings; Markit; Moody's Investors Service; Standard & Poor's; authors' calculations. Table 1

tenor represented in the CDS market. We do the calculation for each rating using the default experience of both sovereign debt and corporate debt. These probabilities we then call "ratings-implied probabilities of default" (RIPDs), and they are presented in Table 2.

The reason we also consider the corporate bond default experience in estimating sovereign default probabilities is the small number of actual sovereign defaults. For example, while Moody's rates the bonds of 92 sovereigns, only 11 have defaulted since 1983 and none rated single-A or higher has done so. It is a natural question, then, whether market participants would rely on such a limited sample to form their estimates of default probabilities for sovereign borrowers and not rely also on the experience of corporate defaults.

One reason to ignore corporate defaults is that these might be very different from sovereign defaults. As Eaton et al (1986), Bulow and Rogoff (1989) and Duffie et al (2003) point out, a sovereign default is largely a political decision, albeit influenced by macroeconomic factors. Rather than defaulting outright, a sovereign issuer usually pursues a restructuring or renegotiation of its debt. In doing so, sovereigns effectively trade off the reduced cost of making debt repayments against the increased costs of reputation effects, asset seizure, increased regulatory monitoring, reduced access to external finance

Sovereign ratings and implied default probabilities

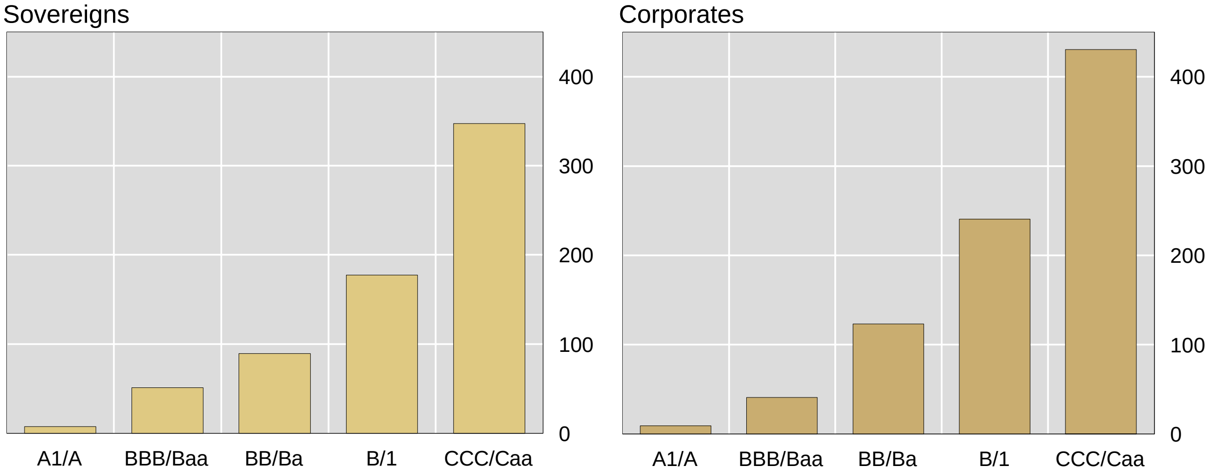
January 2002–June 2006 (in basis points)

Rating category			Moody's		S&P		Fitch		Full RIEL
Moody's	S&P	Fitch	Sovereign	Corporate	Sovereign	Corporate	Sovereign	Corporate	
Investment grade									
Aaa–AA	AAA–AA	AAA–AA	0.0	0.0	0.0	0.0	0.0	0.0	0.0
A	A	A	5.4	6.1	4.5	9.4	13.2	12.0	8.4
Baa	BBB	BBB	48.8	40.6	45.6	41.7	58.8	39.6	45.9
Speculative grade									
Ba	BB	BB	64.0	139.1	111.9	139.3	92.8	90.9	106.3
B	B	B	123.5	280.5	266.4	315.7	142.0	125.3	208.9
Caa down	CCC down	CCC down	273.6	592.9	575.5	469.7	192.7	228.7	388.8

Sources: FitchRatings; Markit; Moody's Investors Service; Standard & Poor's; authors' calculations. Table 2

Average ratings and RIPDs[1]

In basis points



[1] Ratings-implied probabilities of default, average across 26 countries for the period January 2002 to June 2006. The x-axis represents the average ratings across Moody's, S&P and Fitch ratings.

Sources: FitchRatings; Moody's Investors Service; Standard & Poor's; authors' calculations. Graph 1

and international trade disruptions. Nonetheless, rating agencies appear to take all these factors into account and attempt to rate sovereigns and corporates in a consistent manner, so that a given rating represents the same assessment of risk regardless of the nature of the issuer.

Default probabilities are non-linear in ratings

Our calculations show a non-linear relationship between ratings and default probabilities. In Graph 1, we assign a linear scale to ratings, with a AAA/Aaa rating receiving a value of one, a AA/Aa rating a value of two, and so on. The left-hand panel of the graph then shows the relationship of these ratings to RIPDs based on the sovereign default experience and the right-hand panel to RIPDs based on the corporate default experience. As one would expect, in both cases RIPDs rise as ratings decline. In both cases too, the relationship is non-linear, illustrating an important difference of functional form between the two indicators of risk. Amato and Furfine (2003) also find such a non-linear relationship.

In general, default rates have been higher for a given rating for corporates than for sovereigns and this is reflected in the data shown in the two panels. These estimates show an average RIPD for the full sample of countries of 84 basis points a year based on the sovereign default experience. The same average RIPD based on the corporate default experience is 107, about 28% greater than that based on the sovereign default experience.

Are ratings-implied probabilities of default reflected in spreads?

Does our risk measure explain spreads?

To see whether our estimates of RIPDs are indeed relevant measures of sovereign risk from the point of view of market participants, we estimate the extent to which our measure can explain sovereign spreads for our cross section of countries. We also ask whether such estimates can do as well as untransformed sovereign ratings and as well as *Institutional Investor* ratings in explaining sovereign spreads.

For data on sovereign spreads, we use five-year sovereign CDS spreads from the comprehensive Markit database. This database contains monthly quotes on CDS market spreads for 70 developed and emerging market sovereign obligors worldwide. As the sovereign CDS market enables the exchange of sovereign risk between participating financial institutions, Markit compiles quotes from a large sample of financial institutions and aggregates them into a composite spread that is reasonably continuous. We use only spreads of five-year contracts because these contracts are the most liquid and account for a large proportion of the sovereign CDS market.

We compare the explanatory power of three alternative dependent variables: our full RIPD estimates, a simple linear mapping of sovereign foreign currency credit ratings, and the *Institutional Investor* ratings. To ameliorate a possible "peso problem" inherent in our limited sample of sovereign defaults, we propose that our simple RIPD indicator of sovereign default risk be based on the average of sovereign and corporate default rates. This use of corporate default information will not qualitatively change our results. For control variables, we use debt outstanding as a rough measure of liquidity and the VIX index as a measure of global risk (for more discussion on this index, see the special feature by Cairns, Ho and McCauley in this issue). Except for VIX and our risk variables, all the other variables are expressed in natural logarithms. We estimate fixed-effects panel regressions for our sample of countries from March 2002 to end-2005. These estimates use White's correction method so that they are robust to heteroskedasticity.

We avoid a possible "peso problem"

RIPD helps explain spreads

Our results (Table 3) suggest that, as a measure of default risk, RIPD is a significant determinant of sovereign spreads (for a discussion on the determinants of RIPD, see the box). Both the RIPD measure and the agency ratings are statistically significant and economically meaningful for explaining spreads. The *Institutional Investor* country rating appears not to have

Explaining CDS spreads

Explanatory variables	Sovereign risk measures		
	Log (RIPD) (1)	Average agency ratings (2)	*Institutional Investor* ratings (3)
Sovereign risk proxy (1, 2 or 3)	0.262***	–0.274***	0.003
	(0.000)	(0.000)	(0.205)
Bonds outstanding	–0.05	–0.138***	0.111
	(0.231)	(0.011)	(0.163)
VIX	0.062***	0.055***	0.08***
	(0.000)	(0.000)	(0.000)
Time series frequency	quarterly	quarterly	annual
Adjusted R^2	0.95	0.96	0.97

Note: The estimated panel regressions are of the form $log(S_{i,t}) = a_0 + a_1 Sov\_risk_{i,t} + a_2 \, log(Bond_{i,t}) + a_3 VIX_t + \mu_{i,t}$, where $\log(S_{i,t})$ is the natural logarithm of the CDS spread for country *i* at time *t* and *Sov_risk* is the natural logarithm of RIPD, averaged agency ratings, and *Institutional Investor* ratings respectively; *VIX* is the implied volatility index of S&P 500; and $u_{i,t}$ are the iid disturbances. P-values are shown in parentheses, and *, ** and *** denote 10%, 5% and 1% level of significance respectively. Table 3 explanatory power for sovereign spreads. Of the control variables, the liquidity variable and global risk (VIX) are statistically significant with the appropriate signs.

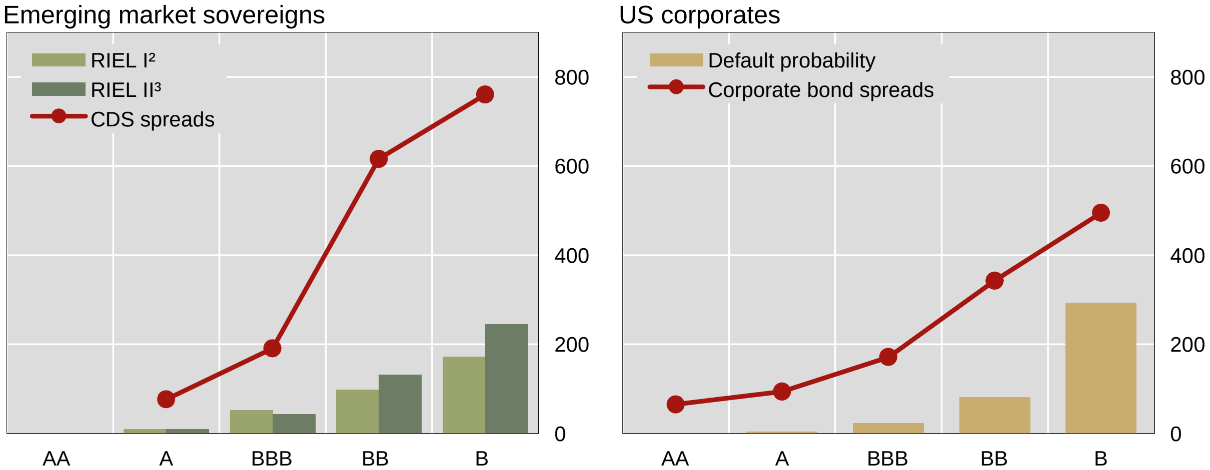
Decomposing sovereign spreads

We now turn to decomposing sovereign spreads into their two components: expected losses and risk premia. In this section, we do so by first calculating expected losses and then subtracting them from averages of spreads over time to arrive at estimates of risk premia for each rating and country. Since our expected losses are based on RIPDs, we call them ratings-implied expected losses (RIELs).

We calculate expected loss by taking the product of the default probability and the average loss-given-default. For the loss-given-default, we rely on historical average recovery rates. Sturzenegger and Zettelmeyer (2005) and Moody's (2006) provide estimates of such recovery rates given default, but the methods for estimating them differ. One method relies on the trading price of a sovereign's bonds 30 days after the first missed interest payment. Another method compares discounted cash flows between the original securities and the new securities received after a distressed exchange. For a given method, the estimated recovery rates also vary widely from one default to another. For example, the recovery rate for the Russian default of 1998 is estimated under the first method at 18% and that for the Dominican Republic default of 2005 at 92%. For the purposes of this article, we take the simple average of recovery rates for the 11 sovereign defaults since 1983 based on the 30-day post default price of the debt. The resulting average recovery rate is 55%, implying a loss-given-default of 45%.

CDS spreads and expected losses[1]

In basis points



[1] January 2002–June 2006. [2] Ratings-implied expected losses based on sovereign default experience. [3] Ratings-implied expected losses based on corporate default experience.

Sources: CreditPro; FitchRatings; Markit; Merrill Lynch; Moody's Investors Service; Standard & Poor's; authors' calculations. Graph 2

What determines sovereign default risk?

In order to investigate the determinants of our measure of expected loss (RIPD), we employ a panel regression framework with fixed effects, using annual data from 1990 to 2005.

We follow the credit risk literature and assume a log-normal functional form, as it is known to fit the fat tails of relevant financial distributions. The models we estimate are of the following specification: $Y_{it} = a_{0i} + a_1 F_{it} + u_{it}$, where Y_{it} represents the natural logarithm of RIPD for country *i* in year *t*. This sovereign risk measure is explained by F_{it}, a vector comprising country-specific fundamentals as well as measures of original sin and currency mismatch created using the international securities statistics of the BIS,[①] u_{it} being the error term.

Explanatory variables	(1)	(2)	(3)	(4)
Log nominal GDP	0.211*** (0.000)	0.324*** (0.000)	0.976*** (0.003)	0.980*** (0.001)
Log GDP per capita	–0.2152*** (0.000)	–0.212*** (0.004)	–0.904** (0.011)	–0.900*** (0.004)
Inflation	0.045*** (0.000)	0.021*** (0.002)	0.019*** (0.004)	0.026*** (0.000)
Current account balance/GDP	0.016*** (0.000)	0.015*** (0.002)	0.014** (0.037)	0.018*** (0.011)
External debt/GDP	0.003** (0.02)	0.002* (0.074)	0.003* (0.077)	–0.000 (0.675)
Political risk		–0.005 (0.188)	–0.012** (0.022)	–0.015*** (0.005)
Years since last default		–0.039*** (0.000)	–0.042*** (0.000)	–0.045*** (0.000)
Original sin			0.309* (0.094)	
Currency mismatch				–0.074*** (0.000)
Adjusted R^2	*0.80*	*0.82*	*0.84*	*0.84*

Note: P-values are shown in parentheses, and *, ** and *** denote 10%, 5% and 1% level of significance respectively; standard errors corrected using White's method. As the currency mismatch variable is simply a scaled version of the original sin measure, they are highly collinear and the panel regressions were estimated separately to ensure robustness (insignificant variable not shown). The political risk variable is constructed so that higher values reflect better conditions.

In regression (1), we only use country-specific fundamentals to explain our RIPD and find that the macroeconomic measures for country size, economic development, inflation, current account balance and external debt are all significant and have the expected signs. Of the qualitative variables added in regression (2), which measure political risk and history of default, only the latter is significant, suggesting that countries with more recent defaults will experience higher expected losses, even after controlling for other fundamentals.

In addition to country-specific fundamentals and debt intolerance perspective, we test whether variables using the BIS data on original sin and currency mismatch help explain our country risk variable (regressions (3) and (4)). The coefficient on the original sin variable, which is meant to measure the inability of a country to borrow abroad in its own currency, is positive and significant, consistent with the concept that countries with a lower capacity to borrow in domestic currency should be riskier. Similarly, the coefficient on the proxy measure for currency mismatch, which is meant to measure the sensitivity of net worth or net income to changes in the exchange rate, is significant with the expected sign, implying that countries whose net asset positions are more vulnerable to exchange rate depreciations have higher expected losses, ceteris paribus.

Overall, while the findings above are consistent with extant sovereign debt studies, they also suggest that the addition of measures of country financial structure using BIS data on the currency denomination of securities issuance significantly contributes to our measurement of sovereign risk.

[①] For details on the creation of these variables, see Borio and Packer (2004). The measure of original sin used here measures the ratio of foreign currency debt to total debt outstanding, assuming that all debt issued in a country's currency should be counted as local currency issuance regardless of the nationality of the issuer. The proxy measure for currency mismatch multiplies the above original sin measure by (reserves – debt) / exports.

Expected losses are a small part of spreads

How large a component of spreads are expected losses? Graph 2 compares average credit spreads to expected losses for varying credit ratings. The left-hand panel does this for emerging market sovereign debt and the right-hand panel for corporate debt. Consistent with the credit spread puzzle in the corporate bond pricing literature, the left-hand panel shows that sovereign spreads are much bigger than measured expected losses. The average RIEL for our sample is 96 basis points. The average CDS spread for our entire sample is 450 basis points, five times the average RIEL based on the sovereign default experience and four times the average RIEL based on the corporate default experience. Even if we made the extreme assumption of a loss-given-default of 100%, the average spread would still be twice the average RIEL.

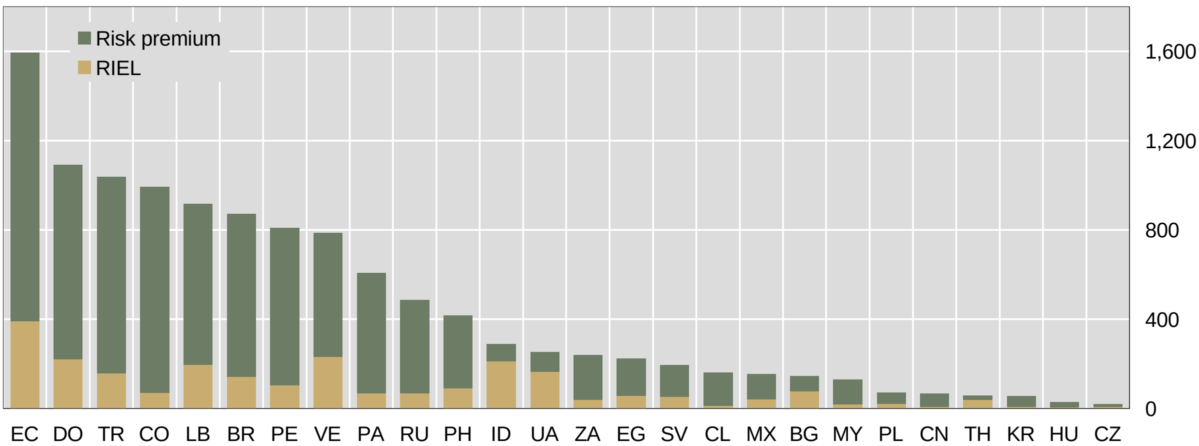
There are clear patterns in the way sovereign spreads and expected losses relate to credit ratings. First, the multiple of spread over expected loss appears to be greater, the higher the country's credit quality. For example, Korea, which is rated single-A, has an average CDS spread of 55 basis points, more than 17 times one estimate of RIEL and seven times the other estimate. Second, average spreads tend to be wider than average RIELs at every letter rating. Third, both average spreads and average RIELs widen as credit ratings decline. Finally, spreads widen more dramatically with lower ratings, and hence the differential between them and expected losses becomes larger.

Credit spread puzzle more pronounced for sovereigns

Comparing the left-hand and right-hand panels, it is evident that spreads on sovereign debt have on average been wider than those on corporate debt for each given rating and relative to estimates of expected losses. In other words, the credit spread puzzle is more pronounced for sovereign debt than for corporate debt. One possible reason for this, as suggested earlier, is that it is more difficult to diversify idiosyncratic default risk for sovereign bonds than for corporate bonds, because there are far fewer issuers of the former than of the latter. Hence, such idiosyncratic risk is priced in the wider spreads on sovereign bonds.

RIEL and risk premia[1]

In basis points



BG = Bulgaria; BR = Brazil; CL = Chile; CN = China; CO = Colombia; CZ = Czech Republic; DO = Dominican Republic; EC = Ecuador; EG = Egypt; HU = Hungary; ID = Indonesia; KR = Korea; LB = Lebanon; MX = Mexico; MY = Malaysia; PA = Panama; PE = Peru; PH = Philippines; PL = Poland; RU = Russia; SV = El Salvador; TH = Thailand; TR = Turkey; UA = Ukraine; VE = Venezuela; ZA = South Africa.

[1] Calculated as the difference between the CDS spread and the ratings-implied expected loss (RIEL).

Sources: FitchRatings; Markit; Moody's Investors Service; Standard & Poor's; authors' calculations. Graph 3

Risk premia account for the larger part of spreads ...

How about sovereign risk premia? Graph 3 shows these risk premia, which are calculated by subtracting expected losses from sovereign debt spreads. In nearly all cases, estimated risk premia are positive. The estimates confirm what one would expect: lower sovereign ratings tend to command higher risk premia. More interestingly, they tend to account for a larger part of the spread than do expected losses. When we calculate risk premia on the basis of the RIEL derived from sovereign defaults, the average risk premium for our sample of countries is 365 basis points, accounting for about four fifths of the spread. When we calculate it on the basis of the RIEL derived from corporate defaults, the average risk premium is 342 basis points, constituting more than two thirds of the spread.

One additional factor is worth noting about our calculation of sovereign risk premia. These premia are derived from averages of sovereign spreads over a period in which such spreads have been relatively low. This factor serves to bias downwards our estimates of risk premia. Even so, these estimates imply that risk premia tend to account for the larger part of sovereign spreads.

Conclusions

To interpret sovereign spreads, we make a clear distinction between sovereign risk and risk premia as the price of that risk. The spreads themselves can be divided into two components: expected losses from default and risk premia.

We propose default probabilities as a measure of sovereign risk and offer illustrative estimates based on information from the historical performance of sovereign and corporate credit ratings. We find our estimated measure of sovereign risk to be a significant determinant of the cross-sectional variation of sovereign spreads. However, it does not fully explain spreads because the price of risk is itself a separate determinant.

... even when spreads are low

We estimate expected losses by taking the product of default probabilities and average sovereign loss-given-default. These expected losses turn out to be a relatively small part of average sovereign spreads. Indeed, they tend to be a smaller part of spreads than are expected losses for corporate bonds, suggesting a “credit spread puzzle” that is more pronounced for sovereign debt than for corporate debt. The size of expected losses implies that risk premia account for the larger part of average sovereign spreads even during a period when such spreads have been relatively low.

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Exchange rates and global volatility: implications for Asia-Pacific currencies[1]

At times of heightened global equity and bond market volatility, high-yielding currencies tend to depreciate while low-yielding ones tend to serve as a "safe haven". The whole spectrum of sensitivity to global volatility is represented among Asia-Pacific currencies.

JEL classification: F3, G1.

Emerging market and industrial currencies offering relatively high yields tended to appreciate in 2006 against lower-yielding currencies, except during the sell-off in May and June. Among Asia-Pacific currencies, the Indonesian rupiah, the Philippine peso and even the Australian dollar offer examples of this pattern. Thus, much of the year's trading added to the increasing body of findings (once considered anomalous) that higher-yielding currencies tend to appreciate against lower-yielding currencies (Hodrick (1987), Froot and Thaler (1990), Lewis (1995), Engel (1996), Remolona and Schrijvers (2003)). The sell-off of high-yielding currencies in May 2006, by contrast, supported Irving Fisher's earlier thesis that the higher-yielding currency would tend to depreciate, over time, against the lower-yielding currency, offsetting the yield advantage.

This alternating currency performance forms part of a broader pattern in which a spectrum of high-yielding currencies tend at times to be stable or firming and at other times to depreciate against their low-yielding counterparts. Kumar and Persaud (2002) used these alternating phases of currency returns to define states of low or high risk aversion. Others have since constructed indicators of risk aversion directly from interest rate spreads or capital market volatility (Tarashev et al (2003); see Illing and Aaron (2005) for a survey). Most large international banks now publish risk/volatility or risk aversion indicators for their clients. Currency strategists like Davies (2005) have often related currency returns to such risk indicators in their daily work.

This special feature investigates the relationship between exchange rates and global capital market volatility, and draws some implications thereof for Asia-Pacific currencies. It first reports patterns of exchange rate responses

[1] Our thanks go to Eric Chan for research assistance and members of the EMEAP Forum for discussion. All errors remain the responsibility of the authors. The views expressed are those of the authors and not necessarily those of the BIS.

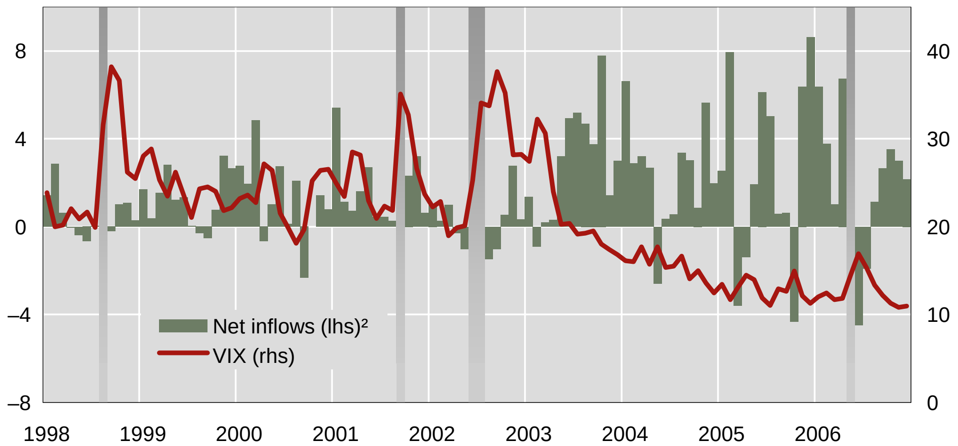
among a broad range of currencies during specific recent episodes of heightened volatility. It then considers exchange rate sensitivity to volatility more generally by relating weekly changes in these currencies to the corresponding weekly changes in capital market volatility over the period 2000–06 as a whole. These two analyses find that certain currencies tend to be stable or to appreciate as volatility rises, while others tend to weaken at such times. The third part of this article relates the differences in currency responses to certain country characteristics. This cross-sectional analysis finds that economies offering higher short-term interest rates tend to see their currencies depreciate against lower-yielding currencies in periods of rising capital market volatility. This regularity poses challenges to Asian exchange rate stability that are discussed in the conclusion.

Exchange rate movements in episodes of higher volatility

In four episodes of high global volatility ...

Is there any discernible pattern across currencies in their responses to bouts of rising volatility? Looking back over the past decade, there were several notable episodes of heightened global volatility, as indicated by sharp spikes in indicators such as the widely used VIX index (Graph 1).[2] These episodes occurred in August 1998, September 2001, June–July 2002 and May

Global volatility and Asian net equity inflows[1]



Note: The shaded areas refer to four episodes of rising volatility; see footnote 3 in the main text.

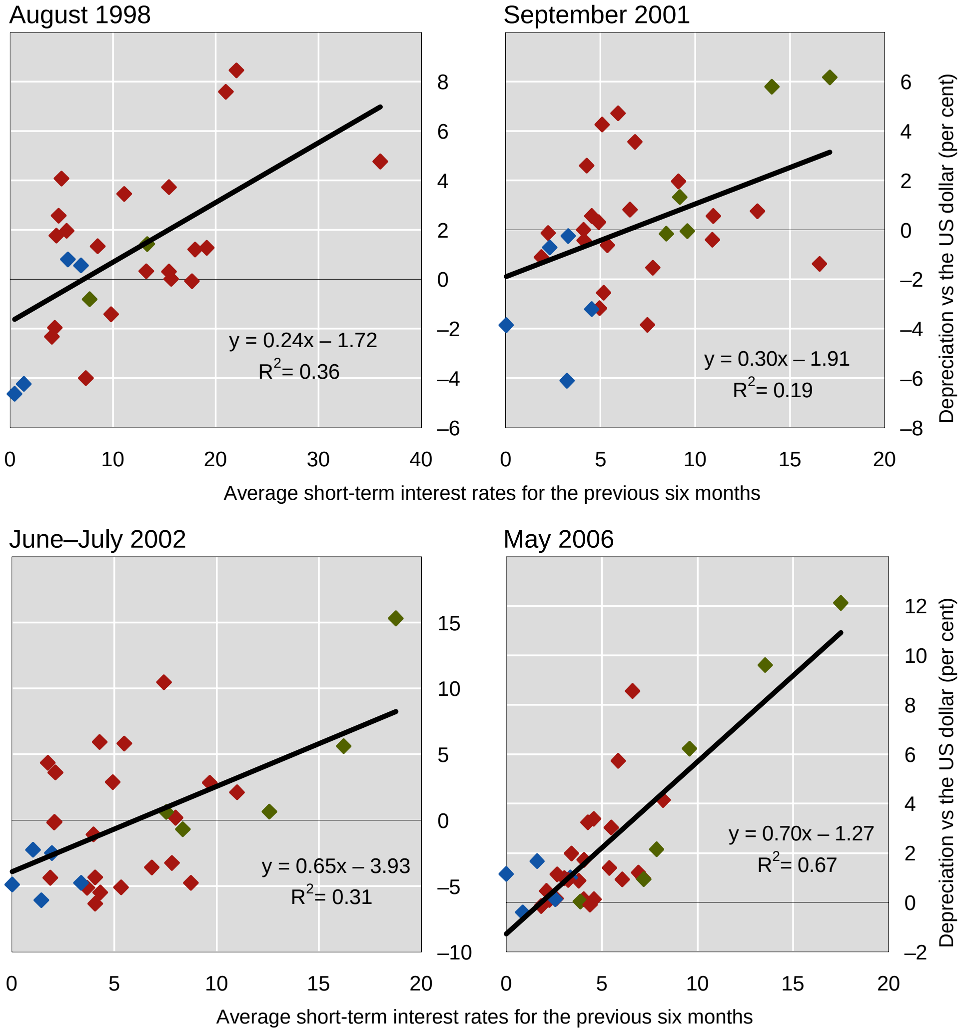
[1] Net foreign purchases of equities in India (data start in 1999), Indonesia, Korea, the Philippines, Taiwan (China) and Thailand. [2] In billions of US dollars.

Sources: Bloomberg; CEIC; Korean Stock Exchange; Philippines Stock Exchange; BIS calculations.

Graph 1

[2] The Chicago Board Options Exchange Volatility Index (VIX) is a measure of market expectations of near-term volatility, as conveyed by S&P 500 stock index option prices. VIX has been used by many as a barometer of investor sentiment and market volatility since its introduction in 1993.

Currency depreciation and interest rates in volatile periods



Note: Blue markers represent generally low-yielding currencies (JPY, CHF, SGD, TWD and EUR), while green markers represent relatively high-yielding ones (TRY, IDR, BRL, RUB, PHP and INR). Currencies with interest rates above 40% and those fixed to the USD are excluded. The HKD 12-month forward and CNY 12-month NDF are used to represent HKD and CNY respectively. For August 1998, the inclusion of RUB, IDR, TRY, ARS and BRL results in a slope of 0.1266 and R^2 of 0.0393. For September 2001, the inclusion of TRY, ARS and MYR results in a slope of 0.1201 and R^2 of 0.2390. For June–July 2002, the inclusion of TRY, ARS and MYR results in slope of 0.2399 and R^2 of 0.2735. Interest rates are either money market rates (60b) or treasury bill rates (60c) from the IMF.

Sources: IMF, *International Financial Statistics*; Bloomberg; BIS calculations. Graph 2

2006.[3] This last episode did not show up as an especially sharp spike in absolute terms, but it nonetheless represented a larger than normal rise in the volatility index relative to the low levels prevailing at the time.[4] In emerging

[3] The episodes are defined as starting with either a discrete event (eg the Russian default, the 11 September terrorist attacks) or when the VIX index first deviated by more than one standard deviation from its three-month moving average. The episodes are defined as ending with the first peak of the VIX index. The four episodes are thus dated: 17–31 August 1998 (Russian default), 10–20 September 2001 (terrorist attacks), 3 June–23 July 2002 (multiple factors, including geopolitical tensions and the WorldCom accounting scandal and bankruptcy) and 11–23 May 2006 (multi-market sell-off).

[4] The May 2006 episode saw only a 5.8 point increase in the VIX, compared to the 12.4, 11.9 and 21.6 point rises in the three earlier episodes.

Asia, the May 2006 episode also saw heavier net sales of equities by non-residents than in the earlier episodes.

… high-yielding currencies tended to depreciate …

During these four episodes, currencies performed in a manner qualitatively consistent with Irving Fisher's hypothesis. Relatively low-yielding currencies such as the Swiss franc (traditionally a "safe haven" currency), the yen and the euro generally appreciated against the US dollar (Graph 2), while higher-yielding currencies such as the Russian rouble, Brazilian real and Turkish lira tended to depreciate. The responses of Asia-Pacific currencies in these episodes offer some further evidence in support of this global dichotomy: low-yielding currencies such as the New Taiwan dollar and the Singapore dollar depreciated relatively little or appreciated in some cases, while higher-yielding ones such as the Indonesian rupiah and the Philippine peso weakened in most episodes. The moderately high-yielding Australian and New Zealand dollars also tended to depreciate. However, the pattern among other currencies is not as clear-cut. For instance, the Indian rupee and the Korean won reacted in a mixed fashion across episodes.

… especially in May 2006

Looking across the episodes, the link between currency performance and average interest rate levels prior to the episode was the tightest during the relatively mild (in terms of the point increase in the VIX) May 2006 episode. Along the least-squares line, currency depreciation over eight business days cost investors about eight months of interest rate premium. Admittedly, industrial economy currencies tended to rise less against the US dollar compared to the experience in the three earlier episodes. In particular, the weakness of the yen against the dollar in May 2006 left emerging Asian currencies especially exposed to the rise in global volatility, given these currencies' tendency to co-move with the US dollar/yen rate (Kawai (2002), Ho et al (2005)).

Regression analysis of volatility and currency performance

Weekly changes in bilateral and effective exchange rates show …

Stepping back from specific episodes, how does currency performance relate to changes in global volatility in general? To assess a currency's overall sensitivity, we regress the percentage change in the currency's exchange rate on the change in global volatility. To control for the regular response of the currency to the movements among the major currencies, the percentage changes in the yen and the euro against the US dollar are included as additional explanatory variables. 34 currencies, including 13 Asia-Pacific currencies, are included in the analysis.[5] Both the bilateral US dollar exchange rates of these currencies and their nominal effective exchange rates (NEERs) were assessed. Two different volatility indicators are considered: the VIX and a

[5] These economies' currencies are covered: Argentina (ARS), Australia (AUD), Brazil (BRL), Canada (CAD), Chile (CLP), China (CNY), Colombia (COP), the Czech Republic (CZK), Denmark (DKK), Hong Kong SAR (HKD), Hungary (HUF), India (INR), Indonesia (IDR), Israel (ILS), Japan (JPY), Korea (KRW), Malaysia (MYR), Mexico (MXN), New Zealand (NZD), Norway (NOK), the Philippines (PHP), Poland (PLN), Russia (RUB), Singapore (SGD), Slovakia (SKK), South Africa (ZAR), Sweden (SEK), Switzerland (CHF), Taiwan, China (TWD), Thailand (THB), Turkey (TRY), the United Kingdom (GBP), the United States (USD) and the euro area (EUR).

composite implied volatility index ("composite index"). While the VIX is derived from US stock market volatility only, the composite index is a more global indicator averaging eight measures of equity and bond market volatility in four major economies.[6] The regressions were performed on weekly changes (Wednesday to Wednesday) over the period January 2000 to December 2006.

Table 1 reports the two sets of estimated coefficients, which indicate the percentage change in the bilateral US dollar exchange rate that is associated with a 1 point change in the two volatility indicators, controlling for changes in the euro's and yen's value against the US dollar.[7] For instance, the estimated sensitivity of the Indonesian rupiah towards the VIX of 0.112 means that, on average, the currency would depreciate by 0.56% in the presence of a 5 point rise in the VIX. Unsurprisingly, this period average result is an order of magnitude smaller than the rupiah's actual movement in May 2006, when the VIX rose by about 5 points.

Regression coefficients relating volatility to US dollar exchange rates

Currency	VIX	Composite	Currency	VIX	Composite	Currency	VIX	Composite
ARS	0.065	0.074	HKD F	0.005	0.011*	PHP	**0.054****	0.085*
AUD	0.140***	**0.347*****	HUF	**0.053****	0.050	PHP NDF	0.035	0.097
BRL	0.144**	**0.307*****	IDR	**0.112*****	0.219***	PLN	0.138***	0.299***
CAD	**0.069*****	0.177***	IDR NDF	0.159***	0.281***	RUB	–0.004	0.024
CHF	–0.057***	–0.114***	ILS	0.059***	0.121***	SEK	0.089***	0.211***
CLP	0.155***	0.298***	INR	0.020**	0.041**	SGD	0.014	0.024
CNY	0.005	0.042	INR NDF	0.045***	0.099***	SKK	0.045***	0.108***
CNY NDF	0.004	0.029*	JPY	0.002	–0.068	THB	0.010	0.056*
COP	0.087***	0.186***	KRW	0.092***	0.182***	TRY	0.285***	0.614***
CZK	0.014	0.071**	KRW NDF	0.033*	0.128***	TWD	0.031**	0.056**
DKK	0.000	0.003	MXN	0.065**	0.116**	TWD NDF	0.020	0.054*
EUR	–0.046	–0.121*	MYR	–0.015	0.029	ZAR	0.044	0.236**
GBP	0.004	–0.016	NOK	–0.000	0.080**	PAIF	0.019**	0.059***
HKD	–0.001	0.000	NZD	0.076**	0.204***	ADXY	0.016**	0.053***

Note: The coefficients indicate the percentage change in the corresponding currency associated with a 1 percentage point change in the volatility index. * Significant at the 10% level. ** Significant at the 5% level. *** Significant at the 1% level. NDFs are three-month rates except for CNY NDF (12-month). HKD F is the 12-month forward rate. PAIF and ADXY are composite Asian currency indices. The sample period is 2000–06, except ARS (from June 2003) and spot CNY, HKD and MYR (from 22 July 2005). Exchange rates and volatility indicators are New York closes. Spot rates that trade only in Asian hours (CNY, IDR, INR, KRW, MYR, PHP and TWD) enter the regressions with a one-day lead. For the EUR and JPY regressions, only the volatility indicator is used as explanatory variable. A full version of this table including the estimated coefficients on JPY and EUR changes is available upon request.

Sources: Bloomberg; BIS calculations. Table 1

[6] The composite index is the simple average of the equity and 10-year swap implied volatilities for each of the United States, the United Kingdom, the euro area and Japan, ie eight series in total, with the VDAX serving as the equity volatility for the euro area. Here no attempt is made, as in Tarashev et al (2003), to decompose investors' risk aversion from the level of risk per se. Regression analysis of currency changes against their monthly index of risk aversion produced similar results.

[7] In terms of direction, a positive coefficient on the volatility indicator means the currency tends to depreciate when the volatility indicator rises in value (ie more risk or risk aversion).

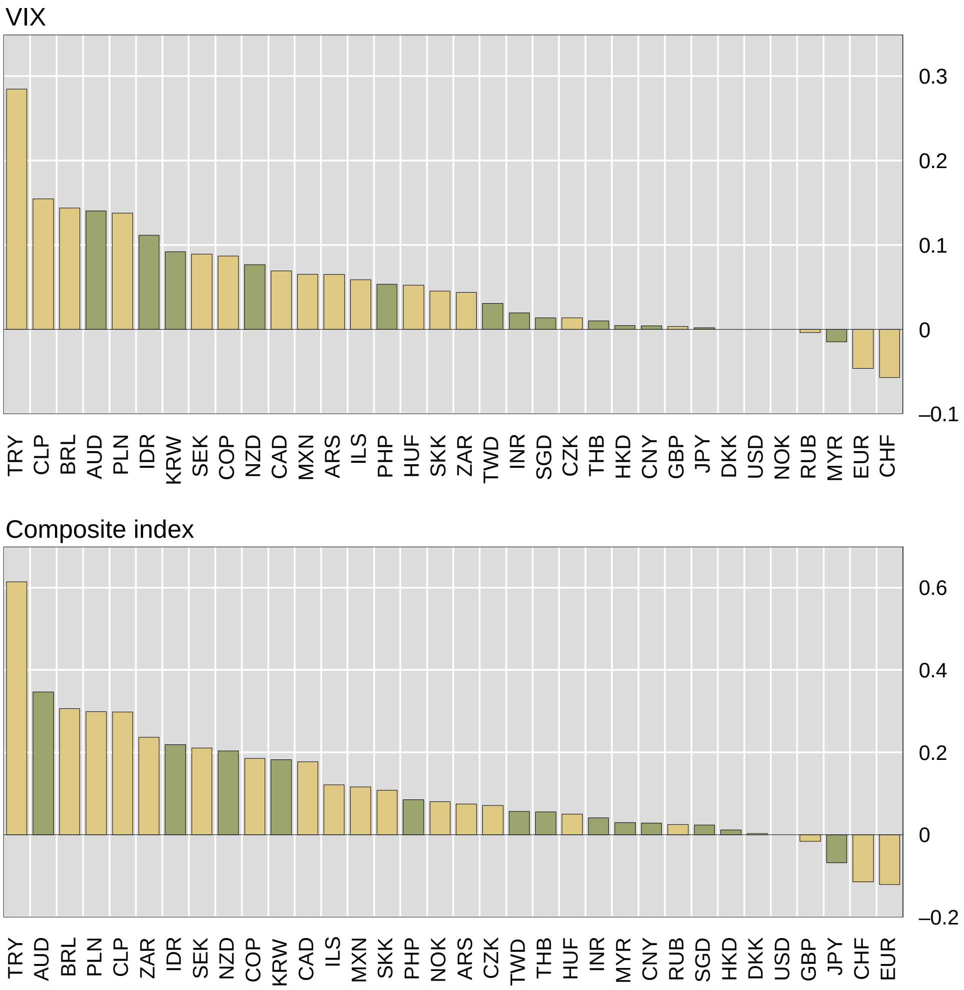
There are several notable observations. First, even when the often significant influences of yen and euro movements are controlled for, most of the currencies still exhibit significant sensitivity towards at least one of the two volatility indicators. Estimated sensitivities to the composite index tend to be more statistically significant than those to the more volatile VIX. However, the differences between the two sets of estimated sensitivity should not be overstated – the correlation between the two is 0.96 while the Spearman rank correlation is 0.91.

… patterns of significant responses of currencies …

Second, the regression results are generally in line with observations from the episodic analysis above. The Swiss franc, the euro and to some degree the yen tend to have negative sensitivities towards the volatility indicators, meaning that they tend to strengthen against the dollar when volatility rises (Graph 3).[8] By contrast, emerging market currencies generally depreciate in an environment of elevated volatility. Overall, the Turkish lira stands out for its

… in line with experience during specific episodes

US dollar exchange rate sensitivity to global volatility, 2000–06



Note: HKD is the HKD 12-month forward, while CNY is the CNY 12-month NDF.

Source: See Table 1. Graph 3

[8] Including the euro and yen as controls in the non-major currency equations leaves their ranking of estimated sensitivities invariant to the choice of numeraire.

high sensitivity. Among the Asia-Pacific currencies, the Australian, Indonesian, Korean, New Zealand and Philippine currencies show relatively high sensitivities to changes in global volatility.

Mixed evidence of exchange rate management constraining sensitivity

Third, there is some, albeit mixed, evidence that currency management constrains exchange rate responses to changes in global capital market volatility. If exchange rate management by the authorities, as in much of Asia for example, constrains the response of the spot exchange rate, it is potentially informative to try to measure the response of forward rates or offshore non-deliverable forward (NDF) rates. Ma et al (2004) show that, owing to capital restrictions, Asian NDFs are generally not tightly bound by arbitrage to the more controlled spot exchange rates. Consequently, NDF volatilities tend to be higher than spot rate volatilities. Accordingly, the Indian rupee and Indonesian rupiah NDFs have higher estimated sensitivities than the respective spot rates (Table 1). Even for the Hong Kong dollar, whose pegged spot rate hardly responds to changes in volatility, the more volatile one-year forward rate shows a small but statistically significant sensitivity to the composite indicator. However, stronger responses are not obtained for the NDFs of the Chinese renminbi, the New Taiwan dollar and the Philippine peso, for which spot market intervention is generally thought to be quite frequent and capital controls still effective.

Bilateral dollar and effective exchange rates results are consistent

Finally, the effective exchange rates of most currencies tend to be less sensitive to volatility than their bilateral rates, owing to the collective weight of trading partners' currencies that also depreciate when volatility rises. For the same reason, currencies with low or negative bilateral exchange rate sensitivities to volatility tend to have effective exchange rates that appreciate even more than their bilateral dollar rates for a given rise in the volatility indicator. The US dollar depreciates very slightly in effective terms in response to rises in the VIX or composite index. Overall, the results using the bilateral US dollar exchange rates and the effective exchange rates are quite similar.[9]

The determinants of currency sensitivity to global volatility

Out of 11 factors tested …

What factors underlie these measured sensitivities to global volatility? As seen already in the episodic analysis, currency reactions seem to relate to the prevailing short-term interest rate levels. To answer this question more systematically, we perform a strictly cross-sectional analysis, relating the estimated sensitivities to various economic characteristics over the entire 2000–06 period.[10]

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[9] The bilateral-NEER correlation is 0.92 in the case of the VIX and 0.96 in the case of the composite indicator. The Spearman rank correlation coefficients are very close at 0.91 and 0.95, respectively.

[10] The estimated sensitivities for 33 currencies are included in this analysis (MYR is excluded). For Hong Kong SAR and China the 12-month forward rate and 12-month NDF respectively are used. Short-term interest rates are money market rates as defined by the IMF *International Financial Statistics*. IMF data are used for current account as a percentage of GDP (2000–06 average), GDP per capita (in USD terms at market exchange rates) and inflation. The net international investment position (NIIP) as a percentage of GDP (2000–06 average) is from Lane and Milesi-Ferretti (2006). 2004 data are used for 2005 and 2006. Foreign exchange

These variables are chosen to capture four broad types of factors that could potentially affect currency sensitivity to changes in global volatility: "carry" (relative interest rates), depreciation and credit risks, external financing requirements and liquidity. For "carry", both short-term interest rates and the inflation rate are included to determine whether international investors are attracted by nominal or real returns.[11] Depreciation risk is proxied by the ratio of reserves to imports, while creditworthiness is proxied by the credit rating and GDP per capita. Financing requirements are captured on a stock basis by the net international investment position (NIIP) and on a flow basis by the current account.[12] Liquidity is represented by each currency's turnover, both in US dollar terms and in relation to trade, to non-resident portfolio investment and to non-resident equity portfolio investment.[13]

Some high bivariate correlations between these economic variables and the estimated currency sensitivities are observed. The short-term interest rate variable shows the strongest correlation (over 0.75), followed by inflation (over 0.6) and NIIP as a ratio to GDP (stronger than –0.44). The credit rating, GDP per capita, current account balance and FX market liquidity show correlations between 0.25 and 0.4 in absolute value.

When these variables showing strong bilateral correlations are made to compete against each other in a multiple regression framework, a remarkably parsimonious empirical account of the sensitivities emerges (Table 2).[14] Two findings stand out.

… short-term interest rates dominate …

First, even after controlling for other economic variables, the short-term interest rate dominates, showing a very significant positive association with currency sensitivity. One way of reading this finding is that investment strategies that target high-yielding currencies (eg carry trades) are vulnerable to rises in global volatility. Inflation, which is highly correlated with the level of interest rates across countries, seems to play no independent role.

Admittedly, high-inflation and high-interest-rate currencies in the sample (the Brazilian real and the Turkish lira) contribute to this strong cross-sectional relationship between interest rate level and currency sensitivity. But even if

(FX) market liquidity is proxied by the FX turnover of each currency from the 2004 triennial survey (BIS (2005)).

[11] It might seem that interest rate spreads, rather than levels, would be the appropriate regressor. However, the correct base currency for calculating the spread would have to be fine-tuned currency by currency, taking into account the "betas" with respect to the yen and the euro. Recall that the regression analysis in effect works on the difference between each currency's interest rate and the sample average. If our not fine-tuning each currency's spread is considered an error in the variable, then the usual analysis applies: the coefficient on short-term interest rates would be biased towards zero.

[12] See IMF (2006, p 14) for the relationship between emerging market currency performance in May–June 2006 and the current account deficit.

[13] IMF (2006, p 13) suggests a variant of the latter two variables that includes only investment in local currency bonds and equities as an operationalisation of the notion of "crowded trade", that is, a position with potentially large reversals in relation to the liquidity of one of the underlying markets.

[14] The remaining four variables with low bilateral correlations were tested and found to be jointly insignificant.

US dollar exchange rate sensitivity and macroeconomic variables: estimation results								
	VIX				Composite index			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
Intercept	0.008	–0.017	0.055	0.011	–0.051	–0.041	–0.110	0.025
	(0.916)	(0.790)	(0.188)	(0.310)	(0.769)	(0.778)	(0.250)	(0.297)
Short-term interest rate	0.008***	0.006***	0.006***	0.007***	0.018***	0.014***	0.014***	0.014***
	(0.002)	(0.000)	(0.000)	(0.000)	(0.003)	(0.000)	(0.000)	(0.000)
Inflation	–0.003	-	-	-	–0.005	-	-	-
	(0.369)	-	-	-	(0.517)	-	-	-
Credit rating (log)	–0.010	-	-	-	0.040	-	-	-
	(0.751)	-	-	-	(0.581)	-	-	-
GDP per capita (log)	0.013	0.013	-	-	0.019	0.028	-	-
	(0.216)	(0.150)	-	-	(0.435)	(0.188)	-	-
NIIP/GDP	–0.025	–0.023*	–0.019	–0.022*	–0.059	–0.046	–0.037	–0.042
	(0.208)	(0.060)	(0.120)	(0.073)	(0.193)	(0.105)	(0.185)	(0.124)
Current account/GDP	0.000	-	-	-	0.002	-	-	-
	(0.986)	-	-	-	(0.735)	-	-	-
FX market liquidity (log)	–0.009	–0.010*	–0.004	-	–0.021	–0.019	–0.009	-
	(0.117)	(0.077)	(0.271)	-	(0.119)	(0.120)	(0.357)	-
Adjusted R^2	0.591	0.621	0.605	0.602	0.550	0.578	0.566	0.568

Note: Specifications (2) – (4) exclude variables with the highest p-values in the previous specification.

Sources: Lane et el (2006); IMF, *International Financial Statistics*; Bloomberg; BIS (2005); BIS calculations. Table 2

these extreme observations are removed from the sample, the positive relationship still holds, indeed, to the exclusion of the other surviving variable in the full-sample case (Graph 4).

… while external financing also matters

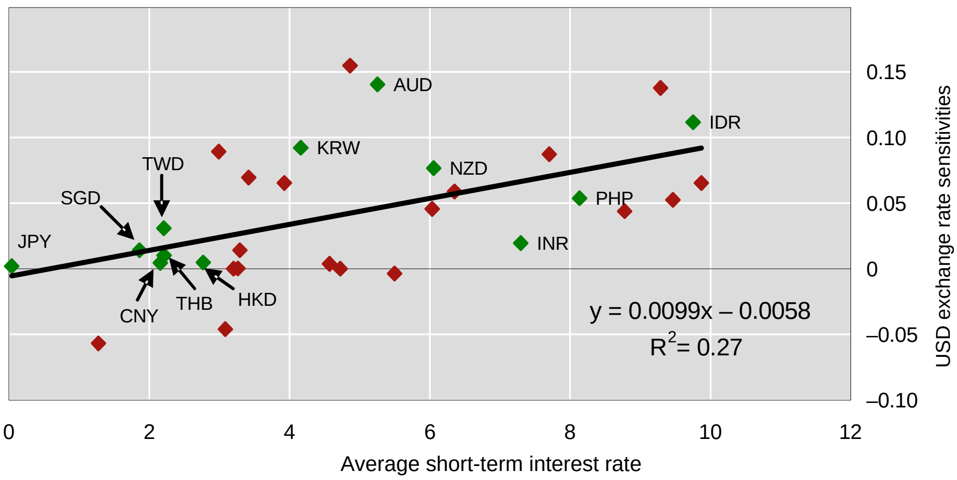
Second, balance of payments fundamentals are found to have some, but less consistent, influence over currency sensitivity. The NIIP in relation to GDP (though not the current account) survives the multiple regression analysis of sensitivity to the VIX. The larger an economy's net international liabilities, the more prone its currency is to depreciation in volatile times. This result lends some support to the widespread view that long currency positions tend to be cut back in periods of rising global volatility, leading to potentially larger declines in currencies with heavier debt burdens to roll over.

The two main findings above help to make sense of the different sensitivities among Asia-Pacific currencies. The Australian and New Zealand dollars, with relatively high interest rates and large external liability positions, are hit hard by upsurges in global volatility. In contrast, even though interest rates are also high in Indonesia and the Philippines, the influence of rising global volatility may be offset to some extent by the ongoing contribution of the two economies' current account surpluses to their external positions. In the rest of Asia, generally lower interest rates and external surpluses tend to limit currency sensitivity to changes in global volatility.

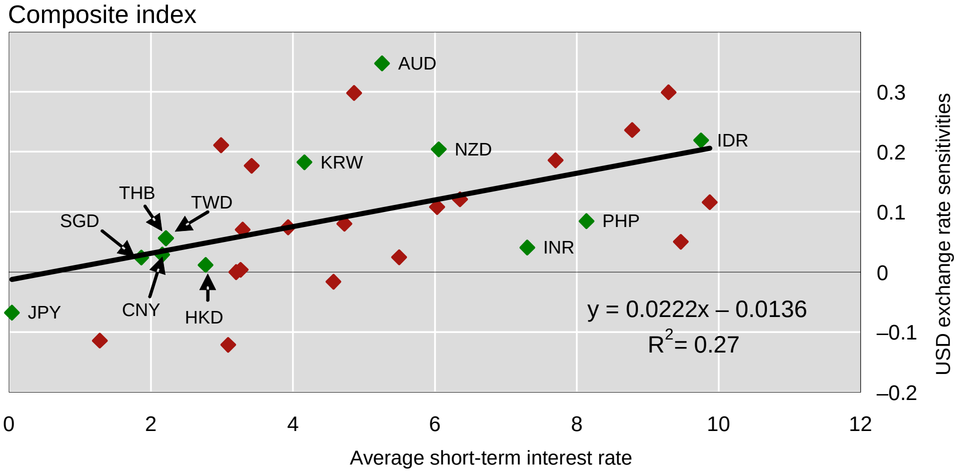
Currency sensitivities and interest rate levels

2000–06; sample excluding Brazilian real and Turkish lira

VIX



Composite index



Note: The p-value on the interest rate is 0.0026 in the case of the VIX and 0.0027 in the case of the composite index.

Sources: IMF, *International Financial Statistics*; Bloomberg; BIS calculations. Graph 4

Caveats may apply

Still, the statistical link traced above between interest rate levels and balance of payments fundamentals, on the one hand, and currency sensitivity, on the other, may not represent the final word. For instance, threshold effects and non-linearities may play an unexplored role. Moreover, differences in the style and intensity of exchange rate management by the authorities have not been formally accounted for in the cross-sectional analysis. To some extent, the currencies that respond to volatility may be the ones that are allowed to do so. One approach to account for exchange rate management would be to include a measure of exchange rate flexibility as an explanatory variable. However, such measures (eg realised currency volatility) could approximate the currency sensitivities that are to be explained, so that their use would risk circularity. The mixed results above from comparing NDF and spot rate sensitivities suggest that our omission of exchange rate management may not be too harmful. Still, caution in interpreting these results is called for.

Conclusions

Both episodic and regression analysis of the years 2000–06 provides evidence of a systematic pattern of sensitivities of various currencies to changes in global capital market volatility. Much of this pattern of currency sensitivities can be accounted for by the level of short-term interest rates and, to a lesser extent, the scale of net international liabilities.

Looking across the Asian currencies, there is some prospect for them to respond more similarly to changes in global volatility. Thus far in this century, the higher interest rate currencies, the Indonesian rupiah and the Philippine peso, have been somewhat sheltered from changes in global volatility by their responsiveness to the yen. Nevertheless, shifts in global volatility tend to strain cross rates between such currencies and lower-yielding Asian currencies. Going forward, the convergence of inflation rates in the region would tend to reduce interest rate differentials. This would in turn tend to narrow the current differences in the response of various currencies in the region to a change in global volatility.

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Financial investors and commodity markets[1]

Commodities have attracted considerable interest as a financial investment in recent years. This article discusses the factors behind their growing appeal and assesses the extent to which market characteristics, such as price volatility, have changed as a result. The feature concludes that commodity markets have become more like financial markets in terms of the motivations and strategies of participants, but that the physical characteristics of commodity markets are still important.

JEL classification: G11, G15, Q41.

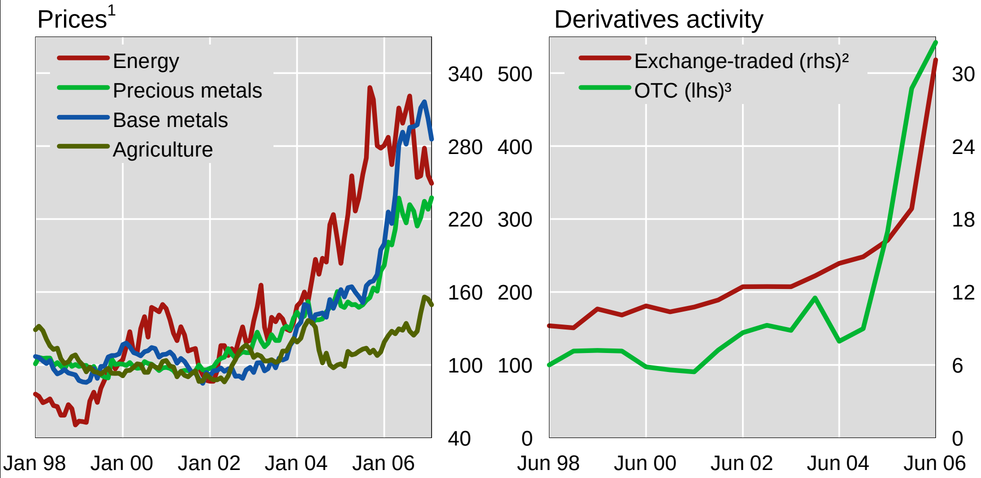
The sharp increase in commodity prices, especially for energy and base metals since 2002, has gone hand in hand with growing derivatives market activity (Graph 1). The number of contracts outstanding in exchange-traded commodity derivatives almost tripled from 2002 to 2005. Over-the-counter (OTC) trading of commodity derivatives also grew rapidly. According to BIS statistics, the notional value of OTC commodity derivatives contracts outstanding reached $6.4 trillion in mid-2006, about 14 times the value in 1998 (BIS (2006)). At the same time, the share of commodities in overall OTC derivatives trading grew from 0.5% to 1.7%.

Along with the rapid increase in commodity derivatives trading, the presence of financial investors in commodity markets has grown rapidly over the past few years. While commodity market investment is still small relative to overall managed funds, it is large relative to commodity production. In addition, there are indications that the types of financial investors and the strategies they employ have changed.

These developments raise the question of whether growing investor presence has altered the character of markets that are of key importance for the global economy. Understanding the nature of the changes in investor types and strategies is an important step in this regard. The first part of this article documents the increasing role of financial investors in commodity markets, while the second presents some evidence about changes in the motivations of market participants. The third section looks at the effect these changes may have had on the dynamics of commodity prices. The feature concludes that

1 The views expressed in this article are those of the authors and do not necessarily reflect those of the BIS. We are grateful to Anna Cobau and Emir Emiray for excellent research assistance.

Commodity prices and derivatives activity



[1] Goldman Sachs Commodity Index (GSCI) sub-indices, monthly averages; 1998–2002 average = 100. [2] Number of contracts outstanding, in millions. [3] Notional amounts deflated by the GSCI, June 1998 = 100.

Sources: Datastream; BIS. Graph 1

while physical characteristics, such as inventory levels and the marginal cost of production, remain important, commodity markets have become more like financial markets in terms of the motivations and strategies of participants.

The presence of financial investors in commodity markets

Financial activity is large relative to physical markets

Financial activity in commodity markets is large compared with the size of physical production and has grown much faster in recent years. For gold, copper and aluminium, the volume of exchange-traded derivatives was around 30 times larger than physical production in 2005 – a significant increase in this ratio from 2002 (Table 1). The much lower ratio for crude oil may understate the relative size of financial activity, given that OTC markets are particularly important for this commodity. Bank of England market contacts suggest that up to 90% of swaps and options trading in oil is done over the counter, reflecting the need for tailored contracts and a lack of organised derivatives markets for certain types of crude oil (Campbell et al (2006)).

Traditional arbitrage ...

Traditionally, specialised financial traders in commodity markets focused on exploiting arbitrage opportunities (Kolb (1997)). Typically, such opportunities arise as the consequence of commercial investors seeking to hedge their production or consumption in futures markets. These arbitrage trades, usually conducted by specialised commodity traders, typically involve taking long or short positions in forward markets for specific commodities and offsetting positions in spot markets. In doing so, financial investors provide liquidity in commodity derivatives markets.

... limited by constraints on short selling

Normally in financial markets, opportunities for (risk-free) arbitrage exist when the futures price deviates from the relevant spot price plus the cost of carry, eg the cost of financing a position in the spot market. However, the scope for arbitrage in commodity markets may be limited by constraints on short selling. In particular, the stock of commodities available for lending is generally small for energy and base metals. This limitation allows the futures price to fall below the spot price – a situation known as backwardation (Duffie (1989)).

Passive investment strategies ...

The current upturn in commodity prices has been accompanied by greater variety in the types of financial investors and investment strategies in commodity markets (Holmes (2006)). One rapidly growing area is passively managed investment and portfolio products, which is consistent with investors now viewing commodities as an attractive separate asset class. By mid-2006, around $85 billion of funds were tracking the Goldman Sachs Commodity Index (GSCI) and the Dow Jones/AIG Index, two important commodity indices (Holmes (2006)).

... can provide diversification benefits

Passively managed investments often pursue a fully collateralised long-only futures strategy. This can be attractive to institutional investors with a longer-term investment horizon, such as pension funds, for several reasons (Beenen (2005)). First, this strategy allows diversification into commodities at a relatively low cost. Historically, commodity prices have had a relatively low correlation with prices in other asset classes and a high correlation with

Indicators of financial and physical activity in selected commodity markets in 2005

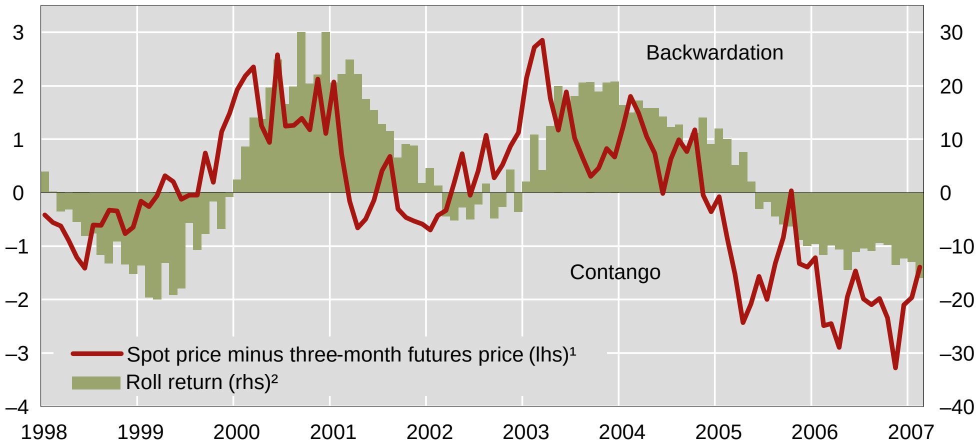
	Financial activity				World production[2]		Ratio[3]	
	Futures		Options					
	Volume[1]	% chg since 2002	Volume[1]	% chg since 2002	2002	2005	2002	2005
Crude oil	93.0	34.4	14.8	27.2	67.0	73.6	3.2	3.9
Of which: NYMEX	59.7	30.6	14.7	28.5				
ICE	30.4	41.5	0.0	–69.7				
Gold	34.5	16.8	2.9	49.7	2.6	2.5	21.8	32.0
Of which: TOCOM	18.0	–12.4	0.3	.				
COMEX	15.9	76.2	2.9	48.3				
Aluminium	33.3	25.2	4.1	368.3	26.1	23.0	22.7	27.3
Of which: LME	30.4	36.3	4.1	368.3				
SME	2.1	–9.0	.	.				
Copper	35.5	41.1	2.2	140.0	15.3	16.5	30.5	36.1
Of which: LME	19.2	16.0	2.1	134.5				
SME	12.4	113.1	.	.				

Note: NYMEX = New York Mercantile Exchange; ICE = IntercontinentalExchange, United Kingdom; TOCOM = Tokyo Commodity Exchange; LME = London Metal Exchange; SME = Shanghai Metal Exchange.

[1] Number of contracts, in millions. [2] Oil: millions of barrels per day; gold: millions of kilograms; aluminium and copper: millions of tonnes. [3] Defined as financial activity in the two largest contracts converted to units of physical production, divided by production.

Sources: Commodity Research Bureau, *The CRB Commodity Yearbook*; Energy Information Agency, *Annual Energy Review*; GFMS; US Geological Survey. Table 1

Crude oil prices and roll returns



[1] In US dollars per barrel. [2] Annual returns from rolling over consecutive three-month futures at maturity in excess of spot price returns.

Sources: Bloomberg; BIS calculations. Graph 2

inflation (Gorton and Rouwenhorst (2004)).[2] Second, these authors also provide evidence that, historically, the return on a diversified basket of long commodity futures has been comparable with the return on other asset classes with similar risk features, such as equities.

Positive roll returns have been important ...

Several authors have emphasised the importance of the so-called roll return from a long position in commodity futures as a component of total returns (Erb and Harvey (2005), Feldman and Till (2006)). Indeed, roll returns are an important explanation for why the average return on commodity futures has exceeded the average return from holding spot commodities (Gorton and Rouwenhorst (2004)). Investors earn a positive roll return if they can roll over a futures contract that is close to expiry into a new contract at a lower price. This occurs when the spot price (to which the price of the original futures contract converges over time) is higher than the price of the new futures contract, ie in a backwardated market.

... but depend on the persistence of backwardation

Roll returns can be considerable. For example, in the crude oil market, the roll yield from purchasing three-month futures was about 14% per annum over 2003–04 (Graph 2). However, roll returns became negative when the price of the futures contract rose above the spot price, ie the market moved into contango, in 2005. Essentially, the profitability of strategies aimed at generating positive roll returns depends on the persistence of the factors that cause markets to backwardate, including low levels of commodity stocks available for short selling and positive returns received by owners from holding the physical commodity (the so-called convenience yield).

Growing presence of investors with shorter-term focus

The presence of investors with a shorter-term focus, such as hedge funds, has increased considerably during the past three years. The number of hedge

[2] It is important to note that these calculations are all in US dollars and therefore the correlation between commodity prices and exchange rate movements is not a consideration. To the extent that commodity prices are in US dollars and other assets in the portfolio under consideration are not, currency hedging may be important for obtaining diversification benefits.

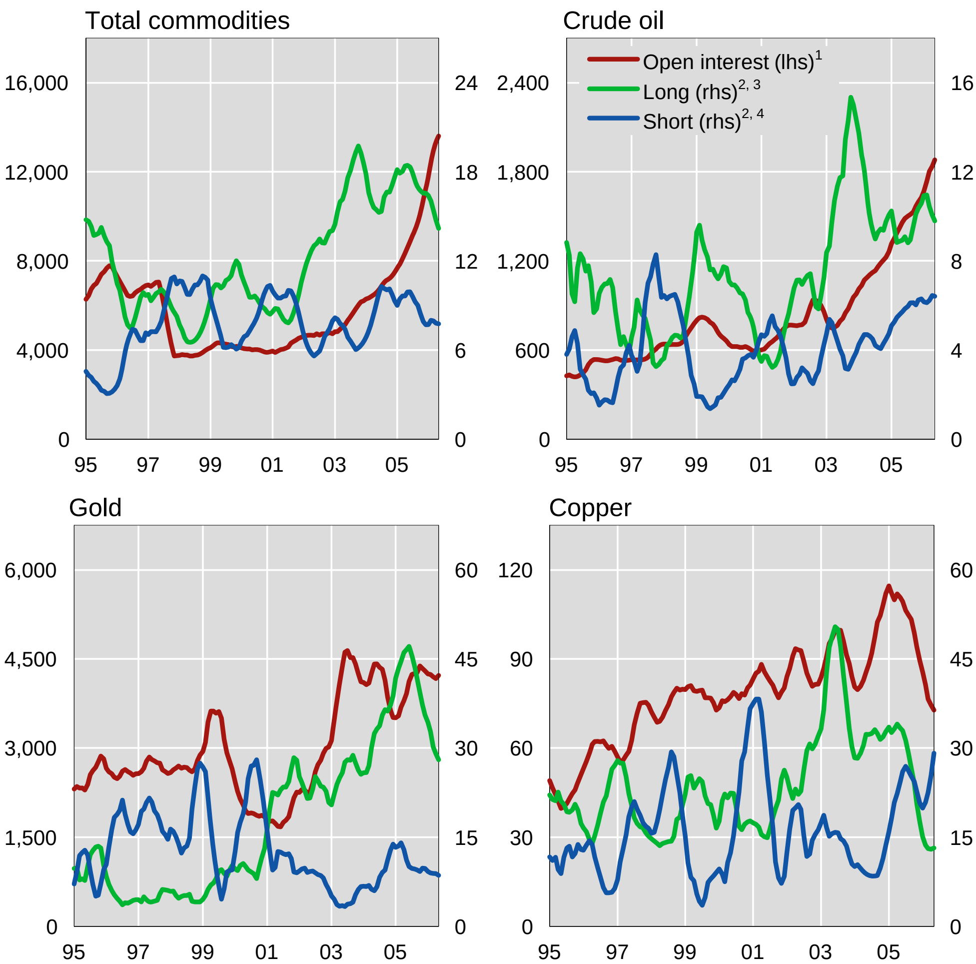
funds active in energy markets has reportedly tripled to more than 500 since the end of 2004, with an estimated $60 billion in assets under management (Fusaro and Vasey (2006)). The $6 billion loss on natural gas derivatives that the hedge fund Amaranth reportedly incurred in September 2006 is a further indication of the size of positions that hedge funds take in commodity markets. Partly as a result of increased demand from financial investors following shorter-term strategies, the number of exchange-traded funds (ETF) for commodities has increased since the first ETF for gold was opened in 2003. A related area of growth is the development of instruments that facilitate the implementation of more complex strategies, including cross-market arbitrage or taking positions on volatility. A specific example is the rapid expansion in structured commodity notes (McNee (2006)).

CFTC data ...

An important source of quantitative information on trading activities in commodity markets is the Commodity Futures Trading Commission (CFTC), which publishes weekly data on the open positions in US futures markets of commercial and non-commercial traders (Graph 3). The non-commercial trader

Total open interest and shares of non-commercial holdings

Futures and options combined; six-month moving averages



[1] Number of contracts, in thousands. [2] Ratio to total open interest, in per cent. [3] Non-commercial long positions. [4] Non-commercial short positions.

Source: CFTC. Graph 3

Activity of managed money traders in selected commodity markets

Market	Number of MMTs holding positions[1]			MMT open interest as % of total open interest[2]		
		1994[3]	2003–4[4]		1994[3]	2003–4[4]
Crude oil	Average	40	80	Long	6.4	14.0
	Maximum	48	100	Short	2.2	6.9
Natural gas	Average	33	66	Long	2.3	11.9
	Maximum	44	81	Short	7.0	15.4

[1] Daily averages and maximums. [2] In futures and options markets. [3] April–September 1994. [4] August 2003–August 2004.

Sources: CFTC (1996); Haigh et al (2005). Table 2

group includes participants who are not primarily using the market for hedging, and encompasses a variety of subgroups. In 2003–04, the non-commercial trading category for both natural gas and oil was dominated by managed money traders (MMTs) (Haigh et al (2005)). This group includes specialised investors such as commodity pool operators and funds advised or operated by commodity trading advisers. Hence, it is likely to capture most financial investors who are operating in centralised commodity markets.

... confirm growing importance of financial investors

The importance of MMTs seems to have grown significantly since 1994. Data available for the crude oil and natural gas markets show that the average number of MMTs trading has roughly doubled and their share of total open interest in each of these markets has increased sharply (Table 2). In addition, assets under management by commodity trading advisers are significant and rose from about \$20 billion in 2002 to about \$75 billion by end-2005 (IMF (2006)).

Share of financial traders varies across markets

The share of non-commercial traders in aggregate has gone up from about 17% in the second half of the 1990s to about 25% in the past three years. This increase is mainly attributable to an upward trend in the share of long positions held by non-commercial investors. While this broad pattern holds across markets, the share of non-commercial positions varies considerably. Since spring 2006, the share of open interest attributed to non-commercial traders has fallen by almost 3 percentage points. This is consistent with a withdrawal of investors during the period of falling commodity prices since May last year, but also with an increase in the hedging activity of commercial producers (JPMorgan Chase (2007)).

Limited information on OTC markets

As regards OTC commodity derivatives markets, the available evidence also supports the notion of a rapidly growing presence of financial investors.

Participants in OTC trading on the ICE

OTC participants' trading (as % of total commissions)	2003	2004	2005
Commercial companies	64.1	56.5	48.8
Banks and financial institutions	31.3	22.4	20.5
Hedge funds, locals and proprietary trading shops	4.6	21.1	30.7

Source: ICE (2006). Table 3

IntercontinentalExchange (ICE) reports that hedge funds, locals and proprietary trading shops accounted for almost one third of trading commissions paid on OTC transactions conducted through ICE in 2005, compared to less than 5% in 2003 (Table 3). However, this increase might in part reflect the higher propensity of institutional investors, in particular hedge funds, to use electronic trading platforms (Davidson (2006)). It may therefore overstate the increase in financial investor participation in commodity markets as a whole.

An empirical examination of investor activity

Empirical approach

To obtain a general sense of the changes in the motivations underlying investment activity, we next estimate the relationship between the activity of financial investors and possible motivating determinants. The results of this simple, illustrative exercise are broadly consistent with the view that the motivations for investing in commodity markets have changed along with the growing presence of financial investors. Given data limitations, this exercise is constrained to using CFTC data on non-commercial open interest in US exchange-traded commodity markets. The dependent variable is defined as the share of net long open interest of non-commercial traders in four somewhat heterogeneous commodity markets that have experienced particularly large price movements since 2002: crude oil, natural gas, gold and copper.[3]

Explanatory variables

To capture the effect of expected returns on the share of non-commercial traders, we include the percentage changes in spot commodity prices and a variable capturing the size of the roll return over the previous 12 months.[4] The standard deviation of monthly percentage changes in three-month futures prices is included to capture any response there may be to volatility in returns. A priori, the effect of such volatility on the position-taking of financial investors is ambiguous. On the one hand, rising volatility may discourage position-taking because it lowers risk-adjusted returns, all else equal, particularly for strategies such as carry trades. On the other hand, volatility is likely to attract more activity if traders are actively taking exposure to it. Another shorter-run return consideration may be the opportunity cost of investing in commodities. To account for this, a world short-term interest rate has also been included. The longer-term demand for commodities arising from their diversification properties is proxied in two ways: by the correlation between percentage changes in commodity prices and a measure of world equity prices over the

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[3] Net long positions of non-commercial traders are frequently used as a variable to capture financial investor activity in commodity markets; see eg IMF (2006) and Micu (2005). By defining the dependent variable as a share, factors that increase net long positions for commercial and non-commercial traders have been controlled for. However, the dependent variable cannot distinguish an increase in non-commercial net long open interest arising from factors that have increased financial activity across all financial markets from an increase arising from a portfolio shift towards commodity markets as a whole, or portfolio shifts between individual commodity markets. These issues serve as qualifications to the interpretation of the estimates.

[4] This variable is defined as the difference between the spot price and the three-month futures price, normalised by the spot price, averaged over the previous 12 months. To the extent that roll returns encourage investor activity, the estimated coefficient on this variable should be positive. All explanatory variables are included with a lag of one month.

previous five years; and by inflation expectations, defined as the difference between nominal and real bonds.

Shorter-term factors seem to have become more important ...

Two broad observations can be made by comparing the results of estimating this model for the period 1998–2001 with those for the period 2002–06 (Table 4). First, shorter-term factors reflecting return considerations appear to have become, on balance, more important over time. Past increases in spot prices have a significant positive effect on the share of non-commercial net long positions across both periods, as expected. Higher roll returns have a more positive effect on the share of non-commercial net long positions in the second period than in the first in the natural gas and oil markets, which have been backwardated for considerable periods since 1998, as well as in the copper market, although the estimated coefficient is not significant.[5] The volatility of futures returns has a negative effect across markets in the second period, which is particularly significant in the copper market. This pattern is consistent with a growing importance of leveraged investors speculating on short-term price trends, as this group is particularly sensitive to short-term price fluctuations.

Regression results[1]

Dependent variable: non-commercial long minus short positions, as a share of total open interest

	Return[2]	Roll[3]	Volatility[4]	Interest[5]	Correlation[6]	Inflation[7]	Adjusted R^2
Expected sign	+	+	–	–	–	+	
1998–2001							
Crude oil	0.04	**–0.45****	**3.30****	**2.88****	–0.01	–2.12	0.67
Natural gas	**0.11****	–0.19	1.15	**–2.47***	**0.53****	**11.17****	0.60
Gold	**1.09****	**18.97***	–1.06	–3.17	**–0.58****	5.19	0.39
Copper	–0.03	**–26.30****	4.10	–4.86	**–2.19****	**24.24****	0.59
2002–06							
Crude oil	**0.11****	**1.35****	**–1.61***	**4.50****	**0.30****	**3.01***	0.42
Natural gas	**0.02***	**0.15***	–0.26	**1.44***	0.06	0.92	0.15
Gold	**0.53***	**–23.10***	–1.75	**–11.77***	0.22	**8.03***	0.41
Copper	0.24	1.14	**–9.56****	**–36.51****	–0.63	1.50	0.81

Note: * indicates significance at the 10% level, ** at the 5% level; bold red indicates expected sign and significance; light red indicates expected sign and non-significance; bold black indicates incorrect sign and significance; light black indicates incorrect sign and non-significance.

[1] The seemingly unrelated regression methodology was used to estimate these results on monthly data in order to allow for contemporaneous correlation in the errors across equations. All variables are lagged once. Other lag structures were tested, but the effectiveness of this strategy was limited by the relatively short sample period. [2] Monthly percentage change in the spot price. [3] Twelve-month moving average of the spot price minus the three-month forward price, divided by the spot price. [4] Twenty-month rolling standard deviation of the monthly percentage change of the three-month futures price. [5] Average of three-month interest rates of Canada, Germany, Japan, Sweden, the United Kingdom and the United States. [6] Correlation between the percentage changes in the spot price and in the Morgan Stanley world equity price index over a rolling period of five years. [7] The difference between nominal and real US 10-year bonds.

Sources: Bloomberg; CFTC; Datastream; Goldman Sachs Research; national data; BIS calculations.

Table 4

[5] The crude oil futures curve has been backwardated around half the time since 1998. Over this period, the natural gas market has been backwardated only 15% of the time, while copper has been backwardated 34% of the time. The futures curve for gold has almost always been in contango due to the large level of above-ground inventories. Since 1975, the gold market has backwardated only four times (in August 1976, May 1983, March 1986 and January 1993).

... although differences across markets appear considerable

The coefficient on the interest rate is more significant across markets in the second period, although with different signs. This supports the view that the size and character of financial investor activity differ considerably across markets. The negative sign for the gold and the copper markets, where the shares of non-commercial positions are three to four times larger than the share for crude oil markets (Graph 3), might indicate that the interest rate variable reflects opportunity costs of financial investors with a shorter time horizon. In energy markets, the positive coefficient might capture a trend increase in net long positions resulting from passive tracking of commodity indices, which tend to place a high weight on energy commodities. For example, oil had an average weight of 27% in the second subperiod in the GSCI index. However, no separate role could be found in the regression for the GSCI commodity weights.

Diversification benefits less significant

The second observation is that the share of non-commercial net long positions appears to have been less influenced by perceived diversification benefits than in the past. In the earlier subperiod, before prices started to accelerate, there is a negative relationship between investor activity and the correlation between returns on commodities and world equities in most cases. In the second subperiod, this relationship is either statistically insignificant, or has a perverse sign. One possible alternative explanation for this outcome is that short-term strategies have been more important than before and dominate the variation in the data. Another possibility is that the correlation variable does not capture the full range of assets which have been relevant for the assessment of diversification benefits in the recent period, although including the correlation between commodity returns and other asset classes such as high-yield credit does not change the result. Commodity investment might also have been motivated by long-term historical correlations that are not apparent in the relatively short span of the second subperiod. The relationship between the share of non-commercial long positions and expected inflation is generally positive, although not always significant, consistent with commodities being purchased as a hedge against future inflation.

Financial investors and market dynamics

Questions

Changes in the scale and character of involvement of financial investors in commodity derivatives markets may have affected the price dynamics of these markets. The first question in this regard is whether the exploitation of perceived profit opportunities by financial investors has fundamentally changed the relationship between prices and the physical characteristics of commodity markets. The second issue is whether the broadening of the investor base has led to significant market deepening and hence affected features such as short-term price fluctuations.

The relationship with physical commodity markets

Investor activity and commodity prices

Intuitively, one might expect large inflows of funds into commodity markets to cause prices to rise sharply, possibly to higher levels than are justified by economic fundamentals. The prima facie evidence seems to support this view, as financial activity has broadly increased in parallel with prices during the past four years. However, the results of empirical work on the impact of the growing presence of financial investors on commodity prices are less clear-cut. Several recent studies, which explore the relationship between investor activity and commodity prices, indicate that price changes have led to changes in investor interest rather than the other way around (Haigh et al (2005), IMF (2006)).

Physical characteristics as a benchmark

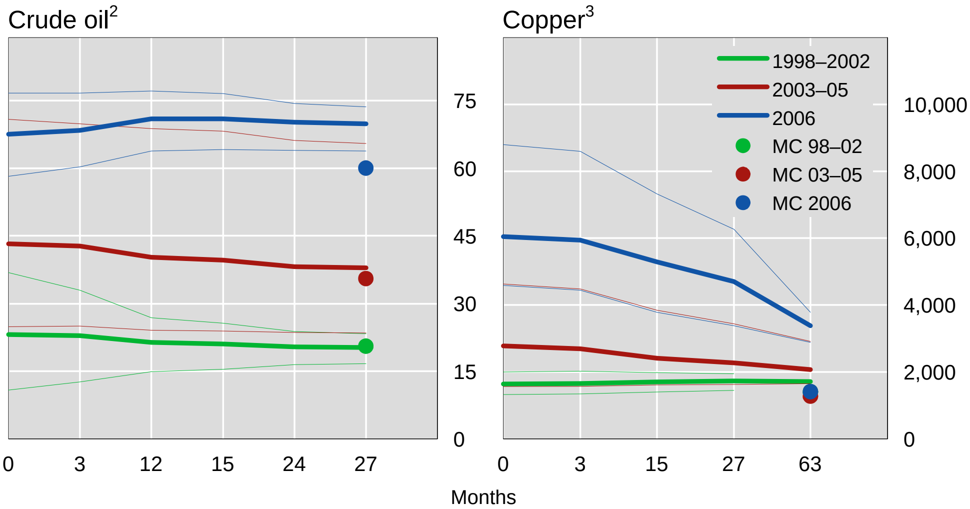
This section uses the physical characteristics of specific commodities as a rough benchmark for assessing whether the increased presence of financial investors has altered price dynamics. Constraints on supply and storability affect the prices of commodity derivatives. In the longer run, production can be changed and the elasticity of commodity supply depends on the marginal costs of production. In the short run, supply from production is relatively inelastic and depends more on above-ground stocks. With the exception of gold, above-ground commodity stocks are small relative to demand. For example, it is usual for four to six weeks of demand to be held in inventories for base metals. For gold, in contrast, stocks either available for production or for lease represent close to 45 years' worth of demand, depending on how this is measured (O'Connell (2005)).

Marginal costs of production have been a strong anchor for long-dated futures prices ...

In efficient markets, the expected marginal costs of commodity production should act as an anchor for longer-run futures prices. Consistent with this, the long ends of oil and copper futures curves have overall tended to fluctuate much less than spot and short-dated futures prices (Graph 4). The tenors that are affected by this "anchoring" may vary, depending on the time needed to adjust production. For instance, from 1998 to 2002, a period of ample spare capacity, marginal costs were steady and production could be expanded at relatively short notice. Indeed, futures prices at tenors from about one year were quite closely aligned with estimates of marginal costs of production in

Prices and marginal costs[1]

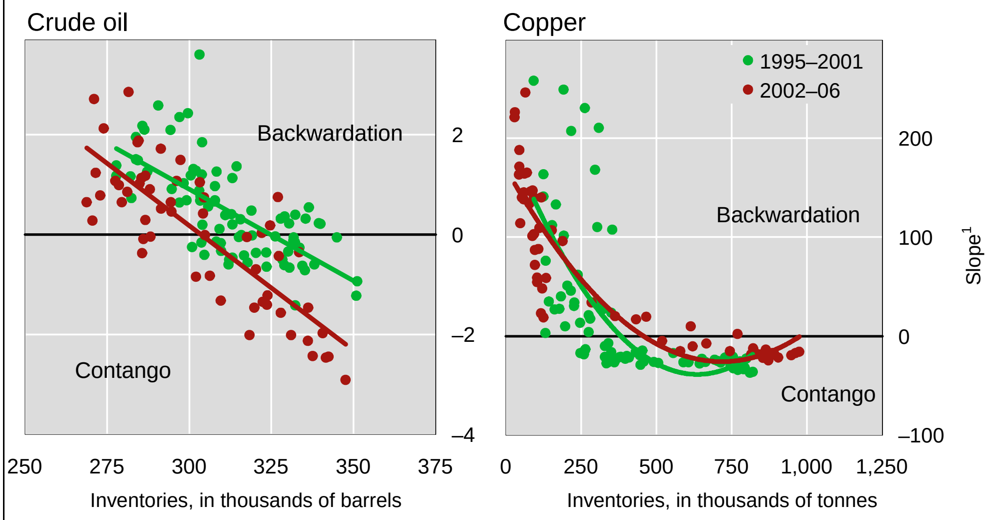
Daily data, 1998–2006



[1] The bold lines have been constructed by averaging the prices at each tenor within each time period. The thin lines represent the futures curves associated with the minimum and maximum spot prices within each time period. [2] In US dollars per barrel. [3] In US dollars per tonne.

Sources: Bloomberg; Goldman Sachs Research; JPMorgan Chase; BIS calculations. Graph 4

Inventories and the slope of the futures curve



[1] Spot price minus three-month futures price; for oil, in US dollars per barrel; for copper, in US dollars per tonne.

Sources: Bloomberg; London Metal Exchange; BIS calculations. Graph 5

both oil and copper markets over this period.

... but seem to have lost power since 2003 ...

Since 2003, however, long-dated futures prices have increasingly diverged from estimates of current marginal costs. In 2006, prices for two-year oil futures were on average about 20% higher than the measure of marginal costs shown in Graph 4. In the case of copper, the deviation was much larger. Several factors related to economic fundamentals could cause such a deviation. For example, a sharp increase in expected marginal costs owing to buoyant demand growth and uncertainty about the costs of further expansion of production in the face of capacity constraints may have been a factor in the oil market. Moreover, the need to explore and develop new sources has probably lengthened the time required to extend production.

... to a degree which is difficult to reconcile with fundamentals

In addition, futures prices are likely to embody risk premia, not least because long-dated futures markets are typically relatively thinly traded. Reluctance by producers to forgo upside opportunities through hedging in an environment of rising prices might have further reduced liquidity. In contrast, there is some tentative evidence that the size of the risk premium in oil futures markets is positively related to the share of net non-commercial long positions in the oil market, controlling for other factors (Micu (2005)). Notwithstanding all these factors, it still appears difficult to reconcile the increases in futures prices until mid-2006 with economic fundamentals, especially in the case of copper.

Inventory-slope relationship has remained intact ...

A second physical anchor is inventories, which link current and future supply and consequently connect the spot price and expected spot prices in the future (Gorton and Rouwenhorst (2004)). It is not clear that growing investor activity can have a systematic direct effect on inventory decisions: the convenience that producers derive from holding stock importantly depends on factors related to real activity such as production smoothing. Indeed, the strong historical relationship between the slope of the futures curve for non-gold commodities and the level of physical inventories has remained intact (Graph 5).

It is more likely that financial investors could indirectly affect inventory decisions through futures prices. To the extent that taking long positions in futures markets increases futures prices, the value of holding inventories for future delivery increases. The effect on the slope of the yield curve remains open, depending on how spot prices respond to possible inventory decisions.

... but indirect effect on inventory decisions possible

Market depth

The second question is whether the increase in the size and diversity of financial investors has increased market depth. Greater market depth would imply that transactions of a given size cause smaller fluctuations and, other things equal, that short-term price volatility should decline. The prima facie evidence on changes in commodity price volatility is mixed. Price volatility has declined in the oil market, especially in the shorter maturities of futures contracts where trading is particularly active (Graph 6). In contrast, it has increased in the copper market.[6]

Financial investors and market depth

Another approach is to look at the interaction of the trading behaviour of commercial and non-commercial traders. Non-commercial traders will add to market depth if they contribute to a two-sided market. This is the case if they act as counterparties to commercial traders' hedging transactions or if they take positions offsetting other financial investors.

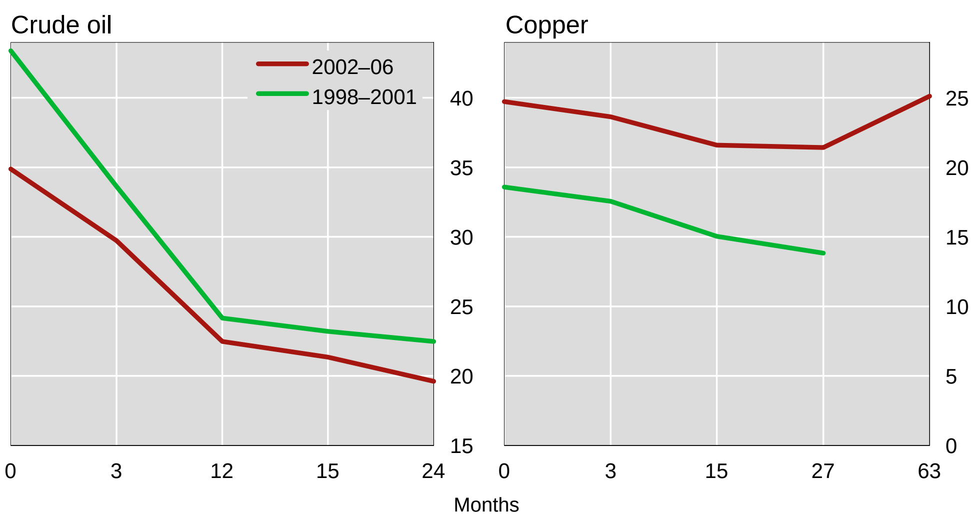
Interaction of commercial and non-commercial traders ...

The pattern of changes in the open positions of commercial and non-commercial traders supports the view that financial investors have, overall,

... seems to have reduced volatility

Volatility of commodity futures prices

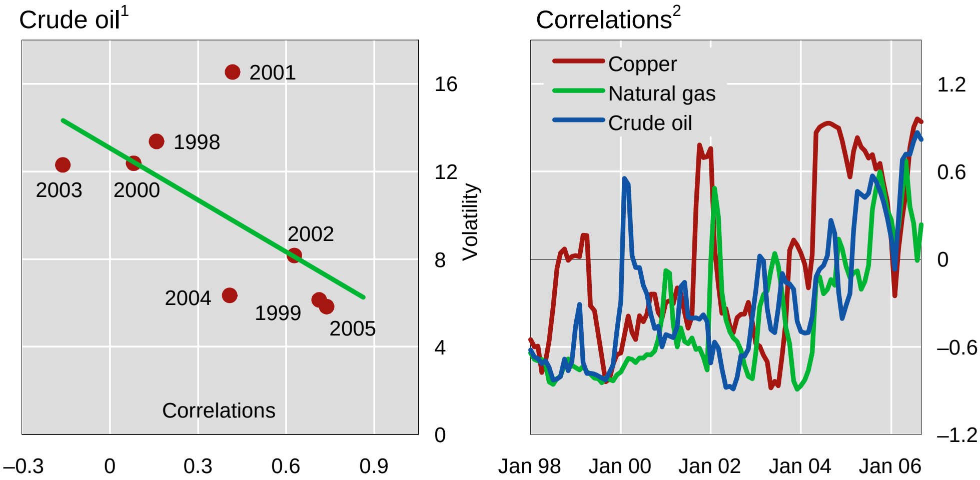
Annualised standard deviation of daily changes in the log of prices



Sources: Bloomberg; BIS calculations. Graph 6

[6] This highlights one of the limitations of the econometric work done earlier, insofar as changes in investor activity cause changes in variables, such as volatility, that we have included as explanatory variables.

Volatility and correlation



[1] The x-axis is the correlation between the changes in non-commercial long positions and commercial short positions: the y-axis is the standard deviation of the spot oil prices. [2] Correlation between the changes in non-commercial long positions and non-commercial short positions; 12-month moving average.

Sources: Bloomberg; CFTC; BIS calculations. Graph 7

contributed to deeper markets.[7] First, a higher correlation between changes in non-commercial long and commercial short positions has been associated with lower volatility in oil markets (Graph 7, left-hand panel). However, the correlation has not significantly increased since 2002, suggesting that a growing presence of financial investors may have accommodated increased hedging needs, but not fundamentally altered the character of the market.

Non-commercial traders increasingly active on both sides of commodity markets

Second, there is also evidence that non-commercial traders have, as a group, increasingly taken positions on both sides of commodity markets. Prior to 2002, changes in long and short positions of non-commercial traders were highly negatively correlated for copper, oil and natural gas: an increase in long positions typically went hand in hand with a reduction of short positions and vice versa. There is also some evidence that MMTs tended to act on the same side of the market at similar times in the past (CFTC (1996)). In the past few years, however, the correlation between changes in long and short positions of non-commercial traders has increased and become positive (Graph 7, right-hand panel). Evidence that non-commercial players are increasingly trading between each other is also provided by the growing share of spread positions, which arise when a trader takes long and short positions in the same commodity at different tenors of the futures curve.

Growing similarities with financial markets

The emergence of trading among financial investors in commodity markets on a substantial scale suggests that the determinants of market liquidity may become more similar to those in traditional financial markets. These determinants include the amount of risk capital that financial investors allocate to commodities trading and the heterogeneity of opinions of market participants. One key risk in both regards is a high concentration of trading

[7] In order to gauge the position-taking of the investor groups on both sides of the market, we consider correlations of long and short positions separately (ie we do not calculate net long or short positions).

activity. The demise of Amaranth, which led to a sharp deterioration in liquidity conditions in those tenors of the natural gas futures market where the firm held extensive positions, provides a clear indication of these challenges.

Conclusion

The presence of financial investors in commodity markets has increased considerably during the past four years or so. While it is difficult to be precise about the exact magnitude and composition of inflows, there is much evidence that the investor base, and with it the range of instruments and strategies employed in commodity trading, has broadened substantially. It is not clear to what extent these changes reflect structural shifts in investor behaviour or a temporary boom supported by a "search for yield". In any case, a full reversal of the trend towards a greater role of financial investors appears unlikely against the backdrop of greater investor sophistication and a broadening range of commodity-related financial instruments.

Commodity markets have become more like financial markets in some respects. Financial investors are increasingly active on both sides of trades, creating a kind of financial trading sphere. Yet the characteristics of physical markets, such as inventory levels and the marginal cost of production, are still important. A lack of liquidity especially in the long tenors of commodity derivatives markets and physical limits to short selling in the spot market may at times significantly affect market dynamics. These effects require further investigation.

While the increase in investor activity can be expected to bring benefits in terms of market efficiency, the ongoing "financialisation" of commodity markets raises issues similar to those in other financial markets. Among these is the question of how to ensure robust market liquidity.

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Economic derivatives[1]

Economic derivatives allow traders to take direct positions on the outcomes of macroeconomic data releases. In contrast to survey-based measures, the prices of economic derivatives provide information on the entire probability distribution underlying these expectations, not just point estimates. Measures for uncertainty derived from such distributions offer valuable information on how uncertainty about the economy evolves and affects financial markets.

JEL classification: E44, G13.

Economic derivatives are financial contracts that allow market participants to take positions on macroeconomic data releases. They are different from the "macro securities" proposed by Shiller (1993), which are meant to insure households and corporations against changes in macroeconomic conditions that may affect their livelihood. By contrast, the economic derivatives which are the subject of this feature focus on short-term data surprises.

Macroeconomic data announcements play an important role in price discovery in financial markets, as has been documented by a large body of literature.[2] They are usually scheduled regularly, with a precise date and timing that are known well in advance. The US employment report, for example, is generally released on the first Friday of the month at 08:30 Eastern Time. The importance of data announcements is underscored by the fact that volatility tends to be markedly higher on release days than on other days. For example, the standard deviation of the yield changes of 10-year US Treasuries is almost twice as high on announcement Fridays (Table 1). Higher volatility than usual is also observed for a whole range of other financial instruments, both in the United States and in the euro area.[3] Prices move primarily in

1 The views expressed in this article are those of the authors and do not necessarily reflect those of the BIS. The authors would like to thank Emir Emiray for help with the data and graphs.

2 Andersen et al (2005) estimate the effects of US macroeconomic announcements on a large variety of financial prices. See also Fleming and Remolona (1997, 1999) and Balduzzi et al (2001) for US Treasuries, Andersson et al (2006) for euro area bonds and Brooke et al (1999) for the sterling market.

3 Many financial prices in the euro area react more strongly to US announcements than to national or area-wide releases; see Goldberg and Leonard (2003) and Andersson et al (2006).

Non-farm payroll announcements and asset price volatility			
Instrument	Volatility[1]		p-value[3]
	Announcement Fridays	Other Fridays[2]	
Federal funds futures	1.8 bp	0.9 bp	0.00
Ten-year Treasury note	9.9 bp	5.2 bp	0.00
Ten-year bund	5.3 bp	3.7 bp	0.00
S&P 500	0.84%	0.85%	0.58
EURO STOXX	1.26%	1.16%	0.17
USD/EUR exchange rate	0.87%	0.59%	0.00

[1] Standard deviation of daily price changes (interest rates) or returns (equities, exchange rate), September 2002–December 2006. [2] Last working days of a month have been dropped since they may be affected by window-dressing. [3] Likelihood-ratio test for equal volatility on announcement and non-announcement Fridays.

Table 1

intervals of just a few minutes around the announcements, reflecting market participants' forceful and instantaneous reaction to the new information.[4]

This special feature describes economic derivatives and the market where they are traded. It investigates the motives for trading these contracts and explores the use of their prices for measuring market expectations. For illustrative purposes, it concentrates on US non-farm payrolls (NFPs), which are released by the Bureau of Labor Statistics as part of its monthly employment report. NFPs rank among the most influential macroeconomic data releases worldwide.

What are economic derivatives?

Position-taking on macroeconomic data releases

Economic derivatives are financial instruments which allow traders to take positions directly on the outcome of macroeconomic data releases. For instance, they may trade a combination of digital options to bet that NFPs would increase by between 150,000 and 200,000.[5] Alternatively, they might purchase plain vanilla options if they believed that NFPs would be much higher than the market consensus, but did not want to commit to a specific range.[6] Unlike conventional financial instruments, economic derivatives separate the surprise component of announcements (the difference between the outcome and the prior expectation by market participants) from the channel through which such news is transmitted to asset prices.

Gravelle and Moessner (2001) find that Canadian interest rates also react more to US macroeconomic announcements than to Canadian ones.

4 See, for instance, Fleming and Remolona (1999) and Gadanecz (2003).

5 Digital call options pay out a fixed amount if the data outcome is higher than the strike price and nothing otherwise. The range of 150,000 to 200,000, for example, can be traded by purchasing call options with a strike price of 150,000 and selling an equal amount of call options with a strike of 200,000. See Kolb (2003) for a discussion of digital options.

6 Plain vanilla call options pay out an increasing amount as the outcome of the release falls further above the strike price. The payouts from vanilla options are capped based on the highest and lowest strikes in the auction.

Traded in auctions …

Economic derivatives were introduced in October 2002 by Deutsche Bank and Goldman Sachs, first on US NFPs and subsequently also on other releases such as the ISM manufacturing index, US initial jobless claims and retail sales, the euro area harmonised index of consumer prices, and US GDP and international trade balance data. They were initially traded over the counter in Dutch auctions (also known as uniform price auctions), with Goldman Sachs acting as the counterparty. Auctions were subsequently moved to the Chicago Mercantile Exchange (CME) in September 2005, where the clearing house offers the usual services and central counterparty guarantees that are available on an organised exchange.

For each data release, one or more auctions are held, usually on the day of and during the week preceding the announcement. Customers can submit sell and buy offers at a limit price which depends in part on their assessment of the volatility of the underlying macroeconomic data or, in other words, on their estimate that the option will expire in the money. Indicative prices and filled orders are given during the auction, while the final pricing and filled orders are determined at the end of the auction.[7]

… economic derivatives are traded by a number of market participants

The main participants in the market are macro and relative value hedge funds, large banks, dealers, proprietary traders and portfolio managers. They follow strategies similar to those in other markets, including directional and volatility-based trading and relative value strategies. Trading of options can be done individually, or in combinations such as spreads, straddles, strangles and risk reversals (Beber and Brandt (2006)).

Limited market size

The market for economic derivatives still remains very small relative to conventional futures and options traded on exchanges. In 2006, the nominal value of all auctions on one month's NFP release was equivalent to less than 5% of the value-at-risk at the end of that month for the 10-year US Treasury note futures contract of the Chicago Board of Trade (CBOT) (Graph 1), and on average 5,700 times less than the end-month open interest outstanding on the same contract.[8] Trading has been strongest in NFP derivatives, with an average nominal value of approximately $9 million per auction (Graph 1).

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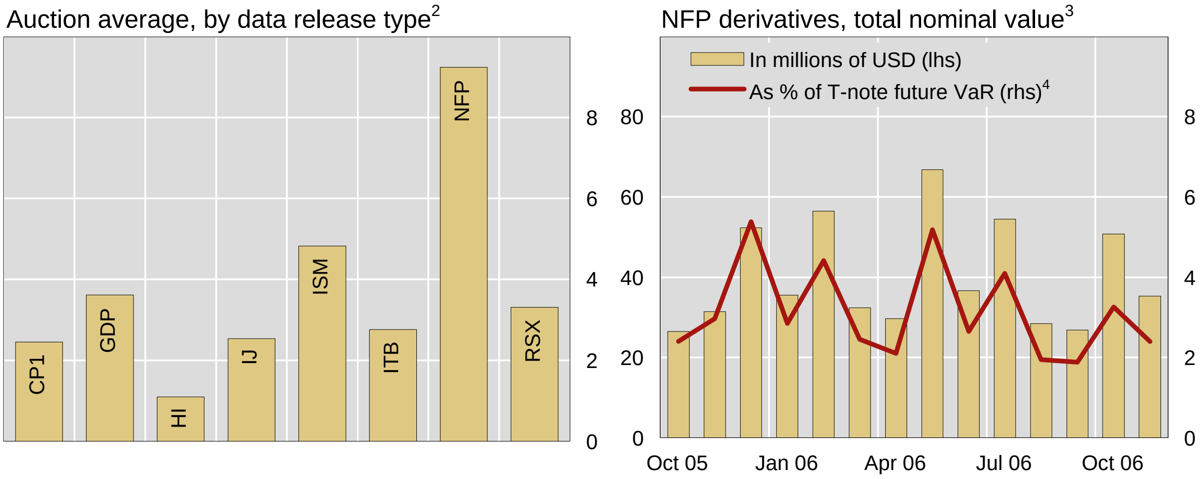
[7] Contrary to traditional order-matching systems, where sell orders are matched with buy orders for the same contract, digital options are traded using a pari-mutuel system similar to the one common in sports betting. Under such a system, the premium collected from the holders of out-of-the-money options is paid out to the holders of in-the-money options. This allows prices to be formed even in the absence of matching buy and sell orders, considering the overall inventory of buy and sell orders as one pool of liquidity.

[8] The nominal value of economic derivatives is obtained by picking the highest total payments of all possible announcement outcomes. Since this amounts to the largest gains or losses, it is more appropriate to compare this measure to large gains or losses in conventional contracts rather than to notional amounts.

Size of the market for economic derivatives

Nominal values,[1] in millions of US dollars



[1] Maximum gain/loss on filled orders. [2] All data releases are for the United States, except for euro area HICP. All releases are monthly, except where otherwise indicated. CP1: core CPI; GDP: gross domestic product (quarterly); HI: euro area harmonised index of consumer prices excluding tobacco (average of auctions one month and two months before the release); IJ: initial jobless claims (weekly, average of all weeks of the month); ISM: ISM manufacturing PMI; ITB: international trade balance; NFP: non-farm payrolls; RSX: retail sales excluding autos. [3] Sum of all auctions on the month's NFP release. [4] Value-at-risk (VaR) proxy calculated as the end-month open interest of the CBOT's 10-year US Treasury note futures contract multiplied by the fifth percentile over the period October 2005–October 2006 of the daily returns based on the underlying price.

Sources: CME; BIS calculations. Graph 1

Motives for trading NFP announcements

There are several motives for trading economic derivatives. Some market participants may use them to express a view on the outcome of a specific data release while others may want to hedge against the impact of adverse data surprises on their portfolios.

Attractive for speculating on announcements …

Speculating on the outcomes of data releases is perhaps the most common motive for trading economic derivatives. Economists working in financial markets spend substantial resources on trying to predict such announcements, so it seems natural that they would like to trade on these predictions. That said, there are alternatives to economic derivatives for taking positions on data releases. Using conventional financial instruments to trade announcement risk may enable traders to access a deeper pool of liquidity than the one available in the market for economic derivatives. On the downside, strategies using other instruments may run into difficulties if payoffs react in a different way than predicted to the outcome of an announcement. Such risk is often referred to as "basis risk".

… due to absence of basis risk …

A measure of the basis risk involved in taking positions on NFPs can be obtained by regressing the return of an instrument on the surprise component of the announcement. For economic derivatives, the basis risk is zero by construction. For other contracts, the amount of basis risk depends on the presence of a stable (or at least predictable) relationship between their prices and the surprise component of announcements, as well as on the absence of price changes because of factors other than NFPs.

… in contrast to conventional financial contracts

Estimates of the basis risk for a broad variety of financial instruments collected in Table 2 (first regression) show that basis risk is substantial. Although the coefficients on announcement surprises are often statistically highly significant, the fit of the equations tends to be relatively poor. For example, less than one half of the changes in short-term swap rates on announcement Fridays can be attributed to surprises in NFPs. For other contracts, the proportion of returns explained by the release is even lower.

There are several reasons for the poor performance of financial contracts for taking positions on NFPs. First, market participants might also react to the other variables included in the employment report, not just NFPs. However, the fit of the regressions increases only slightly when other indicators released at the same time are included (second regression in Table 2).

A second reason might be the focus on daily returns. However, while moving to a shorter time interval would reduce the likelihood of events other than the announcement affecting prices, it would not eliminate basis risk completely. At least, this is the result of several studies estimating announcement effects using very high frequency data.[9]

Third, and most importantly, basis risk may reflect the fact that, unlike in the case of economic derivatives, the returns on financial derivatives depend both on the announcement surprises and on the sensitivity of asset prices to macroeconomic data, which could vary over time. For example, the Federal

Estimates of basis risk

Indicator	First regression[1]		Second regression[2]					
	Payrolls	R^2	Payrolls	Unemp rate	Man emp	Hourly earnings	Weekly hours	R^2
Fed funds second contract	0.104***	**0.18**	0.127***	–0.044**	–0.022	0.007	0.004	**0.29**
S&P 500	1.928	**0.01**	2.072	0.481	–1.130	–0.756	–0.240	**0.05**
EUR/USD exchange rate	–4.810***	**0.22**	–6.215***	1.531**	0.186	–2.405***	–0.235	**0.41**
US 10-year note	0.708***	**0.38**	0.795***	0.004	–0.058	0.151**	0.001	**0.44**
Swap rate 1-year	0.623***	**0.44**	0.755***	–0.081	–0.088**	0.188***	0.026	**0.58**
Swap rate 2-year	0.876***	**0.44**	1.004***	–0.077	–0.056	0.209***	0.010	**0.52**
Swap rate 4-year	0.937***	**0.42**	1.074***	–0.042	–0.050*	0.257***	0.005	**0.51**
Swap rate 10-year	0.748***	**0.37**	0.851***	–0.004	–0.065	0.178**	0.003	**0.44**

Note: ***, ** and * denote significance at the 1%, 5% and 10% levels, respectively.

[1] LHS variable is the change in the indicator from the day before the release of non-farm payroll data for all variables except S&P 500 and EUR/USD; for these indicators, the growth rate is used. RHS variable: actual minus average Bloomberg analyst forecasts for changes in non-farm payrolls, in millions; constant not reported. Sample ranges from January 2002 to December 2006. [2] LHS variable: as in the first regression. RHS variable: difference between actual and Bloomberg analyst forecasts for: changes in non-farm payrolls, in millions; unemployment rate, in per cent; number of employees on US non-farm payrolls, manufacturing industry, month-on-month net change seasonally adjusted (SA), in hundreds of thousands; US average hourly earnings, private non-farm payrolls, in nominal US dollars, month-on-month change SA, in per cent; US average weekly hours, private non-farm payrolls, total private services SA; constant not reported.

Sources: Bloomberg; BIS calculations. Table 2

[9] Balduzzi at al (2001) report an R^2 of 0.56 for the regression of the 35-minute (five minutes before and 30 minutes after the announcement) returns on 30-year Treasury bonds on NFP surprises. For a similar regression using five-minute returns, Andersen et al (2005) report a value of R^2 of 0.36.

Reserve might react much more strongly to higher than expected payrolls when inflation is high than it would in times of low inflationary pressures. This would be reflected in very different reactions of the prices of federal funds futures. Similar changes in the slope of market reactions to announcements were observed in 2003, when the impact on US Treasury yields of NFP surprises steepened substantially (BIS (2004)). In extreme cases, even the sign of the response may vary over time.[10]

Limited attractiveness for hedging portfolio risk ...

The attractiveness of economic derivatives for speculating on data releases does not necessarily imply that they are the most appropriate instrument for hedging the announcement risk of a portfolio. This is because hedgers are much less likely than speculators to be interested in unbundling data surprises from the sensitivity of asset prices to macroeconomic data since they presumably care primarily about how the value of their portfolio is affected by releases rather than about announcements per se.

… might constrain growth potential of the market

A limited attractiveness of economic derivatives to hedgers could constrain the growth potential of the market. Hedgers might be willing to lose money on average in a market, due to being less well informed than speculators, in order to obtain protection for other positions in their portfolio. Theoretical and empirical research in finance suggests that a certain amount of uninformed trading is often necessary to sustain a market.[11] In the absence of significant demand by uninformed agents, trading might also be sustained by differences of opinion between highly sophisticated traders.[12] Such differences of opinion may arise from differences in information (although macroeconomic data tend to be publicly available), but might also result from traders having different ways of processing these data.

In the case of economic derivatives, it is not clear which, if any, of the two explanations – the one based on uninformed trading or the one based on differences in opinion – provides a better characterisation of the motives underlying trading in that market. The limited attractiveness of these instruments for hedgers would suggest a restricted role for informational advantages in the sense that some actors are better informed than others. Similarly, differences of opinion would suggest that trading volumes tend to be high when there is a lot of disagreement, which is at odds with the negative correlation between volumes and the dispersion of analyst expectations in the data.[13] However, this might also be due to the short time span of the data.

[10] See Furfine (2001) for an example concerning US Treasury bonds and Andersen et al (2005) for evidence from the stock market.

[11] Another source could be trading by uninformed, but overconfident, participants (Wolfers and Zitzewitz (2006a)).

[12] See Harris and Raviv (1993).

[13] The correlation coefficient of the standard deviation of responses to the Bloomberg survey (see below for a discussion) and volumes is –0.24.

Economic derivatives as indicators of market expectations

Indicators for traders' economic outlook

The market for economic derivatives is of interest to a much broader audience than the limited group of immediate market participants. This is because the prices of these instruments provide useful information about traders' views of the economy. In addition to obtaining market-based mean expectations of data outturns, under some assumptions it is possible to compute the probability distribution underlying expectations. Such information is not available from analyst surveys, which report the *dispersion* of economists' views about data releases but not the uncertainty surrounding those estimates.

Timelier and less prone to misrepresentation than surveys …

In principle, there are two main reasons to believe that the information contained in the prices of economic derivatives is superior to that from surveys. First, it is timelier. Auctions are generally conducted on the day of the release or on the previous day, which contrasts with a lag of one week or longer in the case of surveys. Second, trading economic derivatives involves real money and is therefore much less likely to be affected by economists misrepresenting their views in order to position themselves relative to consensus forecasts.

… but potentially distorted by risk and liquidity premia

On the other hand, market-based forecasts might be distorted by risk premia or by the limited liquidity of the market. Both premia could introduce a wedge between implied expectations and true expectations of market participants, which would distort any inferences of market participants' expectations from prices. Evidence on the existence of such premia may be obtained by running tests for forecast accuracy. These indicate that we cannot fully rule out their presence since the prices of economic derivatives appear to overpredict outturns on average (Box 1).[14] A similar result has been obtained by Gürkaynak and Wolfers (2006) for a shorter sample period but a broader set of contracts. However, the fact that the mean forecast error based on surveys is also non-zero and close to that of the auction-implied mean forecast error (Table 3) indicates that the overprediction may also be due to overoptimistic expectations or to the particular sample period used in the analysis.

Limited importance of risk premia in practice

In practice, the differences between forecasts implied by auctions of economic derivatives and survey-based mean expectations appear to be relatively small. Both indicators are comparable in terms of their mean forecast errors and correlation with actual NFP data outturns (Table 3). This suggests that neither the potential staleness of survey data, nor any strategic misrepresentation of those data, nor the presence of risk or liquidity premia in the market-based indicators is a particularly significant issue.[15]

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[14] Risk premia can be ruled out if forecasts are unbiased and efficient in the sense that the forecast error is not correlated with other information available to the forecaster. However, the reverse need not apply. Biased or inefficient forecasts could reflect irrational expectations as well as risk premia.

[15] In any case, research by Wolfers and Zitzewitz (2006b) suggests that the distortions arising from risk premia are likely to be small and that the prices of economic derivatives therefore efficiently aggregate market participants' beliefs, at least approximately.

Box 1: Unbiasedness and efficiency

This box examines the unbiasedness and efficiency of market-based NFP forecasts. Forecasts are unbiased if the mean of their forecast errors is zero, ie if the forecast errors are zero on average. Forecasts are efficient if forecast errors cannot be systematically explained. They are efficient in a "weak" sense if forecast errors are uncorrelated with past forecast errors.[①]

A standard test of unbiasedness consists in regressing actual data outturns, d_t, on the market forecasts, d_t^e (Joyce and Read (1999)):

$$d_t = a + bd_t^e + u_t$$

If the market forecasts are unbiased, then we expect that a = 0 and b = 1 and that the residuals are serially uncorrelated. Table A shows that there is no evidence for serial correlation, but that the hypothesis of (a = 0, b = 1) can be rejected at the 5% level, though not at the 1% level. It can therefore not be fully ruled out that market-based NFP forecasts show a systematic bias, perhaps reflecting a risk premium.

Test for unbiasedness of market-based NFP forecasts[1]

a	–17.1
b	0.88
R^2	0.51
Durbin-Watson statistic	2.15
LM test for serial correlation of residuals[2]	1.16 [0.35] [3]
Wald test $\chi^2(2)$: $(a,b) = (0,1)$	7.15 [0.03] [3]

[1] Changes in non-farm payrolls, in thousands; NFP data for September 2002 to September 2006. [2] Breusch-Godfrey LM test with 12 lags, F-statistic. [3] p-values in brackets.

Sources: Bloomberg; Goldman Sachs; BIS calculations. Table A

One test of weak efficiency consists in testing directly whether the forecast errors exhibit no first-order autocorrelation, ie testing whether the coefficient b in the equation

$$d_t - d_t^e = a + b\,(d_{t-1} - d_{t-1}^e) + u_t$$

is zero. As Table B shows, b is not significantly different from zero. Another test of weak efficiency consists in testing whether past actual values have no explanatory power for the forecast errors, ie whether the coefficients on lagged data in the following equation are all equal to zero (Joyce and Read (1999)):

$$d_t - d_t^e = a + \Sigma^{12}_{i=1} b_i d_{t-i} + u_t$$

As also shown in Table B, we cannot reject the hypothesis that the coefficients b_i are all equal to zero. These results suggest that market-based forecasts of NFP outturns are weakly efficient.

Test for efficiency of market-based NFP forecasts[1]

Test for absence of autocorrelation of forecast errors	
b	–0.09 [0.53] [2]
R^2	0.01
Durbin-Watson statistic	2.02
Test for weak efficiency	
R^2	0.24
Durbin-Watson statistic	1.93
LM test for serial correlation of residuals[3]	0.99 [0.51] [2]
Wald test $\chi^2(12)$: $b_i = 0$, i = 1 to 12	7.63 [0.81] [2]

[1] Changes in non-farm payrolls, in thousands; NFP data for September 2002 to September 2006. [2] p-values in brackets. [3] Breusch-Godfrey LM test with 12 lags, F-statistic.

Sources: Bloomberg; Goldman Sachs; BIS calculations. Table B

① They are efficient in a "strong" sense if forecast errors are uncorrelated with any information available at the time the forecasts are made.

Comparison of expectations with actual NFP data outturns

	Mean[1]	Correlation with actual	Mean surprise[1]	Standard deviation of surprises[1]
Auction-implied	124.1	0.71	–31.8	88.7
Bloomberg survey	123.5	0.70	–31.2	90.6
Actual	92.3	.	.	.

[1] Changes in non-farm payrolls, in thousands; NFP data for September 2002 to September 2006.

Sources: Bloomberg; Goldman Sachs; BIS calculations. Table 3

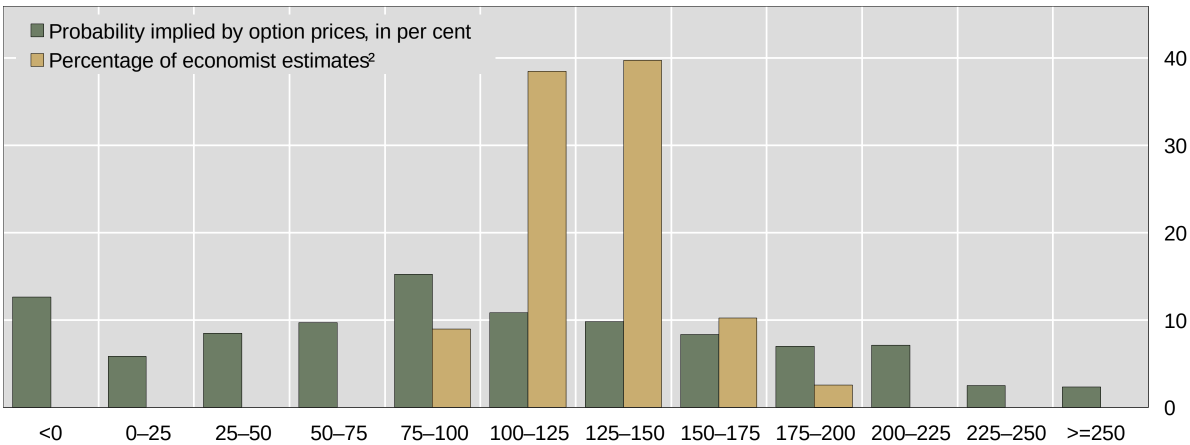
Implied probabilities from economic derivatives …

Given the small difference between the point estimates derived from the prices of economic derivatives and surveys, the main gain from looking at the former arises if one is interested in their uncertainty and the distribution underlying these forecasts. The prices of options with different strike prices can be used to compute implied uncertainty measures and probability distributions of data outturns (see Kolb (2003)).

An example of a probability distribution implied by prices of digital options on NFPs and of the responses of individual analysts to the Bloomberg survey is shown in Graph 2. This distribution is derived at the strike prices with the approximation that the discount factor equals one, which seems reasonable given the proximity of the auction date and the release date. For comparison, Graph 2 also shows the percentage of economists surveyed by Bloomberg forecasting data outturns in the corresponding intervals. Interestingly, the so-called risk neutral probability distribution implied by economic options is wider than the histogram of economists' forecasts. This is to be expected since the market-based probability distribution captures the whole range of market participants' beliefs, including probabilities assigned to tail events, whereas the surveys only capture the central expectations of those surveyed, and contain no information about their expectations of tail events.

Probabilities implied by option prices and economists' forecasts

US non-farm payrolls, August 2006[1]



[1] Bureau of Labor Statistics data releases; changes, in thousands (x-axis). [2] Based on Bloomberg survey.

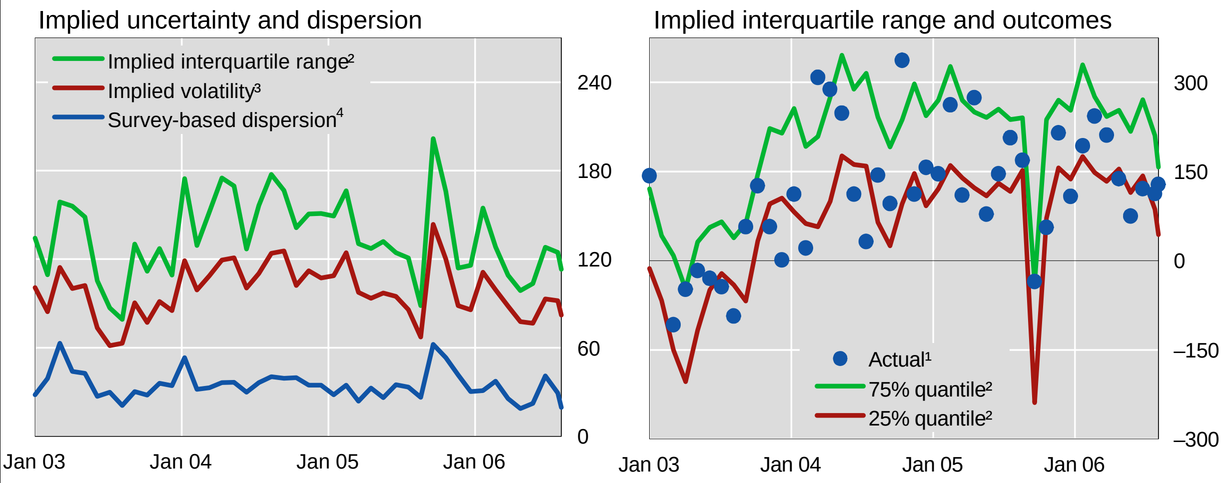
Sources: Bloomberg; Goldman Sachs; BIS calculations. Graph 2

Implied probability distributions such as the one shown in Graph 2 provide a wealth of information on the market's view prior to a particular release, but they are difficult to track over time. Measures that show the evolution of uncertainty include implied volatility and the implied interquartile range[16] (Graph 3, left-hand panel). Graph 3 (right-hand panel) shows the actual NFP data outturns in relation to the implied interquartile range, for the auction closest in time to the data release each month. Of the data outturns, about 50% fell within the implied interquartile range, as would be expected if the market-based uncertainty measure was an accurate measure. This suggests that economic derivatives provided useful information on the market's uncertainty about NFP data outturns, in line with the findings in Gürkaynak and Wolfers (2006). Survey-based dispersion measures, by contrast, tend to provide only a very noisy measure of uncertainty, as is reflected in the low correlation between the auction-based interquartile range and the survey-based dispersion measure of 0.68. In the past, forecast dispersion has often been used as a proxy for uncertainty (Zarnowitz and Lambros (1987)).

… can be used to assess traders' uncertainty

Since surprises in macroeconomic data announcements can affect financial market prices, one would expect that the larger the economic uncertainty about important data releases, the greater the reduction in financial market uncertainty once the data are released. Some evidence suggests that this is indeed the case (Beber and Brandt (2006)). Box 2 shows that when economic uncertainty about NFP outturns is larger, market-based uncertainty about future interest rates, as measured by implied volatilities of options on interest rate swaps, is reduced by more following the announcement of the NFP data.

Uncertainty about US non-farm payrolls[1]



[1] Bureau of Labor Statistics data releases; changes, in thousands. [2] Derived from the cumulative probability distribution implied by option prices. [3] Implied by option prices. [4] Standard deviation of economists' point estimates based on Bloomberg survey.

Sources: Bloomberg; Goldman Sachs; BIS calculations. Graph 3

[16] While the derivation of implied volatility needs to assume a normal probability distribution, the derivation of the interquartile range does not rest on such an assumption. Between the discrete strike prices, the implied cumulative distribution is interpolated linearly in calculating the interquartile range.

Box 2: Impact of economic uncertainty on financial market uncertainty

Since surprises in NFP data releases can affect market interest rates, one might expect that the larger the uncertainty about NFP data releases, the greater the reduction in financial market uncertainty once the data are published (Beber and Brandt (2006)).[①] Here, we investigate this issue using the interquartile range measure, IR_t, of uncertainty about NFP data described in the main text, and using implied volatilities, IV_t, of options on interest rate swaps (swaptions[②]) before ($t–1$) and after (t) the announcement as a measure of financial market uncertainty about interest rates. We consider swaptions with a time to expiry of one month, and with maturities of the underlying interest swap rates of one to 10 years. An advantage of using swaptions is that they have a fixed time to expiry, rather than a fixed expiry date, so that the period over which events can take place and affect uncertainty does not decrease over time. The following regression is estimated on the dates of NFP releases (Beber and Brandt (2006)):

$$(IV_t - IV_{t-1}) \,/\, IV_{t-1} = a + b\, IR_t + u_t$$

Table C shows that swaption-implied volatilities have fallen by significantly more when uncertainty about NFP data releases has been larger and that the effect is greatest when the maturities of the underlying swap rates are two years or less.

Uncertainty about NFP data releases and financial market uncertainty[1]

Maturity	One year	Two years	Five years	Ten years
a	0.12*	0.07	0.02	0.03
b	–0.0012**	–0.0008**	–0.0005*	–0.0006*
R^2	0.16	0.13	0.06	0.10

[1] NFP data for January 2003 to August 2006. ** and * denote significance at the 1% and 5% levels, respectively; Newey-West adjusted standard errors.

Sources: Bloomberg; Goldman Sachs; BIS calculations. Table C

① Beber and Brandt (2006) have found evidence for such a relationship, using an implied volatility measure of uncertainty about NFPs, and using options on US Treasuries and eurodollar futures. ② See Fornari (2005) for a description of swaptions.

Conclusions

Economic derivatives allow market participants to trade directly on macroeconomic data releases and unbundle the news component of announcements from the basis risk contained in financial assets traditionally used as proxies.

Policymakers can use the prices of economic derivatives to obtain information on the perceptions of market participants about the state of the economy. In contrast to survey-based measures, they are true density forecasts, covering the whole distribution of the “market’s view”, not just point estimates. This information could be used to track the uncertainty of market participants about the state of the macroeconomy and to monitor the probabilities they attach to tail events. However, so far this has mainly been possible for US data releases only, with euro area HICP being the exception.

When interpreting the information contained in the prices of economic derivatives, one has to bear in mind that it refers to market participants’ perceptions of the *current* economic situation and not to their expectations of outcomes further ahead. While this may be a limitation when analysing issues such as the transmission mechanism of monetary policy, it may not matter in other settings. For example, the impact of central bank communications might depend on the views of market participants about the current state of the economy, not just on their expectations for the future. That said, it would be useful to have more forward-looking indicators, eg on inflation and growth in the short and medium term, which could complement the information contained in longer-term instruments such as inflation-linked securities.

The potential size of the market for economic derivatives might be limited. In particular, it is not clear whether the market is able to attract a substantial amount of hedging demand, which could serve as a counterweight to highly sophisticated informed traders. In the absence of hedging activity, it is possible that liquidity may dry up in times of limited disagreement between a relatively small number of informed participants.

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Measuring portfolio credit risk: modelling versus calibration errors[1]

A model-based assessment of credit risk is subject to both specification and calibration errors. Focusing on a well known credit risk model, we propose a methodology for quantifying the relative importance of alternative sources of such errors and apply this methodology to a large data set. We find that flawed calibration of the model can substantially affect the measured level of portfolio credit risk. By contrast, a model misspecification generally has a limited impact, especially for large, well diversified portfolios.

JEL classification: C15, G13, G21, G28.

In the wake of recent advances in risk management, models of portfolio credit risk have attracted increasing attention. The validation of such models is of interest to both market practitioners and supervisors, not least because errors in the measurement of credit risk could translate into errors in financial institutions' desired capital buffers. Such errors have different sources. They can be due to a violation of key modelling assumptions (ie *misspecification*) or to wrong estimates of key parameters (ie *flawed calibration*). Thus, a quantification of the relative importance of alternative sources of error in model outcomes would address issues of particular interest to the financial industry.

In this article, we tackle such issues in the context of the well known asymptotic single-risk factor (ASRF) model. This model implies that capital buffers can be set at the level of individual credit exposures and, thus, are *portfolio invariant*. This implication, which is incorporated in the internal ratings-based approach of the Basel II Framework, limits the data and operational requirements on users of the model. However, portfolio invariance rests critically on two assumptions, namely that the systematic component of credit risk is governed by only one common factor and that the portfolio is so finely grained that all idiosyncratic credit risk is diversified away. These assumptions have been criticised in the literature as being too strong and hence as being potential sources of misspecification errors.

[1] The views expressed in this article are those of the authors and do not necessarily reflect those of the BIS. The authors thank Marcus Jellinghaus for valuable help with the data.

By contrast, flawed calibration of the ASRF model, another potential source of error in the measurement of credit risk, has received comparatively little attention. This unbalanced focus in the literature overlooks the fact that accurate estimation of key parameters of this model imposes substantial data requirements. In fact, users who would be attracted by the model's simplicity, embodied in the "portfolio invariance" implication, are likely to also face difficulties in meeting those requirements.

We propose a general methodology for identifying different sources of misspecification and calibration errors in the measurement of portfolio credit risk. Using a data set that contains estimated probabilities of default (PDs) and asset return correlations for a large cross section of firms, we quantify the relative impact of such errors on outcomes of the ASRF model. Our illustrative exercise suggests that model-implied measures of portfolio credit risk are more sensitive to plausible calibration errors than to misspecification errors. This is especially true for larger portfolios where the violation of the granularity assumption is less pronounced.

The rest of this article is organised in four sections. After a brief overview of the related literature, the first section discusses at a conceptual level alternative sources of error in the calculation of portfolio credit risk. Then, the second section spells out the empirical methodology, which quantifies the relative importance of these sources in a unified framework. The third section describes the data and reports the empirical findings. The final section summarises the contribution of this analysis and identifies directions for future research.

Calculated versus target capital: conceptual issues

The ASRF model postulates that an obligor defaults when the value of its assets falls below a particular threshold. In addition, the model assumes that credit risk is related across obligors owing to the dependence (or the "loading") of their assets on a single common risk factor and that the portfolio consists of a large number of small exposures (ie is of fine "granularity"). In this model the capital that covers all portfolio losses with a desired probability can be calculated at the level of individual exposures. In turn, an exposure-specific capital depends solely on the corresponding PD and common-factor loading.

Model misspecification and flawed calibration …

The literature has paid closer attention to misspecification of the ASRF model than to the potentially flawed calibration of its inputs. Analysis of violations of ASRF assumptions has led to proposals of ways to mitigate the impact of such violations on capital calculations. The various proposals, which have been reviewed in BCBS (2006), attempt to strike a balance between the reduction of errors and the associated data or computational burden.[2] However, users who rely on the stylised ASRF model in order to

[2] Adjustments to capital measures that correct for violations of the ASRF granularity assumption have been derived in Martin and Wilde (2002), Emmer and Tasche (2003) and Gordy and Lütkebohmert (2006). In turn, violations of the single-common-factor assumption have been the focus of Pykhtin (2004), Düllmann (2006), Garcia Cespedes et al (2006) and Düllmann and Masschelein (2006). In addition, Heitfield et al (2006) and Düllmann et al (2006)

alleviate such a burden are also likely to face constraints in estimating the model's parameters. Thus, they are prone to an imperfect calibration of the model, which would generate additional bias in the assessment of portfolio credit risk.[3]

... affect popular measures of portfolio credit risk

In this special feature, we extend previous analyses by examining four sources of error in a model-based assessment of portfolio credit risk. Two of these sources relate to the possible misspecifications of the ASRF model that have received much attention in the literature. The other two relate to an erroneous calibration of the correlation of credit risk across exposures.[4]

In the remainder of this section, we define and discuss each of these sources of error at a conceptual level. The metric we use to compare different model outcomes is credit value-at-risk (net of expected losses), which is equivalent to the capital buffer necessary to cover default losses with a desired probability.[5] The comparison is based on a hypothetical benchmark assessment of risk, which assumes knowledge of all relevant parameters and is referred to as "target capital", and alternative assessments, which are subject to one or more of the above-mentioned sources of error.

Multiple factors of credit risk

The obligors in a credit portfolio are likely to be affected not only by aggregate economic conditions but also by conditions relevant for specific business lines. Conceptually, these various conditions can be summarised in mutually independent and often unobservable risk factors. If several of these factors are of material importance, the single-factor assumption of the ASRF model would be violated. This would entail systematic errors in model-implied measures of credit risk and, consequently, in the capital set aside to compensate for it.

Violations of the single-factor assumption ...

A violation of the single-factor assumption is conceptually different from a failure to measure the impact of multiple factors on the correlation across obligors. Such a failure is independent of a modelling misspecification and can arise, for example, when higher concentration in a particular industrial sector is not captured in the estimated average correlation. However, even if the average correlation across obligors is measured accurately, an erroneous single-factor assumption ignores the fact that there are multiple sources of default clustering. This leads to an underestimation of the probability of a large number of defaults and, consequently, to an underestimation of the target capital.

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have examined both assumptions in the context of representative portfolios of US and European banks, respectively.

3 For a discussion of the impact of estimation errors on capital calculations, see Löffler (2003).

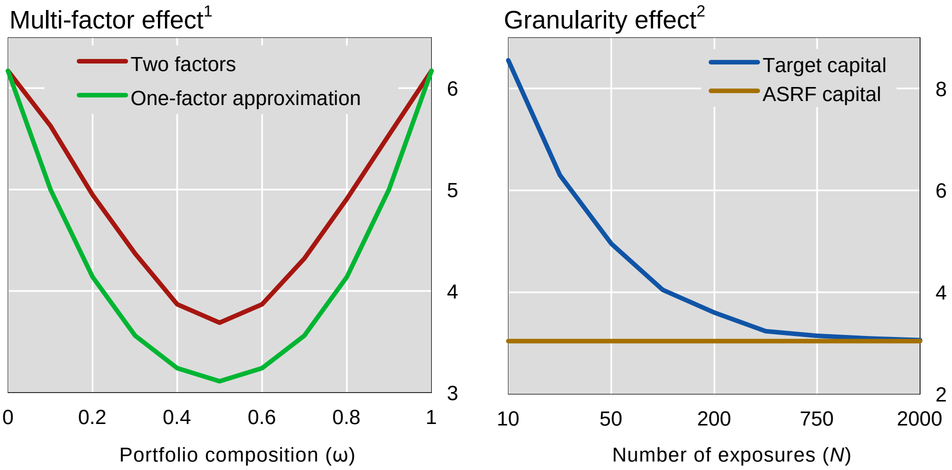
4 In order to focus on "pure portfolio" characteristics of credit losses, we abstract from errors in the measurement of PDs and losses-given-default (LGDs).

5 In this article the terms "model outcome", "assessment of portfolio risk" and "capital" are used interchangeably. Importantly, our capital calculations do not correspond to "regulatory capital", which reflects considerations of bank supervisors, or to "economic capital", which reflects additional strategic and business objectives of financial firms.

Graph 1 (left-hand panel) illustrates such an effect in the context of a stylised portfolio, in which exposures are the same across obligors and have homogeneous PDs and losses-given-default (LGDs). In this portfolio, a fraction ω of the obligors belong to group 1, while 1–ω belong to group 2. The associated asset returns are uncorrelated across groups but are affected by a group-specific common factor. Thus, increasing the value of ω between 0 and 1 increases the relative importance of the first common factor for the overall credit risk in the portfolio. The red line plots the target capital as a function of ω,[6] while the green line portrays an alternative capital calculation implied by a single-factor structure. This structure matches exactly the average level of asset return correlations across exposure pairs.

The difference between the red and green lines illustrates the multi-factor effect that we analyse empirically below. This difference is largest when the role of multiple factors is greatest and, hence, when a single-factor structure approximates most poorly the dispersion of asset return correlations. In our example, this occurs at ω = 1/2.

Modelling misspecifications and capital, illustrative examples



Note: Calculated capital, in per cent, per unit of aggregate exposure, on the vertical axes. For each panel, PD = 1% and LGD = 45% are the same across exposures.

[1] The portfolio consists of two groups of exposures, with ω denoting the weight of the first group. Within each group, the asset return correlation equals 20% for all exposure pairs. The inter-group correlation is zero. The red line plots the target capital level, which incorporates two common factors. The green line plots the capital calculated under a one-common-factor approximation of the correlation structure. This approximation imposes the same common-factor loadings across firms and does not change the average asset return correlation. [2] The blue line plots target capital for a portfolio in which all pairwise asset return correlations equal 10%. The number of exposures in this portfolio varies across the horizontal axis. The brown line plots the corresponding capital estimates when *N* is assumed to be infinite. Graph 1

[6] Target capital is lowest at ω = 1/2 because at this value the portfolio is evenly diversified between the two common factors, which minimises the probability of large losses. The dependence of the desired capital buffer on the relative exposure to multiple factors is studied by Düllmann and Masschelein (2006), at both a theoretical and an empirical level.

Granularity

… and the granularity assumption …

The "perfect granularity" assumption of the ASRF model postulates that all exposure-specific, ie idiosyncratic, risk is diversified away. Since this cannot be attained by any real-world portfolio, the granularity assumption leads to an underestimation of the overall credit risk and, consequently, to an underestimation of target capital. Given an overall correlation across exposures, this underestimation is smaller for a portfolio that comprises more obligors and, thus, benefits from greater diversification gains. This is illustrated in the right-hand panel of Graph 1.

Correlation level

Even if the ASRF model is correctly specified, calculated capital may be affected by errors in the inputs to this model. For instance, a user of the model who has data constraints may choose to rely on readily available external estimates of the asset return correlations in popular credit indices. Such estimates would lead to a discrepancy between target and calculated capital to the extent that the underlying indices were not representative of the user's own portfolio.

… and erroneous correlation estimates …

One driver of the potential discrepancy is a bias in the average level of the asset correlations underpinning the calculation of capital. When this bias is positive, for instance, it inflates the probability that a large number of defaults might occur simultaneously and leads to an overestimation of the target capital. Conversely, a negative bias leads to an underestimation. This is the correlation level effect we examine empirically below.

Dispersion in pairwise correlations

The effect of a bias in the overall level of calibrated correlations could be augmented by errors in the dispersion of these correlations across exposure pairs. Such errors are likely to emerge either if users of the ASRF model rely on external estimates of asset return correlation or if they apply the average of internal estimates to all exposures. This would lead to a correlation dispersion effect on calculated capital. Even though in practice this effect is likely to be tightly related to the correlation level effect, our empirical methodology will disentangle the two in order to quantify their separate roles.

The effect of correlation dispersion on calculated capital is seen most clearly in a stylised example. Suppose that all firms in one portfolio have homogeneous PDs and exhibit homogeneous pairwise asset return correlations. Suppose further that a second portfolio is characterised by the same PDs and average asset return correlation but includes a group of firms that are more likely to default together. The second portfolio is more likely to experience several simultaneous defaults and, thus, requires higher capital in order to attain solvency with the same probability.

However, this result can be weakened or even reversed if PDs vary across firms. To see why, suppose that the strongly correlated firms in the second portfolio are the ones that have the lowest individual PDs. In other words, the firms that are likely to generate multiple defaults are less likely to default. This may lower the probability of default clustering, depressing the target capital level below that for the first portfolio.

Empirical methodology

Our empirical methodology comprises two general steps. In the first step, we construct a large (small) hypothetical portfolio that comprises equal exposures to 1,000 (200) firms.[7] The sectoral composition of this portfolio is designed to be in line with the typical loan portfolio of large US wholesale banks.[8] Given the constraints of such a composition, the portfolio is drawn at random from our sample of firms. Since each draw could be affected by sampling errors, we examine 3,000 different draws for both large and small portfolios.

… can be analysed in a unified framework …

For a portfolio constructed in the first step, the second step calculates five alternative capital measures, which differ in the underlying assumptions regarding the interdependence of credit risk across exposures.[9] Each of these alternatives employs the same set of PD estimates, and assumes that LGD equals 45% for all exposures and that asset returns are normally distributed.[10] We order the measures so that each measure differs from a previous one owing to a *single* assumption.[11]

The first measure is the *target* capital, which incorporates data on asset return correlation estimates. Assuming that these and the other risk parameter estimates we adopt (as well as our distributional assumption) are accurate, we conduct a Monte Carlo simulation to construct the probability distribution of default losses at the one-year horizon. Then, we set the target capital to a level that covers unexpected default losses with a probability of 99.9%,[12] recognising that our methodology also applies to alternative definitions of target capital.

… as sources of deviations from target capital

The second capital measure differs from the target one owing to a restriction on the number of common factors governing asset returns. Specifically, we use a correlation matrix of asset returns that can emerge in the presence of a single common risk factor but fits as closely as possible the

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7 In this analysis, the distinction between large and small portfolios is not based on the size of the aggregate exposure but reflects different degrees of diversification across individual exposures.

8 Such a portfolio does not incorporate consumer loans and, thus, may not fully represent all aspects of credit risk. To construct a large portfolio, we apply the 40 sector-specific weights provided by Heitfield et al (2006). For a small portfolio, we rescale the 10 largest sectoral weights so that they sum up to one and set all remaining weights to zero.

9 The box on page 89 provides further detail on the calculation of each of the five capital measures.

10 The particular data used in this article are described in the next section.

11 We choose one of several possible orderings of the five capital measures. On the basis of background analysis, we are confident that an alternative ordering would not change our main conclusions significantly, even if it altered specific numerical results.

12 The covered level of unexpected losses equals the 99.9th percentile of the distribution of credit losses minus the mean of this distribution.

Calculating capital measures: technical details

This box outlines three general methods for calculating the distribution of default losses. These methods are used to derive the five capital measures considered in the text.

The first method relies on Monte Carlo simulations and delivers the target capital level. This method can be applied to any portfolio comprising *N* equally weighted exposures, provided that the associated probabilities of default, *PD*, losses-given-default, *LGD*, and correlation matrix of asset returns, *R*, are known. The method consists of three general steps. In the first step, one uses the vector of PDs and the assumption that asset returns are distributed as standard normal variables to obtain an *N* x 1 vector of default thresholds. In the second step, one draws an *N* x 1 vector from *N* standard normal variables whose correlation matrix is *R*. The number of entries in this vector that are smaller than the corresponding default threshold is the number of simulated defaults for the particular draw. In the third step, one repeats the second step a large number of times to derive the probability distribution of the number of defaults. Denoting this distribution's (1 – α)th percentile by β and the average *PD* in the portfolio by *A*(*PD*), the target capital for a credit value-at-risk confidence level of (1 – α) equals *LGD**(β – *A*(*PD*)) per unit of exposure.[1]

The second method relies on the so-called Gaussian copula, which is outlined in detail in Gibson (2004). This method delivers the second capital measure in the main text and rests on the assumptions that (i) the portfolio consists of *N* equally weighted exposures with identical *LGD*[2] and (ii) only one common factor drives credit risk. To apply this method, one needs to calibrate the LGD, obtain values for the firm-specific PDs and estimate firm-specific loading coefficients, l_i , which are defined by the following equation:

$$V_i = l_i M + \sqrt{1 - l_i^2}\, Z_i \qquad (1)$$

where V_i is the asset value of firm *i*, *M* is the common risk factor and Z_i is the idiosyncratic risk factor. Equation (1) implies that the correlation between the asset returns of firm *i* and *j* equals $l_i * l_j$. A particular estimate of l_i is obtained by fitting the single-common-factor assumption to the original correlation structure in a mean-squared-error sense, ie by minimising the following sum:

$$\sum_{i=1}^{N-1} \sum_{j>i} \left(\rho_{ij} - l_i l_j\right)^2 \qquad (2)$$

where ρ_{ij} is an element of the correlation matrix *R*. Andersen et al (2003) provide an efficient algorithm for solving this minimisation problem. Estimated in this way, the loading coefficients l_i account almost exactly for the average pairwise correlation in *R*.

The third method, which applies to the other three measures in the main text, is a special case of the ASRF model. In comparison to the second method, it makes the additional assumption that all idiosyncratic risk is diversified away. This implies that the capital buffer for the portfolio is the sum of exposure-specific capital buffers, κ_i, which are calculated as follows:

$$\kappa_i = LGD * \left[\Phi\left(\frac{\Phi^{-1}(PD_i) - l_i \Phi^{-1}(\alpha)}{\sqrt{1 - l_i^2}} \right) - PD_i \right] \qquad (3)$$

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[1] This article sets *N* = 200 or 1,000, *LGD* = 45%, α = 0.1% and *PD* and *R* as estimated by Moody's KMV. In addition, the third step of the first method carries out 500,000 Monte Carlo simulations. [2] For a calculation of the probability distribution of credit losses when individual exposures have different LGDs and portfolio weights, see Hull and White (2004).

unrestricted correlation matrix underpinning the target capital calculation. The difference between the resulting capital estimate and the target level is denoted the "*multi-factor effect*".

The third measure differs from the second one in that it assumes, in addition, that all idiosyncratic credit risk is diversified away. In other words, it ignores the impact of imperfect granularity in the portfolio. This assumption allows one to apply the ASRF model, which delivers an analytic solution for capital calculations. The difference between the third and second measures is the "*granularity effect*".

The fourth measure differs from the third one in that it is based on the assumption that asset return correlations are the same across all exposures in the portfolio. The common correlation, which is set equal to the average of the pairwise correlations underpinning the third capital measure, is used as an input to the ASRF model. The difference between the fourth and third capital measures is denoted the "*correlation dispersion effect*".

Finally, the fifth measure incorporates a bias in the estimates of asset return correlations. This measure differs from the fourth one in that it relies on a standard rule-of-thumb value of the (common) asset return correlation. The difference between the fifth and fourth measures is denoted the "*correlation level effect*".

An important feature of this methodology is that it allows one to quantify the relative importance of alternative drivers of capital miscalculations. Specifically, the methodology can be applied to dissect the difference between the fifth capital measure, which we henceforth dub the "shortcut" one, and the first, ie target, capital measure. By construction, this difference equals exactly the sum of the multi-factor, granularity, correlation dispersion and correlation level effects.

Empirical results

In this section, we employ the methodology outlined above to investigate the discrepancy between target and shortcut capital for simulated portfolios with realistic features. In addition, we derive the relative importance of alternative drivers of this discrepancy.

Data

Data on risk parameter estimates suggest that …

Our empirical analysis relies on two data sets provided by the commercial service Moody's KMV. One data set consists of one-year expected default frequencies (EDF™) that are point-in-time estimates of individual PDs, while the other comprises estimates of pairwise asset return correlations implied by the GCorr model. Both EDF and GCorr models are based on an extended and operational version of the seminal framework of Merton (1974), which is broadly consistent with the ASRF model.[13]

Combining these two data sets, we obtain a pool of 10,891 non-financial firms, for which individual PD and pairwise correlation estimates are available.

[13] See Crosbie and Bohn (2003), Das and Ishii (2001) and Crosbie (2005) for a description of these proprietary models and related references. The Moody's KMV sample comprises only firms with publicly traded equities. A multi-factor loading structure is employed for the estimation of the GCorr model.

Characteristics of simulated portfolios[1]					
In per cent					
A. Large portfolios (1,000 exposures)					
	Mean	Standard deviation	Median	Minimum	Maximum
Average PD	2.42	0.19	2.42	1.79	3.12
Standard deviation of PDs	5.16	0.26	5.16	4.25	6.14
Median PD	0.26	0.03	0.26	0.18	0.36
Average correlation[2]	9.78	0.22	9.77	9.14	10.73
Standard deviation of loadings[3]	9.33	0.31	9.32	8.33	10.47
Corr (PD, loadings)[4]	–20.00	2.04	–20.10	–26.70	–12.80
B. Small portfolios (200 exposures)					
	Mean	Standard deviation	Median	Minimum	Maximum
Average PD	2.28	0.36	2.26	1.24	3.68
Standard deviation of PDs	5.05	0.53	5.06	3.01	6.89
Median PD	0.24	0.05	0.23	0.11	0.55
Average correlation[2]	10.49	0.44	10.48	8.99	12.00
Standard deviation of loadings[3]	10.54	0.70	10.55	7.80	12.79
Corr (PD, loadings)[4]	–19.80	4.59	–20.20	–31.80	–1.20

Note: The calculations in this table use 3,000 simulated portfolios for each portfolio size and are carried out in two steps. First, portfolio-specific statistics specified by row headings are calculated for each simulated portfolio. Second, summary statistics specified by column headings are calculated for each of the portfolio-specific statistics calculated in the first step.

[1] Based on Moody's KMV estimates of PDs and asset return correlations for July 2006. [2] Based on all pairwise correlations. [3] The derivation of common-factor loadings assumes that there is a single common factor and implements the procedure outlined in the box (page 89). [4] The sample correlation between PDs and loadings on the single common factor. Table 1

The vast majority of the firms in the sample are headquartered either in the United States (52% of the total number) or in western Europe (40%). For illustrative purposes, we use EDF and GCorr estimates in July 2006 as the "true" PDs and correlations underpinning the target capital level. Of course, any error in these estimates would warrant a revision of the target capital.

Table 1 summarises the characteristics of the simulated portfolios. For both large and small portfolios, the distribution of EDFs has a long right tail, with the median values much lower than the mean. Correlation estimates are clustered mainly between 5 and 25%, with their mean standing at 9.78% for large portfolios and 10.49% for small ones. In addition, reflecting the benign credit conditions during the sample period, more than 10% of the sample firms have the lowest PD estimates permitted by the EDF model (ie 0.02%).

Target versus shortcut capital levels

… deviations from target capital can be large …

We proceed to quantify and decompose the difference between target and shortcut capital measures (Table 2). For illustrative purposes, the constant correlation underlying shortcut calculations is set at 12%. This is about

Four sources of error in estimated capital[1] Per unit of aggregate exposure, in per cent					
A. Large portfolios (1,000 exposures)					
	Mean	Standard deviation	Median	95% interval	50% interval
Target capital[2]	2.95	0.16	2.95	[2.64, 3.27]	[2.84, 3.05]
Deviation from the target due to:[3]					
Multi-factor effect	–0.03	0.03	–0.05	[–0.09, 0]	[–0.05, 0]
Granularity effect	–0.11	0.01	–0.11	[–0.14, –0.09]	[–0.12, –0.10]
Correlation dispersion effect	0.35	0.04	0.35	[0.27, 0.43]	[0.32, 0.38]
Correlation level effect	0.55	0.06	0.55	[0.44, 0.66]	[0.52, 0.59]
"Shortcut" capital (correlation = 12%)	3.71	0.18	3.71	[3.37, 4.06]	[3.59, 3.83]
Memo: correlation level effect if:					
Correlation = 6%	–0.96	0.07	–0.96	[–1.11, –0.83]	[–1.00, –0.91]
Correlation = 18%	2.01	0.09	2.01	[1.84, 2.18]	[1.95, 2.07]
Correlation = 24%	3.47	0.13	3.47	[3.23, 3.72]	[3.39, 3.56]
B. Small portfolios (200 exposures)					
	Mean	Standard deviation	Median	95% interval	50% interval
Target capital[2]	3.35	0.30	3.34	[2.78, 3.94]	[3.15, 3.53]
Deviation from the target due to:[3]					
Multi-factor effect	–0.04	0.10	0	[–0.23, 0]	[0, 0]
Granularity effect	–0.53	0.07	–0.53	[–0.65, –0.41]	[–0.59, –0.47]
Correlation dispersion effect	0.38	0.11	0.37	[0.17, 0.58]	[0.30, 0.45]
Correlation level effect	0.36	0.11	0.36	[0.15, 0.61]	[0.29, 0.44]
"Shortcut" capital (correlation = 12%)	3.52	0.34	3.51	[2.85, 4.23]	[3.28, 3.75]
Memo: correlation level effect if:					
Correlation = 6%	–1.07	0.12	–1.07	[–1.31, –0.85]	[–1.15, –0.99]
Correlation = 18%	1.76	0.19	1.75	[1.41, 2.15]	[1.63, 1.88]
Correlation = 24%	3.15	0.27	3.14	[2.65, 3.70]	[2.97, 3.33]

[1] Summary statistics for the simulated portfolios underpinning Table 1 (3,000 for each portfolio size). The column entitled "95% interval" reports the 2.5th and 97.5th percentiles of the statistics specified in the particular row heading. The column entitled "50% interval" reports the corresponding 25th and 75th percentiles. [2] Based on Moody's KMV estimates of PDs and asset return correlations and a Monte Carlo procedure for calculating the probability distribution of default losses. [3] Four sources of deviation from the target capital level; a negative sign implies underestimation. The sum of the target capital level and the four deviations equals the shortcut capital level. Each deviation is based on the assumptions underlying previous deviations plus one additional assumption: (a) for the multi-factor effect, the correlation matrix underpinning the target capital level is approximated under the assumption that there is a single common factor; (b) for the granularity effect, there is the additional assumption that the number of firms is infinite; (c) for the correlation dispersion effect, the additional assumption is that the loadings on the single common factor are the same across exposures; (d) for the correlation level effect, the additional assumption imposes a different level for the constant pairwise correlation. See the box on page 89 for further detail on alternative capital measures. Table 2

2 percentage points higher than the average asset return correlation in the simulated portfolios (recall Table 1) and is close to the 12.5% rule-of-thumb correlation suggested by Lopez (2004).

The results show that the shortcut capital measure can be significantly higher than the corresponding target level. The difference is much more pronounced in the context of large portfolios, for which it amounts on average to 76 basis points (per unit of exposure, or 26% of the target capital). This is predominantly the result of the correlation dispersion and correlation level effects. By contrast, these two effects are almost fully offset by the granularity effect in small portfolios, for which the shortcut capital estimate is 5% higher than the target level.[14] The following subsections discuss in some detail the four alternative effects behind the overall discrepancies between target and shortcut capital.[15]

Multi-factor effect

The multi-factor effect lowers the model-implied capital measure, for the reasons outlined above, but its quantitative impact is almost negligible. Imposing a single-factor structure on asset returns leads to a capital allocation on large portfolios that is, on average, 1% lower than the target level. At 1.2%, this decline is only slightly larger for small portfolios.

The low importance of the multi-factor effect is a result of the fact that a single-factor framework approximates quite well the multi-factor structure in the data. For the portfolios used in this exercise, the single-factor framework outlined in the box on page 89 explains almost perfectly the level of the original correlations (with an error of less than 2 basis points) and accounts for the bulk (76% on average) of the cross-sectional variation in pairwise correlations. The robustness of this finding to alternative portfolio specifications and alternative estimates of risk parameters is an important question for future research.

Granularity effect

As expected, the granularity assumption of the ASRF model leads to an underestimate of the target capital ratio. Importantly, the underestimation increases when the size of the portfolio decreases. As Table 2 reports, the granularity effect leads to a 4% underestimation of the target capital for large, diversified portfolios and a 16% underestimation for small, less diversified portfolios. These results are in line with previous analyses of the granularity effect.[16]

In practice, the size of the exposures would vary across obligors, which would complicate the analysis of the granularity effect. For example, a portfolio that consists of a large number of exposures but is highly concentrated in a subset of them can be associated with a larger granularity effect than a portfolio with a smaller number of equally weighted exposures. Our

[14] Even though the exercise focuses on a particular sectoral distribution of exposures, credit risk does differ across the simulated portfolios. Accordingly, columns 2 to 5 in Table 2 report descriptive statistics of the distribution of the portfolio-specific capital estimates.

[15] In quantifying the magnitude of each effect, the adopted sign is such that an effect can be added to the target capital level or subtracted from the corresponding shortcut level.

[16] For instance, Gordy and Lütkebohmert (2006) propose an adjustment formula to correct for the granularity effect. An application of this formula (equation (6) in their paper) matches exactly a granularity effect that leads to a 5.4% underestimate of the target capital for large portfolios and a 24% underestimate for small portfolios.

methodology could also accommodate such cases, but we abstract from them in this special feature in order to simplify the exposition.[17]

Correlation dispersion effect

… especially if the model is erroneously calibrated

Equalising the asset return correlations across exposure pairs causes calculated capital to be more conservative than the target level. In particular, target capital is overestimated by 12% for large portfolios. At 11% for small portfolios, this overestimation classifies the correlation dispersion effect as the most important of the four considered sources of discrepancies between shortcut and target capital. In line with the intuition presented above, the positive sign of the correlation dispersion effect is due to the fact that, in our sample, higher-PD exposures tend to be less correlated among themselves (Table 1).[18]

Correlation level effect

A mismatch between the average correlations underpinning target and shortcut capital calculations would also have a substantial effect. Increasing the average asset return correlation from 9.8% (the level estimated by Moody's KMV for the simulated portfolios) to 12% leads to a 19% overestimation of the target capital for large portfolios. For small portfolios, an average asset return correlation of 12% implies a smaller but still significant overestimation of 11%.

This result is not surprising, because a higher average asset return correlation translates into a higher probability of default clustering, which raises the estimated capital. Alternatively, however, the average level of asset return correlations may be underestimated, which would lead to insufficient capital. Table 2 reports that setting this level to 6% would lead to underestimating the target capital level by about 32% for both portfolios.[19]

Conclusion

In this article, we developed an approach to evaluating errors in the measurement of portfolio credit risk. In particular, we used this approach to quantify the magnitude of different sources of a discrepancy between a predefined target capital level and a shortcut alternative, which is based on the ASRF model and rule-of-thumb correlation estimates. On the basis of simulated

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[17] Accommodating disparate exposures would introduce an additional dimension in portfolio characteristics, requiring the simulation of a greater variety of hypothetical portfolios and making it more difficult to interpret the multi-factor and correlation dispersion effects.

[18] The negative relationship between PDs and correlations (ie loading coefficients) is likely to be a general phenomenon. For example, Dev (2006) finds that global factors often play bigger roles for firms of better credit quality.

[19] Table 2 reports the correlation level effect based on alternative levels of average correlations: 6%, 18% and 24%. These alternatives correspond to rule-of-thumb correlation values reported in previous studies (between 5 and 25%) and to plausible estimation errors. As regards such errors, Tarashev and Zhu (2007) show that, for a true constant correlation of 9.78% and five years of monthly data on asset returns, the 95% confidence interval for the estimated average correlation is between 6.4 and 13.3%.

portfolios, we found that plausible errors in estimated asset return correlations could lead to substantial deviations from the target capital levels for both large and small portfolios. By contrast, the violation of key assumptions of the ASRF model, ie the single-factor or the perfect granularity assumption, tend to result in relatively smaller errors in calculated capital. The only exception is that the granularity assumption does have a significant impact for small portfolios.

The illustrative nature of our analysis identifies different avenues for future research. For one, it would be valuable to analyse the robustness of our empirical results to alternative portfolio specifications and to different (realistic) values of PDs, LGDs and asset return correlations. In addition, it would be important to derive rigorously the range of plausible estimation errors in the parameters used to calculate portfolio credit risk and to study the implications of alternative assumptions as regards the distribution of asset returns. Tarashev and Zhu (2007) attempt to address this latter set of issues.

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Recent initiatives by the Basel-based committees and groups

During the period under review, the Basel Committee on Banking Supervision (BCBS) published a report on international developments in banking supervision and a working paper on credit risk concentration. The Committee on the Global Financial System (CGFS) released a report on institutional investors, and the Committee on Payment and Settlement Systems (CPSS) published the final version of its report (prepared jointly with the World Bank) on general principles for international remittance services. The Financial Stability Institute (FSI) published a paper highlighting the key findings of its updated Basel II implementation survey. Table 1 provides a summary of these and other initiatives.

Basel Committee on Banking Supervision

In the fourth quarter of 2006, the Committee published a report on international developments in banking supervision, as well as a working paper on credit risk concentration. Moreover, it clarified the risk weighting for the International Finance Facility for Immunization and issued a comment on a discussion paper of the International Accounting Standards Board.

BCBS publishes 15th *Report on International Developments in Banking Supervision*

On 26 October the Committee released the 15th *Report on International Developments in Banking Supervision.* The aim of the report was to brief bank supervisors of international supervisory developments in advance of the 14th biennial International Conference of Banking Supervisors (ICBS), held in Mérida, Mexico on 4 and 5 October 2006. For the first time since the inception of this biennial series, the 15th report was made publicly available.

The report is mainly devoted to providing an overview of the work of the Basel Committee and other international groups of banking supervisors since the 13th ICBS in Madrid in September 2004. Commentary on the work of the Basel Committee over the past two years is provided in Chapter II. Chapter III presents a summary of the results of a recent Quantitative Impact Study (QIS 5), which was designed to evaluate the effects of the Basel II Framework in comparison with the current capital standards established in 1988. Chapter IV contains details of the Basel Committee’s publications and website. The work of the Financial Stability Institute (FSI) is discussed in Chapter V, while Chapter VI describes the work of the regional supervisory groups. The

Main initiatives by Basel-based committees and other bodies

Press releases and publications over the period under review

Body	Initiative	Thematic focus	Release date
BCBS	*Report on International Developments in Banking Supervision*	• Work of the Basel Committee over the past two years • Summary of the results of QIS 5 • Details of the Basel Committee's publications and website, contact information for the international groups of supervisors • Work of the FSI and of the regional supervisory groups	October 2006
	Risk weight for International Finance Facility for Immunization (IFFIm)	• Confirmation of a 0% weighting on the IFFIm	
	Comments on the IASB's discussion paper on preliminary views on an improved conceptual framework for financial reporting	• Assessment of stewardship as an objective for financial reporting • Replacement of the concept of reliability with that of faithful representation • Definition of verifiability	November 2006
	Studies on credit risk concentration: an overview of the issues and a synopsis of the results from the Research Task Force project	• Overview of the issues and current industry practice; related policy issues • Deviations of economic capital from Pillar 1 capital charges in the IRB • Fit-for-purpose tools that can be used in the quantification of concentration risk	
FSI	*Implementation of the new capital adequacy framework in non-Basel Committee member countries*	• Update of the FSI's Basel II implementation survey, detailed statistics on implementation plans by pillars, themes and countries	October 2006
CGFS	*Institutional investors, global savings and asset allocation*	• Observed trends • Impact of regulatory and accounting changes • Policy implications	February 2007
CPSS	*Statistics on payment and settlement systems in selected countries*	• Preliminary statistics for 2005	November 2006
	General principles for international remittance services	• Payment system aspects of international remittance services, general principles for improving this market	January 2007

Source: Relevant bodies' websites (www.bis.org). Table 1

report also provides contact information for the international groups of supervisors.

Working paper highlights credit risk concentration as a major source of bank distress

On 30 November, the BCBS released a working paper entitled *Studies on credit risk concentration: an overview of the issues and a synopsis of the results from the Research Task Force project*. Historical experience shows that concentration of credit risk in asset portfolios has been one of the major causes of bank distress. This is true both for individual institutions and for banking systems at large. It is therefore important to measure the concentration risk arising in banks' credit portfolios from two sources, systematic and idiosyncratic. Systematic risk represents the effect of unexpected changes in macroeconomic and financial market conditions on the performance of borrowers. Idiosyncratic risk represents the effects of risks that are peculiar to individual firms.

RTF analytical project had three main angles:

The Concentration Risk Group of the Research Task Force of the BCBS undertook a principally analytical project with the following objectives: (i) to provide an overview of the issues and current practice in a sample of the more advanced banks as well as to highlight the main policy issues that arise in this context; (ii) to assess the extent to which "real world" deviations from the "stylised world" behind the assumptions of the IRB model can result in important deviations of economic capital from Pillar 1 capital charges in the IRB approach; and (iii) to examine and further develop fit-for-purpose tools that can be used in the quantification of concentration risk.

the current state of the art …

… the effect of departure from model assumptions …

… and the role of stress tests

The paper provides an overview of the work conducted by this group and its findings. The work of the group was divided into three workstreams. The first workstream collected information about the current "state of the art" both in terms of industry best practice and in terms of the developments in the academic literature. The second workstream focused on gauging the impact of departures from the ASRF model assumptions on economic capital and examined various methodologies that can help to bridge the gap between underlying risk and risk measured by the specific model. The workstream had two sub-themes that focused on name concentration risk (imperfect portfolio granularity) and sector concentration risk (imperfect diversification across risk factors). The third workstream dealt mostly with the ability of stress tests to detect excessive concentration (of either type) and to provide estimates of economic capital in stress scenarios.

In a *newsletter* published on 24 October, the BCBS agreed that supervisors may allow banks to apply a 0% risk weight to claims on the International Finance Facility for Immunization, similar to claims on a multilateral development bank, in accordance with paragraph 59 of the Basel II Framework.

BCBS comments on IASB discussion paper on financial reporting

The discussion paper published by the International Accounting Standards Board (IASB) in July 2006 entitled "Preliminary views on an improved conceptual framework for financial reporting: the objective of financial reporting and qualitative characteristics of decision-useful financial reporting information" has generated a lot of interest in the international supervisory community. The BCBS formulated *comments* on the IASB consultative document, with particular focus on three aspects: (i) the assessment of stewardship as an objective for financial reporting; (ii) the replacement of the concept of reliability with that of faithful representation; and (iii) the definition of verifiability.

Financial Stability Institute

On 25 October 2006, the FSI published an occasional paper on the *Implementation of the new capital adequacy framework in non-Basel Committee member countries*.

In view of the ongoing challenges and opportunities presented by the Basel II implementation process, the Financial Stability Institute decided to follow up on its original Basel II implementation survey, conducted in 2004, in order to take stock of developments since then.

FSI publishes results of its latest survey on Basel II implementation in non-BCBS member countries

Responses to the survey were received from 98 jurisdictions outside the BCBS, of which 82 intend to implement Basel II. For Pillar 1 (minimum capital requirements), 85% of the respondents intending to adopt Basel II expect to adopt the standardised approach for calculating capital requirements for credit risk. The foundation internal ratings-based (FIRB) approach and the advanced IRB approach are expected to be implemented by 67% and 55% of those adopting Basel II, respectively. With regard to operational risk under Pillar 1, the basic indicator approach is anticipated to be widely employed across regions – by 79% of those adopting Basel II – followed by the standardised approach, at 70%. About 51% of those adopting Basel II expect to adopt an advanced measurement approach for operational risk. Of the 82 jurisdictions, 61 expect to implement Pillar 2 (supervisory review process) by 2008, with an additional nine jurisdictions expecting to implement it during 2009–15. As regards Pillar 3 (market discipline), 54 jurisdictions expect to implement it by 2008 and an additional 12 jurisdictions during 2009–15.

Committee on the Global Financial System

On 28 February 2007, the CGFS released a report on *Institutional investors, global savings and asset allocation,* prepared by one of its working groups. The report identifies some current trends in the institutional investor industry, examines the impact of recent and prospective regulatory and accounting changes, and identifies some policy implications.

CGFS report highlights growing importance of institutional investors for financial systems and economies globally

Among the current trends indentified, institutional investors are becoming more important in global financial markets, with their assets under management rapidly catching up with those of the banking system. Institutional investors help to ensure deeper and better functioning markets, thus contributing to a more efficient allocation of savings, and their growth may help to counter the decline of household saving ratios associated with ageing populations. Global institutional investors have increased their exposure to emerging market economies (EMEs) in recent years. Domestic institutional investors in EMEs, although small in an absolute sense, are significant relative to the size of local markets and have considerable growth potential. The working group also analysed the importance of alternative investments and the effect of different objectives and strategies (for instance the defined benefit or defined contribution nature of the related pension schemes, or varying asset-liability management strategies) on investment behaviour.

Impact of regulatory and accounting changes affecting institutional investors

The working group was mandated to assess the effect on investment behaviour and financial markets of specific regulatory and accounting changes affecting pension funds and insurance companies. These changes were of a global nature and motivated, at least partially, by the 2000–02 equity downturn, which exposed some serious weaknesses in the regulatory frameworks affecting traditional institutional investors with long-term liabilities and offering guaranteed returns in many countries.The main effect of reforms will be to provide incentives for defined benefit pension funds and insurance companies to reduce their risk profile, either by transferring investment risk to households or by adopting investment strategies that directly incorporate liabilities into asset allocation decisions. The adoption of these techniques by institutional investors may involve shifts in asset portfolios from equities to long-term conventional and index-linked bonds, whose financial characteristics more closely resemble liabilities in terms of duration and the cash flow of obligations. In assessing influences on institutions' asset allocation decisions and on market dynamics, the working group noted that it may be difficult to disentangle the effect of regulatory and accounting changes from other factors, such as the current low level of long-term interest rates observed at a global level. However, developments in the United Kingdom, where long-term yields appear to have been affected by recent changes in institutions' asset allocation strategies, illustrate the potential importance of regulatory policy changes affecting intitutional investors.

Positive long-term effects of institutional investors for financial stability …

… with benefits in emerging markets …

… but a need to educate households

The working group also highlighted some policy implications of the growing importance of institutional investors. Recent regulatory and accounting reforms seem likely on balance to enhance the functioning and stability of the financial system and contribute to a more efficient allocation of resources. They should encourage better risk management by institutional investors, the spreading of investment risk among a larger investor base and improved transparency in corporate accounts. In the case of emerging markets, the growing demand from global institutional investors for emerging market assets is likely to be positive for these economies, and should contribute to the depth of local financial markets. The growing role of global investors in emerging markets might nevertheless alter the transmission mechanism of domestic monetary policy, especially if long-term bond yields become more dependent upon global factors. However, with the shift in the pensions sector from defined benefit to defined contribution plans, and in the insurance sector from guaranteed to unit-linked products, the household sector has become increasingly exposed to financial markets, and prospective retirement income more subject to financial market volatility.

Potential short-term distortions

While the reforms appear beneficial for financial stability in the long term, the implementation of these measures may temporarily distort prices in financial markets, eg through feedback effects with the potential to drive long-term interest rates below the levels justified by macro fundamentals. Therefore, during the transition to the new regimes, policymakers will need to take into account the risk of triggering unnecessary market volatility or distorted valuations. The increased interest in alternative investment strategies (of still limited importance in portfolios) on the part of traditional institutional investors was not perceived as a major problem for financial stability. The working group recognised the potential of regulatory policy changes to increase the transparency of the existing links and channels of risk transfer between banks and institutional investors. In addition, such changes may help to provide increased transparency on the nature and location of the risks facing financial conglomerates, reducing the opacity that has existed in the past within complex financial institutions.

Limitations and challenges

Finally, the working group encountered various limitations and challenges in using balance sheet data to study the investment behaviour of institutional investors. In particular, balance sheet data do not reflect accurately the risk exposures of institutional investors that are significant users of derivatives. If the CGFS wants to continue monitoring how pension funds and insurance companies are responding to regulatory policy changes, it needs to consider in more detail how to improve the information and/or analytical frameworks for assessing financial stability issues.

Committee on Payment and Settlement Systems

The CPSS published its final report[1] on general principles for international remittance services (prepared jointly with the World Bank) in January 2007 and updated its statistics on payments and settlement systems in CPSS countries in November 2006. The CPSS-World Bank report, published on 23 January, provides an analysis of the payment system aspects of remittances, on the basis of which it sets out general principles designed to assist countries that want to improve the market for remittance services.

CPSS-World Bank report analyses the payment system aspects of international remittances …

The flow of funds from migrant workers back to their families in their home country is an important source of income in many developing economies. The total value of these remittances has been increasing steadily over the past decade (estimated in 2005 at over $230 billion equivalent, involving some 175 million migrants). However, the related money transfers are sometimes difficult to make, particularly for low-income workers, because of the costs and logistics involved. Although in recent years a number of reports have been prepared by various organisations on the topic of international remittances, few have been devoted specifically to the practical realities of how the money is transferred.

… and sets out general principles to assist policymakers in improving this market

The report contains five general principles covering: transparency and consumer protection; payment system infrastructure; the legal and regulatory framework; market structure and competition; and governance and risk management. The report also highlights the roles of both public authorities and remittance service providers in implementing the general principles.

The *statistics on payment and settlement systems,* released on 21 November, are part of an annual publication that provides data on payments and payment systems in the CPSS countries. This preliminary release contains

[1] A version of the report had been released for comment in March 2006; see “Recent initiatives by Basel-based committees and the Financial Stability Forum”, *BIS Quarterly Review*, June 2006.

individual country data (partial) and cross-country comparisons for 2005 and earlier years. The CPSS intends to publish a revised version in March 2007.