Derivatives and asset price volatility: a test using variance ratios

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Introduction

Theories as to how organised futures and options markets affect underlying cash markets tend to fall into one of the following three general categories. Firstly, it has been said that the presence of derivatives markets can at times cause sharp price movements in the underlying market that are unrelated to price discovery. The channels that have been proposed for this *"excess volatility"* are many. Some of the more plausible channels are: low margin requirements, which, because they permit market participants to take heavily leveraged positions, may lead to liquidity-related selling at times of large price swings; the ease of short selling in the futures market, which may accelerate price swings as short positions are covered; and "dynamic hedging", the practice whereby market participants, in order to maintain a prescribed price-sensitivity for portfolios containing options, adjust their cash market positions in ways that may reinforce large price swings.[1]

Secondly, it has been suggested, often in response to the excess volatility argument, that derivatives markets in fact add *stability* to cash markets. This could be because hedged participants are less likely to panic and sell into a down market. The above-mentioned ease of taking leveraged positions, both long and short, could also enhance price stability, if it allows informed traders to provide additional liquidity at short notice in support of a price that is threatened by excessive buying or selling pressure on the part of uninformed traders.

Finally, and also because of lower costs of taking positions, it has been suggested, by Cox (1976) among others, that *new information* is incorporated more quickly by derivatives markets than by the cash markets. As a result, the cash price may itself adjust more quickly to new information than it would have done had the futures market been absent.

This paper attempts to assess the presence and relative importance of these three hypothesised effects of organised derivatives markets on cash markets. An examination of the implications of the three hypotheses shows that one cannot distinguish among them merely by looking at a conventional measure of volatility, such as the variance of daily price changes. One would need some way of comparing an observed volatility level with changes in determinants of an asset's fundamental value. If fundamentals are themselves volatile, then an efficient market should reflect this volatility.

Previous studies of the effect of derivatives on cash market volatility have used a variety of techniques to put changes in cash market volatility into a meaningful context. Some, such as Figlewski (1981), Simpson and Ireland (1985) and Esposito and Giraldi (1994), have compared volatility in a market for which a futures market exists with volatility in a related market for which no futures market exists. Yet as Edwards (1988) points out, arbitrage ensures that the related market should be just as sensitive to price movement "spillovers" from the futures market as is the underlying

1 Dynamic hedging can involve options dealers hedging their options with cash market positions, or cash market investors hedging their positions with out-of-the-money options. As an example of the latter, Gennotte and Leland (1990) demonstrate how the use of portfolio insurance and similar strategies (which are greatly facilitated by the availability of standardised options) can lead to market crashes in conditions where a large number of traders are only partly informed about the sources of buying and selling pressure. This result holds even if the number of traders practising these strategies is very small relative to the size of the market.

market. Others have examined the effect of the introduction of derivatives on the parameters of a structural model of cash market volatility. Bortz (1984) regresses volatility in the underlying market on variables such as inflation and money supply growth as well as on volatility in related markets. Antoniou and Holmes (1995) model the volatility process over time using a generalised autoregressive conditional heteroskedasticity (GARCH) specification. In all of these studies, the presumption is that there is a component of volatility that is affected, for good or ill, by derivatives, and a component that is in some way inherent to the market in question.

The present paper may be thought of as using the variance of multi-day price changes - movements that "would have happened eventually" - as a proxy for this information-based component of volatility. The effects of derivatives are then sought by comparing price volatility at daily and multi-day time horizons before and after the introduction of exchange-traded derivatives markets. Such comparisons are informative because our three hypotheses have distinct implications, not only for levels of volatility but also for the relationship that should hold among volatility levels measured at these different horizons.

If futures and options markets create "excessive" price turbulence, then one would expect the introduction of these markets to be accompanied by an increase in short-term price volatility. If these short-term price movements are indeed spurious, however, then they should be reversed in the longer term, resulting in a price process that is mean-reverting. Derivatives markets should therefore cause the ratio of long to short-term variance to fall *below* the scaled relationship characteristic of a random walk, according to which the variance of k-period changes should be k times the variance of one-period changes.

If derivatives markets add stability to cash markets, then one would expect short-term price volatility to fall. The ratio of long to short-term price variance should approach that of a random walk, from a starting-point that is below that level. In other words, the stability hypothesis assumes that cash markets lacking derivative counterparts are initially excessively volatile in the short term, and that the presence of derivatives removes that component of volatility not related to new information about fundamentals.

If derivatives markets facilitate price discovery, then one may well see a higher volatility of short-term price movements as prices start to react more "sharply" to new information. The volatility of longer-term movements, however, should be relatively unaffected, because one assumes that the new information would eventually have been absorbed over the longer term without the benefit of derivatives. As a result, the ratio of long-term to short-term variance should fall, but only as far as the random walk level - a lower bound reflecting conditions of immediate, complete absorption of new information. Whereas the stability hypothesis predicts a variance ratio rising to random walk level from below, the information hypothesis predicts a ratio that falls to that level from above.

Another way of stating the implications of these three hypotheses is in terms of the serial correlation of price movements. If it takes several days for a given piece of information to be incorporated into an asset price, then the volatility of one-day changes will be low, while successive price movements will be positively correlated with one another. The information hypothesis predicts that autocorrelations should decline from these positive levels with the introduction of derivatives. In other words, the "persistence" of price movements should fall. Conversely, if prices jump erratically in the absence of meaningful new information, and if such jumps are consistently reversed over one or several subsequent periods, then successive price movements will be negatively correlated. The stability hypothesis thus predicts that, because options eliminate such jumpiness, autocorrelations should rise, having previously been negative. The excess volatility hypothesis predicts an increased frequency of such jumps, as a result of which autocorrelations should become negative (or more negative than they had been previously). As is shown in the next section, variance ratios can serve as indicators of serial correlation: positive autocorrelations imply variance ratios above one, and negative autocorrelations imply variance ratios below one.

This paper presents tests of the hypotheses discussed above for five financial price series: yields on long-term government bonds in the United States, Germany and Japan, and equity price indices in the United States and Germany. Variances and variance ratios are calculated over time periods preceding and following the introduction of exchange-traded futures and options contracts.[2] It is found that the introduction of these contracts is accompanied in many cases by a higher volatility of short-term price changes and a lower ratio of the variances of multi-day to daily returns. However, the ratio tends to fall only as far as a level exceeding or statistically indistinguishable from the level that would accept the random walk hypothesis. This finding is confirmed by other tests showing the serial correlation of daily returns to have fallen substantially from the pre-introduction to the post-introduction period. This would support the third of the hypothesised effects of derivatives markets on cash markets outlined above, namely that derivatives markets increase market efficiency by facilitating the rapid absorption of new information into prices. The only market which does not offer evidence for these effects is that for long-term Japanese government bonds.

This approach is similar to that of Brorsen (1991), who computes autocorrelation statistics and compares daily and multi-day variances in a test of the information-adjustment hypothesis. He finds sharp declines in measures of the autocorrelation of daily log changes in the Standard and Poor's 500 stock index after the introduction of futures trading on that index in April 1982. He further finds increases in the variance of daily price movements and less or no change in variances of movements over longer periods. I extend his work by computing the variance ratios and their standard errors directly and by examining a broader range of markets over a somewhat longer period of time.

The results also accord with those of Cox (1976) and Antoniou and Holmes (1995). Cox finds significant declines in the serial correlation of a number of commodity price series when futures markets are introduced. Antoniou and Holmes model daily changes in the FT 500 stock index before and after the introduction of the FT-SE 100 futures contract in 1984. They find that, while measured volatility in the cash market did increase, the GARCH parameters suggest less persistence in the effects of shocks to volatility. Antoniou and Holmes interpret this as evidence that "news" was incorporated into cash prices more quickly in the presence of the futures market.

The next section examines the motivation for and characteristics of variance ratio tests and reviews their use in previous studies. Section 2 applies variance ratio tests to the five series mentioned and corroborates the results by applying other tests of serial correlation to the same data. The final section concludes.

1. Variance ratio tests[3]

Consider a time series X_t, from which $n+1$ observations, $X_0 \ldots X_n$, are taken. Suppose that X_t follows a random walk with drift, as follows:

$$X_t = \mu + X_{t-1} + \varepsilon_t$$
$$\varepsilon_t \sim N(0,\sigma_0^2), \text{ i.i.d.}$$

Because the error terms are not serially correlated, the following two estimators are both consistent (if biased) for the true variance of daily changes, σ_0^2, under the null:

2 Over-the-counter markets for most of these derivative securities did exist before exchange-traded versions emerged. However, it is generally acknowledged that the availability of standardised, liquid instruments contributes greatly to their widespread use by market participants.

3 The discussion in this section closely follows that in Lo and McKinlay (1988).

$$\hat{\sigma}_a^2 = \frac{1}{n}\sum_{i=1}^{n}(X_i - X_{i-1} - \hat{\mu})^2$$

$$\hat{\sigma}_b^2 = \frac{1}{n}\sum_{i=1}^{n/k}(X_{ki} - X_{ki-k} - k\hat{\mu})^2,$$

where $\hat{\mu} = \frac{1}{n}\sum_{j=1}^{n}(X_j - X_{j-1}) = \frac{1}{n}(X_n - X_0)$. The first estimator is the sample variance of one-period changes. The second estimator is the sample variance of non-overlapping k-period changes, divided by k.

It is easy to see that these two estimators will be equal if the autocovariances of ΔX_t are all zero. For example, when k=2,

$$\hat{\sigma}_b^2 = \frac{1}{n}\sum_{i=1}^{n/2}((X_{2i} - X_{2i-1} - \hat{\mu}) + (X_{2i-1} - X_{2i-2} - \hat{\mu}))^2$$

$$= \hat{\sigma}_a^2 + \frac{2}{n}\sum_{i=1}^{n/2}((X_{2i} - X_{2i-1} - \hat{\mu})(X_{2i-1} - X_{2i-2} - \hat{\mu})).$$

If the series is a random walk, then both estimators are consistent, while only the first is efficient. Specifically, the asymptotic variance of the first estimator is $2\sigma_0^4/n$ and that of the second is $2k\sigma_0^4/n$. We can thus use the result of Hausman (1978) to express the asymptotic distribution of their difference as:

$$\sqrt{n}(\hat{\sigma}_b^2 - \hat{\sigma}_a^2) \xrightarrow{d} N(0, 2(k-1)\sigma_0^4).$$

Even more conveniently, we can derive the asymptotic distribution of the *ratio* between the two estimators:

$$\sqrt{n}\left(\frac{\hat{\sigma}_b^2}{\hat{\sigma}_a^2} - 1\right) \xrightarrow{d} N(0, 2(k-1)).$$

This permits us to use the variance ratio to test the null hypothesis that the series is a random walk. Lo and McKinlay (1988) construct a more powerful test along these lines using unbiased estimators of the two variances and overlapping observations to construct the multi-period variance estimate. They show that, for the one-period unbiased estimator $\hat{\sigma}_c^2 = \left(\frac{1}{n-1}\right)\sum_{i=1}^{n}(X_i - X_{i-1} - \hat{\mu})^2$ and the overlapping multi-period unbiased estimator $\hat{\sigma}_d^2 = \left(\frac{n}{k(n-k-1)(n-k)}\right)\sum_{i=k}^{n}(X_i - X_{i-k} - k\hat{\mu})^2$, the null hypothesis implies that:

$$\sqrt{n}\left(\frac{\hat{\sigma}_d^2}{\hat{\sigma}_c^2}-1\right)\xrightarrow{d} N\left(0,\frac{2(2k-1)(k-1)}{3k}\right).$$

Lo and McKinlay also show that this variance ratio approximately equals a declining linear combination of the first k estimated autocorrelations of the first differences. Specifically:

$$\frac{\hat{\sigma}_d^2}{\hat{\sigma}_c^2} \approx 1+\left(\frac{2}{k}\right)\sum_{i=1}^{k-1}(k-i)\hat{\rho}_i,$$

where $\hat{\rho}_i$ is the i-th autocorrelation of ΔX_t. This illustrates the fact that if the autocorrelations are generally positive, the variance ratio will be above one, while if they are generally negative it will be below one. The reader is referred to the Lo and McKinlay paper for the derivation of these results.

Variance ratios can thus be used as indicators of the persistence of the effects of one-time shocks to a series. Higher levels of ρ_i generally mean higher variance ratios. Alternatively, if a positive shock at time t leads to higher levels of ΔX_t for the following six periods, so that ρ_i is positive for i=1, 2, ..., 6, this should mean a higher variance ratio than if the autocorrelations are positive for only three periods.

Using standard errors derived from the above results,[4] Lo and McKinlay reject the hypothesis that weekly levels of the CRSP equally-weighted and value-weighted indices of US stock prices follow a random walk. Instead, they find variance ratios that are significantly greater than one and increase with k. The value-weighted index does not reject a random walk as consistently as does the equally-weighted index, and portfolios of high market-value firms do not reject as consistently as do portfolios of smaller firms. This confirms the common finding that trading in large (high market capitalisation) stocks is, by various definitions, more efficient than trading in small stocks.

Poterba and Summers (1988) examine variance ratios of stock returns over longer periods. They find positive serial correlation for one-month returns when their variances are compared with those of twelve-month returns, but negative correlation for twelve-month returns when these are compared with multi-year returns.

Variance ratio tests have also been applied to macroeconomic data. Campbell and Mankiw (1987) use a variance ratio test, among others, in an attempt to determine whether the quarterly GNP process is a random walk or is mean-reverting. Cochrane (1988) uses a variance ratio to measure the quantitative importance of permanent shocks to GNP (the "random walk component") relative to temporary shocks (the "stationary component"). He employs the fact that, if the true process for X_t is stationary (or stationary around a trend), then the variance ratio of the detrended series should go to zero for large values of k. If it does not go to zero, then the value that it "settles down to" is a reasonable indicator of the importance of shocks that, in economic terms, are essentially permanent, such as the shocks that cause GNP to depart from its trend even after twenty-five or thirty years.

Of course, applications of variance ratios and similar tests to macroeconomic data, and interpretation of the results, will differ from applications to financial data, because different economic hypotheses are of interest. For our purposes, the variance ratio test is useful because we are interested not only in testing for market efficiency, but also in analysing changes over time in the multi-period and single-period variances themselves.

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4 Actually, Lo and McKinlay modify their standard errors to take account of possible heteroskedasticity in the underlying series. A later version of the present study will make similar adjustments.

2. Results

Table 1 lists the securities to be analysed and the dates on which the first exchange-traded derivatives contracts related to those securities began to be traded. The tests in this section will investigate, for each price series, whether a significant change in autocorrelation patterns is likely to have occurred on either or both of the corresponding dates. Because the thirty-year US Treasury bond, on which the Treasury bond futures and options contracts are based, was itself only issued for the first time in 1977, the yield on ten-year US Treasury notes is used instead.[5] Yields on ten-year bonds are more appropriate for the German and Japanese cases, where the derivatives contracts are specifically linked to them. The fact that the introduction dates for the corresponding derivatives contracts are spread over a long period of time - from 1977 (US Treasury bond futures) to 1991 (DAX options) - reduces the likelihood that the results derive from contemporaneous changes in market structure that affected markets worldwide but were unrelated to derivatives.

Table 1

Introduction of trading in derivative contracts

Underlying security	Exchange	Futures	Options
US Treasury bond	CBOT/CBOE	22nd August 1977	1st October 1982
German federal bond	LIFFE	29th September 1988	20th April 1989
Japanese govt. bond	TSE	19th October 1985	11th May 1990
S & P 500 Index	CME	21st April 1982	28th January 1983
DAX Index	DTB	23rd November 1990	16th August 1991

Ideally one would want to know the effect of the presence of derivatives on asset returns. For the equity series, the log change in the level of the index serves this purpose well. For the bond series, one would want to know the change in the price of the underlying bond.[6] Lacking such prices, the yield series was used to construct approximate log price changes by means of the following formula:

$$\Delta \ln P \approx \frac{-D\Delta y}{1+y},$$

where D is the bond's Macaulay duration, assuming a maturity of ten years and a coupon rate and discount rate both equal to the yield, y.

Standard deviations of these log price changes for the five series, calculated over time intervals of one, two, five, ten and twenty days but with the multi-day standard deviations "normalised" (divided by the square root of the time interval) to make them comparable with the one-day statistic, are presented in Table 2. In the notation of the previous section, this table shows σ_c in the first column and σ_d in the remaining four columns. In all of the series except the German bonds, the one-day standard deviation (σ_c) rose with the introduction of futures and fell with the introduction

5 The correlation of daily changes in the two series over the period February 1977-October 1995 is 0.94. Contracts on Treasury notes have been in existence since the early 1980s, but the Treasury bond contract is used because it was introduced first.

6 If one were looking at returns over periods lasting several months or years - instead of the maximum twenty trading days considered here - one would also have to take account of changes in, respectively, dividends and interest received.

of options. In all except the German equity index, the one-day standard deviation with both options and futures present was higher than its level with both absent. For example, the standard deviation of daily changes in log prices of ten-year US Treasury notes was 0.29% from 1970 to August 1977, jumped to 0.69% during the period from the introduction of the Treasury bond future to the introduction of the option on that future in September 1982, and fell to 0.49% from September 1982 to the present, a decline from 1977-82 but higher than the pre-1977 level.

Table 2a

Standard deviations of log price changes: US 10-year Treasury notes

	Time period	Interval				
		1 day	2 days	5 days	10 days	20 days
Overall	2.1.70-30.6.95	48.89	51.17	53.41	55.35	58.39
Pre-futures	2.1.70-19.8.77	29.00	32.15	35.90	38.76	42.51
Post-futures, pre-options	22.8.77-30.9.82	68.56	71.07	73.83	78.08	80.01
Post-options	1.10.82-30.6.95	48.93	50.90	52.31	52.46	55.91

Table 2b

Standard deviations of log price changes: German 10-year federal bonds

	Time period	Interval				
		1 day	2 days	5 days	10 days	20 days
Overall	5.7.83-30.6.95	29.60	31.45	33.30	34.82	37.31
Pre-futures	5.7.83-28.9.88	28.36	31.01	32.17	33.65	36.08
Post-futures, pre-options	29.9.88-19.4.89	19.29	22.32	23.84	22.87	26.90
Post-options	20.4.89-30.6.95	31.30	32.80	35.16	36.63	39.26

Table 2c

Standard deviations of log price changes: Japanese 10-year government bonds

	Time period	Interval				
		1 day	2 days	5 days	10 days	20 days
Overall	29.10.84-30.6.95	58.13	59.12	61.17	64.93	70.46
Pre-futures	29.10.84-18.10.85	38.22	42.14	40.78	43.61	45.41
Post-futures, pre-options	19.10.85-10.5.90	75.38	75.43	78.37	83.14	89.30
Post-options	11.5.90-30.6.95	41.37	43.37	44.76	47.45	53.19

Table 2d

Standard deviations of log price changes: Standard & Poor's 500 Index

	Time period	Interval				
		1 day	2 days	5 days	10 days	20 days
Overall	2.1.70-30.6.95	93.89	99.14	99.67	98.79	98.28
Pre-futures	2.1.70-20.4.82	87.50	96.51	101.16	100.73	101.78
Post-futures, pre-options	21.4.82-27.1.83	122.15	127.65	142.85	146.81	143.82
Post-options	28.1.83-30.6.95	97.85	99.67	95.00	93.34	91.67

Table 2e

Standard deviations of log price changes: Deutsche Aktienindex (DAX)

	Time period	Interval				
		1 day	2 days	5 days	10 days	20 days
Overall	2.1.70-30.6.95	104.78	107.47	106.73	107.19	112.10
Pre-futures	2.1.70-22.11.90	105.19	108.14	107.51	108.73	114.59
Post-futures, pre-options	23.11.90-15.8.91	130.72	131.08	125.13	126.77	129.53
Post-options	16.8.91-30.6.95	97.05	98.95	99.05	95.16	95.68

Notes: 1. Figures have been multiplied by 10,000 for clarity.
2. Bond market figures are log price changes, calculated from yields, under the assumption that each is a par bond with a ten-year maturity and annual coupon payments equal to the day's yield.
3. Equity index figures are log changes in the corresponding index.
4. Each multi-day figure is the standard deviation of log changes over that interval, adjusted for the bias induced by using overlapping intervals and divided by the square root of the length of the interval in days.

On this somewhat crude basis, one might have grounds for rejecting at least part of the "stability" hypothesis outright: the introduction of a futures market does not seem to enhance the ability of informed speculators to counteract price swings, at least at first.[7] On the other hand, these figures would support the argument that the introduction of exchange-traded options to a market where futures are present reduces price swings.

At longer time intervals, the picture changes somewhat. At five-day intervals and above the three bonds see a rise in standard deviation from the pre-futures period to the post-options period, seemingly parallel to the rise in daily standard deviation. For example, the standard deviation of twenty-day changes in log prices of US Treasury bonds, divided by the square root of twenty for purposes of comparison, rose from 0.43% pre-futures to 0.56% post-options. For the two equity indices, on the other hand, ten and twenty-day volatilities in the post-options period are not only lower than their own pre-futures levels, but also lower than post-options one-day volatility, indicating that the positive autocorrelation seen in the pre-futures period has been replaced by a slight negative

7 Of course, the high volatility for the US ten-year note in 1977-82 probably has more to do with the Federal Reserve's switch from interest rate targeting to monetary targeting in the last three years of that period. The corresponding daily volatility for September 1977-September 1979 is 0.24 %.

autocorrelation. For example, the standard deviation of daily log changes in the S & P 500 was 0.88% over January 1970 to April 1982, and rose to 0.98% over the period January 1983-June 1995. The standard deviation of twenty-day changes in the index over the same two periods (as before, divided by the square root of twenty for comparability) fell from 1.02% to 0.92%.

The ratios of multi-day to daily variance are presented in Table 3. Below each figure is the z-statistic for a hypothesised value of one, calculated according to the asymptotic distribution discussed in the previous section. It is distributed standard normal under the null hypothesis that the series in question follows a random walk. The confidence band narrows sharply as the number of observations increases and as the length of the time interval for the multi-day variance falls; as a result, an identical variance ratio accepts a random walk in some instances and rejects it in others. For example, the post-futures, pre-options period for the German long bond is less than seven months long. This explains why a variance ratio of 1.41 does not reject the random walk over this period, while rejecting it in the five-year pre-futures period.

The variance ratio for the post-options period is less than that for the pre-futures period for sixteen of the twenty security/time interval combinations studied. The only exceptions to this rule are the five, ten and twenty-day variances of the Japanese bonds and the five-day variance of the German equity index, which rises only slightly. For fourteen of the twenty, the ratio falls when futures are introduced, and for twelve of the twenty it falls when options are introduced. The broad pre-futures to post-options increase in standard deviations apparent for bond prices in Table 2 is seen to have masked an equally broad decline in variance ratios: long-term volatility rose proportionately less than did short-term. For example, while the standard deviation of daily US ten-year Treasury note changes rose from 0.29% pre-futures to 0.49% post-options, the standard deviation of twenty-day changes rose less than proportionately, from 0.43% to 0.56%. Accordingly, the twenty-day to one-day ratio fell from 2.15 to 1.31.[8]

Table 3a

Ratios of variances of log price changes: US 10-year Treasury notes

	Time period	**Interval**			
		2 days/ 1 day	**5 days/ 1 day**	**10 days/ 1 day**	**20 days/ 1 day**
Overall	2.1.70-30.6.95	1.10* (7.46)	1.19* (6.90)	1.28* (6.52)	1.43* (6.70)
Pre-futures	2.1.70-19.8.77	1.23* (9.99)	1.53* (10.61)	1.79* (10.16)	2.15* (10.09)
Post-futures, pre-options	22.8.77-30.9.82	1.07* (2.66)	1.16* (2.60)	1.30* (3.14)	1.36* (2.60)
Post-options	1.10.82-30.6.95	1.08* (4.46)	1.14* (3.54)	1.15* (2.40)	1.31* (3.34)

8 Even if the policy of the Federal Reserve Board is responsible for the unusually high volatility of bond prices in our post-futures, pre-options period, the variance ratio over that period was sharply lower than in the pre-futures period. In other words, policy changes may have led to a greater frequency of shocks (new information), but the market absorbed these shocks more quickly.

Table 3b

Ratios of variances of log price changes: German 10-year federal bonds

	Time period	Interval			
		2 days/ 1 day	**5 days/ 1 day**	**10 days/ 1 day**	**20 days/ 1 day**
Overall	5.7.83-30.6.95	1.13* (6.79)	1.27* (6.37)	1.38* (5.97)	1.59* (6.22)
Pre-futures	5.7.83-28.9.88	1.20* (6.81)	1.29* (4.55)	1.41* (4.20)	1.62* (4.32)
Post-futures, pre-options	29.9.88-19.4.89	1.34* (3.80)	1.53* (2.70)	1.41 (1.35)	1.94* (2.13)
Post-options	20.4.89-30.6.95	1.10* (3.69)	1.26* (4.51)	1.37* (4.14)	1.57* (4.35)

Table 3c

Ratios of variances of log price changes: Japanese 10-year government bonds

	Time period	Interval			
		2 days/ 1 day	**5 days/ 1 day**	**10 days/ 1 day**	**20 days/ 1 day**
Overall	29.10.84-30.6.95	1.03 (1.69)	1.11* (2.41)	1.25* (3.62)	1.47* (4.66)
Pre-futures	29.10.84-18.10.85	1.22* (3.17)	1.14 (0.93)	1.30 (1.32)	1.41 (1.22)
Post-futures, pre-options	19.10.85-10.5.90	1.00 (0.05)	1.08 (1.19)	1.22* (2.06)	1.40* (2.61)
Post-options	11.5.90-30.6.95	1.10* (3.41)	1.17* (2.67)	1.32* (3.21)	1.65* (4.52)

Table 3d

Ratios of variances of log price changes: Standard & Poor's 500 Index

	Time period	Interval			
		2 days/ 1 day	**5 days/ 1 day**	**10 days/ 1 day**	**20 days/ 1 day**
Overall	2.1.70-30.6.95	1.11* (9.22)	1.13* (4.65)	1.11* (2.54)	1.10 (1.54)
Pre-futures	2.1.70-20.4.82	1.22* (12.07)	1.34* (8.56)	1.33* (5.37)	1.35* (3.96)
Post-futures, pre-options	21.4.82-27.1.83	1.09 (1.29)	1.37* (2.36)	1.44 (1.85)	1.39 (1.09)
Post-options	28.1.83-30.6.95	1.04* (2.10)	0.94 (- 1.49)	0.91 (- 1.50)	0.88 (- 1.38)

Table 3e

Ratios of variances of log price changes: Deutsche Aktienindex (DAX)

	Time period	Interval			
		2 days/ 1 day	**5 days/ 1 day**	**10 days/ 1 day**	**20 days/ 1 day**
Overall	2.1.70-30.6.95	1.05* (4.16)	1.04 (1.36)	1.05 (1.10)	1.14* (2.32)
Pre-futures	2.1.70-22.11.90	1.06* (4.11)	1.04 (1.47)	1.07 (1.46)	1.19* (2.71)
Post-futures, pre-options	23.11.90-15.8.91	1.01 (0.07)	0.92 (- 0.51)	0.94 (- 0.24)	0.98 (- 0.05)
Post-options	16.8.91-30.6.95	1.04 (1.24)	1.04 (0.59)	0.96 (- 0.36)	0.97 (- 0.18)

Notes: 1. An asterisk (*) indicates that the ratio is significantly different from one at the 95% level.
2. z-statistics for the hypothesis that the ratio equals one are in parentheses.
3. Values differ slightly from squared ratios of corresponding Table 2 values because of rounding.

In no case does the variance ratio fall to a level that would be conclusive evidence for the existence of negative serial correlation, or "excess volatility". The post-options ratio remains above one for all three bond series. For example, the volatility of US Treasury notes over twenty-day periods is 31% higher than it would be were one to extrapolate from daily volatility. The post-options ratio is below one for the two equity indices at several time intervals, but in no case is the difference from one statistically significant. In fact, a random walk cannot be rejected for either equity index at any time interval in the presence of options, with the exception of two-day changes of the S & P 500. Even in that case, the figure of 1.04 is very small and is substantially below earlier levels.[9]

Derivatives seem to have reduced or eliminated positive serial correlation but not to have introduced negative serial correlation. There is no clear pattern as to whether futures or options markets contributed more to this process. It would in any case be difficult to attribute the falling ratios specifically to options or to futures, because the gap between the introduction of the two markets was probably too small (in three of the five cases, less than a year) for the use of futures to have become sufficiently routine before the introduction of options.

Table 4 reports results of several other tests of random-walk-related hypotheses applied to the same data. The first column contains results of augmented Dickey-Fuller tests for stationarity of the bond yields and of the logs of the equity series levels. In no case does the statistic, which is the t-statistic of the coefficient on X_{t-1} in a regression of ΔX_t on X_{t-1}, twenty lags of ΔX_t, a constant and a time trend, attain a level which would indicate 95% confidence in rejecting the null hypothesis that the series is non-stationary. This can be interpreted to mean that there is little ground for assuming the price series are not random walks at this relatively crude level, allowing us to focus on the first differences (the log returns). Identical tests using first differences (adjusted for duration in the case of the bonds), not reported here, reject non-stationarity over every time period examined, with the exception of two of the brief periods between the introductions of futures and options.

The second column of Table 4 reports the Ljung-Box Q-statistic, defined as:

$$Q = n(n+2)\sum_{i=1}^{20} \frac{\hat{\rho}_i^2}{n-i}.$$

Q is distributed as χ^2_{20} under the null hypothesis that all the autocorrelations are zero.[10] Table 4 shows that this hypothesis is rejected in sixteen out of twenty cases. However, the level of this statistic, which should serve as a rough indicator of the degree of autocorrelation over a twenty-day period, falls substantially from the pre-futures period to the post-options period in every case but that of the Japanese bonds.

The third and fourth columns of Table 4 attempt to capture the degree of autocorrelation more directly. The third column shows the coefficient on ΔX_{t-1} in a regression of ΔX_t on a constant and its first five lags. It might be objected that regressions such as this are too heavily affected by large price movements, such as those that occurred in many financial markets in October 1987, rather than revealing the extent to which autocorrelation occurs on a day-to-day level. To meet this objection, a dummy variable, UP_t, set equal to one if the bond or stock price has risen, is regressed on a constant and five lags using a logit specification. The coefficient on the first lag is reported in the fourth column of Table 4. The results in both the third and the fourth columns are broadly consistent with the results of the variance ratio and Ljung-Box tests: with the introduction of derivatives, autocorrelations fell substantially, remaining statistically significant in bond markets but, according to most specifications, falling to insignificance in stock markets.

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9 It is possible that this figure results from the Scholes and Williams (1977) "infrequent trading" effect, whereby the daily measured level of a broad stock index lags true market sentiment because not all stocks in the index are traded every day. However, it is hard to believe that this was the case for more than a very small number of the stocks in the S & P 500 in the 1980s and 1990s, when trading volumes rose dramatically.

10 Note the "family resemblance" between this statistic and the expression in the previous section of the variance ratio statistic as, approximately, a declining weighted sum of *unsquared* autocorrelations.

Table 4a

Autocorrelation tests of daily log price changes: US 10-year Treasury notes

	Time period	Augmented Dickey-Fuller	Ljung-Box	ρ(1) of 1st diffs.	ρ(1) of "UP"
Overall	2.1.70-30.6.95	- 1.60	96.25*	0.10* (7.24)	0.48* (9.19)
Pre-futures	2.1.70-19.8.77	- 2.88	162.79*	0.22* (9.64)	0.87* (8.92)
Post-futures, pre-options	22.8.77-30.9.82	- 1.94	38.85*	0.07* (2.61)	0.38* (3.32)
Post-options	1.10.82-30.6.95	- 1.99	41.66*	0.08* (4.33)	0.27* (3.63)

Table 4b

Autocorrelation tests of daily log price changes: German 10-year federal bonds

	Time period	Augmented Dickey-Fuller	Ljung-Box	ρ(1) of 1st diffs.	ρ(1) of "UP"
Overall	5.7.83-30.6.95	- 1.65	90.78*	0.12* (6.47)	0.54* (6.67)
Pre-futures	5.7.83-28.9.88	- 1.31	65.95*	0.19* (6.53)	0.75* (5.95)
Post-futures, pre-options	29.9.88-19.4.89	- 2.60	15.45	0.15 (1.59)	0.14* (0.37)
Post-options	20.4.89-30.6.95	- 1.96	48.83*	0.08* (3.06)	0.40* (3.64)

Table 4c

Autocorrelation tests of daily log price changes: Japanese 10-year government bonds

	Time period	Augmented Dickey-Fuller	Ljung-Box	ρ(1) of 1st diffs.	ρ(1) of "UP"
Overall	29.10.84-30.6.95	- 2.06	65.94*	0.03 (1.52)	0.24* (2.93)
Pre-futures	29.10.84-18.10.85	- 0.73	32.29*	0.30* (4.28)	0.58* (2.00)
Post-futures, pre-options	19.10.85-10.5.90	- 2.24	38.64*	0.00 (- 0.12)	- 0.11 (- 0.89)
Post-options	11.5.90-30.6.95	- 1.95	39.42*	0.10* (3.42)	0.45* (3.80)

Table 4d

Autocorrelation tests of daily log price changes: Standard & Poor's 500 Index

	Time period	Augmented Dickey-Fuller	Ljung-Box	ρ(1) of 1st diffs.	ρ(1) of "UP"
Overall	2.1.70-30.6.95	- 2.37	122.33*	0.12* (9.62)	0.33* (6.61)
Pre-futures	2.1.70-20.4.82	- 2.41	304.98*	0.23* (12.78)	0.64* (8.66)
Post-futures, pre-options	21.4.82-27.1.83	- 1.89	20.29	0.08 (1.08)	0.03 (0.09)
Post-options	28.1.83-30.6.95	- 2.97	54.01*	0.04* (2.13)	0.04 (0.57)

Table 4e

Autocorrelation tests of daily log price changes: Deutsche Aktienindex (DAX)

	Time period	Augmented Dickey-Fuller	Ljung-Box	ρ(1) of 1st diffs.	ρ(1) of "UP"
Overall	2.1.70-30.6.95	- 2.92	73.30*	0.05* (4.31)	0.32* (6.39)
Pre-futures	2.1.70-22.11.90	- 2.50	77.34*	0.06* (4.36)	0.40* (7.19)
Post-futures, pre-options	23.11.90-15.8.91	- 3.28	18.57	0.01 (0.08)	- 0.02 (- 0.06)
Post-options	16.8.91-30.6.95	- 2.01	28.20	0.05 (1.42)	0.03 (0.24)

Notes:
1. An asterisk (*) indicates that the null hypothesis is rejected at the 5% level.
2. Augmented Dickey-Fuller test: H(0) is that the coefficient from regressing the first difference on the lagged value equals zero. Test performed on yields (bonds) and log index levels (equities). Each regression included twenty lags, a constant and a trend term. The table shows the t-statistic on the lagged value, with significance levels according with MacKinnon critical values.
3. Ljung-Box statistic: H(0) is that the autocorrelations jointly equal zero. Twenty autocorrelations used.
4. Test performed on log price changes (bonds) and log index changes (equities). The statistic is distributed as chi-squared (20) under the null.
5. 1st diffs.: coefficient on the daily log change lagged one period, in a regression of the daily log change on a constant and five of its own lags. t-statistics for whether this coefficient equals zero are in parentheses.
6. UP*: "UP" equals one if the bond price or stock index rose that day. This column reports the coefficient on the first lagged value of UP, when UP is regressed on a constant and five lags using a logit specification. t-statistics for whether this coefficient equals zero are in parentheses.

Table 5 presents test statistics for tests of whether the time series parameters of our series changed with the introduction of derivatives. The second column reports the F-statistic for a Chow test applied to a regression of the daily log change on a constant and five lags. The parameters are found to have changed in a significant way with the introduction of futures in four of the five series, and with the introduction of options in three of the five. The third column reports the results of a similar test of the UP_t dummy variable. This variable was regressed on a constant, its first five lags, a dummy equalling one if an observation was in the second part of the sample, and interactions of the five lags with this dummy. The Wald test F-statistic, testing the hypothesis that the coefficients on the second-half dummy and the five interaction terms all equal zero, is reported in the third column of the table. This time the parameters changed three out of five times when futures were introduced, and four out of five times for options. The "best" results, in the sense that the coefficients changed according to both tests and both break-points, are achieved by US bonds and equities. Table 5 also shows, however, that it may be difficult to attribute the parameter changes to derivatives markets alone. The first day of 1980 and the first day of 1990 perform just as ably as valid break-points for the data according to both tests.

Table 5a

Tests of stability: US 10-year Treasury notes

	Break-point	**Chow test of 1st diffs.**	**Wald test of "UP"**
Introduction of futures	22.8.77	2.85*	6.60*
Introduction of options	1.10.82	4.16*	8.80*
1980	2.1.80	2.41*	10.81*
1990	2.1.90	2.46*	3.13*

Table 5b

Tests of stability: German 10-year federal bonds

	Break-point	**Chow test of 1st diffs.**	**Wald test of "UP"**
Introduction of futures	29.9.88	2.63*	0.59
Introduction of options	20.4.89	2.60*	0.51
1990	2.1.90	2.08	0.93

Table 5c

Tests of stability: Japanese 10-year government bonds

	Break-point	**Chow test of 1st diffs.**	**Wald test of "UP"**
Introduction of futures	19.10.85	1.81	1.26
Introduction of options	11.5.90	1.08	3.65*
1990	2.1.90	0.94	3.16*

Table 5d

Tests of stability: Standard & Poor's 500 Index

	Break-point	**Chow test of 1st diffs.**	**Wald test of "UP"**
Introduction of futures	21.4.82	10.63*	7.24*
Introduction of options	28.1.83	11.03*	7.60*
1980	2.1.80	13.42*	6.36*
1990	2.1.90	1.35	3.16*

Table 5e

Tests of stability: Deutsche Aktienindex (DAX)

	Break-point	**Chow test of 1st diffs.**	**Wald test of "UP"**
Introduction of futures	23.11.90	3.18*	3.58*
Introduction of options	16.8.91	1.49	2.28*
1980	2.1.80	4.48*	2.71*
1990	2.1.90	2.36*	3.49*

Notes: 1. An asterisk (*) indicates that the null hypothesis is rejected at the 5% level.
2. AR(5) of first differences: Chow test statistic for a change in coefficient values after the date in the first column.
3. AR(5) of "UP": The dummy variable "UP" was regressed on a constant, five lags and interactions of these six terms with dummy variables indicating that the observation occurred after the date in the first column. This column reports the Wald test statistic for the hypothesis that these six coefficients all equal zero.

Conclusion

Of the three hypotheses cited in the introduction, the evidence presented in this paper best seems to support the proposition that derivatives facilitate the incorporation of new information into security prices. The variances of changes in the security price series studied are generally higher after the introduction of exchange-traded derivatives markets than before, casting doubt (at a crude level) on the notion that derivatives make underlying markets more stable. The variance of daily changes tends to rise more than does the variance of multi-day changes. Ratios between variances at different intervals suggest, however, that price movements in the bond markets studied remain positively correlated (if less so than before), while movements of stock indices are indistinguishable from a random walk. This contradicts the hypothesis that derivatives add "excessive" volatility to underlying markets, since such a hypothesis would predict negative serial correlation. The reduction or elimination of positive serial correlation suggests that a given piece of news is now incorporated into securities prices much more quickly than before.

The Japanese bond market, as already noted, does not fit as neatly into this pattern. Part of the problem may be that only about one year of daily data are available preceding the introduction of futures. In any case, the *levels* of the Japanese statistics are broadly in line with those in other markets. The decline of the two-day to one-day ratio in Table 3, and declines in the two first-degree autocorrelation coefficients in Table 4, suggest that a given change in Japanese bond prices has indeed become less predictable on the basis of the previous day's change. Over weekly and monthly periods, however, trends may be persistent, so that one week's change is just as good a signal of the following week's change as it was before. Proving such a conjecture would require further study.

Further study is also needed to determine whether these changes in correlation patterns can indeed be ascribed to the presence of derivatives markets alone, or whether other contemporaneous factors were at work. Certainly it will be difficult to isolate particular changes in the institutional structure of these markets and study their individual effects. For example, sharp movements in some equity markets have been attributed to the effects of computerised "program trading", yet such trading itself developed to facilitate arbitrage between index futures prices and underlying prices. Program trading may thus be an example of how the presence of a derivatives market can spur the rapid incorporation of information into underlying prices.

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