The information content of implied volatilities of options on eurodeposit futures traded on the LIFFE: is there long memory?

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Asset price volatility is playing a growing role in portfolio risk management, option pricing and - more generally - in international financial economics. It tends to change over time and is usually difficult to estimate, and the task of deriving a reliable measure of the market's subjective assessment of future volatility of an asset is therefore of paramount importance. It is for this reason that attention has recently focused on analysis of the volatility deriving from option prices, reproducing previous investigation into the unbiased efficiency hypothesis of futures prices and forward exchange rates.

Previous empirical analyses had come up with contradictory results. Latané and Rendleman (1976), Schmalensee and Trippi (1978), Chiras and Manaster (1978) among others found that implied volatility outperforms historical volatility as a predictor of actual volatility. More recent analyses, however, show mixed results. Scott (1992), Day and Lewis (1992), Canina and Figlewski (1993), Lamoureux and Lastrapes (1993) and, more recently, Ané and Geman (1998) cast doubts on the superiority of implied volatility forecasts, while Scott and Tucker (1989), Xu and Taylor (1994, 1995), Jorion (1995), Siegel (1997), Campa and Chang (1995) and Walter and Lopez (2000) by contrast find that, in spite of their shortcomings, implied volatilities provide reliable forecasts, which cannot be improved upon with the help of additional information proxies derived from manipulation of the underlying asset prices. A common characteristic of these analyses, however, is that they are obtained using currency options (on spot and futures contracts). The liquidity and the homogeneity of these contracts may account for the greater accuracy of the corresponding implied volatilities.

This paper examines the behaviour of implied volatility from options on short-term (three month) interest rate future contracts in sterling, the Three Month Sterling interest rate future contract, and in deutschmark, the Three Month Eurodeutschmark interest rate future contract, traded at the LIFFE. These contracts play a significant role in both interest rate and exchange rate risk hedging, and are influenced by monetary and exchange rate policies. Short-term interest rate implied volatility is an indicator of the dispersion of expectations on future short-term interest rate behaviour and is positively correlated with uncertainty on future monetary policy measures. Financial and exchange rate turbulence – which affects monetary policy – influences volatility forecasts and their term structure.

Previous work by Neuhaus (1995), Bhundia and Chadha (1997) and Bahra (1998) is extended in two ways.

(i) The financial analysis is preceded by close examination of the statistical properties of the relevant time series, and it is these properties that justify the parameterisation adopted in the subsequent analysis. Efficiency analysis *à la* Mincer and Zarnowitz (1969) cannot be easily implemented with long-memory, fractionally integrated time series. A less ambitious approach relating current volatility either to current or to lagged implied volatility in order to assess, respectively, the information content and the relative predictive power of the latter seems to be more promising. It involves the use of GARCH models of the volatility of the return of the underlying.

(ii) Investigation is extended across contracts and is associated with the term structure of implied volatility from options with differing time to expiration. Information content analysis misses some relevant aspects of interest rate volatility forecasting. The relative efficiency of London traders in dealing with both a national and a "foreign" interest rate has not been explicitly assessed. It is not necessarily homogeneous, however, and may well produce a systematic and identifiable pattern of volatility transmission from one contract to the other as news hits the market.

The significance of the empirical investigation of daily implied volatilities is strongly affected by their peculiar time series properties. The choice of the model specification is of paramount importance as overdifferencing and/or underdifferencing biases may bring about totally different economic results from the same set of data. These specification problems could be solved using data sampled at longer time intervals. The monthly or even quarterly informational efficiency of option pricing, however, is of little interest to the financial analyst.

The paper is organised thus: the economic and financial aspects are set out in section one, together with the pitfalls of the estimation methodology; the statistical properties of the time series under investigation and their consequences for volatility modelling are set out in section two; the relative informational content and predictive power of volatility forecasts are analysed in section three; the reaction of implied volatilities to the arrival of news as reflected in the term structures and across contracts is analysed in section four, while section five presents the concluding discussion.

1. Accuracy and economic significance of implied volatility

1.1. Predictive power and information content

Tests of predictive power assess ex post the forecasting accuracy of implied volatility and are derived from the asset price efficiency analyses of the early eighties They are assumed to verify whether the market forecast is an unbiased and efficient predictor of the future dependent variable, in this case the volatility of the return on the underlying asset over the remaining life of the contract. They involve estimation of the following relationship *à la* Mincer and Zarnowitz,

$$\sigma_{t,T} = a + b\sigma^{F}_{t,T} + u_{t,T} \qquad (1)$$

where $\sigma_{t,T}$ is the realised volatility between time t and T and $\sigma^{F}_{t,T}$ is the volatility forecast derived at t over the period from t to t+T. Quantifying $\sigma^{F}_{t,T}$ by IV(t,T), the implied volatility, we would obtain a zero intercept and a slope of one if the latter were to be an efficient and unbiased forecast of future volatility.[1]

The predictive power of implied volatility is typically compared with that of alternative measures of volatility forecast, derived from past returns of the underlying contract. Jorion (1995) suggests two proxies; a moving average estimated over the previous 20 (trading) days of historical volatility and the conditional volatility provided by a GARCH parameterisation. A larger spectrum of alternative volatility proxies can be found in Ané and Geman (1998) and in Ap Gwilym and Buckle (1999).

If the market of interest is informationally efficient, then implied volatility, which incorporates all available information about future asset price behaviour, should be more accurate than the alternative forecast proxies and provide coefficient estimates closer to the canonical unbiased efficiency theoretical values. This result is also verified with the help of the following encompassing regression approach, originally set out by Chong and Hendry (1986) and Fair and Shiller (1990)

$$\sigma_{t,T} = a + b\, IV(t,T) + c\sigma^{P}_{t,T} + u_{t,T} \tag{2}$$

where $\sigma^{P}_{t,T}$ is a realised volatility forecast proxy based on past prices. Under the null of implied volatility informational efficiency, $\sigma^{P}_{t,T}$ should have no predictive power and c estimates should not be significantly different from zero.

An alternative testing strategy set out by Day and Lewis (1992), Lamoureux and Lastrapes (1993) and Amin and Ng (1997), among others, exploits the properties of GARCH parameterization of the volatility of the underlying asset and involves analysis of the following relationship

$$\sigma_t^2 = \sigma^2_{(garch)t} + \delta IV(t,T)^2 \tag{3}$$

Squared implied volatility is added as a regressor; its coefficient should be significantly different from zero and, conversely, the

GARCH terms should have no explanatory power if the market is informationally efficient and the option pricing model is valid. The difficulties associated with estimation and the long-memory properties of realised implied volatility are avoided. Implied volatility, however, refers to a longer horizon than one day, and the maturity mismatch affects interpretation of the results; the δ coefficient should be positive, but no *a priori* theoretical value can be attributed to it. The predictive power of implied volatility is assessed adding the lagged (squared) implied volatility to a GARCH variance equation and estimating

$$\sigma_t^2 = \sigma_{(garch)t}^2 + \delta' IV(t-1,T)^2 \tag{3'}$$

The statistical significance of the implied volatility coefficient is then an indicator of predictive power in addition to the historical forecasts provided by the GARCH parameter components.

1.2. Tests of the transmission of information over time and across assets

Several authors have used implied volatilities derived from option contracts with differing time to expiration to investigate the time profile of news influencing the price of the underlying. This analysis is *a priori* highly informative for portfolio managers as it could provide a measure of the feeling of the market on the future evolution of volatility. A shift in market mood due to a change in scenario would certainly be reflected in a shift in the relationship between short-term and long-term implied volatility quotes.

Implied volatilities on “distant” options are usually larger than those from “nearby” options, but this difference is not constant over time. An initial approach is simply to subtract from the implied volatility estimated for distant time to expiration the corresponding nearby volatility. The evolution over time of this index would provide a rough picture of shifts in the term structure of implied volatilities and of the market forecasts. It should be noted, however, that an increase in the volatility differential could be due to an increase in distant volatility relative to nearby volatility, to a relative decrease in nearby volatility, or to a combination of both.

Implied volatilities are generally observed to be mean reverting as volatility shocks tend to dissipate over time (even if the degree of persistence is a positive function of the time to maturity). If this is in fact the case, then when long-run volatility is high relative to its mean value short-run volatility should be yet higher. Indeed, in a rational expectations context long-term volatility should incorporate the currently higher short rate and future reversion towards the mean. In the same way, if nearby volatility lies below its mean value the distant implied volatility should be closer to its equilibrium value. Any alteration in this relationship would indicate a shift in market mood.

Attempts to provide a formal pattern of the relationship between implied volatilities derived from options with differing times to expiration have been presented by Stein (1989), and Campa and Chang (1995), among others. The aim of these studies is to derive *ex ante* testable relationships between short-run and long-run implied volatilities.

Stein (1989), using weekly data, develops a formal test of the term structure of implied volatilities under the joint null hypothesis of a correct specification of the dynamics of volatility of the price of the underlying asset – and of the option pricing model – and of market efficiency. Assuming that implied volatility is equal to the average expected volatility of the underlying over the remaining life of the option and that instantaneous volatility reverts at a constant rate to its constant mean value, he derives a theoretical relationship between the volatility implied by an option close to expiration and the volatility implied by an option on the same underlying asset that is distant from expiration.[2]

The reformulation of this relationship in terms of daily data set out by Diz and Finucane (1993) reads as follows

$$[IV(t,m_{2t}) - \overline{\sigma}\ ] = \beta(\rho)\ [\ IV(t,m_{1t}) - \overline{\sigma}\ ] \quad (4)$$

where

$$\beta(\rho) = \frac{m_{1t}[\rho^{m_{2t}} - 1]}{m_{2t}[\rho^{m_{1t}} - 1]} \quad (5)$$

$IV(t,m_{1t})$ is the implied volatility of a short maturity option at time t with m_{1t} days to maturity, $IV(t,m_{2t})$ is the implied volatility of a longer maturity option at t, with m_{2t} days left to maturity and $\overline{\sigma}$ and ρ are, respectively, the constant mean value of instantaneous volatility and the first order daily autocorrelation coefficient of short run implied volatility (which is assumed to quantify the unobservable autocorrelation path of instantaneous volatility). It is assumed that $m_{2t} = m_{1t} + \Delta d$ where Δd is a constant difference between the maturity of the two options. For $\rho < 1$, it can be shown that $\beta(\rho) < 1$. A shock to the short-run implied volatility will be associated with a smaller shift in the distant long-run implied volatility.[3]

Campa and Chang (1995), using foreign exchange options, compare squared volatilities quoted at different dates. They follow the strategy originally set out by Campbell and Shiller (1991) for interest rates and test whether the long-run and short-run implied volatilities quoted today are consistent with short-run volatility quoted in the future. Here, too, the theoretical argument is set out using the Hull-White stochastic volatility approach and at-the-money options. Moreover they explicitly model the bias associated with the corresponding Black-Scholes option pricing.[4]

The following testable expectations hypothesis involving squared implied volatilities obtained by inverting the Black-Scholes formula is derived with some algebraic manipulation and is assumed to hold over k time periods Q.

$$IV(0,kQ)^2 = \left(\frac{1}{k}\right) E_0\left[\sum_{i=0}^{k-1} IV(iQ,(i+1)Q)^2 \right] \tag{6}$$

The current long run squared volatility $IV(0,kQ)^2$ is equal to the average of the current and expected future squared short run volatilities $IV(iQ,(i+1)Q)^2$, i = 0, 1, …, k-1. Analysis of the transmission of news across contracts does not yield analogous formal relationships incorporating *ex ante* market efficiency and rational expectations hypotheses. Some recent results on the volatility interlinkages across international equity markets can, however, be used to derive simple testable efficiency hypotheses on the diffusion of news across implied volatilities from different option contracts.

As shown in the "meteor shower-heat wave" literature *à la* Engle et al. (1992) on the transmission of news across international equity and foreign exchange rate markets, we can distinguish between international and country-specific news. In the same asset market, there should be no causality hierarchy across assets denominated in different currencies, such as the Short Sterling, the 3-Month Euromark futures contract and the corresponding option contracts investigated above. They are subject to the same set of international news and, at the same time, country-specific (idiosyncratic) news which affects a national futures contract should not spillover to the other country's interest rate futures contract.

A Granger causality test was applied to a 2-equation VAR system involving the first differences of the implied volatilities. The detection of unilateral causality, i.e. of an international hierarchy in the diffusion of news via implied volatility changes, may indicate the presence of contagion, as defined by Masson (1998), among others. Contagion, in turn, may be due to market inefficiency and irrational (herding) behaviour. It is not, however, synonymous with such behaviour. Kodres and Pritsker (1999) have shown that, in the case of international equity markets, contagion-like behaviour may result from rational portfolio hedging policies.

2. Data, macroeconomic scenario and preliminary statistical analysis

End-of-day data on short-term interest rate derivatives traded in London, the Short Sterling, 3 Month Euromark futures and corresponding option contracts are provided by the LIFFE. These contracts are highly liquid and reflect international portfolio hedging requirements associated with interest rate and exchange rate volatility risk. Indeed, the time interval under investigation, from January 1, 1993 through December 31, 1997, encompasses periods of severe financial and exchange rate turbulence, such as the July - August 1993 French Franc crisis, the December 1994 - March 1995 Mexican crisis and the onset of the Asian crisis in the Summer of 1997.

Contract expiration follows the standard March, June, September and December cycle. For the sake of homogeneity, the auxiliary Euromark contracts introduced from June 1994 onwards are disregarded. Each contract lasts at least nine months. One trading week before expiration the series switches into the next contract in order to minimise the contract expiration biases identified by Day and Lewis (1988). Continuous time series of futures prices, option prices and corresponding at-the-money implied volatility (derived from the appropriate formula set out by Black (1976) and quoted by the LIFFE) are thus obtained.[5]

Realised "future" volatility, too, is painstakingly reproduced. For each option contract the short-term implied volatility is matched with the sequence of future standard deviations of the continuously compounded returns of the underlying futures contract until option expiration.[6] As usual, annualised volatility is obtained by multiplying realised volatility by the square root of 252, the number of trading days per year.

Following Canina and Figlewski (1993) and Jorion (1995), among others, the relative accuracy of implied volatility is assessed using as benchmark an alternative volatility forecast proxy. A twenty-trading-day moving average of standard deviations of past rates of return of the underlying contract seems to provide reliable results.

Estimates of implied volatilities with different time to expiration are drawn up with the help of close screening of the maturity of the option contracts. Short-term options, with a time to maturity between zero (in reality 6 trading days) and three months (a 63-trading-day interval) are used to build the short-term implied volatility time series. Medium-term and long-term implied volatility time series are derived from options with a time to expiration between three to six months and six to nine months respectively. In this way every trading day of the sample implied volatilities are listed coming from options that have always, respectively, 63 and 126 days longer to trade than the short-term options. Figures 1 and 2 reproduce their behaviour - measured in percent per annum - over the time period under investigation.

Non-homogeneous interest rate volatility is clearly discernible from the data and seems to be related with major financial and monetary events. The uncertainty associated with the Mexican crisis results in a generalised increase in volatility as US dollar depreciation and the ensuing international financial turbulence bring about an increase in the British and German official interest rates. The appreciation of the dollar from 1996 onwards eases the tension in monetary markets. A trend towards a decrease in volatility levels in the last two years of the sample seems to characterise both contracts. British and German monetary policies become less restrictive and the revitalisation of the European Monetary Union project results in a reduction in inflationary expectations, as also in short-term interest rates and overall European financial turbulence. In 1997 inflationary fears, promptly reflected in a

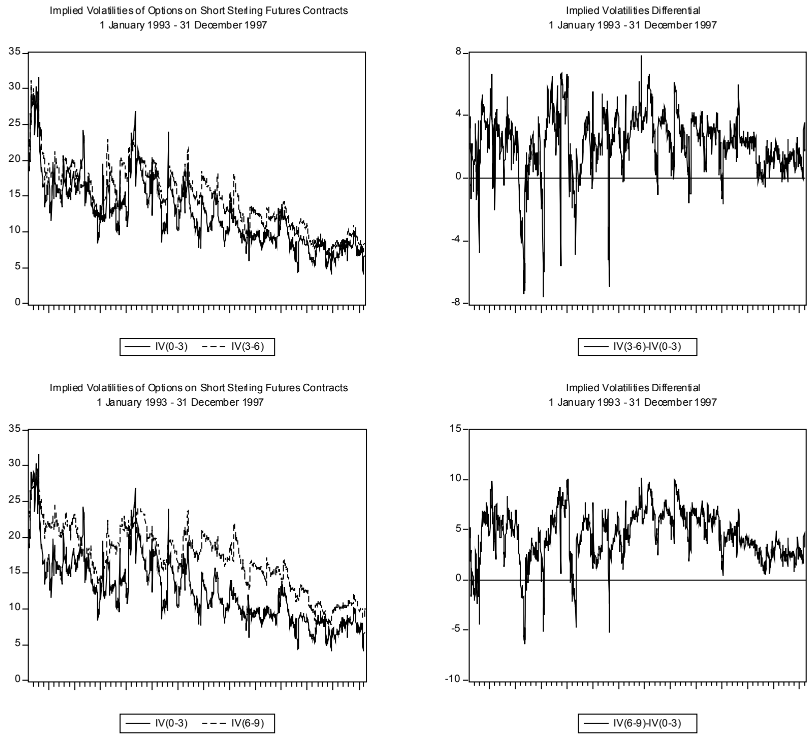

Figure 1

temporary increase in implied volatilities, motivate an increase in the official British and German interest rates. The Asian equity turbulence does not seem to spillover to short-term interest rates as volatility expectations tend to decline at the end of the year.

The paths of implied volatilities of options with different time to expiration provide additional information. The difference between short-term and long-term expectations is at times considerable, which points to a significant term structure, and the presence of crossovers suggests that the slope of the latter may change over time. The dynamics suggest a different degree of mean reversion across the term structure. Indeed, whenever short-term volatility rises, long-term

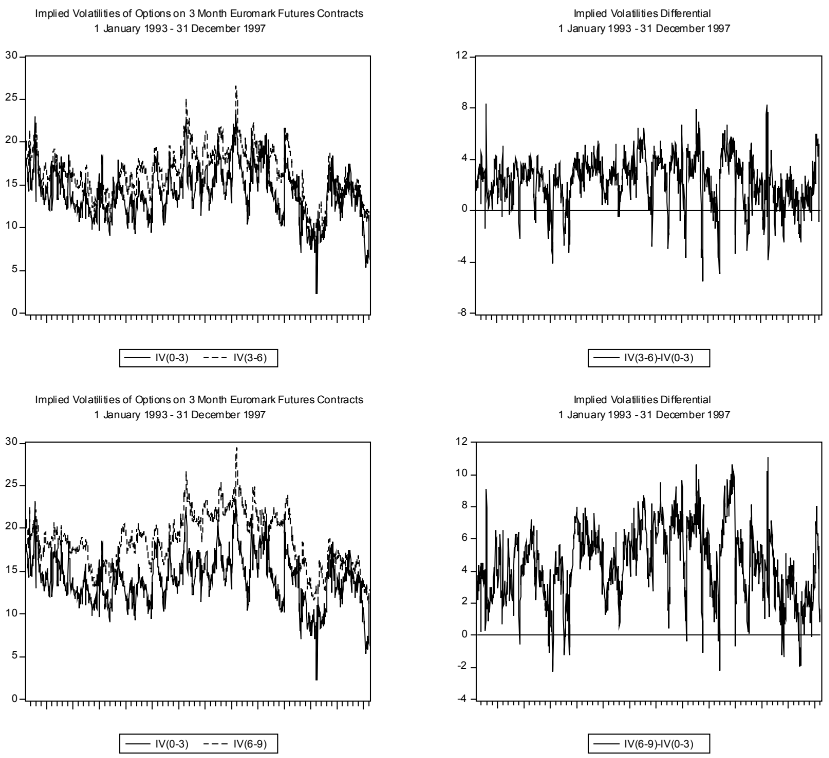

Figure 2

volatility – be it from 3 to 6 month or from 6 to 9 month to expiration contracts – rises too, but less than proportionally, and the distance between the two time series decreases. Conversely, when short-term

Table 1

1 January 1993-31 December 1997

Descriptive Statistics

	Mean	**S.D.**	**Sk.**	**Kurt.**				**A.C. P.A.C.**			
S.Ster-Ling					**1**	**2**	**3**	**4**	**5**	**10**	**20**
σ_t	0.097	0.194	9.698	127.9	0.208 0.208	0.066 0.024	0.037 0.019	0.049 0.038	0.016 -0.003	-0.008 -0.015	0.029 0.028
$\sigma_{t,T}$	0.219	0.202	2.567	11.93	0.966 0.966	0.935 0.038	0.904 -0.015	0.870 -0.070	0.839 0.023	0.715 0.010	0.549 0.004
IV (0-3)	0.128	0.049	0.925	3.804	0.919 0.919	0.887 0.133	0.861 0.020	0.838 0.017	0.817 0.045	0.759 0.018	0.680 -0.012
IV (3-6)	0.150	0.047	0.518	3.155	0.989 0.989	0.980 0.020	0.972 0.034	0.962 -0.036	0.953 -0.011	0.918 -0.031	0.850 -0.028
IV (6-9)	0.172	0.046	-0.023	2.438	0.993 0.993	0.987 0.067	0.980 -0.058	0.974 -0.048	0.968 0.057	0.944 -0.004	0.892 -0.017
Euro-mark											
σ_t	0.092	0.138	8.237	104.0	0.057 0.057	0.026 0.023	0.004 0.001	0.026 0.026	0.061 0.058	0.030 0.023	0.017 0.014
$\sigma_{t,T}$	0.178	0.114	2.198	9.056	0.957 0.957	0.920 0.049	0.886 0.030	0.855 0.029	0.828 0.023	0.717 0.011	0.570 0.012
IV (0-3)	0.142	0.030	0.101	3.898	0.880 0.880	0.817 0.186	0.803 0.214	0.761 -0.019	0.719 0.005	0.591 0.034	0.376 0.025
IV (3-6)	0.165	0.030	0.110	3.363	0.967 0.967	0.942 0.093	0.917 0.010	0.893 -0.009	0.867 -0.020	0.770 0.015	0.621 0.003
IV (6-9)	0.187	0.032	0.092	2.662	0.980 0.980	0.964 0.080	0.947 -0.038	0.931 0.002	0.916 0.038	0.858 -0.005	0.753 -0.026

Notes. S.D. : Standard Deviation; Sk. : Skewness; Kurt. : Kurtosis; A.C. : autocorrelation coefficient; P.A.C. : partial autocorrelation coefficient; σ_t : daily return volatility; $\sigma_{t,T}$: future realised volatility; IV(0-3), IV(3-6), IV(6-9): implied volatilities from options with, respectively, 0 to 3, 3 to 6 and 6 to 9 months to expiration.

implied volatility tends to decline, long-term volatility declines less rapidly and the distance between them tends to rise.

Table 1 presents preliminary statistics for daily volatilities, realised future volatilities and implied volatilities. Daily volatility seems to be affected by a significant first order autocorrelation, higher for the Short Sterling than for the Euromark futures contract. The coefficients of skewness and kurtosis are always very large and do not seem to be compatible with a Gaussian distribution. Future realised volatilities display lower skewness and kurtosis. Their autocorrelation functions remain large, positive and significant at very long lags. They suggest that the time series might have long memory, or even be non-stationary, and call for an accurate investigation of their properties across the frequency and time domains.

The value of the standardized spectral density at zero frequency provides useful information as it is positively correlated with the persistence of deviations from the trend (it is unbounded in the case of a unit root). In table 2 are set out the estimates of scaled and standardized spectral density functions of the levels and of the first differences of the relevant volatilities at various frequencies between 0 and π.[7]

A common characteristic of the estimates is that the spectral densities are concentrated at low frequencies; they are very large at zero frequency and drop rapidly afterwards.[8] These findings validate the hypothesis of a high degree of persistence, especially for the long-term contracts.

The zero frequency long run variances of the first differences corroborate these results; they suggest that implied volatility persistence *à la* Cochrane (1988) increases with the time to maturity of

Table 2

1 January 1993-31 December 1997

Standardized Spectral Density Estimates

Frequency $\omega_j = j\pi/m$ j	**0** **0**	**0.5236** **2**	**1.0472** **24**	**1.5708** **36**	**2.0944** **48**	**2.6180** **60**	**3.1416** **72**
S.Sterling							
$\sigma_{t,T}$	37.0470 (9.9833)	0.4089 (0.0779)	0.0844 (0.0161)	0.0487 (0.0093)	0.0403 (0.0077)	0.0290 (0.0057)	0.0245 (0.0066)
IV(0-3)	26.6241 (7.1745)	0.4918 (0.0937)	0.1830 (0.0348)	0.0873 (0.0166)	0.0784 (0.0149)	0.0481 (0.0092)	0.0526 (0.0142)
IV(3-6)	37.3468 (10.0641)	0.3223 (0.0614)	0.0678 (0.0125)	0.0596 (0.0113)	0.0363 (0.0069)	0.0352 (0.0067)	0.0239 (0.0064)
IV(6-9)	45.9848 (12.3918)	0.2553 (0.0486)	0.0596 (0.0114)	0.0293 (0.0056)	0.0276 (0.0053)	0.0217 (0.0041)	0.0216 (0.0058)
$\Delta\sigma_{t,T}$	0.3725 (0.1004)	1.0740 (0.2047)	0.7540 (0.1437)	0.9400 (0.1792)	1.2439 (0.2371)	1.1141 (0.2124)	0.9401 (0.2534)
ΔIV(0-3)	0.1664 (0.0449)	0.7094 (0.1352)	1.0123 (0.1930)	0.9454 (0.1802)	1.3951 (0.2659)	0.9699 (0.1849)	1.1902 (0.3208)
ΔIV(3-6)	0.2852 (0.0769)	0.9726 (0.1854)	0.5764 (0.1099)	1.2387 (0.2361)	1.3082 (0.2494)	1.6137 (0.3076)	0.9507 (0.2563)
ΔIV(6-9)	0.3480 (0.0938)	0.9731 (0.1855)	0.7817 (0.1490)	0.7932 (0.1512)	1.3059 (0.2489)	1.2000 (0.2287)	1.2682 (0.3419)
Euromark							
$\sigma_{t,T}$	34.9076 (9.4103)	0.5389 (0.1027)	0.1229 (0.0234)	0.0749 (0.0143)	0.0579 (0.0110)	0.0418 (0.0079)	0.0333 (0.0089)
IV(0-3)	29.7026 (8.0041)	0.5114 (0.0974)	0.2019 (0.0385)	0.1561 (0.0297)	0.1211 (0.0231)	0.0809 (0.0154)	0.0727 (0.0196)
IV(3-6)	45.4857 (12.2573)	0.3195 (0.0609)	0.0846 (0.0161)	0.0571 (0.0109)	0.0354 (0.0067)	0.0299 (0.0057)	0.0309 (0.0083)
IV(6-9)	53.3153 (14.3672)	0.2252 (0.0429)	0.0449 (0.0086)	0.0279 (0.0053)	0.0189 (0.0036)	0.0197 (0.0038)	0.0125 (0.0034)
$\Delta\sigma_{t,T}$	0.1575 (0.0425)	0.9535 (0.1818)	0.8364 (0.1594)	1.0725 (0.2044)	1.2422 (0.2368)	1.0895 (0.2077)	0.9102 (0.2454)
ΔIV(0-3)	0.1007 (0.0271)	0.4857 (0.0926)	0.7520 (0.1433)	1.2546 (0.2391)	1.4667 (0.2796)	1.2100 (0.2307)	1.2306 (0.3317)
ΔIV(3-6)	0.2699 (0.0727)	0.9070 (0.1729)	0.9536 (0.1818)	1.4112 (0.2690)	1.2478 (0.2379)	1.2362 (0.2356)	1.5992 (0.4311)
ΔIV(6-9)	0.3290 (0.0889)	1.0080 (0.1921)	0.6570 (0.1252)	1.0118 (0.1929)	1.0001 (0.1906)	1.2153 (0.2310)	0.8117 (0.2188)

Notes. $m = 72 = 2(T)^{0.5}$: bandwidth parameter. The standardized spectral density functions are estimated using the Bartlett kernel. Estimated asymptotic standard errors in parentheses.

the underlying contract.[9] They rise from 0.1664 and 0.1007 for the 0 to 3 month to expiration short Sterling and Euromark contracts (a low value, the time series are probably over-differenced) to 0.3480 and

0.3290 respectively for the 6- to 9- month ones. These estimates justify the hypothesis of a different degree of mean reversion across the term structure identified in the analysis of figures 1 and 2.

Standard unit root tests are reported in the upper half of table 3. They fail to provide homogeneous results. A clear-cut rejection of the null of a unit root is obtained for the 0- to 3- month to expiration option contracts only (and, in the case of the Euromark, also for the 3- to 6- month contract). The pricing of the three remaining contracts shows extreme dependence on the initial conditions (i.e. on the current state of the economy), which seems to contradict the observed pricing behaviour.

Stationary long-memory time series *à la* Granger and Joyeux (1980) have properties that are compatible with those of the volatilities under investigation; autocorrelations that decay slowly as the lags increase and unbounded spectrum at low frequency.[10] Moreover, as shown by Diebold and Rudebusch (1991), unit root tests have low power against fractional alternatives of this kind and lead to the erroneous conclusion that the time series have a unit root. Investigation into the fractionally integrated ARFIMA parameterisation of these volatilities thus seems to be justified. In the lower half of table 3 are given estimates of the ARFIMA(p,d,q) parameterisations of the volatility time series obtained with the Haslett and Raftery (1989) error decomposition procedure and selected according to the maximum LLF criterion. (For a discussion of alternative fractional integration estimation procedures see Baillie, 1996, pages 32-39.)

Parameter d reflects the long-term behaviour, whereas p, q and the corresponding AR and MA coefficients determine the short-term

Table 3

1 January 1993-31 December 1997

ADF Unit Root Tests

S. Sterling						Euro-Mark					
	ϖ	n		ϖ	n		ϖ	n		ϖ	n
$\sigma_{t,T}$	-4.5119	1	$\Delta\sigma_{t,T}$	-38.175	0	$\sigma_{t,T}$	-4.9978	1	$\Delta\sigma_{t,T}$	-23.017	2
IV(0-3)	-3.6517	4	ΔIV(0-3)	-23.412	2	IV(0-3)	-6.1307	1	ΔIV(0-3)	-36.947	1
IV(3-6)	-2.8093	6	ΔIV(3-6)	-27.644	1	IV(3-6)	-4.4810	1	ΔIV(3-6)	-40.786	0
IV(6-9)	-2.4246	2	ΔIV(6-9)	-19.288	3	IV(6-9)	-3.4110	1	ΔIV(6-9)	-39.410	0

Notes. The ϖ test statistics are obtained from the following estimates: $\Delta x_t = \iota + \varpi x_{t-1} + \Sigma_{i=1,\dots,n}\varphi_i \Delta x_{t-i} + e_t$ where n is selected using the AIC. The 5 percent critical value is –2.8642.

1 January 1993-31 December 1997

ARFIMA(p,d,q) Parameter Estimates

	d	Φ_1	Φ_2	Ψ_1	Ψ_2	**LLF**
S. Sterling						
$\sigma_{t,T}$	0.0562 (0.0072)	0.9575 (0.0321)		0.0857 (0.0084)	-0.0127 (0.0248)	2042.407
IV(0-3)	0.4150 (0.0101)	0.8582 (0.0111)		0.4601 (0.0092)	0.0558 (0.0103)	4051.408
IV(3-6)	0.4297 (0.0070)	0.8988 (0.0065)		0.3658 (0.0058)	0.1135 (0.0063)	4762.074
IV(6-9)	0.0327 (0.0112)	0.9915 (0.0060)		0.1159 (0.0029)		4995.290
Euromark						
$\sigma_{t,T}$	0.4431 (0.0095)	0.7334 (0.0265)		0.2811 (0.0193)		2662.581
IV(0-3)	0.1736 (0.0145)	0.9067 (0.0143)		0.4658 (0.0089)		3817.605
IV(3-6)	0.3119 (0.0152)	0.8977 (0.0076)		0.3429 (0.0064)	0.0249 (0.0071)	4610.787
IV(6-9)	0.0000 (0.0000)	0.8924 (0.0004)	0.0901 (0.0043)			4839.220

Notes. The estimates come from the zero mean volatility process $\Phi(L)(1-L)^d (x_t - \mu) = \Psi(L)\varepsilon_t$, where μ is the mean of the x_t time series. Its introduction is justified in Hwang and Satchell (1998). Estimated asymptotic standard errors in parentheses.

correlation structure. Indeed, Hosking (1981) has shown that the long-run behaviour of an ARFIMA(p,d,q) model is analogous to that of an ARFIMA(0,d,0) model with the same value of d. The range of d that is of interest in the context of long-memory modelling is $0 \leq d < 1/2$. In this case the process is mean-reverting.[11] It is stationary with long memory in the sense of McLeod and Hipel (1978) and is appropriate to model long-term persistence. Its correlations and partial correlations are all positive and decay hyperbolically to zero as the lag increases and not exponentially as in standard ARIMA models.

Most time series exhibit long-memory characteristics as evidenced by highly significant d estimates in the 0.4-0.5 interval. As usual, persistence is more marked for the Short Sterling than for the Euromark contract and tends to increase with the time to expiration of the underlying contract. The short-term implied volatilities and the corresponding realised volatilities estimates do not have many points in common, which casts doubts on the forecasting accuracy of the former.

ARFIMA modelling does not seem to be appropriate for long-term implied volatilities. The d estimates of the 6- to 9-month to expiration volatilities are not significantly different from zero, suggesting that deviations from the mean be short-memory. The associated autoregressive parameters, however, imply substantial shock persistence and (being close to one) are compatible with non-rejection of the unit root hypothesis provided by the ADF tests.

The estimates reported in this section suggest that both short-term implied volatilities and realized volatilities are characterised by substantial shock persistence, but do not behave as random walks. Econometric analysis of relationships involving these time series is thus rather difficult. Estimation in terms of levels might lead to a spurious regression bias and, in terms of first differences, to a misspecification bias due to over-differencing. Diebold and Nerlove (1990) point out that the latter may be costly; it tends to discard low frequency information and eliminate cointegration effects.

3. The information content of short-term implied volatility

The long-memory properties of the realised and implied volatility time series determine the choice of the estimation strategy. The estimates of equations (1) and (2) are set forth in table 4. The regressor (T-t) is introduced in order to condition for a time-to-expiration pattern identified in the realised future volatility $\sigma_{t,T}$. The predictive power of implied volatility is compared in equation (2) with that of a moving average - over the previous 20 days – of historical volatility.

The estimates of equation (1) set out in rows 1 and 8 reject the null hypothesis of option pricing informational efficiency. Implied volatilities seem to be biased predictors of future realised volatility as the corresponding null hypotheses (H_0: a = 0, b = 1) are always rejected at the five-percent level of significance. The evidence of bias seems to be larger for the Short Sterling than for the Euromark contract. Conversely, the adjusted coefficients for multiple correlation suggest that the explanatory power of implied volatility be higher for the Short Sterling than for the Euromark contract. The estimates are affected by strong serial correlation and by heteroskedasticity of the residuals, and the standard errors have been adjusted accordingly, using a standard GMM procedure.[12]

Christensen and Prabhala (1998) point out that, because of errors in variables due to the Black and Scholes misspecification of the stochastic volatility, standard OLS b estimates are biased. Estimates of equation (1) obtained with an instrumental-variables (two-stage) approach are set out in rows 2 and 9. The instruments are a constant term, the time-to-expiration dummy and up to two time periods own lagged values of implied volatility. Here, too, the GMM estimator procedure has been used since conditional heteroskedasticity and serial correlation affect the residuals. The overall quality of fit is not significantly altered even if the b estimates are closer to one in the case of the Euromark contract.

The encompassing regression estimates of equation (2) set out in rows 3 and 10 provide mixed results; for both contracts the coefficient of the MA(20) historical volatility is significant and the coefficient of multiple correlation tends to rise. However, inclusion of this regressor reduces the absolute value and significance of the Euromark implied volatility coefficient and does not affect the explanatory power of the Short Sterling implied volatility.

These estimates are not satisfactory, however, because of the very high serial correlation of the residuals. A spurious regression bias *à la* Granger and Newbold (1974) seems to affect the estimation, this bias being due to the long memory properties of the regressands and of the regressors ascertained in tables 1, 2 and 3.[13] (Phillips (1986) has shown that in this case the regression estimates converge to non-degenerate limiting distributions that can be expressed as functionals of Brownian processes and are thus inconsistent.) In rows 4 and 11 are set forth the estimates of equation (1) corrected for first-order serial correlation of the residuals using the Beach-MacKinnon maximum likelihood procedure. The value of the b coefficient drops dramatically to 0.33 in

Table 4

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$$\sigma_{t,T} = a + bIV(t,T) + d(T\text{-}t) + u_{t,T} \qquad (1)$$

$$\sigma_{t,T} = a + bIV(t,T) + c\sigma^{P}_{t,T} + d(T\text{-}t) + u_{t,T} \qquad (2)$$

	a	**b**	**d**	**c**	$\overline{R}^2$	S.D.	DW LM(5)	Arch(1)	W_0
S.Sterling									
OLS [1]	0.1031 (1.3664)	1.5809 (3.2350)	-0.0026 (-2.7175)		0.1895	0.1816	0.0683 1215.1*	1036.4 [0.00]	327.65 [0.00]
CP [2]	0.1005 (1.3081)	1.6015 (3.1527)	-0.0026 (-2.7185)		0.1895	0.1816	0.0685 5333.5*	1039.5 [0.00]	9.660 [0.01]
OLS [3]	0.0938 (1.3065)	1.8891 (3.9083)	-0.0023 (-2.8332)	-0.3995 (-1.9856)	0.1958	0.1809	0.0735 5348.2*	1031.2 [0.00]	10.659 [0.01]
AR [4]	0.1945 (4.0825)	0.3348 (2.6391)	-0.0034 (-21.878)		0.9514	0.0444	2.0498 18.332*	0.6669 [0.41]	55.975 [0.00]
AR [5]	0.2922 (6.0372)	0.3623 (3.3715)	-0.0034 (-21.899)	-0.1456 (-1.5766)	0.9515	0.0444	2.0780 17.575*	0.6755 [0.00]	57.844 [0.00]
FD [6]	0.0176 (6.1491)	0.2627 (2.1304)	-0.0005 (-7.1605)		0.0400	0.0514	2.0870 12.015**	0.0373 [0.84]	71.695 [0.00]
FD [7]	0.0175 (6.0541)	0.2655 (2.1509)	-0.0005 (-7.0407)	-0.0541 (-0.5058)	0.0400	0.0514	2.1093 12.775**	0.0403 [0.84]	71.909 [0.00]
	a	**b**	**d**	**c**	$\overline{R}^2$	S.D.	DW LM(5)	Arch(1)	W_0
Euromark									
OLS [8]	0.1320 (2.7862)	0.7921 (2.0067)	-0.0021 (-3.2202)		0.1313	0.1067	0.0694 1210.9*	918.84 [0.00]	309.59 [0.00]
CP [9]	0.1141 (2.0887)	0.9250 (1.9795)	-0.0021 (-3.2195)		0.1302	0.1068	0.0731 5402.2*	905.85 [0.00]	15.740 [0.00]
OLS [10]	0.1162 (2.5657)	0.3949 (1.1060)	-0.0026 (-3.1566)	0.9669 (3.5212)	0.2163	0.1017	0.0766 4998.6*	962.28 [0.00]	474.22 [0.00]
AR [11]	0.2499 (9.6988)	0.0425 (0.8562)	-0.0026 (-29.229)		0.9510	0.0254	2.0630 2.6320	0.7070 [0.40]	393.13 [0.00]
AR [12]	0.2574 (9.5267)	0.0461 (0.9268)	-0.0026 (-29.231)	-0.0898 (-1.2233)	0.9511	0.0254	2.0937 3.4845	0.7448 [0.39]	394.07 [0.00]
FD [13]	0.1240 (6.8841)	-0.0015 (-0.0237)	-0.0004 (-8.0759)		0.0466	0.0322	2.1140 7.5644	0.0093 [0.92]	296.91 [0.00]
FD [14]	0.0123 (6.8103)	0.0003 (0.0005)	-0.0004 (-7.9709)	-0.0347 (-0.3733)	0.0459	0.0322	2.1266 11.362**	0.0123 [0.91]	296.85 [0.00]

Notes. CP: Christensen and Prabhala (1998) IV estimates; AR: Maximum Likelihood estimates corrected for AR(1) serial correlation of the residuals; FD: First differences OLS estimates; W_0 : Wald test χ^2 statistic for the null hypothesis that a = 0, b = 1(and c = 0); ** : Significant at the 5 % level; * : Significant at the 1 % level. Probabilities are in square brackets, estimated t ratios in parentheses. The t ratios of the levels estimates are robust to heteroskedasticity.

the case of the Short Sterling and to zero in the case of the Euromark. Because of the high degree of serial correlation, the estimation is repeated in terms of first differences and provides analogous results (rows 6 and 13); Short Sterling implied volatility changes only have a significant positive impact on realised future volatility changes.

The estimation of the encompassing regressions set out in rows 5, 7, 12 and 14 corroborates these findings; inclusion of the historical volatility regressor does not affect the (significant) explanatory power of Short Sterling implied volatility, nor does it affect the explanatory power of Euromark implied volatility, which remains insignificant.

The estimates of table 4 suggest that daily short-term implied volatilities fail to predict daily realised future volatility accurately. Does this mean that implied volatilities have to be discarded altogether as having no relevant information content? The answer is, it does not. Financial analysts are mostly concerned with daily volatility forecasts. The fact that implied volatilities are but poor predictors of future realised volatilities does not necessarily imply that they have low predictive power on current volatility too.

GARCH modelling of interest rate volatility provides a useful framework for assessing the relevance of implied volatility as current volatility predictor. The following PGARCH(1,1,1) model seems to provide a reasonable parameterisation of the conditional standard deviation of the underlying and is used as a benchmark

$$ln(i_t / i_{t-1}) = \zeta + \varepsilon_t \ , \tag{7}$$

$$\sigma_t^d = \omega + \alpha(|\varepsilon_{t-1}|)^d + \beta\sigma_{t-1}^d + \gamma i_{t-1} \tag{8}$$

where d = 1 and i_t is the interest rate implied by the underlying futures contract. The estimates are set out in columns 1 and 5 of table 5.[14]

Table 5

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$$ln(i_t/i_{t-1})=\zeta+\varepsilon_t \quad (7)$$
$$\sigma_t^d=\omega+\alpha(|\varepsilon_{t-1}|)^d+\beta\sigma_{t-1}^d+\gamma i_{t-1} \quad (8)$$
$$\sigma_t^d=\omega+\alpha(|\varepsilon_{t-1}|)^d+\beta\sigma_{t-1}^d+\gamma i_{t-1}+\delta IV(t,T) \quad (9)$$
$$\sigma_t^d=\omega+\alpha(|\varepsilon_{t-1}|)^d+\beta\sigma_{t-1}^d+\gamma i_{t-1}+\delta IV(t,T)+\phi IV(t-1,T) \quad (9')$$
$$\sigma_t^d=\omega+\alpha(|\varepsilon_{t-1}|)^d+\beta\sigma_{t-1}^d+\gamma i_{t-1}+\phi IV(t-1,T) \quad (9'')$$

	Short Sterling				Euromark			
Model	(7)-(8) [1]	(7)-(9) [2]	(7)-(9') [3]	(7)-(9'') [4]	(7)-(8) [5]	(7)-(9) [6]	(7)-(9') [7]	(7)-(9'') [8]
ζ	-0.0000 (-0.0000)	-0.0000 (-0.0000)	-0.0000 (-0.0000)	-0.0000 (-0.0000)	-0.0000 (-0.0000)	-0.0000 (-0.0000)	-0.0000 (-0.0000)	-0.0000 (-0.0000)
ω	0.0001 (3.6930)	-0.0001 (-0.0513)	-0.0005 (-0.3199)	-0.0036 (-1.9440)	0.0002 (2.7670)	-0.0009 (-2.0042)	-0.0008 (-1.5510)	-0.0004 (-0.7610)
α	0.0722 (9.5150)	0.0956 (7.3500)	0.0856 (7.8502)	0.0941 (7.1030)	0.0329 (3.7110)	0.0228 (0.8602)	0.0294 (1.0950)	0.0331 (1.2100)
β	0.8919 (7.2550)	0.2129 (1.8670)	0.1001 (0.4607)	-0.1467 (-2.0620)	0.9451 (5.7110)	0.4829 (3.2588)	0.2281 (1.1280)	0.2949 (1.7220)
γ	-0.0001 (-2.1290)	0.0001 (0.7461)	0.0002 (0.9096)	0.0006 (2.2880)	-0.0001 (-0.5076)	0.0002 (2.6197)	0.0002 (2.1510)	0.0002 (2.3640)
δ		0.5979 (5.7720)	0.6154 (6.3392)			0.4195 (3.2758)	0.1893 (1.7330)	
φ			0.0095 (0.5545)	0.9827 (11.7700)			0.4173 (2.5510)	0.5055 (4.3830)
LLF	4540.60	4599.09	4596.88	4606.09	4582.47	4620.49	4599.52	4601.21
AIC	-9069.19	-9184.19	-9177.76	-9198.17	-9152.94	-9226.97	-9183.04	-9188.41
Stand. Resid.								
Sk.	-0.0832	-0.4424	-0.4517	0.8234	-4.2656	-2.5453	-3.2635	-3.4755
Kurt.	99.2949	778.3710	79.787	93.4147	53.2106	37.0455	46.4811	49.4189
LB(12)	11.18 [0.513]	12.24 [0.426]	13.56 [0.329]	10.44 [0.577]	11.30 [0.504]	14.05 [0.298]	13.75 [0.317]	13.280 [0.349]
Arch (12)	5.724 [0.996]	5.047 [0.956]	6.865 [0.866]	3.110 [0.995]	0.795 [0.999]	1.059 [0.999]	0.761 [0.999]	0.714 [0.999]
ν (St.E.)	0.9703 (0.0148)				0.9821 (0.0193)			
$L.R._{\delta,\phi=0}$		116.98 [0.00]	112.56 [0.00]	130.98 [0.00]		76.04 [0.00]	34.10 [0.00]	37.48 [0.00]

Notes. LB(x): Ljung-Box Q statistic for x^{th} order serial correlation of the standardised residuals; L.R.: Likelihood Ratio test statistics of the null hypotheses δ =0, δ =φ = 0 or φ = 0; ν: degrees of freedom parameter of the Ged conditional distribution.

Volatility is highly persistent, especially in the case of the Euromark futures contract. No asymmetry has been identified; good news and bad news seem to have an analogous impact. Lagged interest rates have a small and insignificant coefficient in the Euromark equation; low rates do not seem to exert the dampening effect on volatility identified by Brenner et al. (1996). The standardized and squared standardized residuals, however, show but little evidence of serial correlation and seem to corroborate the choice of the model specification. A Ged conditional distribution of the residuals has been imposed in the estimation, - a choice justified by the strong rejection of conditional normality due to a high degree of kurtosis.

Within sample information content of current implied volatility is assessed estimating the following conditional standard deviation relationship

$$\sigma_t^d = \omega + \alpha(|\varepsilon_{t-1}|)^d + \beta\sigma_{t-1}^d + \gamma i_{t-1} + \delta IV(t,T) \qquad (9)$$

where, here too, it is assumed that d = 1. (IV(t,T) is expressed here on a daily basis.)

The estimates are set out in columns 2 and 6. LR tests of the null hypothesis that $\delta = 0$ are significant at the 5-percent level; implied volatility seems to have a relevant information content. The Short Sterling implied volatility seems to provide a forecast of realised daily volatility that is more accurate (the δ coefficient estimates are closer to one in absolute value), a result that corroborates the findings obtained in the estimation of equation (2) above. The coefficients of the GARCH regressors however, even if smaller in absolute value, do not lose all of their statistical significance. The informational efficiency hypothesis of Black and Scholes option pricing is thus rejected since volatility forecasts from implied volatilities can be improved with the help of historical information. The adjustment suggested by Amin and Ng (1997, page 553) in order to eliminate the implicit lagged implied volatility terms provides mixed results. The estimates of

$$\sigma_t^d = \omega + \alpha(|\varepsilon_{t-1}|)^d + \beta\sigma_{t-1}^d + \gamma i_{t-1} + \delta IV(t,T) + \phi IV(t-1,T) \quad (9')$$

set out in columns 3 and 7 produce evidence of a downward bias in the Short Sterling δ estimates only.[15] The δ coefficient Euromark estimates are smaller, and not larger than the corresponding unadjusted estimates of equation (9).

In order to assess the predictive power of implied volatility, its lagged value is appended to equation (8), producing the following conditional standard deviation parameterisation

$$\sigma_t^d = \omega + \alpha(|\varepsilon_{t-1}|)^d + \beta\sigma_{t-1}^d + \gamma i_{t-1} + \phi IV(t-1,T) \quad (9'')$$

The estimates of the ϕ coefficient set out in columns 4 and 8 are larger in absolute value and more significant than the corresponding α and β coefficient estimates. Implied volatilities thus seem to have greater predictive power than the historical forecasts provided by the GARCH components.

The results of this section suggest that implied volatilities provide reliable predictions of the current volatility of the underlying, - predictions that seem to be more accurate for the Short Sterling than for the Euromark contracts. Even if they are not sufficient predictors of realised volatility (the information efficiency hypothesis is rejected throughout), they seem to outperform alternative historical forecasts.

4. The transmission of news over time and across contracts

4.1. The transmission of news over time

Conditional upon hypotheses reported in section 1.2 above, Stein (1989) derives an *ex ante* relationship between implied volatilities from options with differing time to expiration which is assumed to reflect the transmission of news over time.[16] Using equation (5) and the estimated first-order autocorrelation coefficients of both contracts, the theoretical response of 3- to 6- and 6- to 9-month-to-expiration implied volatility are computed and are set out in the first column of table 6. The smaller value of the $\beta(\rho)$ coefficients – if we compare them with the findings of Stein and of Diz and Finucane – is to be attributed to the longer time to expiration of the options involved rather than to a higher degree of mean reversion of the short-term volatilities. The autocorrelation coefficients are of the same order of magnitude as those reported in the studies mentioned above. This finding is in line with the *ex ante* hypotheses: an increase (decrease) in short-term implied volatility is associated with a smaller increase (decrease) in long-term implied volatility since the latter incorporates a mean-reverting component. The longer the time interval between the short-term and the long-term options involved, the smaller the impact on long run volatility and the greater the degree of mean reversion. (For a discussion of this phenomenon see Tessaromatis (1998).)

Empirical (*ex post*) responses of long-term implied volatilities to shifts in short-term volatility are obtained rewriting equation (4) in the following estimable form

$$\mathrm{IV}(t, m_{2t}) = \eta + \lambda \ \mathrm{IV}(t, m_{1t}) + e_t \tag{10}$$

where it is assumed that $\eta = [1 - \beta(\rho)]\overline{\sigma}$, $\lambda = \beta(\rho)$ and that the residuals have zero mean and are independently and identically normally distributed.

Theoretical and empirical measures of elasticity tend to be larger for the Short Sterling than for the 3 Month Euromark contract. Indeed, as evidenced by the ARFIMA analysis of short-term implied volatilities and inspection of the corresponding autocorrelation coefficients, mean-reversion seems to be more pronounced for the latter contract.

Here, too, the Beach-MacKinnon and first difference estimates are significantly smaller than the corresponding OLS estimates.[17] The evidence of an overreaction of long-term volatility to changes in short-term volatility provided by the OLS estimates in the levels seems to be the result of a spurious regression bias.[18] A comparison of the theoretical $\beta(\rho)$ coefficient with the corresponding adjusted and first difference λ estimates suggests a serious underreaction (and not overreaction) of long-term volatilities across both contracts. Acceptance of the spurious regression hypothesis is thus of paramount importance since it leads to radically different results. It should be noted that first differencing may well introduce an overdifferencing

Table 6
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$$IV(t,m_{2t}) = \eta + \lambda\, IV(t,m_{1t}) + e_t \qquad (10)$$

$\beta(\rho)$ is the sample mean of $m_{1t}[\rho^{m}_{2t} - 1] / m_{2t}[\rho^{m}_{1t} - 1]$ where ρ is the first order autocorrelation coefficient of the short-run implied volatility time series (0-3 months), m_{1t} is the number of days to maturity of the short-term maturity option and m_{2t} is the number of days to maturity of the long-term option.

			OLS			AR			FD	
S. Sterling $IV(m_{2t})$	$\beta(\rho)$	λ	$\bar{R}^2$	DW LM(5)	λ	$\bar{R}^2$	DW LM(5)	λ	$\bar{R}^2$	DW LM(5)
3-6	0.3833	0.8749 (73.31)	0.8112	0.2895 936.6*	0.1517 (9.639)	0.9810	1.9892 4.1795	0.1292 (8.305)	0.0516	2.1719 22.182*
6-9	0.2379	0.8107 (53.51)	0.6961	0.1486 1095.7*	0.1171 (9.025)	0.9870	1.9975 8.8878	0.1021 (7.817)	0.0471	2.2750 31.982*
Euro-mark $IV(m_{2t})$	$\beta(\rho)$	λ	$\bar{R}^2$	DW LM(5)	λ	$\bar{R}^2$	DW LM(5)	λ	$\bar{R}^2$	DW LM(5)
3-6	0.3659	0.7646 (13.91)	0.5959	0.4115 2859.5*	0.1226 (8.523)	0.9386	2.0098 3.8028	0.1152 (8.116)	0.0483	2.2645 33.045*
6-9	0.2281	0.7521 (11.12)	0.4977	0.2568 4084.4*	0.0996 (8.088)	0.9634	1.9980 6.4797	0.0935 (8.016)	0.0465	2.2106 23.345*

Notes. AR: Maximum Likelihood estimates corrected for AR(2) serial correlation of the residuals; FD: First differences OLS estimates. *: significant at the 1% level. The t ratios of the level and first differences OLS estimates are robust to heteroskedasticity.

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$$\left(\frac{1}{k}\right) \sum_{i=0}^{k-1} \left[IV(iQ,(i+1)Q) \right] = \tau + \theta\, IV(0,kQ) + \sum_{i=0}^{k-1} u_{iQ} \qquad (11)$$

S. Sterling	τ	θ	$\bar{R}^2$	S.D.	DW	LM(5)	Arch(1)	W_0
OLS	0.0034 (0.4242)	0.6934 (17.6650)	0.7064	0.0198	0.1139	1051.25*	895.71*	8094.95*
AR	0.0795 (6.7850)	0.2356 (7.4642)	0.9734	0.0059	2.002	9.35	53.55*	603.38*
FD	-0.0000 (-0.3715)	0.1821 (3.5246)	0.0267	0.0060	2.2792	44.97*	38.51*	668.87*
Euro-Mark	τ	θ	$\bar{R}^2$	S.D.	DW	LM(5)	Arch(1)	W_0
OLS	0.0612 (6.5954)	0.4201 (7.9415)	0.3859	0.0159	0.2532	675.55*	1464.94*	532.79*
AR	0.1058 (14.6545)	0.1929 (5.6962)	0.8669	0.0074	1.9887	4.43	49.50*	809.84*
FD	-0.0000 (-0.0985)	0.1512 (4.1277)	0.0133	0.0079	2.5000	110.99*	137.79*	536.70*

Notes. AR: Maximum Likelihood estimates corrected for AR(3) serial correlation of the residuals; FD: First differences OLS estimates; *: Significant at the 1% level; W_0: Wald test statistic for the joint hypothesis that $\tau = 0$ and $\theta = 1$. The t ratios of level and first differences OLS estimates are robust to heteroskedasticity.

misspecification as the estimates have to be corrected for a significant negative first-order autocorrelation of the residuals.

The time series properties of implied volatilities justify the adoption of the same estimation strategy in the term structure investigation set forth in the lower half of table 6. The (null) expectations hypothesis is that the current long-term volatility be equal to the average of the current and expected short-term volatilities. It involves estimation of the following relationship

$$\left(\frac{1}{k}\right) \sum_{i=0}^{k-1} [\, \mathrm{IV}(iQ,(i+1)Q) \,] = \tau + \theta\, \mathrm{IV}(0,kQ) + \sum_{i=0}^{k-1} u_{iQ} \qquad (11)$$

with k = (long-term option maturity)/(short-term option maturity). IV(iQ,(i+1)Q) is the implied volatility quoted at time iQ for an option with expiration date (i+1)Q and Q indicates 3 months i.e. 63 trading days.[19] It follows that k is 3 (3 quarters / 1 quarter). Under the null of rational expectations and of option market efficiency $\theta = 1$ and $\tau = 0$ (the latter is assumed to quantify a risk premium). Wald test statistics suggest that the null is rejected for both contracts. Efficiency in the transmission of news over time, however, seems to be lower for the Euromark contract, irrespective of the estimation procedure.

From an economic point of view, long-term volatility underreaction can be explained using a stale price quotation rationale. If at-the-money options tend to be traded less frequently as their time to expiration recedes over time, new information will affect their price – and the corresponding implied volatility – less frequently.[20] A shock which impacts on short-term contracts will affect only a fraction of the long-term contracts and will result, on average, in a long-term volatility underreaction effect.

4.2. The transmission of news across countries

In order to assess the relevance of linkages between implied volatilities across contracts three bivariate VAR systems have been estimated with OLS. They involve implied volatility daily changes from the short-term, medium-term and long-term option contracts of interest and read as follows

$$\Delta IV(t,T_k)^{UK} = a_{0k}^{UK} + \sum_{i=1}^{6} b_{ik}^{UK} \, \Delta IV(t\text{-}i,T_k)^{UK} + \sum_{i=1}^{6} c_{ik}^{UK} \, \Delta IV(t\text{-}i,T_k)^{D} + u_{tk}^{UK} \qquad (12)$$

$$\Delta IV(t,T_k)^{D} = a_{0k}^{D} + \sum_{i=1}^{6} b_{ik}^{D} \, \Delta IV(t\text{-}i,T_k)^{UK} + \sum_{i=1}^{6} c_{ik}^{D} \, \Delta IV(t\text{-}i,T_k)^{D} + u_{tk}^{D} \qquad (13)$$

$$k = 1,2,3$$

where T_k is the time to expiration of the option. It varies from 0 to 3 months for k =1, from 3 to 6 months for k =2 and from 6 to 9 months for k=3. Implied volatility daily changes are investigated in order to eliminate spurious regression and multicollinearity distortions associated with the strong serial correlation of the time series. The estimates of the VAR system are not reported for lack of space; the VAR order 6 has been selected with the help of the Akaike Information and Schwarz criteria and ensures that any serial correlation of the residuals has been expunged. It is well known that Granger causality tests are essentially tests of the predictive accuracy of time series models. Time series $\Delta IV(t,T_k)^{UK}$ causes time series $\Delta IV(t,T_k)^{D}$ in the Granger sense if current $\Delta IV(t,T_k)^{D}$ can be predicted better by using past values of $\Delta IV(t,T_k)^{UK}$ than by not doing so, conditioning on additional relevant information, including past values of $\Delta IV(t,T_k)^{D}$.

Table 7
1 January 1993-31 December 1997

$$\Delta IV(t,T_k)^{UK} = a_{0k}^{UK} + \sum_{i=1}^{6} b_{ik}^{UK}\,\Delta IV(t-i,T_k)^{UK} + \sum_{i=1}^{6} c_{ik}^{UK}\,\Delta IV(t-i,T_k)^{D} + u_{tk}^{UK} \qquad (12)$$

$$\Delta IV(t,T_k)^{D} = a_{0k}^{D} + \sum_{i=1}^{6} b_{ik}^{D}\,\Delta IV(t-i,T_k)^{UK} + \sum_{i=1}^{6} c_{ik}^{D}\,\Delta IV(t-i,T_k)^{D} + u_{tk}^{D} \qquad (13)$$

χ^2 tests for the null hypotheses
$H_0 : c_{ik}^{UK} = 0$, i= 1,…,6; $H_1 : b_{ik}^{D} = 0$, i=1,…,6.

	Direction of Causality	Null Hypothesis		Time to Maturity (months)	
			0-3	**3-6**	**6-9**
Equation (12)	$\Delta IV^{D} \rightarrow \Delta IV^{UK}$	H_0	7.2634 [0.297]	3.2377 [0.778]	9.5584 [0.145]
Equation (13)	$\Delta IV^{UK} \rightarrow \Delta IV^{D}$	H_1	20.4180* [0.020]	36.6109* [0.030]	50.6992** [0.00]

Notes. Equation (12) tests if 3 Month Euromark implied volatility changes cause Short Sterling implied volatility changes; equation (13) tests if Short Sterling implied volatility changes cause 3 Month Euromark implied volatility changes. *: Significant at the 5% level. **: Significant at the 1% level; probability values are in square brackets.

The test is performed verifying hypotheses H_0 and H_1 set out in table 7.

Column 1 provides the direction of causality; columns 3, 4 and 5 report the statistics for the null hypothesis of no causality for the 3 sets of option contracts. The χ^2 statistics of the second row only are significant and reject the null hypothesis of no causality.

Spillovers across contracts were attributed in section 1.2 above to a contagion-like effect. Internationally relevant news should affect both contracts, and country-specific news should not be transmitted to another contract. The findings of this section suggest that news which brings about implied volatility changes in the Short Sterling contract systematically affects implied volatility from the 3-Month Euromark contract, whereas the opposite is not the case. They are conducive to the overall conclusion – which also emerges from the previous sections – of a reduced efficiency in the pricing of the German interest rate futures option contract on the LIFFE.

5. Concluding remarks

The analysis of the stochastic behaviour of implied volatilities computed by inverting a Black-Scholes-like formula, which postulates a constant volatility, may at first sight seem contradictory. It has been shown, however, that under rather general conditions Black-Scholes implied volatilities from at-the-money options appropriately quantify, in each period, the market expectations of the average volatility of the return of the underlying asset until contract expiration. The efficiency of these expectation estimates is investigated here for options on two major short-term interest rate futures contracts traded at the LIFFE. The analysis is strongly affected by the stationary long-memory characteristics of the implied volatility time series, which may lead to serious specification errors. Even if they are not sufficient predictors of realised volatilities, implied volatilities seem to outperform alternative historical forecasts.

Over the 1993–1997 time interval the performance of implied volatilities is not homogeneous across contracts. Information content and predictive-power tests consistently suggest that implied volatility from Short Sterling contracts is more accurate as a future volatility predictor than implied volatility from 3-Month Euromark contracts. Analysis of the efficiency of news transmission over time and between contracts provides analogous results. Underreaction of long-term volatility to changes in short-term volatility is more relevant to the German interest rate contract than to the British one, and Short Sterling implied volatility changes do “Granger cause” 3 Month Euromark implied volatility changes pointing to a contagion–like interlinkage.

Even in a sophisticated international financial market like the LIFFE implied volatilities have a country-specific pattern as traders seem to be more proficient in predicting domestic interest rate volatility. A possible interpretation is that a (foreign) country risk premium introduces a bias in the Black–Scholes implied volatility estimates. Whether this result is general or, rather, is restricted to the time period and/or to the contracts under investigation provides the scope for future research.

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Notes

[1] Walter and Lopez (2000) point out that equation (1) tests for the partial optimality of $\sigma^F_{t,T}$, i.e. whether forecast errors $u_{t,T}$ are unforecastable with respect to the subset of available information embedded in the forecast $\sigma^F_{t,T}$. If a forecast is partially optimal, the forecast errors should be orthogonal to the forecast itself, producing a = 0 and b = 1.

[2] Stein points out that implied volatility will be an accurate estimator of average expected volatility of the underlying over the remaining life of the option if there is no risk premium and if the price of the option is linear in volatility. He derives his test using a stochastic volatility model *à la* Hull and White (1987). It can, however easily be extended to Black-Scholes implied volatilities. Feinstein (1989) shows that for a stochastic volatility option model the value of an at-the-money option is approximately equal to the Black-Scholes value with the volatility given by the average expected value of the underlying asset over the remaining life of the option. For more details on the interpretation of Black-Scholes implied volatilities as conditional forecasts of the average volatility over the remaining life of the option, see Franks and Schwartz (1991) and Fleming (1993).

[3] Stein's model has been generalised in various ways. Heynen et. al. (1994) adapted it to GARCH parameterisations of the volatility of the returns of the underlying using the option pricing approach of Duan (1994). Xu and Taylor (1994) introduce time-varying long-term expectations and use a Kalman filter to infer the term structure of volatility expectations from implied volatilities from options quoted at six (and no longer two) differing time intervals to expiration.

[4] The theoretical argument is based on the near-linearity in volatility of the Black–Scholes formula in the case of at-the-money options. It involves the ratio of the short-term to the long-term Black-Scholes pricing bias relative to the corresponding Hull-White option prices. The authors do not have to posit that the Black-Scholes bias is nil; they simply assume that it changes little with the time to maturity of the option contract and conclude that disregarding a ratio that is close to one does not seriously affect the analysis .

[5] Black's model for the evaluation of a European option (and - as shown by Lieu (1990) - a margined American option traded on the LIFFE) is set out in standard textbooks, such as Brys et al. (1998), pages 109-110.

[6] Realised future volatilities are computed as follows.
Let $R_t = \log (i_t / i_{t-1})$ where i_t is the implied interest rate (100-the price of the underlying futures contract) at time t. Let T be the time of expiration of the

option contract and T-t the time left to expiration. Realised future volatility reads as follows

$$\sigma_{t,T} = \left\{ \frac{1}{T-t-1} \sum_{i=1}^{T-t} \left(R_{t+i} - \left(\frac{1}{T-t} \sum_{j=1}^{T-t} R_{t+j} \right) \right)^2 \right\}^{0.5}$$

[7] The scaled and standardized spectral density estimate $f(\omega_j)$ reads as follows

$$f(\omega_j) = 1 + 2 \sum_{k=1,\, m} \lambda_k (\gamma_k/\gamma_0)\cos(\omega_j k)$$

where λ_k is the Bartlett kernel, γ_k is the sample k^{th} autocovariance estimate, $\omega_j = j\pi/m$ denotes the j^{th} frequency and $m = 2(T)^{0.5}$ is the bandwidth parameter.

[8] The spectral density estimates are computed using the residuals of OLS regressions on a constant and a time trend. The estimates are very close to those obtained either using residuals of OLS regressions on a constant, a time trend and a time-to-expiration dummy, or using unadjusted volatilities.

[9] It is well known that differencing a trend-stationary time series induces a negative unit root in its MA representation, resulting in a zero spectral density at the origin. Cochrane (1988) interprets the zero frequency value of the scaled spectrum of the first difference of a time series as a measure of shock persistence. The presence of a permanent component in a time series via a unit autoregressive root implies a nonzero spectrum in its difference at the origin. The more persistence induced by the unit root, the larger the zero spectral power. Ouliaris et al. (1989) introduce upper and lower bounds of the distribution of this statistic, which is shown to be asymptotically normally distributed.

[10] A fractionally integrated ARIMA(p,d,q) or ARFIMA(p,d,q) process reads as $\Phi(L)(1-L)^d x_t = \Psi(L)\varepsilon_t$. All roots of $\Phi(L)$ and $\Psi(L)$ lie outside the unit circle and ε_t is iid $(0, \sigma^2)$. The fractional difference operator is defined as $(1-L)^d = \sum_{k=0,\,\infty}\{\Gamma(k-d)L^k / [\Gamma(k+1)\,\Gamma(-d)]\}$ where $\Gamma(.)$ is the gamma function.

[11] For $-1/2<d<1/2$ the process is covariance stationary, while $d<1$ implies mean reversion. This is in contrast to a unit root process which is both covariance non-stationary and not mean-reverting. When $-1/2<d<0$, the process has short memory and, but for the first order one, negative, slowly decaying autocorrelations.

[12] The variance covariance matrix of the residuals is computed using Hansen's (1982) GMM approach. The lag truncation parameter of the Newey-West kernel is selected according to the Andrews (1991) procedure.

[13] Fractional cointegration analysis (Granger, 1981 and Cheung and Lai, 1993) posits that the time series be I(d), d>1/2 and cannot be implemented here.

[14] Previous analyses by Day and Lewis (1992) and Lamoureux and Lastrapes (1993) deal with the conditional variance of the return of the underlying and use it to investigate the properties of squared implied volatilities. The time series properties of the latter, however, differ from those of implied volatilities. We have thus adopted a conditional standard error framework in order to assess the information content and predictive power of implied volatilities.

[15] In the Euromark contract estimates the α coefficient is not significant. Amin and Ng show that this implies that there is no GARCH effect, even if the β coefficient estimate is significant. In this sense Euromark implied volatility informational efficiency holds.

[16] Stein postulates an AR(1) parameterisation of implied volatilities. Examination of the AC and PAC coefficients of table 1 suggests that this might be incorrect, introducing a specification bias in the $\beta(\rho)$ coefficient estimates.

[17] Diz and Finucane (1993) attribute this divergence to the incorrect specification of equation (10), which postulates serially uncorrelated residuals. They point out that volatility shocks and measurement errors tend to introduce a MA component in short-term implied volatility time series and show – using a Monte Carlo simulation – that such an error component biases upwards OLS estimates of λ in equation (10) without altering the estimates adjusted for serial correlation, a spurious regression symptom.

[18] The estimation has been repeated adding a time-to- expiration dummy to the regressors with no significant change in the results. The estimates are not reported for lack of space.

[19] Equation (11) is derived from equation (9) of Campa and Chang expressed on a quarterly basis and replacing squared volatilities with volatilities. Campa and Chang subtract short-term implied volatilities from the regressor and from the regressand in order to eliminate any non-stationary bias. The adjustment does not work in this context, however, as long-term and short-term implied volatilities have differing long-memory properties.

[20] By definition long-term implied volatility quantifies expectations on the volatility of the underlying that are projected farther in the future than the volatility associated with a shorter term implied volatility. Distant future expectations may not react to the arrival of short-term information because agents are not certain of their effect so far off in the future. This hypothesis

may explain why the degree of mean reversion seems to be higher for short-term than for long-term implied volatilities. Random walk behaviour in this case is not to be associated with a market efficiency paradigm but, rather, with hysteresis.