AN INVESTIGATION INTO THE LINKAGES BETWEEN EURO AND STERLING SWAP SPREADS

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Abstract

This paper examines the causal relationship between euro and sterling swap spreads during the period January, 1999 to March, 2003. The absence of any correlation between changes in the two swap spreads would indicate that credit risk factors are country-specific. But euro swap spreads showed some correlation with the interest rate differentials between the two markets. Both spreads follow a GARCH process but sterling swap spreads reacted more intensely to market movements and were more volatile than their euro counterparts. There was evidence of mild volatility transmission from the sterling swap spreads to the euro swap spreads but the causality was one sided.

Keywords: interest rate swaps, swap spreads, bonds, volatility.

JEL classification: G15

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I INTRODUCTION

The observed difference between the swap rate and the government bond yield of corresponding maturity is known as the swap spread. Fixed income securities, including corporate bonds and mortgage-backed securities use interest rate swap spreads as a key benchmark for pricing and hedging. A conventional interest swap is a contract between two companies or counterparties in which one party makes fixed interest payments, calculated on a notional amount, while the other party makes floating-rate interest payments. If swap rates incorporate the risk of default they would be sensitive to the credit ratings of the counterparties. The fixed rate is set at the inception of the contract and the floating-rate is linked to an external reference such as Libor [1] during the life of the swap.

Interest rate swaps, like most bonds are traded over the counter (OTC), rather than, through an organised exchange. Similar to other OTC securities, swaps are characterised by the presence of credit and liquidity risks. Each of the two parties in an OTC transaction is exposed to the default risk of the other. Thus, to compensate for these risks, market swap rates are generally at a premium over the comparable government bond rates. This premium is termed as the *swap spread*. Swap spreads, therefore, reflect the default risk of the interbank market quoting Libor/Euribor rates and the government treasury. However, the spread is not necessarily a pure measure of credit risk, as it can also be indicative of liquidity risk.

The importance of interest swap spreads derives from the dramatic recent growth in the notional amount of interest rate swaps outstanding relative to the government bond markets. After the introduction of the single currency, the euro swap market has nearly doubled in size and grown much faster than the bond market. [2] This can be attributed to the lack of homogeneity in the euro-denominated government securities market inducing a shift to interest rate swaps for hedging and positioning activity.

Swap spreads can be volatile and this has been very much in evidence during recent years. The Russian debt crisis in the autumn of 1998 and the subsequent near collapse of Long-Term Capital Management (LTCM), resulted in a flight to UK, US, and German government bonds which lowered yields and widened swap spreads. This "flight-to-quality" caused by concerns about a systematic meltdown in the financial sector, had a profound effect on the importance of the swap market. A Treasury yield does not incorporate the risk premium that characterises a swap spread. Traditionally, it was the risk-free nature of the Treasury yield curve that necessitated its choice as a benchmark. During the 1998 financial crisis, the flight-to-quality bid that occurred in Treasury bonds, depressed their yields below "true" nominal risk-free rates and resulted in a steep increase in risk premiums. This impinged on the efficacy of Treasury bonds as benchmarks. As the market for Treasury bonds decoupled from other asset classes, market participants who hedged their portfolios with Treasury securities found themselves being adversely affected.

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[1] The reference rate is GBP 6month Libor for sterling swaps and EUR 6 month Libor for euro swaps.
[2] Remolona and Wooldridge (2003)

In the literature, swap spreads have been attributed mainly to two factors: the credit risk of counterparties giving rise to a default premium and, the liquidity of the swap market relative to the government securities market giving rise to a liquidity premium. Sun, Sundaresan and Wang (1993), Sorensen and Bollier (1994), Brown, Harlow and Smith (1994) are among those arguing in favour of default risk as a primary determinant of swap spread changes. While Grinblatt (1995) and Liu, Longstaff and Mandell (2002) support the view that liquidity risk is a more plausible determinant of swap spreads than credit risk. Duffie and Singleton (1997) find that both credit and liquidity risk affect the behaviour of swap spreads but at different time horizons. Liquidity factors are more important in short horizons while credit shocks are more significant over long horizons. However, most studies show that the expected spread between LIBOR rates and the corresponding Treasury bill rates (TED spread) is the most basic determinant of swap spreads.

The purpose of this paper is not on analysing the determinants of swap spreads, but rather the dynamic behaviour of swap spreads. In particular, we are focussing on the transmission of information across the euro and sterling fixed income markets and explore volatility interdependencies. Time series models of asset returns have emphasised stylised facts in the form of volatility clustering, whereby one period of high volatility is followed by more of the same, and then successive episodes of low volatility. Generalised autoregressive conditional heteroskedastic (GARCH) processes which parameterise time-varying conditional variances are able to capture this behaviour.

There have been several studies that have employed GARCH models for examining how news from one international market influences other markets' volatility process. For stock markets, Hamao, Masulis and Ng (1990) use the GARCH-M model to show that volatility spillovers exist from New York to Tokyo, London to Tokyo and New York to London. For currency markets, Engle, Ito and Lin (1990) use a GARCH model to find that Japanese news has the largest impact on the volatility spillovers of the yen/dollar exchange rates.

In the context of fixed income markets, Tse and Booth (1996) use US Treasury bill and Eurodollar futures to investigate volatility spillovers between US and Eurodollar interest rates. A bivariate EGARCH model that allows for the asymmetric volatility influence of the interest differential between markets (Eurodollar minus Treasury rate or the TED spread) as well as that of the domestic market, is used to analyse the volatility spillovers between markets. The results show that although the cross-market volatility effects are insignificant, the lagged TED spread is the driving force of the volatility process.

Eom, Subrahmanyam and Uno (2002) analyse the transmission of credit risk between Japanese yen and U.S. dollar interest rate swap markets between 1990 and 2000. Although they observed low correlations between yen and dollar interest rate swap spreads, they found that dollar interest rate swap spreads "Granger-cause" the changes in the yen swap spreads, for the 10-year maturities. Using a GJR-GARCH model to capture the asymmetric effects in the volatility process, they show that there is a strong transmission of volatility from the dollar swap spread to the yen swap spread.

The methodology used in this paper follows that originally employed by Hamas, Masulis and Ng (1990) and also draws on the framework adopted by Eom, Subrahmanyam and Uno (2002).

The motivation for this paper is driven by the consideration that a comprehensive study on the linkages between euro and sterling swap markets has not been undertaken so far. An investigation of the euro and sterling swap markets would promote a better understanding of the degree of integration, if any, between the fixed income segment of their respective financial markets. The flow of information between financial markets is an issue that has attracted considerable attention in the financial economics literature. Research in this area examines the extent to which a price shock in one market affects returns and volatilities in other markets. But most of these studies focus on inter-linkages between equity markets rather than fixed income markets. Although a lot of research has been devoted to the determinants of swap spreads, the issue of international linkages between them has not been so well addressed.

The paper is organised as follows. Section II provides a description of the data used and makes some inferences on the term structure of euro and sterling swap spreads. Section III attempts to trace the variability in these swap spreads to important economic events marking the euro and sterling fixed income markets. Section IV examines the contemporaneous and causal relationship between euro and sterling interest rate swap spreads. Section V estimates the volatility in euro and sterling swap spreads and investigates the possibility of volatility spillovers between these markets. Section VI concludes the paper.

II DATA AND SUMMARY STATISTICS

The euro swap rates used in this study are quoted rates from the fixed interest branch of a generic interest rate swap of 2-, 3-, 5-, 7-, and 10-years. Daily quoted rates were obtained from *Datastream* which are the average of bid and ask rates. These data cover the period from January 29, 1999 to March 28, 2003. The euro swap spread is calculated by subtracting the swap rate from constant maturity yields of German government bonds with corresponding maturities, which were also obtained from *Datastream*. The dataset consists of 218 weekly observations and 1086 daily observations.

TABLE 1
Summary Statistics of the Euro Swap Spreads

Euro swap spreads defined as the difference between euro swap rates
and constant maturity yields of German sovereign bonds with the corresponding maturity.
Panel A provides the mean, standard deviation, skewness, kurtosis and ADF test for
Non-Stationarity where the critical t-ratio at the 5% level of significance is -2.87.
The tests for integration of order zero or, or I(0), are carried out on the levels of the variables
and the tests for integration of order one, or I(1), are carried out on their first differences.
Daily data are used from 29 January 1999 to 28 March 2003 (total 1086 observations).
Panel B provides the same summary statistics for weekly Euro swap rates
with 218 observations.

Panel A: Daily Observations

Maturity	Mean	Std. Dev	Skewness	Kurtosis	ADF t-stat for I(0) test	ADF t-stat for I(1) test
2 year	0.16	0.06	0.38	3.03	-2.04	-18.36
3 year	0.20	0.07	0.38	2.75	-2.36	-17.99
5 year	0.24	0.10	0.59	2.33	-2.28	-21.23
7 year	0.29	0.12	0.60	2.36	-1.67	-20.80
10 year	0.37	0.15	0.32	2.06	-1.47	-22.08

Panel B: Weekly Observations

Maturity	Mean	Std. Dev	Skewness	Kurtosis	ADF t-stat for I(0) test	ADF t-stat for I(1) test
2 year	0.16	0.06	0.46	2.85	-2.41	-16.94
3 year	0.20	0.07	0.41	2.67	-2.62	-17.19
5 year	0.24	0.10	0.57	2.34	-1.93	-16.94
7 year	0.29	0.12	0.61	2.39	-1.41	-16.86
10 year	0.38	0.15	0.32	2.09	-1.14	-15.92

Table 1 Panel A reports the summary statistics for the daily euro swap spreads on yield basis. Panel B provides the same statistics for the weekly observations in the euro swap spreads. As the table shows, the average spreads of the euro interest rate swaps over the corresponding German government bonds is upward sloping with maturity. The standard deviations of swap spreads increase as the swap maturity increases. Symmetric distributions, such as the normal distribution have a skewness of zero. Kurtosis measures the thickness of the tails and is equal to 3 for a normal

distribution. Euro swap spreads show positive skewness across the term structure but, relatively close to normal kurtosis for the lower maturities. The Augmented Dickey-Fuller (ADF) test is performed to determine whether the various time series of swap rates are non-stationary. This is based on the null hypothesis of non-stationarity. The ADF statistics show that we cannot reject the null-hypothesis at the 5% level of significance. This suggests that euro swap spreads across all maturities are non-stationary.

Table 2 Panel A provides the summary statistics for the sterling swap spreads on a daily basis. Panel B provides the same statistics for the weekly sterling swap rates. As the table shows, the average sterling interest rate swaps slopes upward initially and then flattens out. It is interesting to note that the average swap spreads of sterling interest rate swaps are much larger than those of euro interest rate swaps. This difference can be accounted for by several factors and is discussed in the following section. The average standard deviations of the sterling swap spreads are also larger than those of the euro swap spreads for all maturities. We reject the stationarity of the sterling swap spread and conclude that they follow a random walk.

TABLE 2
Summary Statistics of the Sterling Swap Spreads

Sterling spreads defined as the difference between sterling swap rates and constant maturity yields of UK Treasury bonds with the corresponding maturity. Panel A provides the mean, standard deviation, skewness, kurtosis and ADF test for Non-Stationarity where the critical t-ratio at the 5% level of significance is -2.87. The tests for integration of order zero or, or I(0), are carried out on the levels of the variables and the tests for integration of order one, or I(1), are carried out on their first differences. Daily data are used from 29 January 1999 to 28 March 2003 (total 1086 observations). Panel B provides the same summary statistics for weekly pound swap spreads with 218 observations.

Panel A: Daily Observations

Maturity	Mean	Std. Dev	Skewness	Kurtosis	ADF t-stat for I(0) test	ADF t-stat for I(1) test
2 year	0.41	0.11	0.15	2.44	-2.13	-23.27
3 year	0.54	0.15	0.20	2.46	-1.21	-23.64
5 year	0.61	0.18	0.03	1.82	-0.93	-22.98
7 year	0.62	0.22	0.04	1.81	-0.77	-22.44
10 year	0.65	0.27	0.20	2.00	-0.99	-35.50

Panel B: Weekly Observations

Maturity	Mean	Std. Dev	Skewness	Kurtosis	ADF t-stat for I(0) test	ADF t-stat for I(1) test
2 year	0.40	0.11	0.05	2.33	-1.53	-18.16
3 year	0.54	0.15	0.16	2.38	-0.90	-17.84
5 year	0.61	0.18	0.01	1.82	-0.63	-18.90
7 year	0.62	0.22	0.05	1.81	-0.45	-18.33
10 year	0.65	0.28	0.20	2.00	-0.41	-18.10

III DEVELOPMENTS IN SWAP SPREADS

This section focuses on developments in the sterling and euro swap spreads. Figure 1 shows a time series of 10-year euro and sterling swap spreads using daily observations from January 29, 1999 to March 28, 2003. The figure depicts that the sterling swap spreads were perceptibly wider than euro swap spreads since the launch of the single currency. During the period of observation, the average sterling swap spread was 65.07 basis points as compared to 37.28 basis points for the euro swap spread.

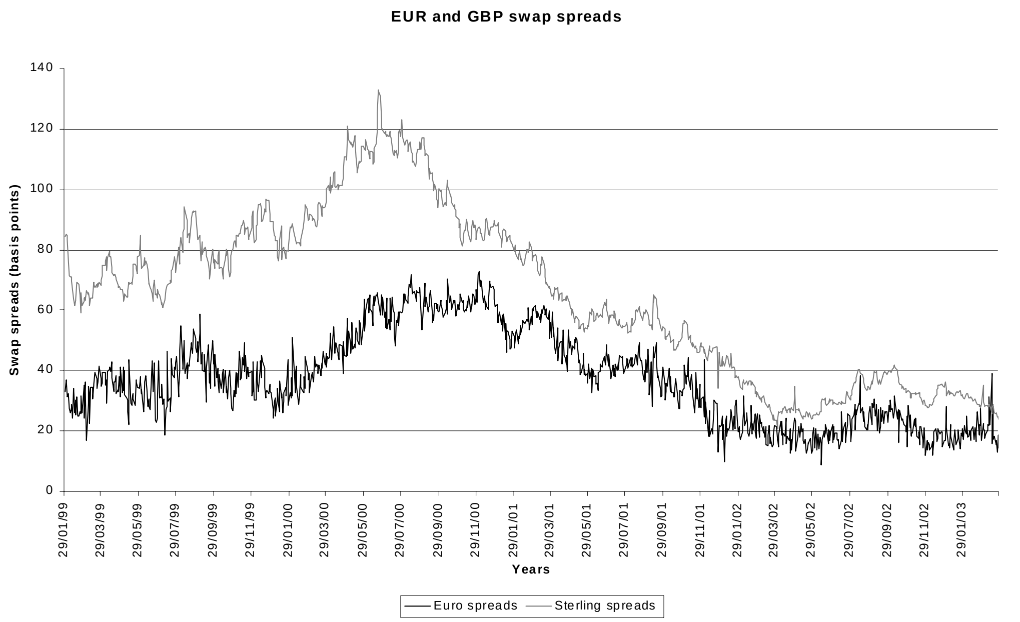

Figure 1 Euro and Sterling Euro Swap Spreads

Although a number of factors may be cited to explain this divergence, the most significant relates to the issuance of bonds by British and other European government bond markets as necessitated by their differing budgetary positions. While in the UK budget surpluses caused the net issuing volume of Treasury bonds to decline in 1999 and 2000, in Europe the issuing activity of governments remained stable due to persistent budget deficits. In the UK, the scarcity of bonds led to a decline in bond yields, causing swap spreads to widen significantly. Cooper and Scholtes (2001) examined the link between swap spreads and net supply of government bonds in the UK and US markets. The results were mixed. In both markets, a very simple regression between these variables suggested a strong negative relationship. But when they incorporated other variables, in particular the slope of the yield curve, net issuance ceased to be statistically significant.

Brooke, Clare and Lekkos (2000) have cited a number of UK-specific supply and demand-side factors that have influenced the shape of the gilt yield curve over the few years prior to 2000. On the supply side, net borrowing by the UK government had been negative between 1998 and 2000 and the outstanding stock of gilts had, therefore, contracted. The heavy demand for gilts from pension funds and insurance companies increased strongly during that phase causing further downward pressure on government bond yields. Pension funds were obliged to buy gilts to comply with the Minimum Funding Requirement (MFR) of the Pensions Act, 1995, designed to ensure that pension fund managers do not take excessive risks with their investments.

Moreover, as yields continued to decline markedly, the UK treasury yield curve inverted. As an illustration, Figure 2 shows the inverted nature of the UK Treasury yield curve on April 28, 2000 where the spot and forward interest rates have been estimated using the Nelson and Siegel (1987) model. [3]

[3] The Nelson and Siegel (1987) model has been used to estimate the zero-coupon yield curve of spot interest rates from observable coupon bonds. Market data on bond prices, coupon rates and yield to maturity have been sourced from *Datastream.*

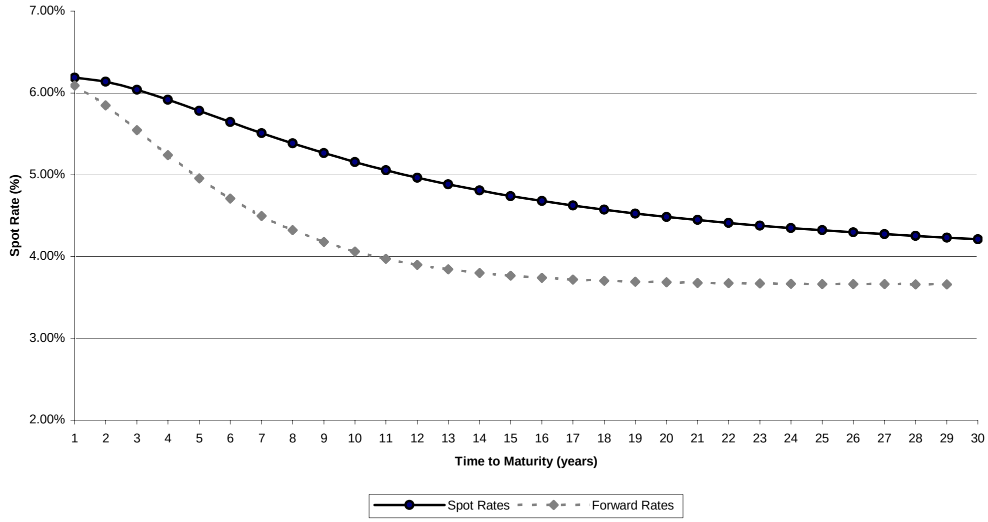

Figure 2. UK Treasury yield curve on April 28, 2000

Another factor that may have impacted on the yields on gilts relates to convergence plays associated with expectations about the United Kingdom joining EMU. Prior to the end of the year 2000 financial markets may have expected that the UK would adopt the single European currency in the near future. Although government bond markets in the Eurozone have remained segmented, integration has been particularly strong at the short-term end of the yield curve. As a result, there is only one short-term interest rate for all EMU member countries, set by the European Central Bank. Thus, a corollary of the UK joining EMU would be the eventual convergence of UK short-term interest rates to the levels prevailing in the Eurozone. According to the expectations theory of the term structure, there should be no expected difference in the returns from holding a long-term bond or rolling over a sequence of short-term bonds. Based on the premise that all bonds will generate a riskless return and ignoring liquidity premia, convergence in future short-term interest rates would entail convergence in long-term bond yields. Therefore, the activities of hedge funds and other market participants betting on the convergence between gilt and bund yields would serve to reduce long-term gilt yields, further inverting the gilt yield curve.

By the year 2001, the UK budget position had moved away from surpluses to deficits with increased spending on public services. The consequent increase in the supply of gilts increased long-term bond yields. Following the release of the Myners' Report,[4] it was announced that the MFR would be abolished. Removing this artificial demand shifted pension fund investment away from gilts to UK corporate debt and with the consequent narrowing of the spread between 5-to 20-year gilts the yield curve flattened. With a high balance of opinion against EMU entry it became apparent that the prospect of the UK joining the single currency in the near future was remote. This may have also contributed to the straightening out of the long end of the yield curve.

Euro swap spreads did not widen to the same degree as did sterling swap spreads. However, a notable feature was the surge in issuance of corporate bonds denominated in euros since the introduction of the single currency. Although, the budgetary situation was not so comfortable in the main Euro-zone countries of France, Germany and Italy they all had upward sloping yield curves. Figure 3 shows the yield curve for German sovereign bonds on 28 April 2000.

[4] The Myners' Report was commissioned to identify the institutionalised obstacles distorting the investment process. In particular, Paul Myners was to determine what prevents the flow of long term savings into the growth points of the economy - namely, new ventures (private equity) and smaller companies.

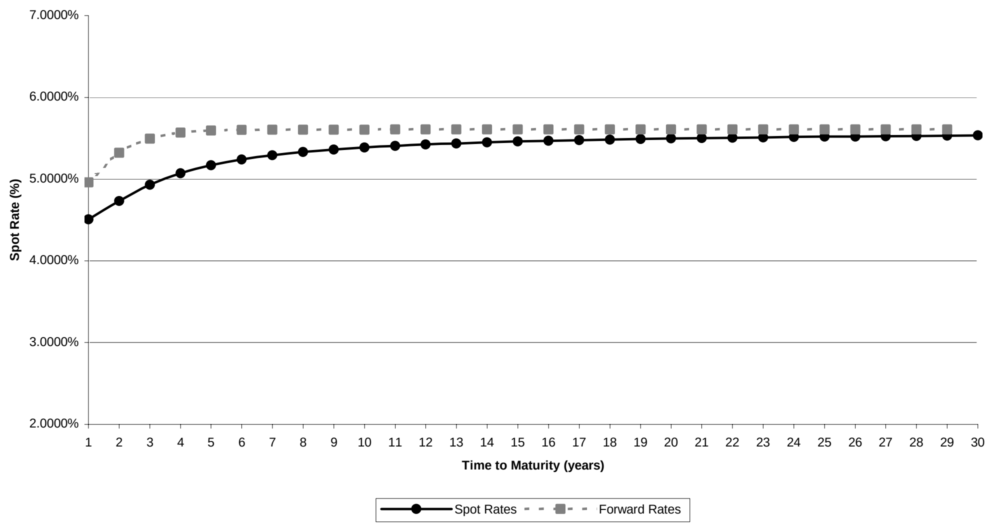

Figure 3. German government Yield Curve on 28th April 2000

From the year 2001 onwards sterling swap spreads trended lower and fell more sharply than euro swap spreads. The UK budget position had also moved away from surpluses to deficits with the increased spending on public services. The consequent increase in the supply of gilts coupled with the increased pension fund demand for UK corporate debt have acted as forces pulling sterling swap spreads lower. In the years 2002 and 2003 public sector borrowing requirements increased and the return to large-scale government debt issuance normalised the longer end of the sterling yield curves, while also helping to narrow spreads between government bonds and interest rate swaps.

The French, German, Italian, Spanish and Dutch governments have all used swaps to reduce the average maturity of their debt. [5] When the swap spread widens, governments find it attractive to receive fixed in the swap market. However, the large

[5] BIS Quarterly Review, March 2003

budget deficits in the main euro-zone countries of France, Germany and Italy have resulted in a narrowing of the spread between euro swaps and their respective government bonds in 2001 and 2002.

IV RELATIONSHIP BETWEEN SWAP SPREADS

This section examines the relationship between euro and sterling swap spreads. Table 3 shows the correlation coefficients between the changes in euro swap spreads, the changes in sterling swap spreads, and the changes in interest rate differentials between the U.K. and Germany. As indicated in preceding section, both the euro and sterling swap spreads are non-stationary. Correlations between such time series data can be partly spurious if they exhibit consistent trends. However, both the variables are stationary if first differences are considered. So in order to avoid spurious correlations, the correlations are analysed for the first differences in these variables and not their levels. Given that swap spreads are a measure of interbank risk and the fact that most international banks have global operations it would be reasonable to expect swap spreads in euros and sterling to be highly correlated. But the coefficients in Table 3 reveal that this correlation is negligible. The correlation between across 2-10 year vertices ranges from -0.04 to 0.17.

TABLE 3
Correlation between Euro and Sterling Swap Spreads

The table indicates the correlation coefficients among changes in the euro interest swap spreads, EURsp, changes in Sterling swap spreads, and the changes in interest rate differentials between UK and Germany (UK-GER).The interest rate differentials are given by the constant maturity yields of government bonds with the same maturity as the swaps.

Maturity	Corr(EURsp, GBPsp)	Corr(EURsp,UK-GER)	Corr(GBPsp,UK-GER)
2 year	-0.04	0.37	0.08
3 year	0.01	0.46	0.22
5 year	0.05	0.55	0.15
7 year	0.08	0.57	-0.01
10 year	0.17	0.56	-0.1

However, the first differences in euro swap spreads are more correlated with the first differences in interest rate differentials between sterling and euro-denominated government bonds. The correlation coefficient between euro interest rate swap spread and the interest rate differentials given by the differences in yields of constant maturity UK and German Treasury bonds has ranged form 0.38 to 0.57. But the sterling swap spread has displayed negligible correlation with these interest rate differentials as indicated by the correlation coefficients ranging from -0.01 to 0.22.

A possible explanation for the high correlation between the changes in euro interest rate swap spread and the interest rate differential is that arbitrageurs go long euro interest rate swaps and go short sterling interest rate swaps to construct a spread position between the government bonds in the two countries. Such a spread position is constructed to take advantage of the differential between the low long-term yields of German sovereign bonds and the high long term yields of UK gilts. Eom, Subrahmanyam and Uno (2000) came to a similar conclusion on observing that changes in yen swap spreads were correlated with the interest differentials between US and Japanese treasury bond yields.

Correlation is intrinsically a short-run measure of co-dependency and reflects the contemporaneous relationship between interest rate swap spreads. The analysis of correlation is significant, in terms of depicting the degree of integration between the swap markets. Additionally, a lead-lag relationship can also be expected if there is some degree of co-dependency in interest rate swap markets. Vector autoregressive models can be used to investigate any lead-lag behaviour between interest rate swap spreads. Granger causality tests are then conducted to see if lagged changes in the spreads for sterling interest rate swaps cause changes in the spreads of euro interest rate swaps.

To illustrate this, let x_t be the first differences in 10-year euro swap spreads and let y_t be the first differences in 10-year sterling swap spreads. Consider the bivariate VAR(2) model:

$$x_t = c_1 + a_{11}x_{t-1} + a_{12}x_{t-2} + b_{11}y_{t-1} + b_{12}y_{t-2} + e_{1t} \qquad (1)$$

$$y_t = c_2 + a_{21}x_{t-1} + a_{22}x_{t-2} + b_{21}y_{t-1} + b_{22}y_{t-2} + e_{2t} \qquad (2)$$

The test for Granger causality from *x* to *y* is an *F*-test for the joint significance of a_{21} and a_{22}, in an OLS regression. Similarly, the test for Granger causality from *y* to *x* is an *F*-test for the joint significance of b_{11} and b_{12}.

Using 216 weekly observations over the sample period from January 29, 1999 to March 28, 2003, each equation has been estimated separately using OLS. Table 4 shows the results of the estimation.

TABLE 4

Bivariate VAR(2) Model using first differences in 10-year swap spreads

Sample period: January 29, 1991 - March 28, 2003

Equation (1)			Equation (2)		
F-statistic	20.07		F-statistic	2.53	
	Coeff.	*t*-stat		Coeff.	*t*-stat
c_1	-0.0014	-0.37	c_2	-0.0035	-1.11
a_{11}	-0.6002	-8.70	a_{21}	0.0569	0.98
a_{12}	-2.2604	-3.79	a_{22}	0.0018	0.03
b_{11}	0.2707	3.15	b_{21}	-0.2280	-3.07
b_{12}	0.0513	0.59	b_{22}	-0.0334	-0.45

The *t*-statistics (in parentheses) indicate that the model coefficients are more significant where the dependent variable is the change in 10-year euro swap spread. The $F_{4,211}$ statistic for goodness of fit is 20.1 for the euro swap spread equation, and this is significant at the 5% level ($F_{4,211}$= 2.37). The *F*-statistic for the euro swap spread to sterling swap spread causality is only 2.53. Although this is just about significant at the 5% level, it is much weaker than the causality from the sterling to euro swap spreads. The results indicate that last week's changes in the 10-year sterling swap spread can have a predictive impact on this week's changes in 10-year euro swap spreads.

TABLE 5
Co-dependency between Euro and Sterling Swap Spreads

The table represents the results of bivariate "Granger causality" tests among changes in euro swap spreads (EURsp), changes in sterling swap spreads (GBPsp) and the lagged changes in interest rate differentials between the euro and the sterling (UK-GER). The numbers in the table are values of the F-statistic of the Granger causality test which have been performed for 2 lags.
Weekly data of changes in swap spreads are from 29 January 1999 to 28 March 2003 providing for a total of 219 observations. The quotations of swap rates and the constant maturity government bond yields were obtained from Datastream.

Maturity	EURsp to GBPsp	GBPsp to EURsp
2 year	4.31717	0.45855
3 year	4.5514	1.25617
5 year	3.23698	1.5803
7 year	1.84414	3.70722
10 year	0.61348	4.97335

Maturity	EURsp to UK-GER	UK-GER to EURsp
2 year	3.8187	0.83146
3 year	4.62935	0.36184
5 year	6.5062	0.81087
7 year	7.47418	0.06039
10 year	3.29147	0.76098

Maturity	GBPsp to UK-GER	UK-GER to GBPsp
2 year	0.91549	0.32706
3 year	0.32628	0.03602
5 year	0.38917	0.38917
7 year	1.6042	1.6042
10 year	0.41198	0.41198

Table 5 reports the Granger causality tests reflecting the lead-lag relationship among changes in euro and sterling swap spreads across the maturities under consideration. Granger causality tests are sensitive to the choice of the number of lags. These tests were performed using 2,3 and 4 lags which all produced qualitatively similar results. The results reported in Table 5 are for 2 lags. As revealed in the table, the nature of the causality depends on whether one is considering the short or long-end of the swap curve. The *F*-value of the Granger causality test for changes in the 10-year sterling swap spread to changes in the 10-year euro swap spread is 4.97, which is statistically significant at the 5% level. This indicates that lagged changes in the sterling swap spreads Granger cause changes in the euro interest swap spread at the 10-year maturity. But this causality is one-sided and does not transmit itself the other way. Lagged changes in 10-year euro swap spreads do not have any significant impact on changes in sterling swap spreads of the same maturity. A similar result emerges for the 7-year maturity. But at the short end of the swap curve the causality again reverses itself. At the 2-, 3- and 5-year maturities, euro-swap spreads Granger cause sterling swap spreads but the causality does not run the other way.

V VOLATILITY IN SWAP SPREADS

In this section we examine the dynamic behaviour of volatility in the euro and sterling swap spreads. We make use of a GARCH framework to capture the time variation and persistence in volatility. The analysis is carried out on the 10-year swap spreads in euro and sterling markets using daily observations over the period January 29, 1999 to March 28, 2003.

GARCH Models

The GARCH (p,q) model expresses the conditional variance of a given time series (σ_t^2) as a linear function of p lagged squared errors and q lagged variances.

$$\sigma_t^2 = \omega + \alpha_1 \varepsilon_{t-1}^2 + ... + \alpha_p \varepsilon_{t-p}^2 + \beta_1 \sigma_{t-1}^2 + ... + \beta_q \sigma_{t-q}^2 \tag{3}$$

$$\omega > 0, \alpha_1, ..., \alpha_p, \beta_1, ..., \beta_q \geq 0$$

Since estimation is difficult for anything other than low values of p and q, in practice the most frequent application is the GARCH (1,1) model. In the context of our analysis , the GARCH (1,1) model would consist of two equations:

$$y_t = c + x_t \delta + \varepsilon_t \qquad \varepsilon_t / I_{t-1} \sim N(0, \sigma_t) \tag{4}$$

$$\sigma_t^2 = \omega + \alpha \varepsilon_{t-1}^2 + \beta \sigma_{t-1}^2 \qquad \omega > 0, \alpha, \beta \geq 0 \tag{5}$$

where in the conditional mean equation (4), y_t represents the swap spread at time *t*, and ε_t their unanticipated component distributed independently over time and assumed to follow a normal distribution with zero mean and conditional variance σ_t^2, x_t as the swap spread of the other currency. The conditional variance equation (5) is a function of the constant term, ω ; news about volatility from the previous period, measured as the lag of the squared residual from the conditional mean equation, ε_{t-1}^2; the previous period's forecast variance, σ_{t-1}^2.

Testing for ARCH effects

Various methods are available to test for the existence of autoregressive conditional heteroskedasticity (ARCH). A test based on the Lagrange multiplier (LM) principle formulated by Engle (1982) is applied here. Let y_t denote the swap spread of one country at time *t* and x_t the swap spread of the other country at time *t*. The process begins by running an OLS regression of y_t on x_t of the following form:

$$\hat{y}_t = a + bx_t \qquad (6)$$

Now the residuals from this preliminary OLS estimation can be tested for ARCH behaviour. The test proposed in Engle (1982) is to regress the squared residuals, e_t^2 (where $e_t = y_t - \hat{y}_t$) on a constant and *p* lagged values of the squared residuals:

$$e_t^2 = \hat{\alpha}_0 + \hat{\alpha}_1 e_{t-1}^2 + \ldots + \hat{\alpha}_{t-p}^2 + \upsilon_t \qquad (7)$$

where υ_t is the error term.

From the results of this auxillary regression in residuals, the LM test statistic is calculated as $(T\text{-}p)*R^2$ where T is the number of observations. As explained in Bollerslev (1986), the LM statistic has an asymptotic chi-square (χ^2) distribution with p degrees of freedom under the null hypothesis of no ARCH effects. If the LM statistic, evaluated under the null hypothesis exceeds the critical value from a chi-square distribution with q degrees of freedom, the null hypothesis is rejected.

The results of the auxillary regression, for one lag, are shown in Table 6. There is strong evidence to reject the null hypothesis of no ARCH effects as the LM test statistic of 519.14 that it returns, far exceeds the critical value of $\chi^2_{0.95}(1)$ = 3.84. A regression residual series was generated by increasing the number of lags to five. But the results for more lag lengths were not qualitatively different from that obtained for one lag and are not reported here.

TABLE 6
ARCH LM Tests on 10-year swap spreads

Sample period: January 29, 1991 - March 28, 2003
Included observations: 1085 after adjusting end points

	Euro Swap Spreads		Sterling Swap Spreads	
	Coeff.	*t*-stat	Coeff.	*t*-stat
α_0	0.0023	8.32	0.0089	10.008
α_1	0.6915	31.52	0.6292	26.67
F-stat	993.59	(0.0000)	711.47	(0.0000)
LM-stat	519.14	(0.0000)	430.18	(0.0000)

Figures in parenthesis show probabilities

The 10-year sterling swap spreads demonstrated similar ARCH effects where the squared residual series for one lag returned an LM test statistic of 430.18. Increasing the number of lags to five did not change the results in so far as the existence of ARCH was concerned.

Testing for an asymmetric effect on volatility

Several studies on the volatility dynamics of asset markets have shown evidence of asymmetry in the response of conditional variances to the type of news revealed to the markets. This is also referred to as the leverage effect in volatility and is often observed in equity markets where downward movements in the market are followed by higher volatilities than upward movements of the same magnitude. In the context of swap spreads the leverage effect would arise if, for instance, the volatility of the swap spread increases more when there is a positive shock, which increases the swap spread, than when there is a negative shock.

The GARCH model specified in equation (5) cannot capture any asymmetric effect, since the conditional variance is a function only of the magnitudes of the lagged residuals and not their signs. The residuals ε_t are specified as a square and so it makes no difference whether they are positive or negative.

In the exponential GARCH (EGARCH) model of Nelson (1991), σ_t^2 depends on both the size and the sign of lagged residuals. The purpose of this EGARCH specification is to try and build in some asymmetry, so that the sign of ε_t matters. The conditional variance equation in the EGARCH model is defined in terms of the standard normal variate z_t :

$$\ln \sigma_t^2 = \omega + g(z_{t-1}) + \beta \ln \sigma_{t-1}^2 \qquad (8)$$

where $g(.)$ is an asymmetric response function defined by

$$g(z_t) = \gamma z_t + \alpha(/ z_t /- \sqrt{2/\pi})$$

The left-hand side of equation (8) shows the log of the conditional variance. This implies that the leverage effect is exponential, rather than the quadratic, and that the forecasts of the conditional variance are guaranteed to be nonnegative. The standard normal variable z_t is the standardized residual ε_t / σ_t. When $\alpha > 0$, and $\gamma < 0$ negative shocks to returns ($z_{t-1} < 0$) induce larger conditional variance response than positive shocks. Therefore, the presence of asymmetric effects can be tested by the hypothesis that $\gamma < 0$. The impact is asymmetric if $\gamma \neq 0$. Formulas for higher order lags in ε_t can be found in Nelson (1991).

To test for the possible existence of this leverage effect in swap spread volatility we applied the EGARCH model to standardized residuals of the conditional mean model using one swap spread as the dependent variable and the swap spread of the other currency as the exogenous variable. We applied the EGARCH model to the swap spreads of both currencies over the sample period. The results are shown in Table 7.

With the 10-year euro swap spread as the dependent variable and the corresponding sterling swap spread as the exogenous variable the asymmetric effect term (γ), is positive and equal to 0.0136. The z-statistic is equal to 0.98 which is not statistically different from zero at the 5% level of significance given by 1.645. We may, therefore, conclude that the volatility in 10-year euro swap spreads do not display asymmetric effects. Performing an identical operation with the sterling swap spreads as the dependent variable and the euro swap spread as the exogenous variable revealed similar results. The asymmetric effect term (γ) was again positive at 0.027. It was also not statistically significant from zero with the z-statistic equal to 0.92.

Eom, Subrahmanyam and Uno (2002) employed a GJR-GARCH model and found that there is an asymmetric volatility effect of dollar swap spreads on yen swap spreads, while the asymmetric effect of the shock on the yen swap spread is insignificant. In their analysis of the swap spreads in Australia, Brown, In and Fang (2002) used an EGARCH approach and found that the asymmetric effects are statistically significant for 3 and 5-year swaps but not for 10-year swaps.

With these tests demonstrating the absence of any asymmetric volatility effect of the shock on 10-year euro and sterling swap spreads, it would be appropriate to confine ourselves to symmetric GARCH models for modelling volatility.

TABLE 7
Testing for the Asymmetric Effect on Volatility

EGARCH Model:

$$\ln\sigma_t^2 = \omega + g(z_{t-1}) + \beta\ln\sigma_{t-1}^2$$

where $g(.)$ is an asymmetric response function defined by

$$g(z_t) = \gamma z_t + \alpha(/\, z_t \,/ - \sqrt{2/\pi})$$

	Euro Swap Spread			Sterling Swap Spread		
	Coeff.	*z*-stat.	Prob.	Coeff.	*z*-stat.	Prob.
EGARCH Model						
Asymmetric effect parameter						
(H_0: $\gamma < 0$)	0.01	0.98	0.3255	0.027	0.93	0.3526

Estimating the GARCH (1,1) model

To assess the appropriateness of the GARCH specification for daily swap spreads, a GARCH (1,1) model based on equations (4) and (5) is used. This specification was found to be the most appropriate for modelling volatility in both euro and sterling 10-year swap spreads.

The model specification also includes a dummy variable for the trading day following a weekend, i.e. Monday, in the conditional variance equation to capture potential "day of the week" effects. The model now has the following form:

$$y_t = c + ax_t + \varepsilon_t \tag{9}$$

$$\sigma_t^2 = \omega + \alpha\varepsilon_{t-1}^2 + \beta\sigma_{t-1}^2 + \delta D_t \tag{10}$$

where D_t represents a dummy variable that takes the value of 1 on Mondays and is 0 otherwise. Panel A of Table 8 shows the results of the estimation of the GARCH(1,1) model for euro-swap spreads. There are no indications of any serious model misspecification.

TABLE 8.
Estimation of GARCH(1,1) model using 10-year swap spreads

Sample period: January 29, 1991 - March 28, 2003

	PANEL A		**PANEL B**	
	Euro Swap Spread		Sterling Swap Spread	
Number of obs.	1086		1086	
Log-likelihood	1408.946		897.005	
	Coeff.	*z*-stat	Coeff.	*z*-stat
Conditional Mean				
c	0.065771	17.46	0.085852	17.08
a	0.441450	70.72	1.204970	84.22
Conditional Variance				
ω	0.000282	3.77	0.000513	4.07
α	0.182550	6.17	0.310197	4.92
β	0.799298	30.53	0.661948	10.53
δ	-0.000838	-2.61	-0.001146	-2.75
Residual Tests	Statistic	Prob.	Statistic	Prob.
Skewness	-0.17		-0.84	
Kurtosis	2.71		3.21	
Jarque-Bera	8.98	(0.011249)	129.78	(0.000000)
LM test statistic	0.004	(0.947156)	2.49	(0.114284)

The parameter estimates for the conditional variance equation (10) correspond to α = 0.1826, β = 0.7993, ω = 0.000282 and δ = -0.000838. The z-statistics reveal that all coefficients are statistically significant. The coefficient of the dummy variable is negative indicating the influence of more subdued trading in government securities on a Monday.

If we put $\sigma_t^2 = \sigma^2$ for all t in equation (10) above, we get an expression for the long-term steady state variance in a GARCH (1,1) model:

$$\sigma^2 = \omega / (1 - \alpha - \beta) \qquad (11)$$

Equation (11) can then be rewritten as:

$$V = \omega / \gamma \qquad (12)$$

where V is the long-term variance which can be calculated as ω/γ. A stable GARCH (1,1) process requires that the sum $\alpha + \beta$ be less than 1. Only then will the GARCH volatility term structures converge to a long-term average level of volatility that is determined by (12). In this estimation the sum of the α and β is equal to 0.981848 which is less than one indicating that volatilities of the 10-year euro swap spread converge to some long-term average level of volatility.

Since $\gamma = 1 - \alpha - \beta$, it follows that γ = 0.081152. And since $\omega = \gamma V$, it follows that V = 0.0155354. In other words, the long-run average variance per day implied by the model is 0.0155354. This corresponds to a volatility of $\sqrt{.0155354}$ = 0.1246414 or 12.46 % per day.

The residual tests display descriptive statistics of the standardised residuals, ε_t / σ_t. Under the null hypothesis of a normal distribution, the observed value of the Jarque-Bera test statistic of 8.975 exceeds the critical value of $\chi^2_{0.95}(2) = 5.99$. So the standardised residuals are not normally distributed. However, we cannot reject the null hypothesis of no ARCH effects in the standardised residuals as the observed LM test statistic of 0.004 is well short of the critical value of $\chi^2_{0.95}(1)$ = 3.84. This clearly indicates that there are no ARCH effects left in the standardised residuals.

The same GARCH (1,1) was then employed to estimate volatility in the 10-year sterling swap spreads. Panel B of Table 3. shows the results of the estimation. In the conditional variance equation, $\alpha = 0.310197$, $\beta = 0.661948$, $\omega = 0.000513$ and $\delta = -0.001146$. As revealed by the z-statistics, all coefficients are statistically significant. The sum of the GARCH coefficients is given by $\alpha + \beta = 0.972145$, which being very close to one indicates that volatility shocks are quite persistent. The value of the coefficient α in the case of sterling swap spreads is much higher than what it is for euro swap spreads. Large GARCH error coefficients α mean that volatility reacts quite intensely to market movements, and so if α is relatively high and β is relatively low then volatilities tend to be more spiky. Using equations (11) and (12) above, the long-term variance V works out to 0.0184168. This means a volatility of $\sqrt{0.0184168}$ = 0.1357085 or 13.57% per day. So we find the volatility of the sterling swap spread to be somewhat higher than that of the euro swap spread.

In the case of both the euro and sterling swap spreads, the distribution of the standardised residuals does not follow a normal distribution. However, the distribution of euro swap spread residuals are relatively closer to a normal distribution, whereas the sterling swap spreads exhibit a much more asymmetric and considerably broader distribution. Accordingly, the sterling swap spreads are more volatile than their euro counterparts.

Volatility Spillovers

Having estimated the volatilities of both the euro and sterling swap spreads over the sample period the paper examines the possibility of a transmission of volatility between them. Although the GARCH (1,1) specification used above was descriptively accurate for estimating volatility in individual markets it did not incorporate the spillover effects from other markets. So it is necessary to introduce an exogenous variable into the conditional variance equation that captures the potential spillover effect from one market into the other. The squared residual from one market is interpreted as a "volatility surprise" and is included in the other market's conditional variance specification:

$$\sigma_t^2 = \omega + \alpha_1 \varepsilon_{t-1}^2 + \alpha_2 \xi_{t-1}^2 + \beta \sigma_{t-1}^2 + \delta D_t \qquad (13)$$

where ε_{t-1}^2 is the lagged squared residual of the domestic swap spread and ξ_{t-1}^2 is the lagged squared shock arising from the foreign market's swap spread.

The results of estimating this model for both the euro and sterling swap spreads are shown in Table 9

TABLE 9

Volatility spillovers between swap spreads

Sample period: January 29, 1991 - March 28, 2003

	PANEL A		**PANEL B**	
	Euro Swap Spread		Sterling Swap Spread	
Number of obs.	1086		1086	
Log-likelihood	1373.686		876.2183	
	Coeff.	*z*-stat	Coeff.	*z*-stat
Conditional Mean				
c	0.065911	17.02	0.084630	16.62
a	0.441040	61.92	1.209475	79.39
Conditional Variance				
ω	0.000300	3.91	0.000482	3.96
α_1	0.219476	5.79	0.307258	4.85
α_2	0.002997	2.04	0.006063	0.80
β	0.734628	19.78	0.661962	10.50
δ	-0.000766	-2.74	-0.001150	-2.85
Residual Tests	Statistic	Prob.	Statistic	Prob.
Skewness	-0.14		-0.84	
Kurtosis	2.64		3.21	
Jarque-Bera	9.27	(0.009724)	129.70	(0.000000)
LM test statistic	0.67	(0.412656)	2.58	(0.108109)

Panel A of Table 9 shows there is evidence of an element of volatility spillover from the sterling swap spreads to the euro swap spreads. The parameter estimate on the sterling swap spread volatility surprise ξ_{t-1} is positive and statistically significant at the 5% level. Therefore, the null hypothesis of no foreign volatility surprise is rejected at the 5% significance level, indicating that there are mild volatility transmissions from the sterling swap spreads to euro swap spreads. However, Panel B of Table 3.9 shows that there is no such volatility spillover from euro swap spreads to sterling swap spreads as the parameter estimate is not statistically significant. These volatility spillover effects are consistent with the findings on Granger causality tests for 10-year swap spreads in Table 6.

VI CONCLUSIONS

This paper empirically examines the case of market integration between euro and sterling swap spreads during the period January, 1999 to March, 2003. The swap spreads are determined by the difference between the swap rates and the constant maturity yields of government bonds with corresponding maturity. Euro swap spreads have been proxied using German sovereign bonds.

To begin with, the main characteristics of the term structure of swap spreads in both the euro and sterling markets were examined. Both swap spreads are non-stationary across the term structure and follow a random walk. However, sterling swap spreads have been perceptibly wider than euro swap spreads since the launch of the single currency. This largely relates to the net supply of government bonds in British and European markets as driven by their respective budgetary positions.

While in the UK, budget surpluses caused the net issuing volume of Treasury bonds to decline in 1999 and 2000, in the main European markets of France, Germany and Italy the issuing activity of governments remained stable due to persistent budget deficits. The sterling swap spreads subsequently trended lower due to the UK budget position moving away from surpluses to deficits and the shift in demand of UK pension funds from gilts to corporate debt.

The correlation coefficient between changes in euro swap spreads and changes in sterling swap spreads is negligible indicating that credit risk can be attributed country specific factors as opposed to global influences. However, the changes in euro swap spreads are correlated, to some degree, with changes in interest differentials between sterling and euro-denominated government bonds. But no evidence is found of sterling swap spreads being correlated with the interest rate differentials. A plausible interpretation for the correlation between the euro swap spread and the interest differential is that arbitrageurs go long euro interest rates swaps and go short sterling interest rates swaps to construct a spread position between the government bonds in the two countries. Such a spread is constructed to take advantage of the low long-term yields of German bunds and the high long term yields of UK gilts.

Granger causality tests, reflecting the lead-lag relationship among changes in euro and sterling swap spreads reveal that the causality depends on whether one is considering the short or long-end of the swap curve. Lagged changes in sterling swap spreads Granger cause changes in euro interest swap spreads at the 10-year maturity. But there is no evidence to suggest that euro swap spreads Granger cause sterling swap spreads at the 10-year maturity. But at the short end of the swap curve the causality again reverses itself. At the 2-, 3- and 5-year maturities, euro swap spreads Granger cause sterling swap spreads but there is no causality in the reverse direction.

The notion of market efficiency dictates that it should not be possible to predict swap spreads in one market using lagged information generated in another market. To the extent that lagged changes in the spreads for sterling interest rates swaps cause changes in the spreads of euro interest swaps, the latter could be characterised as being informationally inefficient. However, interest rate differentials between these two markets do not Granger cause swap spreads in either of the markets.

The analysis of the causal relationship between swap spreads was then extended to the dynamic behaviour of volatility in 10-year euro and sterling swap markets. The time series of both the euro and sterling swap spreads show volatility clustering and reveal strong ARCH effects. An EGARCH model was employed to test for the existence of any asymmetric response in the volatility of 10-year swap spreads. But the volatilities did not display asymmetric effects for either of the swap spread markets.

The GARCH (1,1) specification was found to be the most appropriate for modelling volatility in 10-year swap spreads for both the markets. Volatility shocks were found to be quite persistent in both the markets. But volatility in the sterling swap spreads reacted more intensely to market movements and were more volatile than their euro counterparts. However, both volatility term structures converged to a long-run average level of volatility.

The possibility of volatility spillover effects between 10-year euro and sterling swap spreads were also examined. There was evidence of mild volatility transmission from the sterling swap spreads to euro swap spreads but no spillover effects the other way round. This observation was consistent with the findings on Granger causality.

This investigation into the causal relationship between euro and sterling swap spreads could contribute to an understanding of the degree of financial market integration between the UK and the Eurozone. An awareness of the nature of volatility spillover across the markets could be of importance to economic policy makers from a financial stability perspective. Given our findings that there is no volatility transmission from the euro swap spreads to sterling swap spreads, it seems unlikely that a credit risk shock in the euro fixed income market would have a destabilising effect on the sterling fixed income market. However, the more general conclusions that can be drawn from this paper are somewhat tentative because of the limited period of observation.

REFERENCES

Bollerslev, T. (1986). Generalised autoregressive conditional heteroskedasticity. *Journal of Econometrics,* 31: 307-327

Bollerslev, T. (1987). A conditional heteroskedastic time series model for speculative prices and rates of return. *Review of Economics and Statistics,* 69: 542 - 547

Brooke M., A. Clare and I. Lekkos (2000), "A comparison of long bond yields in the United Kingdom, the United States and Germany", *Bank of England Quarterly Bulletin*. May 2000.

Brown, K., W.V. Harlow and D.J.Smith, (1994). An empirical analysis of interest swap spreads. *Journal of Fixed Income,* 3, 61-78.

Brown, R., In, F., Fang, V, (2002). Modelling the Determinants of Swap Spreads. *Journal of Fixed Income*, 12, 29-50.

Cooper, N., and Scholtes, N., (2001). Government bond market valuations in an era of dwindling supply. *BIS Papers, No.5: The changing shape of fixed income markets*.

Duffie, D., and K. Singleton, (1997). An econometric model of the term structure of interest-rate swap yields, *Journal of Finance,* 52, 1287-1321.

Edwards, F.R., (1999). Hedge funds and the collapse of Long-Term Capital Management. *Journal of Economic Perspectives,* 13, 189-210.

Engle, R.F., (1982). Autoregressive Conditional Heteroskedasticity with Estimates of the Variance of United Kingdom Inflation. *Econometrica*, 50, 987-1007.

Eom, Y.H., Subrahmanyam, M.G., and Uno, J. (2000). Credit risk and the yen interest rate swap market. Unpublished manuscript, *Stern Business School,* New York.

Eom, Y.H., Subrahmanyam, M.G., & Uno. J. (2002). Transmission of Swap Spreads and Volatilities in the Japanese Swap Market. *Journal of Fixed Income,* 12, 1-37.

Grinblatt, M., (1995). An analytic solution for interest-rate swap spreads. Working Paper, *UCLA, Anderson Graduate School of Management*.

Hamao, Y., Masulis, R., and Ng, V., (1990). Correlations in price change and volatility across international stock markets. *Review of Financial Studies,* 3, 281-307.

Hull, J.C. (2000). Options, futures and other derivatives. *Pearson Education Inc.*

Liu, J., Longstaff, F.A., and Mandell, R.E. (2002). The Market Price of Credit Risk: An Empirical Analysis of Interest Rate Swap Spreads. *Working Paper, UCLA.*

Nelson, R., and Siegle, F, (1987). Measuring the term structure of interest rates. *Journal of Business* 44, 19-31.

Remolona, E. M., Wooldridge, P.D., (2003) The euro interest rate swap market. *BIS Quarterly Review*, March 2003.

Sorensen, E.H. and T.F. Bollier., (1994). Pricing swap default risk. *Financial Analysts Journal*, 50, 23-33.

Sun, T.S., Sundaresan, S., Wang, C., (1993). Interest rate swaps - An empirical investigation. *Journal of Financial Economics*, 34, 77-99.

Tse, Y., and Booth, G.G., (1996). Common volatility and volatility spillovers between US and Eurodollar interest rates: Evidence from the futures market. *Journal of Economics and Business,* 48, 299-312.