The UK Stock Market’s Relationship with US and European Stock Markets: Is the UK Stock Market *Snuggling-Up* to the US - or to Europe?

by

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Using 27 years of data this paper considers short-run, bi-lateral, and long-run, linkages between the US, UK and nine European equity markets. Using time-varying parameter and multivariate cointegration modelling techniques, we show that since October 1992, UK stock market returns have become increasingly sensitive to perturbations in the US market relative to those in European markets. We also report evidence to suggest that over the long term stock markets are perfectly correlated and although deviations from this long-term trend do occur, with the possible exception of the US and UK market, they are rarely significant. For the US and UK markets, we found that the 1990s has tended to display a few small, but significant deviations, from the permanent component of stock prices. Overall, however, any gains from international diversification would tend to be short-lived.

Key Words: convergence; random-walk; transitory; persistence; Kalman Filter; multivariate cointegration; moving average representation; stochastic trend.

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1. INTRODUCTION

Historically, policy-makers and finance specialists have given considerable attention to the relationships between national stock markets. However, while the extent of world-wide ‘globalisation’, its source, and its implications for contagion and diversification issues, have long been a topic of professional and academic interest,

the extent of the sensitivity of returns to bi-lateral perturbations, has received somewhat less direct attention in the academic literature. While a number of authors have highlighted a tendency of the UK stock markets to move in tandem with the US stock market,

there has been little attempt to explicitly characterise the dynamic interactions between the UK, US and other markets and to compare these with the long-run trends in national markets.

We add to this strand of the convergence literature by taking a regional approach. First, we examine the extent to which movements in the UK stock market’s bi-lateral relationships with European stock markets are associated with movements in the US stock market, and with movements in other European stock markets. We explore the issue of bi-lateral linkages by the modification of an approach first adopted by Haldane and Hall (1991) in the analysis of exchange rates and which utilises a time-varying parameter model estimated using a Kalman Filter.

The UK-US-European bi-lateral return relationships are of interest not least because the relaxation of UK capital controls in 1979 along with European economic integration and the growth of European stock markets, may have weakened any polarisation of UK and US market movements, and hence the sensitivity of UK returns to US disturbances. As the thrust of European legislation over the last few
decades has been the reduction and ultimate elimination of barriers to free movement of goods and services within the European Union (EU), we might expect any UK-European return differentials (relative to UK-US return differentials) to be diminishing over time. Diminishing differentials can be justified as a response to the demand for a EU-wide pool of capital to invest in EU-wide investment projects and hence aid the development of a pan-European goods market.

Further, if movements in the UK stock market are *snuggling-up*\textsuperscript{iv} to European markets then it implies that the time may be right for an acceleration of merger activity between the UK stock market and its European neighbours - an issue which has received considerable attention in the financial press of late.\textsuperscript{v} If however, UK-European stock market convergence is still some way off, or indeed, if UK market returns are increasingly being associated with movements in the US stock market, then this has implications not only for UK-US contagion effects and the probability of a smooth transition to a single European-wide stock market, but for bi-lateral business cycle alignments and ultimately the UK’s entry to the European single currency.

While the above discussion is concerned with the bi-lateral dynamics of stock market convergence, and hence short-run criteria, we continue the analyses of international stock market behaviour by focusing on the stylised fact that the national stock markets follow an upward trend over long horizons. By characterising stock prices in terms of the importance of any non-random component, we inquire into the nature of this upward movement, and consider the extent to which non-deterministic trending behaviour arises from a single source. From a long-term investor’s perspective, if prices set in national stock markets share a single common trend, and there are no
transitory deviations from this trend, then no long-term gains can be made from international diversification. Further, such a result would imply national responses to a single international common factor – for example, a single common growth factor that, in the long-run, binds national equity markets together.

We begin in section 2 by setting out the empirical framework for this study and go on to discuss the data used and the preliminary statistics in section 3. In section 4 we report and interpret the estimation results and present conclusions on the dynamic bilateral and long-term US-UK-European relationships in section 5.

2. Empirical Framework

First, consider the following two relationships:

\[(r^{us}-r^{uk})_t = \alpha_1 + \beta_1 (r^{eur}-r^{us})_t + e_{1t}\]  
\[(r^{eur}-r^{uk})_t = \alpha_2 + \beta_2 (r^{eur}-r^{us})_t + e_{2t}\]  

where \(r^{us}\) is the continuously compounded return of the US stock market; \(r^{uk}\), the continuously compounded return of the UK stock market; \(r^{eur}\), the continuously compounded return of a European stock market; \(e_t\) is a random error term; and \(\alpha_i\) and \(\beta_i\) are the parameters of interest. If equations (1) and (2) were estimated using the standard OLS fixed parameter estimation, and if the UK stock market was fully integrated with the US stock market, then \((r^{us}-r^{uk})\) would be independent of \((r^{eur}-r^{us})\), and \((r^{eur}-r^{uk})\), \((r^{eur}-r^{us})\) would be perfectly correlated. If this were the case then we would expect the joint restriction of \(\beta_1 = 0; \beta_2 = 1\) to be satisfied. Conversely, if the UK stock market was perfectly pegged to the European stock market, then we would expect the joint restriction \(\beta_1 = -1; \beta_2 = 0\) to be satisfied (see Haldane and Hall (1991)).
However, previous research has indicated that correlations/covariances between national stock markets are not constant over time, and hence provide justification for allowing a gradual adjustment path for the temporal correlation coefficients between the pairs of bilateral spreads. We therefore estimate the following equations:

\[
(r_{us} - r_{uk})_t = \alpha_{1t} + \beta_{1t}(r_{eur} - r_{us})_t + \epsilon_{1t}
\]

(3a)

\[
\alpha_{1t} = \alpha_{1t-1} + \eta_{1t}
\]

(3b)

\[
\beta_{1t} = \beta_{1t-1} + \nu_{1t}
\]

(3c)

\[
(r_{eur} - r_{uk})_t = \alpha_{2t} + \beta_{2t}(r_{eur} - r_{us})_t + \epsilon_{2t}
\]

(4a)

\[
\alpha_{2t} = \alpha_{2t-1} + \eta_{2t}
\]

(4b)

\[
\beta_{2t} = \beta_{2t-1} + \nu_{2t}
\]

(4c)

where \( \eta_{it} \) and \( \nu_{it} \) are white noise processes with variances \( \sigma^2 \). Hence \( \alpha_{it} \) and \( \beta_{it} \) evolve according to a random walk, accommodating the possibility of large shock to the bilateral relationships (see, for example Harvey, 1990).

According to equations (3a) through (4c), if the strength of the UK’s relationship with the European market has increased over time we would expect \( \beta_{1t} \to -1; \beta_{2t} \to 0 \). Conversely, if UK’s relationship with US is becoming stronger, then we would expect \( \beta_{1t} \to 0; \beta_{2t} \to 1 \). Such relationships are linked by an identity: movements in one bilateral spread must necessarily be associated with movements in one or the other bilateral spreads in the system. Therefore, we have an adding up restriction of \( \beta_{2t} - \beta_{1t} = 1 \) which should be satisfied by the data, and will allow us to gauge how robust our procedure is to the weighting between the equations (3a-3c) and (4a-4c).

Further, and as Haldane and Hall (1991) point out, the \( \alpha_{it} \) are stochastic constants and therefore partial out all the systematic influences upon the US-UK relationship and
the European-UK relationship other than those resulting from the movements in the European-US relationship. Hence the procedure will offset any potential model misspecification problems although it will not infer causal linkages, nor will it proffer any economic explanation of what determines return spreads.

We then continue by asking if the upward trend in these markets is due to a stochastic trend. If this is the case, (log) stock prices are said to follow a random walk with drift process. There are however two implications of the random-walk process: the presence of a unit root and uncorrelated disturbances. Given these are both required for a random walk, and unlike much of the existing literature, we focus on both these aspects thus are able to support, or cast doubt on, the random walk hypothesis for international equity prices.

Our first step is to follow Engle and Granger (1987), by testing for unit roots in our data. We apply (the now well known) unit root test statistics developed by Phillips and Perron (1988), which, unlike Dickey-Fuller tests, do not assume independent and identically distributed errors. It is widely accepted however that Phillips-Perron statistics are likely to have low power in detecting all departures from a random walk (see for example, Cochrane, 1988; Lui and He, 1991; Fraser and MacDonald, 1992). As Fraser and MacDonald point out, a series with a unit root is equivalent to a series composed of a unit root and a stationary component and such tests do not give an indication of how important the random walk element is to the behaviour of the series. Cochrane (1988) states that tests for a unit root only, attempt to isolate the random walk component, thus “it is hard to tell a stationary series from a stationary series plus a very small random walk” (p. 895).
Hence, we also consider results from a heteroscedastic consistent variance ratio test (see for example Cochrane (1988) and Fraser and MacDonald (1992)) that focus on the uncorrelated increment aspect of a random walk. Such tests allow us to comment on both the extent and the characteristics of any observed departures from randomness and their implications.

The variance ratio test proposed by Cochrane (1988) and subsequently applied by Lo and MacKinlay (1988, 1989) was extended by Chow and Denning (1993) to control for the overall test size and to define relevant confidence intervals. As Chow and Denning (1993) point out, failure to control the joint-test size for the variance ratio estimates results in substantial incorrect rejections of the random walk null hypothesis (Type I errors).

Our maintained hypothesis is described by:

\[ \ln p_t = \mu + \ln p_{t-1} + \varepsilon_t \]  

(5a)

where, \( \ln p_t \) is the logarithm of the equity price series, \( \mu \) is an arbitrary drift term and \( \varepsilon_t \) is the error term with an expected value of zero. Under the random walk hypothesis the first differences of the series are uncorrelated at all leads and lags, hence changes in \( \ln p_{t+} \), as denoted by:

\[ \Delta \ln p_t = \mu + \varepsilon_t \equiv X_t - X_{t-1} \]  

(5b)

should be uncorrelated and hence innovations are unforecastable from past innovations (see Lo and MacKinlay, 1989, p. 206). This also implies an important property of the random walk described by (5a), which is that the variance of the random walk increments must be a linear function of the time interval. For example, where \( X \) are IID, the variance of bi-monthly increments must be twice as large as the
variance of monthly increments, while the variance of quarterly increments must be three times as large as the variance of monthly increments, and so on. Comparisons of the variance estimates can therefore provide a specification test of the random walk model and a means by which to characterise deviations from randomness.

Following Lo and MacKinlay (1988) and Chow and Denning (1993), if we denote \( q \) as a lag difference of \( X_t \) (where \( q \) can be any integer greater than unity), then, if the series is a random walk:

\[
\frac{1}{q} \left( \frac{\text{VAR} \left( X_{t+q} - X_t \right)}{\text{VAR} \left( X_{t+1} - X_t \right)} \right) = 1
\]

will hold asymptotically even with possible heteroscedastic increments. The Lo and MacKinlay heteroscedastic consistent variance ratio estimate minus one is calculated as

\[
\bar{M}_t(q) = \left[ \frac{\hat{\sigma}^2(q)}{\hat{\sigma}^2(1)} \right] - 1
\]

where

\[
\hat{\sigma}^2(q) \equiv \frac{1}{\omega} \sum_{i=q}^{n} \left( X_i - X_{i-q} - q\hat{\mu} \right)^2
\]

\[
\hat{\sigma}^2(1) \equiv \frac{1}{nq-1} \sum_{i=1}^{nq} \left( X_i - X_{i-1} - \hat{\mu} \right)^2
\]

and

\[
\omega = q(nq - q + 1) \left( 1 - \frac{1}{n} \right)
\]

The number of available return observations is \( nq \) and \( \hat{\mu} \) is the sample mean of
(X_t - X_{t-1}). Equation (8) is the variance of q-period returns scaled by \( \omega \), and equation (9) is the variance of the single-period return.

If a series is a pure random walk, the variance ratio minus one and described by (7) will be zero. If changes in \( X_t \) are positively autocorrelated, the numerator will be larger than the denominator and variances will grow faster than linearly, and if changes in \( X_t \) are negatively autocorrelated, the numerator will be smaller than the denominator and variances will grow slower than linearly (see Campbell, Lo and MacKinlay (1997)). The heteroscedastic consistent test statistic derived by Lo and MacKinlay is:

\[
Z(q) = \frac{M_{r}(q)[V(q)]^{1/2}}{N(0,1)}
\]

\[
V(q) = \sum_{j=1}^{q-1} \left( \frac{2(q-j)}{q} \right)^2 \hat{\rho}(j)
\]

\[
\hat{\delta}(j) = \frac{(nq) \sum_{i=j+1}^{nq} \left( X_{i} - X_{i-1} - \hat{\mu} \right)^2 \left( X_{i-j} - X_{i-j-1} - \hat{\mu} \right)^2}{\sum_{i=1}^{nq} \left( \left( X_{i} - X_{i-1} - \hat{\mu} \right)^2 \right)^2}
\]

where \( \hat{\delta}(j) \) is the heteroscedastic-consistent estimator of the autocorrelation function, \( \hat{\rho}(j) \), of \( \Delta X_i \) under the null hypothesis of zero autocorrelation, and \( V(q) \) is the asymptotic variance of \( M_{r}(q) \) and is the weighted sum of \( \hat{\delta}(j) \).

However, the random walk hypothesis requires that the variance ratio minus one for all the chosen aggregation intervals, q, be equal to zero not just those ratios corresponding to a specific aggregation value. Chow and Denning (1993) extended the Lo and MacKinlay (1988) methodology by testing the all the selected multiple
variance ratio estimates with unity. In order to control the test size, the authors consider a set of variance ratio estimates, \( \{ M_i(q_i) \mid i = 1,2,\ldots,m \} \), corresponding to a set of pre-specified aggregation intervals, \( \{ q_i \mid i = 1,2,\ldots,m \} \), where \( q_i \) is any integer greater than one with \( q_i \neq q_j \) for \( i \neq j \), and test a set of sub-hypotheses (of the random walk null hypothesis), \( H_{0i}: 1 + M_i(q_i) = 1 \) for \( i = 1,2,\ldots,m \). As any rejection of \( H_{0i} \) will lead to the rejection of the random walk null hypothesis, Chow and Denning (1993) let the largest absolute value of the test statistic be

\[
Z^*(q) = \max_{1 \leq i \leq m} |Z(q_i)|
\]  

(14)

They demonstrate that the confidence interval for the extreme statistic, \( Z^*(q) \), is defined as

\[
Z^*(q) \pm \text{SMM}(\alpha; m; \infty)
\]  

(15)

where \( \text{SMM}(\alpha; m; \infty) \) is the asymptotic critical value of the \( \alpha \) point of the Studentised Maximum Modulus (SMM) distribution with parameter \( m \) and \( \infty \) degrees of freedom and that the asymptotic joint confidence interval of at least 100(1-\( \alpha \)) percent for the set of variance ratios is

\[
\sqrt{N M_i(q)} \pm \left[ \sqrt{V(q_i)} \right]^{1/2} \text{SMM}(\alpha; m; \infty) \quad \text{for} \quad i = 1,\ldots,m.
\]  

(16)

where \( N \) is the sample size. The asymptotic SMM critical value can be calculated from the conventional standard normal distribution. Values of \( Z^* \) greater than the absolute SMM critical value indicates significant departures from random behaviour (see for example Poon (1996)).

Hence the variance ratio estimate, \( \overline{M}_i(q) \), with the Chow and Denning (1993) extension, is a long run measure as it utilises all of the autocorrelations of the changes in \( X_t \) and not just those close to the lags zero (which is the traditional way of
determining randomness) and also allows for the multiple comparisons of all the relevant variance ratios with unity. This may be potentially important because the patterns of medium and long autocorrelations may have an important bearing on whether the price series has a tendency to under or overreact to shocks. Further, and in contrast to standard unit root tests, $\bar{M}_r(q)$, can shed light on the extent to which the series deviates from random walk behaviour.

Given our results from the unit root and variance ratio tests detailed above, we utilise the Johansen (1988) multivariate cointegration tests to ask how many common stochastic trends, or equivalently, how many cointegrating vectors, there are in the stock markets in our sample. While Engle and Granger (1987) argue that the existence of cointegration in a speculative market necessarily implies a violation of market efficiency since information in past prices could be used to improve forecasts of current prices (that is, an error correction mechanism exists), it is now generally accepted that evidence of a cointegrating relationship among a vector of variables need not imply market inefficiency if equilibrium returns are time-varying due to perhaps the time-varying nature of a risk premium. The empirical analysis of this study is conducted under the maintained hypothesis of market efficiency.

Essentially, the Johansen maximum likelihood procedure provides a unified framework for the estimation of multivariate cointegrating systems based on the error correction mechanism of the VAR(p) model with Gaussian errors. The benefit of the Johansen technique (compared to the two-step procedure of Engle and Granger (1987)) is that it provides estimates of all the cointegrating vectors that exist within a vector of variables, and offers a test statistic for the number of cointegrating vectors,
which has an exact limiting distribution. This test may therefore be viewed as more
discerning in its ability to reject a false null hypothesis. The Johansen technique can
be briefly described as follows.

Assume the vector $X$ has an autoregressive representation with Gaussian errors, $\varepsilon_t$.

$$
X_t = A_1 X_{t-1} + A_2 X_{t-2} + \ldots + A_k X_{t-k} + \varepsilon_t
$$

where $t = 1,2,\ldots,T$ and $\varepsilon_t$ is a random error term. Equation (17) may be reformulated
into an error correction form:

$$
\Delta X_t = \sum_{j=1}^{k-1} \Gamma_j \Delta X_{t-j} + \Pi_j X_{t-j} + \varepsilon_t
$$

where $\Gamma_j$ is an $N \times N$ coefficient matrix equal to:

$$
- \sum_{j=1}^{k} \Pi_j
$$

and $\Pi$ is an $N \times N$ matrix equal to:

$$
\sum_{i=1}^{k} \Pi_j - I
$$

(\text{where } I \text{ is an identity matrix}) and whose rank determines the number of distinct
cointegrating vectors (or linearly independent combinations of $X_t$ that are stationary)
that exist between the variables in $X$. If $\Pi$ is of full rank there will be no
cointegration or long-run relationship amongst the elements of $X_t$. If however, $\Pi$ is of
reduced rank, $r$, ($r<n$) then there will exist $n \times r$ matrices, $\alpha$ and $\beta$, such that,

$$
\Pi = \alpha \beta'
$$

where $\alpha$ is the adjustment matrix and the columns of $\beta$ are the $r$ linearly independent
cointegrating vectors.
Johansen demonstrates that the likelihood ratio test statistic for the hypothesis that there are at the most $r$ distinct cointegrating vectors is,

$$
LR \text{ or Trace } = T \sum_{i=r+1}^{N} \ln(1 - \lambda_i) \quad (22)
$$

where $\lambda_{r+1}, \ldots, \lambda_N$ are the $n - q$ ($q < n$) smallest squared conical correlations between the $X_{t-k}$ and $\Delta X_t$ series (eigenvalues), corrected for the effect of the lagged differences of the $X$ process. Equation (22) is the Likelihood Ratio or Trace statistic and it tests the restriction $r \leq q$ against the unrestricted model $r \leq n$. An alternative test statistic is the $L$-max statistic, which is based on the comparison of at most $q$ cointegrating vectors, $r \leq q$, and the alternative, $r \leq q + 1$. This is denoted:

$$
L_{\text{max}} = T \sum_{i=r+1}^{N} \ln(1 - \lambda_{q+1}) \quad (23)
$$

The usefulness of this methodology in the current analysis essentially comes down to determining the rank of the matrix, $\Pi$. If $\Pi$ has rank, $r$, then there are $r$ cointegrating relationships between $X_t$ (in our case stock price indices), or, $n-r$ common stochastic trends, where $n$ is the number of stock indices in the sample. The stochastic trends are the linear combinations of the $X_t$, having the ‘common’ feature of not containing the levels of the error correction term in them (Gonzalo and Granger (1995)). In other words they are the permanent forces that create the non-stationary property of the data.

The number of cointegrating vectors reveals the extent of integration across stock markets. If $n-r = 0$, ($r=n$), we have full rank, or null matrix and the absence of any common stochastic trends. The implication here is that there are no stationary long-run relations amongst the random walk elements of stock prices and there are both
short-term and long-term gains to be made from international diversification. If \( n-r>1 \), this implies the existence of more than one common stochastic trend, and implies that long-run integration is not complete and that some long-terms diversification benefits may exist. If however, \( n-r=1 \) (\( r=n-1 \)), there is a single common stochastic trend suggesting the equity markets are perfectly correlated over long horizons. Hence, there are no long-run portfolio diversification gains to be made. This is not to say however, that short-term diversification gains cannot be made but such gains would be subject to the magnitude and enduring nature of any transitory deviations from the stochastic trend as well as the time horizon of the investor.

The procedures described above form the basis of the empirical framework of this paper. We turn now to a description of the data used in the study.

3. DATA AND PRELIMINARY STATISTICS

The data used in this study are sampled on a monthly basis over the period January 1974 to January 2001. The financial data were collected from DATASTREAM Global Indices series and consist of the market price indices of the following countries: Belgium, Denmark, France, Germany, Ireland Italy, Netherlands, UK, US, Spain, Sweden. The stock market indices are updated daily as soon as the closing prices for each market are received in London using Financial Times Actuary classifications.

The exchanges of Paris, Frankfurt, Milan, London and New York represent five of the world’s major centres for the trading and distribution of both domestic and international equities. France, Germany and Italy are viewed as ‘core’ European
economies having been members of the Exchange Rate Mechanism (ERM) since its inception in 1979, and the European common currency – the Euro - since 1999. While, like the UK, Denmark and Sweden are members of the EU, they are not members of the Euro.\textsuperscript{xi}

Continuously compounded returns were calculated as:

\[ \ln P_t - \ln P_{t-1} \tag{24} \]

and therefore comprise of the capital gain component of market returns.\textsuperscript{xii}

\section*{INSERT TABLE 1}

Table 1 provides summary statistics on the sample of market returns. The stock market with the highest return and associated \textit{ex post} risk, as measured by return standard deviated, is the Italian market. The German market has the lowest return while the Netherlands has the lowest return standard deviation over the sample period. Not surprisingly the J-B tests for normality all suggest that continuously compounded returns are highly non-normal.

\section*{INSERT TABLE 2}

Table 2a through 2d provide the contemporaneous correlations between the monthly returns of the various markets for the full sample period and for three sub-sample periods. These correlations show that the degree of association between markets has not been constant over the period. In particular, the correlation of US and UK stock markets declined from 0.942 during 1974-1983 to 0.785 during 1984-1993, and rose to 0.804 over the period 1994-2001 – with the correlation over the full sample period being 0.880. Typically, the contemporaneous correlation between the US market and
the European markets has been falling over the three sub-samples, while the
correlation of the UK market with European markets has tended to fall in the last eight
years of the sample. These results suggest that focusing only on the simple
correlations of international equity prices and returns, as opposed to their adjustment
path and long-term behaviour, can be misleading: such measures of integration are
taken at a point in time over relatively short periods and do not explicitly reflect the
changing relationships between stock markets and the speed of adjustment to any
common permanent component or stochastic trend.

4. EMPIRICAL EVIDENCE

4.1. Short-Run Convergence

Figures 1A through 1I show the convergence measures as described by equations (1)
through (4c).

INSERT FIGURES 1A THROUGH 1I

Recall that our task here is to examine the extent to which movements in the UK stock
market’s bi-lateral relationships with European stock markets are associated with
movements in the US stock market, and with movements in European stock markets.
In this respect both $\beta_1$ and $\beta_2$ show the extent of UK stock market return convergence
with either European markets or with the US market. Recall that if the UK is
converging with the US, $\beta_1$ ($\beta_2$) will tend to zero (unity) but if UK is converging with
European markets, $\beta_1$ ($\beta_2$) will tend to unity (zero). First, the adding-up constraint
discussed above in section 2 is satisfied by the data: in every case we could not reject
the hypothesis that $\beta_2-\beta_1 = 1$. The estimation procedure would therefore appear to
be robust to the weighting between equations (3a-3c) and (4a-4c). In figures 1A
through II, we show $\beta_2$ values only since the $\beta_1$ values show almost identical information.

Second, for all markets, the monthly convergence process has been far from constant and, third, the broad pattern of the UK’s estimated bi-lateral relationships is similar. Pre 1984, we see, if anything at all, greater UK convergence with the US across the sample but between April and December 1985, with the possible exception of Ireland, we witness a definite move toward European markets.

Further investigation has shown that late 1983 saw the US$ appreciation accelerate and its value peak against all other currencies early in 1985. This, in turn, would have the effect of increasing the uncertainty surrounding expected future (net) returns from US investments thus rendering them less attractive to foreign investors. The deregulation of the UK stock market (October 1986) however, is associated with convergence with US but the stock market crash of October 1987 would appear to have ended this along with the fact that sterling was relatively strong against European currencies at this time – sterling was stable at around DM3 over the period 1987-1988.

The most significant reversal however occurs in 1992. Prior to this, the UK had joined the EMS (October 1990) and while this had encouraged a strengthening of UK’s sensitivity to European perturbations, it was short-lived. The sudden shift in the UK’s bi-lateral convergence process from Europe to the US is clearly associated with the UK’s sudden break with the EMS in September 1992. Since then, the convergence trend has been toward the US, with the scale of this increase in the strength of the US-
UK relationship against all the other European countries rising from less than 0.05 in the last quarter of 1992 to around 0.57 at January 2001. An exception to this is the Irish market where no trend is discernable until late 1997, after which the UK and Irish market price dynamics appear to be diverging. Interestingly, it was during this period that the Irish government announced its intentions to join the European common currency.

The implications of these findings are interesting not only because they reflect the sensitivity of movements in the UK market to those perturbations in other markets but because they also suggest such sensitivities may be driven by institutional detail, particularly the exchange rate arrangements of the UK. Clearly the existence of a close (and growing) link between the UK and US stock market returns suggests, not only that the UK’s business cycle is more closely aligned with that of the US, but that the transition to full membership of the Euro by the UK will be less smooth in terms of stock market shocks than if the UK market were already closely linked to the European markets. Given recent evidence on the importance of stock market variability on the real economy (Black, Groenewold and Fraser (2001)), this is important also from the wider policy perspective.

4.2. Long-Run Convergence

4.2(i) Phillips-Perron and Variance Ratio Tests

INSERT TABLE 3

Table 3 provides the results of the Phillips-Perron tests for the (logged) stock market series. Almost in every case, these tests statistics indicate the presence of unit roots in the levels of the series that cancels out on first differencing, and hence are supportive.
of the existence of a random walk element in equity prices. The exception to this is Germany where, with a P-P statistic of –4.001, we can marginally reject the null at conventional significance levels. As discussed earlier however, these tests have low power in detecting all departures from a random walk, and to accommodate this flaw we also consider below the heteroscedastic consistent variance ratio statistics due to Lo and MacKinlay (1988) along with the joint-test adaptation due to Chow and Denning (1993).

**INSERT TABLE 4**

Table 4 provides variance ratio statistics of less than unity at lags of 9 to 24 for the US and at lags 12 to 24 for all markets except Spain and Sweden, with Ireland, Italy and the Netherlands only having one value less than unity at lags 9,12 and 1 respectively. This would suggest that price behaviour at these far horizons is similar across these markets. While a variance ratio of less than unity is indicative of mean reversion, the joint test statistic, $Z^*$ (equation 14), with a critical value of 3.190, implies that the difference between these ratios and unity is insignificant for all markets including Germany.

Statistically significant persistence is however detected at longer lags for Italy, Spain and Sweden, and from lag 3 for Ireland and lag 12 for the Netherlands. Such features are not inconsistent with the view that persistence is associated with thin trading, where agents are slow to react to current information and prices tend to undershoot their equilibrium values, particularly at far horizons.
Overall, the evidence provided in Tables 3 and 4 suggests that, with the exception of Ireland and the Netherlands, our results are supportive of a unit root and uncorrelated disturbances. Due to the extent and significance of the persistence detected in the behaviour of the stock markets of Ireland and the Netherlands, we remove them from the empirical work reported below. However, as the remaining series show some evidence of a stochastic trend and we now ask if such trends are independent of each other.

4.2(ii). Johansen Multivariate Cointegration Tests

The first task here is to determine the model specification for the cointegration tests. We allow for deterministic trends in the data. The chosen lag length is 20, being motivated by the residual analysis and also by the variance ratio tests discussed above, where there was some evidence of similar behaviour in price innovations across markets which we would wish to capture.\textsuperscript{xiv} Tables 5A-5B below reported the results of the multivariate cointegration tests.

**INSERT TABLES 5A AND 5B**

The results clearly suggest that there are eight independent linear combinations of $X_t$ that are stationary ($r=8$) and that the VAR residuals do not exhibit serial correlation. Hence we have a single common stochastic trend, which, in turn, suggests that there are no long-term gains to international diversification. However, the practical relevance of this long-term co-movement depends on the extent of the deviations and how long it takes prices to return to this common trend or permanent component: how short is the short-run? We can provide some insight into this issue by considering the moving average representation of the vector error correction mechanism captured by
equations (17)-(19) above. Johansen (1995) has demonstrated that this has the following form:

\[
X_t = C \sum_{i=1}^{k} \varepsilon_i + C_{\eta t} + C(L)(\varepsilon_i + \eta)
\]  

(24)

where

\[
C = \beta_1 (\alpha'_1(I - \sum_{i} \Gamma_i)\beta_1')^{-1} \alpha'_1
\]

(25)

where \(\alpha_1\) and \(\beta_1\) denote the orthogonal complements to \(\alpha\) and \(\beta\), respectively. \(\alpha_1\) determines the vectors defining the space of the common stochastic trends and \(\beta_1\) gives the loadings associated with \(\alpha_1\). \(\Gamma_1\) is an NxN coefficient matrix from the error correction model. It is this decomposition that allows the elements of \(X_t\) to be explained in terms I(1) factors – the common stochastic trends, \(f_t\), plus some I(0) factors which are the transitory elements, \(X_t\):

\[
X_t = A_1 f_t + \bar{X}_t
\]

(26)

The identification of the common factors can be achieved by assuming that \(f_t\) are linear combinations of the variables, \(X_t\):

\[
f_t = B_1 X_t
\]

(27)

and if \(A_1 f_t\) and \(\bar{X}_t\) form a permanent – transitory decomposition of \(X_t\), then the only combination of \(X_t\) such that \(\bar{X}_t\) has no long run impact on \(X_t\) is:

\[
f_t = \alpha_1'X_t
\]

(28)

where \(\alpha_1'\alpha=0\) (see for example, Granger and Gonzalo (1995)). From this we are able to obtain the following permanent-transitory decomposition of \(X_t\):

\[
X_t = A_1 \alpha_1' X_t + A_2 \beta X_t
\]

(28)

where
\[ A_1 = \beta_1 (\alpha_1' \beta_1)^{-1} \]  
\[ A_2 = \alpha (\beta^\prime \alpha)^{-1} \]  

In our case of a single stochastic trend with \( \alpha_1' \) being a 9x1 vector, we can define the common trend as the weighted average of the nine markets, \( \beta_1 (\alpha_1' \beta_1)^{-1} X_t \), and if we then normalise \( \alpha_1' \) so that each element of \( \alpha_1 \) sums to unity, we can gauge the relative importance of each market to the trend, \( \beta_1 (\alpha_1' \beta_1)^{-1} \), and the relative importance of the trend to each market, \( \alpha_1' \). Further, we take \( X_t - \beta_1 (\alpha_1' \beta_1)^{-1} \alpha_1 X_t \), we can assess the extent to which transitory fluctuations are important. Such deviations will be the result of transitory fluctuations in fundamentals (given the assumption of market efficiency). If returns are mainly driven by innovations to their long-run trends than then the transitory component should be small. The results of this exercise are shown in Table 6, and in Figures 2A through 2I.

**INSERT TABLE 6**

Interestingly, and consistent with the work of Kasa (1992), Table 6 shows that the US and UK markets have relatively smaller weights in the determination of the common trend (8.3% for the US and 9.4% for the UK), despite the fact that these markets have the highest market capitalisation of the sample. In general, the ‘European’ grouping have similar weightings, ranging from 11.1% for Germany to 13.1% for Spain and this is reflected in the \( \alpha_1' \)'s, where four of the seven mainland European markets respond more than proportionately to innovations in the common trend.

**INSERT FIGURES 2A-2I**
Figures 2A through 2I display the (demeaned) transitory component for the sample of equity markets.\textsuperscript{xvi} For the core European markets, and particularly from the late 1970s, there has been only one significant departure from the common trend – Germany in 1994 to 1997. For the US and UK however, the pattern is rather different. The US market appears to be pushing at and frequently breaking the upper significance boundary from mid 1990, with the UK doing likewise since its departure from the ERM in 1992. Correlations of the transitory components (not reported) show that with the exception of the Italian and Swedish markets, movements in the US and UK components had a negative association with the core European markets indicating that over the period actual US and UK prices tended to be above their permanent component or stochastic trend when the core European markets were below their trend.\textsuperscript{xvii}

\textbf{CONCLUSION}

The aim of this paper was twofold: first to consider the short-run bi-lateral linkages between the UK, US and European stock markets and second, to analyse the long-term relationships between those markets.

Our results can be simply stated. First, the correlations between stock market returns have not been constant over time. Second, since autumn 1992, short-run movements in the UK stock market have been increasingly associated with movements in the US market rather than with movements in European markets. Third, the sensitivity of the UK market to perturbations in other markets appears to be driven by institutional arrangements, particularly exchange rate movements and alignments. The main implications are that the time may not be right for an acceleration of merger activity between the UK stock market and its European neighbours and, any transition to the
Euro by the UK, not underpinned by economic fundamentals, may lead to a relatively bumpy stock market ride as the market adjusts to the imposed business cycle conditions.

We also report evidence to suggest that over the long term stock markets are perfectly correlated and although deviations from this long-term trend do occur, with the possible exception of the US and UK market, they are rarely significant. In general, the common stochastic trend in association with factor loadings can track quite precisely these stock markets. For the UK and UK markets, we found that the 1990s has tended to display a few small but significant deviations from the permanent component – a feature that is not surprising given that the long-term trend has relatively less influence on these two markets. Overall, however, gains from international diversification would tend to be short-lived. Future work may consider the extent to which these results hold for other types of portfolios, such as those sorted by value, growth and size.

ENDNOTES

i See for example, Cho et al. (1986); Jorion and Schwartz (1986); Jeon and Fursenberg (1990); Koch and Koch (1991); Stehle (1997); Alexakis et al. (1997); Janakiraman and Lamba (1998); Phylaktis (1999); Bracker et al. (1999). The extant literature on stock market convergence is, in itself, indicative of the lack of consensus amongst researchers of, if, how and why, stock market convergence is occurring. Overall, the reported results are not time invariant and would also appear to be sensitive to the empirical methods employed.

ii Hameo et al. (1990) for example, report conditional mean and variance spillovers from the US and UK to Japan, but not in the opposite direction, while Ammer and Mei (1996), find the correlations of US and UK future expected returns to be high, thus indicating substantial economic and financial integration.

iii From 1996 through 2000, the average yearly growth in stock market value was: Sweden, 23.38%; Ireland, 21.43%; France, 19.20%; Netherlands, 17.84%; Spain, 16.78%; Denmark, 16.44%; US, 16.26%; UK, 14.64%; Germany, 14.40%; Belgium, 13.20%.

iv See ‘Snuggling up to the Euro’ The Sunday Times, February 11th, 2001, p. 18, where the author (David Smith) reports on the degree of UK’s economic convergence with Europe relative to the US.

v At time of writing the exchanges of Amsterdam, Paris and Brussels have agreed to a merger - Euronext. Effectively, this will create a pan-European exchange, which in turn, aims to be the European link in a global exchange encompassing the continents of Asia, Europe and the Americas. For the history of Euronext see: www.euronext.com/en/euronextinfo/history. In 2000, we witnessed a failed consolidation between the London Stock Exchange and the Frankfurt Stock Exchange, and the mount of a take-over bid by the Swedish Stock Market – the OM group.
In the analysis below, we use the description ‘European’ to denote the non-UK European markets.

While in the empirical work below we report results for convergence of continuously compounded returns, we also conducted the empirical analyses on excess returns – returns less the safe rate. As the results were both qualitatively and quantitatively similar, this suggests that movements in equity premiums were, over this period, driven by stock market returns rather than by safe rates of return.

A finding of cointegration could imply either that markets are inefficient processors of information (there are abnormal profits to be made), or that equilibrium returns are time-varying, or both. Alternatively a finding of non-cointegration is a particularly strong result since it indicates the absence of time-varying equilibrium returns and that markets are efficient. There is by now a large body of evidence to support the existence of time-varying equilibrium returns.

Dynamic steady-state equilibrium involves the addition of a term in the constant vector of steady-state growth rates to equation (17), which we omit here for expositional purposes. This does not affect the subsequent discussion.

The Trace statistic will have greater power than the Lmax when the eigenvalues are evenly distributed, while the Lmax statistic will be relatively more robust when the eigenvalues are either large or small.

The Euro currently includes a total of twelve economies – including Belgium, Ireland, Netherlands and Spain. Denmark rejected entry to the common currency by referendum on 28th September 2000. Both Sweden and the UK have yet to decide if and when entry to the Euro will take place.

As the interest in the paper is in the deconstruction of market movements we do not include the income component of market holding-period returns. In practice however, dividends are constant relative to price (Lo and MacKinlay (1990)) and the inclusion of dividends yields is unlikely to affect results.

These statistics are not reported.

Consistent with Kasa (1992) who used seventeen years of data, the cointegration results were sensitive to the chosen lag length. This would support the view that the power of unit root and cointegration tests are mainly a function of the time horizon and not the number of observations (Shiller and Perron (1985)).

As Kasa (1992) points out such a method of extracting the common trend leads to a trend which is not a pure random walk as it will contain any short-run dynamics that are orthogonal to the long-run dynamics.

As zero means were not imposed on the cointegrating relations, the relevant spreads between the actual series and its loaded stochastic trend were demeaned.

The UK had a positive correlation with the Spanish market transitory component while the US had a negative correlation.
REFERENCES


## TABLE 1
### SUMMARY STATISTICS ON MONTHLY MARKET RETURNS

<table>
<thead>
<tr>
<th>Country</th>
<th>Mean</th>
<th>S. D.</th>
<th>J-B</th>
</tr>
</thead>
<tbody>
<tr>
<td>Belgium</td>
<td>0.008</td>
<td>0.067</td>
<td>440.414</td>
</tr>
<tr>
<td>Denmark</td>
<td>0.008</td>
<td>0.067</td>
<td>503.484</td>
</tr>
<tr>
<td>France</td>
<td>0.009</td>
<td>0.066</td>
<td>427.295</td>
</tr>
<tr>
<td>Germany</td>
<td>0.007</td>
<td>0.067</td>
<td>442.930</td>
</tr>
<tr>
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<td>0.011</td>
<td>0.067</td>
<td>317.942</td>
</tr>
<tr>
<td>Italy</td>
<td>0.012</td>
<td>0.069</td>
<td>443.830</td>
</tr>
<tr>
<td>Netherlands</td>
<td>0.009</td>
<td>0.047</td>
<td>326.608</td>
</tr>
<tr>
<td>UK</td>
<td>0.009</td>
<td>0.060</td>
<td>1184.353</td>
</tr>
<tr>
<td>US</td>
<td>0.008</td>
<td>0.066</td>
<td>472.4554</td>
</tr>
<tr>
<td>Spain</td>
<td>0.011</td>
<td>0.067</td>
<td>660.032</td>
</tr>
<tr>
<td>Sweden</td>
<td>0.010</td>
<td>0.066</td>
<td>511.228</td>
</tr>
</tbody>
</table>

S.D. denotes the standard deviation of the return series and J-B is the Jarque-Bera test for normality. Figures in parenthesis under the J-B are marginal significance levels. The data covers the period January 1974 to January 2001.
<table>
<thead>
<tr>
<th>TABLE 2a</th>
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<tbody>
<tr>
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</tr>
<tr>
<td></td>
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<tr>
<td>Belgium</td>
</tr>
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<tr>
<td>France</td>
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<tr>
<td>Germany</td>
</tr>
<tr>
<td>Ireland</td>
</tr>
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<td>Italy</td>
</tr>
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<td>Netherlands</td>
</tr>
<tr>
<td>UK</td>
</tr>
<tr>
<td>US</td>
</tr>
<tr>
<td>Spain</td>
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<td>Sweden</td>
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<table>
<thead>
<tr>
<th>TABLE 2b</th>
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</tr>
<tr>
<td></td>
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</tr>
<tr>
<td>Denmark</td>
</tr>
<tr>
<td>France</td>
</tr>
<tr>
<td>Germany</td>
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<td>Ireland</td>
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<td>Italy</td>
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<td>Netherlands</td>
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</tr>
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<td>US</td>
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<td>Spain</td>
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<table>
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<th>TABLE 2c</th>
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<td>CORRELATION MATRIX OF RETURNS: January 1984 to December 1993</td>
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<td></td>
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<tr>
<td>Belgium</td>
</tr>
<tr>
<td>Denmark</td>
</tr>
<tr>
<td>France</td>
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<tr>
<td>Germany</td>
</tr>
<tr>
<td>Ireland</td>
</tr>
<tr>
<td>Italy</td>
</tr>
<tr>
<td>Netherlands</td>
</tr>
<tr>
<td>UK</td>
</tr>
<tr>
<td>US</td>
</tr>
<tr>
<td>Spain</td>
</tr>
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<td>Sweden</td>
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</table>

<table>
<thead>
<tr>
<th>TABLE 2d</th>
</tr>
</thead>
<tbody>
<tr>
<td>CORRELATION MATRIX OF RETURNS: January 1994 to January 2001</td>
</tr>
<tr>
<td></td>
</tr>
<tr>
<td>Belgium</td>
</tr>
<tr>
<td>Denmark</td>
</tr>
<tr>
<td>France</td>
</tr>
<tr>
<td>Germany</td>
</tr>
<tr>
<td>Ireland</td>
</tr>
<tr>
<td>Italy</td>
</tr>
<tr>
<td>Netherlands</td>
</tr>
<tr>
<td>UK</td>
</tr>
<tr>
<td>US</td>
</tr>
<tr>
<td>Spain</td>
</tr>
<tr>
<td>Sweden</td>
</tr>
</tbody>
</table>
* If the UK is converging with the US, $\beta_2$ will converge to unity. If the UK is converging with Europe, $\beta_2$ will converge to zero.
### TABLE 3

**TESTS FOR STOCHASTIC TRENDS**

**UNIVARIATE UNIT ROOT TESTS\(^a\)**

<table>
<thead>
<tr>
<th></th>
<th>P-P Levels</th>
<th>P-P First Differences</th>
</tr>
</thead>
<tbody>
<tr>
<td>Belgium</td>
<td>-3.284</td>
<td>-15.687</td>
</tr>
<tr>
<td>Denmark</td>
<td>-3.072</td>
<td>-15.505</td>
</tr>
<tr>
<td>France</td>
<td>-2.924</td>
<td>-15.829</td>
</tr>
<tr>
<td>Germany</td>
<td>-4.001</td>
<td>-15.613</td>
</tr>
<tr>
<td>Ireland</td>
<td>-3.041</td>
<td>-15.452</td>
</tr>
<tr>
<td>Italy</td>
<td>-3.006</td>
<td>-16.604</td>
</tr>
<tr>
<td>Netherlands</td>
<td>-2.635</td>
<td>-17.030</td>
</tr>
<tr>
<td>UK</td>
<td>-3.414</td>
<td>-16.264</td>
</tr>
<tr>
<td>US</td>
<td>-3.933</td>
<td>-16.679</td>
</tr>
<tr>
<td>Spain</td>
<td>-2.597</td>
<td>-15.719</td>
</tr>
<tr>
<td>Sweden</td>
<td>-3.625</td>
<td>-15.540</td>
</tr>
</tbody>
</table>

\(^a\)P-P is the Phillips-Perron test for a unit root in the series -- it incorporates an intercept and a trend. The critical values for the P-P test are –3.990 (1%), -3.425 (5%), and –3.135 (10%). The data covers the period January 1974 to January 2001.
### TABLE 4
TESTS FOR STOCHASTIC TRENDS: VARIANCE RATIO TESTS

<table>
<thead>
<tr>
<th>Country</th>
<th>q=2</th>
<th>q=3</th>
<th>q=6</th>
<th>q=9</th>
<th>q=12</th>
<th>q=18</th>
<th>q=24</th>
<th>q=30</th>
<th>q=36</th>
</tr>
</thead>
<tbody>
<tr>
<td>Belgium</td>
<td>1.14</td>
<td>1.17</td>
<td>1.16</td>
<td>1.02</td>
<td>0.91</td>
<td>0.81</td>
<td>0.81</td>
<td>1.01</td>
<td>1.15</td>
</tr>
<tr>
<td>Denmark</td>
<td>1.16</td>
<td>1.18</td>
<td>1.17</td>
<td>1.05</td>
<td>0.95</td>
<td>0.86</td>
<td>0.87</td>
<td>1.08</td>
<td>1.24</td>
</tr>
<tr>
<td>France</td>
<td>1.14</td>
<td>1.14</td>
<td>1.12</td>
<td>1.00</td>
<td>0.91</td>
<td>0.89</td>
<td>0.94</td>
<td>1.17</td>
<td>1.35</td>
</tr>
<tr>
<td>Germany</td>
<td>1.15</td>
<td>1.15</td>
<td>1.10</td>
<td>0.96</td>
<td>0.85</td>
<td>0.72</td>
<td>0.68</td>
<td>0.84</td>
<td>0.94</td>
</tr>
<tr>
<td>Ireland</td>
<td>1.17</td>
<td>1.26</td>
<td>1.40</td>
<td>1.43</td>
<td>1.47</td>
<td>1.53</td>
<td>1.59</td>
<td>1.78</td>
<td>2.00</td>
</tr>
<tr>
<td>Italy</td>
<td>1.10</td>
<td>1.10</td>
<td>1.13</td>
<td>1.00</td>
<td>0.94</td>
<td>1.00</td>
<td>1.10</td>
<td>1.35</td>
<td>1.58</td>
</tr>
<tr>
<td>Netherlands</td>
<td>1.08</td>
<td>1.17</td>
<td>1.26</td>
<td>1.26</td>
<td>1.34</td>
<td>1.62</td>
<td>1.79</td>
<td>2.06</td>
<td>2.38</td>
</tr>
<tr>
<td>UK</td>
<td>1.12</td>
<td>1.12</td>
<td>1.14</td>
<td>1.02</td>
<td>0.93</td>
<td>0.92</td>
<td>0.98</td>
<td>1.20</td>
<td>1.40</td>
</tr>
<tr>
<td>US</td>
<td>1.08</td>
<td>1.08</td>
<td>1.06</td>
<td>0.93</td>
<td>0.86</td>
<td>0.82</td>
<td>0.88</td>
<td>1.07</td>
<td>1.18</td>
</tr>
<tr>
<td>Spain</td>
<td>1.16</td>
<td>1.21</td>
<td>1.28</td>
<td>1.22</td>
<td>1.16</td>
<td>1.20</td>
<td>1.27</td>
<td>1.52</td>
<td>1.80</td>
</tr>
<tr>
<td>Sweden</td>
<td>1.16</td>
<td>1.19</td>
<td>1.20</td>
<td>1.13</td>
<td>1.07</td>
<td>1.01</td>
<td>1.07</td>
<td>1.32</td>
<td>1.52</td>
</tr>
</tbody>
</table>

*a q denotes lag length in months. \( \overline{M}_q(q) \) is the variance ratio estimate of equation (7). \( Z^* \) is the Chow and Denning (1993) joint-test size heteroscedastic-consistent statistic. The critical value for the joint tests of \( Z^* \) is 3.742 (1%), 3.190 (5%) and 3.103 (10%), calculated at nine lag intervals on the SMM tables (Stoline & Ury (1979)). * and ** denote significant values at 1% and 5% respectively.
### TABLE 5A
**MULTIVARIATE COINTEGRATION TESTS**
**JOHANSEN TRACE TESTS**
US, UK, GERMANY, FRANCE, ITALY, BELGIUM, SPAIN, DENMARK, SWEDEN

<table>
<thead>
<tr>
<th>Possible Number of Stationary Linear Combinations</th>
<th>Eigenvalues</th>
<th>LR or Trace Statistic</th>
<th>L-max Statistic</th>
<th>Critical Value LR or Trace</th>
<th>Critical Value L-max</th>
</tr>
</thead>
<tbody>
<tr>
<td>None</td>
<td>0.2978</td>
<td>457.30</td>
<td>107.85</td>
<td>185.83</td>
<td>35.84</td>
</tr>
<tr>
<td>At most 1*</td>
<td>0.2531</td>
<td>349.45</td>
<td>89.00</td>
<td>149.99</td>
<td>32.26</td>
</tr>
<tr>
<td>At most 2*</td>
<td>0.2251</td>
<td>260.45</td>
<td>77.77</td>
<td>117.73</td>
<td>28.36</td>
</tr>
<tr>
<td>At most 3*</td>
<td>0.1854</td>
<td>182.68</td>
<td>62.54</td>
<td>89.37</td>
<td>24.63</td>
</tr>
<tr>
<td>At most 4*</td>
<td>0.1276</td>
<td>120.14</td>
<td>41.63</td>
<td>64.74</td>
<td>20.90</td>
</tr>
<tr>
<td>At most 5*</td>
<td>0.1140</td>
<td>78.51</td>
<td>36.91</td>
<td>43.84</td>
<td>17.14</td>
</tr>
<tr>
<td>At most 6*</td>
<td>0.0883</td>
<td>41.60</td>
<td>28.21</td>
<td>26.70</td>
<td>13.39</td>
</tr>
<tr>
<td>At most 7*</td>
<td>0.0429</td>
<td>13.39</td>
<td>13.38</td>
<td>13.31</td>
<td>10.60</td>
</tr>
<tr>
<td>At most 8</td>
<td>0.0000</td>
<td>0.01</td>
<td>0.01</td>
<td>2.71</td>
<td>2.71</td>
</tr>
</tbody>
</table>

*aThe null hypothesis is no-cointegration. * denotes rejection of the null hypothesis of no cointegration. The number of lags in the VAR was 20. Critical values are from Johansen and Nielsen (1993). The tests indicate 8 cointegrating equations.

### TABLE 5B
**RESIDUAL AUTOCORRELATION TESTS**

<table>
<thead>
<tr>
<th></th>
<th>L-B(12)</th>
</tr>
</thead>
<tbody>
<tr>
<td>US</td>
<td>1.003(0.999)</td>
</tr>
<tr>
<td>UK</td>
<td>1.421(0.999)</td>
</tr>
<tr>
<td>Germany</td>
<td>1.822(0.999)</td>
</tr>
<tr>
<td>France</td>
<td>1.793(0.999)</td>
</tr>
<tr>
<td>Belgium</td>
<td>1.956(0.999)</td>
</tr>
<tr>
<td>Italy</td>
<td>3.485(0.991)</td>
</tr>
<tr>
<td>Spain</td>
<td>0.765(0.999)</td>
</tr>
<tr>
<td>Denmark</td>
<td>2.030(0.999)</td>
</tr>
<tr>
<td>Sweden</td>
<td>1.480(0.999)</td>
</tr>
</tbody>
</table>

*a L-B(12) denotes the Ljung-Box test for serial correlation. Figures in parenthesis are marginal significance levels.*
<table>
<thead>
<tr>
<th></th>
<th>$\alpha_1$</th>
<th>$\beta_1 (\alpha_1 \beta_1)^{-1}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>US</td>
<td>0.735</td>
<td>0.083</td>
</tr>
<tr>
<td>UK</td>
<td>0.835</td>
<td>0.094</td>
</tr>
<tr>
<td>Germany</td>
<td>0.980</td>
<td>0.111</td>
</tr>
<tr>
<td>France</td>
<td>1.005</td>
<td>0.113</td>
</tr>
<tr>
<td>Belgium</td>
<td>1.000</td>
<td>0.113</td>
</tr>
<tr>
<td>Italy</td>
<td>1.099</td>
<td>0.124</td>
</tr>
<tr>
<td>Spain</td>
<td>1.158</td>
<td>0.131</td>
</tr>
<tr>
<td>Denmark</td>
<td>0.994</td>
<td>0.112</td>
</tr>
<tr>
<td>Sweden</td>
<td>1.055</td>
<td>0.119</td>
</tr>
</tbody>
</table>

*Elements of $\alpha_1$ are coefficients to the common trend which indicates the relative importance of the trend to the market. The vector $\beta_1 (\alpha_1 \beta_1)^{-1}$ are weights, with $\alpha_1$ normalised so that the elements of $\beta_1 (\alpha_1 \beta_1)^{-1}$ sum to unity and indicates the relative importance of each market to the trend.*
FIGURES 2A-2I: THE TRANSITORY COMPONENT