*Applied Financial Economics*, 1999, **9**, 73–85

# *Macroeconomic determinants of long-term stock market comovements among major EMS countries*

YIN-WONG CHEUNG and KON S. LAI[‡]

*Department of Economics, University of California, Santa Cruz, CA 95064, USA* and
[‡] *Department of Economics, California State University, Los Angeles, CA 90032, USA*

Long-term comovements of national stock markets in three EMS (European Monetary System) countries – France, Germany and Italy – are examined. The EMS stock markets are found to display long-term comovements governed by two common permanent components. To identify some interpretable sources of such long-term market comovements, the study explores whether they can be linked to similar comovements in macroeconomic variables, including the money supply, dividends and industrial production. Like stock prices, two common permanent components are found driving the comovements in each of these variables. Further analysis suggests that the long-term comovements in stock prices can be partly attributable to those in the macroeconomic variables, especially for the post-1987 period. The results confirm at least a limited role of these macroeconomic variables in accounting for the stock market comovements among the EMS countries.

## I. INTRODUCTION

Many studies have been devoted to investigating the empirical relationship between national stock markets, for which findings of comovements are sometimes interpreted as a loose form of international capital market integration. Early studies by Agmon (1972), Hilliard (1979), Levy and Sarnat (1970), Panton, *et al.* (1976), and Ripley (1973) mostly found little covariation in national stock price indexes based on data from the 1960s and 1970s. Motivated by the trend towards internationalization of equity markets, recent studies (Eun and Shim, 1989; Hamao *et al.*, 1990; Koch and Koch, 1991; Schollhammer and Sand, 1987; Von Furstenberg and Jeon, 1989) examine dynamic linkages among stock markets in the 1980s. These studies consistently find significant short-term comovements of national stock indexes, and the findings may be attributed to the growing importance of international investors and the substantial improvements in telecommunications and computer technology in the 1980s.

This study explores the potential existence and sources of long-term comovements of stock markets among three major member countries of the European Monetary System (EMS), namely, France, Germany and Italy. The aforementioned studies focus exclusively on short-term market comovements and examine how stock markets in the same or different time zones are linked and transmit information in the short run. Long-term comovements between national stock markets can also exist, however (Lai *et al.*, 1993). Ripley (1973) observes that relative movements in stock indexes can reflect general changes in national income across countries. It follows that the presence of strong economic ties and policy coordination among the relevant countries can indirectly link their stock prices over time. Through international capital flows, moreover, real interest rates of some countries tend to move together over time. To the extent that stock prices are affected by interest rate movements, such interest rate linkage may contribute to long-term market comovements.

The EMS provides an interesting economic arrangement for studying comovements of financial markets, not just for their geographical proximity. Established in mid-March 1979 with exchange rate stabilization as a major goal, the EMS yields a hybrid system of pegged and adjustable exchange rates. To achieve exchange rate stability, a certain degree of monetary policy coordination among EMS countries is required. A long-term objective of the EMS arrangement is to foster trade and facilitate a process of economic integration, involving commodity-, capital-, and factor-market integration among member countries. Fratianni and von Hagen (1990) find substantial short-term correlation between the growth rate of the monetary base in Germany and that in other EMS countries. Based on data for the period 1979–1988, MacDonald and Taylor (1991) further suggest the presence of long-term comovements in both exchange rates and the money supply of France, Germany and Italy. The results, obtained from cointegration analysis, are interpreted as evidence of monetary policy convergence among the EMS members.

Kasa (1992) estimates the number of common stochastic trends in the stock markets of Germany and four non-EMS countries (Canada, Japan, the UK and the US) over the period 1974–1990. Evidence is presented in support of the presence of long-term comovements, with a common trend driving the five national stock markets. A single common trend is also found in GNP series, but two common trends in dividend series. No formal analysis is pursued to study whether the common trend in stock prices can be linked to those in dividends and GNP. On the other hand, research focusing on European stock markets appears sparse, albeit some studies (Corhay and Rad, 1994; Corhay *et al.*, 1993; Fraser *et al.*, 1994; Gallagher, 1995; Mathur and Subrahmanyam, 1990; Taylor and Tonks, 1989) have examined common trends among European stock markets and produced mixed results.

This study analyses long-term stock market comovements among EMS countries. The analysis not only provides evidence of market comovements but also explores possible sources of the comovements, which may be the result of some systematic forces influencing the markets. Many models of asset pricing suggest that stock returns can be related to the states of macroeconomy. Financial securities represent claims against the current and future output of the economy, and their price movements may reflect changes in economic activity in the long run. Chen (1991), Chen *et al.* (1986) and Fama (1981, 1990), among others, have investigated the potential links between US stock returns and the states of US macroeconomy. The present study investigates not movements of a single stock market in isolation but *relative* movements of different national stock markets. For countries with their economies increasingly linked together, economic and financial developments in one country can affect the markets in others. The present analysis examines if relative long-term movements of different EMS stock markets can be linked to similar relative long-term movements in their macroeconomic variables.

The empirical study employs Gonzalo and Granger's (1995) analysis of common permanent components, which permits a decomposition of the dynamics of a multivariate system into permanent and transitory components and allows for long-term comovement analysis between subsystems under a cointegration framework. For a system or subsystem of variables, its long-run behaviour is governed by a relatively small set of common permanent components. These permanent components represent the underlying forces driving long-term comovements among the variables. By obtaining common permanent components in different subsystems of variables, researchers can analyse whether there are long-term comovements within individual variable groups and whether the comovements of a specific group of variables can in turn be linked to those of others. In this study, we investigate the potential linkage between the common permanent components in four groups of variables, those of stock prices, the money supply, dividends, and industrial production of three EMS countries.

Several studies, including King *et al.*, (1994), King and Wadhwani (1990), and Hamao *et al.*, (1990), have examined transmission of volatility between national stock markets. These studies analyse short-term comovements in conditional variances of stock market returns across countries. King *et al.* (1994), specifically, report that only a small proportion of the short-term market covariations can be explained by observable economic variables. In the present study, long-term comovements between levels of stock market indexes are examined, instead. The analysis abstracts from short-term market dynamics and focuses on permanent components of stock prices. Permanent and temporary components in stock prices may show different properties, with the former having much higher predictability than the latter (Fama and French, 1988; Poterba and Summers, 1988).

The paper is organized as follows. Section II briefly discusses the empirical methodology. Section III describes the data. Section IV contains empirical results. Section V provides further results of subperiod analysis. Section VI concludes.

## II. STATISTICAL METHODOLOGY

Gonzalo and Granger's (1995) analysis provides a decomposition of the dynamics of multivariate time series into permanent and transitory components, different from those obtained from others, e.g. Kasa (1992) and Stock and Watson (1988). The permanent component obtained from Gonzalo and Granger's method can be explicitly expressed as a function of observable variables (see Escribano and Peña (1994) for a discussion on the relationship between alternative common trend representations). Let $X_t = (x_{1t}, \dots, x_{nt})'$ be a vector of $n$ time series modelled by

$$C(L)X_t = \mu + \varepsilon_t, \qquad (1)$$

where $L$ is the lag operator, $C(L)$ is a matrix polynomial in $L$, $\mu$ is a constant vector, and $\varepsilon_t$ is a vector of white noises. When the series in $X_t$ are I(1), i.e. integrated of order one, their linear combinations are generally also I(1). Under the hypothesis of $r(0 \leqslant r < n)$ cointegrating vectors, however, $C(1) = \gamma\alpha'$ with $\gamma$ and $\alpha$ being $n \times r$ matrices of rank $r$ such that $\alpha' X_t$ is I(0) and represents the long-run equilibrium relationships underlying $X_t$. Consider $\gamma_\perp$ a $n \times (n-r)$ matrix orthogonal to $\gamma$ so that $\gamma_\perp \gamma = 0$. Gonzalo and Granger (1995) show that the dynamics of $X_t$ can be decomposed as follows:

$$X_t = A_1 f_t + A_2 z_t, \tag{2}$$

where $f_t = \gamma'_\perp X_t$ is I(1); $z_t = \alpha' X_t$ is I(0); and

$$A_1 = \alpha_\perp(\gamma'_\perp \alpha_\perp)^{-1} \tag{3}$$

and

$$A_2 = \gamma(\alpha' \gamma)^{-1} \tag{4}$$

are normalization matrices. $X_t$ is thus expressed in terms of a simple sum of the common permanent component ($f_t$) and the transitory component ($z_t$).

Estimates of $\alpha$ and $\gamma_\perp$ are obtained from reduced rank regression analysis, as applied by Johansen (1991). Let $\Delta = 1 - L$ and $k$ be the lag order of the estimated system in $X_t$. Regress $\Delta X_t$ on a constant and $\Delta X_{t-1}, \ldots, \Delta X_{t-k+1}$, giving the residual $\hat{u}_{1t}$. Regress $X_{t-k}$ on a constant and $\Delta X_{t-1}, \ldots, \Delta X_{t-k+1}$, giving the residual $\hat{u}_{2t}$. Form the product moment matrices of the residuals as $S_{ij} = \Sigma_{t=1}^{T} \hat{u}_{it}\hat{u}'_{jt}/T$ for $i,j = 1,2$. Solving the equation $|\lambda S_{22} - S_{21}(S_{11})^{-1}S_{12}| = 0$ yields eigenvalues $\lambda_1 > \ldots > \lambda_n$ and their associated eigenvectors $V = (v_1, \ldots, v_n)$ normalized such that $V'S_{22}V = I$. The estimate of $\alpha$ is given by $(v_1, \ldots, v_r)$. Solving also the equation $|\lambda S_{11} - S_{12}(S_{22})^{-1}S_{21}| = 0$ gives eigenvalues $\lambda_1 > \ldots > \lambda_n$ and their corresponding eigenvectors $M = (m_1, \ldots, m_n)$ with the normalization $M'S_{11}M = I$. The estimate of $\gamma_\perp$ is then given by $(m_{r+1}, \ldots m_n)$.

The rank of cointegration $r$ can be determined using Johansen's (1991) analysis. The test statistic for the hypothesis of at most $r$ cointegrating vectors is

$$-2 \ln Q = -T\, \Sigma_{j=r+1}^{n} \ln(1 - \lambda_j) \tag{5}$$

This statistic is termed the trace or $\lambda_{\text{trace}}$ statistic. An alternative statistic, given by

$$-2 \ln Q_{r|r+1} = -T \ln(1 - \lambda_{r+1}) \tag{6}$$

and referred to as the maximal eigenvalue or $\lambda_{\max}$ statistic, examines the null hypothesis of $r$ cointegrating vectors against the alternative of $r+1$ ones. Using response surface analysis, Cheung and Lai (1993) provide finite sample critical values for these tests by correcting for the effects of the sample size ($T$), the number of variables ($n$), and the lag order ($k$).

## III. THE DATA

All stock price data series consist of the monthly Morgan Stanley Capital International (MSCI) market indexes, covering the period from April 1979 to June 1992 (Italy pulled out of the EMS in September 1992). The MSCI indexes are fully comparable across countries since they are constructed on a consistent basis. These indexes are value-weighted, computed with dividends reinvested, and expressed in terms of the US dollar. In constructing the MSCI indexes, the market value of investment companies and of foreign domiciled companies are excluded to avoid double-counting. All stock indexes are converted into real dollars based on the US consumer price index. Data descriptions are contained in issues of the *Morgan Stanley Capital International Perspectives.*

The macroeconomic data under study include monthly money supply (M1) and industrial production series taken from the OECD's *Main Economic Indicators* data bank. These series are converted into real US dollars using relevant spot exchange rates and consumer price indexes. Month-end spot exchange rates are drawn from the IMF's *International Financial Statistics* data tape; consumer price indexes are obtained from the OECD's *Main Economic Indicators* data bank. The real dividend data are extracted from the difference between series of dividend-inclusive and dividend-exclusive MSCI indexes (adjusted for rights issues, stock dividends, and stock splits). These dividend series represent the gross dividends an investor would receive if the investor were investing in the portfolios defined by the MSCI stock price indexes.

As a preliminary analysis, each series is first checked for a unit root using the augmented Dickey–Fuller (ADF) test that allows for a linear time trend. The null hypothesis is that the relevant series has a unit root. The test is conducted on the levels and first differences of the logarithm series of the stock price index, the money supply, dividends, and industrial production for each of the three EMS countries. The test results are reported in Table 1. In no case can the unit root hypothesis be rejected at any standard significance level for level series; whereas, tests performed on first differences strongly indicate that each of the first-differenced series is stationary. The evidence supports that all the level series contain a single unit root. Figures 1(a) to 1(d) contain plots of the data series under examination. Although individual series across countries can wander extensively, they appear to move in close association over time.

## IV. EMPIRICAL RESULTS

Long-term comovements among stock markets are first analysed. The empirical analysis focuses mainly on three national stock markets in France (FR), Germany (GM) and Italy (IT), though some evidence on comovements for

Table 1. *Unit root test results on individual data series*

| Series | Lag $p$ | ADF($p$) statistic | Critical value 10% | Critical value 5% | Lag $p$ | ADF($p$) statistic | Critical value 10% | Critical value 5% |
|---|---|---|---|---|---|---|---|---|
| | Level series: | | | | Differenced series: | | | |
| $SP_{FR,t}$ | 1 | − 1.741 | − 3.135 | − 3.427 | 1 | − 8.963** | − 3.135 | − 3.427 |
| $SP_{GM,t}$ | 1 | − 2.120 | − 3.135 | − 3.427 | 1 | − 8.947** | − 3.135 | − 3.427 |
| $SP_{IT,t}$ | 2 | − 1.238 | − 3.129 | − 3.421 | 1 | − 8.171** | − 3.135 | − 3.427 |
| $MS_{FR,t}$ | 1 | − 1.622 | − 3.135 | − 3.427 | 1 | − 8.560** | − 3.135 | − 3.427 |
| $MS_{GM,t}$ | 1 | − 1.777 | − 3.135 | − 3.427 | 1 | − 8.358** | − 3.135 | − 3.427 |
| $MS_{IT,t}$ | 1 | − 1.532 | − 3.135 | − 3.427 | 1 | − 8.117** | − 3.135 | − 3.427 |
| $DI_{FR,t}$ | 1 | − 1.255 | − 3.135 | − 3.427 | 1 | − 10.378** | − 3.135 | − 3.427 |
| $DI_{GM,t}$ | 3 | − 2.336 | − 3.124 | − 3.414 | 2 | − 11.217** | − 3.129 | − 3.421 |
| $DI_{IT,t}$ | 2 | − 1.248 | − 3.129 | − 3.421 | 1 | − 8.868** | − 3.135 | − 3.427 |
| $IP_{FR,t}$ | 1 | − 1.435 | − 3.135 | − 3.427 | 1 | − 8.765** | − 3.135 | − 3.427 |
| $IP_{GM,t}$ | 1 | − 1.712 | − 3.135 | − 3.427 | 1 | − 8.786** | − 3.135 | − 3.427 |
| $IP_{IT,t}$ | 2 | − 1.394 | − 3.129 | − 3.421 | 1 | − 9.197** | − 3.135 | − 3.427 |

$SP_{FR,t}$, $SP_{GM,t}$ and $SP_{IT,t}$ are the respective real stock price series for France, Germany and Italy. $MS_{FR,t}$, $MS_{GM,t}$ and $MS_{IT,t}$ are the respective real money supply series for France, Germany and Italy. $DI_{FR,t}$, $DI_{GM,t}$ and $DI_{IT,t}$ are the respective real dividend yield series for France, Germany and Italy. $IP_{FR,t}$, $IP_{GM,t}$ and $IP_{IT,t}$ are the respective real industrial production series for France, Germany and Italy. The lag parameter $p$ is selected using the Schwarz information criterion, with a maximum lag of 8 allowed. Finite sample critical values are calculated based on response surface estimates from Cheung and Lai (1995), which correct for the effect of lag order. The last four columns correspond to results from tests on first-differenced series. Statistical significance is indicated by an asterisk (*) for the 10% level or a double asterisk (**) for the 5% level.

non-EMS stock markets in Canada, Japan, the UK and the US will also be reported.

### *Common permanent components in stock prices*

For the lag specification in cointegration analysis, the lag order $k$ was first estimated using a model selection procedure based on the Schwarz information criterion. The maximum lag length considered is $k = 8$. The corresponding residuals were tested for serial correlation using the Ljung–Box test. If the residuals could not pass the serial correlation test, the lag length would be increased until the serial correlation was removed statistically. In the case for stock prices, $k = 2$ is selected. Residual diagnostic statistics (not reported) were also computed, and they supported the adequacy of the lag specification.

The cointegration test statistics, $\lambda_{trace}$ and $\lambda_{max}$, are provided in Table 2, along with their corresponding finite sample critical values. The use of proper finite sample critical values is particularly important for our analysis with EMS data, for which the sample length available is not long. According to the results given in Table 2, both the $\lambda_{trace}$ and $\lambda_{max}$ statistics suggest the presence of long-term comovements among the real stock prices. Specifically, the rank of cointegration $r$ equals one, so there are two common permanent components governing the long-term comovements of the variables. Following the Gonzalo–Granger analysis, these permanent components are estimated and reported as $f_{1t}(SP)$ and $f_{2t}(SP)$.

The cointegration evidence implies that long-term comovements exist among at least some countries, but it does not necessarily involve all of the countries. For example, the finding of a single cointegrating vector among French, German and Italian stock markets could be due to existence of a cointegrating relationship between the French and German stock markets only. Italy could be totally out of the picture. Fortunately, this can be tested formally. To check if every country actually belongs to the cointegration system, a $\chi^2(r)$ test for exclusion restrictions on the cointegrating vector $\alpha$ has been conducted. Table 2 contains the test results: every exclusion restriction is rejected by the data, confirming that France, Germany and Italy all belong to the cointegration system.

Since it may be interesting for comparison purposes to study long-term comovements of stock markets in countries outside the EMS, real MSCI stock indexes for Canada, Japan, the UK and the US are examined. These four countries, together with France, Germany and Italy, constitute the so-called G7 countries. For these four countries, the null hypothesis of no cointegration cannot be rejected at any standard level of significance ($\lambda_{trace} = 22.370$ and $\lambda_{max} = 13.650$ based on $k = 2$). The UK had been a member of the EMS briefly for a short time period. Excluding the data for the UK similarly yields no statistically significant evidence of cointegration. In addition, when five stock markets in Canada, Germany, Japan, the UK and the US are considered together, as by Kasa (1992), we also fail to find significant evidence of cointegration. The difference in result between Kasa (1992) and this study may be explained by the difference in the sample period and the correction for finite sample bias in applying the cointegration test here.

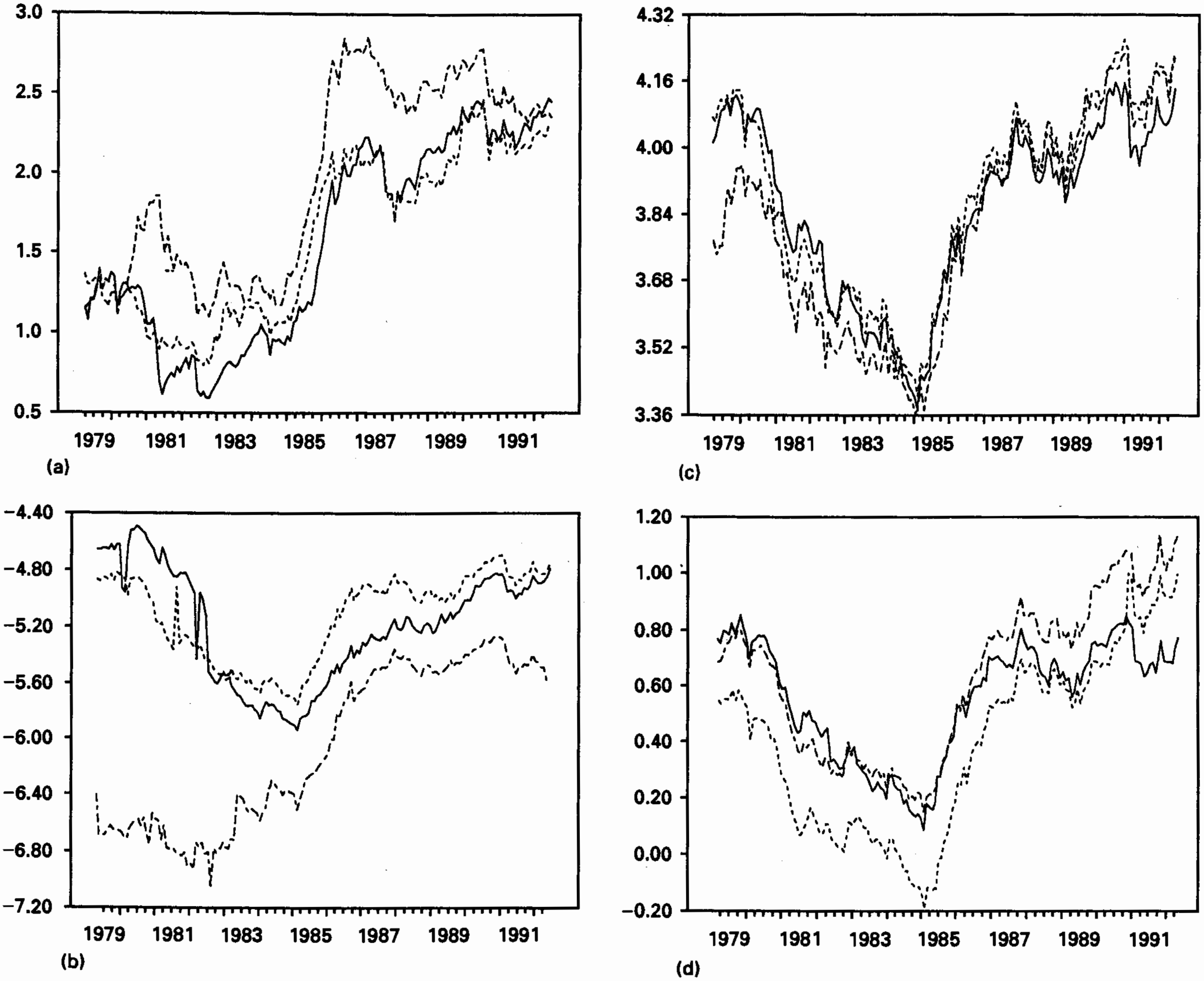


Fig. 1. *Data sources: (a) real stock prices; (b) real stock dividends; (c) real industrial production indexes; (d) real money supply.* (——) *FR*, (- - - -) *GM*, (– – –) *IT*

*Common permanent components in macroeconomic variables*

The finding of long-term comovements among the EMS stock markets naturally leads to the question of what the sources or driving forces of these comovements are. Standard financial valuation models show that stock prices are influenced by fundamental factors determining the state of the economy. To identify the potential sources of stock market comovements, long-term comovements of the real money supply, real dividends and industrial production among the three EMS countries are explored. All these variables are important indicators of general economic activity. The money supply and industrial production variables are considered since relative stock market movements may reflect changes in national output and/or policy coordination across countries (Ripley, 1973). MacDonald and Taylor (1991) have presented some evidence of monetary policy convergence among the EMS members. The money supply can be related to the stock market in various ways. According to the portfolio-balance model, for instance, an increase in money supply leads to portfolio re-balancing towards other assets including securities. Changes in money supply can also affect the stock price through their effects on inflationary expectations and interest rates. Previous studies have reported international evidence showing that money growth rates can have systematic impacts on real stock returns (Mandelker and Tandon, 1985)

With respect to industrial production, this variable is a major indicator of real economic activity and it characterizes general business conditions. Chen (1991) and Fama (1981, 1990) observe that the growth rate of industrial production is a major determinant of long-horizon stock

Table 2. *Cointegration and permanent components in stock prices*

Testing for cointegration:

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 4.845 | 16.150 | 18.485 | 4.315 | 13.227 | 15.117 |
| $r = 0$ | 30.571* | 29.470 | 32.379 | 25.726** | 19.618 | 22.040 |

Restriction tests on the cointegrating vector:

| | $\chi^2$-Test | [p-value] |
|---|---|---|
| Exclude $SP_{FR,t}$ | 16.709 | [0.000] |
| Exclude $SP_{GM,t}$ | 21.410 | [0.000] |
| Exclude $SP_{IT,t}$ | 2.898 | [0.089] |

Common permanent component estimation (equation (2)):

$$\begin{pmatrix} SP_{FR,t} \\ SP_{GM,t} \\ SP_{IT,t} \end{pmatrix} = \begin{pmatrix} 0.098 & 0.041 \\ 0.077 & 0.021 \\ 0.068 & -0.041 \end{pmatrix} \begin{pmatrix} f_{1t}(SP) \\ f_{2t}(SP) \end{pmatrix} + \begin{pmatrix} 0.190 \\ 0.039 \\ 0.083 \end{pmatrix} z_t(SP)$$

where

$$f_{1t}(SP) = -4.822SP_{FR,t} + 16.036SP_{GM,t} + 3.467SP_{IT,t}$$

$$f_{2t}(SP) = 5.375SP_{FR,t} + 6.767SP_{GM,t} - 15.446SP_{IT,t}$$

$$z_t(SP) = 6.616SP_{FR,t} - 9.770SP_{GM,t} + 1.532SP_{IT,t}$$

$\{SP_{FR,t}, SP_{GM,t}, SP_{IT,t}\}$ are the respective real stock price series for France, Germany, and Italy. The lag parameter, $k$, used for model estimation is selected to be 2, using the standard Schwarz information criterion. $r$ denotes the number of the cointegrating vector. The eigenvalue vector, $\lambda$, is estimated to be (0.151, 0.027, 0.003). The 10% and 5% critical values for the cointegration tests, as given next to the relevant test statistics, are computed based on the response surface estimates provided by Cheung and Lai (1993). An asterisk (*) indicates significance at the 10% level. A double asterisk (**) indicates significance at the 5% level. The exclusion test is a likelihood ratio test (Johansen, 1991) examining the null hypothesis that the relevant variable does not belong to the cointegration relationship.

returns, explaining more return variation than other measures of real activity such as real GNP and private investment. The consideration of the dividend variable follows from conventional financial theories. Campbell and Shiller (1987), for example, show that under the present value model, stock prices and dividends are cointegrated, i.e. they share a common permanent component.

Cointegration of the real money supply series across countries is analyzed, and the results are summarized in Table 3. The $\lambda_{trace}$ and $\lambda_{max}$ tests for cointegration seem to yield different results. Although the $\lambda_{trace}$ test fails to find significant evidence of cointegration, the $\lambda_{max}$ test rejects the hypothesis of no cointegration at the 5% significance level. The difference in result may be attributed to the power property of the cointegration tests in small samples. In the analysis below, the hypothesis of one cointegrating vector for the money supply series will be maintained, a conclusion consistent with the previous finding documented by MacDonald and Taylor (1991). Results of exclusion restriction tests on the cointegrating vector further confirm that the three countries all enter into the cointegration relationship. Accordingly, the implied common permanent components are estimated, as given by $f_{1t}(MS)$ and $f_{2t}(MS)$ in Table 3.

Table 4 and Table 5 display the results of both cointegration and common permanent component analyses for dividends and industrial production, respectively. These two cases yield qualitatively similar results. The $\lambda_{trace}$ and the $\lambda_{max}$ tests uniformly suggest the rank of cointegration is one and, again, exclusion restriction testing supports that the cointegration system includes all three EMS countries. The findings imply that there are two common permanent components lying behind the long-run behaviour of both the dividend series and the industrial production series among the EMS countries. The two permanent components in dividends are labelled as $f_{1t}(DI)$ and $f_{2t}(DI)$ in Table 4, and those in industrial production are represented by $f_{1t}(IP)$ and $f_{2t}(IP)$ in Table 5.

In general, the results reported so far indicate that the money supply, dividends, and industrial production across the three EMS countries have a cointegration structure similar to that for these countries' stock prices. Specifically, these individual groups of series display long-term comovements such that they are all cointegrated with the rank equal to one. The evidence, nevertheless, has not yet established any explicit linkage between the long-term comovements among the different groups of variables.

### *Long-term linkages between stock prices and macroeconomic variables*

The question examined next is whether the long-term comovements of stock prices can be linked to those of the

Table 3. *Cointegration and permanent components in money supply series*

Testing for cointegration:

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 4.382 | 15.869 | 18.161 | 3.315 | 12.990 | 14.842 |
| $r = 0$ | 26.077 | 28.959 | 31.812 | 21.695** | 19.266 | 21.639 |

Restriction tests on the cointegrating vector:

| | $\chi^2$-Test | [p-value] |
|---|---|---|
| Exclude $MS_{FR,t}$ | 4.931 | [0.026] |
| Exclude $MS_{GM,t}$ | 16.744 | [0.000] |
| Exclude $MS_{IT,t}$ | 18.289 | [0.000] |

Common permanent component estimation (equation (2)):

$$\begin{pmatrix} MS_{FR,t} \\ MS_{GM,t} \\ MS_{IT,t} \end{pmatrix} = \begin{pmatrix} -0.011 & 0.042 \\ -0.030 & 0.029 \\ -0.025 & 0.031 \end{pmatrix} \begin{pmatrix} f_{1t}(MS) \\ f_{2t}(MS) \end{pmatrix} + \begin{pmatrix} -0.049 \\ -0.060 \\ -0.020 \end{pmatrix} z_t(MS)$$

where

$$f_{1t}(MS) = 37.107MS_{FR,t} - 19.326MS_{GM,t} - 32.758MS_{IT,t}$$

$$f_{2t}(MS) = 26.587MS_{FR,t} - 29.405MS_{GM,t} + 23.815MS_{IT,t}$$

$$z_t(MS) = -5.906MS_{FR,t} - 21.237MS_{GM,t} + 28.280MS_{IT,t}$$

$\{MS_{FR,t}, MS_{GM,t}, MS_{IT,t}\}$ are the real money supply series for France, Germany, and Italy, respectively. The lag parameter, $k$, used for model estimation is selected to be 1, using the standard Schwarz information criterion. The eigenvalue vector, $\lambda$, is estimated to be (0.128, 0.021, 0.007). An asterisk (*) indicates significance at the 10% level. A double asterisk (**) indicates significance at the 5% level. See also Table 2 for other notes.

Table 4. *Cointegration and permanent components in dividend series*

Testing for cointegration:

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 15.536 | 16.150 | 18.485 | 9.029 | 13.227 | 15.117 |
| $r = 0$ | 72.768** | 29.470 | 32.379 | 57.232** | 19.618 | 22.040 |

Restriction tests on the cointegrating vector:

| | $\chi^2$-Test | [p-value] |
|---|---|---|
| Exclude $DI_{FR,t}$ | 38.467 | [0.000] |
| Exclude $DI_{GM,t}$ | 48.049 | [0.000] |
| Exclude $DI_{IT,t}$ | 29.151 | [0.000] |

Common permanent component estimation (equation (2)):

$$\begin{pmatrix} DI_{FR,t} \\ DI_{GM,t} \\ DI_{IT,t} \end{pmatrix} = \begin{pmatrix} 0.079 & -0.022 \\ 0.047 & -0.031 \\ -0.010 & -0.063 \end{pmatrix} \begin{pmatrix} f_{1t}(DI) \\ f_{2t}(DI) \end{pmatrix} + \begin{pmatrix} 0.014 \\ -0.137 \\ 0.009 \end{pmatrix} z_t(DI)$$

where

$$f_{1t}(DI) = 11.650DI_{FR,t} + 0.861DI_{GM,t} - 4.438DI_{IT,t}$$

$$f_{2t}(DI) = -1.228DI_{FR,t} - 1.088DI_{GM,t} - 14.971DI_{IT,t}$$

$$z_t(DI) = 4.253DI_{FR,t} - 6.742DI_{GM,t} + 1.827DI_{IT,t}$$

$\{DI_{FR,t}, DI_{GM,t}, DI_{IT,t}\}$ are the corresponding real dividend series for France, Germany, and Italy. The lag parameter, $k$, used for model estimation is selected to be 2, using the standard Schwarz information criterion. The eigenvalue vector, $\lambda$, is estimated to be (0.305, 0.056, 0.041). An asterisk (*) indicates significance at the 10% level. A double asterisk (**) indicates significance at the 5% level. See also Table 2 for other notes.

macroeconomic variables for these countries. To establish such linkage or the lack of it, possible relationships among the common permanent components of these various cointegration subsystems are analysed. Since the common permanent components fully characterize the long-run behaviour of cointegrated series, these permanent components

Table 5. *Cointegration and permanent components in industrial production series*

Testing for cointegration:

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 8.900 | 15.869 | 18.161 | 8.600 | 12.990 | 14.842 |
| $r = 0$ | 38.143** | 28.959 | 31.812 | 29.243** | 19.266 | 21.639 |

Restriction tests on the cointegrating vector:

| | $\chi^2$-Test | [p-value] |
|---|---|---|
| Exclude $IP_{FR,t}$ | 10.858 | [0.000] |
| Exclude $IP_{GM,t}$ | 19.495 | [0.000] |
| Exclude $IP_{IT,t}$ | 15.803 | [0.000] |

Common permanent component estimation (equation (2)):

$$\begin{pmatrix} IP_{FR,t} \\ IP_{GM,t} \\ IP_{IT,t} \end{pmatrix} = \begin{pmatrix} 0.020 & -0.035 \\ 0.029 & -0.025 \\ 0.041 & -0.011 \end{pmatrix} \begin{pmatrix} f_{1t}(IP) \\ f_{2t}(IP) \end{pmatrix} + \begin{pmatrix} 0.019 \\ -0.008 \\ 0.038 \end{pmatrix} z_t(IP)$$

where

$$f_{1t}(IP) = -24.376 IP_{FR,t} + 25.930 IP_{GM,t} + 17.811 IP_{IT,t}$$

$$f_{2t}(IP) = -33.514 IP_{FR,t} - 0.525 IP_{GM,t} + 16.944 IP_{IT,t}$$

$$z_t(IP) = 16.744 IP_{FR,t} - 28.532 IP_{GM,t} + 12.186 IP_{IT,t}$$

$\{IP_{FR,t}, IP_{GM,t}, IP_{IT,t}\}$ are the real industrial output series for France, Germany and Italy, respectively. The lag parameter, $k$, used for model estimation is selected to be 1, using the standard Schwarz information criterion. The eigenvalue vector, $\lambda$, is estimated to be (0.169, 0.053, 0.002). An asterisk (*) indicates significance at the 10% level. A double asterisk (**) indicates significance at the 5% level. For other notes, see Table 2.

(two of them exist for each variable group) are useful candidates for studying the long-run relationship of one set of variables with other sets of variables.

In examining cointegration between the common permanent components in stock prices and the common permanent components in the macroeconomic variables, the two common permanent components in stock prices are considered separately. The reason is that there can be different underlying economic forces governing these two common permanent components. As a result, additional information may be gained by studying each common permanent component of stock prices individually.

Table 6a contains the results of cointegration analysis between the first common permanent component in stock prices, $f_{1t}(SP)$, and the six permanent components corresponding to those of the money supply, dividends and industrial production. The results from the $\lambda_{trace}$ and $\lambda_{max}$ tests suggest that the rank of cointegration is equal to two. Since finding cointegration in a multivariate system as a whole can obscure the fact that some variable may not actually belong to any cointegration relationship with the others, formal exclusion restriction tests are performed. In our case, for example, it is possible that cointegration may exist among the money supply, dividends and industrial production only, without including stock prices. The results of the restriction tests are reported in Table 6a as well; they confirm that the common permanent component in stock prices does indeed enter into a cointegration relationship with those in the macroeconomic variables. In addition, every macroeconomic variable under study contributes to the cointegration relationship.

Table 6b gives the results of cointegration analysis between the second common permanent component in stock prices, $f_{2t}(SP)$, and all those in the macroeconomic variables. According to both the $\lambda_{trace}$ and $\lambda_{max}$ tests, the rank of cointegration is equal to one, in contrast to the results for the first common permanent component, $f_{1t}(SP)$. The cointegration test results thus suggest the presence of one cointegration relationship among the variables. Further exclusion restriction tests reveal, however, that the stock price variable does not belong to the cointegration relationship. Hence, while the first common permanent component in stock prices is found to be related to the common permanent components in the money supply, dividends and industrial production, the second common permanent component in stock prices is not found to be so.

An interpretation of the above findings is that the long-term comovements of the money supply, dividends and industrial production among the three EMS countries may all contribute to the long-term comovements of these countries' stock prices. However, the relative movements of the stock markets in the long run cannot be fully explained by the relative movements of the set of macroeconomic variables considered.

Table 6. *Long-term linkages between stock prices and macroeconomic variables*

(a) For the 1st common permanent component in stock prices, $f_{1t}(SP)$

Testing for cointegration:

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 2$ | 60.442 | 72.315 | 76.796 | 25.479 | 33.551 | 36.303 |
| $r \leqslant 1$ | 99.244* | 98.317 | 103.534 | 38.802* | 38.796 | 42.957 |
| $r = 0$ | 154.148** | 129.421 | 135.162 | 54.904** | 45.757 | 48.934 |

Restriction tests on the cointegrating vector:

| | $\chi^2$-Test | [p-value] |
|---|---|---|
| Exclude $f_{1t}(SP)$ | 10.415 | [0.005] |
| Exclude $f_{1t}(MS)$ | 18.660 | [0.000] |
| Exclude $f_{2t}(MS)$ | 6.191 | [0.045] |
| Exclude $f_{1t}(DI)$ | 7.585 | [0.023] |
| Exclude $f_{2t}(DI)$ | 13.684 | [0.001] |
| Exclude $f_{1t}(IP)$ | 26.331 | [0.000] |
| Exclude $f_{2t}(IP)$ | 7.753 | [0.021] |

(b) For the 2nd common permanent component in stock prices, $f_{2t}(SP)$

Testing for cointegration:

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 2$ | 56.345 | 72.315 | 76.796 | 21.341 | 33.551 | 36.303 |
| $r \leqslant 1$ | 93.723 | 98.317 | 103.534 | 37.378 | 38.796 | 42.957 |
| $r = 0$ | 150.436** | 129.421 | 135.162 | 56.714** | 45.757 | 48.934 |

Restriction tests on the cointegrating vector:

| | $\chi^2$-Test | [p-value] |
|---|---|---|
| Exclude $f_{2t}(SP)$ | 1.331 | [0.249] |
| Exclude $f_{1t}(MS)$ | 8.130 | [0.000] |
| Exclude $f_{2t}(MS)$ | 2.470 | [0.116] |
| Exclude $f_{1t}(DI)$ | 0.294 | [0.588] |
| Exclude $f_{2t}(DI)$ | 2.841 | [0.092] |
| Exclude $f_{1t}(IP)$ | 17.879 | [0.000] |
| Exclude $f_{2t}(IP)$ | 6.492 | [0.011] |

The lag parameter, $k$, used for model estimation is selected to be 2 in both (a) and (b), using the standard Schwarz information criterion. The eigenvalue vector, $\lambda$, is estimated to be (0.295, 0.219, 0.150, 0.096, 0.080, 0.029, 0.008) in (a) and (0.303, 0.212, 0.127, 0.103, 0.063, 0.039, 0.010) in (b). An asterisk (*) indicates significance at the 10% level. A double asterisk (**) indicates significance at the 5% level. See also Table 2 for other notes.

## V. FURTHER SUBPERIOD ANALYSIS

Studies by Giavazzi and Giovannini (1988, 1989) and Weber (1991), among others, have suggested that the economic and financial linkages among EMS countries are stronger and more evident in the post-1987 period than the pre-1987 period. The relatively frequent and large size of exchange rate realignments in the pre-1987 period reflects the lack of credible macroeconomic policy coordination over this time period. For the post-1987 period, on the other hand, strengthening policy coordination among EMS countries can be observed, as notably signified by the adoption of the September 1987 Basle–Nyborg agreement. Accordingly, the post-1987 period may generate stronger comovement results than the pre-1987 period.

The empirical analysis is repeated for two subperiods: 1979–86 and 1987–92. The results of the subperiod analysis are summarized in Table 7. For the 1979–86 period, no significant evidence of long-term comovements between stock prices and economic fundamentals can be found. For the 1987–92 period, however, the hypothesis of no cointegration between stock prices and economic fundamentals can be rejected at the 10% level. The results of exclusion tests strongly confirm that the common permanent component in stock prices enters into a cointegration relationship with those in the macroeconomic variables. Hence, in contrast to the pre-1987 period, economic fundamentals play a significant role in contributing to the long-term comovements of stock markets among major EMS countries over the post-1987 period.

To further check the robustness of our results for the post-1987 period, two additional issues are examined: (1) the potential effects of the international stock market crash of October 1987 and (2) the sensitivity of the results to the

Table 7. *Results of subperiod analysis*

*1. The 1979–86 period*

(1.a) For the 1st common permanent component in stock prices, $f_{1t}(SP)$

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 91.023 | 105.009 | 110.654 | 34.087 | 42.324 | 46.017 |
| $r = 0$ | 133.595 | 138.230 | 144.458 | 42.572 | 48.979 | 52.420 |
| | | $\chi^2$-Test | [p-value] | | | |
| Exclude $f_{1t}(SP)$ | | 1.008 | [0.315] | | | |

(1.b) For the 2nd common permanent component in stock prices, $f_{2t}(SP)$

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 83.536 | 105.009 | 110.654 | 34.456 | 42.324 | 46.017 |
| $r = 0$ | 129.716 | 138.230 | 144.458 | 46.180 | 48.979 | 52.420 |
| | | $\chi^2$-Test | [p-value] | | | |
| Exclude $f_{2t}(SP)$ | | 2.255 | [0.133] | | | |

*2. The 1987–92 period*

(2.a) For the 1st common permanent component in stock prices, $f_{1t}(SP)$

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 91.658 | 112.839 | 118.985 | 35.132 | 45.581 | 49.597 |
| $r = 0$ | 144.723* | 144.537 | 155.334 | 53.066* | 52.749 | 56.498 |
| | | $\chi^2$-Test | [p-value] | | | |
| Exclude $f_{1t}(SP)$ | | 11.927 | [0.001] | | | |

(2.b) For the 2nd common permanent component in stock prices, $f_{2t}(SP)$

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 99.094 | 112.839 | 118.985 | 31.854 | 45.581 | 49.597 |
| $r = 0$ | 153.732* | 144.537 | 155.334 | 54.638* | 52.749 | 56.498 |
| | | $\chi^2$-Test | [p-value] | | | |
| Exclude $f_{2t}(SP)$ | | 15.518 | [0.000] | | | |

The lag parameter, $k$, used for model estimation is selected to be 2 in all the cases, using the standard Schwarz information criterion. An asterisk (*) indicates significance at the 10% level. A double asterisk (**) indicates significance at the 5% level. See also Table 2 for other notes.

choice of the base country (all the variables for each country have been expressed in real US dollars in the above analysis). The first issue is addressed by conducting cointegration analysis with the inclusion of a dummy variable to control for the possible impact of the October 1987 crash. The second issue is addressed by evaluating results obtained from using Germany, France and Italy as the base country, alternately.

Panel 1 of Table 8 gives the results when the October 1987 crash dummy variable is added to the cointegration analysis. Although the cointegration results are somewhat weaker than in the case of no dummy variable, especially for one of the common permanent component of stock prices, the results in general support that the comovements in stock prices are at least partly linked to those comovements in macroeconomic variables. Panels 2 to 4 of Table 8 report the results of cointegration analysis when Germany, France and Italy are used in turn as the base country, not the US. The results show that our findings are not sensitive to the choice of base country.

Finally, although the analysis has so far focused on dynamic relationships among large EMS countries, it may be interesting to consider some smaller countries, like the Benelux countries, in the analysis. Stock index data for Belgium and the Netherlands are available from our MSCI data set. The following question is explored: Can the influences of the macroeconomic fundamentals of large EMS countries be so strong that they govern the stock market movements of smaller countries, like Belgium and the Netherlands? The foregoing subperiod analysis is repeated to include the stock price series for these two countries. In general, the results indicate no significant evidence of cointegration ($\lambda_{trace} = 134.173$ and $\lambda_{max} = 45.299$ for Belgium and $\lambda_{trace} = 134.790$ and $\lambda_{max} = 47.927$ for the Netherlands) at any standard levels of significance. Of course, the statistical failure to find long-term comovements does not absolutely show the complete absence of such comovements. Moreover, the results in no way preclude the presence of other forms of interactions between large and small EMS countries, on which future research is certainly warranted.

Table 8. *Further analysis of the 1987–92 period*

*1. Including a dummy variable for the October 1987 crash*

(1a) For the 1st common permanent component in stock prices, $f_{1t}(SP)$

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 85.511 | 112.839 | 118.985 | 32.992 | 45.581 | 49.597 |
| $r = 0$ | 138.268 | 144.537 | 155.334 | 52.757* | 52.749 | 56.498 |
| | | $\chi^2$-Test | [p-value] | | | |
| Exclude $f_{1t}(SP)$ | | 11.641 | [0.001] | | | |

(1b) For the 2nd common permanent component in stock prices, $f_{2t}(SP)$

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 95.954 | 112.839 | 118.985 | 31.857 | 45.581 | 49.597 |
| $r = 0$ | 150.868* | 144.537 | 155.334 | 54.914* | 52.749 | 56.498 |
| | | $\chi^2$-Test | [p-value] | | | |
| Exclude $f_{2t}(SP)$ | | 16.950 | [0.000] | | | |

*2. Using Germany as the base country*

(2a) For the 1st common permanent component in stock prices, $f_{1t}(SP)$

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 86.133 | 112.839 | 118.985 | 32.244 | 45.581 | 49.597 |
| $r = 0$ | 142.242 | 144.537 | 155.334 | 56.109* | 52.749 | 56.498 |
| | | $\chi^2$-Test | [p-value] | | | |
| Exclude $f_{1t}(SP)$ | | 11.123 | [0.001] | | | |

(2b) For the 2nd common permanent component in stock prices, $f_{2t}(SP)$

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 91.017 | 112.839 | 118.985 | 30.008 | 45.581 | 49.597 |
| $r = 0$ | 149.761* | 144.537 | 155.334 | 58.744** | 52.749 | 56.498 |
| | | $\chi^2$-Test | [p-value] | | | |
| Exclude $f_{2t}(SP)$ | | 14.535 | [0.000] | | | |

*3. Using France as the base country*

(3a) For the 1st common permanent component in stock prices, $f_{1t}(SP)$

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 86.594 | 112.839 | 118.985 | 31.930 | 45.581 | 49.597 |
| $r = 0$ | 141.927 | 144.537 | 155.334 | 55.333* | 52.749 | 56.498 |
| | | $\chi^2$-Test | [p-value] | | | |
| Exclude $f_{1t}(SP)$ | | 11.095 | [0.001] | | | |

(3b) For the 2nd common permanent component in stock prices, $f_{2t}(SP)$

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 94.175 | 112.839 | 118.985 | 31.651 | 45.581 | 49.597 |
| $r = 0$ | 151.381* | 144.537 | 155.334 | 57.206** | 52.749 | 56.498 |
| | | $\chi^2$-Test | [p-value] | | | |
| Exclude $f_{2t}(SP)$ | | 13.468 | [0.000] | | | |

*4. Using Italy as the base country*

(4a) For the 1st common permanent component in stock prices, $f_{1t}(SP)$

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 89.562 | 112.839 | 118.985 | 31.639 | 45.581 | 49.597 |
| $r = 0$ | 142.296 | 144.537 | 155.334 | 52.783* | 52.749 | 56.498 |
| | | $\chi^2$-Test | [p-value] | | | |
| Exclude $f_{1t}(SP)$ | | 11.248 | [0.001] | | | |

(4b) For the 2nd common permanent component in stock prices, $f_{2t}(SP)$

| $H_0$ | $\lambda_{trace}$ | 10% | 5% | $\lambda_{max}$ | 10% | 5% |
|---|---|---|---|---|---|---|
| $r \leqslant 1$ | 97.279 | 112.839 | 118.985 | 32.647 | 45.581 | 49.597 |
| $r = 0$ | 158.344** | 144.537 | 155.334 | 61.065** | 52.749 | 56.498 |
| | | $\chi^2$-Test | [p-value] | | | |
| Exclude $f_{2t}(SP)$ | | 14.535 | [0.000] | | | |

Statistical significance is indicated by an asterisk (*) for the 10% level or a double asterisk (**) for the 5% level. The lag parameter, $k$, used for model estimation is selected to be 2 in all the cases, using the standard Schwarz information criterion. See also Table 2 for other notes.

## VI. CONCLUDING REMARKS

The advent of the EMS since March 1979 has revived and noticeably expanded the literature on international monetary systems. The experience of the EMS is interesting to study, and it can have impacts on policy and economic analyses. This study explores the issue concerning the potential existence of long-term comovements of national stock markets in light of the experience of the EMS. Although long-term comovements of national stock markets have been investigated in previous empirical work, little has been done to study those of EMS stock markets and also investigate the possible sources of their comovements. Why can national stock markets exhibit long-term comovements? Are such long-term market comovements rational? To what extent do they reflect similar long-term comovements in economic fundamentals? All these questions point to the need to analyse and identify the possible sources of long-term market comovements.

In this study the long-term comovements of stock markets in three EMS countries – France, Germany and Italy – have been examined based on monthly data for the period 1979–1992. Empirical results support that the stock markets in these countries display long-term comovements. The results also suggest the presence of two common permanent components driving the long-run dynamics of these stock markets. To identify some interpretable sources that can explain the long-term stock market comovements among these EMS countries, this study further explores whether the common permanent components found in stock prices can be linked to the relative dynamics of macroeconomic variables among the EMS countries. The macroeconomic variables under study include the money supply, dividends and industrial production. Interestingly, empirical evidence suggests that, like the stock price variable, there are two common permanent components governing the long-term comovements among the EMS countries in each of these macroeconomic variables. To investigate the possible linkage between the long-term comovements of these variables and those of stock prices, cointegration among the common permanent components in stock prices and the macroeconomic variables is analysed. It is found that although part of the long-term comovements of stock prices can be attributed to those comovements of several macroeconomic variables (especially for the post-1987 period), the explanatory power of the latter is far from strong. Nonetheless, the results confirm at least a limited role of these macroeconomic variables in accounting for the relative stock market movements among the three EMS countries. It should be emphasized that the findings here apply to relative movements of different stock markets, not movements of individual stock markets, which prior studies have primarily examined.

Can the missing or weak link between the comovements of EMS stock markets and those of economic fundamentals be explained in terms of market psychology? Possibly because of their geographical proximity or the potential monetary link implied by the EMS exchange rate mechanism, investors may simply follow price movements in different major EMS stock markets. Such behaviour, which is motivated by non-fundamentals, can arguably be 'rational' as long as most investors behave in the similar manner. However, even though the market psychology approach may be interesting in explaining short-term market comovements, its ability to explain long-term market comovements appears questionable unless the market psychology can eventually be supported by similar movements in economic fundamentals.

Another interpretation of the results of this study is the missing-variable perspective. The variables examined here are important macroeconomic indicators, but they by no means exhaustively characterize the macroeconomy. To the extent that stock prices can be influenced by a wide variety of economic factors, it is possible that some other variables not considered in this study may also contribute to the long-term comovements of stock prices. Future research to identify the relevant omitted variables should be of interest for a better understanding of the dynamic behaviour of stock markets. The present study can be viewed as a preliminary step in exploring the rational basis of long-term stock market comovements.

## REFERENCES

Agmon, T. (1972) The relations among equity markets: A study of share price comovements in the United States, United Kingdom, Germany and Japan, *Journal of Finance*, **27**, 839–55.

Campbell, J. Y. and Shiller, R. J. (1987) Cointegration and tests of present value models, *Journal of Political Economy*, **95**, 1062–88.

Chen, N. F. (1991) Financial investment opportunities and the macroeconomy, *Journal of Finance*, **46**, 529–54.

Chen, N. F., Roll, R. and Ross, S. A. (1986) Economic forces and the stock market, *Journal of Business*, **59**, 383–403.

Cheung, Y. W. and Lai, K. S. (1993) Finite-sample sizes of Johansen's likelihood ratio tests for cointegration, *Oxford Bulletin of Economics and Statistics*, **55**, 313–28.

Cheung, Y. W. and Lai, K. S. (1995) Lag order and critical values of the augmented Dickey–Fuller test, *Journal of Business and Economic Statistics*, **13**, 277–80.

Corhay, A. and Rad, A. T. (1994) Daily returns from European stock markets, *Journal of Business and Finance and Accounting*, **21**, 271–81.

Corhay, A., Rad, A. T. and Urbain, J.-P. (1993) Common stochastic trends in European stock markets, *Economic Letters*, **42**, 385–90.

Dwyer, G. P. and Hafer, R. W. (1988) Are national stock markets linked? *Federal Reserve Bank of St. Louis Review*, **70**, 3–14.

Escribano, A. and Peña, D. (1994) Cointegration and common factors, *Journal of Time Series Analysis*, **15**, 577–86.

Eun, C. S. and Shim, S. (1989) International transmission of stock market movements, *Journal of Financial and Quantitative Analysis*, **24**, 241–56.

Fama, E. F. (1981) Stock returns, real activity, inflation, and money, *American Economic Review*, **71**, 545–65.

Fama, E. F. (1990) Stock returns, expected returns, and real activity, *Journal of Finance*, **45**, 1089–1108.

Fama, E. F. and French, K. R. (1988) Permanent and temporary components of stock prices, *Journal of Political Economy*, **96**, 246–73.

Fraser, P., Helliar, C. V. and Power, D. W. (1994) An empirical investigation of convergence among European stock equity markets, *Applied Financial Economics*, **4**, 149–57.

Fratianni, M. and von Hagen, J. (1990) The European Monetary System ten years after, *Carnegie-Rochester Conference Series on Public Policy*, **31**, 173–242.

Gallagher, L. (1995) Interdependencies among Irish, British and German stock markets, *The Economic and Social Review*, **26**, 131–47.

Giavazzi, F. and Giovannini, A. (1988) Can the EMS be copied outside Europe? Lessons from ten years of monetary policy coordination in Europe, NBER Working Paper No. 2786.

Giavazzi, F. and Giovannini, A. (1989) *Limiting Exchange Rate Flexibility: The European Monetary System*, MIT Press; Cambridge, MA.

Gonzalo, J. and Granger, C. W. J. (1995) Estimation of common long-memory components in cointegrated systems, *Journal of Business and Economic Statistics*, **13**, 27–35.

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

Hilliard, J. E. (1979) The relationship between equity indices on world exchanges, *Journal of Finance*, **34**, 103–14.

King, M., Sentana, E. and Wadhwani, S. (1994) Volatility and links between national stock markets, *Econometrica*, **62**, 901–33.

King, M. and Wadhwani, S. (1990) Transmission of volatility between stock markets, *Review of Financial Studies*, **3**, 5–33.

Johansen, S. (1991) Estimation and hypothesis testing of cointegrating vectors in gaussian vector autoregressive models, *Econometrica*, **59**, 1551–80.

Kasa, K. (1992) Common stochastic trends in international stock markets, *Journal of Monetary Economics*, **29**, 95–124.

Koch, P. D. and Koch, T. W. (1991) Evolution in dynamic linkages across daily national stock indexes, *Journal of International Money and Finance*, **10**, 231–51.

Lai, M., Lai, K. S. and Fang, H. (1993) Dynamic linkages between the New York and Tokyo stock markets: a vector error correction analysis, *Journal of International Financial Markets, Institutions and Money*, **3**, 73–96.

Levy, H. and Sarnat, M. (1970) International diversification of investment portfolios, *American Economic Review*, **60**, 668–75.

MacDonald, R. and Taylor, M. P. (1991) Exchange rates, policy convergence, and the European Monetary System, *Review of Economics and Statistics*, **73**, 553–8.

Mandelker, G. and Tandon, K. (1985) Common stock returns, real activity, money, and inflation: some international evidence, *Journal of International Money and Finance*, **4**, 267–86.

Mathur, I. and Subrahmanyam, V. (1990) Interdependencies among the Nordic and the U.S. stock markets, *Scandinavian Journal of Economics*, **92**, 587–97.

Panton, D. B., Lessig, V. P. and Joy, O.M. (1976) Comovement of international equity markets: A taxonomic approach, *Journal of Financial and Quantitative Analysis*, **11**, 415–32.

Poterba, J. M. and Summers, L. H. (1988) Mean reversions in stock prices: Evidence and implications, *Journal of Financial Economics*, **22**, 27–59.

Ripley, D. M. (1973) Systematic elements in the linkage of national stock market indices, *Review of Economics and Statistics*, **55**, 356–61.

Schollhammer, H. and Sand, O. C. (1987) Lead–lag relationships among national equity markets: An empirical investigation, in S.J. Khoury and A. Ghosh (eds), *Recent Developments in International Banking and Finance*, Vol. 1, Lexington Books, Lexington, MA.

Stock, J. H. and Watson, M. W. (1988) Testing for common trends, *Journal of the American Statistical Association*, **83**, 1097–1107.

Taylor, M. P. and Tonks, I. (1989) The internationalization of stock markets and the abolition of UK exchange control, *The Review of Economics and Statistics*, **71**, 332–6.

Von Furstenberg, G. M. and Jeon, B. N. (1989) International stock price movements: Links and messages, *Brookings Papers on Economic Activity*, **1**, 125–67.

Weber, A. A. (1991) Reputation and credibility in the European Monetary System, *Economic Policy*, **12**, 58–102.