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Assessing financial integration in European Union equity markets, 1990-2006: Panel unit root and multivariate cointegration and causality evidence

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Assessing financial integration in European Union equity markets, 1990-2006: Panel unit root and multivariate cointegration and causality evidence

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Abstract. This paper measures financial integration among selected European Union equity markets over the period July 1990 to June 2006 using daily data. Eleven markets (Austria, Belgium, Denmark, France, Germany, Greece, Ireland, Italy, Netherlands, Spain and the United Kingdom) are included in the analysis. Panel unit root tests are used to test for non-stationarity, and multivariate cointegration, Granger causality and level VAR procedures and variance decompositions are conducted to examine the equilibrium and causal relationships among these markets. The results indicate that there is a stationary long-run equilibrium relationship among, and significant and substantial short and long-run causal linkages between, these markets. The findings offer complementary evidence that a high level of financial integration prevails in the region.

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1. Introduction

Financial integration is the process by which a country's or region's financial markets – including its money, bond, bank credit and equity markets – become more closely integrated with those in other countries or regions. More particularly, the market for a given set of financial instruments and/or services is said to be fully integrated if all potential market participants with the same relevant characteristics: (i) face a single set of rules when they deal with these financial instruments and/or services; (ii) have equal access to the set of financial instruments and/or services; and (iii) are treated equally when they are active in the market (Baele et al. 2004: 6).

Three benefits are thought to accrue from the process of financial integration: more opportunities for risk sharing and diversification, the better allocation of capital across investment opportunities, and the potential for higher economic growth. First, sharing risk across regions enhances specialisation, increases the set of financial instruments and/or services available, and thereby provides additional possibilities for diversification by investors. Second, the elimination of barriers to trading, clearing and settlement allows firms

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to choose the most efficient location for their financing activities. Investors too are free to invest their funds where they will be allocated to their most productive end-use. Finally, the improvement in capital allocation enhances financial development, thereby assisting the process of economic growth, with additional funds flowing to (often less-developed) countries or regions with more (and often better) productive opportunities.

The European Union, currently celebrating the fiftieth anniversary of its establishment, is a potential exemplar of the process of financial integration. Starting with the six-member European Economic Community in 1957, the European Union is now the world's largest economic entity with a nominal GDP of €11.5 ($15.0 USD) trillion spread across twenty-seven member states, including the thirteen members of the single-currency euro area. Obviously, financial integration has been an ongoing goal of the European Union with an early emphasis placed on the elimination of cross-border restrictions on the activities of firms and investors within the region, as well as the harmonisation of rules, taxes and regulations among member states. More recently, however, the pace of these changes has accelerated, alongside a surge in cross-border trading. For instance, in the last few years the Financial Services Action Plan has been established as the vehicle for developing a single market in financial services in the European Union, with more than forty measures to be implemented in the areas of banking, securities, insurance and pensions, and asset management (European Commission 2007). At the same time, the European System of Central Banks and the European Central Bank have since 1998 focused on financial integration as a means of achieving their primary objective of price stability alongside a high level of employment and sustainable and non-inflationary growth. This has resulted in series of regular updates on the pace and progress of financial integration by both the European Commission (2006) and the European Central Bank (2007).

In a recent European Central Bank occasional paper, Baele et al. (2004) identify several developments, particularly in equity markets, that suggest that financial integration has increased substantially in the European Union. First, equity market participation by all types of investors has increased considerably, with equity as a share of financial assets held almost doubling between 1995 and 1999 (almost certainly associated with aging populations and the supplementation of public pensions with personal retirement savings). Second, the convergence of interest rates across euro area countries to historically low levels has prompted a reallocation of investments towards equity markets. Third, a number of European

Union directives have removed many of the few remaining barriers to international equity investment. Fourth, rapid growth in the number of investment funds has made it easier for investors to construct well-diversified portfolios.

Finally, with the introduction of the single currency in 1999, a structural shift occurred in the portfolio allocation paradigm, with investors increasingly convinced that the traditional first step of the international asset allocation decision in terms of country selection should give way to industry or sector selection, at least in the European Union (Baele et al. 2004). In turn, the heightened interest in cross-border equity trading has led the region's stock exchanges to expand across national borders, with the consolidation of existing exchanges and attempts to create pan-European exchanges: complicated in part by cross-country regulatory differences and the fragmentation in clearing and settlement systems (Baele et al. 2004).

Baele et al. (2004) use this evidence to argue for the monitoring and understanding of financial market integration. The reasons are as follows. First, while the benefits of financial integration are expected to be positive overall, less positive effects may arise where, say, excessive consolidation in a market segment hinders competition. Second, it is important to accurately measure the state of integration in various segments of the market so that areas where further initiatives are required are identified. Third, since monetary policy is implemented through the financial system, this system must be as efficient as possible in order to guarantee the smooth and effective transmission of monetary policy. Finally, financial integration affects the structure of the financial system, which in turn may have implications for financial stability. Monitoring integration is therefore important for regulators and central banks.

In Baele et al. (2004), the relative importance of sector and country effects, the proportion of local equity market variance explained by common factors, and changes in equity home bias are used separately to assess the degree of financial integration. But a complementary approach exists in the form of multivariate cointegration, causality and variance decomposition methods to examine these sorts of pricing relationships. This builds upon a continuously evolving literature concerned with financial market integration, comprising studies addressing the integration of European member-states with global markets [see, for instance, Arshanapalli and Doukas (1993), Abbott and Chow (1993), Espitia and Santamaria (1994), Kwan et al. (1995), Richards (1995), Longin and Solnik (1995), Malliaris and Urrutia

(1996), Solnik et al. (1996), Darbar and Deb (1997), Meric and Meric (1997), Shawky et al. (1997), Yuhn (1997), Francis and Leachman (1998), Ramchand and Susmel (1998), Masih and Masih (1999) and Cheung and Lai (1999)] and a relatively recent body of work focusing on the institutional and regulatory aspects and outcomes of integration within Europe, especially the role of European Monetary Union (EMU) and the single currency [see, for example, Cheung and Lai (1999), Rouwenhorst (1999), Frantzscher (2002), Worthington et al. (2003), Hartmann et al. (2003), Jian et al. (2003), Reszar (2005), Batten and Kearney (2006), Schotman and Zalewska (2006), Fonteyne (2006), Hardouvelis (2006), Kim et al. (2006) and Papadogonona and Stouraras (2006)].

Accordingly, the purpose of this paper is to present a quantitative method for assessing financial integration in European Union equity markets. The paper itself is divided into four main areas. The second section presents the data employed in the analysis. The third section explains the methodology. The results are dealt with in the fourth section. The paper ends with some brief concluding remarks.

2.2 Data

The data employed in the study is composed of value-weighted equity market indices for eleven European markets, namely, Austria (AUS), Belgium (BEL), Denmark (DEN), France (FRA), Germany (GER), Greece (GRE), Ireland (IRL), Italy (ITL), Netherlands (NTH), Spain (SPN) and the United Kingdom (UK). While the sample of member states is not exhaustive, it does include the largest eleven of the fifteen members in place before the 2004 and 2007 waves of accession (with ten and two new members, respectively). All index data specified is obtained from Morgan Stanley Capital International-Barra (2007) (hereafter MSCI) in US dollar terms and encompasses the period 1 January 1993 to 31 June 2006. The construction of these indices is as follows:

> In constructing a country index every listed security in the market is identified. Securities are free float adjusted, classified in accordance with the Global Industry Classification Standard (GICS®), and screened by size and liquidity. MSCI then constructs its indices by targeting for index inclusion 85% of the free float adjusted market capitalization in each industry group, within each country. By targeting 85% of each industry group, the MSCI Country Index captures 85% of the total country market capitalization while it accurately reflects the economic diversity of the market.

MSCI indices are widely employed in the financial integration literature given the degree of comparability, the avoidance of dual listing and the breadth and reflectivity of index coverage [see, for instance, Meric and Meric (1997), Yuhn (1997), Cheung and Lai (1999) and Worthington *et al*. (2003)]. The daily data used comprise the longest continuous time series for the eleven European equity markets. Each market encompasses 4,175 daily observations; the eleven markets together provide a balanced panel of 45,925 observations.

3. Empirical methodology

This paper investigates the integration among European Union equity markets as follows. Panel unit root tests are first conducted as a means of informing subsequent techniques. Multivariate cointegration, Granger causality, level VAR and variance decomposition methods are then employed to examine the integration among markets.

3.1 Panel unit root tests

Panel unit root tests comprise a multivariate analogue to standard univariate unit root tests, including the Augmented Dickey-Fuller (ADF), Phillips-Perron (PP) and Kwiatkowski, Phillips, Schmidt and Shin (KPSS) tests. The main purpose in extending the application of purely time-series unit root tests to panel unit root tests is to use the increase in sample size from pooling cross-sectional data to improve the power of the tests. Three panel unit root tests are examined, namely: the Levine, Lin and Chu (2002), Im, Pesaran and Shin (2003) and Hadri (2000) tests.

(i) A basic model

Assume the time series $\{y_{i,0}, \ldots, y_{i,T}\}$ on the cross section units (or markets) $i = 1, 2, \ldots, M$ over T time periods are generated for each i by a simple first-order autoregressive, AR(1), process:

$$y_{i,t} = (1-\rho_i)\mu_i + \rho_i y_{i,t-1} + \varepsilon_{i,t} \qquad i = 1, 2, ..., M, \quad t = 1, 2, ..., T \tag{1}$$

where $y_{i,t}$ denotes the observed cross section for the i-th unit at time t and $\varepsilon_{i,t}$ is white noise for the i-th unit at time t. The errors $\varepsilon_{i,t}$ are identically and independently distributed (*i.i.d*) across i and t with $E(\varepsilon_{i,t}) = 0$, $E(\varepsilon_{i,t}^2) = \sigma_i^2 < \infty$ and $E(\varepsilon_{i,t}^4) < \infty$. Under the null hypothesis of a unit root, ρ_i = 1 for all i, equation (1) can be rewritten as the following basic ADF specification:

$$\Delta y_{i,t} = \alpha_i + \phi_i y_{i,t-1} + \sum_{j=1}^{q_i} \gamma_{i,j} \Delta y_{i,t-j} + \varepsilon_{i,t} \tag{2}$$

where $\alpha = (1 - \rho_i)\mu_i$, $\phi_i = (\rho_i - 1)$ and γ_i are coefficients to be estimated for the i-th unit, q_i is the number of lagged terms for the i-th unit $\Delta y_{i,t} = y_{i,t} - y_{i,t-1}$ and all other parameters are as previously defined.

(ii) Levine, Lin and Chu test

One of the first panel unit root tests was proposed by Levine and Lin (1992) and subsequently formalised in Levine et al. (2002) (hereafter LLC). The LLC test permits the intercept, time trend, residual variance and higher-order autocorrelations to vary across individual markets. The LLC test is based on a pooled panel estimator which assumes a common $\phi_i = \phi$ but allows q_i to vary across the cross sections. It also requires the independently generated time series to have a common sample size. The LLC test may then be viewed as a pooled ADF test potentially with different lag lengths across the cross sections of the panel. The main limitation of this test is that it imposes a cross-equation restriction on the first-order autocorrelation coefficients. Under the LLC, the null and alternative hypotheses are given as:

$H_{0,LLC}$: $\phi_1 = \phi_2 = \ldots = \phi_M = 0$

$H_{1,LLC}$: $\phi_1 = \phi_2 = \ldots = \phi_M = \phi < 0$

Under the null hypothesis, each cross section has a unit root (or is non-stationary) while under the alternative each cross section unit is stationary. The LLC test statistic under the null hypothesis is a modified t-statistic.

(iii) Im, Pesaran and Shin test

The Im, Pesaran and Shin (2003) test (herafter IPS) is introduced to take account of the major weakness of the LLC test where it is assumed that all individual AR(1) series have a common autocorrelation coefficient. It allows for individual processes by permitting ϕ_i to vary across the cross sections. The IPS test begins by specifying a separate ADF regression for each cross section unit specified by equation (2). The null and alternative hypotheses for the IPS test are:

$H_{0,IPS}$: $\phi_i = \phi = 0$ for $\forall\ i$

$H_{1,IPS}$: $\phi_i < 0$ for $i = 1, 2, \ldots, M_1$ and $\phi_i = 0$ for $i = M_1 + 1, \ldots, M$

Under the null hypothesis, all cross section units in the panel are non-stationary. The IPS test assumes that under the alternative at least one cross section unit, but not all cross section units is stationary. This differs from the LLC test which presumes all cross section units are stationary under the alternative hypothesis.

The IPS test is based on M independent tests on M cross section units while the LLC test combines the test statistics. The random errors, $\varepsilon_{i,t}$, are assumed to be serially correlated with different serial correlation properties and different variances across each cross section unit. The core of the IPS test is based on a group-mean t-bar statistic where the t-statistics are drawn from each ADF test and averaged across the panels. Adjustment factors are used to standardise the t-bar statistic into a standard normal IPS W-statistic under the null hypothesis.

(iv) Hadri test

The Hadri (2000) panel unit root test parallels the well-known KPSS unit root test with the null hypothesis of no unit root in any of the cross section units in the panel. As with the KPSS unit root test, the Hadri test is based on the residuals from individual OLS regressions of $y_{i,t}$ on a constant or a constant and a trend. The test statistic is distributed as standard normal under the null. The error process may be assumed to be homoskedastic across the panel or heteroskedastic across the cross section units. Two Z-statistics are presented. One Z-statistic is derived from the Lagrange Multiplier (LM) statistic where the residuals from the ADF regression are associated with the homoskedasticity assumption across the panel and the other using the LM statistic that is heteroskedasticity consistent.

3.2 Multivariate cointegration

Following Engle and Granger (1987), suppose the set of M market index series $y_t = [y_{1t,} \, y_{2t}, \cdots, y_{Mt}]'$ are all I(1) and $\beta' y_t = u_t$ is I(0), then β is said to be a cointegrated vector and $\beta' y_t = u_t$ is called the cointegrating regression. The components of y_t are said to be cointegrated of order d, denoted by $y_t \sim \mathrm{CI}(d, b)$ where $d > b > 0$, if (i) each component of y_t is integrated of order d, and (ii) there exists at least one vector $\beta = (\beta_1, \beta_2, \ldots, \beta_M)$, such that the linear combination is integrated of (d - b). By Granger's theorem, if the indices are cointegrated, they can be expressed in an error correction model (ECM) encompassing the notion of a long-run equilibrium relationship and the introduction of past disequilibrium as explanatory variables in the dynamic behaviour of current variables.

In order to implement the ECM, the order of cointegration must be known. A useful statistical test for determining the cointegration order proposed by Johansen (1991) and Johansen and Juselius (1990) is the trace test. For example, to test for no cointegrating relationship, *r* is set to zero and the null hypothesis is $H_0 : r = 0$ and the alternative is $H_1 : r > 0$. However, the Johansen (1991) test can be affected by the lag order. The lag order is determined by using both the likelihood ratio test and information criteria in VAR. The optimum number of lags to be used in the VAR models is determined by the likelihood ratio test statistic:

$$LR = (T - K)\ln(|\Sigma_0| / |\Sigma_A|) \tag{3}$$

where *T* is the number of observations, *K* denotes the number of restrictions, Σ denotes the determinant of the covariance matrix of the error term, and subscripts *0* and *A* denote the restricted and unrestricted VAR, respectively. LR is asymptotically distributed χ^2 with degrees of freedom equal to the number of restrictions. The test statistic in (3) is used to test the null hypothesis of the number of lags being equal to *k*-1 against the alternative hypotheses that *k* = 2, 3, … and so on. The test procedure continues until the null hypothesis fails to be rejected, thereby indicating the optimal lag corresponds to the lag of the null hypothesis.

3.3 Multivariate Granger causality and level VAR tests

To examine the short-run relationships among the markets, Granger (1969) causality tests are specified. Essentially tests of the prediction ability of time series models, a market index causes another index in the Granger sense if past values of the first index explain the second, but past values of the second index do not explain the first. When the indices in question are cointegrated, Granger causality is tested using the ECM:

$$\Delta y_t = \gamma_0 + \sum_{i=1}^{r} \psi_i \Theta_{t-1} + \sum_{i=1}^{m} \gamma_i \Delta y_{t-i} + \varepsilon_t \tag{4}$$

where Θ contains *r* individual error-correction terms, *r* are long-term cointegrating vectors via the Johansen procedure, ψ and γ are parameters to be estimated, and all other variables are as previously defined.

One problem with a Granger causality test based on (4) is that it is affected by the specification of the model. ECM is estimated under the assumption of a certain number of lags and cointegrating equations, which means that the actual specification depends on the pre-test unit root and cointegration (Johansen) tests. To avoid possible pre-test bias, Toda and Yamamoto (1995) propose the level VAR procedure. Essentially, the level VAR procedure is based on VAR for the level of variables with the lag order p in the VAR equations given by $p=k+d_{max}$, where k is the true lag length and d_{max} is the possible maximum integration order of variables. Therefore, the estimated VAR is expressed as:

$$y_t = \hat{\gamma}_0 + \hat{\gamma}_1 t + \cdots + \hat{\gamma}_q t^q + \hat{J}_1 y_{t-1} + \cdots + \hat{J}_k y_{t-k} + \cdots + \hat{J}_p y_{t-p} + \hat{\varepsilon}_t, \tag{5}$$

where $t = 1, \ldots, T$ is the trend term and $\hat{\gamma}_i, \hat{J}_j$ are parameters estimated by OLS. Note that d_{max} does not exceed the true lag length k. Equation (5) can be written as:

$$Y' = \hat{\Gamma}\Lambda + \hat{\Phi}X + \hat{\Psi}Z' + \hat{\mathrm{E}}' \tag{6}$$

where $\hat{\Gamma} = (\hat{\gamma}_0, \ldots, \hat{\gamma}_q)$, $\Lambda = (\tau_1, \ldots, \tau_T)$ with $\tau_t = (1, t, \ldots, t^q)$, $\hat{\Phi} = (\hat{J}_1, \ldots, \hat{J}_k)$, $\hat{\Psi} = (\hat{J}_{k+1}, \ldots, \hat{J}_p)$, $X = (x_1, \cdots, x_T)$ with $x_t = (y'_{t-1}, \ldots, y'_{t-k})'$, $Z = (z_1, \cdots, z_T)$ with $z_t = (y'_{t-k-1}, \ldots, y'_{t-p})'$ and $\mathrm{E} = (\hat{\varepsilon}_1, \ldots, \hat{\varepsilon}_T)$. As restrictions in parameters, the null hypothesis $H_0 : f(\phi) = 0$ where $\phi = vec(\Phi)$ is tested by a Wald statistic defined as:

$$W = f(\hat{\phi})' \left[F(\hat{\phi}) \left\{ \hat{\Sigma}_\varepsilon \otimes (X'QX)^{-1} \right\} F(\hat{\phi})' \right]^{-1} f(\hat{\phi}) \tag{7}$$

where $F(\phi) = \partial f(\phi) / \partial \phi'$, $\hat{\Sigma}_\varepsilon = T^{-1}\hat{\mathrm{E}}'\hat{\mathrm{E}}$, $Q = \hat{Q}_\tau - \hat{Q}_\tau Z(Z'\hat{Q}_\tau Z)^{-1} Z'\hat{Q}_\tau$ and

$\hat{Q}_\tau = I_T - \hat{\Lambda}(\hat{\Lambda}'\hat{\Lambda})^{-1}\hat{\Lambda}'$ where I_T is a $T \times T$ identity matrix. Under the null hypothesis, the Wald statistic (7) has an asymptotic Chi-square distribution with m degrees of freedom that corresponds to the number of restrictions. Although Toda and Yamamoto (1995) present this method principally for the purpose of Granger-causality testing, tests based on level VAR equations can also be used to examine long-run relationships. Test results based on the ECM can then be regarded as an indicator of short-run causality, while the causality tests by the level VAR can complement the result of the cointegration tests in terms of long-run information.

3.4 Variance decomposition

One limitation of these tests is that while they indicate which markets Granger-cause another, they do not indicate whether yet other markets can influence a market through other equations in the system. Likewise, Granger causality does not provide an indication of the dynamic properties of the system, nor does it allow the relative strength of the Granger-causal chain to be evaluated. However, decomposition of the variance of forecast errors allows the relative importance of the variance in causing fluctuations in that market to be ascertained. The decomposition process therefore allows the variance of the forecast errors to be divided into percentages attributable to innovations in all other markets and a percentage attributable to innovations in the market of interest. One problem here is that the decomposition of variances is sensitive to the assumed origin of the shock and the order it is transmitted to other markets. To overcome this problem, a generalised impulse response analysis, which is not subject to any arbitrary othogonalisations of innovations in the system, is applied.

The variance decomposition analysis illustrates the system dynamics by decomposing the random variation of one market into component shocks and analysing how these shocks in turn affect prices in other markets. Consider the following VAR model of *m* market indices proposed by Eun and Shim (1989: 243):

$$y_t = \alpha + \sum_{S=1}^{n} A(S) y_{t-S} + e_t \tag{8}$$

where y_t is a $m\times1$ vector of indices, α and *A(S)* are respectively $m\times1$ and $m\times m$ coefficients, n is the lag length, and e_t is a $m\times1$ column of forecast errors of the best linear predictor of y_t using past values of y. By construction, if the forecast error e_t is uncorrelated with all past values of y and is also a linear combination of current and past y_t, then e_t is serially uncorrelated. The i,j component of *A(S)* measures the direct effect of the jth market on the ith market in S periods. As shown by Sim (1980), by the successive substitution of $e_t s$ into y_{t-S}, the VAR model becomes the following moving average representation where the price of each market is a function of past innovations of other markets:

$$y_t = \sum_{S=0}^{\infty} B(S) e_{t-S} \tag{9}$$

Since e_t is serially uncorrelated, the components of e_t may be contemporaneously correlated. To observe the structure of the response of each market to a unit shock in another market within S periods, the error term is transformed by the triangular orthogonalisation procedure. Let $e = Vu$ where V is a lower triangle matrix and u is an orthogonalised innovation from e such that $Eee' = S$ and $VV' = S$ and the transformed innovation u_t has an identity covariance matrix. Equation (9) can then be re-written as:

$$y_t = \sum_{S=0}^{\infty} B(S)Vu_{t\ S} = \sum_{S=0}^{\infty} C(S)u_{t\ S} \tag{10}$$

where $C(S) = B(S)V$. The $i.j$th component of $C(S)$ represents the impulse response of the ith market in S periods to a shock of one standard error in the jth market. From the orthogonalised innovations, the forecast variance of each market can also be decomposed into portions accounted by shocks or innovations from other markets. The orthogonalisation generates the quantity $\sum_{S=0}^{T} C_{ij}^2(S)$, which is the proportion of forecast error variance of y_i due to innovations in y_j. This variance decomposition provides a measure of the overall relative importance of the markets in generating fluctuations in their own and other markets.

4. Empirical results

Table 1 provides the panel unit root tests comprising statistics for the LLC t, IPS W and Hadri homoskedastic and heteroskedastic Z-tests and corresponding p-values at price levels and first differences for the eleven European markets. The LLC t test statistic and p-value for the price level series are 1.2728 and 0.8985, respectively. This indicates that the sample evidence on the whole panel of eleven European markets does not provide sufficient evidence to reject $H_{0,LLC}$. This suggests that there is insufficient evidence to conclude that each individual price level series is stationary. The LLC t-test for the first-differenced price series on the whole panel produced a t-statistic of -234.4400 and a p-value of 0.0000, which concludes the rejection of $H_{0,\ LLC}$ at the five percent level of significance. The rejection of the null hypothesis indicates that each price differenced series is stationary.

With the IPS test at price levels across the eleven European markets, the IPS W-statistic of 2.1629 and p-value of 0.9847 show that the null hypothesis, $H_{0,IPS,}$ that all cross section units in the panel are non-stationary cannot be rejected. The IPS panel unit root test indicates that at the price level all eleven European markets are non-stationary. The first-differenced series across all eleven European markets gives a IPS W-statistic of -196.1210 and a p-value of

0.0000 thus rejecting the null, $H_{0,IPS}$ which concludes that at least one of the price-differenced series in the eleven European markets is stationary. Turning to the Hadri homoskedastic and heteroskedastic Z tests of the null hypothesis that all series in the panel are stationary; for the price level series, the null hypothesis is rejected with a homoskedastic Z statistic of 68.0786 and a p-value of 0.0000 and a heteroskedastic Z statistic of 53.8621 and a p-value of 0.0000. This suggests that the price level series for all European markets tend to be non-stationary. With respect to the first-differenced series, the Hadri homoskedastic Z-statistic of 0.2629 and p-value of 0.3963 and the heteroskedastic Z-tests of 0.3778 and p-value of 0.3528 fail to reject the required null, thus indicating that all price differenced series are stationary.

According to the panel unit root tests, analysis of the price level series indicates non-stationarity while the first-differenced price forms exhibit stationarity for all eleven European markets. The finding of non-stationarity in levels and stationarity in differences suggests that each index price series is integrated of order I(1). The finding of non-stationarity in levels and stationarity in differences provides comparable evidence to other studies of European equity markets using less-powerful univariate unit root tests. In terms of subsequent modelling procedure, the differenced series are then used to carry out lag length selection, causality tests and decomposition of the forecast error variance for the markets to be analysed.

<TABLE 1 HERE>

Johansen cointegration tests are used in order to obtain the cointegration rank. The eigenvalues and trace test statistics are detailed in Table 2 for the various null and alternative hypotheses. As the multivariate cointegration tests cover all eleven markets rather than the simple bivariate combinations found in much of the earlier work, they consider the full scope of financial integration relationships that may be found. The trace test statistic is greater than the critical value for the null hypotheses of $r = 0$ thereby rejecting the null hypothesis. However, the null hypothesis of $r \leq 1$ fails to be rejected in favour $r > 1$ indicating the order of cointegration is 1. However, similar hypothesis are rejected up to, but not including, $r \leq 4$ thereby suggesting an order of integration of four. The primary finding obtained from the Johansen cointegration tests is that a stationary long-run relationship exists between the eleven European equity markets. Thus, there is a tendency for the eleven markets in the long run not to drift too far apart (or move together).

<TABLE 2 HERE>

Since cointegration exists, Granger causality tests are performed on the basis of equation (4). *F*-statistics are calculated to test the null hypothesis that the first index series does not Granger-cause the second, against the alternative hypothesis that the first index Granger-causes the second. The calculated statistics and *p*-values for the markets are found in Table 3. Among the eleven European markets fifty significant causal links are found (at the 0.10 level or lower). For example, as shown Greece, Ireland, Spain and the United Kingdom markets affect the Austrian market (column 1) and Spain (column 10) is found to have a Granger causal relationship with Austria, Belgium, France, Germany and Ireland.

Further insights are gained by examining the rows in Table 3 indicating the effects of a particular market on all markets. In the short-run it is evident that the most influential markets are Austria, Belgium, Germany, Ireland, Spain and the United Kingdom. Germany, for example, influences seven European markets, including France, Greece, Ireland, Italy, Netherlands, Spain and the United Kingdom. The least influential markets in terms of Granger-causality are the Netherlands, which has no influence across any other European markets, and Italy, which only affects Ireland. There is also an indication that there is feedback at play in several pairwise combinations. For example, the United Kingdom market Granger-causes the Irish market and Ireland Granger-causes the United Kingdom market. This suggests these markets have a common pricing factor and are thereby very closely integrated. Using the total number of causal and caused relationships as one indicator of integration, Austria, Ireland, Spain, Germany, the United Kingdom and Belgium are relatively more integrated, while Denmark, France, Greece, Italy and the Netherlands are less integrated.

<TABLE 3 HERE>

The long-run causality Wald test statistics and *p*-values based on Toda and Yamamoto's (1995) level VAR procedure are presented in Table 4. The model is estimated for the levels, such that a significant Wald test statistic indicates a long-term relationship. This serves to supplement the findings obtained from the Granger causality (short run) results in Table 3. Among the eleven markets, fifty-three significant causal links are found (at the 10 percent level or lower). For example, column 7 shows that the markets in Austria, Belgium, France Germany, Greece, Spain and the United Kingdom affect the Irish market; and the German market (column 5) is influenced by Belgium, Ireland, Spain and the United Kingdom. The rows in Table 4 indicate the effects of a particular market on all markets. The least influential market is Italy which does not have any long-run influence on any other European markets.

<TABLE 4 HERE>

However, these results should be interpreted with the qualification that short and long-run causality tests only indicate the most significant direct causal relationship. For example, it may be that some markets influence non-Granger caused markets indirectly through other markets. In order to address this concern, Table 5 presents the decomposition of the forecast error variance for 2-day, 5-day, 10-day and 15-day ahead horizons for the eleven equity markets. Each row indicates the percentage of forecast error variance explained by the market indicated in the first column. For example, at the 2-day horizon, the variance in the Austrian market explains 99.56 percent of its own innovations, whereas 0.13 percentage of the variance is explained by innovations in the German market and 0.12 percent by the Spanish market. Five European home markets, namely Austria , Denmark, Germany, Greece and Ireland explain at least 70 percent of their own innovations, while with the remaining markets domestic influences on innovation range from 21.57 (France) percent to 47.10 (Belgium) percent. The United Kingdom market significantly influences the German market by 19 percent, even after 15 days.

It is readily apparent from the decomposition of the forecast error variance in Table 5 that sizeable differences in the percentage of variance explained by domestic and international markets prevail across the European Union. In terms of their average influence on forecast error variance across other European markets at the 15-day horizon, Austria and Germany account for 16.4 percent and 19.3 percent, respectively, while Italy and Denmark account for a mere 0.1 percent and 0.2 percent respectively. From a different perspective, Austria accounts for 98.7 percent of its own variance and Greece 87.0 percent, down to the Netherlands at just 19.0 percent and France with 22.1 percent.

5. Concluding Remarks

Financial integration is a long-standing policy goal in the European Union, potentially benefiting its many member-states and their citizens through more opportunities for risk sharing and diversification, the better allocation of capital across investment opportunities, and the potential for higher economic growth. The results of this study are just one indication of a more integrated European equity market, in both the euro area and beyond, signalling that national stock market returns in the European Union are increasingly driven by common (regional) news. This would provide prima facie evidence that institutional and regulatory change in the European Union implemented through a variety of policy mechanisms, along with the changing behaviour of investors and financiers at the market level, has been successful in promoting the desired objective.

Of course, this analysis does suffer a number of limitations, all of which provide possible directions for future research on European financial integration. First, while the equity market is clearly an important dimension of the financial system, along with the money, bond and banking markets, as well as market infrastructures, it is just one part. Ample evidence suggests that the degree of integration varies depending on the market segment, with financial integration usually more advanced in market segments other than equity. For example, it is generally recognised that since the money market lies closer to the single monetary policy in the euro area, it is relatively more integrated than the equity market. It would then be interesting to use similar techniques to those used in this paper to compare the level of integration in different market segments in the European Union.

Second, while there is ample allowance in this study for local and regional factors in pricing equity in Europe, there is no recognition of global factors. This makes it difficult to gauge the relative impact of global, regional and local factors in European equity markets, and thereby make a more complete assessment of financial integration. Finally, this study provides a broad assessment of financial integration for the entire period and across all markets. It therefore is unable to comment on the relative pace of integration over this period, the role of the various institutional and regulatory changes in this process, especially the introduction of the single currency, and the differential impacts on the member-states. By splitting the sample period into, say, a period before and after a major structural or institutional change, it may be possible to illustrate the impact of this change on financial integration.

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TABLE 1. *Panel unit root tests*

	Levels series		First-differenced series	
	Statistic	*p*-value	Statistic	*p*-value
Levin, Lin & Chu *t**	1.2728	0.8985	-234.4400	0.0000
Im, Pesaran and Shin *W*-statistic	2.1629	0.9847	-196.1210	0.0000
Hadri Homoskedastic *Z*-statistic	68.0786	0.0000	0.2629	0.3963
Hari Heteroskedastic *Z*-statistic	53.8621	0.0000	0.3778	0.3528

Notes: Period 2/7/1990–30/6/2006; hypotheses $H_{1,LLC}$: each series is stationary, $H_{1,IPS}$: at least one series is stationary, H_1 (Hadri homoskedastic and heteroskedastic Z-stat) each series is non-stationary**;** the lag orders are determined by the significance of the coefficient for the lagged terms; for the price levels series intercepts and trends are included; for the first differenced price series only intercepts are included.

TABLE 2. *Johansen cointegration tests*

H_0	H_1	Eigen-value	Trace test	Critical value
$r = 0$	$r > 0$	0.0346	**507.7293	310.8100
$r \leq 1$	$r > 1$	0.0216	**360.9169	263.4200
$r \leq 2$	$r > 2$	0.0194	**269.7053	222.2100
$r \leq 3$	$r > 3$	0.0134	**187.8190	182.8200
$r \leq 4$	$r > 4$	0.0085	131.4605	146.7600
$r \leq 5$	$r > 5$	0.0079	95.7882	114.9000
$r \leq 6$	$r > 6$	0.0046	62.8146	87.3100
$r \leq 7$	$r > 7$	0.0037	43.4112	62.9900
$r \leq 8$	$r > 9$	0.0034	27.8572	42.4400
Accepted			4	

Notes: Period 2/7/1990–30/6/2006; 0.05 percent level critical values from Osterwald-Lenum (1992); the optimal lag order of each VAR model selected using LR tests for the significance of the coefficient for maximum lags and Schwarz's Bayesian Information Criterion; in each cointegrating equation, the intercept and trend are included.

TABLE 3. *Granger (short-run) causality tests*

Market	AUS	BEL	DEN	FRA	GER	GRE	IRL	ITL	NTH	SPN	UNK	Causes
AUS	-	2.8775	0.0403	6.9258	6.4687	3.4153	16.2037	3.2945	3.6133	4.6302	6.5114	9
		0.0899	0.8410	**0.0085**	**0.0110**	**0.0647**	**0.0001**	**0.0696**	**0.0574**	**0.0315**	**0.0108**	
BEL	0.0011	-	0.6480	6.4561	3.6017	0.0979	1.3436	4.6061	14.1388	2.7154	12.0870	6
	0.9735		0.4209	**0.0111**	**0.0578**	0.7544	0.2465	**0.0319**	**0.0002**	**0.0995**	**0.0005**	
DEN	0.5668	4.0443	-	1.5453	0.0546	0.0947	2.1862	1.5336	3.9339	0.6663	1.3032	2
	0.4516	**0.0444**		0.2139	0.8152	0.7584	0.1393	0.2156	**0.0474**	0.4144	0.2537	
FRA	0.2535	1.6394	0.0318	-	0.4223	0.0365	0.9811	0.7732	0.3647	10.1735	2.0033	1
	0.6147	0.2005	0.8584		0.5158	0.8485	0.3220	0.3793	0.5459	**0.0014**	0.1570	
GER	0.1508	0.8510	2.1828	27.6069	-	5.4352	23.8636	2.7274	8.0321	13.4407	3.6336	7
	0.6978	0.3563	0.1396	**0.0000**		**0.0198**	**0.0000**	**0.0987**	**0.0046**	**0.0002**	**0.0567**	
GRE	9.5215	0.5931	4.3850	1.9003	0.8514	-	3.3586	3.0319	1.8592	2.3358	1.4820	4
	0.0020	0.4413	**0.0363**	0.1681	0.3562		**0.0669**	**0.0817**	0.1728	0.1265	0.2235	
IRL	8.8577	24.0211	0.9978	36.2855	25.2748	0.2197	-	15.1257	35.7305	11.6243	15.0112	8
	0.0029	**0.0000**	0.3179	**0.0000**	**0.0000**	0.6393		**0.0001**	**0.0000**	**0.0007**	**0.0001**	
ITL	0.0677	0.4723	0.0139	0.0000	2.1370	0.5665	2.9701	-	0.0186	0.0260	0.2927	1
	0.7948	0.4920	0.9062	1.0000	0.1439	0.4517	**0.0849**		0.8916	0.8718	0.5885	
NTH	1.8930	0.8316	0.0090	0.6268	0.0000	0.2124	2.5590	1.0949	-	0.3596	1.9657	0
	0.1689	0.3619	0.9243	0.4286	1.0000	0.6449	0.1097	0.2954		0.5488	0.1610	
SPN	5.5212	12.5830	0.7121	0.8630	15.9029	1.6156	10.1582	3.1065	5.3337	-	8.4140	7
	0.0188	**0.0004**	0.3988	0.3530	**0.0001**	0.2038	**0.0014**	**0.0781**	**0.0210**		**0.0037**	
UNK	6.2849	0.1540	2.9987	2.9360	2.5961	0.8148	16.0253	1.1946	2.8613	0.0196	-	5
	0.0122	0.6948	**0.0834**	**0.0867**	0.1072	0.3668	**0.0001**	0.2745	**0.0908**	0.8887		
Caused	4	4	2	5	4	2	6	6	7	5	5	50

Notes: Granger causality tests conducted by adjusting the long-term cointegrating relationship by the ECM; The figures in the second row for each market are *p*-values; tests indicate Granger causality by row to column and Granger caused by column to row. For example, in the period 2/7/1990–30/6/2006 Denmark (row) Granger causes two markets (Belgium and Netherlands) and is Granger-caused by Greece and the United Kingdom. Significant values ($p \leq 0.10$) are in bold.

TABLE 4. *Long-run causality tests by level-VAR*

Market	AUS	BEL	DEN	FRA	GER	GRE	IRL	ITL	NTH	SPN	UNK	Causes
AUS	-	19.1517	51.2433	11.8869	9.1537	12.5420	19.8368	4.4451	7.3285	14.7364	14.5401	7
		0.0018	**0.0000**	**0.0364**	0.1031	**0.0281**	**0.0013**	0.4873	0.1973	**0.0116**	**0.0125**	
BEL	9.3728	-	15.8705	21.8089	26.5413	1.1245	17.6701	16.7832	24.3238	18.2103	31.2184	9
	0.0951		**0.0072**	**0.0006**	**0.0001**	0.9519	**0.0034**	**0.0049**	**0.0002**	**0.0027**	**0.0000**	
DEN	4.1465	5.2530	-	3.5774	2.6979	45.9024	4.2056	3.0735	4.1642	8.3081	6.2007	1
	0.5285	0.3858		0.6117	0.7464	**0.0000**	0.5202	0.6887	0.5260	0.1401	0.2872	
FRA	3.6743	10.4406	20.6097	-	5.1518	7.6468	9.5436	2.4328	4.2439	14.3631	10.1284	5
	0.5972	**0.0637**	**0.0010**		0.3976	0.1768	**0.0892**	0.7866	0.5149	**0.0135**	**0.0717**	
GER	7.3379	4.9341	12.8234	33.3345	-	8.3184	29.1488	4.4652	14.2813	21.8116	9.5473	6
	0.1967	0.4240	**0.0251**	**0.0000**		0.1395	**0.0000**	0.4846	**0.0139**	**0.0006**	**0.0891**	
GRE	15.4974	5.7350	20.2962	6.5946	2.3835	-	16.2591	8.7250	6.6917	6.8899	11.3059	4
	0.0084	0.3329	**0.0011**	0.2526	0.7939		**0.0061**	0.1205	0.2446	0.2290	**0.0456**	
IRL	7.1180	23.5513	1.4377	28.8990	22.2911	1.3599	-	11.4911	27.2761	7.6627	11.9725	6
	0.2120	**0.0003**	0.9201	**0.0000**	**0.0005**	0.9286		**0.0425**	**0.0001**	0.1758	**0.0352**	
ITL	2.7212	6.5619	3.9878	3.6668	3.9126	2.2692	4.3567	-	1.7325	6.1276	2.3038	0
	0.7429	0.2553	0.5512	0.5983	0.5621	0.8108	0.4993		0.8848	0.2940	0.8057	
NTH	8.6514	2.8433	6.5932	15.2500	8.1486	5.0710	7.1534	18.0115	-	14.8676	16.6881	4
	0.1238	0.7241	0.2527	**0.0093**	0.1482	0.4073	0.2095	**0.0029**		**0.0109**	**0.0051**	
SPN	8.6394	21.6071	17.9425	4.2997	20.6386	7.8976	18.5474	6.0106	10.6800	-	10.7128	6
	0.1243	**0.0006**	**0.0030**	0.5071	**0.0009**	0.1620	**0.0023**	0.3052	**0.0581**		**0.0574**	
UNK	11.7893	5.2362	5.4203	7.6200	10.6451	8.8088	26.1819	8.2876	19.2408	17.2831	-	5
	0.0378	0.3877	0.3668	0.1785	**0.0589**	0.1169	**0.0001**	0.1411	**0.0017**	**0.0040**		
Caused	3	4	6	5	4	2	7	3	5	6	8	53

Notes: Unbracketed figures in table are Wald statistics for Granger causality tests. The figures in the second row for each market are *p*-values. The level VAR are estimated with a lag order of $p = k + d_{max}$; k is selected by the LR test and d_{max} is set to one. Tests indicate Granger causality by row to column and Granger caused by column to row, for example, Greece (row) Granger causes four markets (Austria, Denmark, Ireland and the United Kingdom) and is Granger-caused by Austria and Denmark. Significant values ($p \leq 0.10$) are in bold.

TABLE 5. *Generalised variance decomposition*

	Period	S.E.	AUS	BEL	DEN	FRA	GER	GRE	IRL	ITL	NTH	SPN	UNK
AUS	2	20.0880	99.5595	0.0013	0.0095	0.0086	0.1321	0.0647	0.0000	0.0001	0.0302	0.1189	0.0753
	5	33.8036	99.1699	0.1004	0.0705	0.0418	0.1177	0.0361	0.0058	0.0257	0.1581	0.2289	0.0450
	10	48.8391	98.9775	0.1086	0.0696	0.0874	0.1262	0.0195	0.0130	0.0566	0.3248	0.1468	0.0700
	15	60.0813	98.7583	0.0936	0.0487	0.0975	0.1042	0.0315	0.0340	0.0857	0.5043	0.1112	0.1311
BEL	2	17.8366	22.6932	47.1023	0.0520	0.0118	15.5033	0.0000	10.1451	0.0024	0.0013	4.4882	0.0005
	5	28.6220	24.6366	48.0857	0.0856	0.0484	14.5748	0.0144	7.3308	0.1184	0.0034	5.0777	0.0242
	10	39.7426	26.6880	48.3222	0.0613	0.0728	13.5932	0.0623	6.3547	0.1536	0.0075	4.5823	0.1022
	15	48.1194	27.6343	47.8301	0.0435	0.0608	13.1505	0.1362	6.3927	0.1682	0.0207	4.3927	0.1705
DEN	2	40.2585	12.1294	0.5716	71.2303	0.0202	5.1899	4.8856	4.5161	0.0001	0.0008	1.2579	0.1980
	5	56.8819	17.8701	1.1107	63.0992	0.0801	6.9170	3.3240	4.8091	0.0152	0.0167	2.4424	0.3155
	10	75.9165	22.9633	1.2556	56.8990	0.2998	7.5913	2.3912	4.7980	0.0258	0.0097	3.1725	0.5937
	15	90.1619	26.1291	1.1900	52.5848	0.5367	7.9126	1.8711	4.9298	0.0543	0.0074	3.9276	0.8567
FRA	2	18.3578	17.1558	1.0313	0.0551	21.5741	40.1790	0.0220	9.9158	0.0001	0.0213	7.5860	2.4596
	5	28.2541	17.8527	2.4086	0.0308	21.9749	39.4470	0.0167	8.2694	0.0090	0.2134	7.5285	2.2490
	10	38.9098	18.8260	2.7828	0.0689	22.3181	38.0890	0.0151	7.8762	0.0057	0.3078	7.0223	2.6883
	15	46.9661	19.1034	2.8448	0.1342	22.1148	37.4418	0.0300	7.9836	0.0073	0.4156	6.9029	3.0217
GER	2	20.6271	18.7082	0.0657	0.0001	0.0374	70.7748	0.0166	9.7285	0.0265	0.0006	0.5777	0.0639
	5	32.1831	19.8205	0.9961	0.0558	0.1434	68.6853	0.0079	9.1643	0.0130	0.0373	1.0296	0.0467
	10	44.7604	20.8904	1.4774	0.0454	0.2270	67.1756	0.0086	8.9750	0.0070	0.0271	0.9400	0.2265
	15	54.0921	21.3597	1.5789	0.0330	0.2339	66.2547	0.0178	9.0987	0.0068	0.0359	0.9026	0.4780
GRE	2	23.9226	1.6385	0.1598	0.0241	0.0119	0.6134	96.2677	1.1479	0.0017	0.0054	0.1061	0.0236
	5	31.9924	2.3513	0.2970	0.2883	0.0580	1.6563	92.7448	1.7768	0.0127	0.0842	0.6964	0.0344
	10	40.1188	3.0438	0.4112	0.4697	0.1631	2.6580	89.6822	2.1667	0.0103	0.2011	1.0818	0.1119
	15	46.1146	3.6664	0.4555	0.8033	0.2301	3.3239	87.0521	2.4544	0.0125	0.3609	1.3399	0.3012
IRL	2	4.7052	17.5366	0.0040	0.0403	0.0261	0.9498	0.0014	81.1328	0.0286	0.0350	0.1162	0.1293
	5	7.5075	18.3049	0.3197	0.1051	0.0238	1.8261	0.1225	78.2344	0.0423	0.1470	0.8141	0.0602
	10	10.4249	19.6873	0.6205	0.0831	0.0163	2.2423	0.3131	75.7245	0.0373	0.1552	0.9957	0.1249
	15	12.5869	20.3320	0.8066	0.0583	0.0373	2.4816	0.5256	74.1615	0.0370	0.1818	1.1207	0.2577
ITL	2	6.1646	15.9276	0.6078	0.0123	1.4475	25.1087	0.0357	7.9107	37.9288	0.0268	10.2631	0.7309
	5	9.7293	15.8916	1.3078	0.0130	1.6567	24.7790	0.0427	7.1566	37.6156	0.2536	10.6891	0.5943
	10	13.5418	16.4880	1.4931	0.0103	1.8405	25.0521	0.0481	6.9669	36.3984	0.3153	10.5212	0.8660
	15	16.3317	16.8517	1.5499	0.0190	1.8346	25.1134	0.0500	7.0777	35.2788	0.4001	10.7053	1.1195
NTH	2	26.5391	14.6505	3.7521	0.0281	1.1547	34.0869	0.0027	12.3273	0.1144	24.3618	5.9722	3.5493
	5	40.2887	16.1814	6.0187	0.0136	1.4508	33.4070	0.0476	11.2141	0.0857	21.8296	6.1476	3.6038
	10	55.3666	17.6121	6.6149	0.0138	1.6419	32.4464	0.1006	11.1563	0.0549	20.2956	5.3361	4.7275
	15	66.7739	18.1575	6.5459	0.0246	1.6670	31.9479	0.1427	11.7292	0.0385	19.0041	4.9152	5.8274
SPN	2	5.4916	17.9114	0.0047	0.0069	0.0904	27.2180	0.0280	8.7079	0.0005	0.0204	45.9985	0.0134
	5	8.6063	19.5398	0.4250	0.0359	0.0455	25.2443	0.0270	7.6967	0.0125	0.1913	46.7060	0.0762
	10	11.8549	21.4212	0.6990	0.0346	0.0304	24.4943	0.0152	7.4407	0.0071	0.3483	45.4355	0.0736
	15	14.2628	22.1228	0.7566	0.1217	0.0213	23.8746	0.0106	7.7528	0.0049	0.5688	44.5564	0.2095
UNK	2	13.2889	14.2175	0.8546	0.0045	0.0236	20.1771	0.0155	13.8305	0.0075	0.0304	4.4911	46.3478
	5	20.0731	15.8828	2.2430	0.0409	0.0465	19.7345	0.0476	13.3501	0.0394	0.2284	5.8464	42.5404
	10	27.3949	17.3660	2.6960	0.1415	0.0608	19.3454	0.0400	13.0956	0.0891	0.1845	6.0137	40.9676
	15	32.9124	17.8485	2.8848	0.2557	0.0458	18.9133	0.0287	13.6135	0.1441	0.1789	6.1025	39.9842

Notes: The ordering for the variance decomposition is based on the number of 'causes' in Table 3; the four rows for each market are in order of forecast periods of 2, 5, 10 and 15 days, respectively.