On the Stability of Equity Price Relations: Evidence from Australia, the U.S. and the European Union countries *

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Abstract

An empirical analysis of stock price linkages between the equity markets of the Australia, the U.K., Germany, France and the U.S. is presented, using a cointegrating VAR framework. The existence of one cointegrating vector is ascertained. Restrictions on the cointegrating vector suggest that all markets enter the equilibrium relation significantly and that the equity markets of the U.S., Germany and France appear to be fully integrated. Results from restrictions imposed on the loadings vector suggest that the U.S. and German equity markets are weakly exogenous, thus not affected by the equilibrium relation. An examination of the time–plots of the parameter estimates of the recursively computed restricted model based on Hansen and Johansen (1992) reveal substantial instabilities during two periods: the first at the end of 1990, enduring for approximately two months and the second at the end of 1995 lasting until the end of 1995.

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1. Introduction

Since the conception of Markowitz's (1952) portfolio diversification theory, studies investigating portfolio diversification benefits across international financial markets have grown rapidly. The vast number of research papers that examined the benefits of internationally diversified portfolios relied heavily upon the analysis of correlation structures of equity market returns (see for example Grubel (1968), Levy and Sarnat (1970), Grubel and Fadner (1971), Agmon (1972), McDonald (1973), Ripley (1973), Lessard (1974), Panton *et al.* (1976), and Hilliard (1979) among others). Mixed results were obtained, so that no general consensus on the effectiveness of internationally diversified portfolios could be reached.

However, several problems intrinsic to studies analysing correlation structures of equity returns were identified. The process of differencing removed all possible common trend components from the data, imposed too many unit roots onto the system as a whole and thus information about a possible long–run equilibrium relationship was generally lost (see Engle and Granger (1987), Baxter (1994) and Leachman and Francis (1995)). Eun and Resnick (1984) also raised the point that unconditional correlation matrices need not remain stable over the sample periods analysed.

The introduction of the concept of co–integrated variables by Granger (1981, 1983, 1986) and Engle and Granger (1987) has led to a re–emergence of studies investigating portfolio diversification benefits across national equity markets within a cointegration framework. Nevertheless, a few important flaws seem to be consistent throughout the literature.

Firstly, it appears that conclusions about the degree of equity market integration and thus the validity of portfolio diversification are made without consideration of the relative size of the coefficients of the cointegrating vector. Only a limited number of studies impose restrictions on the cointegrating space (commonly zero restrictions), while still fewer impose restrictions to test for equity price homogeneity or even attempt to interpret the coefficients of the cointegrating relation.[1]

Secondly, it is interesting to observe that structural breaks in the data set are not explicitly examined. Hitherto, it seems to have been common practice that prior knowledge of major events such as the October 1987 stock market crash, the Asian financial crisis in 1997 and the Russian crisis in 1998 were used to partition the data into sub–samples to investigate the stability structures over these sub–samples (recent exceptions are the studies by Fernández–Serrano and Sosvilla–Rivero (2001) and Leong and Felmingham (2003)). However, such an approach will only prove useful in determining structural changes that may have occurred when the break points are known *a priori*.

Lastly, it is also apparent that a substantial number of studies do not estimate the lag length of the vector autoregressive (VAR) model appropriately.[2] Consequently, various lag length specifications are used, with some studies' results being fairly robust to different lag lengths (see for

[1] Equity price homogeneity shall here refer to a one–to–one correspondence in the long–run equilibrium relation of stock prices.

[2] See also Hall (1991), Boswijk and Franses (1992) and Reimers (1992) who further investigate the robustness of results of the Johansen–Juselius ML procedure with different lag lengths.

example Jeon and Chiang (1991), Choudhry (1994), Masih and Masih (1999) and others) while other studies' outcomes vary remarkably (see for example Kasa (1992), Chung and Liu (1994), Richards (1995), among others).[3] Additionally, some studies do not specify the choice of deterministic components to be included in the Vector Error Correction Model (VECM) consistently (see for example Kasa (1992), Chan*, et al.* (1997) to name a few), although Johansen (1991) showed that the asymptotic distribution of the rank test statistic, and thereby the finding of cointegration, is sensitive to the choice of deterministic components included in the VECM.

Hitherto, research papers including the Australian equity market in inquiries of stock price linkages appear to be limited and portray rather controversial findings. Research studies utilising cointegration techniques such as by Blackman*, et al.* (1994), Allen and MacDonald (1995), Corhay*, et al.* (1995) and Kwan*, et al.* (1995) found, depending on the estimation period and the markets included, between one and two cointegrating vectors between Australia and a number of European and Asian equity markets, while analyses by Richards (1995) and Roca (2000) failed to do so. The overall consent, nonetheless, seems to be that global equity markets have become more integrated, suggesting that little gains can be reaped from diversifying a portfolio internationally.

In the majority of these inquiries, however, conclusions about the validity of the cointegrating relationship are drawn without imposing further restrictions on the cointegrating space. Weak exogeneity, the relative size of the coefficients of the cointegrating vector and the stability of the cointegrating relation over time are not considered at all.

The objective of the present study is to investigate the extent of equity market integration between the stock markets of Australia, the U.K., the U.S., Germany and France. The definition of equity market integration follows Kenen's (1976) view, where integration is defined in terms of stock price interdependence or co–movement. We formulate a cointegrated vector autoregressive (VAR) process *à la* Johansen (1988, 1991) to model the dynamic interaction pattern between stock returns, allowing for the possibility of a long–run equilibrium correction. Additionally, linear restrictions on the cointegrating space will be imposed to determine the statistical significance of all markets in the cointegrating relationship. Also, tests on long–run equity price homogeneity and weak exogeneity are conducted. The stability of the cointegrating relationship will be analysed by examination of the plots of the recursively estimated $\boldsymbol{\beta}$ constancy test, the coefficients of the cointegrating vector, the speed of adjustment coefficients and the non–zero eigenvalues, as advocated by Hansen and Johansen (1992, 1999). The short–run dynamics will be investigated by use of impulse responses and forecast variance decomposition within a standard VAR framework.

The remainder of the paper is structured as follows. A brief description of the data is outlined in Section 2. Section 3 gives a concise overview of the methodological approach taken in this study, while the empirical results of the research paper are presented in Section 4. Section 5 contains concluding remarks.

[3] Kasa's study in particular is very sensitive to the lag length specification and the frequency of data employed. When two and 15 lags with monthly frequencies are used, one cointegrating vector is found, while two and 10 lags with quarterly data find up to four cointegrating vectors and one common trend.

2. Data

We utilise logged closing price data from Morgan Stanley Capital International (MSCI) indices for the countries of the U.K., Germany, France, Australia and the U.S., denominated in U.S. Dollars, covering the period from 6 January 1988 to 5 September 2001. Two different frequencies of data are utilised. Weekly (Wednesday) closing prices are used to model the long–run relation between the series, while daily closing prices are employed to estimate the short–run dynamics. The rational for using two different frequencies is as follows: Weekly data avoids issues relating to "day of the week" effects, commonly a problem when dealing with daily data, which still appears to be unresolved empirically in the literature. Jaffe and Westerfield (1985) and Smirlock and Starks (1986) suggest that evidence for "day of the week" effect exists, while Koch and Koch (1993) found no support for it. Weekly data also avoids problems of trading hour inconsistencies and, as Bailey and Stulz (1990) have put it, "too much trading noise" in the data in general. Daily data, on the other hand, highlights the flow of short–run dynamics between the equity markets that would remain undisclosed if weekly data were used (see Eun and Shim (1989), Chowdhury (1994), Hassan and Naka (1996) and Choudhry (1997) and others) It thus seems reasonable to utilise both frequencies in such a way as to benefit from their respective advantages in applied analysis.

We have chosen 1988 as the starting point of our sample period to, on the one hand side, avoid the inclusion of such an extreme event as the October 1987 stock market crash and, one the other hand, because the October 1987 crash represented to many market participants a clear starting point of an increase in interdependence between global financial markets. In the words of Lin, *et al.* (1994: p. 508) "many traders in Tokyo recall that the day after Black Monday they sold stocks on information about the market crash in New York, probably without assessing exactly what fundamental links the New York price declines had on Tokyo stock prices".

The preference for the MSCI database arose from its ease of comparability, consistency in index calculation and avoidance of cross–listings of stock between the indices analysed. The broad based, general market indices provided by local stock exchanges are usually comprised of a different number of stocks, different firm sizes, and do not explicitly adjust for cross–listings of stock across financial markets.[4] Due to its consistent computation using a market capitalisation value weighted formula, and hence easy comparability, the MSCI database is highly regarded in the finance literature.[5]

Plots of the raw data series in levels and the empirical densities of the logged differenced series are displayed in Figure 1 and 2, respectively. Except for Australia, all series are showing strong trending behaviour, particularly after 1995. Towards the end of 1999 some erratic movements are

[4]If, for example, local equity market indices such as the FTSE100 and the All Ordinaries were utilised, the problem of the stocks of British–American Tobacco, Rio Tinto and Cable & Wireless being cross–listed on both stock exchanges would have been encountered. Additionally, equity market indices are not computed homogeneously. The Dow Jones, for example, consists of 30 stocks not weighted by their market capitalisation, while both the FTSE100 and the All Ordinaries are capitalisation–weighted indices.

[5]For a more detailed description of the index computation, see the MSCI index compilation manual, available at `www.msci.com/stdindex/index.html`.

visible, amplified particularly for the German series. Similar erratic behaviour seems to occur once again at the beginning of 2000, also seeming to affect the German series heavier than the others. At the beginning of 1990, there appears to be lasting upward shift in the German series, however, returning previous trend levels towards the end of 1990. The empirical densities of the rates of return series in Figure 2 appear to approximate normal distributions reasonably well, although normality is rejected for all of them. This is mainly due to longer tails, that is excess kurtosis. The densities in panels b), d) and e) show also slight negative skewness.

3. Methodological and Econometric Approach

We investigate the potential of portfolio diversification between the equity markets of Australia Germany, France the U.S. and the U.K. by adopting Kenen's (1976) definition of equity market integration in terms of interdependence or long run co–movement in stock prices. We formulate a vector autoregressive (VAR) process for this group of markets and test for the existence of an equilibrium long–run relation.[6] If such an equilibrium relation exists, following the framework put forward by Johansen and Juselius (1994) and Johansen (1995a), additional restrictions on the cointegrating space will be imposed to test for statistical significance and homogeneity of equity prices. The homogeneity restrictions that are impose are conceptually equivalent to those commonly employed in studies testing for Purchasing Power Parity (PPP) (see for example Corbae and Ouliaris (1988), Johansen and Juselius (1992), Kim (1990), Pesaran and Shin (1996) and others for applications), that is, restrictions determining the existence of a one–to–one cointegrating relation between the series. A portfolio will be deemed desirable in the long–run, and hence diversification gains will exist, if the prices of assets held across national markets are not homogenous, that is, do not form a one–to–one (Masih and Masih 1997).

The short–run dynamic interaction pattern between the markets will be investigated by means of response functions of each of the series to shocks originating in the others. Markets that are in general more sensitive to shocks coming from foreign markets tend to react heavily to these shocks. Excess sensitivity to these shocks will increase the volatility of the portfolio in the short–run and hence diminish the benefits from diversifying a portfolio. Each market's sensitivity to shocks in any of the foreign markets will be gauged by the size of its impulse response function. The degree of short–run exogeneity can be identified from the decomposition of the forecast error variance. Markets that account for most of their own forecast error variance are more exogenous and hence deemed desirable for diversification.

A general consensus has emerged in the existing literature that stock prices are characterised by a stochastic trend and are hence commonly referred to as non–stationary or integrated series. Depending on the extent of integration, statistical inference in hypothesis testing procedures are

[6] It is well accepted in the literature that the Johansen procedure performs superior to other techniques of cointegration, see for example Moore and Copeland (1994) and Gonzalo (1994) for an empirical examination of a number of different techniques.

invalid and are likely to yield "spurious" or nonsense results (see Yule (1926) and Granger and Newbold (1974)). Pitfalls in applications using differenced data series are also well known (see also Engle and Granger (1987) and Baillie and Bollerslev (1989) for an exposition) and thus do not provide a viable alternative. We adopt the cointegrating VAR approach to test for the existence of long–run equilibrium relations, thereby avoid the problems of using differenced data series.[7] Fundamentally, this approach entails estimating a VECM of a k–dimensional VAR process of lag order p, taking the form:

$$\Delta\mathbf{x}_t = \mathbf{\Pi}\mathbf{x}_{t-1} + \sum\nolimits_{i=1}^{p-1} \mathbf{\Gamma}_i \Delta\mathbf{x}_{t-i} + \mathbf{\Psi}\mathbf{D}_{t-1} + \boldsymbol{\epsilon}_t \tag{1}$$

where $\Delta\mathbf{x}_t$ is a differenced vector of variables, $\mathbf{D}_t$ is a vector of deterministic terms, $\mathbf{\Psi}$ is a matrix of coefficients of the deterministic terms, $\mathbf{\Gamma}_i$ is a matrix of coefficients on $\Delta\mathbf{x}_{t-i}$ representing the short–run dynamics, $\mathbf{\Pi}$ is a matrix whose rank (r) denotes the number of independent linear combinations that render non–stationary vector $\mathbf{x}_t$ stationary and $\boldsymbol{\epsilon}_t$ is a vector of $i.i.d.$ disturbance terms. The matrix $\mathbf{\Pi}$ is further decomposed into two matrices $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$, each having dimension $k \times r$, so that

$$\mathbf{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'. \tag{2}$$

The matrix of loadings $\boldsymbol{\alpha}$ measures the average speed of convergence to the long–run equilibrium $\boldsymbol{\beta}'\mathbf{x}_t$ and $\boldsymbol{\beta}$ is a matrix of coefficients that renders $\boldsymbol{\beta}'\mathbf{x}_t$ stationary, thereby ensuring that $\mathbf{x}_t$ converges to its long–run equilibrium. The number of stationary long–run solutions that exist in $\boldsymbol{\beta}$ is equal to r, while $k - r$ represents those components of $\boldsymbol{\beta}$ that do not become stationary by $\boldsymbol{\beta}'\mathbf{x}_t$, thereby continue to contain a unit root and are referred to as the common stochastic trends.

The short–run dynamics are examined through means of impulse response functions (IRF) and forecast error variance (VDC) decompositions. In general, a stationary VAR of the following form is utilised

$$\Delta\mathbf{x}_t = \boldsymbol{\mu} + \sum\nolimits_{i=1}^{p} \mathbf{A}_i \Delta\mathbf{x}_{t-i} + \boldsymbol{\epsilon}_t \tag{3}$$

which can be inverted into a moving average representation using Wold's (1938) theorem, yielding

$$\Delta\mathbf{x}_t = \sum\nolimits_{i=0}^{\infty} \mathbf{B}_i \boldsymbol{\upsilon}_{t-i}. \tag{4}$$

It is then common practice to apply a Cholesky decomposition by defining $\boldsymbol{\epsilon} = \mathbf{V}\boldsymbol{\upsilon}$ where $\mathbf{V}$ is a lower triangular matrix and $\boldsymbol{\upsilon}$ are orthogonalised innovations with an identity covariance matrix, giving

$$\Delta\mathbf{x}_t = \sum\nolimits_{i=0}^{\infty} \mathbf{B}_i \mathbf{V} \boldsymbol{\upsilon}_{t-i} \tag{5}$$

where $\mathbf{B}_i\mathbf{V}$ is the matrix of decomposed impulse responses of dimension $n \times n$. However, one drawback of using the Cholesky decomposition to obtain orthogonalised innovations is that the results can be quite sensitive to the ordering of the series in the VAR process, particularly when contemporaneous correlations between the residual series are large, that is, the residual covariance

[7] This approach is well documented in the literature. See Johansen (1995b) for a textbook treatment.

matrix $\boldsymbol{\Omega_\epsilon}$ is not diagonal. This is generally circumvented by imposing some theoretical restriction on the ordering of the series.[8] In the current context it proves difficult to justify one ordering over another. We therefore adopted the generalised approach to derive impulse responses and variance decompositions as proposed by Lee and Pesaran (1993) and Pesaran and Shin (1996, 1998) to avoid this problem.

4. Empirical Findings

We initially tested the individual series for unit roots, applying the standard Augmented Dickey Fuller (ADF) (see Dickey and Fuller (1979, 1981) and Said and Dickey (1984)) and Phillips–Perron (PP) (see Phillips (1987) and Phillips and Perron (1988)) frameworks, following the unit root testing sequence proposed by Dolado, *et al.* (1990). The results are not presented here to conserve space.[9] *A priori*, we expect all series to be integrated of order one. As anticipated, all five series contain a unit root, that is, become stationary after being differenced once.

The lag order of the VAR was determined using three standard testing procedures; the Aikake Information Criterion (AIC), the Schwartz–Bayesian Criterion (SBC) and a Likelihood Ratio (LR) test.[10] Four weekly lags seem sufficient to capture the dynamics of the VAR. The choice of deterministic components to be included in the estimation of the VECM was determined following the Pantula (1989) principle, as advocated by Johansen (1992b). Model 2, a model with an intercept in the cointegrating relation as the only deterministic component was found to be appropriate.[11] It seems imperative to stress, once more, the importance of the deterministic components in the determination of the cointegrating rank. A recent Monte Carlo study by Hjelm and Johansson (2002: p.3) found that "the 'Pantula principle' is heavily biased towards choosing model 3 when model 4 is the correct one", where model 4 refers to a model allowing for a linear time trend in the cointegrating relation. Although this does not seem to be of immediate relevance to the results of our study, as model 2 rather than model 3 was selected by the Pantula principle, in order to avoid any uncertainties with regards to the deterministic components included in the system and hence the cointegrating rank, we have estimated a model with a linear time trend as well. From a methodological standpoint, we embraced the view taken up by Doornik, *et al.* (1998: p.549), that "adopting a model that includes a trend in the cointegrating relation has low cost even when the DGP does not have one". However, not only was the coefficient estimate on the trend term

[8] It would seem reasonable to suggest that the U.S. is, from a theoretical point of view, the most exogenous market. However, it appears difficult to establish the ordering of the EU markets.

[9] The results are available from the authors upon request.

[10] The AIC reached its minimum at two weekly lags, while the SBC reached it at one weekly lag. The LR test, however, suggests the use of four weekly lags, with a calculated $\chi^2 = 65.4025$. Both estimations with one and two weekly lags lead to problems with serial correlation. Since it is common that both the AIC and the SBC tend to choose too short a lag specification, we re–estimated the model with four weekly lags, consistent with the LR test results, to avoid the residual problems encountered with the shorter lags.

[11] The `proc=rank` command in CATS was used to jointly test for the cointegrating rank and the choice of deterministic components to be included.

very small (–0.001), inferences with regards to the cointegration rank and the α and β coefficients were very similar as well. A likelihood ratio (LR) test restricting the trend term to zero yielded a p–value of 0.29 ($\chi^2(1) = 1.12$), thereby not rejecting the restriction. Model 2 was hence deemed appropriate and was used for the remainder of the analysis. The test results of the trace test are presented in Table 1 below.[12]

The results presented in Table 1 hint to the existence of at least one cointegrating vector, significant at the 1 % level. Depending on the choice of the level of significance, the presence of a second vector seems plausible. To avoid any uncertainties with regards to the cointegrating rank, the trace statistic was computed recursively to allow a visual inspection of the time path of the test statistic, displayed in Figure 3. The number of test statistics that are upward trending and above unity correspond to the cointegrating rank at the 10 % level of significance. As can be seen from Figure 3, only one of the trace statistics remains above unity over the full sample period, with the second one being above unity only from 1996 onwards. It appears to be reasonable, thus, to consider only one cointegrating vector in the remainder of the analysis.

Another rational for including only one vector in the analysis is that the speed of adjustment coefficients of the second vector for Australia and the U.K. are insignificant, yielding t–statistics of –0.665 and –0.061 respectively, suggesting that these markets are exogenous, while that of the U.S. is significant, with a t–statistic of –3.675, hinting to an endogenous U.S. equity market. This clearly seems to be inconsistent with any theoretical and empirical literature. Here we would also like to point out that it appears to be conceptually appealing to suppose that the five markets should correspond to only one common trend, i.e. contain four vectors, be pairwise cointegrated, given the nature of these five economies. However, in an earlier survey conducted by Buncic (2001), over thirty empirical studies were reviewed, only two of which found up to four (the study by Kasa (1992)) and up to six (the study by Leachman and Francis (1995)) cointegrating vectors, depending on the number of lags and sample periods considered. All other studies that were reviewed found at most two cointegrating relations. We have also performed pairwise cointegration tests on the series in a number of different ways, however, we fail to find more than at most two pairwise equilibrium relations, again depending on the level of significance chosen.[13]

The estimates of the unrestricted α and β coefficients of the model are displayed in Table 2. Zero restrictions on the coefficient estimates of α and β have been imposed automatically to determine the statistical significance and weak exogeneity properties of each of the individual markets in the system, following Johansen and Juselius (1990). Results of the restrictions are given as p–values in parentheses to each coefficient estimate. The test results in Table 2 suggest that all markets enter the cointegrating relationship significantly as all zero restrictions on the cointegrating vector are rejected. Zero restrictions on the loadings vector cannot be rejected for the markets of Germany and the U.S., suggesting that these two markets are weakly exogenous to the long–run relation,

[12] Only the results of the trace test statistic are presented here, as it is more robust an estimator when the data series are likely to suffer from excess kurtosis and/or skewness, common to financial data (Cheung and Lai 1993).

[13] These results are available from the authors.

that is, they only exert influence onto the long–run relation but are not influenced by it.[14] It is interesting to observe that the estimates of the β coefficients for France, the U.K. and the U.S. seem to be quite close in absolute magnitude, which opens the possibility to test for equity price homogeneity between these series.

Given this observation, we impose the following joint restrictions on the long–run relation: The cointegrating vector is restricted such that $\beta_{U.S.} = -1$ and $\beta_{France} = \beta_{U.K.} = 1$ and the speed of adjustment vector such that $\alpha_{U.S.} = \alpha_{Germany} = 0$, thereby jointly testing for equity price homogeneity between the markets of the U.S., the U.K. and France and for weak exogeneity of the U.S. and Germany. The LR test yields a $\chi^2(4) = 2.43$ with a p–value of 0.66, thus not rejecting the above restrictions.

The coefficient estimates of the restricted long–run cointegrating relation are reported in Table 3. For completeness, we have summarised the residual statistics of the partial model in Table 4. As with the empirical densities of the individual series, the lack of normality of the residuals appears to be largely driven by leptokurtosis and should therefore not pose a significant problem (Doornik, *et al.* 1998). Johansen and Juselius (1990: p.176) also noted that excess kurtosis "is probably less serious than a skewed distribution", and Johansen (1995b: p.29) pointed out further that the "asymptotic properties of the methods only depend on the *i.i.d.* assumption of the errors. Thus the normality assumption is not so serious for the conclusion". The U.K. series appears to contain some mild ARCH effects, depending on the level of significance considered. Nevertheless, Lee and Tse (1996) showed that the performance of the Johansen procedure is only affected marginally when conditional heteroskedasticity is present. We also checked the stability of the restricted model by recalculation of the largest roots in the model. The three largest moduli were 1.00, 1.00 and 0.8445, within the now restricted model with two exogenous variables corresponding to one stable equilibrium relation and two unit roots. The stochastic properties of the restricted model hence seem acceptable.

The coefficient estimates reported in Table 3 are instructive. Note that we have arbitrarily normalised on Australia to focus on its relation to the other four markets. Firstly, it is clear that the estimates of the α coefficients on the markets of the U.S. and Germany are not significantly different from zero, implying weak exogeneity of these markets in the system. This indicates that, although the U.S. and Germany enter the equilibrium cointegrating relation and contribute towards it, they are not determined within the system in the long–run relation, i.e., they seem to lead the other three series and do not respond to any shocks to the equilibrium relation formed by $\boldsymbol{\beta}'\mathbf{x}_t$. The Australian market appears to react to equilibrium shocks from both the German and the U.S. market with the expect sign and quite heavily. The interpretation in the sense of marginal effects of the U.K. and the French markets seems more difficult. Since they are also

[14]Engle, *et al.* (1983) demonstrated that it is possible to conduct statistical inference conditional on weakly exogenous series, without any loss of relevant sample information. Conditioning on weakly exogenous variables can be very advantageous as a means of improving the stochastic properties of the estimated model. By conditioning on weakly exogenous variables, the remainder of the estimated model is more likely to be better behaved statistically (Johansen 1992a).

endogenously determined by the cointegrating relation, there must be some feedback effect in the system, hence giving the unexpected signs. It is interesting to find though that the validity of the restriction imposed on the cointegrating vector shows that the markets of the U.S., the U.K. and France appear to portray price homogeneity within the system, i.e., a stable one–to– relation. Again it may be dangerous to make marginal interpretations regarding this result.

A sensible extension of the above analysis leads to an investigation of the temporal stability of the parameter estimates of the long–run model, by recursive estimation of the restricted partial model.[15] The recursively estimated parameters of the restricted cointegrating relation are plotted in Figures 4–6 (the dashed lines are the 95 % error bands). From the time plots of the estimates of the cointegrating vector and the speed of adjustment coefficients in Figures 4 and 5, two main instabilities are visible. The first, occurring at the end of 1990 seems to be of a short–lived nature, only lasting for a period of approximately two months. The second instability, however, occurring at the end of 1993 appears to show more persistence, as both, the coefficients of the cointegrating vector as well as the speed of adjustment coefficients, do not return to stable levels until the end of 1995. In the plot of the eigenvalue corresponding to the cointegrating relation displayed in Figure 6 only the second instability is visible, indicated by noticeable downward shift. It is curious to note that the first temporary instability that affects the speed of adjustment as well as the cointegrating coefficients at the end of 1990 is not visible from the recursive eigenvalue estimates, while the second shift occurs much more immediately. We test for the stability of the cointegrating vector by following Hansen and Johansen's (1992, 1999) LR test computed from comparing the likelihood function from each recursive sub–sample (i.e. one weekly period) to the likelihood function obtained from the restriction that the cointegrating vector estimated from the full sample falls within the space spanned by the estimated vectors of each recursive sub–sample. A graphical representation of the recursive test is displayed in Figure 57, with unity indicating the 5 % level of significance. Any values above unity indicate a rejection of the constancy of the cointegrating vector hypothesis. Some initial instability in the test statistic is visible in the first part of Figure 7, corresponding possibly to the short–lived instabilities at the end of 1990. Values of the test statistic above unit at the end of 1993 until the end of 1995 indicates a rejection of the null hypothesis of stability of the cointegrating vector. Interpretations of the parameter estimates should thus be treated with caution.

It would be interesting to obtain parameters estimates of the cointegrating relation for periods before, during and after the more permanent structural change occurred. The appropriate periods of estimation are: 6 January 1988, 23 December 1993 and 22 Mai 1996. However, we encountered several problems in our empirical estimation when trying to estimate the above periods. A methodological problem with this approach appears to be that each period is rather short, thus it seems inconsistent conceptually to test for equilibrium long–run relations over such short time spans, particularly with volatile financial data. Nevertheless, since structural changes are clearly visible over the long–run relation, it is still of interest to determine what changes have occurred to the short–run dynamics of these markets. We hence estimated an unrestricted VAR for each of the

[15] The base period for the recursive estimation is from 3 August 1988 to 3 April 1990.

three periods using daily data. Table 5 displays the results of the generalised variance decomposition. Noteworthy changes seem to have occurred. All three European Union (EU) markets seem quite open in the first period, accounting for just over half of their own forecast error variance. The Australian market on the other hand is more exogenous, accounting for more then 75 % of its own forecast error variance. In the second period, less interaction between the EU markets appears to occur as each markets exogeneity increases. In the last period, however, a marked increase in interaction between the EU markets can be seen to occur, to levels higher than were experienced in period 1. The Australian market also experiences a decrease in independence.

The reaction of each of the series to shocks in any of the other markets over the three periods considered are displayed in Figures 8–10. The response path of the U.S. series is, as expected, rather weak, and was hence excluded from the graphs. Also, responses to shocks in the Australian market are rather low and thus of no particular interest. All series consistently portray an amplified response path in the last period, indicating that the short–run dynamics increase markedly after 1996.

5. Conclusion

The paper analysed the portfolio diversification properties between the equity markets of Australia the U.K., Germany, France, the and the U.S. by investigating the long–term financial integration properties within a cointegrating VAR framework. One long–run equilibrium relationship was found to exist. Restrictions on the cointegrating vector, consistent with the hypothesis testing framework proposed by Johansen and Juselius (1990, 1992) and Johansen (1992a, 1995a), indicate that firstly, all markets enter the equilibrium relation significantly, and secondly, the U.S., Germany and France appear to have equal sized coefficients, suggesting that these markets are determined by a one–to–one relation over the long–run. Additionally, restrictions on the loadings vector suggest that the markets of the U.S. and Germany are weakly exogenous to the system and thus do not respond to shocks to the equilibrium relation. We further investigate the stability properties of the non–zero eigenvalue, the coefficients of the loadings vector and the coefficients of the cointegrating vector of the restricted model, following Hansen and Johansen's (1992, 1999) outline. Two substantial instabilities can be identified from the time plots. The first one, which seems to be only of a very short–lived nature, occurs at the end of 1992. The second one, appearing to be a prolonged period of instability, commences at the end of 1993 and endures until the end of 1995.

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Figures

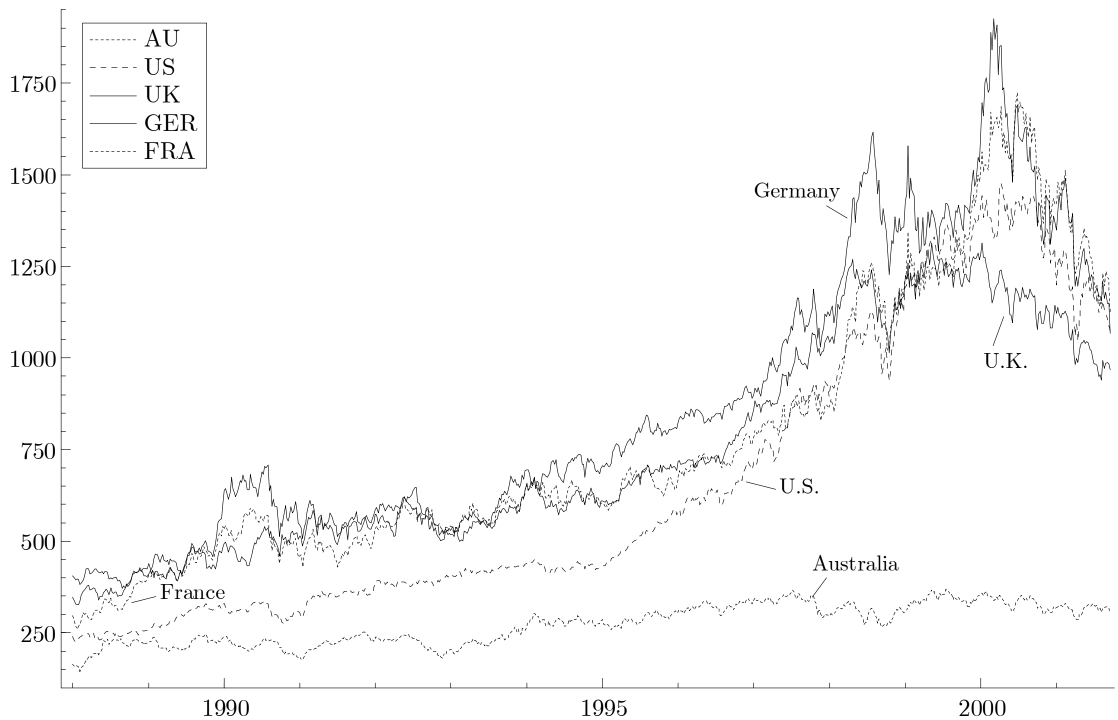

FIGURE 1: Time plots of the raw series.

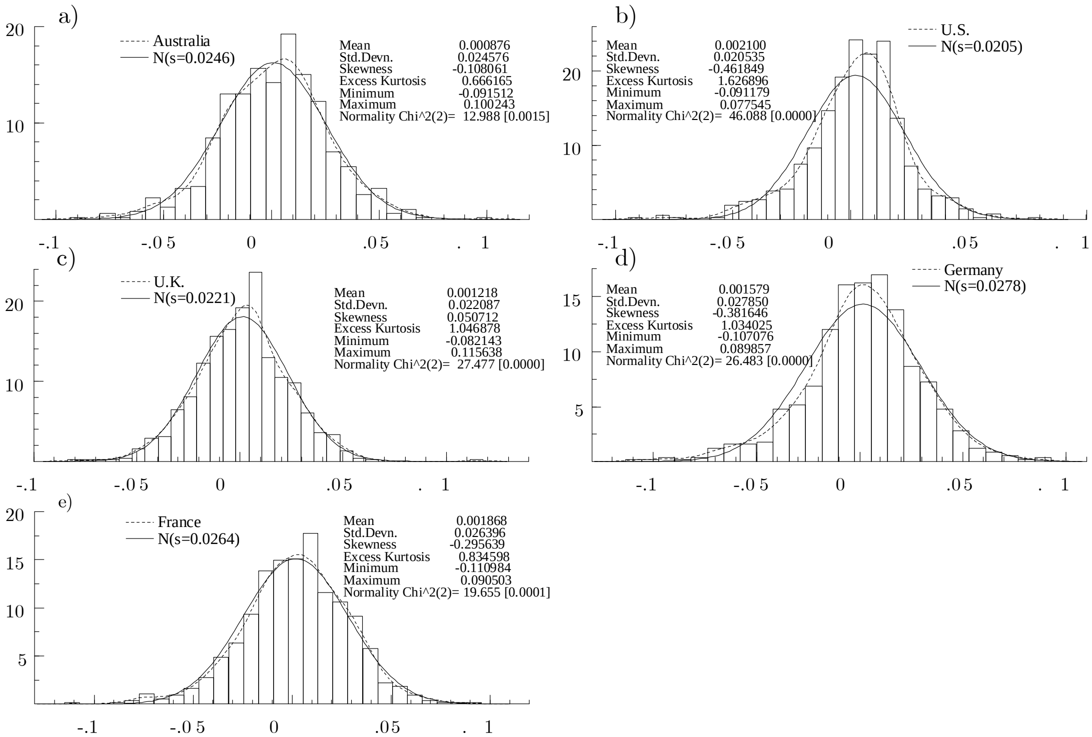

FIGURE 2: Empirical densities.

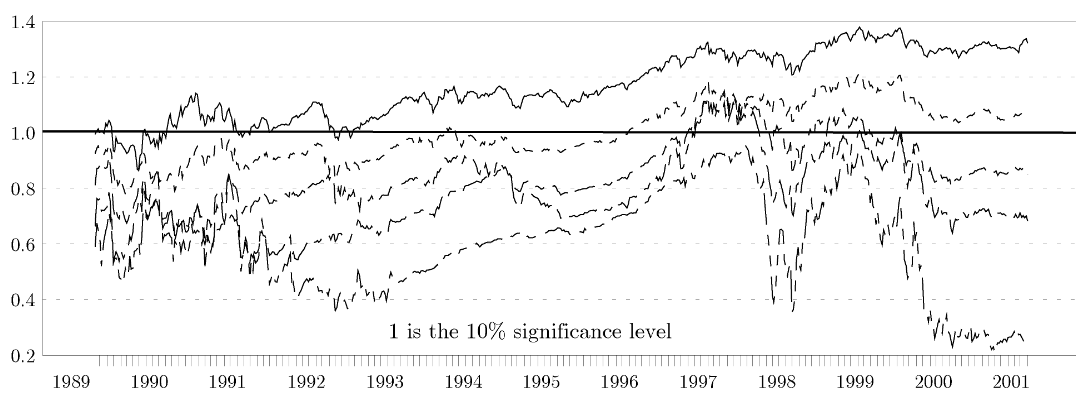

FIGURE 3: Recursively estimated Trace statistic.

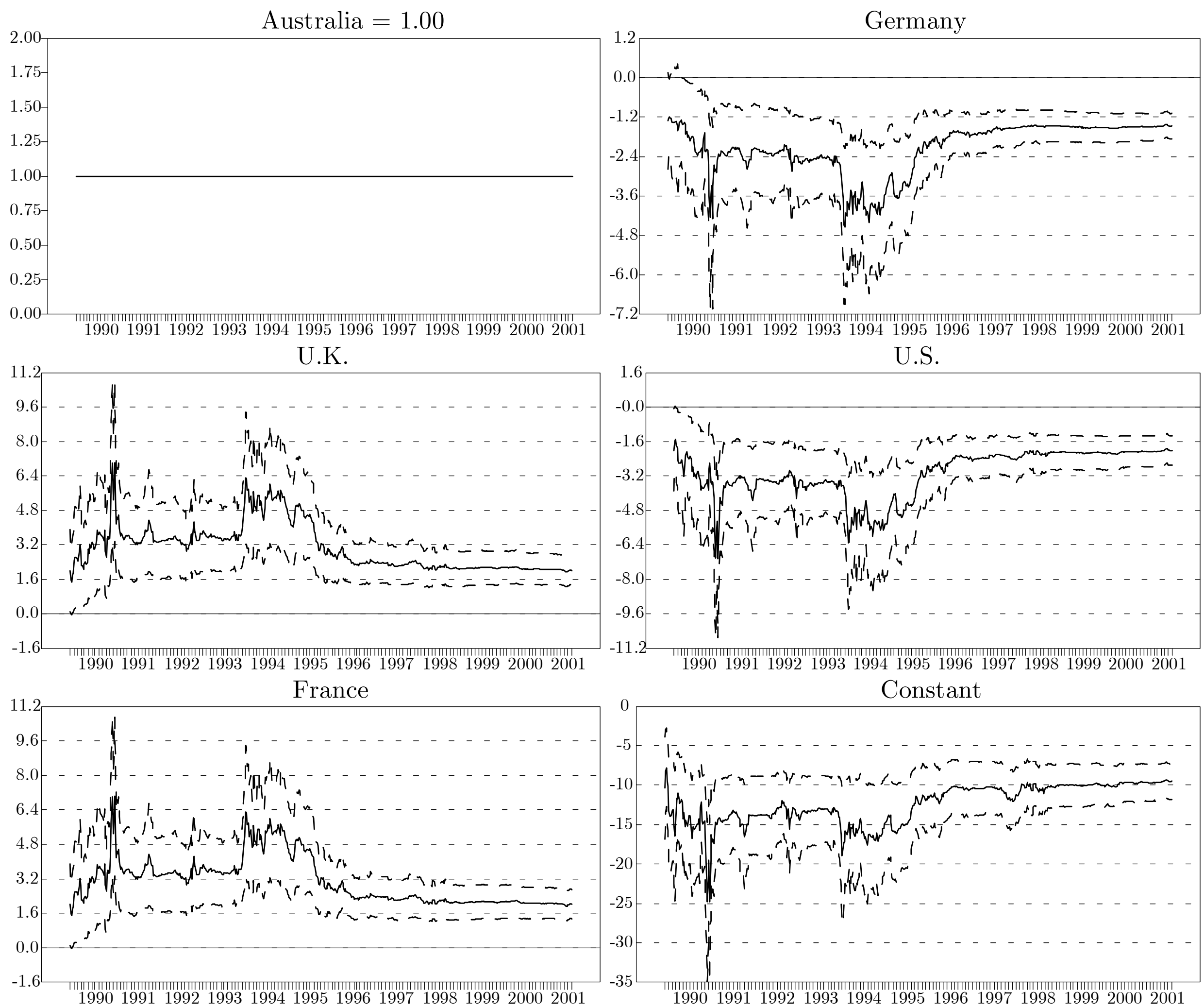

FIGURE 4: Recursive estimates of the coefficients of the cointegrating vector.

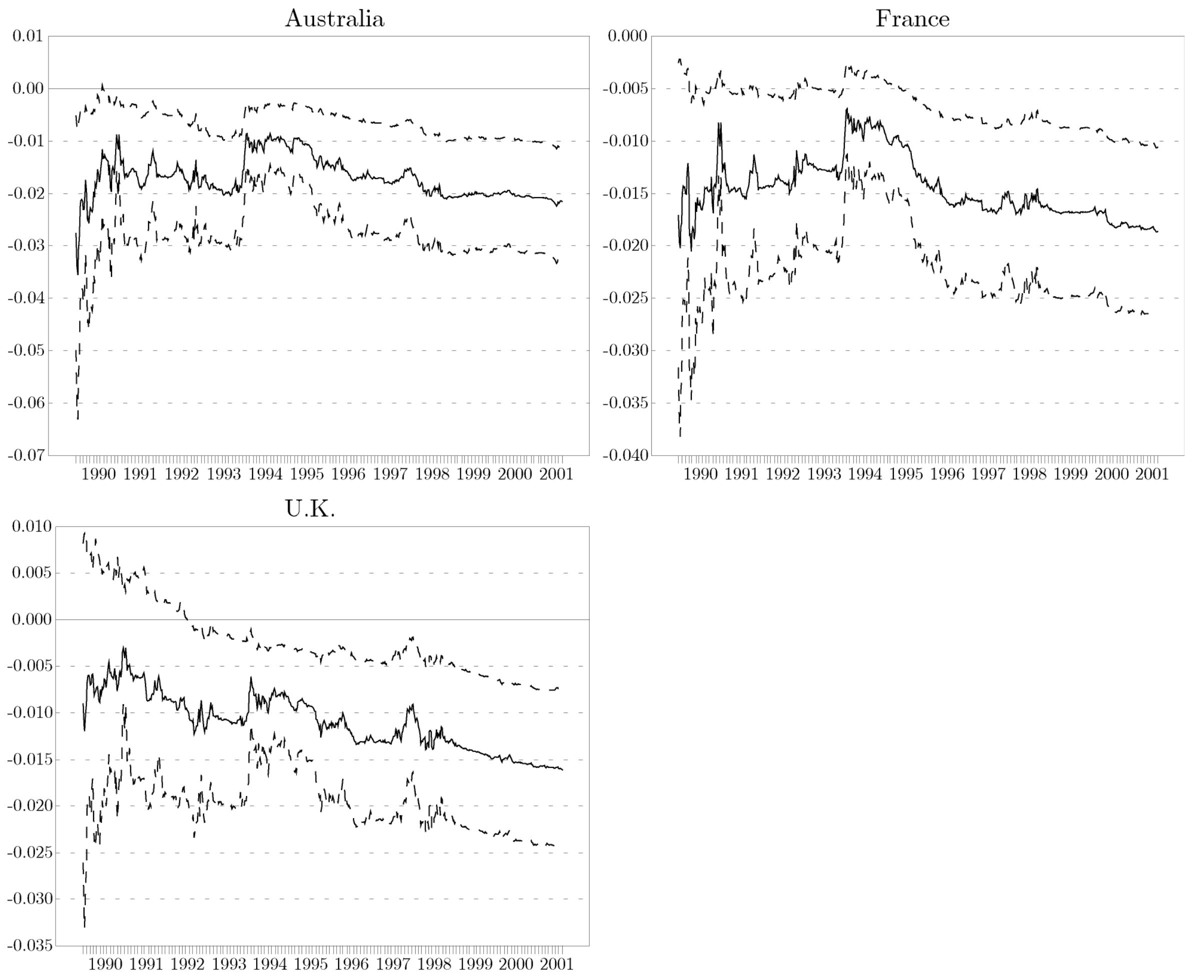

FIGURE 5: Recursive estimates of the speed of adjustment coefficients.

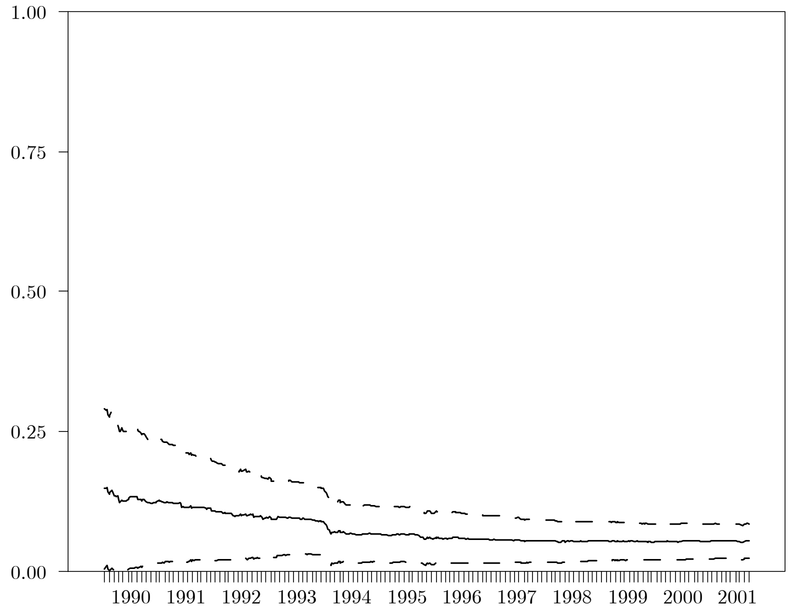

FIGURE 6: Estimate of the non-zero eigenvalue.

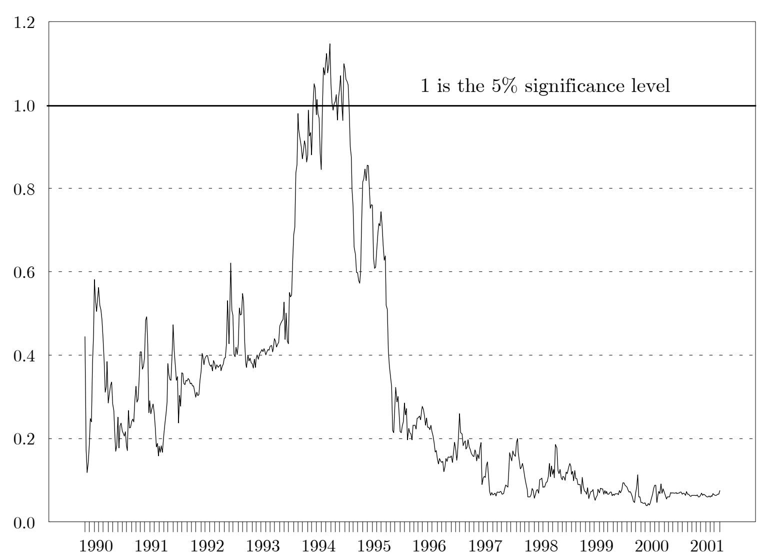

FIGURE 7: Test of know Beta equal to Beta(t)

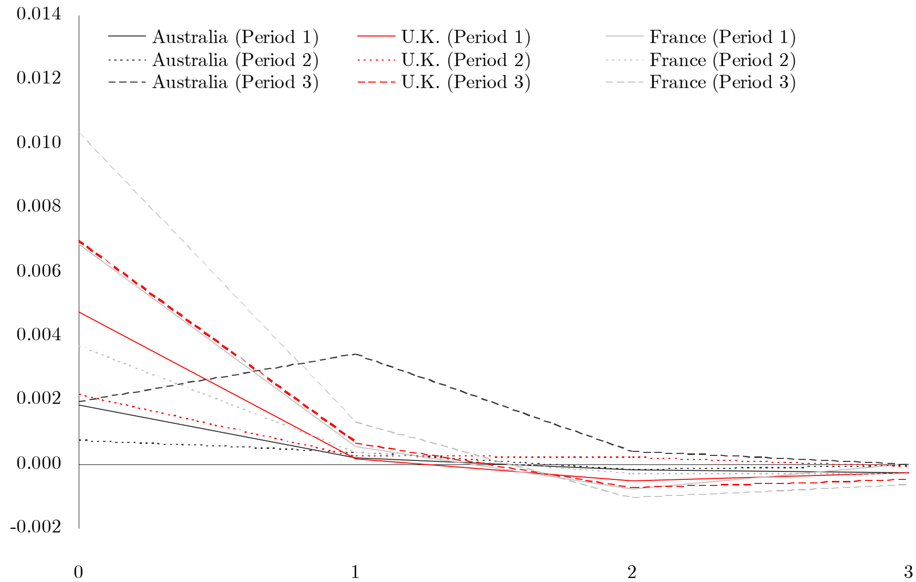

FIGURE 8: Responses to a shock in Germany over all three periods.

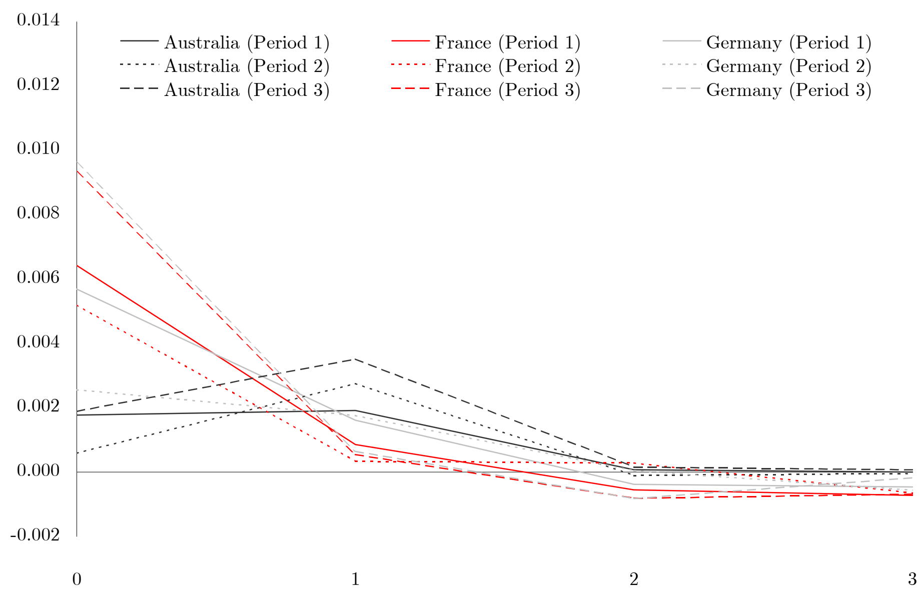

FIGURE 9: Responses to a shock in the U.K. over all three periods.

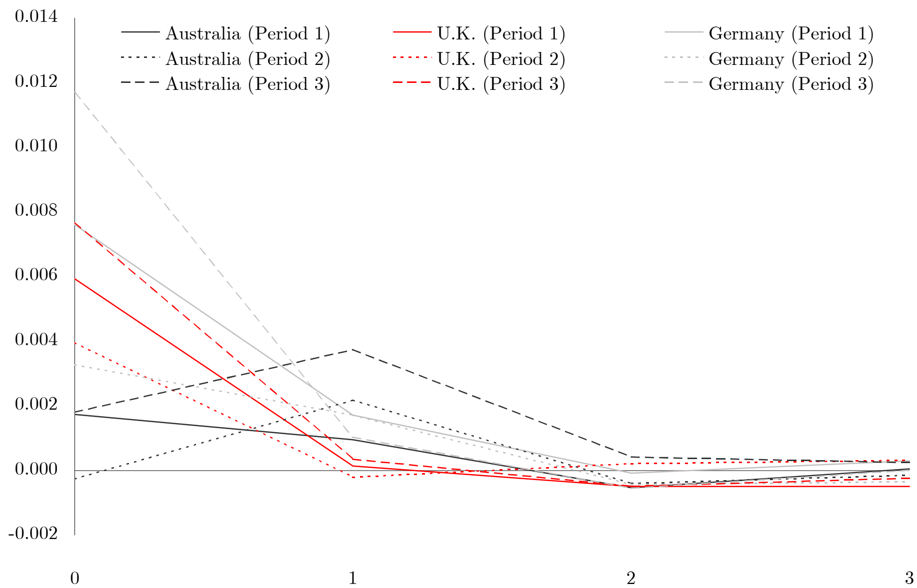

FIGURE 10: Responses to a shock in France over all three periods.

Tables

Table 1

Eigenvalues and Trace Test Results

$\mathcal{H}_0$	$k-r$	$\hat{\lambda}$	$\hat{\lambda}_{trace}$	Critical Values 10%	5%	1%
$r \leq 4$	1	0.0026	1.83	7.50	9.13	12.73
$r \leq 3$	2	0.0144	12.15	17.79	19.99	24.74
$r \leq 2$	3	0.0209	27.16	31.88	34.79	40.84
$r \leq 1$	4	0.0359	53.12[b]	49.91	53.42	60.42
$r = 0$	5	0.0569	94.68[a]	71.66	75.74	83.93

[a] Reject the null at the 1% level.

[b] Reject the null at the 10% level.

Table 2[a]

Coefficient Estimates of the Unrestricted Cointegrating Relation

Market	$\hat{\beta}$	(p–value)[b]	$\hat{\alpha}$	(p–value)[b]
Australia	1.000	(0.00)	–0.021	(0.00)
U.K.	2.350	(0.01)	–0.019	(0.00)
France	2.383	(0.00)	–0.024	(0.00)
Germany	–1.692	(0.00)	–0.010	(0.19)
U.S.	–2.318	(0.00)	–0.005	(0.44)
Constant	–10.708	(0.00)	–	–

[a] The cointegrating vector was normalised on Australia.

[b] The p–values are probability values obtained from restricting each $\hat{\alpha}_i$ and $\hat{\beta}_i$, $\forall i =$ U.S., U.K., France, Germany, Australia equal to zero, using a Likelihood Ratio (LR) test as outlined in Johansen and Juselius (1990).

Table 3[a]

Coefficient Estimates of the Restricted Partial Model[b]

Market	$\hat{\beta}$	$\hat{\alpha}$	(t–values)
Australia	1.000	–0.022	(–4.066)
U.K.	2.013	–0.016	(–3.892)
France	2.013	–0.019	(–4.618)
Germany	–1.464	–0.000	–
U.S.	–2.013	–0.000	–
Constant	–9.501	–	–

[a] The cointegrating vector was normalised on Australia.

[*] This model includes Germany and the U.S. as exogenous variables explicitly.

Table 4

Residual Statistics of the Restricted Model

Univariate Statistics

Market	**Skewness**	**Kurtosis**	R^2	**Normality**[c] (p–value)	**ARCH(4)**[d] (p–value)
Australia	–0.241	3.880	0.227	20.520 (0.0022)	6.854 (0.1438)
U.K.	0.258	5.552	0.420	106.360 (0.0000)	8.907 (0.0635)
France	–0.192	3.969	0.611	23.633 (0.0006)	5.943 (0.2034)

Multivariate Statistics

Normality[c] (p–value)	$\chi^2_{\mathrm{DF}=6} = 171.076$ (0.00)	**Serial Correlation** (p–value)	LB(177)[a] (0.24)	LM(1)[b] (0.89)	LM(4)[b] (0.20)

[a] Ljung and Box (1978) serial correlation test distributed as $\chi^2_{\mathrm{DF}=4345}$.

[b] Lagrange–Multiplier test for first and fourth order serial correlation distributed as $\chi^2_{\mathrm{DF}=9}$. (Godfrey 1988).

[c] Normality tests based on the Shenton and Bowman (1977) test extended to a multivariate framework distributed as $\chi^2_{\mathrm{DF}=6}$.

[d] Univariate Lagrange–Multiplier test for fourth order ARCH effects distributed as $\chi^2_{\mathrm{DF}=4}$ (Engle 1982).

Table 5

Generalised Variance Decomposition [a]

Period 1						
Forecast Variance		**Explained by innovations in**				
Days		**Australia**	**U.K.**	**Germany**	**France**	**U.S.**
1	**Australia**	**0.915**	0.028	0.030	0.027	0.001
5		**0.770**	0.050	0.026	0.031	0.123
1	**U.K.**	0.018	**0.609**	0.128	0.200	0.043
5		0.018	**0.598**	0.127	0.198	0.058
1	**Germany**	0.020	0.129	**0.610**	0.231	0.011
5		0.019	0.128	**0.557**	0.222	0.075
1	**France**	0.016	0.184	0.211	**0.558**	0.031
5		0.017	0.182	0.206	**0.540**	0.056
Period 2						
1	**Australia**	**0.986**	0.004	0.008	0.001	0.001
5		**0.708**	0.073	0.007	0.048	0.164
1	**U.K.**	0.003	**0.668**	0.064	0.205	0.060
5		0.004	**0.644**	0.063	0.200	0.089
1	**Germany**	0.006	0.076	**0.793**	0.124	0.000
5		0.007	0.093	**0.643**	0.131	0.126
1	**France**	0.001	0.204	0.104	**0.665**	0.026
5		0.004	0.202	0.103	**0.645**	0.045
Period 3						
1	**Australia**	**0.901**	0.029	0.033	0.027	0.010
5		**0.583**	0.085	0.086	0.093	0.153
1	**U.K.**	0.016	**0.483**	0.193	0.233	0.074
5		0.015	**0.460**	0.186	0.221	0.118
1	**Germany**	0.016	0.180	**0.450**	0.265	0.090
5		0.017	0.174	**0.432**	0.256	0.121
1	**France**	0.013	0.213	0.260	**0.441**	0.074
5		0.013	0.203	0.251	**0.416**	0.116

[a] The bold diagonal denotes the percentage of the forecast error variance that is explained by innovations accruing in each variables own market. The closer the value is to 1, the more exogenous a variable is considered.