IIIS Discussion Paper

No.25/June 2004

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May 2004

Keywords: Bond Indices, Cointegration, GARCH Models, Integration, Kalman Filter

JEL Categories: F36, G15

Abstract:

In this paper, we use a set of complementary techniques to examine the time-varying level of integration of European government bond markets. We consider daily bond returns and prices over the 1998-2003 period. Strong contemporaneous and dynamic linkages are found between individual European Union (EU) markets and the German market. However, there is no such evidence for the three accession markets of the Czech Republic, Hungary and Poland. The UK's market is also considered. In general, the degree of integration for the accession markets is weak and stable, with little evidence of further deepening despite the increased political integration.

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1. Introduction

The political, economic and monetary developments associated with the European Union (EU) have been major catalysts for regional financial market integration. As such, the next stage of EU enlargement will also have financial implications. Whilst there is substantial evidence of convergence in the present bond markets within the EU, less is known about the level and dynamics of financial integration between accession states and between existing member and accession states. In this study, we focus on integration between the government bond markets of three important accession countries, the Czech Republic, Hungary and Poland, the three largest economies of the accession contries as well as those that have the largest debt markets as well as a subset of countries that already belong to the European Union (EU): Belgium, France, Ireland, Italy, Netherland and Germany. Also the UK market is considered. The choice of countries is determined by data and economic factors. In regard to data the three accession countries chosen represent those that have the longest available series of data. In economic terms these countries represent the most developed economies, with the largest and most liquid government debt markets.

The concept of financial market integration is integral to international finance and it is intuitive that financial market integration changes with economic conditions. The economic explanation that is generally accepted is that the level of risk aversion changes and investors require time-varying compensation for accepting a risky payoff from financial assets. For this reason, recent studies have allowed integration to vary over time and with events, see for example, Aggarwal et al., (2003), Barr and Priestley (2003) and Bekaert and Harvey, (1995). Ilmanen (1995) provided one of the first assessments on time-varying expected bond returns using an asset pricing model. Christiansen (2003) has also recently used the AR-GARCH model of Bekaert, Harvey and Ng (2003) to assess volatility spillovers in European bond markets. She provides empirical evidence that regional effects have become dominant over both own country and global effects in EU markets with the introduction of the euro but not in non-EU countries. Given that Driessen, Melenberg and Nijman

(2003) find that factors relating to the term structure explain most of the variations in excess bond returns, it is conceivable that economic convergence required as part of EU membership has inevitably led to high levels of bond market convergence. However, this remains to be determined. To date there is little evidence of the extent, still less the dynamics, of European bond market integration.

The attention on co-movements across government bond markets pales in comparison with that on stock markets. Smith (2002) is one of the few studies to have tested for cointegration (long-term relationship) in international government bond markets. They apply the Johansen (1988) and Johansen and Juselius (1990) techniques on monthly mixed maturity (greater than one year) bond index prices and detect the presence of cointegrating vectors. In addition, they find mixed evidence on seasonality in government bond markets. However, the literature is silent on the time-varying nature of the European bond market integration in terms of returns, variances and covariances. This paper aims to address this void and provide empirical evidence of the nature of the bond market integration amongst the existing EU members and the three accession countries. Given that yield differences between government bonds in the EU are small (through monetary policy co-ordination), we expect EU bond markets to be more closely integrated overall than new incoming members, and we aim to investigate the extent to which these accession countries bond markets differ in behaviour from the existing markets. This information is vital to gauge the success of the EU's new phase of accession beginning in May 2004. Barr and Priestley (2003) believe that the economic costs and benefits of international bond market integration are likely to be significant, ultimately leading to a lower cost of fiscal funding for governments.

The major finding of this paper are i) although there are strong linkages between individual EU country bond markets and Germany's market, the three accession countries's linkages are relatively weak ii) the UK market's linkage with Germany is relatively weaker than the other EU countries, iii) the accession countries' linkages to the EU do not appear to be growing over time, and iv) the rate of integration appears to be stable with little evidence of increased or decreased integration.

The remainder of this paper is organized as follows: Section 2 discusses the bond index data used in the paper; Section 3 detains the three empirical methodologies employed; the estimation results are discussed in section 4; and finally section 5 concludes.

2. Data description

The data are all-maturity total returns on Morgan Stanley Corporate Index (MSCI) government bond Indices for Belgium, Czechoslovakia, France, Germany, Hungary, Ireland, Italy, Netherlands, Poland, and the United Kingdom. They are sourced at Thompson Datastream. We have chosen these bond indices on the basis that they are available at a daily frequency for the longest time period for the three sample Eastern European countries: Czechoslovakia, Hungary and Poland. The bond indices are denominated in US dollars. We used the data available from 30 June 1998 to 31 December 2003 (rendering 1435 usable observations) to calculate returns as first log differences.[1] In our analyses, we choose to use the German government bond index as the proxy for the entire EU bond market given there is a correlation of 0.995 with the EU-11 tracker government bond index from Datastream International.[2] This will avoid the prospect of spurious integration as individual bond markets will not be a composite of the regional proxy index. We have included data on the UK as well as Eurozone countries, as the three countries under investigation are not expected to adopt the Euro for a number of years. Thus, exclusion of the sterling debt market would be unwarranted.

In Table 1, we provide descriptive statistics on the bond returns. In general, bond returns are higher in the accession countries compared to the existing EU member countries and the UK, and this corresponds to generally higher return variances in these countries due to peceived higher level of credit, political and transfer risks. In addition, it is revealed that the distributions of these bond market returns are statistically non-normal (significant levels of skewness and excess kurtosis). The

[1] We follow the existing literature in applying log-changes of total return government bond indices (eg. Bodart and Reding, (1999), Christiansen, (2003) and Driessen et al., (2003)).

[2] This series is only available starting from the 1st January 1999.

three accession countries have larger (in magnitude) skewness than the rest. Interestingly, Hungary and Poland show significant negative skewness while the others show the opposite. Also, the excess kurtosis of these two countries is considerably larger than that of the other countries. The bond index returns are not serially correlated in the first moment in all cases except for those of Poland. However, significant correlation in the second moments is found in all three accession countries and the UK which is indicative of time-varying volatility in these markets. In addition, the significance of the tests for white noise for each bond market and the German anchor indicates that the first and second moments of all these series move closely together and that the bivariate nature of these distributions need to be accommodated in the modeling of these daily bond market returns.[3] Lastly, Engle-Ng's (1993) sign bias test indicates the existence of asymmetric time-varying volatilities in particularly the Hungarian, Polish and UK government bond markets. As in Cappiello et al. (2003), we find little evidence of asymmetries in government bond index return volatility in the existing EU countries.

3. Dynamic Methodologies

We have noted that it is desirable to consider the dynamic and time varying nature of integration. The techniques we use here expressly allow us to focus in on this element.

3.1 Dynamic Cointegration

The essence of cointegration is that the series cannot diverge arbitrarily far from each other, implying that there exists a long-term relationship between these series and that they can be written in an Error Correction format. By definition, cointegrated markets thus exhibit common stochastic trends. This, in turn, limits the amount of independent variation between these markets. Hence, from the investors' standpoint, markets that are cointegrated will present limited diversification

[3] A bivariate version of the Ljung Box (portmanteau) Q test for serial correlation devised by Hosking (1980) was used on linear and squared market returns.

opportunities. The requirement for assets that are integrated in an economic sense to share common stochastic factors, which is an alternative definition of cointegration, is pointed out in Chen and Knez (1995).

Two primary methods exist to examine the degree of cointegration among indices. As this area is by now well known we do not provide a detailed statistical description of these techniques. For such a description see, for example, Enders (1995). The first is the Engle-Granger methodology (see Engle and Granger (1987)) which is bivariate, testing for cointegration between pairs of indices. The second is the Johansen-Juselius technique (see Johansen (1988) and Johansen and Juselius (1990)), hereafter referred to as the JJ technique which is a multivariate extension and allows for more than one cointegrating vector or common stochastic trend to be present in the data. The advantage of this is that the JJ approach allows testing for the number as well as the existence of these common stochastic trends. In essence, the JJ approach involves determination of the rank of a matrix of cointegrating vectors.

To illustrate, for a given lag length l, and assuming no deterministic components[4], we can write the Vector Autoregression (VAR) representation of the stock indices in levels as

$$\mathbf{E}_t = \mathbf{A}_1\mathbf{E}_{t-1} + \mathbf{A}_2\mathbf{E}_{t-2} + \ldots.. + \mathbf{A}_l\mathbf{E}_{t-l+} + m_t \qquad (1)$$

where $m_t \approx N(0,\Sigma)$ and E represents an $(nx1)$ vector of stock equity indices, **A** is an (nxn) matrix of coefficients. We can represent this relationship more generally in the Vector Error Correction (VECM) format as

$$\Delta\mathbf{E}_t = \prod\mathbf{E}_{t-1} + \Gamma_1\Delta\mathbf{E}_{t-1} + \Gamma_2\Delta\mathbf{E}_{t-2} + \ldots.. + \Gamma_{l-1}\Delta\mathbf{E}_{t-l+1} + \Gamma_l\Delta\mathbf{E}_{t-l} + m_t \qquad (2)$$

[4] The selection of the lag length is important, but more important again is the treatment of deterministic components. In the presence of deterministic elements the estimation of the VAR and the determination of the cointegration vectors, and thus the rank of the system, becomes complex.

Or

$$\prod \mathbf{E}_{t-1} = \Delta\mathbf{E}_{t-1} - \sum_{i=1}^{l} \Gamma_i \Delta\mathbf{E}_{t-i} - m_t \qquad (3)$$

Where the right hand side terms of Equation (3) are stationary, it follows that $\prod \mathbf{E}_{t-1}$ is also stationary. The JJ technique endeavors to ascertain the rank, r, of Π. This gives the number of stable cointegrating vectors in the system, as Π can be demonstrated to be equivalent to ab' where b' is the vector of cointegrating relationships and a a matrix associated with the equilibrium errors $b'\mathbf{E}_t$.[5]

The JJ approach generates two statistics of primary interest. The first is the λ_{trace} statistic, which (in this instance) is a test of the general question of whether there exist one or more cointegrating vectors. An alternative test statistic is the λ_{max} statistic, which allows testing of the precise number of cointegrating vectors. These test statistics can be plotted over time to examine how the nature of market integration is changing over time.[6] This approach is in essence a visual application of the recursive cointegration approach of Hansen and Johansen (1992) that has also been applied in a somewhat different form by Rangvid (2001). The output from the approach which we have taken is twofold: first, the largest value of the λ_{trace} statistic which tests the general hypothesis of no cointegration versus cointegration, and second, the number of cointegrating vectors given by the λ_{max} statistic. A set of series that are in the process of converging should be expected, as in Hansen and Johansen (1992) and Rangvid (2001), to show increasing numbers of cointegrating vectors. Intuitively, this makes sense. Consider a set of p series which have n cointegrating vectors, $n<p$. This implies that there are n linear combinations of the p vectors that are stationary. If we later find that we have k vectors, $n<k<p$, there are additional combinations that can be used in the representation of the p data. If we have a static number of cointegrating vectors then recursive

[5] Serletis and King (1997) used this approach to examine European equity market integration, the BENELUX and France in particular were found to be converging to the US market.
[6] Further details regarding the dynamic cointegration approach can be found in Barari and Sengupta (2002). There -in the process is described whereby the investigator can plot over time the values of selected test statistics from the JJ approach. It concentrates on the λ_{trace} statistic.

estimation will simply lead to an upward trend in the λ_{trace} statistic. It should be noted that in general the λ_{trace} statistic is more powerful and to be preferred to the λ_{max} statistic.

3.3 Haldane and Hall

There are a variety of feasible alternative approaches to the Cointegration methodology. The Haldane and Hall (1991) Kalman Filter based methodology is one that has been used in a number of setting.[7] The Haldane & Hall (hereafter HH) method estimates a simple equation of the following specification

$$\ln\left(\mathbf{E}_{jt}/\mathbf{E}_{Bt}\right) = a + b_t \ln\left(\mathbf{E}_{jt}/\mathbf{E}_{Xt}\right) + e_{jt} \qquad (4)$$

via kalman filter estimation. Here the market subscripted B is the preimposed internal base market and that subscripted X is the preimposed external market. Thus, for example, in testing for integration among South-East (SE) Asian markets, Manning (2002) imposes the US market as the external market (to which the SE Asian markets are assumed to be converging) and Hong Kong as the dominant local market. Here we set the Frankfurt market as the local base and the London market as the external market, and estimate the system. We also invert these relationships, as we are not confident as to which market, over the time period of this study, represents the dominant market towards which the system may be converging. Negative values of b_t indicate divergence, as does a tendency to move further from zero. Markets which are fully converged will show a zero b_t.

The Kalman filter used in this paper works in the following way. The equation is estimated over an initial period, to initialize the coefficients and related information. Thereafter it is updated

[7] Manning (2002) examines Asian stock market integration taking the Haldane and Hall (1991) approach of specifying time varying coefficients via a Kalman filter. Most papers using this time varying approach have examined currency or interest rate relationships (e.g., Zhou, 2003).

with the addition of each daily data point. Let $Y_t = a_t + X_t b_t + e_t, \mathrm{var}(e_t) = h_t$ be the measurement equation of interest. If we set b_t as the coefficient of interest at time t, then the transition equation is given by $b_t = b_{t-1} + n_t, \mathrm{var}(n_t) = \mathrm{M}_t$. Given the estimate of b_{t-1} from information up to that period ($b_{t-1|t-1}$) with the associated covariance matrix Σ_{t-1}, the updated estimate is given by equations (5), (6) and (7).

$$S_t = \Sigma_{t-1} + \mathrm{M}_t \tag{5}$$

$$\Sigma_t = S_t - S_t X_t' (X_t S_t X_t' + h_t)^{-1} X_t S_t \tag{6}$$

$$b_{t|t} = b_{t-1|t-1} + S_t X_t' (X_t S_t X_t' + h_t)^{-1} (Y_t - a_{t-1} X_t b_{t-1|t-1}) \tag{7}$$

3.3 Time-varying Conditional Correlations

Whilst the use of conditional econometric models capable of capturing asymmetric volatility has proliferated in stock market studies, government bond markets have not been dealt with in the same way. However, as revealed by the Engle Ng sign bias test results reported in Table 1, asymmetric volatilities are present in two of the three accession bond markets. As the focus of this study is on the integration of EU accession markets, we utilize an alternative bivariate exponential GARCH framework (incorporating student *t* densities to cater for non-normal error distributions) to derive time-variations in conditional correlations to proxy the dynamic process of bond market integration.

The conditional first moments (means) of the index returns are estimated as a parsimonious restricted bivariate ARMA(*p*,*q*) process as shown in equations (8) to capture the dynamics between mean bond market returns for each individual country (subscripted *N*) and Germany (proxy for the EU, subscripted *E*).

$$R_{N,t} = a_{cN} + \sum_{i=1}^{p_E} a_{rE,i} R_{E,t-p_E} + \sum_{j=1}^{q_N} a_{mN,j} e_{N,t-q_N} + e_{N,t} \tag{8}$$

$$R_{E,t} = a_{cE} + \sum_{i=1}^{p_N} a_{rN,i} R_{N,t-p2} + \sum_{j=1}^{q_E} a_{mE,j} e_{E,t-q_E} + e_{E,t}$$

with

$$e_t = \begin{bmatrix} e_{N,t} \\ e_{E,t} \end{bmatrix} \sim t(0, H_t, d),\ H_t = \begin{bmatrix} h_{N,t} & h_{NE,t} \\ h_{EN,t} & h_{E,t} \end{bmatrix}$$

In essence, $R_{N,t}$ is the national bond index return that is a function of past returns in the German market and past idosyncratic shocks, $e_{N,t}$, and $R_{E,t}$ is the German bond index returns that is a function of past returns in country N and its own past shocks, $e_{E,t}$. Specifically, the regional and country mean spillover effects can be quantified by the sign and magnitude of the estimated coefficients for the lagged Germany and national returns respectively. Note that p_N and p_E are the number of autoregressive terms and q_N and q_E are the number of moving average terms needed to eliminate joint linear and nonlinear serial correlation in the standardized residuals, $\frac{e_{N,t}}{\sqrt{h_{N,t}}}$ and $\frac{e_{E,t}}{\sqrt{h_{E,t}}}$ which are jointly t distributed.

In equations (9), we incorporate volatility spillover effects in the conditional second moments (variances) in modeling joint bond market returns as we are interested in their cross-market volatility interdependencies and this has not been previously investigated for the accession government bond markets.

$$\ln h_{N,t} = b\,c_N + b_{hN} \ln h_{N,t-1} + \left[be_{N1} \frac{e_{N,t-1}}{\sqrt{h_{N,t-1}}} + be_{N2} \left(\frac{|e_{N,t-1}|}{\sqrt{h_{N,t-1}}} - \sqrt{\frac{2}{p}} \right) \right] + \left[b_{E1} \frac{e_{E,t-1}}{\sqrt{h_{E,t-1}}} + b_{E2} \left(\frac{|e_{E,t-1}|}{\sqrt{h_{E,t-1}}} - \sqrt{\frac{2}{p}} \right) \right] \quad (9)$$

$$\ln h_{E,t} = b\,c_E + b_{hE} \ln h_{E,t-1} + \left[be_{E1} \frac{e_{E,t-1}}{\sqrt{h_{E,t-1}}} + be_{E2} \left(\frac{|e_{E,t-1}|}{\sqrt{h_{E,t-1}}} - \sqrt{\frac{2}{p}} \right) \right] + \left[b_{N1} \frac{e_{N,t-1}}{\sqrt{h_{N,t-1}}} + b_{N2} \left(\frac{|e_{N,t-1}|}{\sqrt{h_{N,t-1}}} - \sqrt{\frac{2}{p}} \right) \right]$$

The conditional variance for each return series is determined by its own past variance, its own negative and positive past unanticipated shocks as well as those from the bond index return. In this context, the regional (Germany) and country volatility spillover effects can be measured by the magnitude of the estimated coefficients for the negative and positive lagged external innovations in the latter part of equation (9). Instead of assuming constant correlation between the national and regional stock index return series, as in Bollerslev (1990) and many others, we allow it to vary across time to capture the time varying nature of the stock market integration process. The conditional covariance equation is shown below:[8]

$$h_{NE,t} = d_0 + d_1 \sqrt{h_{N,t}.h_{E,t}} + d_2 h_{NE,t-1} \quad (10)$$

where the dynamics of the conditional correlation coefficient have been modeled based on the cross-product of standard errors of the national and regional stock index returns and past conditional correlations. Hence, the time varying conditional correlations can be computed as the standardized covariance

[8] An alternative covariance structure was estimated as $h_{NE,t} = d_0 + d_1 \sqrt{h_{N,t}.h_{E,t}}$ to ensure that the results obtained were robust to different functional forms for the conditional covariance equation. This alternative specification made no major differences to our parameter estimates for the bivariate EGARCH model. The cross-product of the unexpected returns (shocks) from the conditional mean equations has been omitted due to the complexity of these shock terms in the bivariate EGARCH framework.

$$r_t = \frac{h_{NE,t}}{\sqrt{h_{N,t}.h_{E,t}}} \quad (11)$$

and can be used to indicate the level of co-movement between national and regional (German) bond index returns. Specifically, this measures the contemporaneous conditional correlation between the two series and has been used in this paper to proxy the degree of integration between the individual bond markets and the German market[9].

Finally, the bivariate ARMA-EGARCH-*t* model is implemented for the bond index returns data via maximum likelihood estimation of the following log likelihood function[10]

$$L_T(q_f) = \sum_{t=1}^{T} l_t(q_f) = \sum_{t=1}^{T} \begin{bmatrix} -\left(\frac{k}{2}\right)\log(2p) - \frac{1}{2}\log\left(\frac{d-2}{d}\right) - \frac{1}{2}\log|H_t| - \frac{k}{2}\log\left(\frac{d}{2}\right) \\ +\log\Gamma\left(\frac{d+k}{2}\right) - \log\Gamma\left(\frac{d}{2}\right) - \left(\frac{d+k}{2}\right)\log\left(1+\frac{e_t' H_t^{-1} e_t}{d-2}\right) \end{bmatrix} \quad (8)$$

where k = 2 in the bivariate case, q_f is the vector of parameters to be estimated, T is the number of observations. As discussed above, a conditional bivariate student's *t* distribution with variance-covariance matrix H_t and *d* degrees of freedom has been assumed for the joint distribution of the two error processes instead of the standard bivariate normal distribution in order to account for possible leptokurtosis in the joint conditional densities (see Bollerslev, (1987) and Hamilton, (1994)). The advantage of employing this distribution is that the unconditional leptokurtosis observed in most high-frequency asset price data sets can appear as conditional leptokurtosis and still converge asymptotically to the Normal distribution as *d* approaches infinity (usually in lower-frequency data). As shown below, this is well suited for the dynamics of those stock market returns employed.

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9 Typically, time varying conditional correlations have been used more in the domain of risk management for calculating short-term hedge ratios to reflect current market conditions. However, time varying conditional correlations have been used in recent macroeconomic research papers in recognition that static correlations are too simplistic and are blurred by the transition process (see Babetski et al., 2002 and Sarkar and Zhang, 2002).

10 The Simplex algorithm was first used to determine appropriate starting values for parameter estimates then numerical optimization was based on the Broyden, Fletcher, Goldfarb and Shanno (BFGS) algorithm.

4. Empirical Results

The bond index returns, in levels, contain a unit root with zero drift[11]. Thus the cointegration analysis is possible, in which we find that on the whole, bond markets in the existing EU members are already fully integrated, however, those in the accession countries are not as well integrated with the existing members. Neither is the UK bond market well integrated with the rest of the EU, which perhaps is not surprising given that the series of large public sector surpluses in the UK in the 1980s had greatly reduced the volume of debt outstanding and thus liquidity resulting in the relatively small size of the public bond market. For this reason, the UK gilt market is, relative to the countries size, negligible.

We show results for the dynamic cointegration analyses in Figures 2 and 3. Figure 2 shows results for the global, recursive, analysis. The data are initially estimated over the first 500 observations, equating to approximately end-May 2000. Thereafter 20 observations, 4 weeks data, are added each iteration and the data reanalyzed. The figures show the trace statistics normalized to the asymptotic 90% critical values – thus a value greater than 1 implies cointegration and less than 1 no cointegration. It is clear that over the time period in general there is consistent evidence of cointegration indicating that the markets are in a stable relationship: the bond markets of the accession countries and those of the existing countries form part of a system[12]. However, the number of cointegrating vectors from the λ_{max} statistic settles at between 3 and 4, again indicating that the system is not integrating further. Recall that in a system of 9 variables full integration would be achieved with 1 or 8 cointegrating vectors. What we find here is perhaps a reflection of the near complete integration of the two sets of countries considered independently with a very weak linkage between the two sets of markets. The local plots are shown in Figure 3: the evidence is more favorable to the hypothesis of an integrated system, but again there is little evidence that the system

[11] Details available on request.

[12] Although omitted for space, the evidence from an analogous examination of the existing members is that they are multivariate cointegrated, as are, generally, the three accession countries. Details are available on request.

is increasing in convergence.

The Haldane and Hall convergence factors for the three accession countries are shown in Figure 4. It is clear that these bond markets are not in general close to convergence, with the exception of Poland, which has converged to the UK. On a more close examination, we find that the convergence factors to the UK are declining, indicating that the markets are converging to the UK, while the factors to Germany are increasing. Thus, In so far as there is any evidence it is that the markets are converging to the UK than to Germany. This has obvious implications for monetary integration, which we shall discuss later. In general, this would appear to cast some doubt on the appropriateness of moving towards involvement in the Euro.

The series of conditional correlations generated from the bivariate EGARCH estimations for our sample countries are shown in Figure 5. The first column of figures show the three accession countries' conditional correlations to Germany, as the proxy of the EU, whereas figures in the second and the third columns show the conditional correlations of the other countries against Germany. For the EU country bond markets, the level of integration with Germany is shown to be very high and stable over the sample - the conditional correlation is very close to positive one in all cases. In the case of UK, the conditional correlation with Germany is relatively smaller and there is some evidence of falling co-movement until late 2000 before it started to have an upward trend. As for the three accession countries, all three show considerably lower levels of integration as evidenced by low conditional correlations. In the Czech market, the conditional correlation hovers around 0.024, whereas Hungary and Poland show higher conditional correlations (average of 0.75 and 0.4, respectively), but are still considerably lower than those shown by the EU markets. In summary, the bond market integration, as shown by the time varying conditional correlations, suggest that the existing EU member markets are tightly integrated whereas the UK and the other newly joining member states show relatively lower and no clear sign of increasing degree of integration with Germany.

Table 2 reports a summary of the estimation results of the bivariate EGARCH(1,1) models shown in equations (8) and (9). Specifically it reports only the mean and volatility spillover coefficients between individual country bond market and the German market. It is clear that there is a significantly richer and greater set of both the first and the second moment spillover relationships between the existing EU countries and Germany than between the accession countries and Germany. Belgium, Italy and the Netherlands share bivariate relationships at mean, regular and asymmetric volatility terms with Germany; Ireland has a univariate relationship with Germany, receiving mean and volatility; France has a bivariate relationship with Germany through volatility, while the UK has a bivariate relationship (significant only at 10%) only in asymmetric volatility. However, none of the accession countries has a bivariate relationship with Germany as deep as Netherlands-Italy-Belgium; the country with the deepest relationship is Hungary, then Poland, and then the Czech republic, which has little relationship. The spillover results reinforce the results of the contemporaneous linkages shown by the conditional correlation analyses reported above. Thus, we report the presence of significant contemporaneous and dynamic market linkages between the EU countries and Germany, while those between accession countries and the UK with Germany are considerably weaker, usually unidirectional and there is no evidence of these linkages improving.

5. Conclusions

We have noted that previous research on bond market integration is predominantly static in nature. This paper has examined, from a variety of dynamic perspectives, the evolving nature of the relationship between the MSCI bond indices of selected European countries, distinguishing between those that most recently joined the EU and more established members. We have examined the dynamic nature of the linkages via an dynamic cointegration, Haldane and Hall's Kalman filtering method and bivariate EGARCH modeling perspectives. We provide evidence for strong contemporaneous and dynamic linkages between existing EU member bond markets with that of

Germany. For the UK and the three accession countries of Czech republic, Poland and Hungary, however, we find such linkages relatively weak but stable over the sample. Convergence, so far as it exists, appears to be slow and towards the UK for Poland, the largest of the new members. It appears that the pre-accession measure to achieve economic convergence were insufficient to generate bond market integration. Thus, our results have an important policy implication in that the government bond market convergence requires more than monetary and fiscal policy coordination. That is, bond market convergence requires policies designed specifically to address issues unique to this segment of the financial market.

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Figure 1 Global Trace and Vector

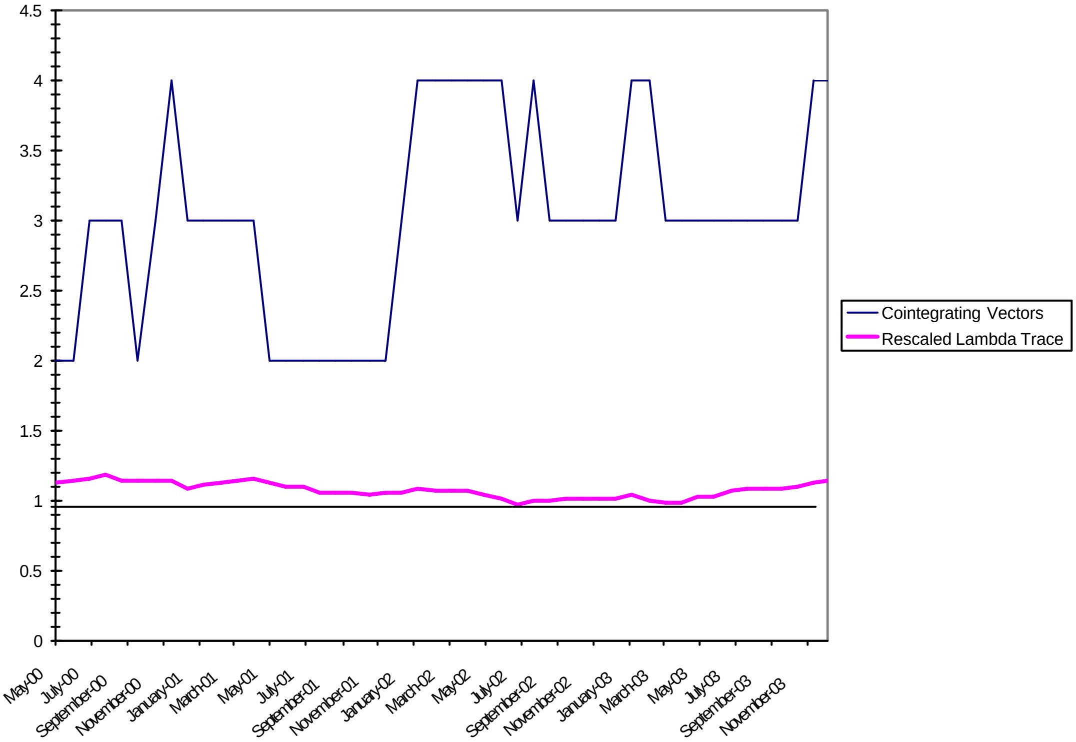

Figure 2:Local Trace and Vector

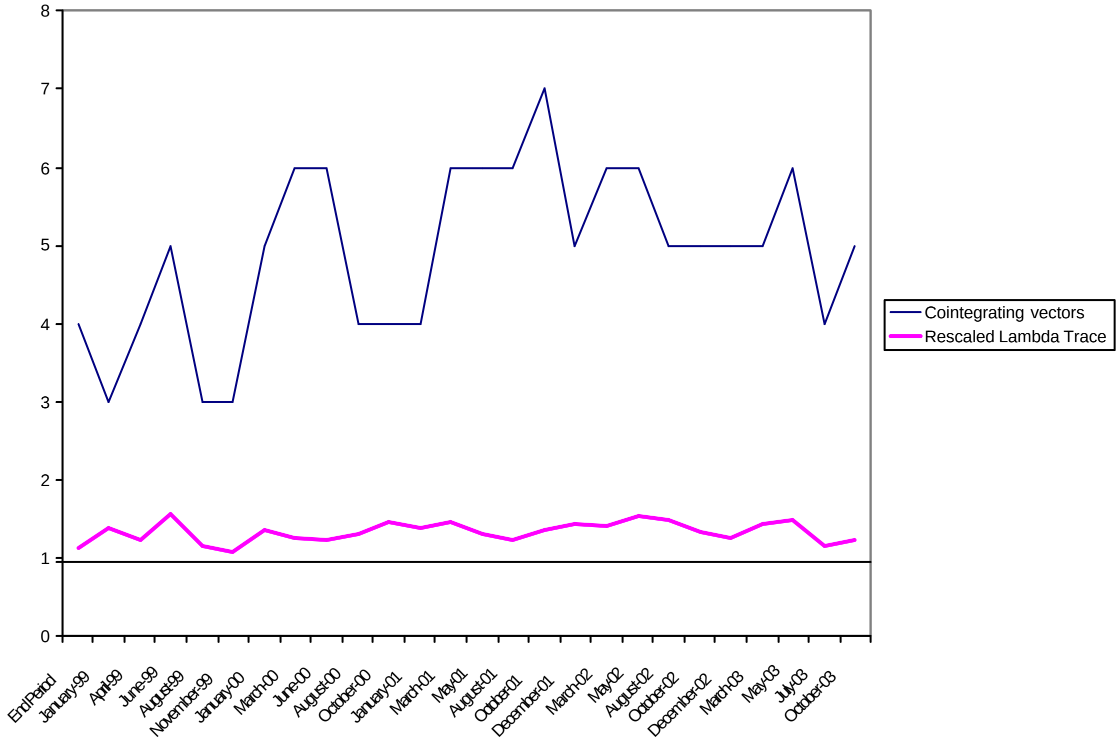

Figure 3: Haldane and Hall convergences

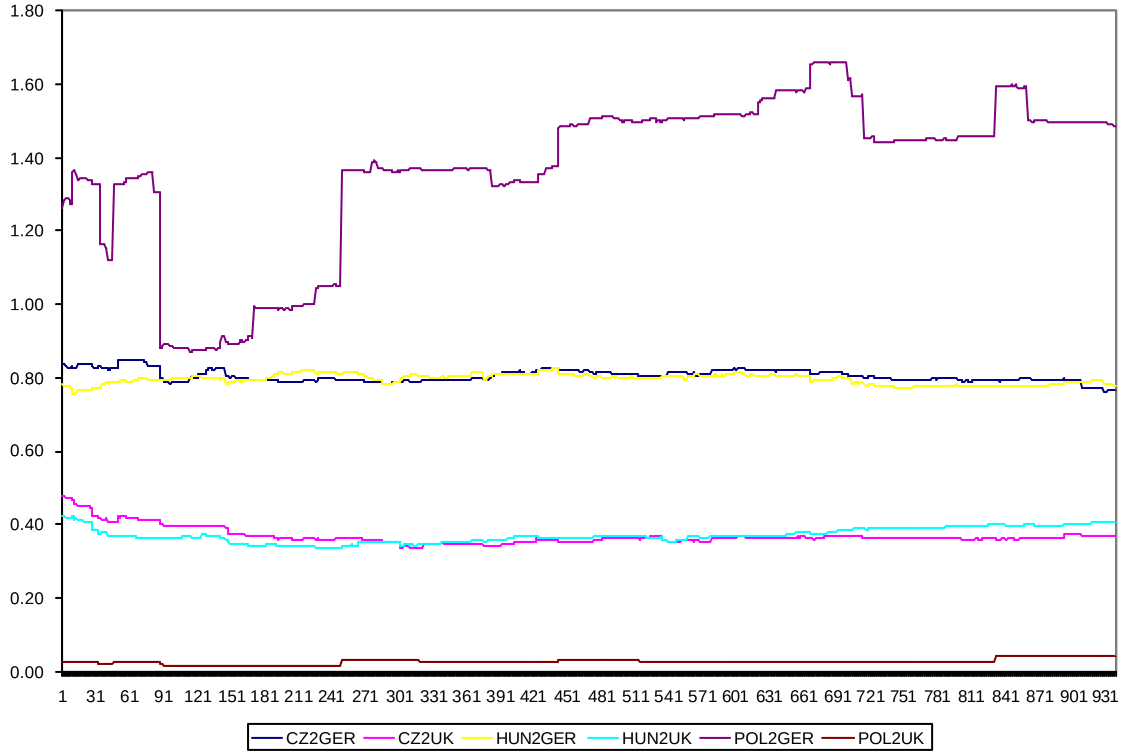

Figure 4. Time-variations in European bond market integration:

1/7/1998-31/12/2003

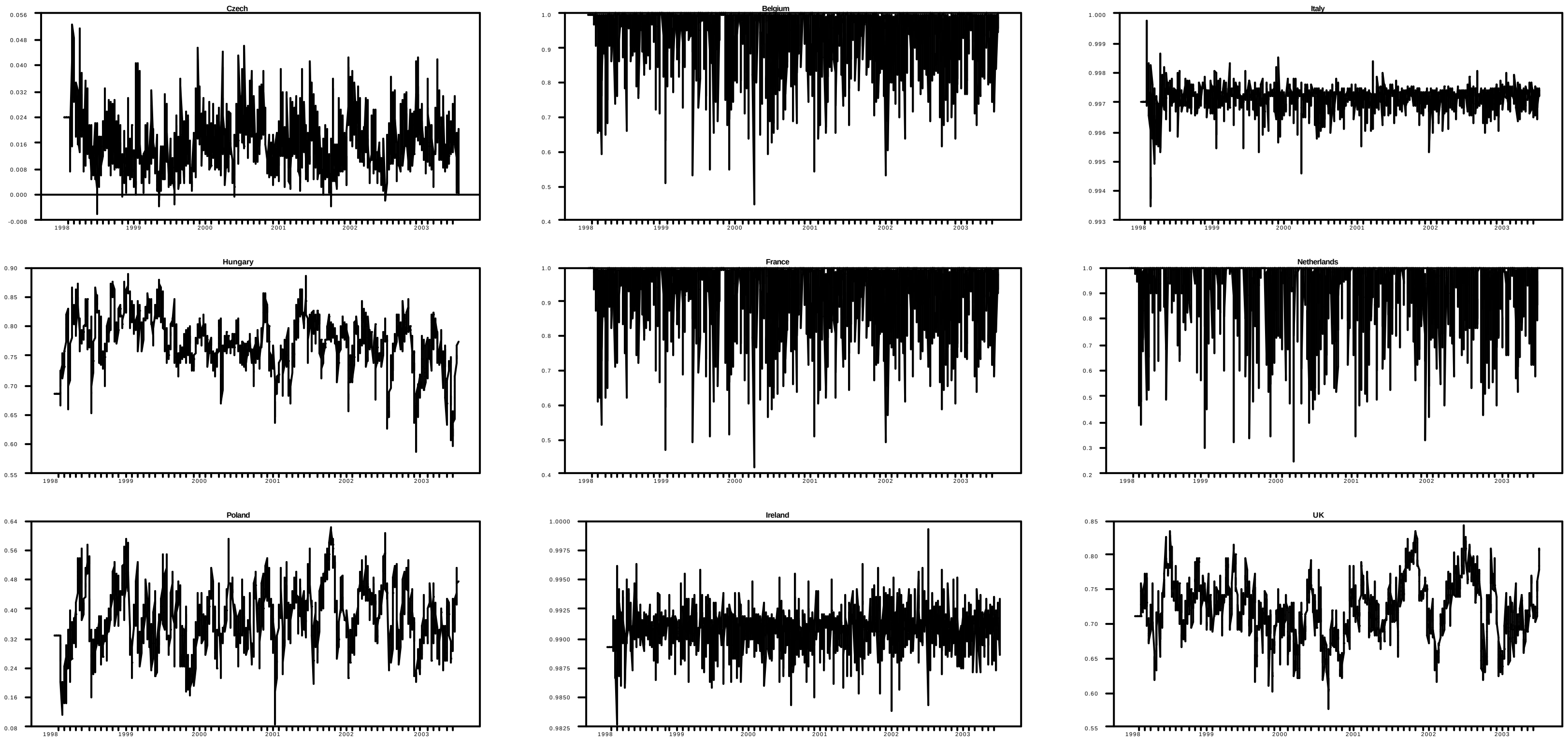

Table 1. Descriptive Statistics on daily MSCI bond index returns (%), 1/7/1998-31/12/2003

This table shows the summary statistics for the bond index returns. Asymptotic p -values are shown in the brackets. *,**,*** denote statistical significance at the 10, 5 and 1% level respectively. Test results for H_0:Skewness=0 and H_0:Excess kurtosis=0 are indicated. Q(40) is the Ljung-Box test statistic for serial correlation up to the 40th order in the return series (since $\sqrt{N = 1435} \approx 40$); $Q^2(40)$ is the Ljung-Box test statistic for serial correlation up to the 40th order in the squared returns. $Q_b(40)$ and $Q^2{}_b(40)$ are the bivariate Ljung-Box tests for joint white noise in the linear and squared returns up to the 40th order. The Engle Ng Sign bias joint test is an LM test on the significance of all three regressors in a regression of z_t^2 on S_{t-1}^- , $S_{t-1}^- \cdot e_{t-1}$ and $S_{t-1}^+ \cdot e_{t-1}$ where e_{t-1} are lagged demeaned bond market returns, $z_t^2 = \left(e_t / \sqrt{s^2} \right)^2$ and s^2 is the unconditional variance of the daily bond market returns.

	Bond Index Return				Test of univariate *iid*		Test of bivariate *iid* (with Germany)		Engle Ng Sign Bias
	Mean	Variance	Skewness	Excess Kurtosis	Q(40): $?^2(40)$	$Q^2(40)$: $?^2(40)$	$Q_b(40)$: $?^2(160)$	$Q^2{}_b(40)$: $?^2(160)$	JointTest:$?^2(3)$
New EU members:									
Czech	0.056	0.543	0.295*** [0.000]	1.033*** [0.000]	45.143 [0.266]	105.839*** [0.000]	140.968 [0.858]	225.174*** [0.001]	0.635 [0.888]
Hungary	0.045	0.611	-0.586*** [0.000]	4.837*** [0.000]	45.432 [0.256]	257.530*** [0.000]	113.522 [0.998]	371.959*** [0.000]	19.766*** [0.000]
Poland	0.052	0.593	-0.377*** [0.000]	3.066*** [0.000]	72.257*** [0.001]	277.988*** [0.000]	127.583 [0.972]	371.230*** [0.000]	83.754*** [0.000]
Existing EU members and the UK									
Belgium	0.032	0.528	0.161** [0.013]	0.991*** [0.000]	30.211 [0.869]	50.912 [0.116]	298.226*** [0.000]	165.817 [0.360]	1.124 [0.771]
France	0.031	0.528	0.173*** [0.007]	1.033*** [0.000]	29.819 [0.880]	48.447 [0.169]	198.438** [0.021]	231.686*** [0.000]	1.313 [0.726]
Ireland	0.033	0.557	0.118* [0.068]	1.015*** [0.000]	31.682 [0.823]	47.539 [0.193]	254.932*** [0.000]	125.563 [0.980]	1.318 [0.725]
Italy	0.031	0.513	0.137** [0.034]	1.034*** [0.000]	30.473 [0. 862]	50.863 [0.117]	338.878*** [0.000]	334.921*** [0.000]	2.477 [0.479]
Netherlands	0.031	0.523	0.178*** [0.006]	1.018*** [0.000]	30.006 [0.875]	49.469 [0.145]	204.418** [0.010]	249.609*** [0.000]	1.284 [0.733]
UK	0.028	0.358	0.074 [0.253]	0.942*** [0.000]	42.183 [0. 377]	70.223*** [0.002]	102.000 [0.999]	118.417 [0. 994]	9.424** [0.024]
Germany	0.030	0.521	0.173*** [0.007]	1.004*** [0.000]	29.832 [0.880]	50.429 [0.125]			2.698 [0.441]

Table 2. Bivariate EGARCH-t summary results on bond market integration:

1/7/1998-31/12/2003

This table provides a summary of the linkages between government bond markets in our sample accession and existing EU countries with the German government bond market. The full models estimated are as in equations (8)-(9)[(a)]

$$R_{N,t} = a_{cN} + \sum_{i=1}^{p_E} a_{rE,i} R_{E,t-p_E} + \sum_{j=1}^{q_N} a_{mN,j} e_{N,t-q_N} + e_{N,t}$$

$$R_{E,t} = a_{cE} + \sum_{i=1}^{p_N} a_{rN,i} R_{N,t-p2} + \sum_{j=1}^{q_E} a_{mE,j} e_{E,t-q_E} + e_{E,t}$$

$$\ln h_{N,t} = bc_N + b_{hN} \ln h_{N,t-1} + \left[be_{N1} \frac{e_{N,t-1}}{\sqrt{h_{N,t-1}}} + be_{N2}\left(\frac{|e_{N,t-1}|}{\sqrt{h_{N,t-1}}} - \sqrt{\frac{2}{p}} \right) \right] + \left[b_{E1} \frac{e_{E,t-1}}{\sqrt{h_{E,t-1}}} + b_{E2}\left(\frac{|e_{E,t-1}|}{\sqrt{h_{E,t-1}}} - \sqrt{\frac{2}{p}} \right) \right],$$

$$\ln h_{E,t} = bc_E + b_{hE} \ln h_{E,t-1} + \left[be_{E1} \frac{e_{E,t-1}}{\sqrt{h_{E,t-1}}} + be_{E2}\left(\frac{|e_{E,t-1}|}{\sqrt{h_{E,t-1}}} - \sqrt{\frac{2}{p}} \right) \right] + \left[b_{N1} \frac{e_{N,t-1}}{\sqrt{h_{N,t-1}}} + b_{N2}\left(\frac{|e_{N,t-1}|}{\sqrt{h_{N,t-1}}} - \sqrt{\frac{2}{p}} \right) \right]$$

	Return Spillovers		Volatility Spillovers			
	From GER	To GER	From GER		To GER	
			Asymmetric	Volume	Asymmetric	Volume
	αr_E	αr_N	$ß_{E1}$	$ß_{E2}$	$ß_{N1}$	$ß_{N2}$
New EU Members:						
Czech	0.029 [0.353]	0.040 [0.158]	0.033 [0.118]	0.074* [0.050]	0.080** [0.021]	-0.015 [0.801]
Hungary	-0.003 [0.798]	-0.047*** [0.002]	0.109*** [0.000]	-0.137*** [0.000]	-0.057** [0.047]	-0.011 [0.586]
Poland	0.118*** [0.000]	-0.012 [0.601]	0.078*** [0.002]	0.072* [0.051]	0.032 [0.343]	-0.001 [0.991]
Existing EU members and the UK:						
Belgium	0.019*** [0.000]	0.020*** [0.000]	0.046*** [0.000]	0.052*** [0.000]	0.037*** [0.000]	0.036*** [0.000]
France	0.010 [0.866]	0.017 [0.797]	0.059*** [0.000]	0.005 [0.300]	0.043*** [0.000]	0.063*** [0.000]
Ireland	0.408*** [0.000]	-0.043 [0.240]	-0.015*** [0.000]	-0.005 [0.809]	-0.040*** [0.000]	-0.024 [0.233]
Italy	0.053*** [0.000]	0.052*** [0.000]	-0.016*** [0.000]	0.129*** [0.000]	-0.013*** [0.000]	0.116*** [0.000]
Netherlands	-0.017*** [0.000]	0.034*** [0.000]	0.063*** [0.000]	0.047*** [0.000]	0.014*** [0.000]	-0.008** [0.035]
UK	0.003 [0.859]	-0.028 [0.194]	-0.030* [0.083]	0.083 [0.137]	0.050* [0.073]	-0.035 [0.429]

(a) The full set of estimation results have not been included for space considerations but they can be provided upon request from the authors.

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