Dynamics of bond market integration between established and accession European Union countries

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Abstract

In this paper, we examine the integration of European government bond markets using daily returns over the 1998–2003 period with a set of complementary techniques to assess the time varying level of financial integration. We find evidence of strong contemporaneous and dynamic linkages between Euro zone bond markets with that of Germany. However, there is much weaker evidence outside of the Euro zone for the three accession markets of Czech Republic, Hungary and Poland, and the UK. In general, the degree of integration for these markets is weak and stable, with little evidence of further deepening despite the increased political integration associated with further enlargement of the European Union (EU).

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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 historical stage of EU enlargement will also have financial implications. Whilst there is substantial evidence of convergence in the present bond markets within the European Union (e.g., see Galati and Tsatsaronis, 2003), less is known about the level and dynamics of financial integration between the accession and established members. In this study, we focus on integration between government bond markets of three major accession countries, Poland, the Czech Republic and Hungary, as well as a subset of countries already belonging to the EU, Belgium, France, Ireland, Italy, Netherlands, UK and Germany. The choice of countries is determined by data and economic factors. In regards to data, the three accession countries chosen represent those that have the longest available time series data comparable to the established EU countries. In economic terms these countries represent the largest, most developed economies amongst the accession and established member groups, 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 (e.g., Aggarwal et al., 2003; Barr and Priestley, 2004; Bekaert and Harvey, 1995). For government bonds, Ilmanen (1995) provided one of the first assessments on time varying expected returns using an asset pricing model. Extending from this, Barr and Priestley (2004) applied a similar framework to assess international bond market integration. Moreover, both Clare and Lekkos (2000) and Cappiello et al. (2003) have found significant variations in international bond market return comovements but they do not interpret this in a financial market integration context. Like Cappiello et al. (2003), Christiansen (2003) has also found some changes in European bond markets since the introduction of the Euro. Christiansen (2003) recently used the AR-GARCH model of Bekaert et al. (2002) 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 European Monetary Union (EMU) bond markets with the introduction of the Euro but not in non-EMU countries where country effect remains strong. Given that Driessen et al. (2003) find factors relating to the term structure to explain most of the variations in international excess bond returns, it is conceivable that economic convergence required as part of EU membership has inevitably led to higher levels of bond market convergence. However, this remains to be determined for EU accession members, as there is little evidence of the extent, still less the dynamics, of bond market integration.

The attention on comovements across government bond markets in the literature pales in comparison to 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 returns 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 European bond market
integration in terms of the returns, variances and covariances. This paper aims to address this void and provide empirical evidence on the dynamic nature of bond market integration amongst the established EU and accession countries. Given that yield differences between government bond issues in the European Monetary Union are small (through monetary policy coordination), we expect EMU 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 from the existing markets. This is vital for the success of the European Union’s new phase of accession beginning in May 2004. Barr and Priestley (2004) believe the economic costs and benefits of international bond market integration are likely to be significant, ultimately leading to lower cost of fiscal funding for governments. This suggests that benefits of integration are likely to outweigh the costs.

The major findings of this paper are: (i) although there are strong linkages between established EU bond markets with that of Germany, the three accession countries’ linkages are weaker and show no evidence of growing integration with the EU core in the near term; (ii) the UK bond market’s linkage with Germany is relatively weaker than the EMU countries’; (iii) of the three accession countries, the Czech Republic is the least integrated with the established EU bloc due to high currency risks.

The remainder of this paper is organized as follows: Section 2 discusses the bond market index data used in this study; Section 3 details the three empirical methodologies employed; the estimation results are discussed in Section 4. Finally, we provide our conclusions in Section 5.

2. Data description

The data that we use are all-maturity total returns on MSCI government bond indices for the Czech Republic, Hungary, Poland, Belgium, France, Ireland, Italy, Netherlands, the UK and Germany sourced from Thompson Datastream.1 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 EU accession countries: Czech Republic, Hungary and Poland. The bond indices are all denominated in US dollars. We used the data available from 30 June 1998 to 31 December 2003 (yielding 1435 usable observations) to calculate returns as first log differences.2 In our analyses, we choose to use the German government bond index as the proxy for the entire EMU given it has a correlation of 0.995 with the EMU-11 tracker government bond index (over the available sample period) which can also be obtained from Datastream.3 This will also avoid spurious integration results as individual bond markets will not be a composite of the EU regional proxy index. We have included data on the UK as well as Euro zone 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.

1 Total return indices account for both price changes and dividends reinvested.
2 We follow the existing literature in applying log-changes of total return government bond indices (e.g., Bodart and Reding, 1999; Christiansen, 2003; Driessen et al., 2003).
3 This series is only available starting from the 1st January 1999.
Table 1
Descriptive statistics on daily MSCI bond index returns (%), 1/7/1998–31/12/2003

<table>
<thead>
<tr>
<th>Bond index return</th>
<th>Test of univariate iid</th>
<th>Test of bivariate iid (with Germany)</th>
<th>Engle Ng sign bias</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Mean</td>
<td>Variance</td>
<td>Skewness</td>
</tr>
<tr>
<td>New EU members</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Czech</td>
<td>0.056</td>
<td>0.543</td>
<td>0.295***</td>
</tr>
<tr>
<td>Hungary</td>
<td>0.045</td>
<td>0.611</td>
<td>−0.586***</td>
</tr>
<tr>
<td>Poland</td>
<td>0.052</td>
<td>0.593</td>
<td>−0.377***</td>
</tr>
<tr>
<td>Existing EU members and the UK</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Belgium</td>
<td>0.032</td>
<td>0.528</td>
<td>0.161***</td>
</tr>
<tr>
<td>France</td>
<td>0.031</td>
<td>0.528</td>
<td>0.173***</td>
</tr>
<tr>
<td>Ireland</td>
<td>0.033</td>
<td>0.557</td>
<td>0.118***</td>
</tr>
<tr>
<td>Italy</td>
<td>0.031</td>
<td>0.513</td>
<td>0.137***</td>
</tr>
<tr>
<td>Netherlands</td>
<td>0.031</td>
<td>0.523</td>
<td>0.178***</td>
</tr>
<tr>
<td>UK</td>
<td>0.028</td>
<td>0.358</td>
<td>0.074 (0.253)</td>
</tr>
<tr>
<td>Germany</td>
<td>0.030</td>
<td>0.521</td>
<td>0.173***</td>
</tr>
</tbody>
</table>

This table shows the summary statistics for the bond index returns. Asymptotic p-values are shown in the brackets. 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 $\sqrt{N} = 1435 \approx 40$; $Q^b(40)$ is the Ljung-Box test statistic for serial correlation up to the 40th order in the squared returns. $Q_b(40)$ and $Q^b(40)$ 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 $S_{t-1}$, $S_{t-1}^*e_{t-1}$ and $S_{t-1}^*e_{t-1}$ where $e_{t-1}$ are lagged demeaned bond market returns, $z_t^2 = (e_t^* / \sigma_t^2)$ and $\sigma_t^2$ is the unconditional variance of the daily bond market returns.

* Denotes statistical significance at the 10% level.
** Denotes statistical significance at the 5% level.
*** Denotes statistical significance at the 1% level.
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 perceived 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 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 are considerably larger than the other countries. The bond index returns are not serially correlated in the first moment in all cases except Poland. However, significant correlation in the second moments is found in all three accession countries and the UK which is clear evidence of time varying volatility in these markets. In addition, the significance of the bivariate 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. 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 conditional variances of government bond index returns in the established EU countries.

3. Dynamic methodologies

We have noted earlier that there is a need to consider the time varying nature of financial market integration. The techniques we use here expressly allow us to capture this important 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 form. 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 opportunities. The requirement for assets that are integrated in an economic sense to share common stochastic factors, is an alternative definition of cointegration, as pointed out in Chen and Knez (1995).

In the literature, two primary methods exist to examine the degree of cointegration among indices. 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); Johansen and Juselius (1990)), hereafter...

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4 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.
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. The JJ approach generates two statistics of primary interest. The first is the $\lambda_{\text{trace}}$ 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 $\lambda_{\text{max}}$ 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.\(^6\)\(^,\)\(^7\) 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 two-fold: first, the largest value of the $\lambda_{\text{trace}}$ statistic which tests the general hypothesis of no cointegration versus cointegration, and second, the number of cointegrating vectors given by the $\lambda_{\text{max}}$ 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 estimation will simply lead to an upward trend in the $\lambda_{\text{trace}}$ statistic. It should be noted that in general the $\lambda_{\text{trace}}$ statistic is more powerful and to be preferred to the $\lambda_{\text{max}}$ statistic.

3.2. 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 settings.\(^8\) The Haldane and Hall (hereafter HH) method estimates a simple equation of the following specification

\[
\ln \left( \frac{E_{jt}}{E_{Bt}} \right) = \alpha + \beta_t \ln \left( \frac{E_{jt}}{E_{Xt}} \right) + \varepsilon_{jt} \quad (1)
\]

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 SE Asian markets, Manning (2002) imposes the US market

\(^{6}\) 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.

\(^{7}\) 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. The Barari and Sengupta (2002) paper concentrates on the $\lambda_{\text{trace}}$ statistic.

\(^{8}\) 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).
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 examine convergence towards the German market, to keep congruence with the other methods. Negative values of \( \beta_t \) indicate divergence, as does a tendency to move further from zero.

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 with the addition of each daily data point. Let \( Y_t = \alpha_t + X_t \beta_t + \varepsilon_t \), var(\( \varepsilon_t \)) = \( \eta_t \) be the measurement equation of interest. If we set \( \beta_t \) as the coefficient of interest at time \( t \), then the transition equation is given by \( \beta_t = \beta_{t-1} + v_t \), var(\( v_t \)) = \( M_t \). Given the estimate of \( \beta_{t-1} \) from information up to that period (\( \beta_{t-1}[t-1] \)) with the associated covariance matrix \( \sum_{t-1} \), the updated estimate is given by Eqs. (2)–(4).

\[
S_t = \sum_{t-1} + M_t
\]

\[
\Sigma_t = S_t - S_t X_t'(X_t S_t X_t')^{-1} X_t S_t
\]

\[
\beta_{t|t} = \beta_{t-1|t-1} + S_t X_t'(X_t S_t X_t')^{-1}(Y_t - \alpha_{t-1} X_t \beta_{t-1|t-1})
\]

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 a bivariate version of Nelson’s (1991) EGARCH framework (incorporating Student’s \( t \) 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 Eq. (5) to capture the dynamics between mean bond market returns for each individual country (subscripted \( N \)) and Germany (proxy for the EMU, subscripted \( E \)).

\[
R_{N,t} = \alpha_{cN} + \sum_{i=1}^{pN} \alpha_{rN,i} R_{N,t-i} + \sum_{j=1}^{qN} \alpha_{mN,j} \varepsilon_{N,t-j} + \varepsilon_{N,t}
\]

\[
R_{E,t} = \alpha_{cE} + \sum_{i=1}^{pE} \alpha_{rE,i} R_{N,t-i} + \sum_{j=1}^{qE} \alpha_{mE,j} \varepsilon_{E,t-j} + \varepsilon_{E,t}
\]

With

\[
\varepsilon_t = \begin{bmatrix} \varepsilon_{N,t} \\ \varepsilon_{E,t} \end{bmatrix} \sim t(0, H_t, d), \quad H_t = \begin{bmatrix} h_{N,t} & h_{NE,t} \\ h_{EN,t} & h_{E,t} \end{bmatrix}
\]
In essence, $R_N, t$ is the national bond index return that is a function of past returns in the German market and past idiosyncratic shocks, $\varepsilon_N, t$, and $R_E, t$ is the German bond market index return that is a function of past returns in country $N$ and its own past shocks, $\varepsilon_E, t$. Specifically, the regional and country mean spillover effects can be quantified by the sign and magnitude of the estimated coefficients for the lagged German 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 non-linear serial correlation in the standardized residuals, $\varepsilon_N, t/\sqrt{h_N, t}$ and $\varepsilon_E, t/\sqrt{h_E, t}$ which are jointly $t$ distributed.

In Eq. (6), 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.

\[
\begin{align*}
\ln h_N, t & = \beta_{cN} + \beta_{hN} \ln h_N, t-1 + \beta_{cN1} \frac{\varepsilon_{N,t-1}}{\sqrt{h_N, t-1}} + \beta_{cN2} \left( \frac{|\varepsilon_{N,t-1}|}{\sqrt{h_N, t-1}} - \sqrt{\frac{2}{\pi}} \right) \\
& + \beta_{E1} \frac{\varepsilon_{E,t-1}}{\sqrt{h_E, t-1}} + \beta_{E2} \left( \frac{|\varepsilon_{E,t-1}|}{\sqrt{h_E, t-1}} - \sqrt{\frac{2}{\pi}} \right) \\
\ln h_E, t & = \beta_{cE} + \beta_{hE} \ln h_E, t-1 + \beta_{cE1} \frac{\varepsilon_{E,t-1}}{\sqrt{h_E, t-1}} + \beta_{cE2} \left( \frac{|\varepsilon_{E,t-1}|}{\sqrt{h_E, t-1}} - \sqrt{\frac{2}{\pi}} \right) \\
& + \beta_{N1} \frac{\varepsilon_{N,t-1}}{\sqrt{h_N, t-1}} + \beta_{N2} \left( \frac{|\varepsilon_{N,t-1}|}{\sqrt{h_N, t-1}} - \sqrt{\frac{2}{\pi}} \right)
\end{align*}
\]

The conditional variance for the national (German) return series is determined by its own past variance, its own negative and positive past unanticipated shocks as well as those from the German (national) bond market index return. In this context, the regional (German) 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 Eq. (6). Instead of assuming constant correlation between the national and regional bond market 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 bond market integration process. The conditional covariance equation is shown below:

\[
h_{NE, t} = \delta_0 + \delta_1 \sqrt{h_N, t} \sqrt{h_E, t} + \delta_2 h_{NE, t-1}
\]

where the dynamics of the conditional correlation coefficient have been modeled based on the cross-product of standard errors of the national and regional bond market index returns.

\[\text{An alternative covariance structure was estimated as } h_{NE, t} = \delta_0 + \delta_1 \sqrt{h_N, t} \sqrt{h_E, t} \text{ to ensure that the results obtained were robust to different functional forms for the conditional covariance equation. This alternative specification made no major differences to our parameter estimates for the bivariate EGARCH model. The cross-product of the unexpected returns (shocks) from the conditional mean equations has been omitted due to the complexity of these shock terms in the bivariate EGARCH framework.}\]
and past conditional correlations. Hence, the time varying conditional correlations can be computed as the standardized covariance

\[ \rho_t = \frac{h_{NE,t}}{\sqrt{h_{N,t}h_{E,t}}} \]  

(8)

and can be used to indicate the level of comovement 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 national bond markets and the German market.\(^{10}\)

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.\(^{11}\)

\[
L_T(\theta_f) = \sum_{t=1}^{T} l_t(\theta_f)
\]

\[
= \sum_{t=1}^{T} \left[ -\frac{k}{2} \log (2\pi) - \frac{1}{2} \log \left( \frac{d-2}{d} \right) - \frac{1}{2} \log |H_t| - \frac{k}{2} \log \left( \frac{d}{2} \right) + \log \Gamma \left( \frac{d+k}{2} \right) - \log \Gamma \left( \frac{d}{2} \right) - \log \left( 1 + \varepsilon_t^2 H_t^{-1} \varepsilon_t \right) \right]
\]  

(9)

where \( k = 2 \) in the bivariate case, \( \theta_f \) 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; 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 bond market returns employed.

4. Empirical results

All bond index returns, in levels, contain a unit root with zero drift.\(^{12}\) Thus a cointegration analysis is possible, in which we find that, on the whole, bond markets in the established EU

\(^{10}\) 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; Sarkar and Zhang, 2002).

\(^{11}\) 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.

\(^{12}\) Details available on request.
members are already fully integrated, these results corroborating existing studies such as Galati and Tsatsaronis (2003) and Cappiello et al. (2003). However, those in the accession countries are not as well integrated with the established EU bloc and this is a new result. Moreover, the UK bond market is not 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 during the 1980s had greatly reduced the volume of debt outstanding and thus liquidity, resulting in relatively small size of the public bond market. For this reason, the UK gilt market is negligible, relative to the country’s size. This is also consistent with lower unconditional correlations found by Cappiello et al. (2003) between EMU and non-EMU bond markets on a regional level.

We show results for the dynamic cointegration analyses in Figs. 1 and 2. Fig. 1 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. For ease of interpretation, the trace statistics are 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.13 However, the number of cointegrating vectors from the

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13 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.
\( \lambda_{\text{max}} \) 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 Fig. 2: the evidence is more favorable to the hypothesis of an integrated system, but again there is little evidence that the system is increasing in convergence.

The Haldane and Hall convergence factors for the three accession countries with Germany are shown in Fig. 3. Again we initialize the system over the first 500 points, and re-estimate daily. It is clear that these bond markets are not in general close to convergence with the German market.

The series of conditional correlations generated from the DCC estimations are shown in Fig. 4. The first column show the three accession countries’ time varying conditional correlations with Germany (proxy for the EU), whereas figures in the second and the third columns show the conditional correlations of the other EU countries against Germany. For individual EMU 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 unity in all cases. In the case of UK, the conditional correlation with Germany is relatively smaller and there is some evidence of divergence until late 2000 before a persistent and strong upward trend commences through 2002. As for the three accession countries all three show considerably lower levels of integration and this correlation is evidently more volatile than that of any of the existing countries. In the Czech market, the conditional correlation hovers
around 0.024, whereas Hungary and Poland show higher conditional correlations (average of 0.72 and 0.33, respectively), but are still considerably lower than those shown by the established EU markets. Despite the low correlation with the EU market however the Czech correlation is statistically different to zero, indicating that although low it is extant. In sum, bond market integration, as shown by these time varying conditional correlations, suggest that the EMU member markets within the EU 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 the EU core.

Table 2 reports a summary of the estimation results of the bivariate EGARCH-t (1,1) models shown in Eqs. (5) and (6). Specifically it reports only the mean and volatility spillover coefficients between sample EU bond markets and the German bond 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, only receiving mean and volatility; France has a bivariate relationship with Germany through volatility, while the UK has a bivariate relationship only

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14 The bivariate Ljung-Box Q statistics indicate that linear and non-linear serial correlations have been removed in most cases. The significance of the asymmetric and volume effects in cross-market volatilities indicates the suitability of a bivariate EGARCH model for modelling the bivariate nature of individual and German bond market returns in their first and second moments.
Fig. 4. Estimated dynamic conditional correlations 8 July 1998 to 31 December 2003.
Table 2
Bivariate EGARCH-t summary results on bond market integration: 1/7/1998–31/12/2003

<table>
<thead>
<tr>
<th>Return spillovers</th>
<th>Volatility spillovers</th>
</tr>
</thead>
<tbody>
<tr>
<td>From GER</td>
<td>To GER</td>
</tr>
<tr>
<td>α_E</td>
<td>α_N</td>
</tr>
</tbody>
</table>

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\}

Established EU members
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\} 0.116*** \{0.000\} −0.013*** \{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\}

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 Eqs. (5) and (6).

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

* Denotes statistical significance at the 10% level.

** Denotes statistical significance at the 5% level.

*** Denotes statistical significance at the 1% level.
in asymmetric volatility (significant only at 10%). 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, then the Czech Republic, which has little linkage. The conditional mean and volatility spillover results reinforce the evidence of the contemporaneous linkages shown by the conditional correlation analyses reported above. Thus, we report the presence of significant contemporaneous and dynamic bond market linkages between EMU countries and Germany, while those between accession countries and the UK with Germany are considerably weaker, usually unidirectional and there is no evidence of deepening in these linkages in the near term.

5. Conclusions

We note that previous research on bond market integration is predominantly static in nature. This paper has examined, from a variety of dynamic perspectives associated with different methodologies, 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 dynamic cointegration, Haldane and Hall’s Kalman filtering method and bivariate EGARCH modeling perspectives. We provide robust empirical evidence for strong contemporaneous and dynamic linkages between existing EU member country 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 weak and relatively 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 preaccession measures to achieve economic convergence were insufficient to generate rapid bond market integration and currency risk remains a problem for the Czech Republic. Thus, our results have an important policy implication in that 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.

References


