Government Bond Market Integration Within European Union

Abstract

This paper examines the extent of linkages among Euro and non-Euro government bond markets in the pre- and post-Euro introduction period. Multivariate cointegration analyses indicate absence of cointegration among the Euro bond markets in the pre-Euro period but a weak one in the post-Euro period. By contrast, there is evidence of strong cointegration in the post-Euro period among the non-Euro bond markets. Further, in the post-Euro period several bivariate linkages among the Euro bond markets exist. Finally, the US bond market appears to uni-directionally Granger-cause all Euro bond markets in both subperiods. The findings have important implications for investors, in terms of diversification benefits, and for policymakers, in terms of the proper conduct of the common monetary policy.

JEL classification codes: G15, C10, F30
Keywords: government bond, EMU, cointegration, Granger causality, policy implications

Nikiforos T. Laopodis
Assoc. Professor of Finance
Dept. of Finance
Dolan School of Business
Fairfield University
N. Benson Rd, Fairfield, CT 08624, USA
Tel. (203) 254 4000 x3273
E-mail: nlaopodis@mail.fairfield.edu

Government Bond Market Integration Within European Union

1. Introduction

The worldwide wave of financial market liberalization, facilitated by political, technological, and financial advances, has led to increased interdependence among global financial markets. This issue has attracted considerable attention in the finance literature which placed particular emphasis on the globalization of national equity markets (see, for instance, some studies on stock market integration by Bekaert and Harvey, 1997, Ng, 2000, Fratzscher, 2002, and Bessler and Yang, 2003). By contrast, the issue of bond market integration pales in comparison to stock market integration, particularly bond integration involving the European Union. It is well-known that the European Monetary Union (EMU), with the introduction of the Euro, has been to large extent responsible for the continuing integration of the European (money and) capital markets. As a result, Euro government bond markets have grown considerably since then and make up more than 55% of the total bonds outstanding within the Euro area (see Pagano and von Thadden, 2004) as well as account for a large percentage of bond investments accessible to international investors in the EMU (see Holder, 1999).

Moreover, during the 1990s the value of sovereign and private European bond markets accounted for a close to 60% of the value in the United States and about 25% of the world total (see Yang, 2005, and the references therein). In recent years, debt issuance in the Euro area more than quadrupled from their mid-1990s level. Since the introduction of the Euro, two major forces are responsible for the growth in sovereign bond issuance (see ECB, 2004). First, broad-based improvements in budgetary balances enabled governments to have very low (net) borrowing requirements and adopted buyback programs and/or bond exchanges. Second, several significant changes in the issuance of sovereign bonds has fostered efficient competition for governments. As a result of these two factors, the market for public debt is now regarded as highly liquid, competitive and most important in the Euro area.

The significance of this study is threefold. First, we investigate the trend in integration/interdependence in the European government bond market both before and after the introduction of the Euro. In that respect, it would be interesting to see if significant gains in integration were achieved following the use of the new common currency (that is, if there was a structural change in sovereign bond market integration in the Euro area after 2001). Second, for the international investor, diversification opportunities/benefits may have considerably diminished, if higher bond integration is present and sovereign bond issues are becoming close substitutes, with a wider consequence of the depreciation of the Euro. Finally, an understanding of the extent of comovement among Euro area bond markets is also important for the conduct of the common monetary policy. In other words, greater interdependence of bond yields among European countries may reduce the capability of the European Central Bank to influence long-term interest rates and thus its intent to attain price stability (see Clare and Lekkos, 2000). In addition, for policymakers, increased comovement among bond yields suggests that financial market shocks in one country are much faster transmitted to other countries, within and outside the Euro area, and may have adverse consequences for the stability of the world financial system.

Following the above considerations, the following questions are addressed in this paper. How closely are the Euro government bond markets associated with each other and has this association increased since the introduction of the common currency? Are there any short-run and/or long-run relationships between each of and among these bond markets (that is, within a bivariate and multivariate setting, respectively)? In addition, are there any structural changes within these bi- or multivariate relationships? Do we observe stronger causal relationships among these bond markets following the use of the Euro? Finally, what are the possible implications of our findings for both international investors and policymakers? These questions will be addressed via the use of correlation, cointegration, error-correction, and Granger causality analyses both within a bivariate and multivariate context for the 1994 to 2006 period.

The remainder of the paper obeys the following order. Section 2 presents a brief review of the empirical literature on bond market integration within the Euro area. Section 3 discuses the data and the methodological design of the study. Section 4 contains the preliminary and main empirical results, while section 5 discusses the findings in terms of policy implications for both investors and policymakers. Finally, section 6 summarizes the study and concludes with some general observations.

2. Literature Review

Empirical work on global bond market integration is very scarce relative to international stock market integration. Researchers have employed a wide variety of frameworks to empirically examine the linkages among international bond markets. One of the early attempts to investigate this issue for government bonds is by Ilmanen (1995), who employed a linear regression model to assess the varying (excess) returns of long-term international bonds. His evidence suggested that excess returns were highly correlated implying, in turn, international bond market integration. Clare et *al*. (1995) provided early insight into the significance of international bond market linkages for bond portfolio diversification. Some later studies, by McCauley and White (1997) and Portes and Rey (1998), emphasized the central role of European bond markets relative to the US bond market and suggested that demand for Euro assets will increase globally.

A more recent study by Clare and Lekkos (2000), using a VAR model with short-rates and terms-structure slopes as endogenous variables, examined the linkages among the US, UK, and German bond markets and found that significant variations in international bond market relationships existed and that transnational factors were most relevant during periods of financial instability. By contrast, Driessen et *al*. (2003), employing principal component analysis, reported that positive correlations were driven by term-structure levels and not by slopes (for the US, Japanese and German bond markets). Similar studies by Cappiello et *al*. (2003) and Christiansen (2003), using variations of the GARCH modeling methodology to examine volatility spillovers across bond markets, noted strong volatility-spillover effects between the US and European bond markets and both concluded that the introduction of the new common currency has strengthened these spillover effects leading to near perfect correlation among bond returns within the EMU countries. Finally, along the same line of research, Skintzi and Refenes (2006) studied the time-varying correlation structure between several individual and aggregate European bond markets indices and the US bond market and found that significant volatility spillovers from the US to all individual European bond markets were present for the 1991 to 2002 period.

Another recent line of empirical research on government bond rate differentials concentrated on liquidity and default risk premiums. Adam et *al*. (2002), using panel data techniques to test indicators designed to measure the evolution of the capital market integration in the European Union, noted that convergence in the Euro-area government bond market occurred before 1999 (the introduction of the Euro year) and not afterwards. Beale et *al*. (2004) similarly reported that bond yields were driven by common rather than local factors implying that Euro area bond integration has been high since the introduction of the Euro. Berben and Jansen (2005), using a different methodology, also arrived at the conclusion that the introduction of the Euro has strengthened the integration process. Bernoth and Schuknecht (2004) stated that the EMU area was characterized by lower liquidity premiums suggesting a higher degree of integration in the Euro area bond markets. Finally, Côté and Graham (2004) found that greater convergence in monetary and fiscal policies within EMU (and outside of it) positively contributed to the harmonization of long-term government bond yields.

3. Methodology and Data

The methodological design of the study will be comprised of the following steps. The first step entails checking if each series is integrated of the same order. This will be accomplished by applying the Kwiatkowski et al. (KPSS, 1992) test to determine if each series has a unit root. The second step is to test for cointegration to see if the bond markets share a common long-run path. Cointegration will be investigated only at the multivariate level and will be carried out via the Johansen and Juselius (JJ, 1990) approach. Next, in case of presence of a long-run relationship, we should also examine the possibility that short-run relationships among the bond markets might exist. This test will be achieved via the Granger (1969) causality test.

3.1. Unit Root, Cointegration, and Granger Causality Tests

The Kwiatkowski, Phillips, Schmidt, and Shin (KPSS, 1992) unit root test postulates that a series y_t is assumed to be (trend-) stationary under the null hypothesis. The KPSS statistic is based on the residuals from the OLS regression of y_t on the exogenous variables x_t:

$$y_t = x_t'\delta + u_t \tag{1}$$

The Lagrangian Multiplier (LM) statistic is defined as:

$$LM = \sum_t S(t)^2 \,/\, (T^2 f_0) \tag{2}$$

where f_0 is an estimator of the residual spectrum at frequency zero and where $S(t)$ is a cumulative residual function:

$$S(t) = \sum_{r=1} \hat{u}_t \tag{3}$$

based on the residuals $\hat{u}_t = y_t - x_t'\delta(0)$ from equation (1).

Although there are several tests of cointegration have been developed, the approach used in this paper is based on Johansen (1988) and Johansen and Juselius (JJ, 1990). To illustrate these approaches briefly, let X_t be a vector of N time-series variables, each of which is integrated of order 1, $I(1)$. Assume that X_t can be modeled by the following vector autoregression (VAR):

$$X_t = \Pi_1 X_{t-1} + \Pi_2 X_{t-2} + \ldots + \Pi_k X_{t-m} + \mu + e_t \tag{4}$$

where Π_1, Π_2, ..., Π_k are N × N matrices of coefficients and μ is the constant (or drift term). This equation can be rewritten as

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \Gamma_2 \Delta X_{t-2} + ... + \Gamma_{k-1} \Delta X_{t-m+1} + \Pi_k X_{t-m} + \mu + e_t \qquad (5)$$

where

$$\Gamma_i = -(I - \Pi_1 - . . . - \Pi_i) \quad \text{for } i = 1, ..., m-1 \qquad (6)$$

and

$$\Pi = -(I - \Pi_1 - ... - \Pi_m) \qquad (7)$$

The last model is expressed as a traditional VAR in first differences except for the term ΠX_{t-m}. The matrix Π is called the long-run impact matrix and its appropriate lag length is determined with the log-likelihood test statistic.

There exist three possible cases for the rank of Π. If rank(Π) = 0, then Π= 0 and all of the processes in X_t are stationary. If rank(Π) = N, then all of the variables in X_t are stationary in levels. If 0 ≤ rank(Π) ≤ N, the components of X_t are cointegrated. Finally, if rank(Π) = r, where $0 \leq r \leq N$, then there are r cointegrating vectors or stationary long-run relationships among the N variables in X_t and N - r common stochastic trends. In this case, there exists matrices α and β of dimension N × r such that $\Pi = \alpha\beta'$. The r cointegrating vectors β have the property that $\beta' X_t$ is stationary even though X_t is nonstationary. The α coefficients can be interpreted as measuring the average speed of adjustment toward the cointegrating relationships.

JJ have suggested the examination of the rank of Π, which is equal to the number of non-zero eigenvalues, as the manner in which to test for the number of cointegrating relations. In this respect, there are two tests of whether the eigenvalues of the estimated Π

matrix are significantly different from zero: the trace test and the maximum-eigenvalue test. The trace test and the maximum eigenvalue test are expressed as follows:

$$\lambda_{trace}\,(r) = -\,T\,\Sigma\,ln\,(1 - \lambda_i) \tag{8}$$

$$\lambda_{max}\,(r, r + 1) = -\,T\,ln\,(1 - \lambda_{r+1}) \tag{9}$$

where λ_i equals the estimated values of the characteristic roots obtained from the estimated Π matrix, r is the numbers of cointegrating vectors, and T equals the number of usable observations. The trace test evaluates the null hypothesis that the number of distinct cointegrating vectors is less than or equal to r against a general alternative. The maximum eigenvalue test examines the number of cointegrating vectors versus that number *plus* one. If the variables in X_t are not cointegrated, the rank of Π is zero and all characteristic roots are zero. Since $ln(1) = 0$, each of the expressions in $ln\,(1 - \lambda_i)$ will equal zero in that case. Finally, the maximum likelihood tests of JJ ensure that coefficient estimates are symmetrically distributed and asymptotically efficient using standard χ^2 tests. The critical values for these tests, for up to six variables, are given by JJ and by Osterwald-Lenum (1992) for at most eleven variables (for multivariate cointegration tests).

Because each return series may have nonzero means, deterministic trends, and/or stochastic trends, the cointegrating equations may have intercepts and deterministic trends. As a consequence, the asymptotic distribution of the likelihood ratio test statistic for cointegration does not have the usual χ^2 distribution and may depend on the assumptions made with respect to deterministic trends. Therefore, we need to make several assumptions regarding the trends underlying the series. We will consider the following five deterministic trend assumptions, namely, *H2*(r), H1 *(r), *H1*(r), H*(r), and *H*(r). These hypotheses are summarized below:

1. $$H_2(\mathrm{r}): \Pi y_{t-1} + \mathrm{B}\, x_{\mathrm{t}} = \alpha\, \beta' y_{t-1} \qquad (10)$$

The level data y_{t} have no deterministic trends and the cointegrating equations do not have intercepts.

2. $$H^*_1(\mathrm{r}): \Pi y_{t-1} + \mathrm{B}\, x_{\mathrm{t}} = \alpha\, (\beta' y_{t-1} + \rho_0) \qquad (11)$$

The level data y_{t} have no deterministic trends and the cointegrating equations have intercepts.

3. $$H_1(\mathrm{r}): \Pi y_{t-1} + \mathrm{B}\, x_{\mathrm{t}} = \alpha\, (\beta' y_{t-1} + \rho_0) + \alpha \perp \gamma_0 \qquad (12)$$

The level data y_{t} have linear trends and the cointegrating equations have only intercepts.[1]

4. $$H^*(\mathrm{r}): \Pi y_{t-1} + \mathrm{B}\, x_{\mathrm{t}} = \alpha\, (\beta' y_{t-1} + \rho_0 + \rho_1 \mathrm{t}) + \alpha \perp \gamma_0 \qquad (13)$$

The level data y_{t} and the cointegrating equations have linear trends.

5. $$H(\mathrm{r}): \Pi y_{t-1} + \mathrm{B}\, x_{\mathrm{t}} = \alpha\, (\beta' y_{t-1} + \rho_0 + \rho_1 \mathrm{t}) + \alpha \perp (\gamma_0 + \gamma_1 \mathrm{t}) \qquad (14)$$

The level data y_{t} have quadratic trends and the cointegrating equations have linear trends.

Although for the purposes of exposition of a comprehensive multivariate cointegration test we included all possible assumptions, we will make the case for selecting assumption number 3 as the most plausible one for the ensuing analysis. Under this case, no trend is assumed in either the cointegrating equation or the VAR because the bond index values are expressed in natural logarithms. As a result, when the first difference in each index is taken, the result is a return series. However, an intercept term is assumed in view of the impositions of any restrictions on the series and the absence of any dummy variables (for capturing seasonality or other events).

Short-run relationships and/or causality are two issues not adequately captured by cointegration tests. We can employ a Granger causality test with an error-correction term (if cointegration among bond index returns is found) or without an error-correction term (if cointegration is not present) to investigate these issues. To illustrate the procedure, for the bivariate case, if a pair of return series, r^b_{1t} and r^b_{2t}, are (co)integrated of order r, we can formulate an unrestricted Vector Error-Correction Model (VECM), shown for the first series only, to include the error-correction term, e_{t-1}, as follows:

$$\Delta r^b_{1t} = \mu + \sum_{i=1}^{n1} \lambda_i \, \Delta r^b_{1t-i} + \sum_{j=1}^{n2} \xi_j \, \Delta r^b_{2t-j} + \gamma \, e_{t-1} + \upsilon_t \tag{15}$$

$$\text{where } e_{t-1} = r^b_{1t-1} - \alpha \, r^b_{2t-1} \tag{16}$$

Parameters λ_i and ξ_j (if statistically significantly different from zero) reflect the short-term impact of own (bond market 1) and the other's (bond market 2) impact, respectively. Parameter α (when different from zero) measures the speed of adjustment to (the long-run) equilibrium within a single period. Obviously, if cointegration between the (or any) two series is not found, then equation (15) should not have the error term, e_{t-1}.

Generalizing equations (15) and (16) to the multivariate case, when several markets are examined concurrently, in order to identify lead/lag relationships, the following VECM specification emerges:

$$\Delta Z_t = \sum_{i=1}^{n} A_i \Delta Z_{t-i} + \sum_{i=1}^{r} \varphi_i \psi_{t-i} + \nu_t \tag{17}$$

where Z_t is an $n \times 1$ vector of bond market returns, A_i's are estimable parameters, φ_i is a vector of impulses reflecting the unexpected movements in Z_t, and ψ_{t-i} contains the r individual error correction terms derived from the r long-run relationships.

3.2. Data Description

The data used in this study are daily, nominal total returns on MSCI 10-year government bond indices from ten Euro area countries (Austria, Belgium, Finland, France, Germany, Ireland, Italy, Netherlands, Portugal, and Spain) and four non-Euro area countries (Denmark, Norway, UK and US) for the 12/31/1993 to 7/27/2006 period. The bond indices are denominated in US dollars and have been collected from *DATASTREAM* international. The return series are calculated as the growth rate of bond indices, $R_{it} = \log(P_t/P_{t-1})$ where i is the bond market. The sample period will be split into two subperiods, from 1/1/1994 to 12/31/2000 and from 1/1//2001 to 7/27/2006, in order to examine the different characteristics of the bond returns before and after the introduction of the new common currency, the Euro, in January 2001.[2]

4. Emprical Results

4.1. Descriptive Statistics

Descriptive statistics for the daily returns of all bond markets are presented in Table 1. For the pre-Euro period we observe that the average daily returns for the Euro area markets range from 0.0126% in Finland to 0.0246% in Italy and for the non-Euro markets from 0.0185% in Denmark to 0.0302% in the United Kingdom. For the Euro bond markets, the standard deviations hovered around an average of 0.62% (with Ireland having the lowest one, 0.60%) while for the non-Euro markets the average standard deviation was lower than that of the Euro bond markets, with the United States's being the lowest, 0.2672%. For the post-Euro period, we see that the returns for all bond markets (with the exception of the United Kingdom's) were much lower, around 0.44%

and the United States' being the lowest again, 0.0177%. By contrast, the standard deviations are much higher now than in the pre-Euro period for all bond markets and several markets experienced sharp increases in risk like Belgium, Ireland, Spain, and the United Kingdom.

The skewness and kurtosis statistics suggest that positive shocks were common in the pre-Euro period but negative shocks were prevalent in the post-Euro period for all bond markets and that large shocks are more frequent than expected in all bond markets in both subperiods, albeit of lower intensity in the post-Euro subperiod. Finally, all bond returns appear to be non-normal (as evidenced by the Jarque-Bera (J-B) statistic for testing for normality in the returns) but there is no evidence of significant persistent linear dependence in the nominal returns in either subperiod (with the exception of the Norwegian and the American bond returns in the pre-Euro period).

4.2. Correlations

In order to obtain an idea about the degree of interdependence among the bond markets, Table 2 exhibits the contemporaneous correlations for both subperiods for the Euro and non-Euro bond markets separately. A clear result from these correlations is that all bond markets have realized much higher intertemporal dependencies in the post-Euro period. For the Euro bond markets, in particular, this is the result of the unifying effect of the European Monetary Union. However, the correlation coefficients for the non-Euro bond markets are also higher in the post-Euro period (although much lower in magnitude than those of the Euro bond markets). This is the result of the increasing pace of the globalization of national bond markets. Obviously, these (still) low correlations between the European bond markets and the United States indicate the potential of profitable international (i.e, inter-European) diversification for US investors.

In the Euro area, the rise in the correlations of bond market returns in the post-Euro subperiod was accompanied by an increase in their standard deviations (as we saw above). It is not clear whether this increase in the level of risk has any causal relationship with EMU (although it might be indeed difficult not to attribute this to the process of economic and monetary integration culminating in the elimination of currency risk) or is simply a reflection of a wider worldwide trend perhaps as a result of greater global financial integration. Hence, the simple conclusion form the above analysis is that the process of economic and monetary integration within Europe appears to be associated with an increase in the correlations of government bond markets with the implication that the benefits of international diversification using country allocation models within the EMU land have diminished. Finally, a similar process of increasing correlations among the European markets and the US market is evident, suggesting that EMU factors may not be the only ones at play. Overall, however, it appears that some short-run gains from diversifying across some European bond markets still exist. The next section on cointegration, however, will determine whether these short-term correlations are also appropriate indicators of diversification benefits for long-term US investors.

4.3. Unit Root Test Results

Table 3 presents the results of the four unit root tests for the first two subperiods for all bond markets. The boldface numbers below each country row refer to the second subperiod. Appropriate lag lengths were selected according to the Akaike Information Criterion. Moreover, *F*- and *t*- tests were conducted to determine lag lengths and produced very similar results. The unit root tests were corrected for possible serial correlation and autoregressive heteroscedasticity. Finally, the Ljung–Box tests (not shown but are available upon request) on the residuals showed lack of serial correlation in all cases. Based on the results, for all bond markets, the null hypothesis of the presence of a unit root cannot be rejected. Further testing on the first differences of each bond price index series does not indicate the presence of a second unit root, which implies that they are stationary. Therefore, for those countries, individual bond markets appear to be *I(1)*. Because all country stock index levels are *I(1)*, all are likely candidates for cointegrating markets, to which we turn next.

4.4. Multivariate Cointegration Results

Before embarking in the presentation and discussion of the multivariate cointegration results, it is worth noting that we checked also for presence of bivariate cointegration between all Euro bond markets and between the Euro markets and the US and the UK, respectively. Presence of bivariate cointegration was only detected between Finland and the US, and Finland and the UK at the 5% level (all these results can be available upon request).[3] Let us now turn to the multivariate cointegration results which are presented in Table 4.

At the outset, we must report that the five assumptions from these tests included a lag interval (in first differences) from one to four. The first step in the JJ cointegration test is to determine the lag length for the vector autoregression (VAR). Intuitively, one can interpret these cointegration results as follows. If there exist two or more shared common stochastic trends in a given group of countries, then it must be the case that some countries' government bond markets behave independently of the others in the long run. By contrast, if we find only one shared common stochastic trend in a given group, then it would mean that these bond markets have a single common long-run path and any one market may be representative of the behavior of the group. Therefore, an investor should only invest in one of these markets and not in all of them.

From the findings in Table 4, under the five cointegration assumptions outlined earlier, we can see that there are mixed results, as far as the first subperiod is concerned. Specifically, it appears that different cointegrating relationships among the Euro bond markets under these assumptions exist. In the post-Euro period, however, we observe cointegration among these markets (even under three different assumptions about the data). To further shed light on the above results, some exclusion tests were performed in an effort to determine which bond markets were participating in the cointegrating space. The rejection of the null hypothesis of exclusion of a variable from the cointegrating space confirms the presence of close links among the variables in the system. In the event of accepting the null hypothesis, the conclusion would be absence of cointegration among all bond markets or simply absence of close linkages among them. The results from these tests, based on the test statistic distributed as a χ^2 with r degrees of freedom, are reported in the bottom of panel B and are applied to the results from the third assumption of an intercept but no trend in the series. The exclusion tests suggest that Belgium, Finland, Germany, Ireland, and Italy can be excluded from the group since they do not participate in the cointegrating space.

A strong corroboration of the above findings is illustrated in Figure 1, where dynamic cointegration is shown. Specifically, the data are estimated over each year (365 observations in each) from 1994 to 2006 and the two cointegration statistics [equations

(8) and (9)], the maximum eigenvalue and the trace, are plotted in the graph. It is clear from the graph that, based on the number of cointegrating vectors, there is no consistent evidence of cointegration indicating that the bond markets are not in a stable relationship. It is important to remember that in a system of ten variables (markets) full integration would require either one or nine cointegrating vectors. From the graph we see that one cointegrating vector was achieved in 1994, 1999, 2005, and 2006. Thus, despite evidence of cointegration among the bond markets in the post-Euro period, we may consider it as weak and (perhaps) incomplete. We will explore this issue further in subsection 5.

Finally, in panel C of Table 4 evidence of cointegration in the post-Euro period among the non-Euro bond markets is presented and a strong one, we might add, given the fact that cointegration is evident under all five different assumptions about the data. Hence, it seems that these markets share a long-run stable relationship and can be viewed as one asset class, for the purposes of diversification in the (bond) portfolio construction.

4.5. Granger Causality Results

In view of the mixed findings about multivariate cointegration and the finding of no cointegration between the US and the Euro bond markets (evidence not presented but available upon request), it might be possible that short-run relationships among these bond markets still exist. We can employ the Granger causality test with an error-correction term (in the post-Euro period) and without an error-correction term (in the pre-Euro period) to investigate the above possibility. The test was, therefore, applied to the Euro and non-Euro bond markets during the pre- and post-Euro periods, and did the same for the non-Euro bond markets. Since this test is highly sensitive to the lag orders of the right-hand side variables, the Akaike Information Criterion was applied to determine the optimal lag length, which was in all cases two. Finally, we examined the short-run linkages between the US and the other bond markets and between the UK and the other bond markets. These results are found in panels A, B, C, and D of Table 5, respectively.

There several clear-cut conclusions from the results. First, in the post-Euro period we observe a greater number of bivariate linkages among the Euro bond markets relative to the pre-Euro period. However, from these results is not clear as to which bond market emerges as a 'leader', in terms of causing (most of the) other bond markets, and thus this finding may be interpreted as the Euro bond markets having a weak relationship among themselves (a result in line with the above multivariate cointegration results). Second, while we see uni-directional causality between the non-Euro bond group running from the United States to the other markets in the pre-Euro period, in the post-Euro period, we see, additionally, reciprocal linkages among Denmark and Norway. Third, the United States bond market appears to uni-directionally Granger-cause all Euro bond markets in both subperiods. Fourth and final, the United Kingdom's bond market does not surface as having significant Granger-causality influences on the Euro or non-Euro bond markets in either subperiod.

Finally, examining the causality relationship(s) in terms of the long-run equilibrium of two cointegrating variables (i.e., bond markets), via en error-correction term [e_{t-1}, in equation (15)], we conclude the following (see panel A of Table 6). First, some of the error-correction terms are statistically significant at the 10% level, thus confirming the fragile outcomes of multivariate cointegration in panel B of Table 4 and bivariate Granger causality in panel A of Table 5 above. Second, in panel B of Table 6, Wald exclusion tests (for any of the lags in the vector error-correction models) suggest that the first lag (among the optimal three, based on the Akaike Information Criterion), within the Euro bond market group, is the most significant one (in essence, the first lag is highly significant at the 1% level). Finally, just like for the Euro bond markets, for the non-Euro bond markets, between the two optimal lags the first one is the most important one.

5. Discussion and Implications of the Findings

What do the above findings suggest about the linkages among the Euro and non-Euro bond markets since the introduction (and actual use) of the Euro? We will discuss the evidence (and the implications of that evidence) presented in the above tables at length and present some additional data on the degree of integration of these markets over time. We begin with some further analysis of the correlations among these bond markets with particular emphasis on the post-Euro period.

It is worth noting that the rise of correlations in the post-Euro period is not an exclusive trait of the European Union but rather a global phenomenon, as comovement between the Euro and non-Euro bond markets markedly increased as well. Besides, it is inferred that Euro bond market integration process has a large component owing to higher trade and investment linkages among countries. Additionally, the rate of increase in pairwise correlations differs from market to market. For instance, the German-Ireland market pair witnessed a correlation of 0.6789 in the pre-Euro period but a 0.9869 one in the post-Euro period, an increase of 45%! By contrast, the pairs of Austria-Belgium or Portugal-Spain bond markets saw a very small increase in their correlations in the same subperiod. Therefore, it is clear that correlations differ across any two Euro bond markets both during the pre- and post-Euro periods but are definitely higher in the second subperiod. This pattern of change in the pairwise correlations may imply that the introduction of the Euro, and the corresponding reduction in exchange-rate volatility, were important drivers of bond market integration. This is also suggested from research by Bodart and Reding (1999) and Baele et *al*. (2004).

Therefore, the impact of EMU on sovereign bond markets is higher correlations and this will eventually eliminate the cross-market bond differences, which constituted a large part of the investing practices within Euroland in the first place. This could further alter the process by which fixed-income strategies are applied in the Euroland suggesting that is not possible anymore to switch between national government bond markets in order to obtain better returns. Presently, however, Euro government bonds are distinct since differences in issuance and pricing practices still exist. These differences make comparing government bonds difficult and make them less than perfect substitutes. And this, in turn, implies that diversification opportunities are still possible within the Euroland.

How can we further solidify the earlier finding that cointegration among the Euro government bond markets is weak? Figure 2 exhibits the yield differentials between the nine Euro bond markets and German 10-year benchmark government bond from 1999 to 2006 (the post-Euro sample period). There are several comments we can make about these graphs. First, we see a considerable variation in these differentials over this seven and a half-year period. For instance, yield differentials range from 2.1/1.48 basis points for Austria/Portugal in the early 2000s to 0.98 basis points for Portugal in the 2005-6 period. Greater convergence with the German bond yield is detected for Austria, Belgium, Finland, France, and Portugal in the 2005-6 period. Interestingly, the Irish and

Italian bond markets never fully come close to convergence with the German one. So, cross-country yield differences still exist for Euro bond markets even five years after the introduction of the Euro.

Second, not all bond markets exhibit the same convergence path. In other words, while some bond markets (e.g., Belgium, Ireland and Spain) converge with their German counterpart in a declining fashion, others (e.g., France and the Netherlands) converge upwards. Besides, we observe that the absolute values of the yield differentials are not vanishing for most of the Euro bond markets. Pagano and von Thadden (2004) consider two reasons for the presence of non-zero yield spreads namely, fundamental risk factors and residual market frictions.

Third, the poor long-run performance of bond markets within Europe is not indigenous to these markets. The same is also prevalent among the yield differentials of the non-Euro bond markets relative to the US 10-year benchmark government bond, as seen in Figure 3. Put differently, this graph implies that the establishment of the European Monetary Union did not really improve the desired convergence among government bond markets despite the elimination of exchange rate risks and the convergence of inflation expectations across countries although it did improve considerably the convergence of private bond markets (see, for instance, Driessen, 2002, and Pagano and von Thadden, 2004).

The fragility in the linkages among the Euro government bond markets may also be explained by several facts. First, by the fact that some countries, namely Italy, have had several episodes of exchange-rate crises in the 1990s and the capital markets did not believe that the Italian government (as well as the Greek government but it is not examined here) would not align its economic policies with its European partners' policies in a successful and timely manner. Second, the existence of financial barriers to the access of the European bond markets like different taxation structures and institutional features, heterogeneous fiscal policies regarding long-term interest rates and (see, for instance, Clare et *al*., 1995, Barassi et *al.*, 2001, and Yang, 2005). Third, following Pagano and von Thadden (2004), some other factors for the lack of a consistent long-run relationship among the bond markets may be country-specific risk differences, i.e., one country's bonds (like those from Italy, Portugal, and Finland) may be more risky than another country's (such as Germany or France) and liquidity differences across bond markets (which implies that Euro area government bonds are not perfect substitutes).

Overall then, what are the general implications of the above findings for the Euro government bond markets for investors and policymakers? For long-range investors, they imply that (bond portfolio) diversification benefits are still possible within the EMU's government bond market, despite increases in (simple) correlations since 2001. Additionally, better diversification benefits are achieved when pairing Euro market bonds with non-Euro government (and corporate) bonds due to the existence of persistent (small yet variable) yield differentials for government debt, again despite increases in Euro and non-Euro bond market correlations. However, this benefit should be gradually deteriorating as the integration process keeps increasing within the Euroland leading to a decrease in this 'home bias' phenomenon (see also Laopodis, 2005).

For European policymakers, higher correlations mean greater shock transmission avenues within the Euroland, with possible adverse consequences for the stability of the European monetary system. This further implies that the European Central Bank will have a more difficult task of controlling/setting the land's monetary policy (as it attempts to influence long-term interest rates) and thus achieve its goal of price stability. Finally, greater bond market integration may have important consequences for fiscal discipline within EMU (see Barr and Priestley, 2004, and Detken et *al*., 2004). Specifically, fiscal discipline is loose within the union as a specific government can draw on the common pool of the union's savings. Therefore, because there is an incentive to abuse such savings, the case for a more coordinated and disciplined fiscal policy within EMU is much more imperative than for governments at the global level.

6. Summary and Conclusions

This paper examines the degree of integration/interdependence in the European government bond market both before and after the introduction of the Euro for ten Euro markets and three non-Euro markets for the 1994-2001 period. Multivariate cointegration analyses indicate absence of cointegration among the Euro bond markets in the pre-Euro period but a weak one in the post-Euro period. By contrast, evidence of strong cointegration in the post-Euro period among the non-Euro bond markets is detected. Additional results pertain to the Granger causality results and their inferences. First, in the post-Euro period we detect a higher frequency of bivariate linkages among all of the Euro bond markets relative to the pre-Euro period. Second, the United States bond market appears to uni-directionally Granger-cause all Euro bond markets in both subperiods. Finally, the United Kingdom's bond market does not surface as having significant Granger-causality influences on the Euro or non-Euro bond markets in either subperiod.

These findings have important implications for investors and policymakers alike. For international investors, an understanding of the extent of comovement among Euro area bond markets is essential for the structure and adjustment of their international portfolios. Implications also exist for portfolio diversification opportunities. Specifically, these may have considerably been diminished if higher bond integration is present and sovereign bond issues are becoming closer substitutes with a wider consequence of the depreciation of the Euro. For European policymakers, knowledge of the extent of integration among sovereign bond markets is significant for the conduct of the common monetary policy. In other words, greater interdependence of bond yields among European countries may reduce the capability of the European Central Bank to influence long-term interest rates and thus attain price stability.

In conclusion, the effects of the EMU, various institutional changes and market actions led to a substantial convergence in sovereign Euro bond yields since 2001. Although yield differences are still present across Euro government bond issues, future developments/ advances both in the public and private bond market will result in greater harmonization of bond yields with further reductions in diversification benefits for global investors. Moreover, the Euro bond area will keep expanding as new (Eastern) European countries join the EMU. Although bonds of these countries are already actively traded, the degree of integration among these new markets is incomplete and thus bonds by these markets still offer significant diversification opportunities for global investors.

Endnotes

1. The terms associated with $\alpha_\perp$ are deterministic trends 'outside' the cointegrating relations. When a deterministic trend appears both inside and outside the cointegrating relationship and the decomposition is not uniquely identified, Johansen identifies the part that belongs inside the error-correction term by orthogonally projecting the exogenous terms on the α space so that $\alpha_\perp$ is the null space of α such that $\alpha'\alpha_\perp = 0$.

2. The sample period was also split in 12/31/1998, to reflect the theoretical Euro introduction period in January 1999, but there was no qualitative change in the post-Euro period results.

3. We also investigated the possibility of bivariate cointegration between Germany and the other bond markets but there was no cointegration suggested in any pair.

References

Adam, K., Jappelli, T., Menichini, A., Padula, M., Pagano, M., 2002. Analyse, compare, and apply alternative indicators and monitoring methodologies to measure the evolution of capital market integration in the European Union. Report to the European Commission.

Barassi, M.R., Caporale, G.M., Hall, S.G., 2001. Irreducibility and structural cointegrating relations: an application to the G-7 long-term interest rates. International Journal of Finance and Economics 6, 127-138.

Barr, D.G., Priestley, R., 2004. Expected returns, risk and the integration of international bond markets. Journal of International Money and Finance 23, 71-97.

Beale, L., Ferrando, A., Hördhal, P., Krylova, E., Monnet, C., 2004. Measuring financial integration in the Euro area. European Central Bank Occassional Paper Series No. 14, April.

Bekaert, G., Harvey, C.R., 1997. Emerging equity market volatility. Journal of Financial Economics 43, 29-77.

Berben, R.P., Jansen, W.J., 2005. Bond market and stock market integration in Europe. Working paper, De Nederlandsche bank, Amsterdam, Netherlands, November.

Bernoth, K., von Hagen, J., Schnuknecht, L., 2004.Sovereign risk premia in the European government bond market, European Central Bank Working Paper No. 369, June.

Bessler, D.A., Yang, J., 2003. The structure of interdependence in international stock markets. Journal of International Money and Finance 22, 261-287.

Bodart, V., Reding, P., 1999. Exchange rate regime, volatility and international correlations on bond and stock markets. Journal of International Money and Finance 18, 133-151.

Cappiello, L., Engle, R.F., Sheppard, K., 2003. Asymmetric dynamics in the correlations of global equity and bond returns. European Central Bank Working Paper No. 204, Frankfurt am Main.

Christiansen, C., 2003. Volatility-spillover effects in European bond markets. Centre for Analytical Finance, University of Aarhus Working Paper No. 162, Denmark, November.

Clare, A., Maras, M. Thomas, S., 1995. The integration and efficiency of international bond markets. Journal of Buisness Finance and Accounting 22, 313-322.

Clare, A., Lekkos, I., 2000. An analysis of the relationship between national bond markets. Working Paper no. 123, Bank of England.

Côté, D., Graham, C., 2004. Convergence of government bond yields in the Euro zone: the role of policy harmonization. Bank of Canada Working Paper 2004-23.

Detken, G, Gaspar, V., and Winkler, B., 2004. On prosperity and posterity: the need for fiscal discipline in a monetary union. European Central Bank Working Paper no. 420, December.

Driessen, J., 2002. I sdefault risk priced in corporate bonds? Discussion paper, Unversity of Amsterdam, Netherlands.

Driessen, J., Melenberg, D., Nijman, T., 2003. Common factors in international bond returns. Journal of International Money and Finance 22, 629-656.

European Central Bank, 2004. The Euro bond market study. December.

Fratzscher, M., 2002. Financial market integration in Europe: on the effects of EMU on stock markets. International Journal of Finance and Economics 7, 165-193.

Holder, M., 1999. The Euro impact on European financial markets. Managerial Finance 25(1), 27-34.

Ihmanen, A., 1995. Time varying expected returns in international bond markets. Journal of Finance 50(2), 481-506.

Johansen, S., 1988. Statistical analysis of cointegration vectors. Journal of Economic Dynamics and Control 12, 251-254.

Johansen, S., Juselius, K., 1990. Maximum likelihood estimation and inference of cointegration with application to the demand for money. Oxford Bulletin of Economics and Statistics 52, 169-209.

Kwiatkowski, D.P., Phillips, P.C.B., Schmidt, P., Shin, Y., 1992. Testing the null hypothesis of stationarity against the alternative of unit root: how sure are we that economic times series have a unit root? *Journal of Econometrics* 54, 159-178.

Laopodis, N.T., 2005. Portfolio diversification benefits within Europe: implications for the US investor. International Review of Financial Analysis 14, 455-476.

McCauley, R.N., White, W.R., 1997. The Euro and European financial markets. Board of Governors of the Federal Reserve System Working Paper, May.

Ng, A., 2000. Volatility spillover effects from Japan and the US to the Pacific Basin. Journal of International Money and Finance 19, 207-233.

Pagano, M., von Thadden, E.L., 2004. The European bond markets under EMU. Oxford Review of Economic Policy 20(4), 531-554.

Portes, R., Rey, H., 1998. The emergence of the Euro as an international currency. Economic Policy, April, 305-332.

Skintzi, V. D., Refenes, A. N., 2006. Volatility spillovers and dynamic correlation in European bond markets. Journal of International Financial Markets, Institutions and Money 16, 23-40.

Yang, J., 2005. Government bond market linkages: evidence from Europe. Applied Financial Economics 15, 599-610.

Table 1. Descriptive Statistics and Correlations on Daily Nominal Bond Index Returns, Pre- and Post Euro Periods

Bond Index Return	Mean	St. Dev.	Skewness	Kurtosis	J-B (prob.)	LB(5)
Pre-Euro period, 1/1/1994 – 12/31/2000						
Austria	0.0135	0.6418	0.3965	5.3901**	481.01 (0.000)	3.657
Belgium	0.0161	0.6249	0.4474	5.8775**	674.00 (0.000)	3.114
Denmark	0.0185	0.6023	0.3150	5.1397**	551.01 (0.000)	3.756
Finland	0.0126	0.6423	0.1655	13.3378**	775.34 (0.000)	3.567
France	0.0147	0.6208	0.4321	5.3700**	670.80 (0.000)	3.556
Germany	0.0127	0.6377	0.3789	5.4322**	747.75 (0.000)	3.678
Ireland	0.0193	0.6005	0.2184	5.6432**	586.00 (0.000)	3.643
Italy	0.0246	0.6438	0.1312	5.4623**	452.13 (0.000)	3.664
Netherlands	0.0127	0.6455	0.3886	5.3316**	692.56 (0.000)	3.562
Norway	0.0126	0.6423	0.1655	13.2634**	876.34 (0.000)	15.6**
Portugal	0.0228	0.6107	0.3869	5.4021**	278.84 (0.000)	3.432
Spain	0.0201	0.6201	0.2119	5.3830**	442.82 (0.000)	3.334
United Kingdom	0.0302	0.5684	-0.3058	4.6543**	352.57 (0.000)	3.667
United States	0.0247	0.2672	-0.3050	5.0022**	361.23 (0.000)	9.62**
Post-Euro period Descriptives, 1/1/2001 – 7/27/2006						
Austria	0.0415	0.7008	-0.0865	3.5431**	18.001 (0.000)	4.657
Belgium	0.0421	0.7019	-0.0934	3.5325**	19.600 (0.000)	4.544
Denmark	0.0417	0.6783	-0.0730	3.5327**	18.301 (0.000)	4.786
Finland	0.0483	0.6933	-0.2055	3.7078**	40.534 (0.000)	4.867
France	0.0407	0.6998	-0.0981	3.5320**	19.080 (0.000)	4.654
Germany	0.0403	0.6987	-0.0959	3.5232**	19.115 (0.000)	4.658
Ireland	0.0433	0.7225	-0.1114	3.4992**	17.800 (0.000)	4.665
Italy	0.0426	0.7028	-0.0982	3.5433**	19.615 (0.000)	4.456
Netherlands	0.0407	0.6965	-0.0876	3.4566**	19.346 (0.000)	4.432
Norway	0.0483	0.6922	-0.2033	3.7034**	40.314 (0.000)	4.860
Portugal	0.0422	0.6935	-0.0861	3.5321**	20.354 (0.000)	4.654
Spain	0.0421	0.7001	-0.0969	3.5230**	18.902 (0.000)	4.554
United Kingdom	0.0341	0.6364	-0.0548	3.4893**	14.767 (0.000)	4.543
United States	0.0177	0.3062	-0.3761	4.4362**	145.23 (0.000)	4.327

Notes : ** means statistical significance at the 5% level ; J-B is the Jarque-Bera statistic for normality in the return series; asymptotic p-values are in parentheses; LB(5) is the Ljung-Box Q-statistic for serial correlation in the return series for up to 5 lags.

Table 2. Correlations of Nominal Bond Index Returns, Pre- and Post-Euro periods

Euroland Markets: **Pre-Euro period:** 1/1/1994 – 12/31/2000

	Aus	Bel	Fin	Ger	Ire	Ita	Net	Por	Spa
Aus	1								
Bel	0.9971	1							
Fin	0.9108	0.9263	1						
Ger	0.8412	0.8303	0.6516	1					
Ire	0.8571	0.8843	0.8146	0.6789	1				
Ita	0.6088	0.6525	0.6883	0.5092	0.9101	1			
Net	0.9984	0.9968	0.8131	0.8574	0.8603	0.6172	1		
Por	0.9053	0.9277	0.8313	0.6995	0.9869	0.8686	0.9062	1	
Spa	0.8753	0.9010	0.7988	0.6992	0.9876	0.9062	0.8791	0.9919	1
Fra	0.9662	0.9777	0.8268	0.8111	0.9513	0.7855	0.9685	0.9771	0.9665

	Aus	Bel	Den	Fin	Fra	Ger	Ire	Ita	Net	Nor	Por	Spa	UK
US	0.089	0.122	0.103	0.031	0.135	0.132	0.145	0.110	0.152	0.031	0.074	0.141	0.270

Post-Euro period: 1/1/2001 – 7/27/2006

	Aus	Bel	Fin	Ger	Ire	Ita	Net	Por	Spa
Aus	1								
Bel	0.9979	1							
Fin	0.9908	0.9963	1						
Ger	0.9717	0.9603	0.8416	1					
Ire	0.9976	0.9943	0.8946	0.9869	1				
Ita	0.9985	0.9925	0.8983	0.9792	0.9901	1			
Net	0.9964	0.9998	0.8931	0.9774	0.9993	0.9972	1		
Por	0.9953	0.9977	0.8913	0.9695	0.9999	0.9986	0.9962	1	
Spa	0.9952	0.9910	0.8988	0.9792	0.9996	0.9962	0.9991	0.9969	1
Fra	0.9961	0.9987	0.8968	0.9761	0.9993	0.9955	0.9985	0.9981	0.9995

	Aus	Bel	Den	Fin	Fra	Ger	Ire	Ita	Net	Nor	Por	Spa	UK
US	0.367	0.369	0.342	0.268	0.368	0.368	0.396	0.369	0.365	0.268	0.362	0.371	0.386

Non-Euroland Markets:

	Pre-Euro period			**Post-Euro period**		
	Nor	UK	US	Nor	UK	US
Den	0.7639	0.5247	0.1031	0.8447	0.8062	0.3288
Nor		0.4108	0.0323		0.6919	0.2687
UK			0.2770			0.3865

Table 3. Unit Root Test Results for Euro and non-Euro Bond Markets

Bond market	LM statistic	Significance Levels 1%	5%	10%
Austria	1.5191	0.7390	0.4630	0.3470
	4.3735			
Belgium	1.9445	0.7390	0.4630	0.3470
	4.3651			
Finland	2.2078	0.7390	0.4630	0.3470
	4.3323			
France	2.5978	0.7390	0.4630	0.3470
	4.4352			
Germany	1.3444	0.7390	0.4630	0.3470
	4.2435			
Ireland	3.6545	0.7390	0.4630	0.3470
	4.4436			
Italy	4.1432	0.7390	0.4630	0.3470
	4.3432			
Netherlands	1.6423	0.7390	0.4630	0.3470
	4.3343			
Portugal	3.3478	0.7390	0.4630	0.3470
	4.2345			
Spain	3.5734	0.7390	0.4630	0.3470
	4.5564			
Denmark	2.8656	0.7390	0.4630	0.3470
	4.3544			
Norway	2.2078	0.7390	0.4630	0.3470
	4.4409			
UK	5.2234	0.7390	0.4630	0.3470
	4.4989			
US	5.3322	0.7390	0.4630	0.3470
	4.1978			

Notes: all statistic values significant at the 5% level; bold numbers refer to the post-Euro period (1/1/2001-7/27/2006); the Lagrangian Multiplier (LM) is the KPSS statistic for testing for unit root with the null hypothesis that a bond market is stationary; LM statistic estimated using the Newey-West approach.

Table 4. Multivariate Cointegration Results and Exclusion Tests, Pre- and Post-Euro Periods

Panel A: Cointegration Among all Euro Bond Markets, *Pre-Euro Period*

Data Trend	None	None	Linear	Linear	Quadratic
Rank of CE(s) Number of CE(s)	No Intercept No Trend	Intercept No Trend	Intercept No Trend	Intercept Trend	Intercept Trend
Statistic	*Number of cointegrating equations by model (columns)*				
Trace	3	4	5	3	3
Max-Eigenvalue	3	2	2	2	2

Panel B: Cointegration Among all Euro Bond Markets, *Post-Euro Period*

Data Trend	None	None	Linear	Linear	Quadratic
Rank of CE(s) Number of CE(s)	No Intercept No Trend	Intercept No Trend	Intercept No Trend	Intercept Trend	Intercept Trend
Statistic	*Number of cointegrating equations by model (columns)*				
Trace	2	1	**1**	1	1
Max-Eigenvalue	2	2	**1**	1	1

Exclusion Tests [$\chi^2(2)$]

Austria: 79.4122*** (0.0000); Belgium: 24.8563 (0.0000); Finland: 0.0946 (0.7656); France: 6.8323** (0.0086); Germany: 0.4152 (0.5123); Ireland: 0.3211 (0.2324); Italy: 1.3667 (0.2068); Netherlands: 20.6546*** (0.0000); Portugal: 15.4323** (0.0006); Spain: 42.2642*** (0.0000)

Table 4. Multivariate Cointegration Results and Exclusion Tests, Pre- and Post-Euro Periods (concl'd)

Panel C: Cointegration Among non-Euro Bond Markets, *Pre-Euro Period*

Data Trend	None	None	Linear	Linear	Quadratic
Rank of CE(s) Number of CE(s)	No Intercept No Trend	Intercept No Trend	Intercept No Trend	Intercept Trend	Intercept Trend
Statistic	*Number of cointegrating equations by model (columns)*				
Trace	0	0	**0**	0	0
Max-Eigenvalue	0	0	**0**	0	0
			Post-Euro Period		
Trace	1	1	**1**	1	1
Max-Eigenvalue	1	1	**1**	1	1

Notes: Probability values in parentheses; *, **, *** refer to 10%, 5%, and 1% levels of significance, respectively; pre-Euro period: 1/1/1994 – 12/31/2000; post-Euro period: 1/1/2001 – 7/27/2006; bold numbers correspond to cointegration assumption number 3.

Table 5. Bivariate Granger Causality Tests for Euro and non-Euro Bond Markets

Panel A: Euro Bond Markets *Pre-Euro Period*: 1/1/1994 -12/31/2000

Null Hypothesis	F-Statistic	Probability
Belgium does not Granger Cause Austria	6.2454	9.84e-06***
France does not Granger Cause Austria	5.3773	6.69e-05***
Germany does not Granger Cause Austria	4.1956	0.0008***
Germany does not Granger Cause Italy	4.0189	0.0007***
Germany does not Granger Cause Spain	3.9089	0.0009***
Austria does not Granger Cause Italy	3.9913	0.0013**
Netherlands does not Granger Cause Austria	10.820	3.27e-10***
France does not Granger Cause Portugal	4.9328	0.0001***
Portugal does not Granger Cause Italy	3.2359	0.0065**
Spain does not Granger Cause Portugal	5.4008	6.36e-05***
Portugal does not Granger Cause Spain	3.7040	0.0024**

Post-Euro period: 1/1/2001 – 7/27/2006

Finland does not Granger Cause Austria	4.0840	0.0010**
Austria does not Granger Cause Finland	3.7007	0.0024**
Germany does not Granger Cause Austria	4.8425	0.0002***
Austria does not Granger Cause Germany	25.745	0.0000***
Belgium does not Granger Cause Finland	3.8967	0.0016**
Germany does not Granger Cause Belgium	4.7502	0.0002***
Belgium does not Granger Cause Germany	22.292	0.0000***
France does not Granger Cause Finland	3.9416	0.0014**
Finland does not Granger Cause France	4.3771	0.0005***
Germany does not Granger Cause Finland	5.8356	2.33e-05***
Finland does not Granger Cause Germany	8.1057	1.39e-07***
Ireland does not Granger Cause Finland	3.8308	0.0018**
Finland does not Granger Cause Ireland	4.2913	0.0006***
Finland does not Granger Cause Italy	5.9437	1.83e-05***
Italy does not Granger cause Finland	4.6278	0.0002***
Netherlands does not Granger Cause Finland	3.9375	0.0014**
Portugal does not Granger Cause Finland	3.9074	0.0015**
Spain does not Granger Cause Finland	3.7566	0.0021**
Finland does not Granger Cause Spain	4.3952	0.0005***

Table 5. Bivariate Granger Causality Tests for Euro and non-Euro Bond Markets (concl'd)

Panel B: Non-Euro Bond Markets	*Pre-Euro Period*: 1/1/1994 -12/31/2000	
US does not Granger Cause Denmark	5.8050	2.61e-05***
US does not Granger Cause UK	5.1264	0.0001***
US does not Granger Cause Norway	5.8760	2.23e-05***
	Post-Euro period: 1/1/2001 – 7/27/2006	
US does not Granger Cause Denmark	13.270	9.68e-13***
US does not Granger Cause UK	10.629	4.29e-10***
US does not Granger Cause Norway	10.790	2.96e-10***
Norway not Granger Cause Denmark	3.8391	0.0018**
Denmark does not Granger Cause Norway	4.1825	0.0008***
Norway does not Granger Cause UK	6.0200	1.54e-05***

Panel C: Euro bond markets with the US	*Pre-Euro Period*: 1/1/1994 -12/31/2000	
US does not Granger Cause Austria	6.8610	2.36e-06***
US does not Granger Cause Belgium	4.9001	0.00011***
US does not Granger Cause Finland	7.6819	3.70e-07***
US does not Granger Cause France	6.0096	1.59e-05***
US does not Granger Cause Germany	3.5606	0.0033*
US does not Granger Cause Ireland	8.9198	2.21e-08***
US does not Granger Cause Italy	6.7575	2.98e-06***
US does not Granger Cause Netherlands	4.6890	0.0002**
US does not Granger Cause Portugal	5.4206	5.89e-05***
US does not Granger Cause Spain	4.5800	0.00037**
Post-Euro period: 1/1/2001 – 7/27/2006		
US does not Granger Cause Austria	10.5410	5.16e-09***
US does not Granger Cause Belgium	10.4301	5.09e-09***
US does not Granger Cause Finland	8.3319	7.50e-09***
US does not Granger Cause France	10.2196	8.49e-09***
US does not Granger Cause Germany	10.7406	7.60e-08***
US does not Granger Cause Ireland	10.2693	4.43e-07***
US does not Granger Cause Italy	10.7245	5.43e-08***
US does not Granger Cause Netherlands	10.3810	6.43e-07***
US does not Granger Cause Portugal	9.8806	7.43e-09***
US does not Granger Cause Spain	10.6830	8.45e-09***

Panel D: Euro bond markets with the UK	*Post-Euro period*: 1/1/2001 – 7/27/2006	
Finland does not Granger Cause UK	5.8260	2.45e-05***

Notes: lags, 5; *, **,*** denote significance at the 10%, 5% and 1% levels, respectively.

Table 6. Multivariate Error-Correction and Lag-Exclusion Test Results

Panel A: Euro and non-Euro Bond Markets, *Post-Euro Period*

Bond Market	Error-Correction Term, e_{t-1}, coefficient	T-ratio
Austria	-0.01262*	-2.0345
Belgium	-0.01471	-1.7637
Finland	0.00112	1.3213
France	-0.01481*	-1.8189
Germany	-0.01432	-1.7189
Ireland	-0.01445	-1.6819
Italy	-0.01482	-1.5431
Netherlands	-0.01341*	-1.8321
Portugal	-0.01412*	-1.8556
Spain	-0.01554*	-1.7890
US	-0.00481**	-2.5617
UK	-0.01832***	-4.8728
Denmark	-0.01824***	-4.5678
Norway	-0.00189*	-1.9819

Panel B: Wald Lag-Exclusion Test Results

	χ^2 statistic tests at Lag			Probability Values at Lag		
	1	2	3	1	2	3
Austria	1313.4***	13.7723	11.1867	0.0000	0.1833	0.3431
Belgium	1311.6***	13.8754	10.9567	0.0000	0.1788	0.3609
France	1440.3***	19.0451	19.8434	0.0000	0.0398	0.0308
Finland	1310.4***	13.5326	10.5878	0.0000	0.1951	0.3854
Germany	1312.4***	13.9901	10.7718	0.0000	0.1766	0.3761
Ireland	1330.2***	7.9514	12.7718	0.0000	0.6342	0.2677
Italy	1299.1***	13.9451	11.7018	0.0000	0.1756	0.3566
Netherlands	1316.8***	13.7718	11.0324	0.0000	0.1860	0.3548
Portugal	1323.9***	13.7654	10.4977	0.0000	0.1839	0.3983
Spain	1310.8***	13.5327	10.5877	0.0000	0.1978	0.3567
US	18.531***	2.1376		0.0009	0.7104	
UK	45.017***	5.7534		3.9E-09	0.2186	
Denmark	58.683***	2.6872		5.4E-12	0.6114	
Norway	47.023***	9.7373		1.5E-09	0.0455	

Notes: *,**,*** refer to 10%, 5%, and 1% levels of significance, respectively; lags represent the optimal lag length based on the Akaike Information Criterion.

Figure 1. Trace and Maximum-Eigenvalue Statistics for Euro Bond Markets, 1/1/1994 – 7/27/2006

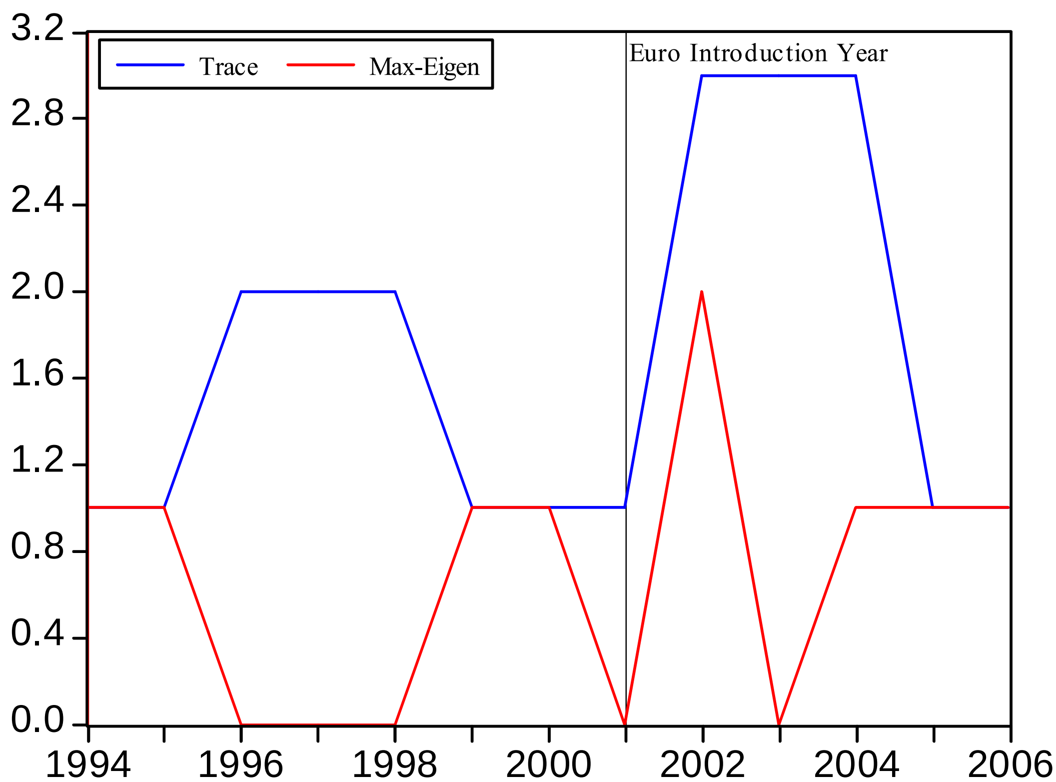

Figure 2. Yield Differentials relative to the German 10-yr benchmark government bond, 1999-2006

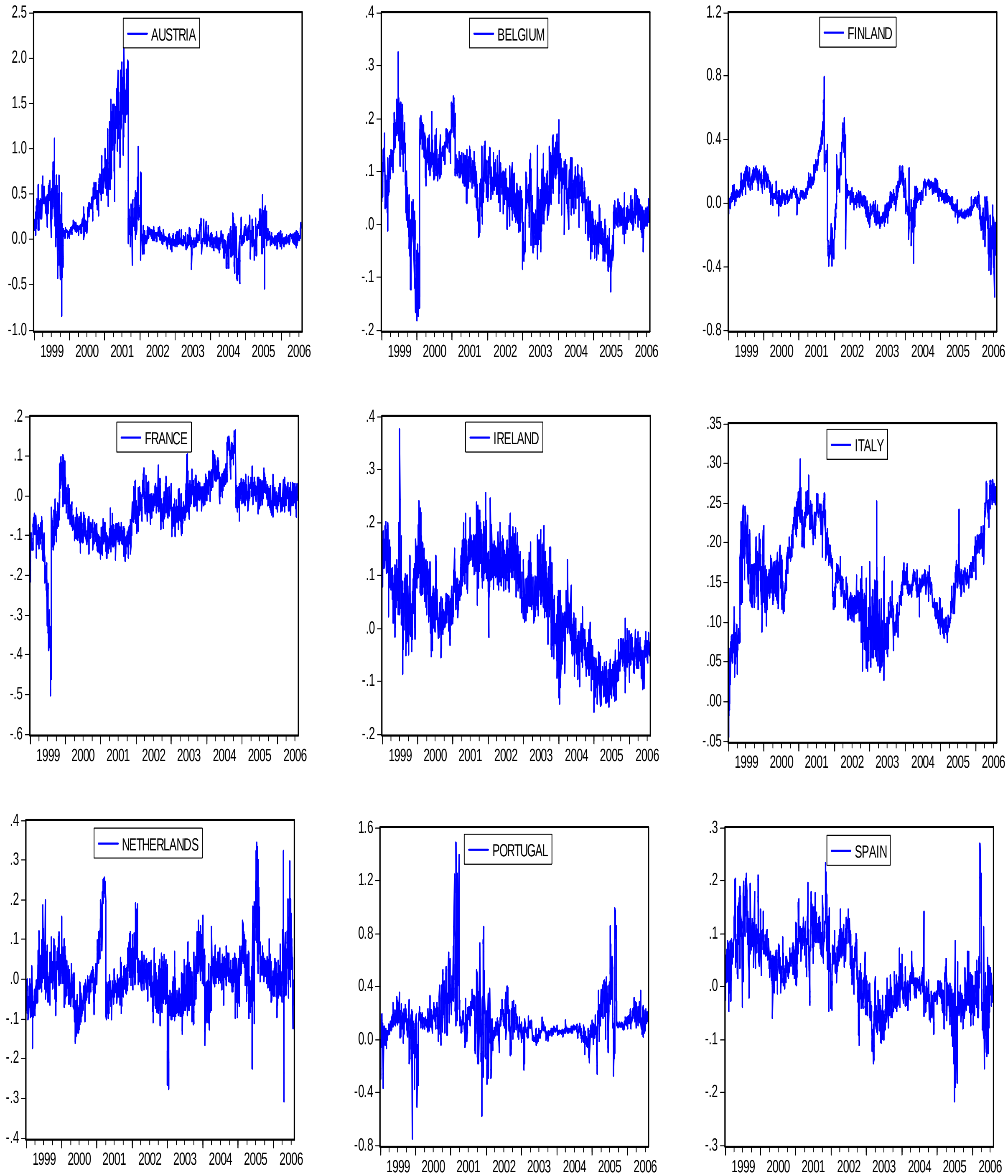

Figure 3. Yield differentials relative to the US 10-yr benchmark government bond, 1999-2006