THE EURO: INTEGRATION AND RISK DIVERSIFICATION OPPORTUNITIES IN EU-15 SOVEREIGN DEBT MARKET *

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Abstract

The objective of this paper is twofold. First we analyze the impact of the Euro on EU-15 sovereign debt market integration (both with world and Eurozone) by means of using a CAPM-based model. Results show that whilst non-EMU markets present a higher vulnerability to world risk factors, EMU markets are more vulnerable to Eurozone factors. However, they are only partially integrated with the German markets. For this reason, the second objective is to apply a cointegration analysis to examine portfolio diversification opportunities in public debt markets after the Euro. Results reject the existence of a unique common trend among the EMU yields, suggesting the possibility of risk diversification across the Eurozone. These findings have important implications for investors, in terms of diversification benefits in a context of a single currency.

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1. Introduction

The market capitalization of international bond markets is much larger than that of international equity markets. However, compared to the large body of literature on international equity market linkages (see Bessler and Yang, 2003), there are far fewer empirical studies of bond systemic risk or international bond market linkages (Smith, 2002 and Barr and Priestley, 2004). Moreover, little has been written on the sources of co-movements in bond markets in the European context. Nevertheless, the extent of international and European bond market linkages merits investigation, as it may have important implications for the cost of financing fiscal deficit, monetary policymaking independence, modeling and forecasting long-term interest rates, and bond portfolio diversification.

Conversely, more has been written on emerging countries and on volatility spillovers in international bond markets (see Cappiello *et al.* (2003), Christiansen (2003), or Skintzi and Refenes (2006), among others). In the context of emerging economies, a very important question in the study of spread co-movements is the analysis of the relative influence of fundamental variables on that behavior (see Cifarelli and Paladino, 2006). Economies are related through trade and financial flows, and shifts in the economic fundamentals of one country may affect its neighbors. However, in periods of growing uncertainty, changes in market sentiment may go beyond fundamentals and generate "contagion phenomena". Calvo and Reinhart (1996), Masson (1999) and, more recently, Bekaert, Harvey and Ng (2005) draw a distinction between fundamentals-based contagion, which arises when the "infected country" is connected to others via trade and/or financial links, and pure contagion, which is due to a shift in market sentiment without (or beyond) links in economic fundamentals.

Therefore, even though there is far more literature on emerging than on developed economies, it is well established in both cases that bond markets in different countries tend to move together, i.e. bond prices and returns are positively correlated across countries. Some of the early attempts to investigate this issue are Clare et *al.* (1995) who provided insights into the significance of international bond market linkages for bond portfolio diversification, and Ilmanen (1995), whose evidence suggested that excess returns of long-term international bonds were highly correlated, implying, in turn, international bond market integration.

More recently, using a bivariate conditional correlation GARCH model, Hunter and Simon (2005) examine the lead-lag and contemporaneous relationships between 10-year US government bond returns and 10-year UK, German, and Japanese government bond returns. They find that the mean and volatility of the US bond market lead the mean and volatility of both German and Japanese bond markets, whilst there is no significant lead-lag relation between the US and UK bond markets. Volatility spillover from world bond markets into individual bond markets has also focused researchers' attention. In particular, Christiansen (2007) finds strong statistical evidence of volatility spillover from the US and aggregate European bond markets into individual European bond markets. For EMU countries, the US volatility spillover effects are rather weak (in economic terms) whereas the European volatility-spillover effects are strong.

Little has been written on the sources of co-movements in Government bond markets in the European context. Studies of this issue include Geyer, Kossmeier and Pischer (2004), Pagano and Von Thadden (2004), and Gómez-Puig (2009a and 2009b).

Geyer et *al*. estimate a multi-issuer state-space version of the Cox-Ingersoll-Ross (1985) model of the evolution of bond-yield spreads (over Germany) for four EMU countries (Austria, Belgium, Italy and Spain). Their main findings are (i) one single ("global") factor explains a large part of the movement of all four processes, (ii) idiosyncratic country factors have almost no explanatory power, and (iii) to a limited extent, the variation in the single global factor can be explained by EMU corporate-bond risk, but by nothing else. The most striking finding in Geyer et *al*. (2004) is the virtual absence of country-specific yield-spread risk. On the other hand, despite the considerable differences in the methodology and data used, Pagano and von Thadden (2004) also agree that yield differentials under EMU are driven mainly by a common risk (default) factor and suggest that liquidity differences have at best a minor role in the time-series behavior of yield spreads.

Gómez-Puig (2009a and 2009b) estimates panel regressions for two groups of EU-15 countries (EMU and non-EMU) including both domestic (differences in market liquidity and credit risk) and international risk factors. Concretely, Gómez-Puig (2009a) analyses the immediate effect of Monetary Union on EU-15 Sovereign yield spreads during the period 1996-2001 (i.e., three years before and three years after the introduction of the euro), whilst Gómez-Puig (2009b) studies the effects of Monetary Union on the different factors that influence sovereign yield spreads in EU-15 countries using a sample that covers the period from January 1999 until December 2005, i.e. the first seven years since the introduction of the common currency. Her results present evidence that it is domestic rather than international risk factors that mostly drive the evolution of 10-year yield spread differentials over Germany in all EMU after the beginning of Monetary Integration. Conversely, in the case of non-EMU countries, adjusted yield spreads are influenced more by world risk factors. The fact that these countries do not share a common Monetary Policy might explain these results, which may also show that government bonds from EMU countries have a better safe-haven status than those of non-EMU countries. Moreover, she concludes that with the introduction of a common currency and in the current context of higher competition between Euro-area government securities markets, the success of these sovereign securities debt markets may be highly dependent on their market size (the removal of the exchange rate barrier seems to have punished Euro-area smaller markets twice, by making them pay both higher liquidity and a higher default risk premium than larger ones). In the case of non-Euro participating countries since they did not suffer an increase in their degree of substitutability and competition, they may have benefited from the fact that market participants consider their risk premium to be low and the investment advantages to be high.

The abovementioned results are consistent with the empirical evidence presented by other authors like Cappiello, Hördahl, Kadareja, and Manganelli (2006), who used a completely different methodology to investigate whether the introduction of the euro had an impact on the degree of integration of European financial markets. Controlling for the impact of global factors, they document an overall increase in co-movements in euro area financial markets, especially in bond markets, suggesting that integration in the euro area has progressed since the introduction of the single currency. In contrast to previous studies, they propose two methodologies to measure integration: one that relies on time-varying GARCH correlations, and the other on a regression quantile-based co-dependence measure (see Cappiello, Gérard, Kadareja and Manganelli, 2005).

Finally, a number of papers have studied financial integration exploiting the implications of asset pricing models. The works by Barr and Priestley (2004) and Hardouvelis, Malliaropulos and Priestley (2006 and 2007) are in this vein. In particular, Barr and Priestley (2004) use a version of Bekaert and Harvey's (1995) CAPM-based model to analyze the degree of integration of the US, UK, Japan, Germany and Canada bond markets, and find strong evidence that national markets are only partially integrated into world markets. Around one quarter of total expected excess returns is related to local market risk, the remainder being due to world bond market risk. A similar methodology is used by Hardouvelis *et al.* (2006 and 2007) to analyze the impact of EMU on European stock market integration. They present evidence linking the process of increased integration of European stock markets to the prospects of the formation of EMU and the adoption of the euro as the single currency. Specifically, these authors show that in the second half of the 1990s, expected stock returns in Europe became increasingly more determined by EU market risk and less by local risk. However, this methodology has not yet been used to study bond markets integration in the European context.

In this scenario, the objective of this paper is twofold. First, we will study financial integration in EU-15 sovereign bond markets, exploiting the implications of asset pricing models. The final goal of the first analysis of the paper is to examine whether participation in the Monetary Union is an important factor which determines the differences in the impact of two sources of systemic risk (world and regional risk) on each EU-15 Government bond market and to establish whether the degree of integration with the world and the regional debt markets has differed within EMU and non EMU-countries and across the different countries of each group since 1999. Secondly, we will analyze the evolution of risk diversification opportunities in EU-15 countries since the introduction of the euro by means of a cointegration analysis. Both objectives are highly related and their relationship will be examined in Section 2. Section 3 is focused in the study of the first objective. The second objective is developed in Section 4. Finally, Section 5 draws the main conclusions of all the analysis.

2. Objectives

More than ten years after the introduction of the euro, the first objective of this paper is to study whether the introduction of the euro had an impact on the degree of integration of European Government bond markets. Therefore, we will carry out a comparative analysis of the degree of integration of Government bond markets in two groups of EU-15 countries: those that joined the European Monetary Union (EMU) and those that stayed out of it. Our sample will span the period since the beginning of Currency Union until June 2008. Our intention is to separate each individual country's Government bond return into three effects: a local (own country) effect, a regional (Eurozone) effect, and a global (world) effect, and to establish whether there are significant differences within EMU and non-EMU participating countries. That is., we analyze whether participation in the Monetary Union is an important factor which determines the differences in the impact of world and regional risk on each EU-15 Government bond market.

To examine financial integration, we exploit the implications of asset pricing models. In particular, following Barr and Priestley (2004) who assess the degree of integration of the US, UK, Japan, Germany and Canada bond markets, we adopt Bekaert and Harvey's CAPM-based model (1995). This model allows partially integrated markets and still has not been used to study bond markets integration in the European context. Moreover, it has only been used to analyze the impact of one kind of common or systemic risk factor over bond or stock returns behavior (see Hardouvelis, Malliaropulos and Priestley (2006 and 2007)). So, the model used in this paper draws on Barr and Priestley (2004), but goes beyond it. As far as we know, this is the first empirical study that applies this methodology to analyze the impact of the euro on European Government bond markets integration with a weekly dataset that covers almost ten years since the introduction of the common currency.

As it will be presented in Section 3, our empirical evidence suggests that the impact of the introduction of the euro on the degree of integration of European Government bond markets was important. The markets of the countries that decided to stay out of the Monetary Union present a higher vulnerability to external risk factors. For their part, the countries that share a monetary policy are less vulnerable to the influence of world risk factors, and more vulnerable to EMU risk factors. However, euro markets are only partially integrated with the German market since their markets are still segmented and present differences in their market liquidity or default risk. In a different context, Laopodis (2008) reached the same conclusion, suggesting that benefits from portfolio diversification are still possible within Monetary Union.

In our opinion, this is a very important conclusion which breaks some fears and supports the idea that portfolio diversification is alive and well in Euroland! (see Adjaouté et *al.* 2004). This is the reason why we think that this topic deserves a deep research and the second objective of this paper will be to analyze the extent to which optimal portfolio allocations might have been modified by the advent of the single currency and the decrease in the cost to investing abroad within the euro area.

Actually, European Monetary Union (EMU) caused large-scale changes in euro-area sovereign securities markets (see Danthine et al. 2001, and the BIS Study group on fixed income markets, 2001). Before the introduction of the euro, yield differentials between European sovereign borrowers were mostly determined by four factors: expectations of exchange rate fluctuations, differences in domestic tax-regimes, differences in credit risk, and differences in market liquidity. The removal of foreign exchange risk in January 1999 and the elimination (or reduction to insignificant levels) of differences in tax treatment during the 1990s eliminated two of these factors, and paved the way for a much more integrated and competitive public debt market.

As a result, euro-area government bond markets began to be considered as a single market, comparable in terms of size to the US or Japanese markets. Nevertheless, segmentation did not disappear completely. In 2010, public debt management is still decentralized under the responsibility of 12 sovereign issuers with differences in rating and a variety of issuing techniques (see Favero, Missale and Piga, 1999). These are features that distinguish the euro-area debt market from its US and Japanese counterparts. One example of this segmentation is the persistence of yield differentials. Gómez-Puig (2006 and 2008) sets out to examine this persistence and to explore what happened to euro-area countries' yield spreads on government bonds after the introduction of the euro[1].

The pre-EMU literature speculated that with the elimination of currency risk, yield spreads would narrow and would primarily reflect default risk. Conversely, market participants and member state debt managers appeared to believe that EMU yield differentials would be due mostly to liquidity

[1] She assesses whether EMU increased credit risk by denying governments the emergency exit of money creation and by forbidding both the ECB and the EU to bail out troubled governments; or whether, conversely, the maximum threshold that countries had for both their budget deficit and their level of public indebtedness (resulting in broad improvements in budgetary balances) and the possibility that markets did not regard the "no-bail-out" clause as credible, especially in the case of large markets (i.e. that the theory "too big to fail" holds), had resulted in a decrease in perceived credit risk.

factors. Therefore, in order to reduce borrowing costs, debt managers introduced substantial innovations that were expected to enhance the liquidity of their bonds.

Actually, the main effects of the introduction of the euro in government bond markets were, on the one hand, an increase in the degree of substitutability among securities issued by different treasuries and higher levels of competition between issuers to attract investors, which led to a certain reorganisation of the market structure[2], and on the other, a gain in the importance of credit risk and market liquidity in yield differentials. Before Monetary Union differences in these factors were perhaps not completely priced due to market segmentation.

Therefore, the introduction of the euro reduced segmentation among euro-area government bond markets. The removal of the exchange rate risk brought down an important barrier that had fostered captive domestic markets and had gone some way to explaining the home bias that existed in cross-border investments in the European Union. Adjaouté et *al.* (2000) traced the extent of the home bias, in both the bond and equities markets, for the major European countries -- the UK, France, Germany, Spain, the Netherlands, and Italy -- during the period 1980-1999[3]. The increased substitutability of sovereign securities after EMU intensified the rivalry between sovereign issuers to attract investors, since they were competing directly for the same pool of funding. In this new scenario, market liquidity differences may have become a more significant component of yield spreads.

Nevertheless, segmentation has not disappeared completely within the EMU and in 2010 public debt management is still decentralized under the responsibility of 12 sovereign issuers. In this scenario, there are many important reasons that justify the study of the second objective of this paper. Concretely, several broad categories of possible portfolio allocation should be considered, domestic versus non-domestic investment, debt versus equity investment, and public debt versus

[2] Blanco (2001) reports that on the side of the issuers, some significant changes were observed such as the harmonisation of market conventions in the computation of yields, the introduction of a single trading calendar and pre-announced auction calendars, or the increase in issue sizes. In some countries, the creation of large issues was facilitated by the introduction of programmes of exchange of old illiquid bonds for new bonds and by the concentration of issuance activity in a smaller number of benchmark securities. With the aim of attracting more investors, some of the smaller issuers such as Austria, Belgium, the Netherlands and Portugal resorted to syndication procedures. Others such as the French Treasury introduced new instruments such as constant maturity and inflation-indexed bonds.

[3] They report that the United Kingdom held the highest share of foreign assets as a function of total financial wealth (24%); Spain had the smallest (5%), and the Netherlands, Germany and Italy had shares around 17%. Moreover, as expected, for bills and bonds, the level of diversification was substantial only for banks in the UK, France and the Netherlands, i.e. the countries where intermediaries played an important role as market-makers in the eurobond markets. These results are consistent with Tesar and Werner (1995), who present evidence on long-term international investment patterns in Canada, Germany, Japan, the UK, and the US during the 1970-1990 period. At the beginning of the 1990s, the UK led this sample in international portfolio diversification, with foreign security holdings of 32% (compared with 10% in Germany).

private debt investment, among them. Adjaouté and Danthine (2004) analyze diversification opportunities in euro area equity investments. We will focus our analysis to public domestic debt versus public non-domestic debt investment in the European Union-15, though. We apply time series techniques in order to detect the existence of common trends in EU-15 daily 10-year sovereign yields during the period 1994-2008. Our aim is to assess whether, after the introduction of the euro, investors can still benefit from portfolio diversification, not only in the three countries that did not join the EMU in 1999, but also within the euro area. To the best of our knowledge, this is also the first empirical analysis that studies government yields co-movements within the EU-15, during this long period of time, with the perspective of investor's portfolio allocation.

3. The Euro and EU-15 Sovereign Debt Market Integration

This section will be focused on the study of the impact of EMU on EU-15 Government Bond Market Integration. With this goal in mind, we assume that Government bond excess returns (r_i) for country i are linearly related to world and local information variables as follows:

$$r_{i,t}=a_i+b^W_i Z^W_{i,t-1}+b^L_i Z^L_{i,t-1}+\varepsilon_{i,t} \tag{1}$$

where Z^W_i represents the world variables, Z^L_i, represents local variables for country i, and $\varepsilon_{i,t}$ is an error term.

Equation (1) is consistent with a range of asset pricing models, and with any level of integration. If a market is fully integrated, the local variables should be absent from Eq. (1). Similarly, if it is completely segmented, the world variables will be absent. We estimate this equation by OLS to identify the relevant world and local instruments.

Once the instruments are identified, we adopt Bekaert and Harvey (1995)'s CAPM-based model and assume that excess returns in country i are generated by the following version of the conditional international CAPM:

$$r_{i,t}= \theta^W \lambda_{w,t-1} \mathrm{cov}_{t-1} (\mathrm{r}_{w,t}, \mathrm{r}_{i,t})+ (1- \theta^W) \lambda_{i,t-1} \mathrm{var} (\mathrm{r}_{i,t}) + e_{i,t} \tag{2}$$

In equation (2), $\boldsymbol{\theta^W}$ is interpreted as a measure of the degree of integration with world bond markets, $\boldsymbol{\lambda_{w,t}}$ is the world price of risk, and $\boldsymbol{\lambda_{i,t}}$ is the local price of risk.

The excess return on the world portfolio Government's bonds is modelled similarly as:

$r_{w,t} = \lambda_{w,t-1} \, \mathrm{var}(r_{w,t}) + e_{w,t}$ (3)

When markets are completely integrated the coefficient θ^W takes the value 1, and the variance term in Equation (2) is reduced to zero. To model the conditional covariance matrix we use a multivariate GARCH model. Specifically, we use the BEKK model proposed by Engle and Kroner (1995). This model can be written as:

$H_t = C'C + A' e_{t-1} e'_{t-1} A + B' H_{t-1} B$ (4)

where **C** is a (NxN) symmetric matrix and A and B are diagonal (NxN) matrices of constant coefficients. By doing this, we allow the variances to depend only on lagged squared errors and lagged conditional variances and the covariances to depend only upon cross-products of lagged errors and lagged conditional covariances (see Bollerslev *et al.* (1988) and De Santis and Gerard (1997, 1998)).

Following the financial literature (see Bekaert and Harvey, 1995 and De Santis and Gerard, 1997, among others), we model the price of risk as a function of a set of information variables. As the price of risk must be positive (see Merton, 1980), the functional form that we assume is:

$\lambda_{w,t-1} = \exp(K'_w Z^w_{t-1})$ (5)

$\lambda_{i,t-1} = \exp(\delta'_L Z^L_{i,t-1})$ (6)

We estimate a system of equations using the Quasi-Maximum Likelihood method. Bollerslev and Wooldridge (1992) show that the standard errors calculated using this method are robust even when the normality assumption is violated. Then, we estimate equations (2), (3) and (4) jointly with the price of risk (equations (5) and (6)), for each of the local Government bond markets, and for the world Government bond market. This estimation is implemented in two steps. First, we estimate the world equation, and then impose the results on the individual countries in 13 bivariate regressions (10 EMU countries, and 3 EU-15 countries that did not join the euro in 1999). We thus restrict the estimates of the world Government bond market price of risk and of the coefficients in the conditional variance of the world market variance to be the same in all countries. Once these estimates are imposed on each bivariate regression, in the second step we will obtain the following for each country: θ_i^W (the estimated level of integration with the world bond market) and δ^L_i (the vector of estimated coefficients for the local price of risk).

As we explained in the previous sections, our analysis goes beyond Barr and Priestley (2004) and Hardouvelis *et al.* (2006 and 2007), who only analyse the impact of one kind of common or systemic risk factor over bond or stock returns behaviour respectively. Unlike them, our aim is to compare the differences in the relative importance of two sources of systemic risk (world and Eurozone risk) on Government bond excess returns since the beginning of Monetary Union, in the two groups of countries (EMU and non-EMU) in EU-15. This is the reason why we also assume that excess returns (r_i) for country i are linearly related to regional (EMU) and local information variables as follows:

$$r_{i,t}=a_i+b^{E}_{i}Z^{E}_{i,t-1}+b^{L}_{i}Z^{L}_{i,t-1}+\varepsilon_{i,t} \quad (7)$$

where Z^{E}_{i} represents the regional (EMU) variables, $Z^{L}_{i,}$ represents local variables for country $i,$ and $\varepsilon_{i,t}$ is an error term.

If we consider that $r_{e,t}$ represents the excess return of the Eurozone Government bond portfolio and replace $r_{w,t}$ by $r_{e,t}$ in equations (2) to (5), we will obtain another system of equations for each of the local bond market and the Eurozone bond market. In particular, analogously to equation (5), the Eurozone price of risk will follow this functional form:

$$\lambda_{e,t-1}= \exp\ (\ K'_{E}\, Z^{E}_{t-1}) \quad (8)$$

We also estimate this system in two steps and obtain, for each country, $\boldsymbol{\theta_i^E}$ (the estimated level of integration with the Eurozone Government bond market), and $\boldsymbol{\delta^L{}_i}$ (the vector of estimated coefficients for the local price of risk).

Hence, two bivariate models will be estimated for each of the countries in our sample: one with world and local risk factors, and the other with European and local risk factors. The final goal is to analyse the impact on each EU-15 country's Government bond return of the three sources of risk: local (own country), regional (Eurozone), and global (world). We also aim to establish whether $\boldsymbol{\theta_i^W}$ and $\boldsymbol{\theta_i^E}$ have differed between EMU and non EMU-countries and across the different countries of each group since the introduction of the euro.

We use weekly data (sampled on Wednesdays) covering the period from January 1999 to June 2008. Using weekly data (compared to, e.g., daily data) partially overcomes the potential problem of non-synchronous data, which may arise because there are instances in which markets are closed in one country and open in another (Burns and Engle (1998) and Lo and MacKinlay (1990) study the effects of this problem). Moreover, we analyse European sovereign bond return behaviour with the perspective given by a long time period (almost 10 years since the beginning of Monetary Union). The empirical analysis makes use of the 10-year Government benchmark yields, and the sample includes 13 countries (all EU-15 countries with the exception of Luxembourg and Greece).[4] Data have been collected from Datastream and Global Financial Data. Bond returns are continuously compounded and are computed with the following formula:

$$R_{it} = p_{it} - p_{it-1} = n(y_{it-1} - y_{it}) \tag{9}$$

Where R_{it} denotes the (weekly) returns on bonds, p_{it} the log price of the bond, $p_{it} \equiv \ln (P_{it})$, y_{it} the log of the gross yield to maturity, $y_{it} \equiv \ln (1+ Y_{it})$, and n the maturity, which in our case is ten years. The dependent variable in our model (r_{it},) is the excess return[5] which is calculated relative to the appropriate 1-month Euro-deposit rate quoted in London.[6]

We use the following instrumental variables to capture the different prices of risk (world, regional and local risk): (1) the slope of the yield curve, as measured by the difference between the 10-year and the 3-year Treasury yield. Several studies (Campbell and Shiller, 1991; Ilmanen, 1996) have found that steeper yield curves are associated with higher subsequent yields on longer-maturity bonds. The interpretation of this finding is that the yield curve steepens primarily because of an increase in the risk premium. Moreover, the slope of the yield curve is also a proxy of the business cycle. (2) Lagged stock indexes returns are included to allow for the possibility that stock returns lead bond returns. In recent years, important cross-asset linkages between stocks, bonds and money market instruments have been observed. Fleming, Kirby and Ostdiek (1998) investigate the nature of volatility linkages between stocks, bonds and money markets and conclude that volatility linkages between the three markets are strong. In particular, stock market weakness has been associated with economic weakness, which has corresponded to bond market strength.[7] If equity market weakness gives rise to subsequent bond market strength, the coefficient on lagged stock indexes returns should be significantly negative (see Hunter and Simon, 2005).[8] (3) Lagged 10-year Government returns are also added to the specification. Taking into account that some aspects of risk premiums (related to domestic factors such as liquidity or credit risk) do not change over the period

[4] Luxembourg's public debt market is negligible and Greece did not join Monetary Union until January 2001. For these reasons, these two countries are not included in the analysis.

[5] International CAPM models (ICAPM) contain additional terms to reward exchange rate risk. Concretely, risk premium are based on the covariances of assets with exchange rates, in addition to the traditional premium based on the covariance with the market portfolio (see Solnik (1974) and Adler and Dumas (1983)). These models are mostly used to analyze the predictability of international stock returns, which is usually examined in terms of a common currency. However, since the volatility of exchange rates greatly exceeds that of interest rates (see Thomas and Wickens, 1993), the predictability of bond returns, however, is more usually analyzed only using local-currency returns (see Barr and Priestley, 2004). Otherwise, results might produce more evidence on the predictability of exchange rates than of bond returns. Consequently, in order to avoid this bias in the results, the dependent variable in this paper is the excess bond returns in local-currency.

[6] Euro-deposit rates are used as a proxy for the risk free rate due to the lack of a liquid Treasury bill market in some of the countries. The excess world return is calculated with reference to the rate on \$US deposits, whilst the excess Eurozone return is calculated with reference to the Euribor rate.

[7] Kim *et al.* (2006) present evidence that the introduction of the monetary union has Granger caused an apparent segmentation between bond and stock markets within Europe. Hence, the EMU has increased benefits of diversification across stocks and government bonds at the country level.

[8] Nevertheless, note that other authors (see McQueen and Roley (1993)) demonstrate that the opposite results are obtained when market participants are concerned about an overheating economy. During these periods, data suggesting a weaker-than-expected economy lead to stronger bond and stock prices as this makes it less likely that the Federal Reserve will be forced to tighten monetary policy aggressively and possibly drive the economy into a recession.

considered, the objective will be to identify their relative importance in explaining fluctuations, rather than returns levels. With this aim a lag of the dependent variable is introduced in the model, which will allow for a slow dynamic adjustment to a long-term equilibrium value of Government returns. (4) Moreover, we include the difference between lagged 10-year Government returns and lagged stock index returns to capture bond markets relative risk compared to stock markets. Finally: (5) the difference between lagged corporate bond returns and 10-year Government returns is also an important information variable that will be included in the specification. It can be interpreted as a proxy of the credit cycle[9] or, more importantly, as a proxy for time-varying credit risk premium in the bond market.[10]

The same five variables are used as information variables to capture the price of regional and world risk. In the case of regional risk, we use German returns (the German 10-year yield is the benchmark in the euro area) as proxies of the behaviour of Euro area debt markets. We think that this is a better way to capture regional risk effect than using the return of a synthetic Euro area bond that will always contain the evolution of its own local market return. Similarly, US data are used to capture the price of world risk.[11]

Therefore, the following regional instruments are used: (1) the slope of the German yield curve, as measured by the difference between the 10-year and the 3-year German Treasury yield. (2) The lagged return of the Eurostoxx50. We think that the use of this index is appropriate as it reflects the price evolution of the 50 most important firms in the euro area (unlike the Eurozone 10-year synthetic Government yield, it is not built up as an average of the different local market indexes). (3) The lagged value of the 10-year German Government return. (4) The difference between lagged 10-year German Government return and lagged Eurostoxx50 return. And (5): the difference between lagged German corporate bonds return and 10-year German Government return. For their part, the world instruments are: (1) the slope of the US yield curve, as measured by the difference between the 10-year and the 3-year US Treasury yield. (2) The lagged return of the Standard & Poor's 500. (3) The lagged value of the 10-year US Government return. (4) The difference between lagged 10-year US Government return and lagged Standard & Poor's 500 return. And (5): the difference between lagged US corporate bonds return and 10-year US Government return.

Then, we will estimate 13 bivariate models (all EU-countries except Luxembourg and Greece) which will contain local and world instruments, and 12 models (all EU-countries with the exception of Germany, Luxembourg and Greece) which will contain local and regional instruments.

[9] Landschoot (2008) provides evidence that both euro and US dollar yield spreads are significantly affected by the credit cycle.

[10] For each individual country (except for Ireland and Portugal due to the lack of available data) we use a corporate bond index which has been built by Lehman Brothers or The Economist, depending on the country. These data have been provided by Datastream.

[11] Barr and Priestley (2004) present evidence that the US-world return correlation is very high, reflecting the relatively large proportion of US bonds in the world portfolio.

First, we investigate the extent and sources of predictability in local bond markets. To do this, we estimate equation (1) using world and local instruments (Panel A in Table 1) and regional and local instruments (Panel B in Table 1). In each case we test the separate hypothesis that the coefficients associated with the world (regional) and local variables are zero. When we use world and local instruments jointly (Panel A) the R^2s range from 53% in Ireland to 93% in the Netherlands, indicating a high degree of predictability. For all countries we reject the null hypothesis that both sets of instruments can be excluded. Then, we estimate equation (1) using the world and local instruments separately. In both cases, the results show clear patterns of predictability in all the local bond markets using local instruments. We observe that when we use only one set of instruments the R^2s are lower than when we use both sets, implying that it is necessary to include all kinds of instruments. Similarly, if we use regional and local instruments jointly (Panel B) the R^2s range from 68% in Denmark to 92% in the Netherlands, also indicating a high degree of predictability. The F-tests reveal that each set of instruments is separately and jointly significant. We also report estimated equations for local returns based on the regional instruments only. The results indicate that regional instruments are able to predict local bond returns in all markets. Overall, these results show that a set of world (regional) and local instruments are useful to predict local bond returns, suggesting incomplete integration.

Tables 2 and 3 present the results of the estimation of the system of equations [(2), (3), (4), (5) and (6)] using the Quasi-Maximum Likelihood method for each of the local Government bond markets jointly with the world (United States) Government bond market (Table 2) and the Eurozone (German) Government bond market (Table 3). Tables 4 and 5 show the standardized residuals analyses. It can be observed (with few exceptions) that the standardized residuals appear free from serial correlation and heteroskedasticity. In all cases, the necessary conditions for the stationarity of the process are satisfied.[12]

All world instruments are relevant in forecasting the world price of risk, as it is shown in the first row of Table 2. The estimates of δ^L s in Table 2 indicate that all local instruments are also important in explaining the local price of risk except for the cases of Sweden and the U.K, confirming its higher degree of dependence on world risk factors[13]. The results of the estimation including world and local risk factors indicate that EMU and US Government bond markets present a low degree of integration. The estimated level of integration with the world bond market ($\boldsymbol{\theta^W}$) displays an average value of 0.052 in the countries that belong to the euro. This level seems low

[12] The BEKK model is covariance stationary if and only if the eigenvalues of the matrix $A \otimes A + B \otimes B$, where $\otimes$ denotes the kronecker product of two matrices, are less than one in modulus [see Proposition 2.7 in Engle and Kroner (1995)].

[13] See Spencer and Liu (2009) for an analysis of the effects of world economic variables on the UK economy and Treasury bond market.

in view of the absence of major impediments to cross-country investment[14]. There are no significant differences within countries: Germany presents one of the highest degree of integration with the US Government bond market (0.067), only surpassed by Belgium (0.069). These results present clear evidence that it is domestic (idiosyncratic) rather than international (systemic) risk factors that mostly drive the evolution of 10-year Government debt returns in all EMU countries throughout almost all of the ten-year period after the beginning of Monetary Integration.

These results also indicate that the degree of integration with US markets clearly differs between euro and non-euro participating countries. Whilst the average value is 0.052 in the case of euro countries, it is 0.468 in the case of non-euro countries. Hence, in the case of non-EMU countries, Government debt returns are influenced more by world risk factors. The fact that these countries do not share a common Monetary Policy may explain their higher vulnerability to international (systemic) risk factors.

Finally, note that there are important differences in $\boldsymbol{\theta^W}$ value in the case of non-euro countries. Denmark is the country that presents the lowest degree of integration (0.086) with US debt markets. Actually, the fact that the exchange rate regime, in that country, links the evolution of its currency to the Euro explains why the Danish Government debt returns present a behaviour that is closer to EMU-countries than to non-EMU countries[15]. Moreover, the degree of integration with US markets is much larger in the case of Sweden (0.936) than in the case of the United Kingdom (0.383). The fact that the British market is one of the most important European debt markets (the fourth biggest, after the Italian, the German and the French markets), could be the reason for its higher degree of independence from world risk factors.

The first row in Table 3 shows that all regional instruments are relevant in forecasting the regional price of risk. Similarly, all local instruments are also important in explaining the local price of risk (with few exceptions) As expected, the estimated level of integration with the German bond market ($\boldsymbol{\theta_i^E}$) differs between euro and non-euro participating countries. The average values are 0.376 and

[14] Differences between bonds in different countries are small, so it seems reasonable to expect a high degree of integration in bond markets (much larger than in equity markets). However, there are some reasons for expecting that bond markets may not be "fully" integrated, which are basically related to "home bias" on both the investors and issuers' side. For instance, one of the major impediments in the debt market is the currency matching rule widely adopted across countries. Pension funds for example are forced to invest a share of their funds in local currency.

[15] On 1 January 1999, the Exchange Rate Mechanism (ERM II) was set up as a successor to ERM (a mechanism that before the introduction of the Euro promoted a stable exchange rate zone in the EU with a monetary policy anchored to the Deutsch mark). The objective of ERM II is to ensure that exchange rate fluctuations between the Euro and other EU currencies do not disrupt economic stability within the single market. Participation in ERM II is voluntary although a country must participate in the mechanism without severe tensions for at least two years before it can qualify to adopt the euro. The Danish Kroner participated in the ERM since its creation in 1979 and joined the ERM II on 1 January 1999. Actually, it observes a central rate of 7.46038 to the Euro with a narrow fluctuation band of ±2.25%, whilst the Swedish Kroner and the Sterling Pound have never belonged to the ERM II. In the case of Sweden, its currency has never belonged to the ERM either, and the Sterling Pound had a very short-lived stay in it (from 1990 until 1992). This fact could explain why the degree of integration with the US of the local government bonds of Denmark, a country that did not join the EMU but belonged to the ERM from its creation and it is now in the ERM II, is so low, similar to that of other EMU countries. However, the coefficient is very large in the case of Sweden, and also for the UK.

0.078, respectively, for EMU and non-EMU countries. Within EMU countries, the Finnish market is the one that presents the lowest degree of integration (0.149), and The Netherlands presents the highest (0.627); the rest of the countries present very similar values, around 0.376 on average, indicating that euro-participating markets are only partially integrated with the German market. This fact captures the idea that European Government bond markets are still imperfect substitutes due to differences in their domestic risk factors (either market liquidity or default risk)[16]. Outside the Monetary Union, Denmark is again the market that presents a behaviour that is much closer to euro than to non-euro participating countries. In addition, Britain and Sweden present a similar degree of integration with the German market, 0.049 and 0.044, respectively.

The introduction of the euro had a major impact on the degree of integration of European Government bond markets. Within the Currency Union, on average, the estimated level of integration with the world ($\boldsymbol{\theta}^{W}$) and German ($\boldsymbol{\theta}^{E}$) bond markets is 0.052 and 0.379 respectively. Conversely, outside the Monetary Union, these levels present the following average values: 0.468 and 0.067. Consequently, the markets of those countries that share a monetary policy are less vulnerable to the influences of world risk factors and more vulnerable to EMU risk factors. However, they are only partially integrated with the German market since their markets are still segmented and present differences in their market liquidity or default risk. The empirical evidence presented by Laopodis (2008) also shows a weak degree of integration among the EMU bond markets since the beginning of the Currency Union. These findings have important implications for investors in terms of portfolio diversification benefits, and are an argument against the issue of a single European bond, a matter that is currently under debate.

For their part, the countries that decided to stay out of Monetary Union and maintain monetary autonomy present a higher vulnerability to external risk factors. These results are in concordance with Gómez-Puig (2009b), who presents empirical evidence that it was mostly idiosyncratic rather than systemic risk factors that drove the evolution of 10-year yield spread differentials over Germany in all EMU countries during the first seven years of Monetary integration. Conversely, in the case of non-EMU countries, adjusted yield spreads (corrected from the foreign exchange factor) are influenced more by systemic risk factors.

Finally, we re-estimate the model to test the restriction of a constant price of risk. We use the likelihood-ratio test procedure to examine whether the reduced model (with a constant price of

[16] Gómez-Puig (2008) provides evidence that market size scale economies seem to have increased with Currency Union and that the smaller the debt market, the higher the rise. Therefore, the removal of the exchange rate barrier seems to have punished smaller countries twice (they are forced to compete in terms of liquidity with larger countries for the same pool of funding, only being able to offer smaller bond issues), by making them pay both higher liquidity and a higher default risk premium than larger ones.

risk) provides the same fit as the full model (with a time-varying price of risk). Results (not reported) imply strong rejection of the null hypothesis of a constant price of risk and justify modelling the price of risk as a time varying function.

4. The Euro and Risk Diversification in EU-15 Sovereign Debt Markets

The second objective of this paper is to examine in depth the impact of Monetary Integration on portfolio diversification opportunities in European public debt markets, since this study has important implications for investors, in terms of diversification benefits in a context of a single currency.

With this goal in mind, we examine the existence of common trends in daily 10 year sovereign yields for EU-15 countries (both EMU and non-EMU participating) during the period 1994-2008 by means of using multivariate cointegration techniques. It is worth to note that 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).

Cointegration techniques have been widely used by the literature to examine co-movements and linkages in international bond markets. Ilmanen (1995) examined the effect of integration of six markets (the US, the UK, Canada France, Germany and Japan) using long-maturity government bonds. He found strong evidence of integration across mature bond markets. On the other hand, Clare, Maras and Thomas (1995) using the daily yield of mature market bonds (the US, UK, West Germany and Japan) found no cointegrating relationship between these markets. Their study focused on bonds with less than 5 year maturity to test market integration through a multivariate cointegration framework. Similarly, Mills and Mills (1991) investigating four major bond markets find no integration. Arshanapalli and Doukas (1994) investigated the temporal relationship between Eurodeposit instruments of five different maturities for different currencies and found several cointegrating factors binding them together for the period between 1986 and 1992. Their multivariate cointegration test for dependency on five maturity sets of seven dimensional system reveals that the cointegrating structure is stronger at the short end rather than at the long end of the maturity spectrum. Kim, Lucey and Wu (2006) examine the time varying level of financial integration of European markets using government bond indices of European economies (Czech Republic, Hungary, Poland, Belgium, France, Ireland, Netherlands, the UK and Germany) in the region. Their test was to see how the Euro zone markets were integrated with Germany. They found strong evidence of linkages between Euro zone markets and Germany. However, a cointegration analysis that studies the nature of financial market integration in EU-15 Government's bond market with the final goal to study risk diversification opportunities is yet to be explored.

This will be then the second objective of this paper. The sample is composed of daily 10-year government yields and includes all EU-15 countries with the exception of Luxembourg (its public debt market is negligible). Data have been obtained from Datastream and correspond to the "on the run" (benchmark) 10-year issue for each market at every moment of time (they are quoted rates at market close)[17] and span the period January 1994 to the end of December 2008. So, we include daily information[18] that covers five years before the beginning of Currency Union and close to ten years after the introduction of the euro. Figure 1 shows the temporal evolution of each of the individual yields. It is important to note that the sample period (1994-2008) is split into two subperiods, namely 1994-1998 and 1999-2008, in order to take into account the introduction of a single currency in January 1999.

Before the introduction of the euro, yield differentials between European sovereign borrowers were mostly determined by four factors: expectations of exchange rate fluctuations, differences in domestic tax-regimes, differences in credit risk, and differences in market liquidity. The removal of foreign exchange risk in January 1999 and the elimination (or reduction to insignificant levels) of differences in tax treatment during the 1990s eliminated two of these factors, and paved the way for a much more integrated and competitive public debt market.

The aforementioned elimination of two of the main components of yield differentials prompted a substantial convergence in EMU 10-year yields during the period January 1999-December 2008. This is clearly reflected in figure 1. Moreover, non-EMU countries yields, also displayed an important decrease and begun to convergence with euro yields after the introduction of the common currency. In particular, Figure 1 shows that the country with the highest spread over Germany is the United Kingdom, while the country whose government's yields follow those of Germany most closely is Denmark.

[17] Datastream creates continuous yield series by taking the yield from the current benchmark in each market and using it to update a separate time series. As a benchmark changes, data are taken from a new stock on the first day of the month.

[18] We don't have the potential problem of non-synchronous data, which may arise because there are instances in which markets are closed in one country and open in another, in this second part of the analysis which allows us to work with daily data.

To obtain a first approximation of the extent of linkages among Government's yields, Table 6 reports the contemporaneous correlation coefficients of all 10 year yields for the whole sample and for each of the two subperiods. A first result from these correlations is the high degree of interdependence among all the Government bond markets along the whole time period. However, the degree of correlation among the UK yields and the rest of European yields appear to be lower, especially for the second subperiod.

Regression results are likely to be spurious if the variables concerned are not stationary. To determine the integration order of each of the variables, we carry out Augmented Dickey-Fuller (ADF)[19] unit root tests in order to check the integration order of each of the Government's yields[20]. That is, we test the null hypothesis $\gamma=0$ based on next equation,

$$\Delta y_t = \alpha_0 + \alpha_1 t + \gamma y_{t-1} + \sum_{j=1}^{k} \theta \Delta y_{t-j} + \varepsilon_t \tag{10}$$

where y is the Government 10-year yields and k is the number of lagged differences included to capture any autocorrelation. Three ADF unit root tests are carried out (i) with no regressors ($\alpha_0=\alpha_1=0$), (ii) with a constant ($\alpha_1=0$) and (iii) with a constant and a linear time trend (see Table 7). The results from the ADF tests are tabulated for the first period (panel 1), second period (panel 2) and the whole period (panel 3). Overall, the findings provide strong evidence in favor of the null hypothesis of unit root for all the interest rates. Furthermore, when ADF unit roots are carried out to the first differences of 10 year yields, we reject the unit root null hypothesis in all the cases, suggesting that 10 year yields follow I(1) processes.

In a next step, we test for multiple cointegration using the Johansen and Juselius (1990) multivariate cointegration technique, that is, we tests whether or not the yields share a common stochastic trend.

In the literature, two primary methods exist to examine the degree of cointegration time series. The first is the Engle–Granger methodology (see Engle and Granger, 1987) which is bivariate, testing for cointegration between pairs of indices. The second is the Johansen–Juselius technique (see Johansen, 1988 and Johansen and Juselius, 1990), which is a multivariate extension and allows for more than one cointegrating vector or common stochastic trend to be present in the data. Although bivariate

[19] Alternative approaches to test for unit roots are Phillips and Perron (1988), Kwiatkowski *et al.* (1992) and recently Ng and Perron (2001). The results for those tests are available upon request.

[20] See Dickey and Fuller (1979).

cointegration procedures capture the dynamics between series, the under dimensionality associated with this testing process introduces bias in the identification and attribution of the common factor. In addition, cointegration is multivariate by nature and the equilibrium dynamics and the common features in the series are best understood when they are analyzed in a multivariate setting (see Hendry and Juselius, 2001). This is the reason why we will use the Johansen-Juselius technique in this study. The main advantage of it this approach is that it allows testing for the number as well as the existence of these common stochastic trends.

As mentioned, cointegration is based on the idea that while a set of variables are individually I(1) nonstationary, a linear combination of these variables might be stationary. While the variables are individually unbounded, the existence of a stationary combination implies that the variables cannot drift arbitrarily far apart. Intuitively, it is the long-run equilibrium relationship that links the cointegrated variables together. We start by considering a vector autoregression of order p:

$$y_t = A_1 y_{t-1} + A_2 y_{t-2} + ... + A_p y_{t-p} + Bx_t + \varepsilon_t \tag{11}$$

where y_t is a k-vector of non-stationary I(1) variables, x_t is a vector of deterministic variables and ε_t a vector of innovations. The above equation can be rewritten as

$$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta y_{t-1} + Bx_t + \varepsilon_t \tag{12}$$

where

$$\Pi = \sum_{i=1}^{p} A_i - I \tag{13}$$

and

$$\Gamma_i = -\sum_{j=i+1}^{p} A_j \tag{14}$$

The number of cointegrating relationships is indicated by the rank of Π. If Π has reduced rank ($r<k$), then there are $k-r=n$ common stochastic trends. There are three possible scenarios depending on the rank of Π. First, if Π is of full rank ($r=k$), then all elements of y_t are stationary, which implies that the cointegration technique is inappropriate. Second, if the rank of Π equals zero, there are no combinations which are stationary, and, thus, there are no cointegrating vectors. Finally, if the rank of Π is r ($0<r<k$), then the y_t variables are cointegrated and there exist r cointegrating vectors. In this

case, and if the number of common trends is exactly one (n=k-r=1), we conclude that the interest rates are integrated completely and perfectly, since they are driven by the same common stochastic trend. If the number of common stochastic trends is more than one, we conclude that interest rates present certain degree of interdependence, although not complete convergence.

In essence, the Johansen-Juselius approach involves the determination of the rank of a matrix of cointegrating vectors. This technique generates two statistics of primary interest. The first is the λtrace statistic, which (in this instance) is a test of the general question of whether there exist one or more cointegrating vectors. The second is the λmax statistic (maximum eigenvalue), which allows testing of the precise number of cointegrating vectors.

Thus, the output from the approach we employ is two-fold: first, the largest value of the λtrace statistic which tests the general hypothesis of no cointegration versus cointegration, and second, the number of cointegrating vectors given by the λmax or maximum eigenvalue statistic.

The trace statistic for the null hypothesis of r cointegrating vectors is defined as

$$LR_{TR}(r|k) = -T\sum_{i=r+1}^{k}\log(1-\lambda_i) \qquad (15)$$

Where λ_i is the ith largest eigenvalue of Π, r is the number of cointegrating relationships and T is the number of observations. The maximum eigenvalue statistic for the null hypothesis of r cointegrating vectors is expressed as

$$LR_{\max}(r|r+1) = -T\log(1-\lambda_{r+1}) \qquad (16)$$

The results for the multiple cointegration tests are reported in Tables 8.1-8.2 for the whole sample of countries and Tables 9.1 and 9.2 for the EMU countries[21]. As in the previous analysis, cointegration tests have been carried out for the whole period and each of the two subperiods.

The results presented in these tables may be summarized as follows. First, the number of cointegrating vectors for the first subperiod is lower (9 or 8 depending on whether we use the trace or maximum eigenvalue statistic) than in the second subperiod, which implies that the degree of cointegration or interdependence among long term interest rates is lower in the period before 1999. Second, the number of cointegrating vectors in the first subperiod is the same when analyzing only EMU countries (Table 9.1) and all European countries (Table 8.1), which suggest that the degree of

[21] Tables 8.1 and 9.1 do not include Greece in the analysis, whilst the results of the estimates including this country are reported in Tables 8.2 and 9.2.

interdependence found among interest rates is limited to EMU countries, while interest rates of non-EMU countries do not share any common trend with the rest of interest rates. Third, the removal of the exchange risk in 1999 increased the number of cointegrating vectors (and the interdependence) among not only EMU countries, but among non-EMU countries as well.

5. Conclusions

The objective of this paper is twofold. First, we analyze the impact of Monetary Union on European debt market integration, looking at integration both with world debt markets and with Eurozone debt markets. Second, we examine the impact of Monetary Integration on portfolio diversification opportunities in public debt markets.

Regarding the first objective, we separate each individual country's Government bond return into three effects: a local (own country) effect, a regional (Eurozone) effect, and a global (world) effect. We examine whether there are significant differences within two different groups of European countries: those that joined the euro in 1999 and those that did not. The objective is to explore whether participation in the Monetary Union is an important factor that determines the difference in the impact of world and regional risk on each European Government bond market.

Our sample period goes from January 1999 to June 2008, covering almost ten years since the introduction of the common currency. We use Bekaert and Harvey's CAPM-based model (1995). This is the first time that this methodology has been used to analyze the differences in the relative importance of two sources of risk, systemic and idiosyncratic. In contrast to the previous literature, which has focused only on one kind of systemic risk, we distinguish between the world and the Eurozone risk.

The most important results of the paper are the following. First, the results show that apart from a set of world (regional) instruments, a set of local instruments are also able to predict local bond returns. This result suggests incomplete integration. Second, we find that EMU and US Government bond markets present a low degree of integration, indicating that it is domestic rather than international risk factors that mostly drive the evolution of government debt returns in EMU countries.

Third, the results show that the degree of integration with the US and German bond markets clearly differs between euro and non-euro participating countries. Government bond returns of non-EMU countries are more influenced by world risk factors. This result agrees with Gómez-Puig (2009b) and indicates that these countries present a higher vulnerability to external risk factors. On the other hand, Government bond returns of EMU countries are more influenced by Eurozone risk factors. In spite of this, EMU countries are only partially integrated with the German market since their markets are still segmented and present differences in their market liquidity or default risk. In a different context, Laopodis (2008) reached the same conclusion, suggesting that benefits from portfolio diversification are still possible within Monetary Union.

Regarding the second objective, which is linked with the third abovementioned conclusion, we analyze (also for the first time) the convergence process of EU-15 countries by means of using time-series tests over the period 1994-2008. So, we study whether there are differences in risk diversification opportunities during the five years preceding the introduction of the euro and the ten years after its implementation. In particular, we examine a multivariate framework, using Johansen and Joselius (1988) cointegration test. The results suggest the following.

First, multiple cointegration tests rejects the existence of a unique common trend among the 14 yields that have been analyzed, suggesting the possibility of risk diversification across Europe. Furthermore, when this analysis is carried out only for EMU countries, we also find the existence of more than one common trend for these countries.

However, it is worth to note that there exist differences between EMU and non-EMU countries and between the two periods that have been analyzed. Interdependence is significantly lower during the period that includes the five years before the introduction of the common currency than during the ten years after the euro. Before EMU, there were high opportunities of portfolio diversification, specially, among the countries that did not join the euro in 1999. Actually, during the period 1994-1998, the UK, Denmark and Sweden do not share a common trend with the rest of EU-15 countries. This result is highly reasonable since the countries that stayed out of the euro did not have to fulfill the convergence criteria in long-term interest rates[22].

Nevertheless, the removal of the exchange risk in 1999 increased the number of cointegrating vectors (and the interdependence) not only among EMU countries, but among non-EMU countries as well (meaning that currency union has prompted a convergence of long term interest rates in all EU-15 countries). So, portfolio diversification opportunities have experienced a reduction with the single currency but multiple cointegration tests still rejects the existence of a unique common trend

[22] In particular, the Treaty stipulated: "the durability of convergence achieved by the Member State ... being reflected in the long-term interest-rate levels". In practice, the nominal long-term interest rate must not exceed by more than 2 percentage points that of, at most, the three best-performing Member States in terms of price stability (that is to say, the same Member States as those in the case of the price stability criterion). The period taken into consideration is the year preceding the examination of the situation in the concerned Member State.

among the 14 yields that have been analyzed, which suggests that diversification benefits have not disappeared with EMU.

So, we can conclude that, even though the countries that do not participate in the common currency present higher benefits from diversification, EMU has not eliminated portfolio diversification opportunities in public debt markets' investments. These results are sound with the empirical evidence presented by Adjaouté and Danthine (2004) who examine diversification opportunities in euro area equity markets during the period 1999-2001[23]. Indeed, the removal of the exchange rate barrier fostered convergence in EMU public debt markets (which has been extended to all EU-15), these markets are not perfectly integrated. The fact that public debt management is still decentralized under the responsibility of 12 Sovereign issuers (with differences in rating and a variety of issuing techniques) and there still exists differences in their domestic risk premium (liquidity and default risk)[24] explains their imperfect integration. Hence, benefits from diversification are still possible within EMU public debt markets. This conclusion supports the results of the first part of our analysis, where we found that EMU sovereign debt markets were only partially integrated with the German market.

[23] Concretely, their results clearly invalidate the hypothesis that diversification opportunities in the euro-area have been permanently impaired as a consequence of the process of economic and monetary integration. Conversely, they strongly confirm the superiority of a model where diversification is sought after simultaneously across country and sector dimensions over the traditional country allocation model

[24] See Gómez-Puig (2006 and 2008)

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7. Tables and Figures

FIGURE 1: 10 year Governments' yields

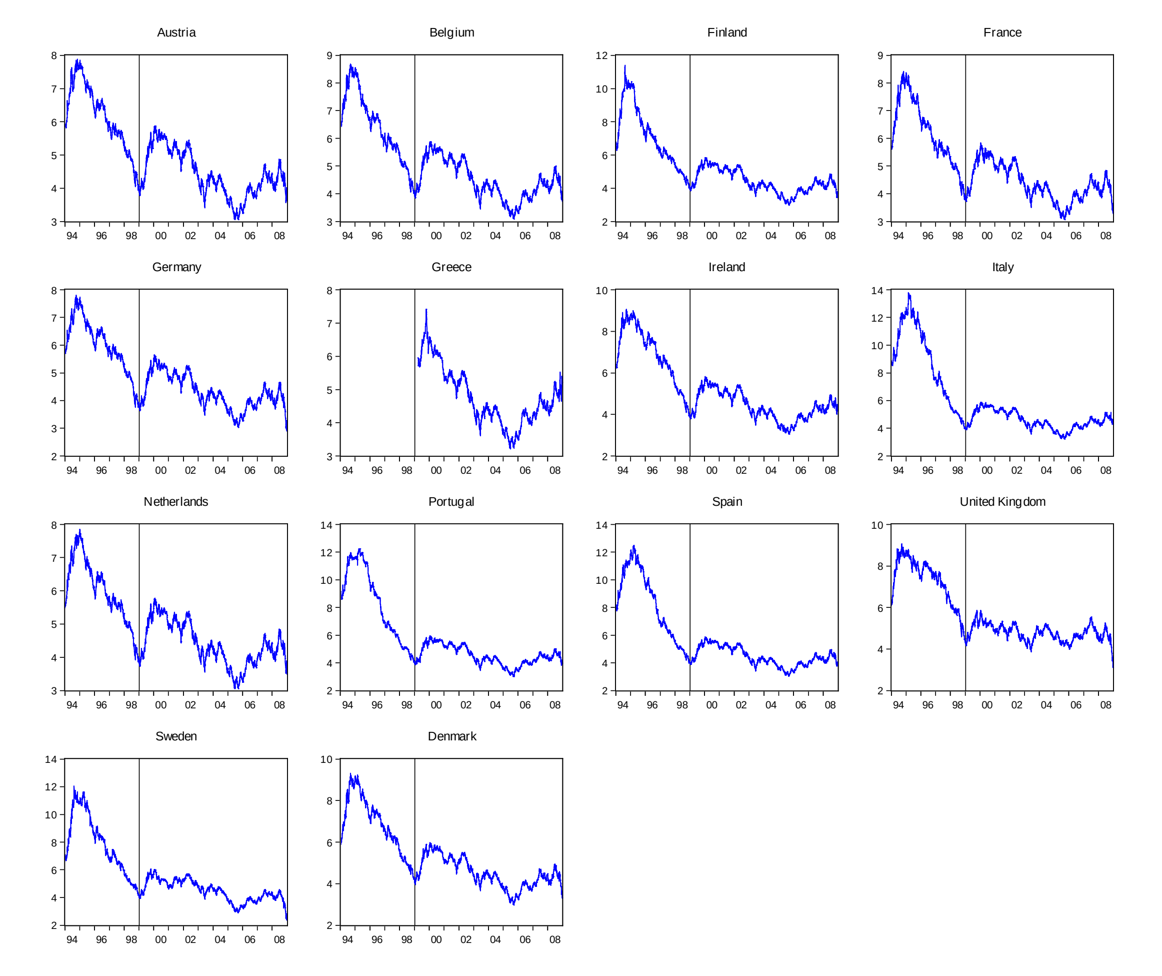

Table 1: Predicting local excess returns

	Germany	Austria	Belgium	Denmark	Spain	Finland	France	Ireland	Italy	Netherlands	Portugal	Sweden	U. K
Panel (A). Word and local instruments													
R2	0.82	0.85	0.86	0.61	0.85	0.65	0.92	0.53	0.75	0.93	0.54	0.67	0.82
F-test	223.62	278.90	292.03	76.23	269.82	89.26	545.55	61.40	148.51	619.21	62.14	98.77	216.50
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
F-test exclude local	167.39	229.70	248.37	27.33	220.06	43.82	553.98	14.10	104.82	570.30	14.09	73.51	204.23
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
F-test exclude world	35.87	31.48	24.27	56.08	24.95	52.96	33.25	59.83	46.12	11.39	65.41	21.64	17.05
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Local instruments only													
R2	0.76	0.80	0.82	0.39	0.81	0.46	0.89	0.24	0.64	0.92	0.22	0.60	0.79
F-test	303.08	401.08	452.03	61.61	413.28	81.93	795.09	39.56	171.58	1109.00	35.00	145.20	357.22
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
World Instruments only													
R2	0.51	0.50	0.49	0.50	0.50	0.49	0.45	0.48	0.49	0.50	0.48	0.42	0.43
F-test	103.46	98.14	94.98	98.55	98.50	93.62	80.58	89.61	93.14	97.78	90.84	71.17	74.22
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Panel (B). Regional and local instruments													
R2	--	0.85	0.85	0.68	0.84	0.74	0.91	0.64	0.73	0.92	0.69	0.76	0.84
F-test	--	272.17	284.12	104.43	254.14	140.99	463.16	95.19	131.90	574.23	117.66	149.58	251.42
	--	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
F-test exclude local	--	85.05	133.47	16.33	90.96	14.75	277.05	4.50	25.56	306.13	2.74	71.49	191.00
	--	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.03)	(0.00)	(0.00)
F-test exclude regional	--	28.84	21.46	90.66	18.96	109.21	15.24	105.79	34.07	4.11	143.10	62.49	32.04
	--	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Regional Instruments only													
R2	--	0.72	0.65	0.63	0.69	0.71	0.63	0.63	0.66	0.68	0.68	0.58	0.52
F-test	--	246.78	184.44	166.39	217.16	234.22	169.59	163.06	190.34	204.13	206.65	132.19	105.83
	--	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)

Note: The table reports OLS estimations of equation (1). R2 denote R-squared statistic. F-test denotes the F-statistic from a test of the hypothesis that all of the slope coefficients (excluding the intercept) in the regression are zero. F-test exclude X denotes the F-statistic from a test of the hypothesis that some coefficients (all excluding the set X) in the regression are zero. P-values are displayed in parentheses.

Table 2: Model estimates for each of the local Government bond market jointly with the world Government bond market

$$r_{i,t} = \theta^{W} \exp(K'_{w} Z^{w}_{t-1}) \, cov_{t-1}(r_{w,t}, r_{i,t}) + (1-\theta^{W}) \exp(\delta'_{L} Z^{L}_{i,t-1}) \, var(r_{i,t}) + e_{i,t}$$

$$r_{w,t} = \exp(K'_{w} Z^{w}_{t-1}) \, var(r_{w,t}) + e_{w,t}$$

$$\mathbf{H}_{t} = \mathbf{C'C} + \mathbf{A'} \boldsymbol{e}_{t-1} \boldsymbol{e}'_{t-1} \mathbf{A} + \mathbf{B'} \mathbf{H}_{t-1} \mathbf{B}$$

	$K0_{w}$	$K1_{w}$	$K2_{w}$	$K3_{w}$	$K4_{w}$	$K5_{w}$		C	A	B		
World	-186.825	-40.500	-4.195	-23.031	-71.391	2.751		0.000	0.085	0.346		
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)		(0.000)	(0.000)	(0.000)		
	$\delta 0_{L}$	$\delta 1_{L}$	$\delta 2_{L}$	$\delta 3_{L}$	$\delta 4_{L}$	$\delta 5_{L}$	θ^{W}	C11	C22	C12	A22	B22
Germany	-97.562	-15.238	-3.877	24.685	10.444	13.597	0.067	0.012	0.000	0.005	-0.205	0.792
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.816)	(0.000)	(0.025)	(0.007)
Austria	-20.903	-22.028	-4.855	1.709	-0.437	8.896	0.041	0.012	0.000	0.005	-0.206	0.798
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.850)	(0.000)	(0.026)	(0.007)
Belgium	-24.183	-4.331	0.614	1.377	0.646	-1.452	0.069	0.012	0.000	0.005	0.243	0.802
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.743)	(0.000)	(0.023)	(0.007)
Spain	-21.889	-455.78	-9.353	-0.298	0.744	0.545	0.054	0.012	0.000	0.005	-0.170	0.803
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.785)	(0.000)	(0.029)	(0.007)
Finland	-51.677	3.918	0.018	-10.936	-0.596	-12.460	0.055	0.012	0.000	0.005	-0.173	0.797
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.804)	(0.000)	(0.026)	(0.006)
France	-20.221	-78.632	24.850	0.496	1.582	-1.557	0.025	0.012	0.000	0.005	0.242	0.819
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.768)	(0.000)	(0.022)	(0.006)
Ireland	-26.289	0.824	0.211	0.055	-7.271	---	0.047	0.012	0.000	0.005	0.215	0.816
	($9.667 \cdot 10^{10}$)	(0.000)	(0.000)	(0.000)	(0.000)	(---)	(0.000)	(0.000)	(2.561)	(0.000)	(0.024)	(0.005)
Italy	-42.269	-1.125	2.512	14.853	1.555	13.208	0.061	0.012	0.000	0.005	0.242	0.819
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.740)	(0.000)	(0.022)	(0.006)
Netherlands	-43.784	0.440	0.422	41.781	0.883	-2.258	0.051	0.012	0.000	0.005	-0.184	0.793
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.876)	(0.000)	(0.026)	(0.007)
Portugal	-22.893	-9.393	-4.333	2.152	0.908	---	0.054	0.012	0.000	0.005	0.223	0.814
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(---)	(0.000)	(0.000)	(0.749)	(0.000)	(0.024)	(0.006)
Denmark	-29.084	1.359	0.874	3.091	-0.148	-1.324	0.086	0.012	0.000	0.005	-0.180	0.829
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.779)	(0.000)	(0.026)	(0.006)
Sweden	5.024	-0.136	6.765	39.481	5.045	28.779	0.936	0.012	0.000	0.004	-0.192	0.850
	(0.108)	(0.163)	(4.253)	(12.271)	(2.772)	(1.004)	(0.007)	(0.000)	(0.529)	(0.000)	(0.021)	(0.005)
U.K.	2.576	0.219	1.831	-22.187	3.797	169.543	0.383	0.012	0.000	0.004	-0.105	0.845
	(0.135)	(0.132)	(4.052)	(11.827)	(3.882)	(8.823)	(0.083)	(0.000)	(0.485)	(0.000)	(0.030)	(0.004)

Note: We estimate a system of equations [(2), (3), (4) and (5)] using the Maximum Likelihood method for each of the local Government bond market jointly with the world Government bond market. Standard errors are displayed in parentheses.

Table 3: Model estimates for each of the local Government bond market jointly with the Eurozone Government bond market

$$r_{i,t} = \theta^{E} \exp\left(K'_{E} Z^{E}_{t-1} \right) cov_{t-1}\left(r_{E,t}, r_{i,t}\right) + \left(1-\theta^{E}\right) \exp\left(\delta'_{L} Z^{L}_{i,t-1} \right) var\left(r_{i,t}\right) + e_{i,t}$$

$$r_{E,t} = \exp\left(K'_{E} Z^{E}_{t-1} \right) var\left(r_{E,t}\right) + e_{E,t}$$

$$\mathbf{H}_{t} = \mathbf{C'C} + \mathbf{A'}\boldsymbol{e}_{t-1}\boldsymbol{e'}_{t-1}\mathbf{A} + \mathbf{B'}\mathbf{H}_{t-1}\mathbf{B}$$

	$K0_E$	**$K1_E$**	**$K2_E$**	**$K3_E$**	**$K4_E$**	**$K5_E$**		**C**	**A**	**B**		
Germany	-282.034	6.968	-8.231	2.878	0.188	-12.081		0.000	0.156	0.812		
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)		(0.000)	(0.001)	(0.001)		
	$\delta0_L$	**$\delta1_L$**	**$\delta2_L$**	**$\delta3_L$**	**$\delta4_L$**	**$\delta5_L$**	**θ^E**	**C11**	**C22**	**C12**	**A22**	**B22**
Austria	0.042	0.456	-6.973	-101.27	-3.530	166.624	0.320	0.005	0.000	0.006	0.089	0.750
	(0.140)	(0.143)	(8.420)	(11.195)	(5.468)	(10.960)	(0.095)	(0.000)	(0.031)	(0.000)	(0.004)	(0.001)
Belgium	-0.747	1.449	-2.915	-93.850	6.511	151.250	0.446	0.005	0.000	0.005	0.004	0.809
	(0.228)	(0.220)	(9.304)	(17.961)	(7.708)	(12.973)	(0.126)	(0.000)	(0.023)	(0.000)	(0.001)	(0.004)
Spain	-221.724	-3.936	0.202	-2.836	1.101	2.831	0.344	0.005	0.000	0.006	0.070	0.758
	($3.568 \cdot 10^{7}$)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.003)	(0.001)
Finland	-48.783	7.830	1.596	-5.830	3.428	0.609	0.149	0.005	0.000	0.006	0.031	0.796
	($2.078 \cdot 10^{9}$)	(951120865)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.003)
France	-47.139	10.942	0.544	-7.688	-0.252	1.977	0.336	0.005	0.000	0.004	0.230	0.859
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.021)	(0.000)	(0.002)	(0.001)
Ireland	-26.544	0.663	1.005	0.026	-0.340	---	0.365	0.005	0.000	0.004	0.171	0.881
	($4.644 \cdot 10^{8}$)	(0.000)	(0.000)	(0.000)	(0.000)	(---)	(0.000)	(0.000)	(0.003)	(0.000)	(0.003)	(0.000)
Italy	-25.503	-3.993	2.637	1.171	5.135	0.603	0.302	0.005	0.000	0.006	0.112	0.722
	($4.239 \cdot 10^{8}$)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.050)	(0.000)	(0.007)	(0.002)
Netherlands	-27.558	-9.929	2.432	-144.99	55.365	1.429	0.627	0.005	0.000	0.006	0.068	0.765
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.034)	(0.000)	(0.003)	(0.002)
Portugal	-30.077	-0.223	0.071	0.938	0.912	---	0.495	0.005	0.000	0.004	0.152	0.877
	($1.685 \cdot 10^{8}$)	(0.000)	(0.000)	(0.000)	(0.000)	(---)	(0.000)	(0.000)	(0.000)	(0.000)	(0.005)	(0.001)
Denmark	-26.485	-0.451	0.496	0.369	0.670	1.333	0.106	0.005	0.000	0.006	0.070	0.758
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.029)	(0.000)	(0.003)	(0.001)
Sweden	-26.247	-3.007	0.931	2.368	0.718	2.440	0.044	0.005	0.000	0.006	0.070	0.758
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.029)	(0.000)	(0.003)	(0.001)
U.K.	-29.729	-1.376	-0.783	61.766	1.209	1.555	0.049	0.005	0.000	0.004	-0.0751	0.916
	($5.712 \cdot 10^{9}$)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.518)	(0.000)	(0.0010)	(0.002)

Note: We estimate a system of equations [(2), (3), (4) and (5)] using the Maximum Likelihood method for each of the local Government bond markets jointly with the Eurozone bond market. Note: Standard errors are displayed in parentheses.

Table 4: Summary statistics for the standardized residuals of the model estimates for each of the local Government bond markets jointly with the world (United States) Government bond market

	Maximum likelihood function value	**Bera-Jarque**	**Q(20)**	**ARCH(20)**
World		3.440 (0.17)	26.025 (0.16)	38.346 (0.01)
Germany	3275.91	17.130 (0.00)	13.368 (0.86)	26.738 (0.14)
Austria	3260.24	20.214 (0.00)	15.209 (0.76)	25.992 (0.16)
Belgium	3282.51	31.298 (0.00)	17.739 (0.60)	33.663 (0.03)
Spain	3280.08	23.861 (0.00)	14.933 (0.78)	22.021 (0.34)
Finland	3269.00	20.520 (0.00)	18.063 (0.58)	27.672 (0.12)
France	3241.14	21.373 (0.00)	14.720 (0.79)	20.281 (0.44)
Ireland	3262.75	29.471 (0.00)	18.990 (0.52)	17.461 (0.62)
Italy	3241.14	21.373 (0.00)	14.720 (0.79)	20.282 (0.44)
Netherlands	3277.97	26.494 (0.00)	17.739 (0.60)	22.818 (0.30)
Portugal	3262.68	27.641 (0.00)	22.062 (0.34)	24.180 (0.23)
Denmark	3235.70	11.115 (0.00)	16.032 (0.17)	36.092 (0.01)
Sweden	3218.97	74.456 (0.00)	16.815 (0.66)	19.232 (0.51)
U.K.	3279.11	71.439 (0.00)	19.273 (0.50)	19.331 (0.50)

Note: The Bera-Jarque statistic tests for the normal distribution hypothesis and has an asymptotic distribution $\chi^2(2)$. Q(20) is the Ljung-Box test for twentieth order serial correlation in the standardized residuals. ARCH(20) is Engle's test for twentieth order ARCH, distributed as $\chi^2(20)$. The p-values of these tests are displayed in parentheses. In all cases the necessary conditions for the stationarity of the process are satisfied.

Table 5: Summary statistics for the standardized residuals of the model estimates for each of the local Government bond markets jointly with the Eurozone (Germany) Government bond market

	Maximum likelihood function value	Normal	Q(20)	ARCH(20)
Germany		15.972 (0.00)	13.645 (0.85)	28.597 (0.09)
Austria	4012.35	23.970 (0.00)	16.769 (0.00)	23.676 (0.26)
Spain	4000.07	21.298 (0.00)	15.820 (0.73)	25.067 (0.20)
Belgium	3977.91	31.645 (0.00)	19.903 (0.46)	42.087 (0.00)
France	3872.16	20.988 (0.00)	14.468 (0.80)	18.372 (0.56)
Finland	4017.46	13.476 (0.00)	18.871 (0.53)	31.067 (0.05)
Ireland	3862.88	27.662 (0.00)	19.039 (052)	16.452 (069)
Italy	3929.65	20.439 (0.00)	13.481 (0.85)	29.673 (0.07)
Netherlands	4008.65	24.039 (0.00)	18.870 (0.53)	27.132 (0.131)
Sweden	3867.52	22.629 (0.00)	19.068 (0.50)	26.124 (0.113)
Portugal	3935.75	24.281 (0.00)	22.758 (0.30)	25.679 (0.18)
Denmak	3712.16	11.435 (0.00)	16.994 (0.65)	41.173 (0.00)
U.K.	3519.85	59.361 (0.00)	13.927 (0.83)	32.623 (0.04)

Note: The Bera-Jarque statistic tests for the normal distribution hypothesis and has an asymptotic distribution $\chi^2(2)$. Q(20) is the Ljung-Box test for twentieth order serial correlation in the standardized residuals. ARCH(20) is Engle's test for twentieth order ARCH, distributed as $\chi^2(20)$. The p-values of these tests are displayed in parentheses. In all cases the necessary conditions for the stationarity of the process are satisfied.

<u>Table 6:</u> Contemporaneous Correlation Coefficients (10-Year Governments' Yields)

First period. 01/03/1994 – 12/31/1998													
	AT	**BE**	**FI**	**FR**	**GR**	**IE**	**IT**	**NL**	**PT**	**SP**	**UK**	**SW**	**DK**
GE	0.99	0.98	0.95	0.98	--	0.98	0.93	0.99	0.93	0.94	0.97	0.95	0.98
AT		0.99	0.97	0.99	--	0.99	0.95	0.99	0.96	0.96	0.95	0.97	0.99
BE			0.98	0.98	--	0.97	0.94	0.99	0.95	0.95	0.92	0.97	0.97
FI				0.97	--	0.95	0.93	0.97	0.95	0.94	0.90	0.98	0.97
FR						0.98	0.97	0.99	0.97	0.97	0.93	0.98	0.99
IE							0.96	0.98	0.96	0.97	0.97	0.97	0.99
IT								0.94	0.99	0.99	0.90	0.97	0.96
NL									0.94	0.95	0.96	0.97	0.98
PT										0.99	0.89	0.97	0.96
SP											0.91	0.98	0.97
UK												0.92	0.96
SW													0.99
Second period. 01/01/1999 – 12/31/2008													
	AT	**BE**	**FI**	**FR**	**GR**	**IE**	**IT**	**NL**	**PT**	**SP**	**UK**	**SW**	**DK**
GE	0.99	0.98	0.99	0.99	0.87	0.96	0.96	0.99	0.98	0.98	0.82	0.94	0.99
AT		0.99	0.99	0.99	0.91	0.99	0.99	0.99	0.99	0.99	0.77	0.91	0.99
BE			0.99	0.99	0.92	0.99	0.99	0.99	0.99	0.99	0.76	0.90	0.99
FI				0.99	0.91	0.99	0.98	0.99	0.99	0.99	0.78	0.92	0.99
FR					0.90	0.98	0.98	0.99	0.99	0.99	0.80	0.92	0.99
GR						0.92	0.91	0.90	0.92	0.92	0.69	0.81	0.92
IE							0.99	0.99	0.99	0.99	0.72	0.88	0.98
IT								0.99	0.99	0.99	0.73	0.87	0.97
NL									0.99	0.99	0.79	0.92	0.99
PT										0.99	0.76	0.89	0.98
SP											0.76	0.90	0.99
UK												0.76	0.78
SW													0.94
Whole sample. 01/03/1994 – 12/31/2008													
	AT	**BE**	**FI**	**FR**	**GR**	**IE**	**IT**	**NL**	**PT**	**SP**	**UK**	**SW**	**DK**
GE	0.99	0.99	0.97	0.99	--	0.98	0.94	0.99	0.95	0.95	0.94	0.96	0.99
AT		0.99	0.97	0.99	--	0.98	0.94	0.99	0.95	0.95	0.92	0.96	0.99
BE			0.99	0.99	--	0.99	0.95	0.99	0.96	0.96	0.93	0.96	0.99
FI				0.98	--	0.98	0.96	0.97	0.98	0.98	0.93	0.98	0.98
FR						0.99	0.96	0.99	0.97	0.97	0.93	0.97	0.99
IE							0.97	0.98	0.97	0.98	0.96	0.97	0.99
IT								0.93	0.99	0.99	0.94	0.98	0.96
NL									0.94	0.95	0.92	0.96	0.99
PT										0.99	0.94	0.98	0.97
SP											0.94	0.98	0.97
UK												0.95	0.95
SW													0.98

<u>Table 7:</u> ADF Unit Root Tests

Testing the integration order of government yields			
	No regressors	With an intercept	With an intercept and a linear time trend
First period. 01/03/1994- 12/31/1998			
Germany	-0.95	0.41	-3.30*
Austria	-1.03	0.55	-3.51**
Belgium	-1.09	0.50	-4.02**
Finland	-0.91	0.13	-3.58**
France	-0.80	0.25	-3.80**
Greece	--	--	--
Ireland	-1.03	0.78	-3.54**
Italy	-1.12	0.56	-2.95
Netherlands	-0.85	0.26	-3.67**
Portugal	-1.58	1.01	-3.02
Spain	-1.17	0.66	-3.30*
United Kingdom	-0.68	0.29	-3.40*
Sweden	-0.81	0.22	-3.78**
Denmark	-0.77	0.18	-3.48**
Second period. 01/01/1999- 12/31/2008			
Germany	-0.59	-1.12	-2.52
Austria	-0.39	-1.46	-2.50
Belgium	-0.35	-1.42	-2.43
Finland	-0.46	-1.35	-2.50
France	-0.46	-1.51	-2.72
Greece	-0.62	-1.55	-0.89
Ireland	-0.14	-1.64	-2.27
Italy	-0.10	-1.79	-2.47
Netherlands	-0.41	-1.46	-2.59
Portugal	-0.28	-1.57	-2.44
Spain	-0.31	-1.47	-2.49
United Kingdom	-0.66	-1.54	-2.37
Sweden	-0.80	-0.51	-2.49
Denmark	-0.59	-1.18	-2.54
Whole sample. 01/03/1994- 12/31/2008			
Germany	-1.14	-0.77	-2.46
Austria	-1.03	-1.04	-2.29
Belgium	-1.14	-1.03	-1.97
Finland	-1.17	-0.93	-1.78
France	-0.95	-0.92	-2.43
Greece	--	--	--
Ireland	-1.00	-1.00	-1.64
Italy	-1.38	-1.01	-1.03
Netherlands	-0.93	-1.05	-2.53
Portugal	-1.93*	-1.16	-0.85
Spain	-1.40	-0.92	-1.22
United Kingdom	-1.00	-0.55	-2.49
Sweden	-1.26	-0.36	-1.92
Denmark	-1.04	-0.76	-2.37

Note: * and ** mean that we can reject the null hypothesis of unit root at the 10 and 5% significance level.

Table 8.1: **Johansen's Multiple Cointegration Tests**

Rank	Eigenvalue	Trace statistic	Max-eigenvalue statistic
First period. 01/03/1994 – 12/31/1998			
0	0.079	529.91	107.55
At most 1	0.064	422.36**	85.81**
At most 2	0.054	336.55**	72.22**
At most 3	0.045	264.34**	59.41
At most 4	0.032	204.93*	42.60
At most 5	0.028	162.33	37.58
At most 6	0.025	124.75	33.60
At most 7	0.022	91.15	29.00
At most 8	0.017	62.15	21.81
At most 9	0.014	40.35	18.64
At most 10	0.010	21.71	13.73
At most 11	0.004	7.98	5.07
At most 12	0.002	2.91	2.91
Second period. 01/01/1999 – 12/31/2008			
0	0.122	1100.656	340.09
At most 1	0.071	760.57**	191.90**
At most 2	0.049	568.67**	131.11**
At most 3	0.044	437.56**	116.76**
At most 4	0.033	320.80**	86.43**
At most 5	0.024	234.37**	63.30**
At most 6	0.019	171.07**	51.00**
At most 7	0.015	120.07**	39.91*
At most 8	0.013	80.16**	35.39**
At most 9	0.009	44.77	24.12
At most 10	0.006	20.66	15.36
At most 11	0.002	5.30	4.00
At most 12	0.001	1.30	1.30
Whole sample. 01/03/1994 – 12/31/2008			
0	0.049	911.15	195.70
At most 1	0.040	715.45**	160.40**
At most 2	0.032	555.05**	128.03**
At most 3	0.031	427.02**	121.38**
At most 4	0.020	305.64**	80.12**
At most 5	0.016	255.52**	61.19**
At most 6	0.014	164.33**	53.97**
At most 7	0.011	110.36**	43.28**
At most 8	0.008	67.08	32.82*
At most 9	0.004	34.26	17.40
At most 10	0.002	16.86	7.49
At most 11	0.002	9.36	7.10
At most 12	0.001	2.27	2.27

Note: * and ** denote that we can reject the null hypothesis of r cointegrating vectors at the 10 and 5% significance level.

Table 8.2: **Johansen's Multiple Cointegration Tests (Including Greece)**

Rank	Eigenvalue	Trace Statistic	Max-Eigenvalue Statistic
Second period. 01/01/1999 – 12/31/2008			
0	0.123	1155.22**	333.49**
At most 1	0.067	821.72**	177.66**
At most 2	0.059	644.07**	153.37**
At most 3	0.047	490.70**	123.16**
At most 4	0.034	367.54**	88.39**
At most 5	0.027	279.15**	65.51**
At most 6	0.022	210.64**	55.37**
At most 7	0.017	155.27**	43.24
At most 8	0.015	112.03**	39.37*
At most 9	0.012	72.66	29.75
At most 10	0.008	42.91	21.27
At most 11	0.006	21.64	14.27
At most 12	0.002	7.36	5.79
At most 13	0.001	1.57	1.57

Note: * and ** denote that we can reject the null hypothesis of r cointegrating vectors at the 10 and 5% significance level.

Table 9.1: Johansen's Multiple Cointegration Tests (EMU Countries)

Rank	Eigenvalue	Trace statistic	Max-eigenvalue statistic
First period. 01/03/1994 – 12/31/1998			
0	0.010	426.25**	136.86**
At most 1	0.062	289.39**	82.82**
At most 2	0.047	206.58**	62.35**
At most 3	0.033	144.23**	43.72
At most 4	0.023	100.51	30.66
At most 5	0.019	69.85	25.04
At most 6	0.016	44.81	21.24
At most 7	0.013	23.57	17.04
At most 8	0.004	6.53	4.94
At most 9	0.001	1.60	1.60
Second period. 01/01/1999 – 12/31/2008			
0	0.113	860.92**	313.41**
At most 1	0.065	547.51**	174.76**
At most 2	0.040	372.75**	106.21**
At most 3	0.031	266.54**	81.15**
At most 4	0.026	185.39**	67.46**
At most 5	0.017	117.93**	45.88**
At most 6	0.014	72.05**	36.52**
At most 7	0.009	35.53	23.68
At most 8	0.004	11.85	9.66
At most 9	0.001	2.19	2.19
Whole sample. 01/03/1994 – 12/31/2008			
0	0.052	755.93**	210.05**
At most 1	0.045	545.87**	179.79**
At most 2	0.030	366.09**	119.66**
At most 3	0.021	246.43**	82.39**
At most 4	0.019	164.04**	74.77**
At most 5	0.013	89.27**	51.29**
At most 6	0.005	37.98	19.43
At most 7	0.003	18.55	11.52
At most 8	0.001	7.03	4.33
At most 9	0.001	2.70	2.70

Note: * and ** denote that we can reject the null hypothesis of r cointegrating vectors at the 10 and 5% significance level.

Table 9.2: Johansen's Multiple Cointegration Tests (EMU Countries, Including Greece)

Rank	Eigenvalue	Trace statistic	Max-eigenvalue statistic
Second period. 01/01/1999 – 12/31/2008			
0	0.115	890.57**	311.47**
At most 1	0.062	579.10**	162.09**
At most 2	0.046	417.02**	118.81**
At most 3	0.032	298.21**	82.24**
At most 4	0.026	215.96**	67.62**
At most 5	0.019	148.35**	48.61**
At most 6	0.015	99.74**	37.48**
At most 7	0.011	62.26**	29.00**
At most 8	0.009	33.26*	21.73*
At most 9	0.003	11.53	8.27
At most 10	0.001	3.26	3.26

Note: * and ** denote that we can reject the null hypothesis of r cointegrating vectors at the 10 and 5% significance level.