Eurozone Debt Crisis and Bond Yields Convergence: Evidence from the New EU Countries

Minoas Koukouritakis*

Department of Economics, University of Crete, Greece

Abstract

This article examines 10-year bond yields convergence between each of the new EU countries and Germany, including a structural break that embodies the effects of the current debt crisis in the Eurozone. The analysis is based on a new definition of bond yields convergence that can be interpreted either as strong or weak monetary policy convergence, depending on whether the conditions of UIP and ex-ante PPP hold or are violated, respectively. The empirical results provide evidence of either strong or weak monetary policy convergence to Germany only for Croatia, the Czech Republic, Lithuania, Romania and Slovakia. For the rest of the new EU countries, the evidence implies lack of monetary policy convergence to Germany. This result could be explained by the increased risk premia in these countries, which caused by the Eurozone debt crisis.

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* Department of Economics, University of Crete, University Campus, Rethymno 74100, Greece. Tel: +302831077411, fax: +302831077404, e-mail: m.koukouritakis@uoc.gr. This study was funded by the Special Account for Research (ELKE – Project KA4449) of the University of Crete.

1. Introduction

Long-term bond yields' convergence between each new EU country and the Eurozone is examined in the present paper, in the framework of the current debt crisis in the Eurozone. As the German dominance was established during the crisis, convergence implies that the long-term bond yield of each new EU country must converge to that of Germany. As shown in this paper, under the conditions of uncovered interest rate parity (UIP) and ex-ante relative purchasing power parity (PPP) long-term bond yield spreads are equal to expected inflation differentials. Thus, evidence of yields' convergence between a new EU country and Germany can be interpreted as monetary policy convergence of this country to Germany. However, lack of yields' convergence does not necessarily imply monetary policy divergence with Germany. There is the possibility that a new EU country has achieved monetary policy convergence to Germany, but its yields to diverge with those of Germany. The reason is that the recent debt crisis in the Eurozone might increase the sovereign default risk of this country and thus, led to large and persistent risk premium. Of course, such information has practical implications regarding the evaluation of each new EU country in order to join the Eurozone.[1] Hence, a proper evaluation of bond yield linkages or, in other words, monetary policy convergence should take the above arguments into account, especially in the period of the debt crisis. Otherwise, invalid conclusions may be drawn.

The empirical literature on interest rate convergence within the EU is extensive, and convergence has been linked to the concepts of unit roots and cointegration in most studies. Among others, Karfakis and Moschos (1990) investigated interest rate linkages between Germany and each of Belgium, France, Ireland, Italy and the Netherlands. Using short rates from the late 1970s to the late 1980s, they found no evidence of long-run interest rates convergence. Evidence against the German leadership hypothesis within the European Monetary System (EMS) for the same period, was also found by Katsimbris and Miller (1993). By including the USA to their sample, they showed that both the US and the German rates have important causal influences on the interest rates of the EMS members. Hafer and Kutan (1994) examined long-run co-movements of short rates and money supplies in a group of five EMS countries from the late 1970s to the early 1990s, and found evidence that implies partial monetary policy convergence.

[1] In fact, Slovenia adopted the euro in January 2007, followed by Cyprus and Malta in January 2008, Slovakia in January 2009, Estonia in January 2011, Latvia in January 2014 and Lithuania in January 2015. All of the remaining new EU countries aspire to apply for Eurozone membership in the future.

Similar evidence was provided by Kirchgässner and Wolters (1995), who used money market rates from mid-1970s to mid-1990s, and showed that Germany has a strong long-run influence within the EMS. Haug *et al.* (2000) tried to determine which of the twelve original EU countries would form a successful monetary union based on the nominal convergence criteria of the Treaty on European Union (TEU). Using data from 1979 to 1995, they found that the formation of a successful monetary union would require significant adjustments in fiscal and monetary policies by several of these countries.

Camarero *et al.* (2002) investigated convergence of long-term interest rate differentials for the EU countries in relation to the TEU criterion, using 10-year bond yields from 1980 to mid-1990s. Departing from the literature, they adopted the definitions of long-run convergence of per capital output and catching-up convergence (Bernard and Durlauf, 1995, 1996),[2] and accounted for structural breaks in the data using the one-break unit root test of Perron (1997). They showed that six countries satisfied the criterion of long-run convergence, seven countries satisfied the conditions of catching-up convergence, and only Italy did not converge in either sense. Holtemöller (2005) studied the degree of monetary integration to the Eurozone for Greece and the Central and Eastern European EU countries, based on interest rate spreads and ex-post deviations from the UIP. Using interbank rates from mid-1990s to the early 2000s, his evidence implied high degree of monetary integration for Estonia and Lithuania, medium degree of monetary integration for Greece and Slovakia, and low degree of monetary integration for the Czech Republic, Hungary, Latvia, Poland and Slovenia.

Jenkins and Madzharova (2008) investigated real interest rate convergence for the original EU countries, using 10-year bond yields from the late 1990s to mid-2000s. Their evidence implied failure of the real interest rate parity, mainly due to inflation rate differences. Gabrisch and Orlowski (2010) departed from cointegration analysis and applied GARCH methodology in order to investigate interest rate convergence for the Czech Republic, Hungary, Poland, Slovakia and Slovenia in relation to the Eurozone yields. They focused on 10-year bond yields from the early to the late 2000s and found evidence of stronger convergence for the Czech Republic, Slovenia, and Poland, in which the macroeconomic fundamentals are solid and the financial markets are stable, and weaker convergence for Hungary and Slovakia. Frömmel and Kruse

[2] Long-run convergence exists when the long-term forecasts of interest rates are equal and catching-up convergence is interpreted as the cointegration between the interest rates along a deterministic time trend.

(2015) studied interest rate convergence by implementing a changing persistence model for Belgium, France, Italy and The Netherlands in relation to Germany as the reference country. Using 3-month treasury bill rates from the early 1980s to the late 2000s, they found evidence of very different convergence periods for the sample countries, and showed that fiscal and monetary policy coordination were the main factors that led to interest rate convergence.

Several limitations of the existing studies can be pointed out, which may have affected the reported results. Firstly, most of the aforementioned studies, with the exception of Camarero *et al.* (2002), did not account for structural shifts in the data. Secondly, the existing studies have not distinguished in a systematic way between stochastic and deterministic trends in the structure of interest rates. This is an important issue because evidence of cointegration between, for example, two interest rates implies the presence of a single common stochastic trend that ties them in the long run. On the other hand, deterministic trends depend on the underlying process that generates the stochastic variables under study. Thus, for two interest rates it is not enough to cointegrate with cointegrating vector $(1,-1)$; it is also required that they are cotrended, so that the deterministic trends cancel out in the differential of the two series. Thirdly, in most of the existing studies, interest rate convergence has been examined without an explicit formal definition of convergence or a data generation process (DGP) for the interest rates. The above omissions make the interpretation of the empirical results less transparent and informative.

The present study attempts to deal with these considerations. Firstly, consistent with the Eurozone's nominal convergence criteria, this study focuses on nominal 10-year bond yields' convergence between each new EU country and Germany, in the framework of an explicit DGP for bond yields and a new definition of convergence that allows for a constant non-negative deviation in each pair of bond yields. The inclusion of these elements leads to explicit testable cointegration and cotrending restrictions that makes the interpretation of the econometric results more informative and meaningful. Furthermore, under the UIP and PPP conditions, deviations from yields' parity are equal to expected inflation differentials. Such deviations can be eliminated in the long run, if monetary authorities (or market forces) in each new EU country contribute in establishing common deterministic and stochastic trends with Germany, regarding the long-term yields or expected inflation rates. This case can be interpreted as *strong convergence* with Germany, which more than satisfies the TEU criterion for yields' convergence. On the other hand, if the UIP and PPP conditions do not hold due to time-varying stationary risk premia, different tax rates (Mark, 1985) or transactions costs (Goodwin and Grennes, 1994) across countries, yields convergence can be defined broader as *weak convergence*, in which yields converge to a non-negative constant. If this constant is less than 2%, the TEU criterion is also satisfied. Hence, the empirical results are interpreted in terms of strong or weak monetary policy convergence between each new EU country and Germany.

Secondly, I employ the cointegration test developed by Lütkepohl, Saikkonen and Trenkler in several papers noted below, in order to capture possible structural shifts in the data. The omission of such shifts in the data when they actually exist can distort substantially standard inference procedures for cointegration. In this analysis, such shifts cannot be omitted as the current debt crisis in the Eurozone has probably altered the deterministic components of the new EU countries' yields. In addition, as the deterministic components of yields are assumed to be independent of the stochastic components, the Gonzalo and Granger (1995) methodology for estimating and testing for the common stochastic trend in each pair of yields has been implemented.

The rest of the paper is organised as follows. Section 2 defines yields' convergence and relates it to monetary policy convergence, using the conditions of UIP and PPP. Section 3 discusses the cointegration methodology in the presence of structural shifts in the data, along with the common trends test. Section 4 describes the data, analyses the empirical results and provides some policy implications. Finally, Section 5 contains some concluding remarks.

2. Yields' Convergence with Structural Breaks

The TEU nominal convergence criterion regarding interest rates requires that the 10-year bond yield of a Eurozone candidate country must converge to a level that is less than 2% of the average 10-year bond yield of the three Eurozone countries with the lowest inflation rates. In this analysis, Eurozone is proxied by Germany as its dominance in the Eurozone was established during the current debt crisis. Apart from the debt crisis, there may be several reasons that the 10-year yields of the new EU countries will not converge to the Eurozone criterion, even in the long run. Transaction costs, different tax rates or failures of the UIP and PPP conditions, may create a 'band of inaction' within which there are no arbitrage opportunities for long-term bonds issued by different countries. In addition, differences in the fiscal positions of the Eurozone countries may cause a wedge in yields. The above considerations are taken into account in the definition of convergence that follows as I allow for a non-negative constant gap $c \geq 0$ between the 10-year yields of each new EU country and Germany.

Hence, convergence exists if $\lim_{k \to \infty} E\left[\left(r_{i,t+k} - r_{G,t+k}\right) | I_t\right] = c$ at any fixed time t and at all horizons $k = 1, 2, ...$, where r_i is the 10-year yield of a new EU country i, r_G is the German 10-year yield and I_t is the information set at time t. *Strong convergence* between these two yields exists when $c = 0$, while *weak convergence* exists when $c > 0$. This definition states that the yields will converge, if their long-term forecasts differ by a non-negative constant. If the yields are $I(1)$, convergence requires cointegration with cointegrating vector $(1, -1)$. Furthermore, if the yields have deterministic trends, they should also be cotrended, so that their differential has no deterministic trends.[3] The above definition is satisfied, if it is probably restricted, by the following data generation process (DGP) for the long-term yield r of any new EU country i:

$$r_{i,t} = \mu_{i,t} + \tilde{r}_{i,t}, \ \tilde{r}_{i,t} = b\tilde{r}_{i,t-1} + u_{i,t}, \tag{1}$$

where $\mu_{i,t}$ is the deterministic component possibly with structural breaks, $\tilde{r}_{i,t}$ is the stochastic component and u_t is an error term. It is clear that if $b = 1$, then $r_{i,t}$ will be an $I(1)$ process. The DGP in equation (1) has been used, among others, by Bhargava (1986) and Schmidt and Phillips (1992) for studying non-stationary time series with no structural breaks. The cointegration test with structural breaks that is used in this paper and analysed in the next section adopts similar representations.

Equation (1) implies that the deterministic component of $r_{i,t}$ is independent of and not affected by its stochastic component. As Schmidt and Phillips (1992) indicate, this property allows for an unambiguous interpretation of the parameters of the DGP. Also, the DGP in equation (1) is economically plausible, because domestic policy actions or other exogenous international events, such as the Eurozone debt crisis, affect directly the deterministic component but not the stochastic component of $r_{i,t}$. The latter is more likely to be influenced by market forces, perceptions of each new EU country's risk, expectations about future government policies and their credibility, yield movements in the dominant economy of Germany. Moreover, the

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[3] This definition is inspired by per capita income convergence of Bernard and Durlauf (1995), which assumes $c = 0$. Pesaran (2007) considers the case of $c \neq 0$ and deals explicitly with the cointegration and cotrending restrictions.

definition of yields' convergence imposes several restrictions on the above DGP. Let $\mu_{i,t} = d_{i,0} + d_{i,1}t + d_{i,2}D_t$, where t is a time trend and D_t is a dummy variable corresponding to a level shift in $\mu_{i,t}$ at some specific time T_B. Using equation (1), one can obtain:

$$E\left[\left(r_{i,t+k} - r_{G,t+k}\right) | I_t\right] = \left(d_{i,0} - d_{G,0}\right) + \left(d_{i,1} - d_{G,1}\right)\left(t+k\right) + \left(d_{i,2} - d_{G,2}\right)D_{t+k} + E\left[\left(\tilde{r}_{i,t+k} - \tilde{r}_{G,t+k}\right) | I_t\right]. \quad (2)$$

Thus, for yields' convergence to be realised, the following restrictions on the parameters of equation (2) must hold: (i) $d_{i,0} - d_{G,0} \geq 0$ if $D_{t+k} = 0$ and $d_{i,0} - d_{G,0} + d_{i,2} - d_{G,2} \geq 0$ if $D_{t+k} = 1$, (ii) $d_{i,1} - d_{G,1} = 0$, and (iii) $E\left[\left(\tilde{r}_{i,t+k} - \tilde{r}_{G,t+k}\right) | I_t\right] = 0$. Restriction (i) is easily satisfied as the yield of each new EU country is larger, in general, than the German yield. Restrictions (ii) and (iii) imply cotrending and cointegration, respectively.

Following the above, I test sequentially for convergence between the yields of each new EU country and Germany as follows: (i) if cointegration exists and in this case, if the cointegrating vector $(1,-1)$ spans the cointegration space, (ii) conditional on (i), if the pairs of yields are cotrended, and (iii) if the regression constant and the level shift in yields are jointly less than 2%, as stated by the TEU criterion. With the absence of transaction costs in asset markets, different tax rates and different fiscal positions across countries, restriction (i) should hold with equality, along with restrictions (ii) and (iii). Hence, 10-year yields should be equalised across countries in the long run and converge strongly. In this case, strong convergence more than satisfies the TEU criterion as the latter allows for a 2% yield differential.

The above definition of convergence also accommodates deviations from the UIP condition:

$$r_{i,t} - r_{G,t} = E\left(\Delta S_{i,t} | I_t\right), \quad (3)$$

and the ex-ante relative PPP condition:

$$E\left(\Delta S_{i,t} | I_t\right) = E\left[\left(\pi_{i,t} - \pi_{G,t}\right) | I_t\right], \quad (4)$$

where, $S_{i,t}$ is the logarithm of the nominal exchange rate (the domestic price of the foreign currency), $\pi_{i,t}$ is the inflation rate of a new EU country i and $\pi_{G,t}$ is the German inflation rate. Substituting equation (4) into equation (3), one gets:

$$r_{i,t} - r_{G,t} = E\left[\left(\pi_{i,t} - \pi_{G,t}\right) | I_t\right]. \quad (5)$$

Equation (5) implies that the 10-year yield of a new EU country i will converge to that of Germany in the long run, if the expected inflation rate of this new EU country converges to that of Germany, or alternatively, if the monetary policy of this new EU country converges to the German monetary policy in the long run.[4] On the other hand, evidence of yields' divergence for a new EU country could be attributed to the probability of large and persistent risk premium due to the Eurozone debt crisis.

3. Cointegration with Structural Breaks

As noted in the introductory section, structural shifts in the data can distort substantially standard inference procedures for cointegration. Thus, it is necessary to account for possible breaks in the data before inference on cointegration can be made. There is a recent large literature on different approaches and techniques for cointegration testing in the presence of structural breaks in the data.[5] For reasons of consistency as the deterministic trends are treated as independent of the stochastic trends in the present paper, I implement the approach developed by Lütkepohl and his co-authors (Lütkepohl and Saikkonen, 2000; Saikkonen and Lütkepohl, 2000; Trenkler *et al.*, 2008). This approach assumes that in the data generating process (DGP) for a vector-valued process y_t, its deterministic part (μ_t) does not affect its stochastic part (X_t). Thus, the deterministic part can be removed in the first stage and a likelihood ratio (LR) cointegration test can be applied on the detrended stochastic part of y_t in the second stage.

Briefly, consider the case of a single exogenous break at time T_B in μ_t, in both the level and the trend of y_t. In this case, the DGP for y_t is

$$y_t = \mu_t + X_t = \mu_0 + \mu_1 t + \delta_0 b_t + \delta_1 d_t + X_t, \quad t = 1,....,T, \tag{6}$$

where t is a linear time trend, μ_i $(i = 0,1)$ and δ_i $(i = 0,1)$ are unknown $(v \times 1)$ parameter vectors, b_t and d_t are dummy variables defined as $b_t = d_t = 0$ for $t < T_B$, and $b_t = 1$ and $d_t = t - T_B + 1$ for

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[4] If expected inflation differential converges to a small non-negative constant π_0, the addition of a stationary 'risk premium' in equation (5) of the form $u_t = \rho_0 + \rho(L)u_{t-1} + v_t$, where $\rho(L)$ is a m-order polynomial in the lag operator L and v_t is a zero mean stochastic process, in order to reflect imperfect substitutability of bonds will still be consistent with the definition of weak convergence.

[5] Perron (2006) provides a comprehensive review of this literature.

$t \geq T_B$. The unobserved stochastic component X_t is assumed to follow a $VAR(k)$ process with the following VECM representation:

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \varepsilon_t, \quad \varepsilon_t \sim iidN(0,\Omega), \quad t = 1,...,T. \tag{7}$$

It is also assumed that the components of X_t are at most $I(1)$ and cointegrated (*i.e.*, $\Pi = \alpha\beta'$) with cointegrating rank r_0. Based on the DGP described in equations (6) and (7), one obtains estimates of μ_0, μ_1, δ_0 and δ_1 using a feasible GLS procedure under the null hypothesis $H_0(r_0): rank(\Pi) = r_0:$ vs. $H_1(r_0): rank(\Pi) > r_0$. Using these estimates, the detrended series $\hat{X}_t = y_t - \hat{\mu}_0 - \hat{\mu}_1 t - \hat{\delta}_0 d_t - \hat{\delta}_1 b_t$ are computed and replacing X_t in the VECM of equation (7). Then, the following LR statistic is computed:

$$LR_{LST} = -T \sum_{i=r_0+1}^{p} \ln\left(1 - \tilde{\lambda}_i\right), \tag{8}$$

where the eigenvalues $\tilde{\lambda}_i$'s are obtained by solving a generalised eigenvalue problem, along the lines of Johansen (1988). Asymptotic results and p-values were derived by Trenkler *et al.* (2008), using response surface techniques. These authors also showed that the asymptotic distribution of the LR statistic in equation (8) depends on the break point location.

Regarding common trends, Gonzalo and Granger (1995) identified, estimated and tested for the significance of common trends in a system of time series. They exploited the duality between cointegration and common trends in a VECM framework, in the sense that if there are r_0 cointegrating vectors in a $p-$dimensional vector of $I(1)$ variables, then there will be $p - r_0$ common trends that induce shifts in the cointegrating relations within the cointegration space. They also showed that the common trends in the zero mean stochastic process X_t are simply the cumulated disturbances $\alpha_\perp' \sum_{i=1}^{t} \varepsilon_t$, where $\alpha_\perp$ is a $p \times (p - r_0)$ matrix that is the orthogonal complement of α (Johansen, 1995, p. 41). By assuming that the common trends are a linear combination of X_t in the form of $f_t = \alpha_\perp' X_t$, one can test for them in different linear combinations of X_t. The null hypotheses is $H_0: \alpha_\perp = G\theta$, where G is a $p \times m$ known matrix of constants and θ is an $m \times (p - r_0)$ matrix of unknown coefficients, such that $p - r_0 \leq m \leq p$. To perform the test, one solves two eigenvalue problems under the null and the alternative hypotheses and obtains the eigenvalues $1 > \hat{\lambda}_1^* > ... > \hat{\lambda}_m^* > 0$ and $1 > \hat{\lambda}_1 > ... > \hat{\lambda}_p > 0$, respectively. The LR statistic for testing H_0 is given by:

$$L = -T\sum_{i=r_0+1}^{p} \ln\left[\left(1-\hat{\lambda}_{i+(m-p)}^*\right)\Big/\left(1-\hat{\lambda}_i\right)\right], \tag{9}$$

which under H_0 is distributed as $\chi^2_{(p-r_0)\times(p-m)}$ asymptotically.

4. Data and Empirical Results

4.1 Data

The data set consists of annualised monthly observations for 10-year government bond yields for each new EU country and Germany. Estonia was left out of the analysis, because Estonian long-term bonds are issued only occasionally and thus, their yields are not disseminated. The time span for each country begins in 1999:01 with the establishment of the Eurozone, or later due to data availability, and ends in 2014:09. Data details and their sources are reported in Table 1.

4.2 Unit Root Tests Results

Before testing for cointegration, I tested each yield for a unit root using the ADF, DF-GLS and KPSS unit root tests. Columns 2 and 3 of table 2 report the results for the ADF and DF-GLS tests, respectively, which both indicate that the unit root hypothesis cannot be rejected for any yield at the 5 per cent level of significance. The results for the KPSS test are presented in column 4 of table 2. Similarly, they provide evidence that the null hypothesis of covariance stationarity is rejected for all yields.

4.3 Convergence of Monetary Policies

This section examines the possibility of monetary policy convergence between each new EU country and Germany by investigating long-run linkages in bond yields and testing for the restrictions implied by the analysis of section 2. For each new EU country i a two-dimensional VECM for $y_t = \left(r_{i,t}, r_{G,t}\right)$ has been used, consisting of the 10-year bond yields of this country and Germany. Initially, a cointegration test for these two yields is applied. If cointegration exists, then I test if the cointegrating vector $\left(1,-1\right)$ spans the cointegration space. Secondly, conditional on the cointegrating vector being $(1,-1)$, I test for cotrending and examine both strong and weak monetary policy convergence.

Furthermore and before testing for cointegration, the detection of the structural breaks that are included in the VECMs is crucial. As suggested by economic theory and indicated by Koukouritakis (2013), these breaks have to be detected exogenously and, of course, must be based on specific economic events that affected the sample countries. Hence for all VECMs, a single break is allowed to be at the beginning of the current financial and debt crisis. According to the U.S. National Bureau of Economic Research, the financial crisis began in December 2007. Figure 1 reports the yields for each sample country, along with the structural shift. One can easily observe that from 2007 onwards, all yields show higher volatility, reflecting the fiscal deficit and sovereign debt problems that several new EU countries faced.

4.3.1 Testing the cointegration hypothesis

For each new EU country, the model described in equations (6) and (7) was estimated. Then, the LR test statistics and the corresponding response surface p-values were computed.[6] The lag length for each VECM was selected using the Akaike information criterion (AIC). Table 3 reports the cointegration results. As shown in the third and fourth column of this table, the 10-year German yield is cointegrated only with the 10-year yield of each of Croatia, the Czech Republic, Lithuania, Romania and Slovakia. In contrast, there is no evidence of cointegration between the 10-year German yield and the 10-year yield of each of Bulgaria, Cyprus, Hungary, Latvia, Malta, Poland and Slovenia. Next, for the five countries for which there is evidence of cointegration, two separate tests were performed. Firstly, I tested the null hypothesis that the cointegrating vector linking the pairs of the 10-year bond yields is $(1,-1)$. Under the null hypothesis, this test is distributed asymptotically as χ_1^2 (Johansen, 1995, p. 104). As shown in the column 6 of table 3, this hypothesis is not rejected for all five countries. Secondly, I tested the null hypothesis that the German 10-year yield is the shared common trend. Column 7 of table 3 gives the *L*-statistics for the null hypothesis that the matrix $G=(0,1)$. As shown, this null hypothesis is not rejected in any case. These results provide significant empirical support for the necessary condition of monetary policy convergence for each of Croatia, the Czech Republic,

[6] The author is grateful to Carsten Trenkler for kindly providing him with the GAUSS codes.

Lithuania, Romania and Slovakia to Germany. Alternatively, Germany (as the dominant country of the Eurozone) sets the long-term trend for expected inflation, and these five new EU countries tend to adjust their monetary policies in order to achieve an expected inflation rate consistent with that of Germany.[7]

Furthermore, the estimated residuals of each VECM were checked for *s*-order serial correlation, using the multivariate versions of the Lung-Box Q and LM tests. Under the null hypothesis of no serial correlation in the error term of the VECM, these test statistics are asymptotically distributed as χ^2 with degrees of freedom $p^2(s-k)$ and p^2, respectively (Johansen, 1995, p. 22). The computed test statistics and associated p-values are reported in table 4. As shown, both the Q and LM tests do not reject the null hypothesis of no serial correlation in the estimated residuals, in all cases.

4.3.2 Testing the cotrending hypothesis and the significance of the constant term

As it was discussed in section 2, yields' convergence requires not only that a pair of yields is cointegrated with cointegrating vector $(1,-1)$, but also that it is cotrending. The latter means that yield spreads have no deterministic trends but they may have a non-negative constant term, including the level shifts, where applicable. If this constant term is insignificantly different from zero, this implies strong convergence and the TEU criterion is more than satisfied. Otherwise, if this constant term is significantly different from zero but insignificantly different from 2%, then weak convergence has been achieved and the TEU criterion for yields' convergence is also satisfied.

Figure 2 reports the yield spreads in relation to the 10-year German yield, for the five countries that their yields are cointegrating with the German yield with cointegrating vector $(1,-1)$. As shown, there is evidence of different trending behaviour for these five countries. A formal test for the cotrending hypothesis in each of these five spreads suggests the regression of each yield spread on an intercept, a linear trend and the respective level and trend shifts using appropriate dummy variables. In each regression, I also included as many lags of the yield spread

[7] As a robustness check for this evidence I also used alternative dates for the structural break, such as the EU or the Eurozone accession date for each new EU country. The cointegration and cotrending results are similar to those reported in this paper.

as necessary for making the residuals white noise. The test results are reported in table 5. As shown in column 7 of this table, the cotrending hypothesis is not rejected for any of the yield spreads of Croatia, the Czech Republic, Lithuania, Romania and Slovakia, at the 5 per cent level of significance. This hypothesis was tested using an *F*-test on the null hypothesis that the linear trend and the trend shift are jointly zero, in each case. Consequently, there is strong statistical evidence of weak monetary policy convergence between each of these five new EU countries and Germany, as far as deterministic cotrending in the 10-year yield spreads is concerned.

Concerning the significance of the constant terms for each of the above five countries, Columns 2 and 3 of table 5 indicate the intercept and the level shift of the yield spreads are statistically insignificant in the cases of the Czech Republic, Lithuania and Slovakia. Based on these results, one can conclude that these three countries have achieved not only weak but also strong monetary policy convergence to Germany, since the TEU criterion is more than satisfied. This is an expected result not only because Lithuania and Slovakia are already Eurozone members, but also because Germany plays a very important role in the economies of these three countries. For the yield spreads of Croatia and Romania, for which there is also evidence of cotrending, the results indicate statistical significance of the level shift coefficient in the former country and of the intercept in the latter country. Hence, in order to determine if these two countries have achieved weak monetary policy convergence, I performed an additional *t*-test on the sum of the intercept and the level shift coefficient being greater than or equal to 2%, against the alternative of being less than 2%. Column 8 of table 5 reports the respective *t*-statistics and indicates rejection of the null hypothesis, at the 5 per cent level of significance. Thus, there is evidence that also Croatia and Romania satisfy the TEU criterion for monetary policy convergence.

4.3.3 Policy implications

In the framework of the debt crisis in the Eurozone, the results reported in tables 3 and 5 indicate that even though Germany is the dominant country in the Eurozone and sets the macroeconomic policies, seven new EU countries, namely Bulgaria, Cyprus, Hungary, Latvia, Malta, Poland and Slovenia (regardless if they are Eurozone members or not) are unable to follow these policies. Even though these new EU countries (a) managed to stabilise their exchange rates during the last decade[8], (b) adopted implicit or explicit inflation targeting polices in order to fight inflation, (c) implemented tight fiscal policies in order to reduce fiscal deficit and public debt, and (d) promoted structural reforms designed to support growth, the Eurozone debt crisis harmed their economies significantly. Especially for Cyprus, Latvia and Slovenia, these results do not necessarily imply monetary policy divergence with Germany. These countries are Eurozone members and their monetary policies are no different from that of Germany. Lack of yields' convergence could probably be attributed to the increased sovereign default risk of these three countries due to the Eurozone debt crisis, which in turn led to large and persistent risk premia. More specifically, the crisis had negative effects on the economic growth of these countries. These effects led Latvia to agree for rescue package with the EU and the International Monetary Fund (IMF) in 2008, while Cyprus and Slovenia were downgraded by the Credit Rating Agencies in 2011. Furthermore, due to the default of its commercial banks in 2013 that led to the need for bailout funds from the EU and the IMF, Cyprus had to proceed to a 'haircut' in bank deposits. It is also worth noting that the credit ratings of the remaining new EU countries remain at moderate risk. On the other hand, the evidence of yields' divergence for Bulgaria, Hungary and Poland, which are not yet Eurozone members, could probably be attributed to expected inflation differentials, as mentioned in section 2.

5. Concluding Remarks

Long-run bond yields' convergence between each new EU country and Germany was investigated in the present paper, in the framework of the Eurozone debts crisis. Because these bond yields are random walks with structural shifts over the sample period, I evaluated these issues using cointegration and cotrending analysis including structural shifts in the data.

The cointegration and cotrending analysis provides useful insights about the degree of monetary policy convergence of each new EU country to Germany, whose dominance in the Eurozone was established during the debt crisis. Based on the empirical results, there is some clear evidence of strong monetary policy convergence for each of the Czech Republic, Lithuania and Slovakia to Germany. Alternatively, under the UIP and ex-ante relative PPP conditions, the expected inflation rate of these three countries has converged to the expected inflation rate of

[8] Cyprus, Latvia, Malta and Slovenia joined the ERM II, Hungary pegged its currency to the euro, Poland implemented a free-floating exchange rate regime, while Bulgaria adopted a euro-based currency board.

Germany. This is an expected result not only because Lithuania and Slovakia are already Eurozone members, but also because Germany plays a very important role in the economies of these three countries. Furthermore, the empirical results provide evidence of weak monetary policy convergence for each of Croatia and Romania to Germany. In contrast, for the remaining seven new EU countries, namely Bulgaria, Cyprus, Hungary, Latvia, Malta, Poland and Slovenia, the empirical evidence suggests yields' divergence for each of these countries in relation to Germany. For Cyprus, Latvia and Slovenia, which as Eurozone members they have common monetary policy with Germany, the empirical evidence could probably be attributed to the increased sovereign default risk of these countries, which in turn led to large and persistent risk premia.

In summary, the empirical evidence indicates that in the context of the Eurozone debt crisis, even though Germany has established its dominance and sets the macroeconomic policies in the Eurozone, several new EU countries are unable to follow these policies. And this conclusion addresses once more the issue of core-periphery in the Eurozone and, thus, the Eurozone's future prospects.

References

Bernard, A.B. and Durlauf, S. (1995). 'Convergence in International Output', Journal of Applied Econometrics, Vol 10, No. 2, pp. 97-108.

Bernard, A.B. and Durlauf, S. (1996). 'Interpreting Tests of the Convergence Hypothesis', Journal of Econometrics, Vol. 71, No. 1-2, pp. 161-173.

Bhargava, A. (1986). 'On the Theory of Testing for Unit Roots in Observed Time Series', Review of Economic Studies, Vol. 53, No. 3, pp. 369–384.

Camarero, M., Ordonez, J. and Tamarit, C.R. (2002). 'Tests for Interest Rate Convergence and Structural Breaks in the EMS: Further Analysis', Applied Financial Economics, Vol. 12, No. 6, pp. 447-456.

Frömmel, M. and Kruse, R. (2015). 'Interest Rate Convergence in the EMS prior to European Monetary Union', Journal of Policy Modeling, Vol. 37, No. 6, pp. 990-1004.

Gabrisch, H. and Orlowski, L.T. (2010). 'Interest Rate Convergence in Euro-Candidate Countries: Volatility Dynamics of Sovereign Bond Yields', Emerging Markets Finance and Trade, Vol. 46, No. 6, pp. 69-85.

Gonzalo, J. and Granger, C.W.J. (1995). 'Estimation of Common Long-Memory Components in Cointegrated Systems', Journal of Business and Economics Statistics, Vol. 13, No. 1, pp. 27-35.

Goodwin, B.K. and Grennes, T.J. (1994). 'Real Interest Rate Equalization and Integration of Financial Markets', Journal of International Money and Finance, Vol. 13, No. 1, pp. 107-124.

Hafer, R.W. and Kutan, A.M. (1994). 'A Long Run View of German Dominance and the Degree of Policy Convergence in the EMS', Economic Inquiry, Vol. 32, No. 4, pp. 684-695.

Haug, A.A., MacKinnon, J.G. and Michelis, L. (2000). 'European Monetary Union: A Cointegration Analysis', Journal of International Money and Finance, Vol. 19, No. 3, pp. 419-432.

Holtemöller, O. (2005). 'Uncovered Interest Rate Parity and Analysis of Monetary Convergence of Potential EMU Accession Countries', International Economics and Economic Policy, Vol. 2, No. 1, pp. 33-63.

Jenkins, M.A. and Madzharova, P. (2008). 'Real Interest Rate Convergence under the Euro' Applied Economics Letters, Vol. 15, No.6, pp. 473-476.

Johansen, S. (1988). ‘Statistical Analysis of Cointegration Vectors’, Journal of Economic Dynamics and Control, Vol. 12, No. 2-3, pp. 231-254.

Johansen, S. (1995). Likelihood-based Inference in Cointegrated Vector Autoregressive Models, Oxford, Oxford University Press.

Karfakis, C.J. and Moschos, D.M. (1990). ‘Interest Rate Linkages within the European Monetary System: A Time Series Analysis’, Journal of Money Credit and Banking, Vol. 22, No. 3, pp. 388-394.

Katsimbris, G.M. and Miller, S.M. (1993). ‘Interest Rate Linkages within the European Monetary System: Further Analysis’, Journal of Money Credit and Banking, Vol. 25, No. 4, pp. 771-779.

Kirchgässner, G. and Wolters, J. (1995). ‘Interest Rate Linkages in Europe before and after the Introduction of the European Monetary System’, Empirical Economics, Vol. 20, No. 3, pp. 435-454.

Koukouritakis, M. (2013). ‘Expectations Hypothesis in the Context of Debt Crisis: Evidence from Five Major EU Countries’, Research in Economics, Vol. 67, No. 3, pp. 243-258.

Kwiatkowski, D., Phillips, P.C.B., Schmidt, P., Shin, Y. (1992). ‘Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root’, Journal of Econometrics, Vol. 54, No. 1-3, pp. 159-178.

Lütkepohl, H. and Saikkonen, P. (2000). ‘Testing for the Cointegrating Rank of a VAR Process with a Time Trend’, Journal of Econometrics, Vol. 95, No. 1, pp. 177-198.

Mark, N. (1985). ‘Some Evidence on the International Inequality of Real Interest Rates’, Journal of International Money and Finance, Vol. 4, No. 2, pp. 189-208.

Perron, P. (1997). ‘Further Evidence on Breaking Trend Functions in Macroeconomic Variables’, Journal of Econometrics, Vol. 80, No. 2, pp. 355-385.

Perron, P. (2006). ‘Dealing with structural breaks’, in Patterson, K. and Mills T.C. (eds), Palgrave Handbook of Econometrics, Vol. 1: Econometric Theory, New York: Palgrave Macmillan, pp. 278-352.

Pesaran, M. (2007). ‘A Pair-Wise Approach to Testing for Output and Growth Convergence’, Journal of Econometrics, Vol. 138, No. 1, pp. 312-355.

Saikkonen, P. and Lütkepohl, H. (2000). ‘Testing for the Cointegrating Rank of a VAR Process with Structural Shifts’, Journal of Business and Economics Statistics, Vol. 18, No. 4, pp. 451-464.

Schmidt, P. and Phillips, P.C.B. (1992). 'LM Tests for a Unit Root in the Presence of Deterministic Trends', Oxford Bulletin of Economics and Statistics, Vol. 54, No. 3, pp. 257-287.

Trenkler, C., Saikkonen, P. and Lütkepohl, H. (2008). 'Testing for the Cointegrating Rank of a VAR Process with Level Shift and Trend Break', Journal of Time Series Analysis, Vol. 29, No. 2, pp. 331-358.

Table 1: Sample of 10-year government bond yields

Country	Time span
Bulgaria	2002:04-2014:09
Croatia	2005:12-2014:09
Cyprus	1999:01-2014:09
Czech Republic	2000:04-2014:09
Hungary	1999:01-2014:09
Latvia	2001:01-2014:09
Lithuania	2001:01-2014:09
Malta	2000:01-2014:09
Poland	1999:05-2014:09
Romania	2005:04-2014:09
Slovakia[a]	1999:01-2014:09
Slovenia	2002:03-2014:09
Germany	1999:01-2014:09

Notes: Almost all of the data are central government bond yields on the secondary market, gross of tax, with a residual maturity of around 10 years. Only for Cyprus primary market yields are reported, while the same applies to Bulgaria and Romania up to 12:2005, Slovenia up to 10:2003 and Lithuania up to 10:2007. Data were obtained by the Eurostat. [a] For the period 1999:01-2000:8, government bond yields for Slovakia were obtained by the National Bank of Slovakia, as the Eurostat data series begins at 2000:9. All data are period average.

Table 2: ADF, DF-GLS and KPSS unit root tests

10-year bond yield	ADF t-stat.	DF-GLS t-stat.	KPSS LM-stat.
	Intercept and trend		
Bulgaria	-2.19 (0.490)	-1.56	0.33*
Croatia	-2.02 (0.582)	-1.80	0.52*
Cyprus	-1.65 (0.768)	-1.46	0.70*
Czech Republic	-2.30 (0.433)	-2.24	0.17*
Hungary	-2.21 (0.480)	-2.02	0.26*
Latvia	-2.19 (0.489)	-1.96	0.61*
Lithuania	-1.84 (0.680)	-1.69	0.56*
Malta	-1.68 (0.756)	-1.68	0.19*
Poland	-2.34 (0.409)	-2.24	0.23*
Romania	-1.82 (0.690)	-1.70	0.26*
Slovakia	-2.44 (0.357)	-0.82	0.30*
Slovenia	-1.81 (0.697)	-1.38	0.42*
Germany	-2.67 (0.249)	-1.58	0.21*
	Intercept		
Bulgaria	-2.26 (0.188)	-0.26	0.73*
Croatia	-2.05 (0.267)	-1.51	0.53*
Cyprus	-2.08 (0.253)	-1.07	0.95*
Czech Republic	-1.35 (0.607)	0.33	1.09*
Hungary	-1.90 (0.332)	-0.29	0.80*
Latvia	-2.23 (0.196)	-1.55	0.58*
Lithuania	-1.83 (0.364)	-0.76	0.58*
Malta	-0.17 (0.939)	1.08	1.74*
Poland	-1.24 (0.656)	-0.55	1.11*
Romania	-1.00 (0.750)	-1.31	0.51*
Slovakia	-1.92 (0.322)	0.62	1.10*
Slovenia	-1.79 (0.384)	0.05	0.59*
Germany	-0.06 (0.950)	-0.37	1.40*

Notes: The null hypothesis for the ADF and DF-GLS tests is the unit root hypothesis, while the null hypothesis for the KPSS test states that a series is covariance stationary. Number of lags in the ADF and DF-GLS tests regression was selected using the AIC criterion. Numbers in parentheses are p-values. The 5% critical value for the KPSS test is 0.146 with intercept and trend as exogenous terms, and 0.463 with only intercept as exogenous term (Kwiatkowski *et al*., 1992). * denotes rejection of the covariance stationarity hypothesis at the 5% level of significance.

Table 3: Cointegration and common trends tests results

Germany with	$(p\text{-}r_0)$	$LR(r_0)$	p-values	$\hat{k}$	$CV=(1,\text{-}1)$	L-statistic
Bulgaria	2	9.30	0.652	3	NA	NA
	1	1.93	0.823			
Croatia	2	19.97**	0.031	5	2.04	2.64
	1	1.06	0.903		(0.153)	(0.104)
Cyprus	2	11.13	0.472	3	NA	NA
	1	2.21	0.768			
Czech Republic	2	17.22*	0.095	5	1.06	0.83
	1	4.57	0.365		(0.303)	(0.363)
Hungary	2	10.19	0.563	3	NA	NA
	1	3.51	0.528			
Latvia	2	8.48	0.733	12	NA	NA
	1	2.28	0.761			
Lithuania	2	18.73*	0.059	4	1.00	1.00
	1	1.76	0.852		(0.316)	(0.317)
Malta	2	10.29	0.554	3	NA	NA
	1	1.57	0.881			
Poland	2	9.91	0.591	1	NA	NA
	1	0.76	0.976			
Romania	2	21.74**	0.017	4	0.60	0.10
	1	2.21	0.717		(0.440)	(0.756)
Slovakia	2	23.31**	0.011	4	0.05	2.27
	1	0.60	0.986		(0.829)	(0.132)
Slovenia	2	8.62	0.719	3	NA	NA
	1	0.81	0.973			

Notes: The value reported at the top of the second column for each panel is for $r_0 = 0$, so that $p\text{-}r_0 = p$ is the dimension of the VECM. $\hat{k}$ is the estimated lag length in the VECM. Sixth column refers to the H_0 that the cointegrating vector is $(1,-1)$. Under the null hypothesis, this test is distributed as χ_1^2, asymptotically. The *L*-statistics are computed under the null hypothesis that the German 10-year bond yield is the common trend. Under the null hypothesis, the *L*-statistic is also distributed as χ_1^2. Numbers in parentheses are p-values. ** and * denote rejection of the null hypothesis at the 5% and 10% level of significance, respectively. NA stands for “Not Applicable”.

Table 4: Residuals-based bivariate tests for autocorrelation

VECM for	Ljung-Box Q-test		
Germany and	Q-statistic	Adj. Q-statistic	LM-test
Croatia	4.82 (0.306)	5.08 (0.279)	5.08 (0.279)
Czech Republic	8.58 (0.072)	8.87 (0.065)	6.70 (0.153)
Lithuania	6.21 (0.184)	6.40 (0.171)	4.24 (0.374)
Romania	3.90 (0.419)	4.08 (0.396)	2.03 (0.730)
Slovakia	2.95 (0.566)	3.02 (0.554)	0.73 (0.947)

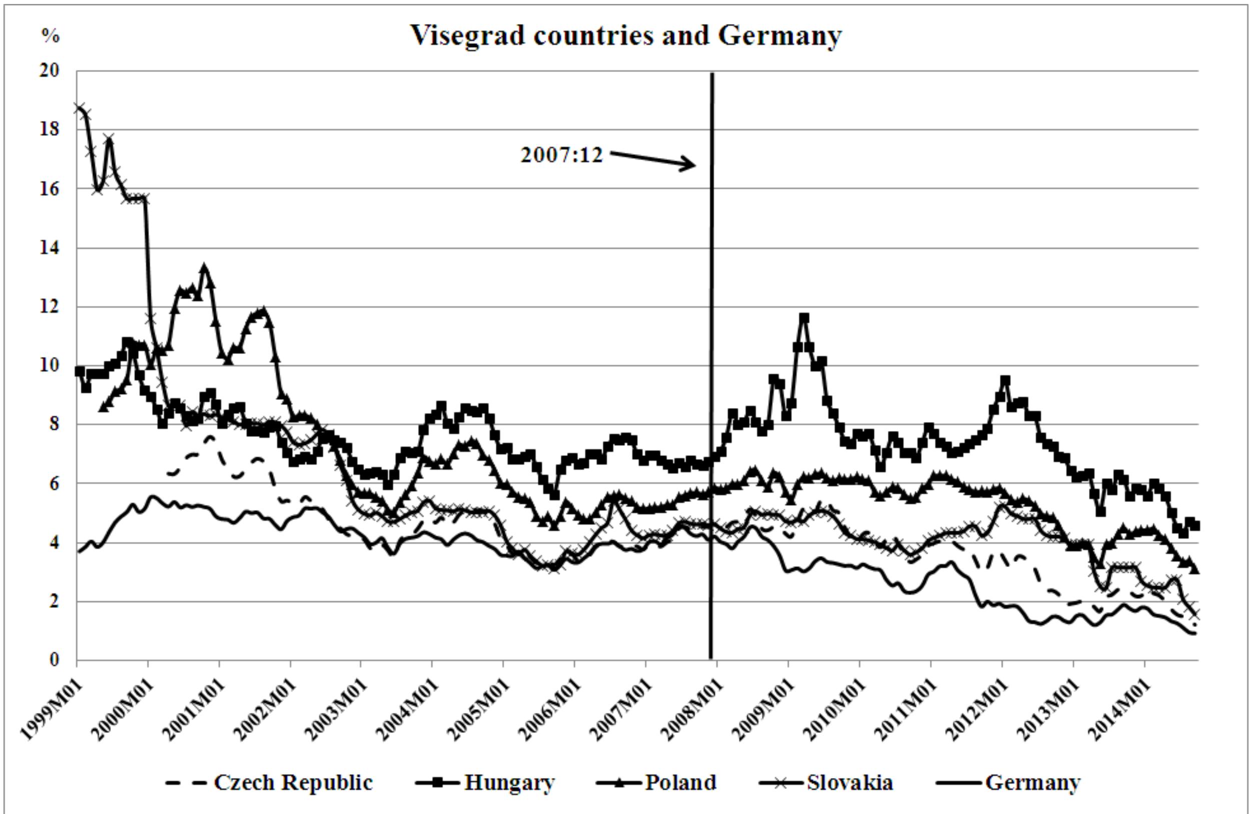
Notes: Under the null hypothesis of no serial correlation, both the Ljung-Box Q and the multivariate LM test statistics are distributed as χ^2 asymptotically, with degrees of freedom $p^2(s-k)$ and p^2, respectively, where $p=2$ is the dimension of the VECM, k is the lag length of the VECM determined by the AIC criterion, and $s=k+1$. The Adjusted Q-statistics correct the Q-statistics for sample size. For all tests df=4. Numbers in parentheses are p-values.

Table 5: Cotrending hypothesis

Country	Constant	Level shift	Linear trend	Trend shift	$\hat{k}$	*F*-test	*t*-test
Croatia	0.007	0.364*	0.006	-0.007	2	0.23	-9.08*
	[0.12]	[2.13]	[1.41]	[-1.51]		(0.635)	
Czech Republic	0.042	0.224	-0.000	-0.001	8	2.05	
	[0.75]	[1.71]	[-0.31]	[-0.95]		(0.155)	
Lithuania	0.157	0.888	-0.002	-0.004	3	1.78	
	[1.16]	[1.42]	[-1.01]	[-0.97]		(0.185)	
Romania	0.608*	0.342	-0.009	0.004	1	2.33	-2.84*
	[2.76]	[1.33]	[-1.66]	[0.84]		(0.130)	
Slovakia	0.310	-0.148	-0.003	0.003	6	0.00	
	[1.84]	[-0.56]	[-1.70]	[1.19]		(0.976)	

Notes: $\hat{k}$ is the lag length in each regression, based on the AIC criterion. Numbers in brackets are *t*-statistics, based on Newey-West standard errors. Null hypothesis for the *F*-test is the cotrending hypothesis (*i.e.* linear trend and trend shift are jointly zero). Numbers in parentheses are *F*-statistic p-values. Null hypothesis for the *t*-test is the weak monetary policy convergence criterion (*i.e.* the sum of intercept and level shift coefficients are greater than or equal to 2%). * denotes rejection of the null hypothesis at the 5% level of significance.

Figure 1. 10-year government bond yields



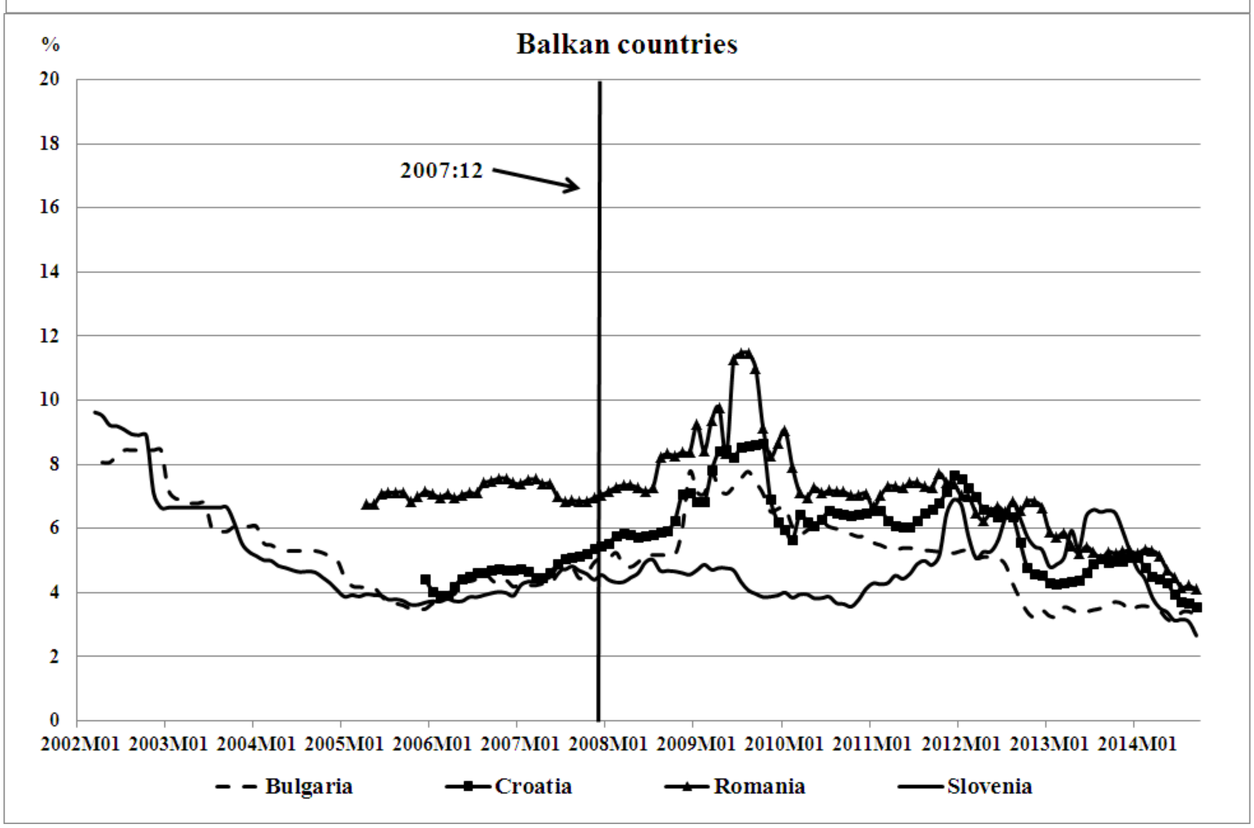


Figure 1. ***(continued)***

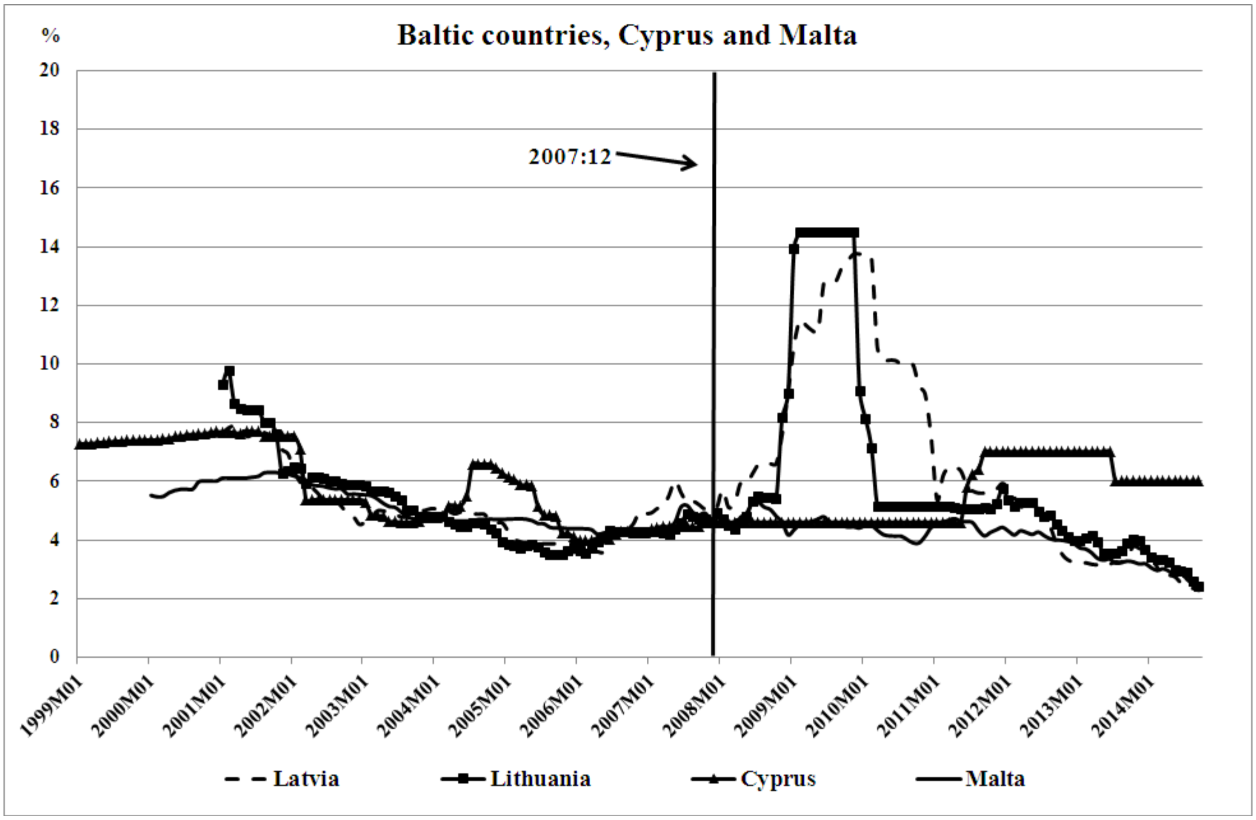


Figure 2: Spreads in relation to 10-year German government bond yield ($r_i - r_G$)