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Determinants of sovereign bond yield spreads in the EMU: An optimal currency area perspective

Mauro Costantini [a], Matteo Fragetta [b,c,1], Giovanni Melina [d,*]

[a] Department of Economics and Finance, Brunel University, Uxbridge, UB8 3PH, United Kingdom
[b] Department of Economics and Statistics, University of Salerno, Via Ponte Don Melillo, 84084 Fisciano, SA, Italy
[c] ISCTE – IUL, Business Research Unit (UNIDE-IUL), Instituto Universitario de Lisboa, Portugal
[d] Department of Economics, Social Sciences Building, City University London, Whiskin Street, London EC1R 0JD, United Kingdom

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ABSTRACT

In the light of the recent financial crisis, we take a panel cointegration approach that allows for structural breaks to the analysis of the determinants of sovereign bond yield spreads in nine economies of the European Monetary Union. We find evidence for a level break in the cointegrating relationship. Moreover, results show that (i) fiscal imbalances – namely expected government debt-to-GDP differentials – are the main long-run drivers of sovereign spreads; (ii) liquidity risks and cumulated inflation differentials have non-negligible weights; but (iii) all conclusions are ultimately connected to whether or not the sample of countries is composed of members of an Optimal Currency Area (OCA). In particular, we establish (i) that results are overall driven by those countries not passing the OCA test; and (ii) that investors closely monitor and severely punish the deterioration of expected debt positions of those economies exhibiting significant gaps in competitiveness.

1. Introduction

The sovereign debt crisis, which escalated in the European Monetary Union (EMU) in 2010, has sparked big debates about its causes and possible solutions, both in academia and in policy institutions. Since the start of the EMU and before the financial crisis, spreads on 10-year sovereign bond yields relative to the German benchmark were small.[2] With the financial crisis the picture completely changed. By the spring of 2009 the Greek sovereign bond spread had reached almost 300 basis points and by 2010 it had skyrocketed to over 1000 basis points (see Fig. 1). Investors started to question the ability of certain EMU governments of meeting their debt obligations and began requiring higher and higher risk premia.

What are the determinants of sovereign bond yield spreads in the EMU? The empirical literature has identified both a common international time-varying factor – commonly dubbed as international risk aversion – and country specific factors – in particular default and liquidity risk – as potential determinants of sovereign bond yield spreads in the EMU.

* Corresponding author. Tel.: +44 20 7040 4522; fax: +44 20 7040 8580.
E-mail addresses: mauro.costantini@brunel.ac.uk (M. Costantini), mfragetta@unisa.it (M. Fragetta), giovanni.melina.1@city.ac.uk (G. Melina).

[1] The author completed part of this work while he was visiting Cass Business School, City University London.

[2] Pagano and von Thadden (2004) argue that the exchange rate risk elimination and institutional factors contributed to the small sovereign spreads observed prior to the financial crisis. Barrios et al. (2009) look at EMU sovereign spread determinants during the financial crisis and argue that in the pre-crisis period there was also an underestimation of risk.

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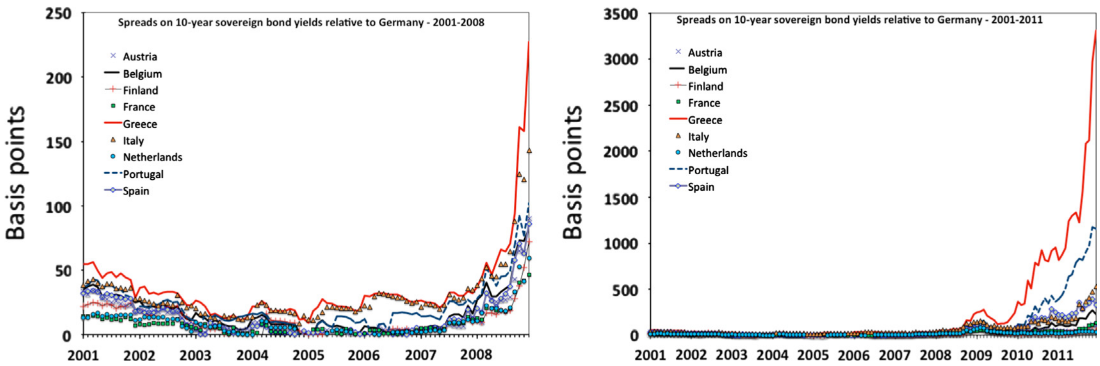

Fig. 1. Spreads on 10-year sovereign bond yields vis-à-vis Germany in the EMU.
Source: Authors' computations on Bloomberg data.

However, both in academic debates and in the context of policy-making, no clear consensus has arisen. As far as the default risk is concerned, Faini (2006), Hallerberg and Wolff (2008) and Bernoth et al. (2012) find that the budget balance and the stock of government debt have, on average, a significant impact on sovereign bond spreads, whereas Codogno et al. (2003) find that public debt plays a role only for Italy and Spain. As regards the liquidity risk component, Codogno et al. (2003) and Sgherri and Zoli (2009) find that liquidity explains only a small fraction of sovereign spreads, while Gomez-Puig (2006) and Barrios et al. (2009) show that liquidity is more important to explain euro area sovereign spreads. With respect to international risk aversion, Attinasi et al. (2010) show that this factor has substantially contributed to the change in sovereign bond spreads during the financial crisis.

This paper adds to this debate by taking a long-run approach to the analysis of the determinants of sovereign bond yield spreads in nine EMU economies (Austria, Belgium, Finland, France, Greece, Italy, the Netherlands, Portugal and Spain) relative to Germany, by looking at the issue from the viewpoint of the theory of optimal currency areas (OCA).[3] In particular, we argue (i) that long-run determinants of sovereign bond spreads are more relevant for policy-makers when they have to decide whether, and to what extent, structural policy interventions are needed to reduce sovereign bond yield differentials and (ii) that investors take OCA issues, and in particular diverging competitiveness among EMU members, seriously into account when they have to assign and price sovereign default risk.

In order to take the first point into consideration – and this is also the first innovation of the paper relative to the existing literature – we base our investigation on recently developed panel cointegration techniques that treat cross-sectional dependence via factor models, allow for potential breaks, and are robust to endogeneity. In fact, as regards cross-sectional dependence, we conjecture – and empirically test – that aspects of country interdependence, such as the economic and financial integration processes, the Maastricht convergence criteria, and the common monetary policy framework, cannot be neglected. In addition, given the evident shift in the level of sovereign bond yield spreads experienced during the financial crisis and the subsequent sovereign debt crisis (reported in Fig. 1), we believe that any analysis dealing with the determinants of sovereign spreads should take potential breaks into account. In this paper, we tackle these issues by testing for panel cointegration with break using the approach of Westerlund and Edgerton (2008).

As far as the second point is concerned – and this is also the second novel feature of the paper – in addition to the standard measures of default and liquidity risk, we include cumulated inflation differentials among our explanatory variables, to capture asymmetric shocks leading to a divergence in competitiveness. In fact, even small differences in inflation rates, if persistent, can lead to sizable changes in relative price levels. As shown in Fig. 2, since the start of the monetary union, cumulated inflation differentials among EMU countries have persistently diverged. As noted by Estrada et al. (2012), in principle, persistent inflation differentials may both be a benign phenomenon explained by a structural convergence process according to a Balassa–Samuelson type of argument, and the source of long-lasting and damaging losses of competitiveness.[4] In order the former type of argument to hold, however, inflation rates should be positively correlated with the difference between labor productivity growth in the traded versus non-tradable sectors. While there is some evidence that this effect can justify some inflation differentials in the euro area, a consensus seems to have emerged around the claim that the Balassa–Samuelson hypothesis cannot be the general explanation of the persistent inflation differentials across EMU members (ECB, 2003; Estrada et al., 2012). In particular, Estrada et al. (2012) argue that the heterogeneous inertial components of price and wage-setting rules across the EMU, such as those caused by wage indexation clauses, play a predominant role. Given the EMU fixed exchange rate regime, countries that have experienced persistent positive inflation differentials have been subject to an appreciation of the real exchange rate. As noted by De Grauwe and Ji (2012), a country experiencing a real appreciation is likely to bump into problems of competitiveness which in turn may lead to current

[3] In order to have a sufficiently large sample we consider founding members of the euro in addition to Greece, which joined the eurozone in 2001. Ireland and Luxembourg are left out as data are not available for all the variables of the empirical model.

[4] The Balassa–Samuelson hypothesis states that in response to a rise in productivity in the tradable goods sector – if labor is sufficiently mobile across sectors – wages will grow both in this sector and in the non-tradable sector. As the wage increase in the non-tradable sector is not matched by an increase in productivity, this will raise costs and prices in the non-tradable goods sector and will lead to a rise in inflation.

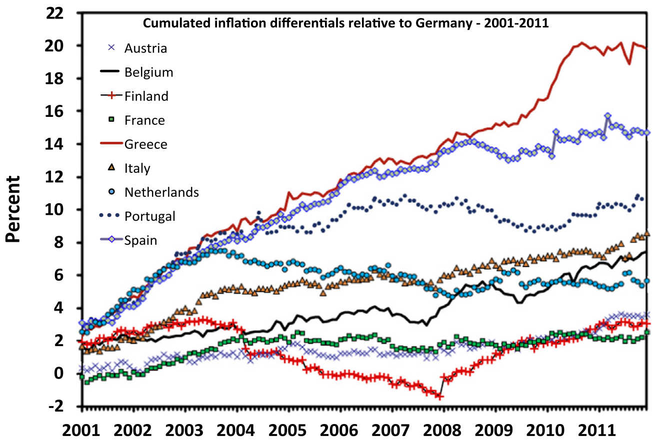

Fig. 2. Cumulated inflation differentials vis-à-vis Germany in the EMU.
Source: Authors' computations on Eurostat data.

account deficits and debt problems.[5] Regardless of the source of the imbalances, the appreciation of the real exchange rate for some EMU members has represented a gradual large asymmetric shock.[6] As a result, one of the theoretical conditions for an OCA, which requires that a shock in one country should be sufficiently correlated with that in the rest of the union, or that the union has put in place measures to balance out asymmetric shocks, has clearly been violated (see Mundell, 1961).[7]

As far as our empirical results are concerned, we find evidence for a level break in the cointegrating relationship, which we ascribe to the EMU sovereign debt crisis. This indicates that, after the crisis, the expected higher risk awareness of investors keeps government bond yield spreads at a higher level than in the pre-crisis period.

Moreover, results point at fiscal imbalances – in particular expected government debt-to-GDP differentials – as the main drivers of sovereign spreads, although liquidity risks have a non-negligible weight. Cumulated inflation differentials turn out to be a significant variable, its importance being of the same order of magnitude as liquidity risk. But, perhaps most importantly, their inclusion among the regressors allows us to establish that the conclusions we draw on sovereign bond yield spreads determinants are closely interlinked to whether or not diverging competitiveness significantly affects sovereign bond yield spreads themselves.

In particular, we argue that a statistical significance attached to cumulated inflation differentials is an indication that the economies included in the sample of countries do not belong to an OCA. In fact, if shocks were sufficiently correlated or if the monetary union were able to absorb and balance out asymmetric shocks, then cumulated inflation differentials would be small and unimportant for sovereign bond yield spread determination.[8] We iteratively run this test by excluding one country at a time from the full sample of countries, starting from that with the highest cumulated inflation differential relative to Germany, and going forward until such a variable becomes statistically insignificant. This process leads (i) to a grouping of countries into two categories corresponding to EMU core (Austria, Finland, France, Germany and the Netherlands) and EMU periphery (Belgium, Greece, Italy, Portugal and Spain), and (ii) to the finding that cointegrated panel regression results are clearly driven by the inclusion of the observations belonging to the peripheral EMU economies considered. In fact, when such observations are excluded, debt-to-GDP differentials turn out to be the least important determinant of the sovereign bond yield spread, while expected budget balance differentials and the liquidity risk carry the highest weights.

It is noteworthy that while in the sample of peripheral countries a one-percent-point rise in the expected public-debt-to-GDP ratio differential leads, on average, to an 8.63 basis points increase in the sovereign bond yield spread; in the restricted sample pooling only core EMU economies, the same increase in the expected public-debt-to-GDP ratio differential leads, on average, to an increase in the sovereign spread of only 0.46 basis points. These results clearly unveil the fact that international investors heavily punish the deterioration of expected debt positions of those countries that face competitiveness gaps and hence are not being perceived as OCA members.

[5] Also other measures of real exchange rate variations are used in the literature (see Arghyrou and Kontonikas, 2012; De Grauwe and Ji, 2012; Gibson et al., 2012).

[6] According to the literature, imbalances within the eurozone may be due to increasing cross-border investment since the Euro (Krugman, 2012), lower short and long interest rates in troubled countries (Wren Lewis, 2012) and different unit labor costs (Levy, 2012).

[7] The other conditions are: (i) mobility in the labor market which can help adjust to asymmetric shocks (Mundell, 1961); (ii) trade integration among countries in a Monetary Union which generates benefits due to the use of the same currency (Mundell, 1961); (iii) a high degree of mutual openness (McKinnon, 1963); (iv) a high degree of diversity of production and fiscal integration (Kenen, 1969).

[8] For an analysis on the impact of economic and institutional asymmetries on the effectiveness of monetary policy in the euro area see Aksoy et al. (2002).

The remainder of the paper is structured as follows. Section 2 describes the data employed in the estimation. Section 3 outlines the econometric methodology. Section 4 reports and discusses the results. Finally, Section 5 concludes and highlights policy implications. Technical details are appended to the paper.

2. Data

Our panel dataset contains monthly data of nine euro-area countries over the period 2001:1–2011:12. The countries, selected on the basis of data availability, are Austria, Belgium, Finland, France, Greece, Italy, the Netherlands, Portugal and Spain. In the remainder of the paper subscript i refers to the cross-sectional dimension (country) and subscript t refers to the time dimension (month).

The variable to be explained is the ten-year government bond yield spread over Germany (r_{it}). Data are taken from Bloomberg. The remainder of this section provides a rationale for the choice of potential explanatory variables.

As a measure of a country's creditworthiness, we use the forecast of the ratios of government budget balance to GDP (GB_{it}) and debt to GDP (DB_{it}) as differences vis-à-vis Germany's counterparts. The rationale behind the use of these two expected fiscal variables is that they represent two of the main sources of information for investors to form expectations on a country's fiscal position and the associated default risk. Given its prominent role in the euro area, we use the European Commission Forecasts which were released on a bi-annual basis over our sample period.[9] In our database, the value of budget balance and debt ratios are updated every time new forecasts are published (also Attinasi et al., 2010; Favero and Missale, 2012 adopt a similar procedure).

To capture the liquidity risk component (also explored by Codogno et al., 2003; Gomez-Puig, 2006; Barrios et al., 2009; Sgherri and Zoli, 2009), we use the bid-ask spread (differences vis-à-vis Germany) on ten-year sovereign bonds available in Bloomberg (BAS_{it}). This component denotes the difference in price between the highest price that a buyer is willing to pay for an asset and the lowest price at which a seller is willing to sell it. A rise in the bid-ask spread represents a deterioration of liquidity conditions and this may affect the corresponding sovereign yield spread. Fleming (2003) analyzes this measure of liquidity risk with US data and shows that the bid-ask spread is a useful measure for assessing and tracking Treasury market liquidity. In fact it can be calculated quickly and easily with data that are widely available and it is highly correlated with episodes of reported poor liquidity in the expected manner.[10]

Given that the literature considers also the international risk aversion as one determinant of sovereign bond yield spreads (see Attinasi et al., 2010 among others), we also consider the inclusion of a variable capturing this phenomenon, e.g. measured by the US corporate Baa-Aaa spread, among our regressors. As this is common to all countries, we subject it to a DF-GLS unit root test. Results show that the null hypothesis of a unit root can be rejected at 5 percent significance level (test statistics $=-2.488$). The stationarity of the time-varying degree of international risk aversion indicates that it is a phenomenon influencing short-run variations in sovereign yield spreads – e.g. Attinasi et al. (2010) find it to be relevant during the financial crisis – but not long-run fluctuations. As a result, we exclude it from our empirical model specification.

Finally, as a measure of competitiveness gaps, we use cumulated inflation differentials with respect to Germany (CID_{it}), derived using the monthly harmonized indices of consumer prices available in the EUROSTAT database. In fact, while there is some evidence that the Balassa–Samuelson hypothesis can justify some inflation differentials in the euro area, a consensus seems to have emerged around the claim that this cannot be the general explanation of the persistent inflation differentials across EMU members (ECB, 2003; Estrada et al., 2012) and that persistent inflation differentials are a source of long-lasting and damaging losses of competitiveness (Estrada et al., 2012). Given the fixed exchange rate regime of the EMU, this variable captures the relative variation in the real exchange rate, reflecting different potential sources of imbalances. For instance, De Grauwe and Ji (2012) argue that a country experiencing a real appreciation is likely to bump into problems of competitiveness, which in turn may lead to current account deficits and debt problems. This is the reason why investors may require, *ceteris paribus*, an additional risk premium if they observe large and persistent inflation differentials.

3. Econometric methodology

The discussion on the variables of interest in Section 2 leads to the following specification for an empirical model on long-run determinants of sovereign bond yield spreads:

$$r_{it}=\alpha_{i1}+\alpha_2 DB_{it}+\alpha_3 CID_{it}+\alpha_4 BAS_{it}+\alpha_5 GB_{it}+\varepsilon_{it},\quad i=1,\ldots,N,\quad t=1,\ldots,T, \tag{1}$$

where $\alpha_j,\ j=1,\ldots,5$ are coefficients to be estimated, the notation on the regressand and the regressors is that described in Section 2 and ε_{it} is an error term.

To estimate Eq. (1) we have to take three important econometric issues into account. First, we cannot ignore the fact that the countries under investigation are closely interconnected, hence it is very likely that the variables feature cross-sectional dependence (CD). Likely sources of such a form of dependence are the economic and financial integration processes, the Maastricht convergence criteria and the common monetary policy framework, among others. Second, the variables in our specification are likely to exhibit a

[9] The forecasts are available at http://ec.europa.eu/economy_finance/publications/european_economy/forecasts_en.htm. In 2012 the European Commission started releasing these forecasts more frequently.

[10] Conversely, Fleming (2003) also finds that quote size, trade size, and on-the-run/off-the-run yield spread are found to be only modest proxies for market liquidity.

unit root. Hence, given that we are interested in establishing a long-run relationship between bond yield spreads and its determinants, if we verify that the variables in Eq. (1) are indeed non-stationary, we need to determine whether these are linked by a cointegrating relationship. Third, it is likely that structural breaks may have taken place around the EMU sovereign debt crisis. For instance, Fig. 1 shows that sovereign bond yields vis-à-vis Germany in the EMU experienced a visible level shift as a result of which, in the case of many countries, their post-2009 average is one or two orders of magnitude higher. This is an indication that breaks might have taken place both in the unit root processes of the variables in questions and within their cointegrating relationship.

The presence of CD – for which we also formally test using the procedures developed by Breusch and Pagan (1980), Pesaran (2004), and Ng (2006) – dictates the choice of appropriate unit root and cointegration tests, as well as of an appropriate estimator. We proceed in three steps.

1. We test for non-stationarity in the data using the testing procedure developed by Bai and Carrion-i-Silvestre (2009) on each variable of the empirical model. In particular, this procedure employs panel unit root statistics pooling the modified Sargan–Bhargava tests for individual series taking structural breaks and cross-dependence into account through the common factors model proposed by Bai and Ng (2004). Details on the test statistics are provided in Appendix A.
2. We investigate on the existence of a cointegrating relationship for our empirical model using the panel cointegration tests proposed by Westerlund and Edgerton (2008), the null hypothesis of which is that of no cointegration. These tests can be used under very general conditions (heteroskedastic and correlated errors, individual-specific intercepts and time trend, cross-section dependence and unknown breaks both in the intercept and slope of the cointegrated regression). Details on these test statistics are provided in Appendix B.
3. If we find cointegration in step 2, we estimate the long-run relationship (1) among the variables of interest using the continuously-updated fully-modified (CUP-FM) estimator developed by Bai and Kao (2006).[11] This estimator takes both cross-sectional dependence and endogeneity into account. While the former is tackled via a common factor structure, the latter is treated via an appropriate variable transformation exploiting long-run covariances, as reported in Appendix C. This is particularly relevant for our analysis. For instance, while it is plausible to think that expectations of worse (better) fiscal positions lead to higher (lower) sovereign spreads, it is also plausible to conjecture that higher (lower) spreads lead to worse (better) fiscal conditions through an increase (decrease) of debt servicing costs. Similar arguments apply to the relationship between sovereign spreads and the measure of liquidity risk. Failure to correct for endogeneity would lead to inconsistent estimates of the coefficients. In this paper, this issue is tackled by employing the CUP-FM estimator that, by construction, corrects for endogeneity.

4. Results

Three tests, the results of which are reported in Table 1, point at a clear-cut evidence of cross-sectional dependence for all variables. This reinforces our prior conjecture based on economic considerations. In particular we run the LM test by Breusch and Pagan (1980), the CD test by Pesaran (2004), and the standardized Spacing Variance Ratio (SPR) by Ng (2006).[12]

In order to check for the presence of unit roots in the data, we apply the panel unit root tests of Bai and Carrion-i-Silvestre (2009) allowing for the presence of a break both in the mean and in the trend. As shown in Table 2, we find that the unit root hypothesis cannot be rejected.

Then, given non-stationarity, we test for panel cointegration among the five variables described in Section 2. In particular, we run three versions of the test of Westerlund and Edgerton (2008): (i) in the absence of any break; (ii) in the presence of solely a level break, i.e. just in the intercept of the cointegrating relationship; and (iii) in the presence of a regime shift, i.e. a break also in the slope of the relationship. The results in Table 3 show that while the null hypothesis of no panel cointegration in the absence of any break and with regime shift cannot be rejected, the two tests Z_τ and Z_ϕ reject the null hypothesis of no panel cointegration with level break at a 5 percent and marginally at a 10 percent significance level, respectively.[13]

Both the Bai–Carrion-i-Silvestre and the Westerlund–Edgerton procedure are very convenient because not only do they allow conducting panel unit root and cointegration testing in the presence of breaks but they also allow estimating the dates in which the breaks have most likely occurred. In Fig. 3 we report the estimated break dates both for the unit root processes and the cointegrating relationship.[14] In the latter, all break dates are estimated to have occurred between 2008 and 2010.[15] Noticeably, for most countries, the estimated level break in the cointegrating relationship is sometime in 2010, the year in which the sovereign debt crisis escalated according to many commentators and institutions such as the European Central

[11] On the use of this estimator see also Costantini and Gutierrez (2013) and Costantini et al. (2013).

[12] For a recent contribution employing the CD test see e.g. Camarero et al. (2013). Details on the SPR are provided by Ng (2006) and Carrion-i-Silverstre and German-Soto (2009). For a recent survey on large panel data models with cross-sectional dependence see Chudik and Pesaran (2013).

[13] Given that the regime shift model nests the level break model, it is possible that the null hypothesis is not rejected in the former case due to lack of power. The residuals (reported in Appendix D) of our panel regression (which takes the level break into account) do not exhibit a worrying behaviour with respect to the presence of further breaks.

[14] Expected government-budget-to-GDP differentials do not display breaks.

[15] In order not to exclude potential breaks that may have occurred during the financial and sovereign debt crisis, in searching for breaks, we set a trimming of 0.10.

Table 1
Cross-sectional dependence test results.

Variable	Breusch–Pagan's LM[a]	Pesaran's CD[b]	Ng's test[c]			
			ϑ	S	L	W
r_{it}	1005.676	55.597	0.083	−1.732	1.793	3.448
	(0.000)	(0.000)		(0.958)	(0.036)	(0.000)
DB_{it}	1523.857	19.554	0.083	−1.732	2.153	5.033
	(0.000)	(0.000)		(0.958)	(0.016)	(0.000)
CID_{it}	518.149	34.748	0.083	−1.732	3.290	1.093
	(0.000)	(0.000)		(0.958)	(0.000)	(0.137)
BAS_{it}	401.074	41.755	0.083	−1.732	3.866	2.495
	(0.000)	(0.000)		(0.958)	(0.000)	(0.006)
GB_{it}	844.278	46.651	0.139	−1.553	1.883	4.180
	(0.000)	(0.000)		(0.939)	(0.030)	(0.000)

Notes: in the three tests the null hypothesis is that of cross-sectional independence. *p*-values are reported in parenthesis.
[a] The Breusch and Pagan (1980) LM test is distributed as a χ^2 under the null hypothesis.
[b] Pesaran (2004) shows that under the null hypothesis of no cross-sectional dependence $CD \xrightarrow{d} N(0,1)$.
[c] The test by Ng (2006) is carried out by splitting the whole (W) sample of (ordered) spacings at $\vartheta \in (0,1)$, so that the group of small (S) correlation coefficients and the group of large (L) correlation coefficients can be defined. The null hypothesis of independence is tested for the small, large and the whole sample using the standardized Spacing Variance Ratio (SVR) in Ng (2006), which under the null hypothesis of independence converges to the standard Normal distribution. For further details on spacing see Ng (2006).

Table 2
Panel unit root test results.

Variables	Break in the mean			Break in the trend					
	Z	$\mathcal{P}$	$\mathcal{P}_m$	Z	$\mathcal{P}$	$\mathcal{P}_m$	Z^*	$\mathcal{P}^*$	$\mathcal{P}^*_m$
r_{it}	−2.136	4.093	42.556	−1.225	4.415	44.493	6.344	−2.075	5.552
DB_{it}	0.571	−1.528	8.830	−1.589	0.776	22.655	−0.736	0.303	19.819
CID_{it}	3.979	−2.056	5.664	−0.674	0.996	23.974	−0.710	1.318	25.908
BAS_{it}	−2.493	26.165	174.992	−1.281	5.420	50.523	14.697	0.227	19.362
GB_{it}	−1.043	−0.406	15.562	−0.418	0.097	18.581	−0.418	0.097	18.581

Notes: (a) Z, $\mathcal{P}$ and $\mathcal{P}_m$ denote the test statistics developed by Bai and Carrion-i-Silvestre (2009), the 5 percent critical values of which are 1.645, −1.645 and 50.998, respectively. (b) Z^*, $\mathcal{P}^*$, and $\mathcal{P}^*_m$ refer to the corresponding statistics obtained using the *p*-values of the simplified MSB statistics. (c) The number of common factors is 1. (d) The maximum number of breaks allowed is 3. To determine the breaks, the procedure of Bai and Perron (1998) is used (for details see Bai and Carrion-i-Silvestre, 2009). Details on the computation of Z, $\mathcal{P}$, $\mathcal{P}_m$, Z^*, $\mathcal{P}^*$, and $\mathcal{P}^*_m$ are provided in Appendix A. (e) No simplified test for model 1 (see Eq. (A.4) in Appendix) is provided by Bai and Carrion-i-Silvestre (2009).

Table 3
Panel cointegration test results.

Model	$Z_\tau(N)$	$Z_\phi(N)$
No break	−0.384	−0.024
	(0.350)	(0.491)
Level break	−1.848	−1.232
	(0.032)	(0.109)
Regime shift	−1.186	−0.705
	(0.118)	(0.240)

Notes: (a) The number of lags in the test regressions for both LM tests is selected using the procedure of Campbell and Perron (1991). (b) The number of common factors is set equal to 1. (c) The breaks are determined by grid-search at the minimum of the sum of squared residuals (see details in Westerlund and Edgerton, 2008). (d) *p*-values in parentheses are for a one-sided test based on the normal distribution.

Bank (ECB, 2010). In particular in Italy and Spain the break is estimated to have occurred in the same month of October 2010, while in Greece and Portugal the estimated break occurs slightly earlier on May 2010. The break in the intercept indicates that, after the crisis, the expected higher risk awareness of investors keeps government bond yield spreads at a higher level than in the pre-crisis period.

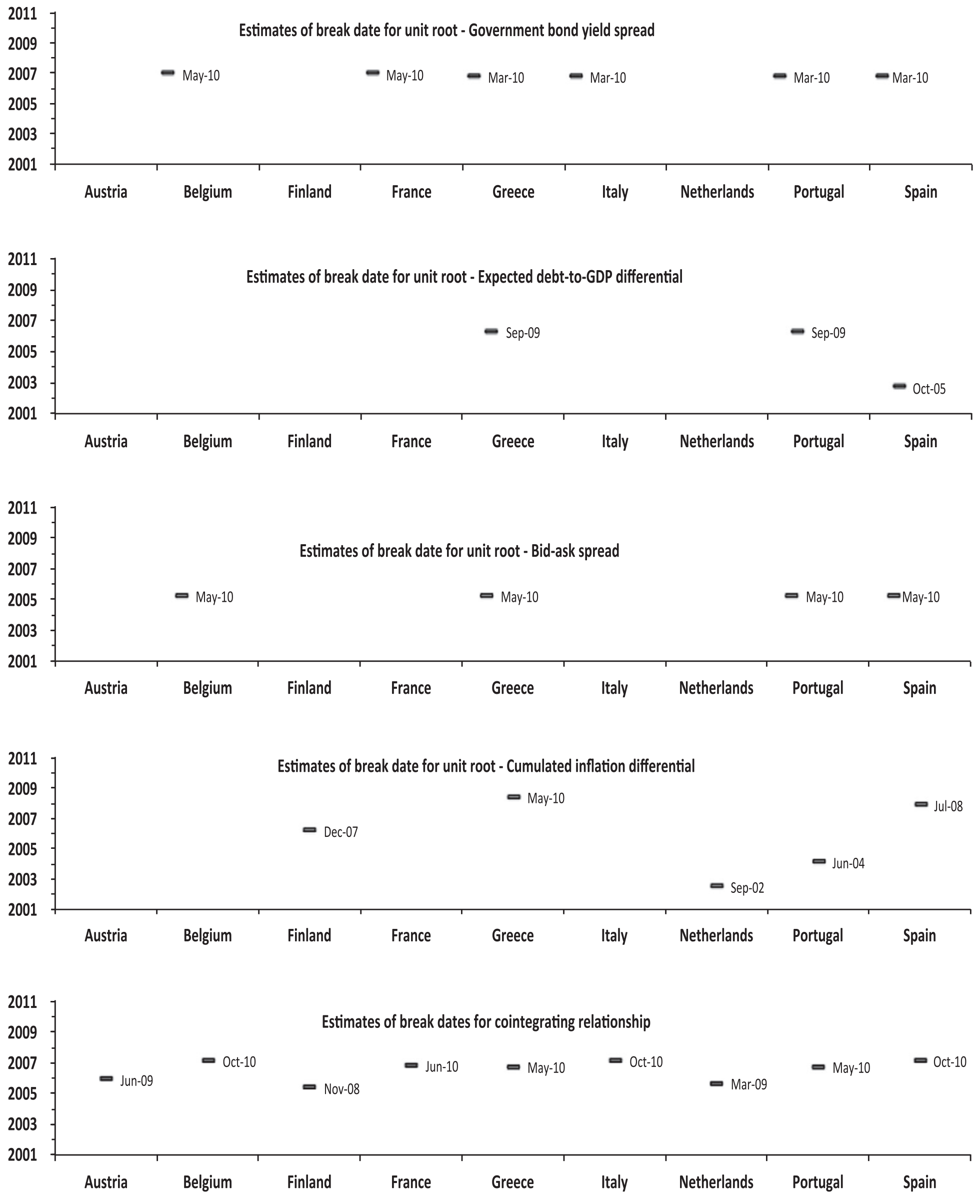

Fig. 3. Estimates of break dates for unit roots and the cointegrating relationship.

Having found evidence of cross-sectional dependence, non-stationarity, and cointegration makes the CUP-FM an appropriate estimator for Eq. (1), which we estimate taking also the level break into account with dummies.

The results for the full sample of countries are reported in Table 4(a). We report both the results obtained using the levels of the variables constructed in line with Section 2 and using standardized variables. We conduct the latter exercise in order to make the variables adimensional and assess the relative importance of the various determinants of sovereign bond yield spreads. The estimated coefficients are statistically significant at any conventional level and carry the expected sign.[16]

In our specification, public debt and the government budget balance represent those determinants most directly linked with sovereign default risk as they capture deterioration in fiscal conditions. Both theory and common sense suggest that spreads should widen when fiscal conditions deteriorate, e.g. when public debt rise and/or the government budget balance falls, and viceversa. The estimated coefficients indeed confirm this hypothesis and highlight a key role for the default risk in the relationship. In fact, on one hand, a one-percent-point rise in the expected public-debt-to-GDP ratio differential yields

[16] In Appendix D we report the standardized residuals from the panel regression. These are typical residuals obtained from a cointegrating relationship, in that they exhibit some persistence but they are stationary (mean-reverting) and do not exhibit any abrupt shifts in the level. Having controlled for the level break, residuals display a comparable behaviour in relative terms.

Table 4
Panel estimation results.

Regressor	Coefficient	(a) Full sample		(b) Greece excluded		(c) Greece and Spain excluded		(d) Greece, Spain and Portugal excluded		(e) Greece, Spain, Portugal and Italy excluded		(f) Greece, Spain, Portugal, Italy and Belgium excluded	
		Level	Standardized	Level	Standardized	Level	Standardized	Level	Standardized	Level	Standardized	Level	Standardized
Expected gov. debt/GDP diff.	α_2	0.0756***	0.8936***	0.0229***	0.2414***	0.0171***	0.1853***	0.0050***	0.2616***	0.0032***	0.2122***	0.0046***	0.0337***
		(0.0034)	(0.0397)	(0.0026)	(0.0262)	(0.0027)	(0.0280)	(0.0017)	(0.0825)	(0.001)	(0.0898)	(0.0013)	(0.0067)
Cumulated inflation diff.	α_3	0.2004***	0.4108***	0.0803***	0.1314***	0.0827***	0.1030***	0.0943***	0.4192***	0.0363***	0.2324***	−0.0007	−0.0102
		(0.0226)	(0.0397)	(0.0145)	(0.0250)	(0.0151)	(0.0204)	(0.0102)	(0.0466)	(0.0092)	(0.0623)	(0.0076)	(0.0075)
Bid-ask spread	α_4	0.0441***	0.5031***	0.0921***	0.2945***	0.0952***	0.3246***	0.0865***	0.1966***	0.1503***	0.4991***	0.1160***	0.0554***
		(0.0004)	(0.0050)	(0.0015)	(0.0048)	(0.0014)	(0.0048)	(0.0035)	(0.0077)	(0.0045)	(0.0150)	(0.0036)	(0.0017)
Exp. gov. balance/GDP diff.	α_5	−0.0933***	−0.1151***	−0.1037***	−0.1207***	−0.0885***	−0.0989***	−0.0583***	−0.2770***	−0.0543***	−0.4236***	−0.0475***	−0.0544***
		(0.0121)	(0.0149)	(0.0077)	(0.0089)	(0.0075)	(0.0083)	(0.0045)	(0.0216)	(0.0039)	(0.0306)	(0.0030)	(0.0035)

Notes: The dependent variable is the ten-year government bond yield spread over Germany (r_{it}). *, ** and *** denote significance at 10, 5 and 1 percent level, respectively. Standard errors are in parenthesis.

an average increase in the sovereign bond yield spread of 7.56 basis points. On the other hand, a surge in the expected budget-balance-to-GDP-ratio differential of one percent point generates a reduction in the spread of 9.33 basis points. Government debt turns out to be the most important determinant of the spread amongst all the determinants encompassed by our specification, while the budget balance carries the smallest weight. This finding seems reasonable if it is looked at through the lens of investors who need to price risk. In fact, it implies that these assign a greater importance to a proxy of a long history of fiscal conditions – the accumulation of the stock of debt over time – rather than to possibly one-off fiscal conditions exemplified by the government budget balance.

Liquidity risk, proxied by the bid-ask spread, is the second most important determinant of the sovereign bond yield spread. In particular, a one-basis-point increase in this measure of liquidity risk leads to a *ceteris paribus* higher liquidity premium of 4.41 basis points. In other words, the higher the level of liquidity in the government bond market, the lower the sovereign spread, and the extent to which this occurs is not only statistically but also economically important.

Cumulated inflation differentials come third in the ranking of determinants, although their weight is of comparable magnitude as liquidity risk. In particular, a one-percent increase in cumulated inflation differentials leads to a rise of 20 basis points in the spread. This result agrees with De Grauwe and Ji (2012).

In a monetary union, significant cumulated inflation differentials unveil competitiveness gaps as they imply an appreciation of the real exchange rate. This in turn represents a failure of the economies belonging to the monetary union in what should be their aim of moving towards higher and higher degrees of economic integration, in order to ultimately tick the boxes necessary for OCA membership. Our estimates show that investors do take these considerations into account when pricing sovereign risk. Therefore, we argue that a statistical significance attached to cumulated inflation differentials is an indication that the economies included in the sample of countries do not belong to an OCA. We run this OCA test iteratively by excluding the economies with the highest cumulated inflation differential from the full sample of countries one at a time, until the estimated coefficient attached to the variable becomes statistically insignificant. The residual countries will then represent economies that may appertain to an OCA in the eyes of investors who have to price sovereign risk. Among the economies included in the full sample of countries, the procedure (see Table 4(b) to (f)) excludes, in the order, Greece, Spain, Portugal, Italy and Belgium (arguably peripheral EMU economies) from the OCA. In the restricted sample – including Austria, Finland, France, the Netherlands along with Germany (arguably core EMU economies) – inflation differentials, although they display considerable variation (see Fig. 2), do not play a significant role in sovereign bond yield spreads determination. We argue that this is due to the fact that these countries face much smaller competitiveness gaps and investors perceive them as belonging to an OCA.

A comparison between Table 4(a) and (f) also highlights that the results obtained with the full sample of countries are mainly driven by those economies that are not perceived as belonging to an OCA. In particular, in the restricted sample of countries, expected government debt-to-GDP differentials turn out to be the least important determinant of the sovereign bond yield spread, while expected budget balance differentials and the liquidity risk carry the highest weights. Table 5 contrasts the results obtained using observations of the restricted sample of peripheral countries (Greece, Spain, Portugal, Italy and Belgium) with those obtained using only observations of core EMU economies (Austria, Finland, France, the Netherlands along with Germany). The same one-percent-point rise in the expected public-debt-to-GDP ratio differential yields an average increase in the sovereign bond yield spread of 8.63 basis points in peripheral EMU countries and only 0.46 basis points in the restricted OCA (core) sample. The results obtained using standardized variables confirm the difference in the ranking of determinants identifiable in Table 4. Such findings are informative on the fact that expected debt positions of those countries facing problems of competitiveness due to a sustained appreciation of the real exchange rate, and not being perceived as OCA members, are closely monitored by investors and their deterioration is much more heavily punished.

5. Conclusions and policy implications

This paper provides useful information for policy makers facing the difficult task of tackling high sovereign bond spreads with the aim of fostering greater public finance stability and ultimately guaranteeing EMU survival. Results primarily point at expected fiscal imbalances (namely expected government debt-to-GDP differentials) and liquidity risks as the main determinants of sovereign bond yield spreads in the long run. We find evidence for a level break in the relationship, occurring during the sovereign debt crisis.

These results suggest that some EMU countries do need fiscal consolidation in order to remove imbalances and bring sovereign spreads to acceptable levels. Across the EMU, however, this is still a time of weak private demand and fiscal tightening may worsen economic conditions even further with a perverse effect on government debt-to-GDP ratios themselves. In the literature, there are ongoing debates on the appropriate timing (slow versus fast) and composition (expenditure versus tax-based) of fiscal consolidations and on whether high levels of public debt harm economic growth (see e.g. Batini et al., 2012; Cantore et al., 2013; Panizza and Presbitero, 2013 among others). Our paper does not take a stance in these particular debates. Nevertheless, by looking at the issue through the lens of the OCA theory, it is able to establish that this is only one important side of the coin.

The other side, which we deem as equally important, is the extent to which EMU countries do form an OCA and, above all, whether investors take this information into account when they have to assess and price sovereign default risk. Our empirical analysis finds that cumulated inflation differentials have non-negligible weights in sovereign bond yield spread determination. If our full sample of countries comprised only OCA members, then cumulated inflation differentials would be

Table 5
Panel estimation results: peripheral vs. core countries.

Regressor	Coefficient	(a) Peripheral		(b) Core	
		Level	Standardized	Level	Standardized
Expected gov. debt/GDP diff.	α_2	0.0863***	1.0143***	0.0046***	0.0337***
		(0.0046)	(0.0536)	(0.0013)	(0.0067)
Cumulated inflation diff.	α_3	0.2253***	0.4609***	−0.0007	−0.0102
		(0.0305)	(0.0639)	(0.0076)	(0.0075)
Bid-ask spread	α_4	0.0444***	0.5454***	0.1160***	0.0554***
		(0.0005)	(0.0066)	(0.0036)	(0.0017)
Exp. gov. balance/GDP diff.	α_5	−0.0796***	−0.1014***	−0.0475***	−0.0544***
		(0.0205)	(0.0251)	(0.0030)	(0.0035)

Notes: The dependent variable is the ten-year government bond yield spread over Germany (r_{it}). *, **, and *** denote significance at 10, 5 and 1 percent level, respectively. Standard errors are in parenthesis.

negligible and, most importantly, they would have an immaterial effect on sovereign bond yield spreads. In fact, substantial and protracted cumulated inflation differentials (i) derive from a failure of the EMU to work as an OCA and hence to absorb and balance out asymmetric shocks and (ii) lead to a divergence of real exchange rates and competitiveness. Therefore, a statistical significance associated to cumulated inflation differentials can be interpreted as an indication that the economies in the sample do not constitute an OCA. Using this criterion, we are able to group the countries in our sample into peripheral EMU economies (countries that do not pass the OCA test) and core EMU economies (countries that do pass the test) and this allows us to establish that the above results are driven by peripheral EMU countries. In particular, within core EMU, (i) debt-to-GDP differentials cease to be the main long-run drivers of sovereign bond yield spreads; and (ii) the same increase in the debt-to-GDP differential leads to a dramatically smaller increase in the sovereign spread.

Such findings are noteworthy because they highlight that investors closely monitor and severely punish the deterioration of debt positions of those economies exhibiting significant competitiveness gaps. This suggests that policy-makers willing to reduce the burden of high sovereign spreads in the EMU should embrace structural policies aiming at a higher level of coordination of prices and wages across the union, besides well-designed consolidations programs.

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Appendix A. Panel unit root test

Bai and Carrion-i-Silvestre (2009) propose panel unit root statistics that pool the modified Sargan–Bhargava (hereafter MSB) tests for individual series taking into account structural breaks and cross-dependence through a common factors model proposed by Bai and Ng (2004). The common factors may be non-stationary processes, stationary processes or a combination of both. Bai and Carrion-i-Silvestre (2009) consider the following panel data model:

$$X_{it} = D_{it} + \mathbf{F}_t' \boldsymbol{\pi}_i + e_{it.} \tag{A.1}$$

$$(I-L)\mathbf{F}_t = C(L)u_t \tag{A.2}$$

$$(1-\rho_i L)e_{it} = H_i(L)\varepsilon_{it}, \tag{A.3}$$

$t = 1, \ldots, T$ and $i = 1, \ldots, N$, where $C(L) = \sum_{j=0}^{\infty} C_j L^j$, $H_i(L) = \sum_{j=0}^{\infty} C_{ij} L^j$, L is lag operator, and ρ_i is the autoregressive parameter in the univariate model. The component D_{it} indicates the deterministic part of the model, $\mathbf{F}_t$ is an $(r \times 1)$ vector that accounts for the common factors of the panel, e_{it} is the idiosyncratic error term, $\mu_t \sim i.i.d.(0, \Sigma_u)$ and $\varepsilon_{it} \sim i.i.d.(0, \Sigma_{\varepsilon_i})$. Despite the operator $(1-L)$ in Eq. (A.2), $\mathbf{F}_t$ does not have to be $I(1)$. In this regard, $\mathbf{F}_t$ can be $I(0)$, $I(1)$, or a combination of both, depending on the rank of $C(1)$. If $C(1) = 0$, then $\mathbf{F}_t$ is $I(0)$. If $C(1)$ is of full rank, then each component of $\mathbf{F}_t$ is $I(1)$. If $C(1) \neq 0$ but not full rank, then some components of $\mathbf{F}_t$ can be $I(1)$ and others $I(0)$.[17] As regards the deterministic component, D_{it}, Bai and Carrion-i-Silvestre (2009) propose two specifications:

$$\text{Model 1:} \; D_{it} = \mu_i + \sum_{j=1}^{l_i} \theta_{ij} DU_{ijt} \tag{A.4}$$

[17] For further details on assumptions regarding the panel data model see Bai and Carrion-i-Silvestre (2009).

$$\text{Model 2: } D_{it} = \mu_i + B_i t + \sum_{j=1}^{l_i} \theta_{ij} DU_{ijt} + \sum_{k=1}^{m_i} \gamma_{ik} DT_{ikt} \tag{A.5}$$

where l_i and m_i denote the structural breaks affecting the mean and the trend of a time series, respectively, and l_i is not necessarily equal to m_i. The dummy variables are defined as follows: $DU_{ijt} = 1$ for $t > T^i_{aj}$ and 0 elsewhere, and $DT_{ikt} = (t - T^i_{bk})$ for $t > T^i_{bk}$ and 0 elsewhere. T^i_{aj} and T^i_{bk} indicate the j-th and k-th dates of the break in the level and in the trend, respectively, for the i-th individual, with $j = 1, \dots, l_i$ and $k = 1, \dots, m_i$. Bai and Carrion-i-Silvestre (2009) propose to combine individual MSB test statistics to test the null hypothesis of $\rho_i = 1$ for all $i = 1, \dots, N$ against the alternative $|\rho_i| < 1$ for some i. This approach is suitable since e_{it} are cross-sectionally independent (the individual statistics are free from the common factors). Bai and Carrion-i-Silvestre (2009) provide two approaches for pooling individual statistics. The first approach is based on the use of the average of the individual statistics:

$$Z = \sqrt{N}\frac{\overline{MSB(\lambda)} - \overline{\xi}}{\overline{\varsigma}} \xrightarrow{d} N(0,1) \tag{A.6}$$

with $\overline{MSB(\lambda)} = N^{-1}\sum_{i=1}^{N} MSB_i(\lambda_i)$, $\overline{\xi} = N^{-1}\sum_{i=1}^{N}\xi_i$ and $\overline{\varsigma}^2 = N^{-1}\sum_{i=1}^{N}\varsigma_i^2$, where ξ_i and ς_i^2 denote the mean and the variance of the individual modified Sargan–Bhargava ($\text{MSB}_i(\lambda_i)$) statistics respectively and $\lambda_i = T^i_b/T$ the break fraction parameter. In order to purge the break fraction parameter in the limiting distributions in the case of Model 2, Bai and Carrion-i-Silvestre (2009) propose another test based on the simplified MSB statistics:

$$Z^* = \sqrt{N}\frac{\overline{MSB^*(\lambda)} - \overline{\xi}^*}{\overline{\varsigma}^*} \xrightarrow{d} N(0,1) \tag{A.7}$$

with $\overline{MSB^*(\lambda)} = N^{-1}\sum_{i=1}^{N} MSB^*_i(\lambda_i)$, $\overline{\xi}^* = N^{-1}\sum_{i=1}^{N}\xi^*_i$ and $\overline{\varsigma}^{*2} = N^{-1}\sum_{i=1}^{N}\varsigma_i^{*2}$, where ξ^*_i and ς_i^{*2} denote the mean and the variance of the individual $MSB^*_i(\lambda_i)$ statistics, respectively.[18]

The second approach is based on the method developed by Maddala and Wu (1999) and Choi (2001) that pools the p-values associated with the individual tests:

$$\mathcal{P} = -2\sum_{i=1}^{N} \ln p_i \xrightarrow{d} \chi^2_N \tag{A.8}$$

$$\mathcal{P}_m = \frac{-2\sum_{i=1}^{N} \ln p_i - 2N}{\sqrt{4N}} \xrightarrow{d} N(0,1) \tag{A.9}$$

where p_i denotes the individual p-value. Bai and Carrion-i-Silvestre (2009) also propose a version of $\mathcal{P}$ and $\mathcal{P}_m$ tests based on the p-values of the individual simplified MSB. They are denoted as $\mathcal{P}^*$ and $\mathcal{P}^*_m$, respectively.[19]

Appendix B. Panel cointegration tests

Westerlund and Edgerton (2008) propose two versions of a simple test for the null hypothesis of no cointegration that can be used under very general condition (heteroskedastic and correlated errors, individual-specific intercepts and time trend, cross-section dependence and unknown breaks both in the intercept and slope of the cointegrated regression). The test is derived from the Lagrange multiplier (LM)-based unit root tests (see Schmidt and Phillips, 1992; Ahn, 1993; Amsler and Lee, 1995). In our empirical analysis, we consider the following model:[20]

$$y_{it} = \alpha_i + \delta_i D_{it} + x'_{it}\beta_i + (D_{it}x_{it})'\gamma_i + z_{it}, \tag{B.1}$$

$$x_{it} = x_{it-1} + \omega_{it}, \quad t = 1, \dots T, \quad i = 1, \dots, N. \tag{B.2}$$

The k-dimensional vector x_{it} contains the regressors and is modeled as a pure random walk. The variable D_{it} is a scalar break dummy such that D_{it}=1 if $t > T^b_i$ and zero otherwise. α_i and β_i represent the intercept and slope before the break, while δ_i and γ_i represent the change in these parameters at the time of the shift. ω_{it} is an error process with mean zero and independent across i. In Eq. (B.1), the error term z_{it} is generated by the following model:

$$z_{it} = \lambda'_i F_t + \nu_{it}, \tag{B.3}$$

$$F_{jt} = \rho_j F_{jt-1} + u_{jt}, \quad j = 1, \dots, k, \tag{B.4}$$

$$\phi_i(L)\Delta\nu_{it} = \phi_i \nu_{it-1} + e_{it}, \tag{B.5}$$

where $\phi_i(L) = 1 - \sum_{j=1}^{p_i}\phi_{ij}L^j$ is a scalar polynomial in the lag operator L, F_t is an r-dimensional vector of unobservable common factors F_{jt} with $j = 1, \dots, r$, λ_i is a conformable vector of loading parameters, u_t is independent of e_{it} and ω_{it} for all i

[18] As regards the individual simplified MSB statistics, see Bai and Carrion-i-Silvestre (2009).

[19] Bai and Carrion-i-Silvestre (2009) point out that there is no need to construct a simplified test for Model 1 since this test does not depend on break fractions in limits.

[20] Westerlund and Edgerton (2008) also consider the deterministic trend in the model. We consider the constant only.

and t and e_{it} is an error term with mean zero and independent across both i and t. The relationship in (B.1) is cointegrated if $\phi_i < 0$ and it is spurious if $\phi_i = 0$. The tests proposed Westerlund and Edgerton (2008) are

$$Z_{\phi}(N) = \sqrt{N}(\overline{LM}_{\phi}(N) - E(B_{\phi})) \tag{B.6}$$

$$Z_{\tau}(N) = \sqrt{N}(\overline{LM}_{\tau}(N) - E(B_{\tau})), \tag{B.7}$$

where $\overline{LM}_{\phi} = (1/N)\sum_i^N LM_{\phi}(i)$, $\overline{LM}_{\tau} = (1/N)\sum_i^N LM_{\tau}(i)$, $LM_{\phi}(i) = T\hat{\phi}_i(\hat{\omega}_i/\hat{\sigma}_i)$, $LM_{\tau}(i) = \hat{\phi}_i/SE(\hat{\phi}_i)$, $\hat{\phi}_i$ is the least square estimates of ϕ_i in the equation (9) in Westerlund and Edgerton (2008), $\hat{\sigma}_i$ and $SE(\hat{\phi}_i)$ are the estimated standard errors of the same regression (9).

Appendix C. Panel estimation

In the third step of our analysis, we use the CUP-FM estimator:

$$\hat{\beta}_{CUP} = \left[\sum_{i=1}^{n}\left(\sum_{i=1}^{T}\hat{y}_{i,t}^{+}(\hat{\beta}_{CUP})(x_{i,t}-\overline{x}_i)' - T(\hat{\lambda}_i'(\hat{\beta}_{CUP})\hat{\Delta}_{F\varepsilon i}^{+}(\hat{\beta}_{CUP}) + \hat{\Delta}_{\mu\varepsilon i}^{+}(\hat{\beta}_{CUP}))\right)\right]$$
$$\left[\sum_{i=1}^{n}\sum_{t=1}^{T}(x_{i,t}-\overline{x}_i)(x_{i,t}-\overline{x}_i)'\right]^{-1} \tag{C.1}$$

This estimator makes corrections for endogeneity and serial correlation. The endogeneity correction is achieved by modifying the original variable y_{it} as follows: $\hat{y}_{i,t}^{+} = y_{i,t} - (\hat{\lambda}_i'\hat{\Omega}_{F\varepsilon i} + \hat{\Omega}_{\mu\varepsilon i})\hat{\Omega}_{\varepsilon i}^{-1}\Delta x_{i,t}$. The CUP-FM is constructed by estimating parameters, long-run covariances matrix (Ω) and factor loadings (λ_i). Thus $\hat{\beta}_{FM}$, $\hat{\Omega}$ and $\hat{\Lambda}_i$ are estimated repeatedly, until convergence is reached.

Appendix D. Panel regression residuals

Standardized residuals from the panel regression are shown in Fig. 4.

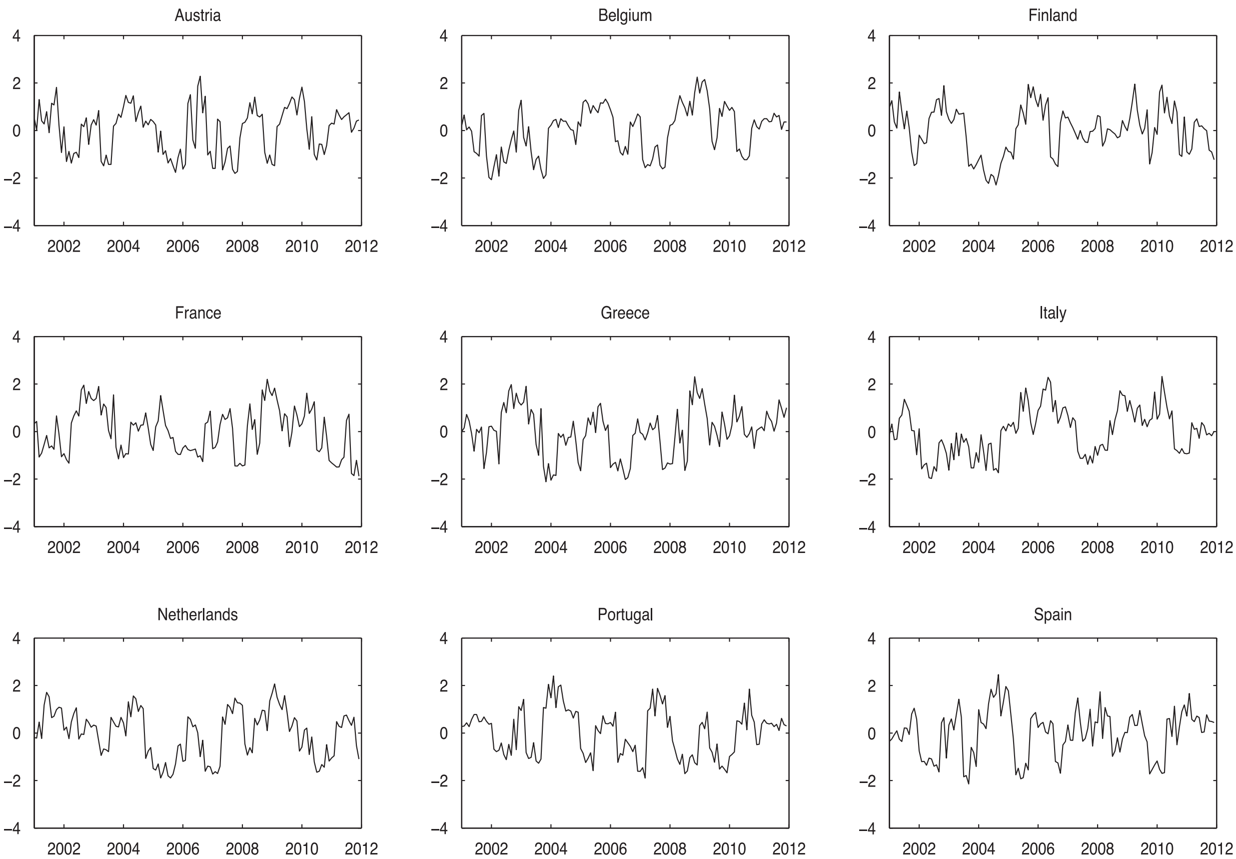

Fig. 4. Standardized residuals from the panel regression.

Appendix E. Supplementary data

Supplementary data associated with this article can be found in the online version at http://dx.doi.org/10.1016/j.euroecorev.2014.06.004.

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