Int. Fin. Markets, Inst. and Money 22 (2012) 278–304

Contents lists available at SciVerse ScienceDirect

Journal of International Financial Markets, Institutions & Money

journal homepage: www.elsevier.com/locate/intfin

The determinants of sovereign credit spread changes in the Euro-zone☆

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ARTICLE INFO

Article history:
Received 13 November 2010
Accepted 30 September 2011
Available online 14 October 2011

JEL classification:
C22
E45
E43
C15

Keywords:
Credit risk
Structural models
Term structure of interest rates
Gaussian HJM multi-factor models

ABSTRACT

Using a database of Euro-denominated government bonds covering the period from January 2000 to December 2010, this paper provides an empirical analysis of the determinants of government credit spreads in the Euro-area. The analysis is divided into two sub-periods delimited by the global financial crisis that started in August 2007. We find evidence of a clear shift in the behavior of market participants from a convergence-trade expectation, based on market related factors, before August 2007, to one mainly driven by macroeconomic country-specific variables and an international common risk factor. There is no evidence of a significant role for the liquidity risk before or during the financial crisis period. Overall, our results give support to the Merton-type structural credit risk models and confirm that there are considerable similarities between the factors explaining the dynamics of the credit risk spreads and the factors driving the prices on the government bond markets.

1. Introduction

Before the introduction of the Euro, the Euro-zone capital markets, specially the sovereign bond markets, were very heterogeneous. Bond issues were made primarily in domestic currency and there were substantial differences in liquidity and technical, legal, regulatory and supervisory practices

☆ The authors would like to thank the Editor and the anonymous referee for very useful suggestions and comments. We also thank In Choi for providing us the Gauss code to compute the subsampling versions of the LLC and IMC panel unit root tests.

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doi:10.1016/j.intfin.2011.09.007

within the pre-EMU sovereign debt markets. Sovereign issues with quite similar characteristics commonly exhibited high yield spreads.

The introduction of the Euro had different consequences on the yield curves of Euro-zone countries. New euro-denominated debt was issued and, at the same time, all outstanding debt was redenominated in the new currency, leading to the disappearance of currency risk premiums of inflation or devaluation. Markets widely expected that the sovereign risk premiums would narrow, especially those for the long-term interest rates. Since the start of third stage of the EMU and until the beginning of the recent financial crisis, investors believed that without any country-specific risk and with a high standardization of a large Euro-denominated bond market, the different government debt issues of the same maturity would become perfect substitutes. Hence, the liquidity risk would tend to be eliminated and the sovereign credit risk should narrow even further. Effectively, at the short end of the curve, yields converged due to suppression of money market competition and the long term government bond spreads for the Euro-zone countries relative to the German zero bond curve converged and stabilized at very low levels.

Despite the strong convergence of the long term yields on public debt observed in the period of transition for the EMU regime, mainly arising from the convergence of the non-core EMU participants, yield differentials have not disappeared completely. After September 2008, with the intensification of the financial crisis and the deterioration of the Euro-area macroeconomic fundamentals, in particular for the peripheral countries, sovereign bond spreads started to wide significantly. Due to the fact that the US economy has the largest and most mature bond market in the world, the great majority of empirical studies on the determinants of credit spreads are concentrated on the US bond market. However, since the inception of EMU, in January 1999, a large range of empirical literature concerning the determinants of Euro-zone government bond yield spreads has been developed, and increased substantially following the onset of the global financial turmoil in August 2007, and particularly after September 2008.

Credit risk is expected to explain a considerable portion of EMU sovereign credit spreads, even though its role changed over time and especially since the beginning of EMU. Lemmen and Goodhart (1999) work with differentials between redemption yields and swap rates to identify factors that explain the default risk between EU countries. They find a strong significant correlation between credit spreads and the ratio of debt-to-GDP, inflation, government expenditures, and level of taxation. Using a credit risk model, which measures the sovereign credit spreads on 10-year zero-coupon yields estimated on a panel dataset of seven EMU countries, Van Landschoot (2004b) investigates the relationship between sovereign credit spreads and the composition of the government budget. The main finding is that governments with a broader tax base, higher investments, and less spending on consumption, social security and subsidies have significantly lower credit spreads.

Collin-Dufresne et al. (2001) investigate which factors determine the changes of credit spreads on industrial US corporate bonds. They find that the determinants within structural bond pricing models can explain only a small portion of credit spread changes and, therefore, conclude that the credit spreads are mostly determined by a single common factor, that is not covered by the bond pricing model. In their own words, this dominant component is driven by local/demand shocks that are independent of credit and liquidity risks.

Boss and Scheicher (2002) study the determinants of the pricing process in corporate bond markets, focusing the analysis on the Euro-area and, for comparative purposes, in the US. They define credit spreads as the difference between yields of corporate bonds (from industrial and financial issues) and government bonds. Although they identify the presence of a sizeable unobserved explanatory component, they conclude that factors based on yields of German government bonds play an important role in explaining the movements of Euro credit spreads. In addition, they find that stock returns, the volatility of stock returns and liquidity drive a significant influence in credit spread changes.

Huang and Kong (2003) examine the determinants of corporate bond credit spreads using option-adjusted spreads for nine Merrill Lynch corporate bond indexes, and confirm that credit spread changes for high yield bonds are more closely related to equity market factors than to macroeconomic factors.

Codogno et al. (2003) highlight the role of the international risk in determining spreads against Germany, and conclude that the risk of default is small but an important component of yield differentials because it imposes the market discipline in countries' fiscal policies. Bernoth et al. (2004) study the interest differentials between bonds issued by European countries and bonds issued by Germany or by the US. Using a sample data from before and after the start of EMU, they estimate the effects of the monetary union on the risk premiums paid by European governments. Their results point out that the default risk premium is positively affected by debt and debt service ratios of the issuer country.

The importance of the international risk factor and the investors risk aversion in explaining the EMU sovereign spreads have been also revealed in several studies—see, for instance, Geyer et al. (2004), Barrios et al. (2009), Sgherri and Zoli (2009), Attinasi et al. (2009) or Favero et al. (2010). As noted by Haugh et al. (2009), this international effect is particularly strong during times of financial distress and for countries with high levels of public debt and budget deficits—see also Barrios et al. (2009), as well as Schuknecht et al. (2010).

Liquidity is another possible explanation for government residual bond yield differentials across EMU members, but the conclusions are somewhat controversial. Some authors found that liquidity has, at best, a minor explanatory power in the behavior of yield spreads—see Geyer et al. (2004), Bernoth et al. (2004) or Jankowitsch et al. (2006). By contrast, other studies provide strong evidence in favor of a prominent liquidity effect—see Gómez-Puig (2006), Manganelli and Wolswijk (2009) or Beber et al. (2009). Favero et al. (2010) argue that the liquidity effect cannot be the full answer for the bond yield differentials' behavior. Hence, they include in their model a term that interacts liquidity variables with the international risk factor, and provide empirical evidence of a negative and significant impact for that interactive term.

The contributions of this paper to the existing literature are threefold. First, to evaluate the impact of the financial crisis on sovereign credit spreads, we analyze separately two sub-periods: before the financial crisis period—January 2000 to July 2007—and the financial crisis period—August 2007 to December 2010.[1] Second, our paper estimates the sovereign credit spreads as the difference between the term structure of interest rates of seven EMU-member governments—Austria, Belgium, France, Italy, the Netherlands, Portugal and Spain—and the German government term structure of interest rates. For this purpose, zero-coupon spot yield curves are derived for each country using a term structure specification that is consistent with an arbitrage-free Heath et al. (1992) Gaussian multifactor term structure model. Working with zero-coupon spot rates has the advantage that, unlike an yield to maturity, they are not affected by bond characteristics—like the coupon rate or maturity. Third, we analyze the determinants of credit spread changes for 5-, 10- and 15-year maturity buckets, even though the majority of the empirical studies focus solely on the 10-year yields.

Prior to the financial crisis, our results seem to confirm that there are considerable similarities between the factors explaining the dynamics of the credit risk spreads and the factors driving the prices on the government securities markets, giving support to the structural credit risk models. During the crisis period, our results point to a shift to country's macroeconomic fundamentals as the main determinants in explaining the changes in the credit spreads.

Our paper proceeds as follows. Section 2 presents the theoretical framework for the determinants of sovereign default risk. In this section we identify the main determinant factors and predict their expected influence on the sovereign default risk premium. Section 3 describes the term structure extraction methods used to estimate each sovereign yield curve. Section 4 presents the estimates of the term structure of credit spreads for the seven EMU-countries under analysis. Section 5 describes the econometric models used and reports the empirical results. Finally, Section 6 concludes.

2. Sovereign credit spreads' determinants: theoretical background

The factors that we use as explanatory variables can be divided into two categories: common explanatory variables and country-specific explanatory variables. In the first set we include equity related variables and interest rate risk sensitivity variables. These are factors that reflect the systematic or market component of the credit risk level and are partly motivated by the structural models of credit risk introduced by Merton (1974). As mentioned by Collin-Dufresne et al. (2001, p. 2179), Boss and

[1] The onset of the financial crisis is generally accepted to be late July 2007. On August 2007, the European Central Bank provided the first large emergency loan to banks in response to increasing pressures in the euro interbank market.

Scheicher (2002, p. 183), and Van Landschoot (2004a, p. 10), the factors that are currently included in regression equations to investigate the behavior of the credit spreads are partly motivated by the intuitive framework of the so-called "structural models of default risk", which relate credit risk to fundamental corporate variables, namely to the value of the firm's assets. These models were primarily introduced by Merton (1974) and further investigated by other authors including Black and Cox (1976), Longstaff and Schwartz (1995), Leland (1994) and Collin-Dufresne and Goldstein (2001). It is assumed that the firm value follows some stochastic process and default is triggered when the firm value falls bellow some threshold that is a function of the amount of outstanding debt. This framework allows the application of the option pricing theory to the pricing of a risky bond. The fundamental insight of Merton's model is that the issue of a risky zero-coupon bond by the firm has the same payoff structure as a risk-free bond plus a put option on the firm's value, with a strike price equal to the face value of the firm's debt. In other words, the value of the put option is the cost of eliminating the credit risk.

The second group of variables were chosen to better capture the country-specific risk level and were based on the economic literature and/or empirical reasoning. In this set of explanatory factors we include the liquidity risk measure and six macroeconomic variables to measure the fundamental risk—namely the variables that account for the level of public debt and current account deficit, the government expenditure and revenues. To measure the economic climate we use a proxy for the business cycle—output growth—and, to control for the inflation effect, we consider the level of consumer prices. In the following subsections we briefly describe the variables that we will use in our regression models.

2.1. *Equity-related variables*

2.1.1. *Asset value and the firm's leverage ratio*

Structural models consider the firm's asset value as one of the most important variables that accounts for the level of the firm's credit spread. Since default is triggered when the leverage ratio[2] equals unity, we expect the assets value and the credit spread to be negatively correlated. In fact, firms that experiment an increase in their assets market value should present a higher return on equity and are more unlikely to default. As pointed out by Collin-Dufresne et al. (2001), all else equal, credit spreads should be a (decreasing) function of the firm's return on equity. Since the market assets value is not an observable variable, it is usually replaced by the equity return of publicly traded and most liquid companies included on a representative stock index. The evolution of the returns of those companies should be reflected by the evolution of the Dow Jones EURO STOXX 50 index.[3]

2.1.2. *Asset volatility*

The market value of debt has features that are similar to a short position in a put option on the firm value. Since the value of the put option increases with volatility, we expect a positive relation between the asset volatility and the credit spread. This prediction is very intuitive: an increasing volatility reflects an increase in the market risk level and, hence, in the probability of default. Furthermore, the role of a common international risk factor has been referred in several studies as an important explanatory variable—Codogno et al. (2003), Geyer et al. (2004), Barrios et al. (2009), Attinasi et al. (2009), Arghyrou and Kontonikas (2011). To proxy the Euro-area implied common volatility, we use the Dow Jones EURO STOXX50 Volatility Index (VSTOXX) which measures the Euro-area stock market implied volatility.[4]

[2] In practical terms, the leverage ratio is the market value of debt divided by the market value of the firm's assets.

[3] This index measures the performance of the Euro-zone equity market and is computed in Euros. The index covers the 50 largest sector leaders in the Euro-zone based on free float market capitalization. For further details about the index components and calculation issues, please see STOXX (2006).

[4] Arghyrou and Kontonikas (2011) use, as a general indicator of a common international risk factor, the Chicago Board of Options Exchange volatility Index (VIX). For specific details on the VSTOXX index methodology, please see STOXX (2005).

2.2. *Interest rate sensitive variables*

According to structural models of credit risk, the value of a put option on the firm value—computed under the Black–Scholes–Merton framework—largely depends on the level of the risk-free rate. In order to capture the impact of the dynamics of the benchmark (i.e., the German) yield curve, and in accordance with Litterman and Scheinkman (1991), we use three factors: the level, the slope, and the curvature of the yield curve.

2.2.1. *Interest rate level*

As pointed out by Longstaff and Schwartz (1995), the marginal effect of a higher level of the spot interest rate is the narrowing of the credit spreads. The reason for this is that the drift of the risk-neutral process for the firm value increases as the spot rates increase, turning the probability of default lower. Hence, the price of the put option on the firm's asset value—i.e., the cost of default insurance—diminishes, implying a lower credit spread. We use the level of interest rates, instead of the short-term rate, because the part of the government debt that is financed with short-term maturities is relatively limited. As a proxy for this variable we use the German 10-year spot rate.

2.2.2. *Slope of the term structure*

Although the riskless spot interest rate is the unique interest rate sensitive factor identified in the process of determining the firm's asset value, other factors are relevant when describing the dynamics of the term structure. The importance of the slope of the riskless term structure can be supported by two arguments. First, an increase in the yield curve slope reflects an increase in the expected future short-term interest rates, which should lead to a decrease in the sovereign spreads. Second, according to the interest rate expectations theory, a negative yield curve slope reflects a diminishing maturity premium, which may imply a weakening economy, which in turn is conducive to a lower growth rate of the firm's assets and to an increase of the credit spread. The slope of the German yield curve is the proxy we include in the empirical analysis and is defined as the difference between the German 15-year and 1-year spot rates.

2.2.3. *Curvature of the term structure*

The third factor in the category of interest rate sensitive variables is the curvature of the yield curve as a measure of the convexity of the term structure. We define the curvature factor as the difference between the 3.5-year spot rate and a synthetic spot rate for a maturity of 3.5 years interpolated between the 15-year and the 1-year spot rates. It is not clear if the sign of the coefficient should be positive or negative, i.e., the magnitude of the credit spreads may fall or rise if the curvature increases or decreases. As pointed out by Collin-Dufresne et al. (2001), an increase in the squared yield of the 10-year government bonds (used as a proxy for the convexity of the term structure) has a negative effect on the credit spreads of bonds with short maturities and a positive effect on bonds with long maturities.

2.3. *Liquidity measure*

Liquidity risk is a portion of the total spread between a risky spot rate and a riskless spot rate of a certain maturity. Frequently, the bid-ask spread is stated as the appropriate measure of liquidity in the market. For instance, if one considers two bonds with the same credit risk level and with similar characteristics—like maturity, coupon, embedded options and provisions—they could trade at a different price and, thus, at different yields since their liquidity is different. Amihud and Mendelson (1991) focus on liquidity effects in fixed-income US markets and show that the yield of the treasury notes is higher than the ones on the more liquid treasury bills with the same residual maturities. Therefore, the spread of these bonds can be interpreted as the liquidity premium demanded by the investors for the immediate trade execution. In fact, although structural models of credit risk do not assume the existence of liquidity premiums, several studies find evidence that liquidity strongly influences the yield differentials dynamics—see for instance Collin-Dufresne et al. (2001), Boss and

Scheicher (2002), Codogno et al. (2003), Van Landschoot (2004a), Gómez-Puig (2009), Attinasi et al. (2009), Favero et al. (2010), and Arghyrou and Kontonikas (2011).

The literature acknowledges a high degree of collinearity between empirical measures of liquidity and other risk factors. To avoid possible collinearities, we capture the liquidity effect directly from the interest rate model used to estimate the spot yield curves for each country. Following Theobald et al. (1999), we use a liquidity measure that implicitly reflects the liquidity information of the whole term structure. This measure corresponds to the relative pricing error derived from the difference between market bond prices and the model-fitted bond prices. Hence, we measure the liquidity risk using the Weighted Mean Absolute Pricing Error (*WMAPE*) obtained in the estimation of the yield curve for each country. This measure should be a good proxy for the level of liquidity on each day because it uses indirectly the bid-ask spreads and is measured along the entire yield curve. If the liquidity is high, the bonds' bid-ask spreads tend to narrow and the fitted prices tend to be closer to the market prices, showing a smaller pricing error. On the other hand, if the liquidity dries, the bid-ask spreads tend to widen and, hence, the fitted prices tend to show evidence of larger pricing errors. Therefore, we expect the regression coefficient of the liquidity measure to be positive.

2.4. Debt service

In structural models, the distance to default measures the firm's default risk and is a function of the firm's leverage ratio. If the firm's assets value drops to the level of the firm's outstanding debt, default occurs. This probability is relatively low as long as the debt stays below a critical level. If the firm continues borrowing and the debt exceeds that critical level, the interest rate begins to rise as the credit spreads go up. In an extreme situation, a firm that is excessively in debt can be denied to assess the credit markets.

The same reasoning mechanism can be applied to the level of sovereign indebtedness and to the country's probability of default. As pointed out by Goldstein and Woglom (1991) and in line with the market-based fiscal discipline hypothesis, the same credit constraints can play a positive role in disciplining sovereign borrowers. In particular, the advocates of the market discipline assume that yields will rise smoothly at an increasing rate with the level of borrowing when sovereign debt stays bellow a certain critical level. If this critical level is exceeded, the credit markets will respond by imposing an increasing default risk premium until, eventually, denying the irresponsible borrower access to additional credit. The increase in the cost of borrowing would then provide the incentive to correct irresponsible fiscal behavior.[5] Market discipline of this kind is expected to be particularly relevant in the EMU because each member state can issue debt, but none has the ability to monetize, and hence, to inflate away the excessive debt.

Because the credit spreads are expected to increase with leverage, it is reasonable to believe that sovereign credit spreads should be an increasing function of the debt level. As a proxy for this variable we use the debt, including interest payments, relative to GDP.

2.5. Current account deficit

The current account balance, and the respective deficit or surplus relative to GDP, gives an indication of the change in the indebtedness position of a country. A persistent deficit in current account implies that private domestic investment is being persistently externally financed. It is reasonable to expect that the greater the current account deficit, the higher the perceived probability of default. Moreover, the widening of the current account balance deficit reflects a loss of country competitiveness that is penalized by higher credit spreads.

Cantor and Packer (1996) find a positive (but statistically insignificant) correlation between the current account deficit and sovereign credit ratings. Aßmann and Boysen-Hogrefe (2009) show that

[5] Advocates of the market approach, such as Bishop et al. (1989), recognize that the market discipline approach will work only if certain conditions are satisfied. Bayoumi et al. (1995) report empirical evidence in US states on the key question of whether credit markets impose sufficient default premia to restrain irresponsible borrowing.

the fiscal budget balance as well as the current account balance are very important to explain the risen ten-year spreads among Euro-area countries. Barrios et al. (2009) highlight the strong significance role played by the deterioration of the current account balance relative to Germany which, in their own estimates, can lead to a rise of 1.3 basis points (b.p.) in the spread of a given country. Arghyrou and Kontonikas (2011) also confirm that an improvement of the expected fiscal position (movement from net borrowing towards net lending) leads to lower spreads, with the respective coefficient being negative and statistically significant. We use the current account deficit, instead of the fiscal budget balance, to capture simultaneously the effects of domestic saving capability and the competitiveness level dynamics of each given country.

2.6. Government expenditures

Public-policy endogenous growth models highlight the impact that the composition of government expenditure has on both the level of output and the long-run growth rate. Several empirical studies show the relative importance of the productive expenditures in determining the long-term growth rate. For instance, Kneller et al. (1999) use a panel of 22 OECD countries and find that productive government expenditures enhance growth.[6] Following the main hypothesis of Van Landschoot (2004b), we also test three components of the government budget as potential determinants of the level of sovereign credit spreads: government consumption, government investment, and government social expenditures.

One of the most robust findings, frequently documented in the existing literature on the effects of fiscal policy on growth, is the negative relation between government consumption and output growth, which is induced by the reduction of private savings. Therefore, government consumption is expected to increase sovereign spreads because it tends to reduce future growth. On the other hand, government investments, which are considered as productive expenditures, have a positive impact on productivity and on future long-run growth—see, for instance, Easterly and Rebelo (1993), Evans and Karras (1994) or de la Fuente (1997). In accordance with the endogenous growth models, government investment is expected to reduce sovereign credit risk because it favors future growth. By contrast, Kneller et al. (1999) identify social expenditures as the main component of the non-productive expenditures, and document a positive, although not significant, correlation with economic growth.

2.7. Other determinants

2.7.1. Level of taxation

Theory predicts that the impact of fiscal policy on growth depends on the structure as well as the level of taxation and expenditure. Kneller et al. (1999) find that distortionary taxation on investment and income significantly reduces growth by reducing the private returns on capital accumulation, whereas nondistortionary taxation has an insignificant impact. From this point of view, it seems that the level of taxation has a negative correlation with the long-term growth and, hence, may increase the sovereign credit spreads. On the other hand, and in accordance with Lemmen and Goodhart (1999), because governments could use tax collection instead of printing money as means of raising funds to repay debt service obligations, the level of taxation should be negatively correlated with sovereign spreads. We use, as a proxy for the level of taxation, the total government tax receipts relative to GDP.

2.7.2. Business cycle

Since a higher output growth denotes a favorable business cycle and anticipates an economy expansion, the government's creditworthiness will improve because the future debt burden becomes lower. Inversely, if a country has experienced a slow or negative growth, it will likely find it more difficult to meet payments on its external debt. Alesina et al. (1992), based on monthly percentage changes

[6] Government expenditures are differentiated according to whether they are included as arguments in the productive function or not. If they are, then they are classified as productive and hence they have a direct effect on the growth rate—see Kneller et al. (1999) for details.

of the industrial production index, find a significant negative correlation between the growth rate of industrial production and the spread differential between public and private debt. Geyer et al. (2004) point out the importance of the growth rate of industrial production as a measure of the state of the business cycle and find a positive, but not statistically significant, effect for the respective coefficient. As a proxy of the business cycle, we use the Industry Production Index—NACE Rev. 2).[7]

2.7.3. Level of consumer prices

Another variable that may affect sovereign spreads is the level of inflation. In fact, a high rate of inflation points to structural problems in government's finances since it reduces the real value of the accumulated debt. If investors anticipate future inflation, however, they will demand higher interest rates on government debt, making public borrowing more expensive. Cantor and Packer (1996) find a negative, although not significant, correlation between inflation and sovereign credit ratings. Lemmen and Goodhart (1999) find similar results with respect to the rate of inflation, and a significant positive correlation between the variation of the rate of inflation and the changes of sovereign spreads. To capture the evolution of prices in each country, we use the respective Harmonized Index on Consumer Prices: HICP—all items—index (2005 = 100).[8]

3. Term structure extraction method

In order to obtain and compare the credit spread changes of bonds issued by different governments and with different maturities, we estimate the term structure of spot interest rates for the eight Euro sovereign debts under analysis. The sovereign spreads are calculated as the difference between the term structure of spot interest rates on each considered sovereign debt, and the German term structure, which is fixed as the benchmark.[9] The use of zero-coupon spot rates affords several advantages to our analysis. First, it eliminates the coupon effect and the ambiguous phenomenon whereby the yields to maturity of bonds with the same maturity and credit risk, but different coupons, may vary considerably. Second, it does not require any assumptions about the reinvestment rates applicable to capitalizing intermediate cash flows. Finally, all maturities are identified with a unique interest rate.

All bonds are issued by the treasury and provide a stream of certain cash flows at known times in the future. The ith observed time-0 bond price is denoted by $B_i(0)$, while the fitted bond price is expressed as $\widehat{B}_i(0)$. This bond provides future cash flows $c_{i,j}$ at times t_j for $j = 1, \ldots, m_i$. The discount function $P(0, t_j)$, and henceforth the spot rates, is extracted from these observed bond prices by imposing the static no-arbitrage condition appropriate to a world without taxes, embedded options, or other frictions. This simple pricing function reflects the present value of the promised future cash flows:

$$\widehat{B}_i(0) = \sum_{j=1}^{m_i} c_{i,j} P(0, t_j). \tag{1}$$

As noted by Bliss (1997), because real markets (from which we collect our bond data) are not frictionless, in practice we do not use an exact pricing relation such as Eq. (1), but rather the following inexact relation:

$$B_i(0) = \widehat{B}_i(0) + \varepsilon_i, \tag{2}$$

where ε_i is a random error term.

[7] Eurostat publishes, on a monthly basis, the industrial production index which is a business cycle indicator that measures monthly changes in the price-adjusted output of industry. Currently the indices for industrial production are calculated with 2005 as the base year (=100).

[8] The HICP for each EMU-member consists of a breakdown of final individual consumption of goods and services, and covers the monetary expenditures of households in the economic territory of the European Union countries.

[9] This choice could seem arbitrary since, between 2000 and 2010, the French government bonds are regarded as the proper Euro benchmark in several periods. Geyer et al. (2004) also use the German yield curve as the reference and pose the same question. They check whether the choice of the reference curve would have any effect on their results and conclude that, due to the remarkably high correlation between the two spot yield curve levels, the choice of the benchmark should not affect the results.

3.1. HJM Gaussian and multifactor term structure model

Many different functional forms can be used to estimate discount factors from the observed market prices of treasury coupon-bearing bonds—see, for instance McCulloch (1971, 1975), Nelson and Siegel (1987), Fama and Bliss (1987), Vasiček and Fong (1982), Fisher et al. (1995), Waggoner (1997), or Jeffrey et al. (2006) for a survey. We use a parametrization proposed by Björk and Christensen (1999) which is consistent with a Gaussian and multifactor Heath et al. (1992) term structure model.

Such a model can be formulated in terms of risk-free pure discount bond prices, which are assumed to evolve over time (under the risk-neutral martingale measure $\mathcal{Q}$) according to the following stochastic differential equation

$$\frac{dP(t,T)}{P(t,T)} = r(t)dt + \underline{\sigma}(t,T)' \cdot d\underline{W}^{\mathcal{Q}}(t), \tag{3}$$

where $P(t, T)$ represents the time-t price of a (unit face value and default-free) zero-coupon bond expiring at time T, $r(t)$ is the time t instantaneous spot rate, $\cdot$ denotes the inner product in $\Re^k$, and $\underline{W}^{\mathcal{Q}}(t) \in \Re^k$ is a k-dimensional standard Brownian motion. The k-dimensional volatility function, $\underline{\sigma}(\cdot, T) : [0, T] \to \Re^k$, is assumed to satisfy the usual mild measurability and integrability requirements—as stated, for instance, in Lamberton and Lapeyre (1996, Theorem 3.5.5)—as well as the "pull-to-par" condition $\underline{\sigma}(u,u) = \underline{0} \in \Re^k$, $\forall u \in [0, T]$, where 0 denotes the current time. Moreover, for reasons of analytical tractability, such volatility functions are assumed to be deterministic.

Following, for instance, Musiela and Rutkowski (1998, Proposition 13.3.2), it is well known that if the short-term interest rate is Markovian and the volatility function $\underline{\sigma}(\cdot, T) : [0, T] \to \Re^k$ is time-homogeneous, then the volatility function must be restricted to the analytical specification

$$\underline{\sigma}(t,T)' := \underline{G}' \cdot a^{-1} \cdot [I_k - e^{a(T-t)}], \tag{4}$$

where $I_k \in \Re^{k\times k}$ represents an identity matrix, while $\underline{G} \in \Re^k$ and $a \in \Re^{k\times k}$ express the model' time-independent parameters. The Gauss-Markov time-homogeneous HJM model to be estimated is defined by Eqs. (3) and (4).

As shown in Nunes and Oliveira (2007, Proposition 4), under the assumption that matrix a is diagonal, the minimal consistent family (manifold) $\mathcal{G}$ of forward rate curves which is invariant under the dynamics of the HJM model (3) and (4), is defined by the mapping $\gamma : \Re^{2k} \times \Re_+ \to \Re$, such that

$$\gamma(\underline{z}, x) \equiv f(t, t+x) = \sum_{j=1}^{k} z_j \exp(a_j x) + \sum_{j=1}^{k} z_{k+j} \exp(2a_j x), \tag{5}$$

where z_j represents the jth element of vector $\underline{z} \in \Re^{2k}$, a_j defines the jth principal diagonal element of matrix a, and $f(t, t+x)$ corresponds to the time-t instantaneous forward interest rate for date $(t+x)$, with $x \in \Re_+$.

Under the specification of the forward rate curve provided by Eq. (5), parameters a and $\underline{z}$ can be estimated by minimizing the absolute percentage differences between a cross-section of market treasury coupon-bearing bond prices and the corresponding discounted values obtained by decomposing each government bond into a portfolio of pure discount bonds, which are parametrized as

$$P(t,T) = \exp\left\{ \sum_{j=1}^{k} \frac{z_j}{a_j}[1 - e^{a_j(T-t)}] + \sum_{j=1}^{k} \frac{z_{k+j}}{2a_j}[1 - e^{2a_j(T-t)}] \right\}. \tag{6}$$

Under this general formulation, the HJM model dimension is set at three factors. Hence, nine parameters are used in the discount factor specification of Eq. (6).

3.2. Bond data description

The bond data used to estimate the spot yield curve for each country was collected from Bloomberg (Bloomberg Generic, BGN). They consist of coupon-bearing treasury bonds bid and ask prices recorded

Table 1
Distribution of the number of bonds by country and by maturity buckets.

	Total	Avg	Stdev	Min	Max	<2 years	2–5 years	5–10 years	>10 years
Austria	84	25	11	8	57	7	8	7	3
Belgium	127	29	20	12	76	5	11	10	3
France	75	26	7	15	37	4	4	8	10
Germany	170	40	9	27	63	13	13	8	6
Italy	150	49	5	29	62	14	15	9	11
The Netherlands	54	26	5	12	35	6	8	8	5
Portugal	31	13	2	8	16	3	4	5	2
Spain	64	20	4	9	29	5	6	5	5

Table 1 reports the sample bonds distribution based on country and residual maturity. Total refers to the total number of bonds that remain after imposing the outlying and liquidity filters. Avg, Stdev, Min and Max are the average, standard deviation, minimum and maximum number of bonds that are considered per country. The number of bonds in each maturity bucket reflects the average number of bonds considered per country.

on the last day of each month (end of session) during the period between January 2000 and December 2010. We compare government bonds issued by the following eight EMU-countries: Austria, Belgium, France, Germany, Italy, the Netherlands, Portugal, and Spain.[10]

To mitigate problems associated with distorted prices from bonds not actively traded during the sample period, the robust outlier identification procedure proposed by Rousseeuw (1990) is used to exclude from the sample all illiquid issues.[11] Also, to avoid illiquidity problems usually associated with issues that are close to the maturity date, treasury bonds with a residual time-to-maturity of less than three months are excluded from the sample. To preserve the homogeneity of the data, we only consider fully-taxable and non-callable issues. These filters leave us with a dataset of 755 sovereign bonds issued by the eight countries under analysis. Germany (170), Italy (150) and Belgium (127) are the largest contributors for the total. Portugal (31) has the lowest number of bonds in the dataset. Table 1 reports the distribution of the bonds' sample size by country and residual maturity.

As Table 1 shows, after eliminating the bonds that do not pass the outlying and liquidity filters, bonds with a residual maturity beyond 10 years are sometimes sparse. On average, Italy and France are the main contributors for the long maturity buckets.

On each sample day, the term structure of interest rates is estimated by finding the values of the parameters that minimize the duration weighted mean absolute percentage pricing error (*WMAPE*), which reflects the weighted differences between fitted and market coupon-bearing bond prices. As noted by Bliss (1997), pricing errors for longer maturities tend to be larger. Because of the observed heteroskedastic behavior of fitted pricing errors, and acknowledging the theoretical relation between bond prices and interest rates, Bliss (1997) suggests weighting the percentage pricing errors by the inverse of the corresponding bond's duration to prevent that the errors of long-term bonds dominate the results:

$$WMAPE_t = \sum_{i=1}^{N_t} \frac{\left|e_i(t)\right| \times \omega_i(t)}{N_t}, \tag{7}$$

where $\omega_i(t) = 1/D_i(t)$, $e_i(t) = \frac{B_i(t) - \widehat{B}_i(t)}{B_i(t)}$, $D_i(t)$ denotes the duration of the ith bond at time-t, and N_t is the total number of bonds sampled at time-t. As stated in Eqs. (1) and (2), $\widehat{B}_i(t)$ and $B_i(t)$ represent the fitted price and the "mid-quote", respectively, of the ith bond.

[10] As there was a lack of market depth and/or an insufficient number of liquid government issues traded in the market during most of the period under analysis, Finland, Greece, Ireland and Luxembourg were excluded from the sample.

[11] On each day, the bid-ask spreads of all traded bonds for each country are standardized using the sample median (location estimator) and the median of all absolute deviations from the sample mean (scale estimator). Whenever the standardized score of a specific bond is higher than a pre-specified cutoff value (defined here as 2.5), the bond is automatically excluded from the panel data under analysis.

Table 2
Performance of the estimated yield curves.

	Austria	Belgium	France	Germany	Italy	The Netherlands	Portugal	Spain
Panel A: full period								
Average	10.8	7.1	9.2	6.1	15.8	7.8	12.1	11.5
St. deviation	12.8	4.0	11.6	5.7	12.7	7.9	10.0	7.8
Minimum	1.2	1.2	0.6	0.6	2.2	0.9	1.0	1.6
Maximum	40.7	31.5	24.6	16.4	43.1	27.5	45.8	38.0
Panel B: before crisis period								
Average	6.0	5.6	4.2	3.4	11.1	5.6	6.9	7.8
St. deviation	6.3	2.4	3.1	2.1	11.7	4.9	3.1	3.8
Minimum	1.2	1.2	0.6	0.6	2.2	0.9	1.0	1.6
Maximum	28.1	17.4	19.3	11.2	23.3	15.2	21.6	23.4
Panel C: crisis period								
Average	12.6	10.6	12.5	8.2	21.5	12.2	14.3	13.0
St. deviation	11.9	4.8	6.6	6.3	27.2	7.7	9.1	8.0
Minimum	6.1	3.1	5.7	4.3	10.4	4.4	6.1	5.8
Maximum	40.7	31.5	24.6	16.4	43.1	27.5	45.8	38.0

This table presents the mean absolute percentage pricing errors (expressed in basis points) resulting from the daily term structure estimation model given by Eq. (6). Panels A, B, and C report the performance of the estimated yield curves during the full period, from January 2000 to December 2010, the before crisis period, extending from January 2000 to July 2007, and the crisis period, covering the period from August 2007 to December 2010.

4. Estimation of the term structure of credit spreads

To obtain the yield curves for the eight EMU countries under analysis, the parameters contained in Eq. (6) are estimated by minimizing the *WMAPE* statistic. Throughout the empirical analysis, all optimization routines are based on the quasi-Newton method, with backtracking line searches, described in Dennis and Schnabel (1996, section 6.3).

4.1. Empirical results for the term structures of interest rates

The parametrization adopted fits the discount functions implicit in the eight government bond markets rather well, resulting in reliable and smooth yield curves for the sample period under analysis.[12] In order to validate the previous assertion, the mean absolute percentage pricing error (*MAPE*), obtained by minimizing the *WMAPE* measure, for the period January 2000 to December 2010, is presented in Panel A of Table 2. The model reaches a maximum in-sample *MAPE* value of 46 b.p. for the Portuguese sovereign debt and a minimum *MAPE* of 0.6 b.p. for the French and the German sovereign debts. The average *MAPE* value ranges from 6 b.p. for the German government bonds to approximately 16 b.p. for the Italian yield curve.

To show the impact of the financial crisis in our yield curve estimations, we divide the entire period into two sub-periods: the before crisis period, which was the period preceding the beginning of the global credit crunch (January 2000 to July 2007), and the during crisis period (August 2007 to December 2010). Before the credit crisis the pricing errors stabilized at very low levels and the term structure estimations for each country are more robust than in the crisis period. During the pre-crisis period, the average *MAPE* ranged from about 3 b.p. in Germany to 11 b.p. in Italy, with five countries—Austria, Belgium, France, Germany, and the Netherlands—exhibiting an average *MAPE* of no more than 6 b.p.

After the beginning of the credit crisis, the average pricing errors and the respective standard deviations of the estimated yield curves increased by two or three times, reflecting the augmented risk exposure and the increasing price volatility that have been observed in European bond markets. In

[12] To test how well the models describe the underlying term structures, we compute two in-sample statistics—the percentage of bonds with a pricing error outside a 95% confidence interval, and the persistence of the errors that is captured by estimating the conditional and the unconditional frequency of pricing errors—as well as an out-of-sample performance (*WMAPE*) test. To save space the results are not reported here but are available upon request.

Table 3
Estimated average credit spreads for each EMU-member and maturity bucket.

		Maturity buckets		
		5-year	10-year	15-year
Austria	Full period	17.7	23.0	18.8
(AAA)	Before crisis	8.6	9.3	15.1
(AAA)	During crisis	37.4	41.2	30.1
Belgium	Full period	24.2	29.0	21.4
(AA+)	Before crisis	5.6	10.1	14.8
(AA+)	During crisis	56.9	54.0	35.5
France	Full Period	8.9	14.4	12.9
(AAA)	Before crisis	2.7	3.7	8.0
(AAA)	During crisis	21.1	28.7	23.3
Italy	Full period	36.3	51.4	44.9
(AA)	Before crisis	8.5	17.9	25.7
(A+)	During crisis	79.1	93.3	89.4
The Netherlands	Full period	9.1	15.1	16.8
(AAA)	Before crisis	3.1	5.0	13.8
(AAA)	During crisis	18.3	27.2	23.3
Portugal	Full period	42.8	51.6	65.7
(AA)	Before crisis	8.8	15.8	36.9
(BBB+)	During crisis	114.4	118.6	109.3
Spain	Full period	26.8	34.7	33.2
(AA+)	Before crisis	2.8	10.1	12.8
(AA)	During crisis	74.8	77.9	61.6

This table presents the credit spreads term structure, expressed in basis points, for different EMU members. The full period covers January 2000 to December 2010, the before crisis period extends from January 2000 until July 2007, while the crisis period covers August 2007 to December 2010. Credit ratings were obtained from S&P and reflect the end of period notations on long term bonds.

fact, during the crisis period, the German average *MAPE* increases to about 8 b.p., and the average *MAPE* of Belgium reaches approximately 11 b.p., being the second lower. France, Austria, the Netherlands, and Spain present an average *MAPE* between 12 and 13 b.p. Italy has the highest average *MAPE* of 22 b.p., and Portugal has the second highest average *MAPE* of 14 b.p.

4.2. Term structure of sovereign spreads

Table 3 reports the spreads relative to the German yield curve for the seven EMU-countries being analyzed. The credit ratings of Euro-denominated EMU sovereign bonds remained relatively high for Austria, Belgium, France and the Netherlands, driving the spreads among EMU government bonds by a relatively low level, particularly in the maturities lower than ten years. On the other hand, the "high yielders" Portugal, Spain, and Italy have substantially higher credit spreads than the "low yielders" the Netherlands, France, and Austria.

As Table 3 indicates, and considering the full sample period, all EMU-countries show increasing spreads relative to longer maturity buckets. As expected, the "higher rated" EMU-countries have lower spreads than "lower rated" EMU-countries. Mean spreads across issuers range from about 9 b.p. for the 5-year maturity to approximately 66 b.p. for the 15-year maturity. France presents the lower mean spread curve and, in line with being the lowest rated issuer in our sample, Portugal has the highest mean spread curve for all maturities. The 5-year spreads range from 9 b.p. in the case of France and the Netherlands to 43 b.p. in Portugal. For the long maturity buckets the spreads are higher. The 10-year maturity spreads range from about 14 b.p. in the case of France, to more than 51 b.p. in the cases of Italy and Portugal. France has the minimum spread in the 15-year maturity bucket (13 b.p.), while Portugal and Italy present the maximum 15-year spread of 66 b.p. and 45 b.p., respectively. As expected, the dispersion of sovereign spreads among countries is higher for longer maturities. The differences on the means on 10- and 15-year credit spreads are not statistically different from zero at 5% level only for France, the Netherlands and Spain.

To make sure that our credit spread estimates are robust, we check the results obtained with the quoted 5- and 10-year credit default swaps (CDS) available for the countries under analysis. The comparison indicates that, although our estimated spreads are not of the same order of magnitude, they are highly correlated with the CDS' market quotes. In fact and except for the Netherlands, the linear correlation coefficients between our estimated credit spreads and the CDS' market are higher than 95% for all other countries. For the Netherlands we obtain the lowest correlation coefficients of 71% and 79% for 5- and 10-year buckets, respectively. On average, the 5- and 10-year spreads vs. CDS' quotes reach a maximum difference of about 12 b.p. for Spain and, at a significance level of 10%, the null difference could not be rejected in any country under analysis.[13]

Our results present two markedly different patterns of credit spreads corresponding to the before and during crisis periods. In the pre-crisis period, the average short maturity spreads ranged from 3 b.p. in France, the Netherlands, and Spain to approximately 9 b.p. in Italy, Portugal, and Austria. The 10-year maturity bucket shows average credit spreads ranging from 4 b.p. and 5 b.p. in France and the Netherlands, respectively, to 18 b.p. in Italy. The 15-year spreads are higher than the 10-year spreads but show a similar pattern. The long maturity spreads have somewhat widened due to increased liquidity problems in the long maturity issues, because small EMU members that are committed to improving public finances were not in a position to provide the necessary liquidity across the yield curve. This constraint explains an important part of the long term sovereign spreads detected in the present analysis. The somewhat lower spreads observed between January 2000 to July 2007 evidence the market perception of a fully credible EMU commitment to the bail-out of its member states, resulting in a detachment of the macroeconomic fundamentals from government bonds interest rates.

After the beginning of the financial crisis, the credit spreads increase substantially, reflecting the deterioration of the macroeconomic fundamentals, particularly the strong debt levels as well as the budgetary and the current account deficits of the peripheral Euro-area countries. Effectively, this situation has been particularly pronounced in Greece, Ireland, Portugal, and Spain where the fiscal budget deficits increase significantly. Since the onset of the financial crisis, all Euro-area economies have experienced a substantial increase in their spreads relative to Germany. Table 3 shows that, after the beginning of the financial crisis, average credit spreads increased in all countries and for all maturity buckets. For some countries the average spread increases dramatically by three or four times. Portugal is the country where the credit spread relative to Germany most rises during this period. Fig. 1 plots the 10-year spread relative to Germany for the seven EMU-countries under analysis. Before the credit crisis, spreads have stabilized at very low levels despite the evident deterioration of the macroeconomic fundamentals in most countries. This pattern is no longer present after the beginning of the credit crisis, when the spreads take increasingly high values reaching a peak at the end of 2008. After the end of 2008, the 10-year spreads have somewhat stabilized in Austria, France, and the Netherlands, but continue to increase in Belgium, Italy, Portugal, and Spain.

During the crisis period, German government bonds have experienced a so called "flight-to-quality/liquidity" effect with the bond prices rising sharply during this period. This effect implied an upward pressure in the Euro-area government debts, causing an increasing perturbation in the behavior of prices and yields for the more peripheral EMU countries. Beber et al. (2009) show that the bulk of sovereign yield spreads is explained by differences in credit quality, though liquidity plays a non-trivial role especially for low credit risk countries and during times of market uncertainty. Fig. 2 shows a clear negative correlation between the German 10-year yield and the VSTOXX index during the financial turmoil periods suggesting that investors have rebalanced their portfolios toward less risky and liquid German government bonds after the start of the financial market turmoil in summer 2007, causing a substantial increase in credit spread levels that clearly exceeded those observed in the early years of EMU.

[13] Barrios et al. (2009) use CDS' quotes as a measure of credit risk but argue that CDS' spreads may be an imperfect proxy because they are affected by liquidity.

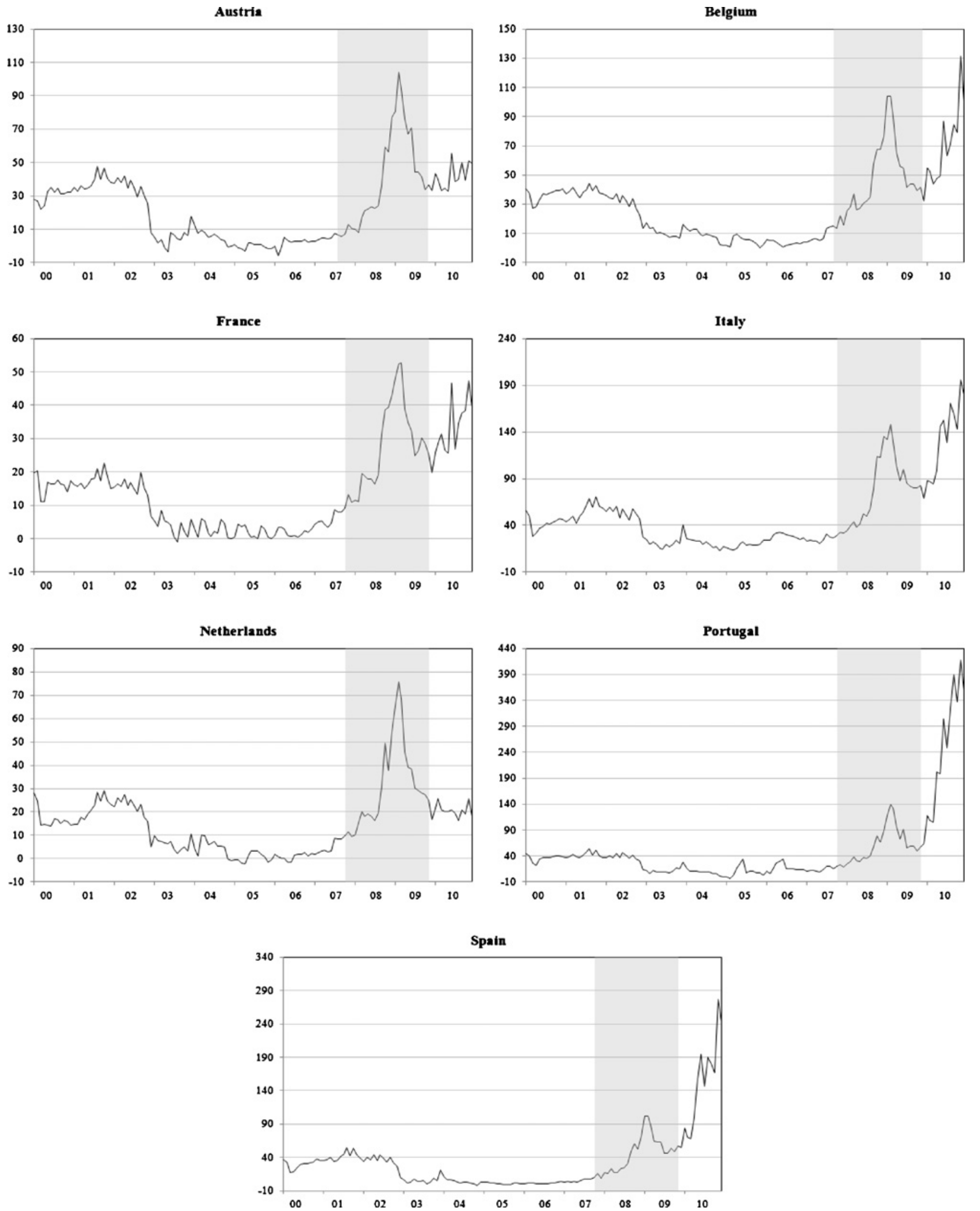

Fig. 1. Estimated 10-year spread relative to Germany. This figure plots the 10-year sovereign spread relative to Germany. The shaded area corresponds to the period between August 2007, the beginning of the financial crisis, and November 2009, when sharp increases in CDS spreads were observed in Greece, Portugal, and Spain, prenouncing the beginning of the EMU sovereign debt crisis-contagion.

5. Empirical analysis of the determinants of sovereign spread changes

Before running the linear regressions, we must be sure that all the variables included in the model are stationary. Hence, time series properties of the data were examined in order to test for the existence of unit roots using the sub-sampling version of the Levin et al. (2002) (LLC) and Im et al. (2003) (IPS)

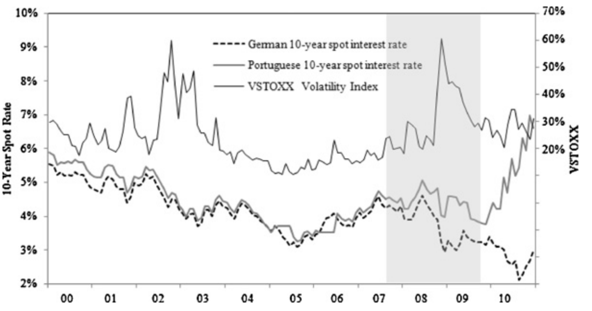

Fig. 2. German and Portuguese 10-year spot rates and European volatility. This figure plots the evolution of the German and the Portuguese 10-year spot rates relative to the VSTOXX Index. The shaded area corresponds to the period between August 2007, the beginning of the financial crisis, and November 2009, when sharp increases in CDS spreads were observed in Greece, Portugal, and Spain, prenouncing the beginning of the EMU sovereign debt crisis-contagion.

Table 4
Subsampling-based panel unit root tests.

Variable	LLC		IPS	
	Test statistic	Critical value (5%)	Test statistic	Critical value (5%)
SPRD 5-year	−9.58	−9.41*	−3.96	−4.47
SPRD 10-year	−10.22	−12.58	−3.40	−3.30*
SPRD 15-year	−8.16	−10.47	3.01	3.18
STOXX50	−6.40	−20.39	−2.33	−7.04
VOL	−8.06	−13.23	−2.95	−4.64
YLD	−6.58	−8.94	−2.41	−3.14
SLP	−5.03	−12.85	−1.84	−4.39
CURV	−9.11	−13.98	−3.34	−4.87
LIQ	−4.15	−6.20	−2.81	−3.39
DEBT	−4.51	−11.14	−2.31	−3.47
GCONS	−5.29	−5.55	−1.93	−2.19
GINV	−0.65	−4.45	−7.01	−8.96
SEXP	−4.27	−9.84	−1.80	−4.05
GROW	−15.95	−16.90	−5.78	−6.68
TAX	−6.58	−6.75	−2.21	−2.63
HICP	−3.62	−14.70	−1.33	−4.83
DEF	−5.29	−5.55	−1.93	−2.19

This table displays the test values and the sub-sampling-based critical values of the Levin et al. (2002) (LLC) and Im et al. (2003) (IPS) panel unit root tests when applied to the levels of variables. The AR order is chosen by the AIC, and data includes 44 quarterly observations from January 2000 to December 2010. In applying the minimum volatility rule, the minimum block size is $T^{0.45}$ and the maximum block size is $T^{0.85}$, where $T = 44$. Critical values marked with "*", mean that the null hypothesis brings rejection. SPRD 5-year is the 5-year maturity sovereign spread; SPRD 10-year is the 10-year maturity sovereign spread; SPRD 15-year is the 15-year maturity sovereign spread; STOXX50 is the DJ Euro Stoxx50 equity index; HICP is the Harmonized Index on Consumer Prices. Remain variables follow Table 5 description.

panel unit root tests to account for cross-sectional dependence in the data—for details see Choi and Chue (2007). The results are reported in Table 4. The unit root null hypothesis cannot be rejected for most of the variables no matter the test adopted (the only exceptions are the 5 and 10-year spreads for LLC and IMP tests, respectively). Therefore, as the estimation requires stationarity, all the variables are entered in first differences (indicated by Δ).

Table 5 presents the description of all the explanatory variables that we use in our model's estimations. Panel A presents the common explanatory variables, while Panel B presents the country-specific explanatory variables.

Armed with the sovereign spreads estimated in Section 4, and the data collected on the theoretical determinants of the sovereign spreads presented in Section 2, we can now quantify the impact that each factor has on the term structure of sovereign spreads for the seven EMU-countries under analysis.

Table 5
Description of the explanatory variables included in the regressions.

Panel A: common variables	
RET	Quarterly returns based on the EURO STOXX50 stock index
ΔVOL	Quarterly changes in implied volatilities of EURO STOXX 50 options traded at the Eurex and reflected by the Dow Jones EURO STOXX 50 Volatility Index (VSTOXX). The volatility is calculated as the average of the puts' and the calls' implied volatilities at a fixed time to maturity of 30 days
ΔYLD	Quarterly changes in the zero-coupon rate of the 10-year German government bond
ΔSLP	Quarterly changes in the slope of the default-free term structure defined by the German yield curve. The slope is defined as the difference between the 15-year and the 1-year spot rates
$\Delta CURV$	Quarterly changes in the curvature of the default-free term structure defined by the German yield curve. The curvature factor is defined as the difference between the 3.5-year spot rate and a synthetic spot rate for a maturity of 3.5 years interpolated between the 15-year and the 1-year spot rates
Panel B: country-specific variables	
ΔLIQ_j	Quarterly changes in the Weighted Mean Absolute Pricing Error (*WMAPE*). This statistic represents a proxy for the liquidity risk measured along the yield curve for each country
$\Delta DEBT_j$	Quarterly changes in the government debt, excluding debt interest payments, relative to GDP
$\Delta GCONS_j$	Quarterly changes in the Government intermediate consumption relative to GDP
$\Delta GINV_j$	Quarterly changes in the Government capital investment level relative to GDP
$\Delta SEXP_j$	Quarterly changes in the Government social expenditures and subsidies relative to GDP
GROWj	Quarterly changes of the industrial production index
ΔTAX_j	Quarterly changes of total Government receipts relative to GDP
$INFL_j$	Quarterly changes of the Harmonized Index of Consumer Prices (HICP)—all items
ΔDEF_j	Quarterly changes in the current account deficit relative to GDP

Panel A of this table presents the common explanatory variables and Panel B presents the country-specific explanatory variables in the analysis. The subscript *j* refers to a country-specific explanatory variable (*j* = Austria, Belgium, France, Italy, the Netherlands, Portugal and Spain). The data for EURO STOXX50 and VSTOXX were obtained from Datastream. Government Euro-denominated bond prices for each country were obtained from Bloomberg. The remaining data were obtained from Eurostat.

5.1. Model specification and general estimation results

Our sample data can be considered a balanced panel with 7 EMU-countries and 44 quarterly observations. We expect strong correlation in the dependent variable observations for each country and the usual consequences for the consistency of the OLS estimators. Thus, before estimating the regression models, we test first which one of the methodologies—pooled or panel data (random or fixed effects) regression—is more appropriate to describe the relationship between the changes in credit spreads and the explanatory variables included in the regressions.

To determine the appropriate methodology for each of the regression models, we perform two statistical tests: the *F*-test and the Breusch–Pagan LM test. In the *F*-test the pooled regression (in the null) is tested against the fixed effects regression model. The Breusch–Pagan LM test can be used to test the pooled regression (in the null) against the alternative random effects model. If the null hypothesis is rejected, the random effects regression is more appropriate when compared to the pooled regression and the Hausman test can be used to test the random effects model (in the null) against the alternative fixed effects model—see Baltagi (2001). The three bottom line sections of Table 6 present the results of both the *F*-test (pooled vs. fixed effects), Breusch–Pagan (pooled vs. random effects) test, and Hausman (random vs. fixed effects) test, which were computed through the software STATA 9.0.

For the pre-crisis period and the 15-year spread, we cannot reject the null hypothesis in both *F* and Breusch–Pagan tests (for a 5% significance level). Hence, the fixed and random effects models were discarded in favor of the pooled regression for all estimated equations. The basic model of sovereign credit spreads among the seven EMU-countries is estimated for a balanced (without country or time effects) panel dataset. All further regressions use fixed effects since the pooled and random effects regressions were reject (for a significance of 5%) by using the *F*, Breusch–Pagan and Hausman tests, respectively (as the null is always rejected).

The default risk measure is the change in spread of each country for three maturity buckets that constitute the dependent variable. The above listed determinants are the explanatory variables. Assuming

Table 6
Pooled regressions estimations on a panel data of the seven EMU-countries.

	Regression (1)		Regression (2)		Regression (3)		Regression (4)	
	Before crisis	During crisis	Before crisis	During crisis	Before crisis	During crisis	Before crisis	During crisis
Panel A: dependent variable is the 5-year maturity sovereign spread								
Intercept	1.448	3.075***	0.408***	8.415***	1.981	7.202**	1.028**	3.501***
RET (−1)	−0.095***	−0.125	−0.095***	−0.311			−0.097***	−0.034*
Δ*VOL*	0.122	2.02***	0.323	1.467***			0.123	1.857***
Δ*YLD*	−0.653	−0.897	−0.535	−1.888			−0.600	−5.763
Δ*YLD* (−1)	−0.359***	−2.746*	−0.319***	−1.823*			−0.381***	−4.822***
Δ*SLP*	−0.027**	0.171**	−0.027***	0.117*			0.025***	0.086**
Δ*SLP* (−1)	−0.017**	0.170**	−0.018***	−0.255***			−0.02***	0.287***
Δ*CURV*	0.003	0.147	0.001	0.014			0.009	0.159
Δ*CURV* (−1)	0.068***	0.184	0.071***	0.466*			0.077***	0.739***
Δ*LIQ*	0.064	−0.876	0.023	−0.83	0.097	−0.874	0.032	−0.71
Δ*DEBT* (−1)	0.063*	1.22**			0.213	0.817**	0.044	1.112**
Δ*GCONS* (−1)	0.018*	0.449*			0.034	0.336*		
Δ*GINV* (−1)	−0.005*	−1.733**			−0.029	−1.309***	−0.049*	−1.793**
Δ*SEXP* (−1)	0.003	0.195			0.003	0.021		
GROW (−1)	−0.010	−3.23**			−0.082	−1.651***	−0.015	−3.934**
Δ*TAX* (−1)	−0.791	−2.231*			−0.192	−2.503		
INFL	0.466	0.816*			0.761	0.933***	0.402	2.05**
Δ*DEF* (−1)	0.913*	2.259***			0.427	1.333***	0.818*	2.294***
R-squared	0.548	0.602	0.539	0.489	0.133	0.491	0.510	0.571
Adjusted *R*-squared	0.457	0.465	0.447	0.423	0.071	0.397	0.459	0.464
No. observations	216	91	216	91	216	91	216	91
S.E. of regression	6.29	8.88	6.29	9.17	6.76	13.32	6.28	11.67
F-statistic	2.74	3.79	4.40	3.52	0.94	4.38	3.74	4.94
Prob (*F*-statistic)	(0.000)	(0.000)	(0.000)	(0.001)	(0.485)	(0.000)	(0.000)	(0.000)
Durbin–Watson	1.75	1.79	1.85	1.98	1.84	1.92	1.71	1.77
F-test (pooled vs. fixed effects)	3.66	10.61	4.34	7.22	3.13	10.02	3.93	10.82
(*p*-Value)	(0.002)	(0.000)	(0.000)	(0.000)	(0.006)	(0.000)	(0.001)	(0.000)
Breush–Pagan-test (pooled vs. random effects)	13.57	44.54	23.4	53.24	7.92	55.26	17.35	50.33
(*p*-Value)	(0.000)	(0.000)	(0.000)	(0.000)	(0.005)	(0.000)	(0.000)	(0.000)
Hausman-test (random vs. fixed effects)	47.26	28.21	36.29	28.48	35.29	13.49	39.45	27.14
(*p*-Value)	(0.000)	(0.005)	(0.000)	(0.000)	(0.000)	(0.009)	(0.000)	(0.007)

Panel B: dependent variable is the 10-year maturity sovereign spread

Intercept	1.605	4.207***	1.169***	8.182 ***	1.141*	8.031***	1.86***	5.023***
RET (−1)	−0.041**	−0.319**	−0.128***	−0.526**			−0.03**	−0.439**
Δ*VOL*	0.077*	1.554***	0.092	1.043***			0.095	1.423***
Δ*YLD*	−1.034***	−2.935***	−0.918***	−2.269**			−0.924***	−1.345***
Δ*YLD* (−1)	−0.579***	−5.446***	−0.634***	−3.406***			−0.669***	−6.054***
Δ*SLP*	0.023	0.047	0.020*	0.142**			0.027**	−0.037**
Δ*SLP* (−1)	−0.038***	0.180*	−0.037***	0.197***			−0.041***	0.268**
Δ*CURV*	0.163**	0.452*	0.163***	0.036*			0.171***	0.238**
Δ*CURV* (−1)	0.125***	0.828**	0.019***	0.051**			0.019***	1.045**
Δ*LIQ*	0.565*	−0.564	0.320*	−0.547	0.736	−0.825	0.468*	−0.864
Δ*DEBT* (−1)	0.298*	0.434**			0.389	1.625***	0.216*	0.436***
Δ*GCONS* (−1)	0.021*	0.324*			0.026	0.123**		
Δ*GINV* (−1)	−0.008	−1.442***			−0.044	−1.263***	−0.004	−1.507***
Δ*SEXP* (−1)	0.001	0.164*			0.005	0.035**		
GROW (−1)	−0.074	−2.102**			−0.184	−0.925***	−0.042	−2.76***
Δ*TAX* (−1)	0.32*	1.713			0.314*	2.231		
INFL	0.232**	0.527**			0.236	0.734***	0.205	0.644***
Δ*DEF* (−1)	0.079*	1.976***			0.189	1.139***	0.252	2.028***
R-squared	0.526	0.617	0.442	0.539	0.277	0.562	0.551	0.597
Adjusted *R*-squared	0.430	0.485	0.392	0.447	0.225	0.464	0.484	0.496
No. observations	216	91	216	91	216	91	216	91
S.E. of regression	8.84	8.34	8.940	9.31	9.72	11.22	8.95	10.60
F-statistic	3.91	3.33	6.13	3.27	1.10	3.83	4.45	4.38
Prob (*F*-statistic)	(0.000)	(0.000)	(0.000)	(0.002)	(0.368)	(0.001)	(0.000)	(0.000)
Durbin–Watson	2.25	2.25	2.24	1.91	2.36	2.48	2.21	2.24
F-test (pooled vs. fixed effects)	13.56	12.27	14.46	9.6	11.65	12.16	13.46	12.93
(*p*-Value)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Breush–Pagan-test (pooled vs. random effects)	199.90	53.37	248.90	81.30	149.94	71.28	205.91	62.40
(*p*-Value)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Hausman-test (random vs. fixed effects)	28.69	27.15	29.31	30.00	13.94	21.28	20.99	21.85
(*p*-Value)	(0.0044)	(0.0074)	(0.0006)	(0.0004)	(0.0075)	(0.0003)	(0.0072)	(0.0052)

Table 6 (*Continued*)

	Regression (1)		Regression (2)		Regression (3)		Regression (4)	
	Before crisis	During crisis	Before crisis	During crisis	Before crisis	During crisis	Before crisis	During crisis
Panel C: dependent variable is the 15-year maturity sovereign spread								
Intercept	1.084	3.883**	1.044	8.052***	1.038	7.498***	1.162	4.322***
RET (−1)	−0.251**	−0.649*	−0.265***	−0.361**			−0.269***	−0.49**
Δ*VOL*	0.479*	1.875***	0.491	1.143***			0.505	1.617***
Δ*YLD*	−0.617**	−2.21**	−0.643**	−3.485***			−0.725**	−1.824**
Δ*YLD* (−1)	−2.195***	−7.086**	−2.256***	−5.393***			−2.377***	−8.389**
Δ*SLP*	0.189**	0.021	0.185***	0.089*			0.191***	0.127*
Δ*SLP* (−1)	−0.110**	−0.431**	−0.107***	−0.104*			−0.116***	−0.551**
Δ*CURV*	0.575**	0.838**	0.577***	0.520*			0.584***	0.665**
Δ*CURV* (−1)	0.129**	1.365*	0.156***	0.438**			0.163***	1.77**
Δ*LIQ*	0.967*	−0.418	0.763*	−0.326	2.36*	−0.732	1.081	−0.561
Δ*DEBT* (−1)	0.534	1.88**			0.749	3.87***	0.441	0.966***
Δ*GCONS* (−1)	0.078*	0.625***			0.724	0.320*		
Δ*GINV* (−1)	−0.0417	−2.170***			−0.079	−1.725***	−0.044	−2.180***
Δ*SEXP* (−1)	0.0005	0.209*			0.011	0.185**		
GROW (−1)	−0.579*	−3.239***			−0.972	−1.729***	−0.602*	−3.873***
Δ*TAX* (−1)	0.589*	2.186*			0.815*	2.698**		
INFL	0.319*	0.452*			0.355*	0.804***	0.308	0.667***
Δ*DEF* (−1)	0.190*	3.084***			0.077	2.821***	0.300	3.093***
R-squared	0.581	0.619	0.555	0.481	0.045	0.589	0.560	0.683
Adjusted *R*-squared	0.525	0.569	0.488	0.361	0.023	0.476	0.506	0.597
No. observations	216	91	216	91	216	91	216	91
S.E. of regression	8.11	11.43	8.07	11.93	12.82	13.79	8.100	11.330
F-statistic	3.03	3.14	5.20	3.65	1.02	4.35	3.88	4.03
Prob (*F*-statistic)	(0.000)	(0.000)	(0.000)	(0.001)	(0.424)	(0.000)	(0.000)	(0.000)
Durbin–Watson	2.38	2.11	2.30	1.81	2.44	2.29	2.27	2.09
F-test (pooled vs. fixed effects)	0.25	6.54	0.24	3.18	0.31	5.85	0.21	6.58
(*p*-Value)	(0.959)	(0.000)	(0.962)	(0.008)	(0.931)	(0.000)	(0.974)	(0.000)
Breush–Pagan-test (pooled vs. random effects)	2.22	20.68	2.20	10.04	1.95	22.24	2.41	22.43
(*p*-Value)	(0.136)	(0.000)	(0.138)	(0.002)	(0.162)	(0.000)	(0.121)	(0.000)
Hausman-test (random vs. fixed effects)	n.a.	37.71	n.a.	22.29	n.a.	30.43	n.a.	27.43
(*p*-Value)	n.a.	(0.003)	n.a.	(0.008)	n.a.	(0.000)	n.a.	(0.007)

Panels A, B, and C of this table present the estimated coefficients and their significance level (*10%; **5%; ***1%) for the regression models tested, using, as dependent variable, the sovereign spread (with respect to the German spot yield curve) for the 5-year, 10-year and 15-year maturity buckets, respectively, presented in Table 3. Explanatory variables with a (−1) are lagged one period and follow Table 5 description. Models marked with a "n.a." indicate that the respective test is not applicable.

that the benchmark German government spot yield curve is default free, the change of the sovereign spread in country j at time-t may be written as

$$\Delta CR_{jt} = \alpha + \beta X_{jt} + \epsilon_{jt}, \tag{8}$$

where ΔCR_{jt} is the change in sovereign spread, X_{jt} is a k-vector of explanatory variables, and $\beta = (\beta_1, \ldots, \beta_k)'$ are the coefficients for the explanatory variables. The constant (α) in the model represents the EMU common market-wide level of default risk. Further, it is assumed that the explanatory variables influence the country default risk measure in a similar way, and therefore the reaction coefficients are the same for all countries. Finally, ϵ_{jt} are the error terms of the $j = 1, 2, \ldots, M$ countries for dated periods $t = 1, 2, \ldots, T$.

Table 6 summarizes the results generated by the pooled or the fixed effects regressions when applied to our balanced panel data of seven EMU-countries during the before crisis period, from the first quarter of 2000 to the third quarter of 2007, and the during crisis period, from the fourth quarter of 2007 to the fourth quarter of 2010.[14] Panels A, B, and C contain the estimation results for the maturity buckets of 5-, 10- and 15-year sovereign spreads, respectively.

We estimate four regressions for each period and maturity bucket. The first regression—labeled as regression (1)—includes as independent variables all the explanatory factors described above. Regression (2) only uses as independent variables the common variables and the liquidity measure. Regression (3) considers as explanatory factors all the fundamental country-specific variables. Regression (4) is a somewhat "parsimonious model" which only takes into account the most significant variables. Hence, the selected fourth model corresponds to the first regression excluding changes in government consumption ($\Delta GCONS$), changes in government social expenditures ($\Delta GSEXP$), and changes in the level of taxation (ΔTAX). Since the residuals of the regressions show evidence of heteroskedasticity, all the standard errors and associated t-tests are computed using the White's heteroskedasticity-consistent procedure.

In line with Kwan (1996), Collin-Dufresne et al. (2001), and Van Landschoot (2004a), we consider the stock equity returns (RET) lagged by one-quarter. Additionally, and to represent the state of the corporate bond market, we include the changes in interest rate sensitive variables (ΔYLD, ΔSLP, and $\Delta CURV$), contemporaneously and lagged by one-quarter, in the first, second, and fourth regressions. As Collin-Dufresne et al. (2001, p. 2196) argue, "if there is mean-reverting behavior in spot rates, leverage, volatility, or credit spreads, then the beginning-of-month levels of those variables should contain information about the current month's change in credit spreads." Since some country-specific variables are available only on a quarterly basis and during the following quarter, the market participants only perceive the changes in these variables after two or three months. To control for this lagged frequencies, we include the country-specific variables—$\Delta DEBT$, $\Delta INTP$, $\Delta GCONS$, $\Delta GINV$, $\Delta GSEXP$, ΔTAX, $GROW$, and ΔDEF—lagged by one-quarter in the respective regressions.[15] The regressions are tested for jointly insignificant coefficients (F global-test) and, except for the third regression run for the before crisis period, the null hypothesis is strongly rejected for all maturity buckets and regressions.

Regressions (1), (2) and (4) show that the loadings of the explanatory variables are generally correct and provide a reasonably good explanation for government default risk innovations, since they present an adjusted R-squared between 35% and 60% of the variance of the dependent variable,[16] depending on the model and the length of the maturity bucket. Regression (3) extends the empirical specification by considering exclusively the role played by the country-specific variables, but this regression is not statistically significant for all the performed estimations during the pre-crisis period.

[14] Because macroeconomic data is available from Eurostat on a quarterly basis, we choose to run the regressions in a quarterly frequency, avoiding the mismatch of observations. This procedure favours the measurement of the fundamental risk, which is better measured at lower frequencies, but has the drawback of a less effective perception of the liquidity risk, which is best captured with relatively high-frequency data.

[15] We run the regressions with the contemporaneous country-specific observations and the significance of the estimated coefficients shrank considerably or even disappeared.

[16] This global explanatory power is satisfactory given that the regressions are run in differences and at low frequencies.

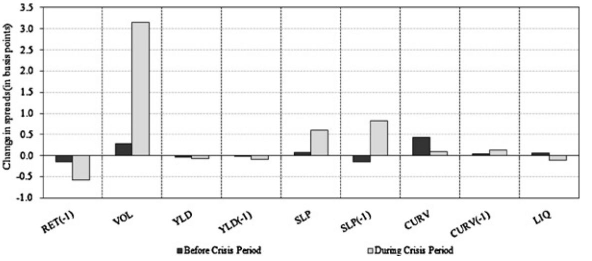

Fig. 3. Credit spreads sensitivity under the second regression model. This figure shows the impact on the 10-year credit spread of a one standard deviation change in the second regression model determinant factors, before and during the crisis periods.

In general, the statistical significance of the coefficients increases with the maturity of the credit spread buckets, and nearly all variables have the expected sign, particularly those that are more significant. The Durbin–Watson test shows no errors' serial correlation.

5.2. Evidence from the before crisis period

As Table 6 shows, during the pre-crisis period the estimates for the coefficients of the common explanatory variables generally present high significance and have the expected signs, particularly for the longer maturity buckets. Inversely, the country-specific variables show a weaker explanatory power during this period.

The equity-related variables allow us to evaluate whether the predictions of the Merton (structural) model are validated in the European sovereign debt market. The return on the equity stock index has a clear influence on sovereign default risk and could proxy the general economic sentiment of the agents. Table 6 shows that the lagged equity stock returns (*RET*) have a strong and significant effect on credit spreads for all regression models and maturity spreads. Fig. 3 shows that the impact that a one standard deviation change in lagged stock returns has on credit spreads is quite small. The negative signs mean that, on average, for higher stock prices, the required risk premium falls because the firms financial strength improves, and the capacity to serve their debts is higher.

The second equity-related factor is the equity volatility. We observe that the estimated coefficients for changes in our proxy for future equity volatility (ΔVOL) is significant at the 10% level only in the first regression and for the 10- and 15-year maturity spreads. This could suggest that the European financial risk was not-priced during the pre-crisis period. According to Fig. 3, and as with equity stock returns, the impact of a one standard deviation change in volatility is small. The positive signs of the coefficients are always correct but the corresponding *p*-values show weak evidence of their statistical significance.

In the literature on the determinants of the yields on government bonds, the level, slope and curvature of the yield curve are the three most frequently documented factors. Our results are generally in line with those presented by Bliss (1997), Collin-Dufresne et al. (2001), Boss and Scheicher (2002) and Schuknecht et al. (2010). In general, the interest rate sensitivities, namely the lagged variables, are highly significant for all maturities and regression models and show the expected sign, being in accordance with economic and financial theory, which suggests the existence of strong similarities between credit markets and government debt markets.

As shown in Table 6, and contrary to the results pointed out in the literature, our estimates cannot support the evidence that changes in spreads between EMU issuers are largely explained by changes in the liquidity factor (ΔLIQ). Actually, the liquidity coefficients are significantly different from zero at a 10% level and the signs are correct[17] in the first, second and fourth regressions and for the 10- and

[17] Confirming that an increase in liquidity, i.e., a decrease in our liquidity proxy (*WMAPE*), causes a decrease in the credit spread due to an increase in the liquidity premium.

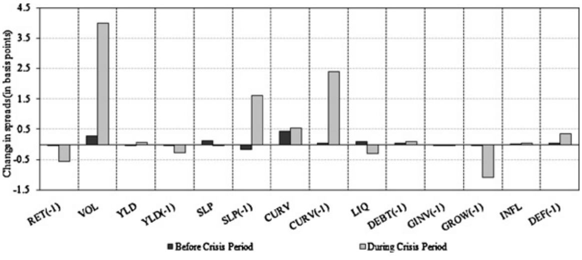

Fig. 4. Credit spreads sensitivity under the fourth regression model. This figure shows the impact on the 10-year credit spread of a one standard deviation change in the fourth regression model determinant factors, before and during the crisis periods.

15-year maturities. Similar evidence in terms of limited explanatory power of the liquidity factor is provided by Geyer et al. (2004), Favero et al. (2010) and Arghyrou and Kontonikas (2011).

Although correct signs are generally present in all regressions and spreads, the coefficients of the debt to GDP ($\Delta DEBT$) and the current account deficit (ΔDEF) have, at best, a minor role in the behavior of credit spread's differences. The deterioration of the macroeconomic fundamentals, already observed over the period from 2000 until middle 2007, seems not to be accounted by the market participants. Arghyrou and Kontonikas (2011) argue that, in this period, EMU sovereign bond markets were operating under an implicit guarantee that there was very little default risk. The $\Delta DEBT$ coefficient is statistically different from zero at a 10% level in the first regression for the 5-year maturity, and in the first and fourth regressions for the 10-year maturity. The ΔDEF coefficient is statistically significant at a 10% level in the first regression for all maturities, and in the fourth regression only for the 5-year credit spread.

The coefficients of the expenditure side of the government budget ($\Delta GCONS$, $\Delta GINV$, and $\Delta GSEXP$) show also a minor significance in explaining the 5-year spread. Only the first regression shows a 10% significance for the $\Delta GCONS$ in all maturities, and $\Delta GINV$ is only statistically significant (at a 10% level) for the 5-year credit spread in the first and in the fourth specifications. These results are not in accordance with those presented in Alesina et al. (1992), Lemmen and Goodhart (1999) and Van Landschoot (2004b), who show that these variables significantly influence sovereign default risk. By contrast, our results are in-line with those reported by Faini (2006) for the period 1979–2002, who also finds no significant effect for the government consumption neither at the country nor at the EMU levels. The same low explanatory power is observed for the loadings of the changes in the level of taxation (ΔTAX) and in the percentage change in industrial production (*GROW*). The effect of the rate of inflation (*INFL*) is statistically different from zero at a 5% level in the first regression for the 10-year bucket, and at a 10% level in the first and third regressions for the 15-year maturity bucket.

5.3. *Evidence from the crisis period*

As Table 6 reports, in the crisis period the equity index return (*RET*) presents the correct sign and remains highly significant for the 10- and 15-year maturity spreads, indicating a highly persistent influence in the long term credit spreads. Except for the first regression, where the *RET* loading is significant at a 10% level, the equity index returns coefficient stands significantly different from zero at a 5% level in all regressions. For the 5-year maturity, and even though the coefficients are correctly signed, only in the fourth regression the coefficient is statistically different from zero at a 10% significance level. This result suggests that, during the financial crisis, the equity returns have a rather limited influence on the short term sovereign bond markets evolution.

The estimated volatility coefficient is positive and strongly significant for all maturities and specifications. European risk aversion seems to be the largest contributor to the widening of spreads after July 2007—see Figs. 3 and 4. In fact, the crisis period has been characterized by the investors' increased uncertainty and our proxy for the international risk aversion seems to capture this effect quite well, confirming the important role played in the linkage between credit spreads and financial risk. Moreover, and as Fig. 2 documents, after the onset of the financial crisis, investors' have seek the high credit quality of German bunds as a safe-haven investment, causing the increasing of spreads for other countries. The 10-year yield on Portuguese government debt has been one of the most affected by the increased international risk perception. Similarly, all the other high indebted EMU-countries have experimented this same effect during the crisis period.

Before the crisis, interest rate factors appear to have played an important role in explaining the credit spreads dynamics. After July 2007, however, this role becomes somewhat less significant. The estimated coefficients have generally the expected sign and are often significant but at a lower level. For the longer maturities the influence of the three interest rate factors is stronger than in the short term, and the lagged terms provide a more important contribution to explain the credit spreads evolution. Nevertheless, the results confirm the basic picture from the pre-crisis period: German interest rates' behavior has strongly conditioned the creditworthiness of the Euro-area government financial sector.

The liquidity variable (ΔLIQ) remains unsuccessful in explaining sovereign yield differentials. Furthermore, during the crisis period liquidity presents the wrong sign for all maturity buckets and regressions. This somewhat controversial result could have been influenced by the lack of sample observations since August 2010 and/or by the low-frequency used in the estimated regressions. These results are in line with those of Geyer et al. (2004), Pagano and von Thadden (2004), Favero et al. (2010), Arghyrou and Kontonikas (2011) who also conclude that liquidity plays a smaller role in explaining yield differentials. However, our estimates are clearly in contrast with the findings of Beber et al. (2009)—that postulate that, while credit risk matters for bond valuation in normal times, liquidity becomes more important in times of financial distress—of Barrios et al. (2009)—who find a large explanatory power for the liquidity factor during the period August 2007–April 2009 in France, Greece, and Italy—and of Attinasi et al. (2009)—who report, during the period between July 2007 and March 2009, a strong statistical significance for the liquidity risk.

The importance of the government default risk variables is found to increase significantly in the crisis period. Changes in Debt to GDP ratio ($\Delta DEBT$) show a high significance in all regressions and maturities, suggesting that a higher government debt level is strongly associated with higher government bond yield spreads for the Euro-zone countries analyzed in our sample. Accordingly, increasing current account deficits (ΔDEF), reflecting a deterioration in international competitiveness and a higher future fiscal burden, strongly widen the sovereign credit risk and, hence, the credit spread premium. In-line with previous studies—Goldstein and Woglom (1991), Alesina et al. (1992), Lemmen and Goodhart (1999), Van Landschoot (2004b), Bernoth et al. (2004), Faini (2006) and Attinasi et al. (2009)—our estimates for the current account deficit coefficient are statistically significant at a 1% level for all specifications and maturities. Thus, in the crisis period our estimates are clearly in accordance with the above postulated influence in credit default risk of the level of public debt and current account balances, being consistent with the notion that the credit market monitors fiscal performance and exerts disciplinary pressure on governments. Furthermore, these findings support the hypothesis that, prior to the global financial crisis, spreads were not linked to macroeconomic risk factors due to the fact that market participants did not react to the prevailing effect of domestic fundamentals deterioration. In fact, countries such as Portugal, Greece and Spain with large current account deficits in the pre-crisis period and large increases in government debts after the beginning of the financial crisis, have supported substantial increases on their bond yield spreads only after the onset of the credit crunch.[18] Moreover, the downgrade of credit ratings of weaker issuers had a significant impact on institutional investors asset allocation decisions, deteriorating even more the sustainability of these countries' public finances. Fig. 4 shows the impact on credit spreads in response to a one standard deviation change in the determinant factors presented in the fourth regression model.

The changes in government investments ($\Delta GINV$), business cycle proxy ($GROW$) and taxes (ΔTAX) confirm the importance of the macroeconomic fundamentals in the crisis period. The estimated coefficients of these explanatory variables become more robust and strongly significant, exerting an increased influence on the Euro-zone spreads. Moreover, the larger importance of these country-specific factors in explaining the credit spreads dynamics suggests that significant shifts have taken

[18] Barrios et al. (2009) document a snowball effect in the Euro-area: the increasing debt burden, arising from higher interest rates and/or low GDP growth, implies the worsening of Euro-zone governments' financing conditions.

place in the Euro-area government bond markets after the start of the financial crisis. This macroeconomic shift effect has been pointed out in the literature by, among other authors, Barrios et al. (2009), Attinasi et al. (2009), Arghyrou and Tsoukalas (2010), or Arghyrou and Kontonikas (2011). The rate of inflation (*INFL*) shows a positive and significant effect on sovereign credit spreads behavior in all regressions and particularly for the 10- and 15-year credit spreads. These results are in line with the findings of Cantor and Packer (1996) and confirm the theoretical assumption that an unexpected increase in inflation is indicative of structural problems in the government's finances, representing an additional channel which raises risk aversion and leads to a rise in default risk premiums. Moreover, Manasse et al. (2003) argue that episodes of high inflation rates or high inflation volatility imply a larger default risk and are associated with an increased probability of entering and remaining in distress.

5.4. Robustness and additional tested variables

Arghyrou and Kontonikas (2011) use the Chicago Board Options Exchange Volatility Index (*VIX*) as a general indicator of common international risk. Furthermore, Haugh et al. (2009) use US corporate bond spreads (instead of EMU corporate bond spreads) as a proxy for general risk aversion to reduce the likelihood that the risk variable is endogenous to the dependent variable (which corresponds to the differences in 10-year EMU government bond yields against Germany). To test whether the international risk factor could affect our results, we include the VIX volatility index for the 10-year spread in all specifications. As the correlation between the two indices is around 95%, the volatility coefficient remains positive and strongly significant after the crisis period. This result is also reported by Manganelli and Wolswijk (2009), and suggests that spreads in the Euro-area are determined not only by local variables, but also by international factors.

As the role of the liquidity variable is typically related to local market features and frictions, the absence of explanatory power may be due to the fact that liquidity is best captured in relatively high-frequency data. We have also run all regressions using data in a monthly basis, by linearly interpolating the macroeconomic variables. The results are broadly similar to those obtained using quarterly observations. Additionally, and since the literature does not provide a clear-cut evidence about the liquidity proxy factor in use, we reestimate the models using two other proxies for liquidity: the average bid-ask spread on government bond prices and, following Codogno et al. (2003), the trading volumes expressed in percentage of the total traded volumes. The liquidity factor proxied by the bid-ask spread yields the same sign and significance level as before but the sizes of the coefficients are now smaller. When we proxied the liquidity factor by the percentage of trading volumes, the volatility coefficient became statistically insignificant in all estimated models.

To test the predictions of the market-based discipline hypothesis and the non-linear relation between government debt levels and government bond yields, and similarly to Barrios et al. (2009), the squared of the debt to GDP ratio is added to the 10-year spread in all specifications. In order to avoid multicollinearity, the current account deficit was omitted. We found a nonsignificant (and negative for the 5-year maturity bucket) influence on the sovereign credit spreads for the 10- and 15-year maturities. This result is in line with those reported by Goldstein and Woglom (1991) as well as Lemmen and Goodhart (1999), who also found a nonsignificant coefficient for the squared debt term. However, Barrios et al. (2009) report a positive and statistically significant coefficient at 5% and 10% levels for the square debt to GDP ratio.

6. Conclusion

This paper analyzes whether the changes in European sovereign credit spreads are influenced by changes in financial and macroeconomic variables during the period between January 2000 and December 2010. This 11 years sample is splitted into two sub-samples covering the "before crisis period", ending in July 2007, and the "crisis period", from August 2007 to December 2010.

Using a database of European government bonds, we first estimate the term structure of credit spreads for different EMU-countries by applying a yield curve parameterization that is consistent with an HJM-Gaussian multifactor term structure model. Our data consists of a balanced panel dataset for seven EMU-countries—Austria, Belgium, France, Italy, the Netherlands, Portugal, and Spain. For all these seven EMU-countries, we examine the economic and statistical significance of a set of explanatory variables, and find that they are generally in accordance with economic and credit risk theories. For each regression model tested, we analyze the quarterly changes of three maturity buckets of credit spreads: 5-year, 10-year, and 15-year. Credit spreads are measured as the difference between the zero-coupon rates of government bonds from each country and the German Treasury spot rates, which we consider the EMU reference.

Our main findings can be summarized as follows. First and during the “before crisis period”, the stock returns and the interest rate sensitive variables—level, slope and curvature of the benchmark yield curve—are the most important determinants of credit spreads. Second, during this period, the country-specific macroeconomic fundamentals and the European risk conditions show, at most, a limited explanatory power. This finding is consistent with the so-called “convergence-trading hypothesis” according to which market participants were betting in a scenario of full real convergence to the German economy of all Euro-zone countries. In fact, it is commonly accepted that, until the beginning of the financial crisis and despite a generalized deterioration of Euro-macroeconomic fundamentals, markets were operating under a perceived implicit guarantee according to which there was very little default risk associated to investments in EMU sovereign bonds, and, therefore, there was simply no market discipline pressure on EMU governments to improve public finances.

Third, during the “crisis period” the market behavior shifted to a different regime that is strongly determined by the international volatility and country-specific macroeconomic fundamentals. The European volatility risk, the credit related macroeconomic variables—namely, public debt level, current account deficit, government investment, and the state of the business cycle—as well as the inflation rate now seem to play an important role in explaining the sovereign credit spreads dynamics among EMU-countries. More specifically, the role played by macroeconomic fundamentals (such as the level of government debt and the current account deficit) in explaining credit risk is shown to increase with the level of risk aversion. High debt countries and countries with large current account deficits are found to experience the highest credit spread increases. The government consumption, social expenditures, and taxes received yield coefficients with the correct signs but evidence a null or, at the best, minor ability to explain the changes in sovereign spreads.

Finally and somewhat controversially, we find an insignificant role for the liquidity risk factor in explaining the evolution of the credit spreads. In the pre-crisis period the liquidity coefficient is positive but not significantly different from zero, indicating that during this period spreads were not affected by liquidity considerations. During the crisis period, the liquidity coefficient remains statistically insignificant and negative for all credit spread maturities, indicating clearly that market participants continue to misprice the intra EMU liquidity risk.

Despite the estimation techniques and the low-frequency changes used, our regressions present an overall adjusted R-squared that reaches a maximum of 60%. Even though this is a relatively high portion of the credit spreads’ variance that is explained by our models, there still remains a relevant percentage of unexplained variance that can indicate the presence of another common factor that drives the joint variation of EMU spreads (particularly in the short-term maturity) and that is not included in our models’ specification. This result is also reported in previous studies, such as Collin-Dufresne et al. (2001), Boss and Scheicher (2002) or Geyer et al. (2004).

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