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Discussion Paper No. 289

November 2010

ISSN 1860 0921

Sovereign bond yield spreads: A time-varying coefficient approach*

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Abstract

We study the determinants of sovereign bond yield spreads across 10 EMU countries between Q1/1999 and Q1/2010. We apply a semiparametric time-varying coefficient model to identify, to what extent an observed change in the yield spread is due to a shift in macroeconomic fundamentals or due to altering risk pricing. We find that at the beginning of EMU, the government debt level and the general investors' risk aversion had a significant impact on interest differentials. In the subsequent years, however, financial markets paid less attention to the fiscal position of a country and the safe haven status of Germany diminished in importance. By the end of 2006, two years before the fall of Lehman Brothers, financial markets began to grant Germany safe haven status again. One year later, when financial turmoil began, the market reaction to fiscal loosening increased considerably. The altering in risk pricing over time period confirms the need of time-varying coefficient models in this context.

JEL Classification Codes: C14, E43, E62, G12, H62, H63

Keywords: sovereign bond spreads, fiscal policy, euro area, financial crisis, semiparametric time-varying coefficient model, nonparametric estimation

* The authors thank Konstantin Kholodilin, Vladimir Kuzin, John Lewis, Andreas Pick and Georg Stadtmann for helpful comments and suggestions.

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1 Introduction

After the start of the European Monetary Union (EMU), financial markets barely differentiated between sovereign borrowers. Sovereign bond yield spreads across EMU member states relative to Germany converged and were generally smaller than fifty basis points. However, with the 2007/2008 global financial crisis, government bond yield spreads began to increase considerably, reaching values around 250 basis points for Greece and Ireland in Q4/2008.

Analyzing the driving forces of sovereign yield differentials within the euro area is attracting a lot of interest in the literature. The general consensus is that bond yield differentials are significantly affected by both international and country-specific risk factors such as liquidity or default risk premia.¹ Recent evidence shows that the sharp increase of government bond yield spreads during the financial crisis can not purely be attributed to changes in macroeconomic fundamentals, but also to the fact that the general pricing of government credit risk has increased over time, in the sense that financial markets reacted more strongly to different risk variables than they did before. Thus, the relationship between the variables proxying default and liquidity risk and government bond yield spreads may be time-varying. Most studies analyzing the determinants of bond yield spreads rely on simple linear regression models, which assume a constant relationship between the explanatory variables and bond yield spreads. These linear models, however, are not an appropriate approach to accurately model these non-linear dynamics.

We contribute to the literature by estimating time-varying coefficients in an additive nonparametric fixed-effects panel model framework. Estimating time-varying coefficients allows us to identify to what extent an observed change in the yield spread is due to a shift in macroeconomic fundamentals such as a country's fiscal position and to what extent it reflects a change in markets' pricing of these fundamentals. Further, we are able to endogenously identify the timing and patterns of any changes in the pricing of the

¹Note, that exchange rate risk have been eliminated in EMU.

different risk components. In this form of semiparametric models, a separate nonparametric regression function is fitted to each explanatory variable. An appealing feature of this approach is that additivity of the individual predicting variables is the only assumption on the functional form of the model and hence no further assumptions about the specific functional form for the path of coefficients are imposed on the data. This is a major advantage compared to parametric approaches and is especially relevant for our data set, where the bond yield spreads show no clear convergence or divergence path over the entire time span of the data sample.

Our model is based on Sun et al. (2009), who develop a semiparametric fixed effects panel data model with varying coefficients using a local linear regression approach. Their methodology has the nice feature that the fixed effects are removed by applying a one-step estimation approach based on kernel weights without the need of back-fitting techniques. We adapt their model into a smooth time-varying coefficient model.

We find that the impact of fiscal policy variables and general investors' risk aversion on sovereign yield spreads is not constant over time, which confirms the need of time-varying coefficient models in this context. At the beginning of EMU in 1999, the debt level of a country and the general investors' risk aversion significantly explained interest differentials. In the subsequent years, however, the safe haven status of Germany diminished, while sovereign debt differentials continued to play an important role in explaining yield differentials. By the end of 2006, two years before the fall of Lehman Brothers, financial markets began to grant Germany a safe haven status again, which signals that financial markets started worrying about risk long before the start of the financial crisis. With the financial crisis, also the market reaction to fiscal loosening increased considerably, indicating that fiscal discipline imposed by financial markets has become stronger.

The rest of the paper is organized as follows. Section 2 gives an overview about the related literature. Section 3 discusses the methodology that we apply for our estimations. Section 4 details the data and presents some descriptive analysis. Section 5 reports the main results and Section 6 concludes.

2 Literature Review

Analyzing the determinants of sovereign yield spreads in the euro area is attracting a lot of interest in the literature. A number of studies find that part of the interest differentials across EMU countries are significantly affected by fiscal imbalances, which indicates that interest rates are subject to a default risk premium. Codogno et al. (2003) find in a sample of nine EMU countries that for Italy and Spain the fluctuations in yield differentials can be attributed to domestic fiscal fundamentals, while Heppke-Falk and Hübner (2004) find that expected deficits have a positive impact on yield spreads in Germany, France, and Italy. Hallerberg and Wolff (2008), Bernoth et al. (2006), and Gerlach et al. (2010) find that interest differentials among EU countries vary depending upon the debt and deficit level of the issuing country. A similar result is found by Gomez-Puig (2008), who shows that yield spreads respond positively to a rise in debt relative to Germany. Bernoth and Wolff (2008) focus on the accuracy of officially reported fiscal variables and find that fiscal transparency and budget deficit levels have a significant impact on yield spreads.

Several studies show that sovereign bond yield spreads are driven not just by country-specific risk factors but also international factors and global investors' risk aversion. Codogno et al. (2003), Geyer et al. (2004), Favero et al. (2010) and Pozzi and Wolswijk (2008) find that yield spreads across EMU countries are significantly affected by global risk factors. Similarly, Sgherri and Zoli (2009) and Manganelli and Wolswijk (2009) find that a substantial part of EMU yield spreads can be explained by a common international factor that reflects the investors' risk aversion. The explanation is that in times of uncertainty, investors become more risk averse and re-structure portfolios accordingly. This flight-to-safety motive favors bonds of countries that are generally regarded to have a low default risk.

Another potential determinant of yield differentials is a liquidity risk premium. It is important to extract the liquidity component from yield spreads, because it might signal to a lack of financial market integration rather than

discrepancies in fiscal positions. In theory, illiquidity is priced by financial markets owing to the trading costs it creates. The empirical evidence for the existence of a liquidity premium in bond yields, however, is mixed. Gomez-Puig (2006), Barrios et al. (2009) and Gerlach et al. (2010) confirm that a liquidity risk premium is a significant element of euro area bond yield spreads. Favero et al. (2010) show that liquidity risk is priced only in a subset of the euro area bond markets, while Beber et al. (2009) find evidence that liquidity matters only in times of heightened market stress. Codogno et al. (2003) conclude that liquidity differences appear to play at most a minor direct role. This result is confirmed by Pagano and von Thadden (2004), who add, however, that liquidity gains more significant role through the interaction with changes in fundamental risk. Geyer et al. (2004), and Bernoth et al. (2006) cannot find a significant liquidity effect on yield differentials across EMU countries.²

Several studies, using linear regression models, test for discrete coefficient shifts and find that the strength of market discipline varies over time. Bernoth et al. (2010) find that after the start of EMU the impact of debt and deficits on yield differentials, while remaining significant, weakened. After the intensification of the financial crisis in August 2008, however, financial markets began penalizing fiscal imbalances more than they did before and, at the same time, the impact of global investor risk aversion to yield spreads increased significantly. This result is confirmed by Sgherri and Zoli (2009). Barrios et al. (2009) add that the role of government debt on yield differentials became significantly more important with the greater level of general risk aversion observed during the global financial crisis. This finding is supported by Haugh et al. (2009) who show that the general increase in the risk aversion magnified the gravity of fiscal performance.

However, it might be more plausible to think of coefficients changing gradually over time, rather than having a discrete break-point between regimes. To our knowledge, only two papers follow this idea by applying a time-varying

²Before 1999, Bernoth et al. (2006) find that yield spreads are affected by liquidity premia. However, these liquidity premia have largely vanished with the start of EMU.

coefficient approach to analyze the dynamics of bond yield spreads within EMU. Both, Aßmann and Boysen-Hogrefe (2009) and Pozzi and Wolswijk (2008), estimate a state space model employing the Kalman filter. Aßmann and Boysen-Hogrefe (2009) focus on a data from January 2001 through March 2009. They find that between 2003 and 2007, the debt to GDP ratio is the most important variable explaining the sovereign bond spreads, while budget balance and liquidity are insignificant. During the financial turmoil, liquidity and both fiscal variables gained in importance. However, they do not control for the impact of general investors' risk aversion or global risk factors, which might bias their results. Pozzi and Wolswijk (2008) examine bond risk premia in five EMU countries over the period 1995 - 2006. They find that a common risk factor is always relevant for explaining bond risk premia, whereas country-specific factors were almost eliminated by the end of 2006 for all countries but Italy. Moreover, they show that country-specific exposures to the international risk factor decreased and converged during the observed time period. Both papers have the major caveat that in order to estimate the state space model, a specific process for the path of the model coefficients must be assumed. Aßmann and Boysen-Hogrefe (2009) use a random walk process, while Pozzi and Wolswijk (2008) implicitly assume convergence by introducing a convergence operator into the coefficient paths. The latter process seems not best suited to model the dynamics and observed divergence of yield spreads during the financial crisis. As stated by Cai (2007), a misspecification of the underlying coefficient function leads to serious bias in the estimation results.

3 Methodology

We estimate the time-varying determinants of EMU yield spreads by applying a semiparametric model in form of an additive nonparametric regression approach. In such semiparametric models, a separate nonparametric regression function is fitted to each explanatory variable. An appealing feature of this approach is that additivity of the individual predicting variables is the

only assumption on the functional form of the model and hence no further assumptions about the specific functional form for the path of coefficients are imposed on the data. A nonparametric estimator to model time-varying coefficients is initially proposed by Robinson (1989) and is further developed by Orbe et al. (2000; 2005; 2006), Park and Phillips (2001), Phillips (2001). The underlying idea is that each sequence of coefficients lies on a smooth function of the time index. The literature extending this methodology to panel data, however, is scarce.³

Our model is motivated by Sun et al. (2009), who develop a general varying coefficient panel data model with fixed effects using a local linear regression approach. Their methodology has the nice feature that the fixed effects are removed by applying a one-step estimation approach based on kernel weights without the need for back-fitting techniques. We modify their model in the sense that we introduce smoothly time varying coefficients. Let's assume that the sovereign bond yield spread is denoted by y_{it} and has the following functional form:

$$y_{it} = x'_{it}\beta_t + \mu_i + \nu_{it} \quad (1)$$

where $i = 1, \dots, N$; $t = 1, \dots, T$ and $x_{it} = (x_{it,1}, \dots, x_{it,k})'$ is a vector of explanatory variables of dimension k , which consists of variables measuring default, liquidity and global risk factors. $\beta_t = (\beta_{1t}, \dots, \beta_{kt})'$ are time-varying coefficients and μ_i are country specific fixed effects. The random errors, ν_{it} , are assumed to be i.i.d with zero mean and finite variance $\sigma_v^2 > 0$, which are independent of μ_j and x_{js} for all i, j, t and s . Rewriting equation (1) in matrix form yields:

$$Y = B \{X, \beta(t)\} + D\mu + V \quad (2)$$

³Hoover et al. (1998) develop nonparametric estimators for longitudinal data and propose a method for the selection of smoothing parameters and establish asymptotic properties. Wu et al. (1998) suggest a time varying coefficient estimator that minimizes a local least squares criterion and construct a class of approximate pointwise and simultaneous confidence regions for the coefficients.

with $Y = (Y'_1, \dots, Y'_N)'$ and $V = (\nu'_1, \dots, \nu'_N)'$ are $(NT) \times 1$ vectors; $Y'_i = (y_{i1}, \dots, y_{iT})$ and $\nu'_i = (\nu_{i1}, \dots, \nu_{iT})$. $B\{X, \beta(t)\}$ is an $(NT) \times 1$ vector which stacks all $x'_{it}\beta_t$ and $\mu = (\mu_2, \dots, \mu_N)'$ is an $(N-1) \times 1$ vector. For identification purpose, we impose similar to Su and Ullah (2006), $\sum_{i=1}^N \mu_i = 0$, such that $D = [-e_{N-1} \ I_{N-1}]' \otimes e_T$ is an $(NT) \times (N-1)$ matrix, where I_{N-1} denotes an identity matrix of dimension $N-1$ and e_{N-1} a $(N-1) \times 1$ vector with all elements being 1.

The basic idea of the local linear regression approach is to fit locally a straight line through the observations around a specific point in time. Thus, to estimate the slope coefficient at time t , we give those observations that lie in a close neighborhood around time t more weight than observations that are measured much earlier or later than at time t . This is done by introducing a Kernel function into the regression equation, which weights the observations according their distance to the specific point in time under consideration. Let $K_{h,i}(t, s)$ be a kernel for all i and t , which is defined by $K_{h,i}(t, s) = K((t-s)/h)$ with h being the bandwidth parameter and $s = 1, \dots, T$; and define a $(T \times T)$ diagonal matrix $K_{h,i}(t) = \text{diag}\{K_{h,i}(t, 1), \dots, K_{h,i}(t, T)\}$ for each i and $(NT) \times (NT)$ diagonal matrix $W_h(t) = \text{diag}\{K_{h,1}(t), \dots, K_{h,N}(t)\}$. The local weight matrix $W_h(t)$ ensures the locality of our nonparametric fitting. We then solve the following optimization problem:

$$\min_{\beta_t, \mu} [Y - B\{X, \beta(t)\} - D\mu]' W_h(t) [Y - B\{X, \beta(t)\} - D\mu] \quad (3)$$

Solving equation (3) for $\hat{\mu}$ and replacing μ with its estimator, the optimization problem modifies to:

$$\min_{\beta_t} [Y - B\{X, \beta(t)\}]' S_h(t) [Y - B\{X, \beta(t)\}] \quad (4)$$

where $S_h(t) = M_h(t)' W_h(t) M_h(t)$ and $M_h(t) = I_{NT} - D \{D' W_h(t) D\}^{-1} D' W_h(t)$. Note that the fixed effects are removed in model (4) since $M_h(t) D \mu \equiv 0_{NT \times 1}$

for all t . This optimization problem provides:

$$\hat{\beta}_t = \left[\sum_{s=1}^T \sum_{i=1}^N S_{h,i}(t, s) x_{is} x'_{is} \right]^{-1} \sum_{s=1}^T \sum_{i=1}^N S_{h,i}(t, s) x_{is} y_{is}, \quad (5)$$

We calculate the confidence intervals of the estimated coefficients pointwise by:

$$CI = \left[\hat{\beta}_{kt} \pm z_{1-\frac{\alpha}{2}} \hat{\sigma}_t m_{kk}^{-1} \right], \quad (6)$$

where

$$\hat{\sigma}_t^2 = \frac{1}{N} \sum_{s=1}^T \sum_{i=1}^N \frac{K_{h,i}(t, s)}{\hat{f}_h(t)} (y_{is} - \hat{y}_{is})^2, \quad (7)$$

and

$$m^2 = \sum_{s=1}^T \sum_{i=1}^N K_{h,i}(t, s) x_{is} x'_{is}, \quad (8)$$

where \hat{y}_{is} denotes the estimate of y_{is} , m_{kk} is the kk^{th} element of the matrix m and $\hat{f}_h(t) = \frac{1}{N} \sum_{s=1}^T K_{h,i}(t, s)$ (compare Härdle (1990) for details.).

A problem that is very often neglected in the literature is that the estimations based on smoothing methods are biased and less accurate near the boundaries of the observation interval. The reason is that the kernel is truncated at the starting and end point, wherefore the estimates are based on one-sided data information. Several methodologies are proposed in the literature to correct for these ‘boundary effects’. One solution is to modify the kernel, which is called the ‘boundary kernel approach’.⁴ Another solution is the so called ‘transformation method’ as proposed by Wand and Ruppert (1991), Marron and Ruppert (1994) and Yang (2000). And, finally, one can solve the ‘boundary problem’ by modifying the bandwidth near the edges⁵, which is the methodology we apply in this paper.

⁴See Gasser and Müller (1979), Gasser et al. (1985), Granovsky and Müller (1991), Müller (1991) and Jones (1993).

⁵Compare Rice (1984), Gasser et al. (1985), Müller (1991), Hall and Wehrly (1991).

Based on the idea presented in Dai and Sperlich (2010), we reduce the bandwidth in the boundary region to reduce the ‘boundary effect’. We use locally reduced bandwidths in the boundary areas and the optimal (global) bandwidths h^* in the interior time period for the estimation. This methodology is advantaged in that it is easy to implement and, as Dai and Sperlich (2010) show, is more efficient than alternative bandwidth correction methodologies. For $t = 1, \dots, T$, the bandwidth used for the estimation is defined by:

$$h = \begin{cases} \max(t - 1, \epsilon) & \text{if } 1 \leq t < (1 + h^*), \\ \max(T - t, \epsilon) & \text{if } (T - h^*) < t \leq T, \\ h^* & \text{otherwise} \end{cases} \quad (9)$$

with $\epsilon > 0$. According Dai and Sperlich (2010), ϵ approaches zero as $N \rightarrow \infty$. Thus, for a large number of cross-sections, the slope coefficients at the boundaries are estimated on basis of the observations in the very close neighborhood around the first and last time period of the data sample, i.e. $t = 1$ or $t = T$. However, our panel consists of only ten countries, thus $N = 10$. To get consistent estimates at the boundaries, we must choose a larger ϵ to take the observations in the neighborhood into account. However, ϵ should be smaller than h^* in order to reduce the boundary effect.

Before estimating the time-varying slope coefficients according equation (4) and (9), we need to determine the optimal bandwidth, h^* . If the bandwidth parameter h^* is small, the modeling bias or approximation error is small, whereas the variance of the estimated parameter is large, since few data points fall in the small local neighborhood around t . However, if h^* is large, the variance of the estimations is smaller compared to the case of small bandwidth, but the estimation is biased. Thus, there is a bias and variance trade-off in the choice of the smoothing parameter. We apply the method of cross-validation for choosing the optimal bandwidth h^* , as explained in Hoover et al. (1998). Let $\hat{\beta}^{(-i)}(t)$ be the slope coefficient estimated according equations (4) and (9), where we leave out all the observations of the i^{th} subject. The optimal smoothing parameter is then given by minimizing the

following cross-validation average predictive squared error criterion:

$$\min_{h^*} CV(h) = \frac{1}{NT} \sum_{i=1}^N \sum_{t=1}^T \left\{ Y_{it} - X'_{it} \hat{\beta}^{(-i)}(t) - \hat{\mu}_i \right\}^2 \quad (10)$$

4 Data and Descriptive Analysis

We analyze the sovereign bond yields of ten euro area countries: Belgium, Finland, France, Greece, Ireland, Italy, the Netherlands, Austria, Portugal and Spain. The time covered runs from Q1/1999, the beginning of EMU, until Q1/2010, such that our data sample consists of in total 440 observations. The yield spreads of the individual countries are calculated as the end-quarter yield differential of their 10-year benchmark bonds relative to the 10-year German Bund. Since Greece joined EMU only in 2001, its yield differentials up to Q4/2000 also contain an exchange rate risk premium that compensates the investor for a possible devaluation of the Greek drachma with respect to the euro. We therefore adjust the Greek yield spreads recorded before Q1/2001 by subtracting the yield spread between Greek and German 10-year interest rate swaps, which is regarded as a measure for exchange rate uncertainty (compare e.g. Gomez-Puig (2006)).

Figure 2 in the appendix plots the sovereign bond yield spreads over the time period analyzed. After the introduction of the euro in 1999, or 2001 in case of Greece, yield differentials across member states were small and hardly any differentiation across countries was visible. Between Q1/1999 and Q1/2005 the yield spreads narrowed even further reaching the trough in Q4/2004. Thereafter, they started to diverge slightly. With the start of the global financial crisis in Q4/2007, yield spreads started to increase considerably, reaching spreads of more than 250 basis points. In the subsequent quarters, the yield differentials decreased again somewhat, signalling an easing of global financial market tensions. The only exception is Greece, for which the interest differential started to increase considerably again in Q4/2009 as a consequence of worries over a possible default, reaching a record high of

almost 350 basis points in Q1/2010.

We explain government bond yield spreads with variables proxying credit risk, liquidity risk and general risk aversion. To estimate credit risk, we focus on variables measuring the fiscal performance of a country. We expect that if the fiscal position of a country deteriorates relative to the benchmark country, the bond spread increases, as the market asks for a higher default risk premium. Thus, we use the debt to GDP ratio and the projected (12-months ahead) deficit to GDP ratio as explanatory variables, which are commonly used in the literature.⁶ Both variables are expressed in differences to the corresponding debt and deficit figures of the benchmark country, Germany. Eurostat provides the debt to GDP ratios on a quarterly basis starting with 2000. Before 2000, the debt data is only available annually. Hence, we interpolate the debt in 1999 to a quarterly frequency. The projected deficit to GDP ratios for the preceding year are published, semi-annually, in the OECD Economic Outlook. Thus, we allocate the projected deficit figures published in the mid-year report to the first two quarters of the year, and the projected deficit figures published in the end-year report to the third and fourth quarters of the year.

We include levels and quadratic terms of the fiscal variables to test for 'credit punishing' effects, meaning that interest rate spreads grow non-linearly with the level of fiscal variables (compare e.g. Bayoumi et al. (1995)). The inclusion of squared fiscal variables assures that the variation of estimated coefficients over time is indeed attributable to an alteration in risk pricing and does not indirectly reflect a non-linear reaction of interest rates with respect to the debt or deficit level.

Empirical papers examining market liquidity in bond markets use both direct measures based on transaction data, such as trading volume or bid-ask spreads, and indirect measures based on bond characteristics, such as the

⁶Using the projected rather than the current deficit to GDP ratio has the advantage that one takes the forward looking behavior of financial markets into account. Further, it avoids a potential endogeneity bias in the estimation results, which could arise from the fact that the current deficit figure contains government interest payments.

outstanding amount of debt securities and the issue size of the specific bond. Several studies show that indirect and direct liquidity measures are closely related to each other.⁷ We focus on the bid-ask spread as a measure of liquidity risk, which is commonly considered as the best measures for liquidity.⁸ The bid-ask spread measures the cost associated with bond trading and is influenced by market depth. A deep market is considered to have low bid-ask spreads, which reduces the liquidity premium contained in bond yield spreads. We calculate the end-quarter bid-ask spreads for the same benchmark bonds used also for the calculation of the yield spreads. We measure the bid-ask spreads relative to the bid-ask spread of the German benchmark bond.

Finally, we use the corporate bond yield spread as a proxy for general investors' risk aversion, which is a conventionally used measure in the related literature.⁹ The corporate bond spread measures the spread between low grade corporate bonds (Merrill Lynch BBB) and government bonds. In times of greater uncertainty, the corporate bond yield spread widens because of a shift in investor preference from riskier corporate bonds to safer government bonds. Thus, assuming that the benchmark country Germany is a 'safe haven' among EMU countries, we expect a positive relationship between the corporate bond yield spread and sovereign bond yield differentials. Since this study focuses on euro area sovereign bond yield spreads, ideally we would use the corporate bond spread measured for the European Union. However, such a variable is provided only from 2002 onward. Therefore, we use the

⁷Ejsing and Sihvonen (2009) show that there is a highly significant relationship between bid-ask spreads, trading volume and the log issue size for German and French bond markets. Korajczyk and Sadka (2008) find that there is a common component among different measures of liquidity, which in turn suggests that one can view each individual measure of liquidity as an approximation to the underlying liquidity factor. Gerlach et al. (2010) show that their results are robust to various liquidity measures, i.e. bid-ask spreads, total amount of outstanding bonds and actual turnover.

⁸See e.g. Flemming (2003) or Barrios et al. (2009), who argue that bid-ask spreads are better indicators for gaging liquidity conditions in bond markets than traded volumes, since data on volume can be affected by multiple trading operations between bank's affiliates to meet balance sheet requirements.

⁹Compare e.g. Codogno et al. (2003), Favero et al. (2010) and Bernoth et al. (2006).

corporate spread measured for the USA.¹⁰

Figure 3 in the appendix illustrates the development of the US and EU corporate bond yield spread in quarterly frequency for the time period analyzed. The high correlation of the two series is obvious, indicating that data on US corporate government bond yield spreads can be used as a good proxy for investor risk attitude in the euro area. Between Q1/1999 and Q4/2002 the spread hovered around a value of 250 basis points. Thereafter, the corporate bond yield spread steadily decreased reaching a trough in Q3/2004. With the start of the financial crisis in autumn 2007 the corporate bond yield spread again widened considerably. After the collapse of Lehman Brothers, the spread more than tripled, reaching values close to 700 basis points in Q4/2008. In 2009, it decreased substantially, probably due to massive government and central bank interventions.¹¹

Detailed summary statistics of all variables and information about the data sources are provided in Table 2 in the appendix.

5 Estimation Results

As a starting point, we ignore the fact that the determinants of euro area sovereign bond yield spreads may be time-varying by estimating a standard OLS fixed effects panel data model.¹² We use the corporate bond yield

¹⁰Following Codogno et al. (2003), Haugh et al. (2009), Attinasi et al. (2009) and Barrios et al. (2009), we test whether the interaction terms between the fiscal variables and the general risk aversion indicator play a role in explaining yield differentials. Similar to Codogno et al. (2003), we do not find a significant amplified effect of the fiscal variables in times of high risk aversion. Moreover, following Beber et al. (2009) and Favero et al. (2010) we add a variable interacting the variable measuring global risk aversion with the liquidity variable to test whether the level of the general risk factor significant impacts the effect of liquidity on bond yields. This interaction variable is also highly insignificant.

¹¹Gerlach et al. (2010) use also three alternative measures for aggregate risk: the VIX (implied equity market volatility); the US agency spread, Refcorp; and the Treasury-to-T-Bill (Ted) spread. They find that that the VIX and Refcorp are both highly correlated with the corporate bond yield spread and that all variations of the aggregate risk measure yield comparable results in the analysis of sovereign bond yield spreads.

¹²Note that these results can also be obtained when estimating equation (4) for very large values of the bandwidth parameter h .

spread, the bid-ask spread and the linear and squared debt and projected deficit ratio as explanatory variables. Table 1 shows the results.

Table 1: Fixed effects panel data model with constant coefficients

	Beta	Std. Dev.
US BBB spread	12.45	0.73***
Debt	0.51	0.14***
Debt ²	0.002	0.00
Proj. Deficit	9.08	0.61***
Proj. Deficit ²	1.02	0.14***
Bid-ask spread	4.11	1.50***
R ²	0.71	
N	440	

*, **, *** indicate significance at the 10, 5, and 1% significance levels respectively.

In line with previous findings in the literature, we find that sovereign yield spreads display default and liquidity risk premia as well a global risk premium. The US corporate bond spread ('US BBB-spread') and both fiscal variables have a significant positive and the bid-ask spread a significant negative effect on interest differentials. The coefficient of the squared debt variable turns out to be insignificant, while the coefficient of the squared projected deficit variable is significantly positive, which suggest that the marginal effect for higher deficit ratios increases with higher fiscal imbalances. Thus, we find some 'credit punishing' effects, indicating that the effect of a worsening of the fiscal position on the yield spread increase with a country's deficit level. However, if the real underlying coefficients are time varying, as several previous studies have shown¹³, the estimated coefficients in Table 1 are inaccurate and may also be misleading. For instance, if the deficit levels of all countries start to rise at the same time, as it was the case during the 2008-9 financial crisis, one cannot identify with this static panel model, whether the significance of the squared deficit variable is due to a non-linear reaction with respect to the deficit level, or due to varying coefficients over time. This is

¹³Compare e.g. Bernoth et al.(2006; 2010), Sgherri and Zoli (2009), Haugh et al. (2009).

the reason why we refrain from drawing any conclusions from this estimation result. Instead, it serves as reference for the following analysis.

We overcome the deficiencies of the static linear panel model by estimating a time varying coefficients framework using a semiparametric fixed effects panel data model as described in section 3. For the estimations of equation (5) and (6), we use the Gaussian kernel as our smoothing function:

$$K(u) = \frac{1}{\sqrt{2\pi}} \exp\left(-\frac{1}{2}u^2\right),$$

where $u = (t - s)/h$. Thus, to estimate the slope coefficients at a specific point in time, t , all observations contained in our data sample measured between $s = 1, \dots, T$ and $i = 1, \dots, N$ are given a positive weight in the estimations. However, the observations in the close neighborhood around time t get a larger weight than observations measured much earlier or later. Thus, the weights given to each observation are a decreasing function of the time distance between the point of time under consideration and the actual time of observation, $|t - s|$.

Similar to the static panel, we start with using the corporate bond yield spread, the bid-ask spread and the linear and squared debt and projected deficit ratio as explanatory variables. According the ‘leave-one-out cross-validation’ methodology described in equation (10), the optimal smoothing parameter is $h^* = 1.8$ in the search interval $[0.005, 5]$. Thus, referring to equation (9), the estimates of the first two and the last two quarters of our data sample are based on reduced bandwidths to eliminate the ‘boundary effect’. In our estimations, we set $\epsilon = 0.8$.

Our estimation results show that the coefficient on the squared projected deficit variable is always insignificant. The coefficient of the squared debt variable is very small in magnitude and significant in only five out of 45 quarters (Q1/1999, Q2/2001-Q4/2001, and Q1/2010). Thus, our estimation results suggest that ‘credit punishing’ effects seem not to play a significant role. Accordingly, the time-variation observed in the linear fiscal variables is only attributable to an altering in risk pricing over time and does not reflect

indirectly a non-linear reaction of interest rates with respect to the debt or deficit level. Since all other variables seem unaffected by the inclusion or omission of the two quadratic fiscal variables, we save degrees of freedom by excluding the squared fiscal variables from our model.¹⁴

Figures 1(a)-1(d) present the estimated time-varying coefficients together with the 95% pointwise confidence intervals when regressing the sovereign yield spreads on the reduced set of regressors. In this case, the optimal smoothing parameter is $h^* = 1.5$, meaning that again the first two and the last two quarters of our data sample are estimated with reduced bandwidths ($\epsilon = 0.8$).¹⁵

Our estimation results indicate that the degree of general investors' risk aversion plays an important role in explaining sovereign bond yield spreads in the euro area. Figure 1(a) shows that over the entire observed time period the coefficient on the US corporate bond yield spread ('US BBB-spread') was with only one exception always positive, indicating that in periods of high global risk aversion, the interest differentials of EMU countries versus Germany rose. At the beginning of EMU, the interest differential significantly increased by around 10 basis points for every percentage point increase in the corporate bond spread. In the proceeding years, however, between Q1/2001 and Q3/2006, the impact of the global risk factor became much weaker and turned insignificant, indicating that Germany lost its safe haven status in this period. From Q4/2006 onwards, however, two years before the fall of Lehman Brothers, the impact of the global risk factor on euro area yield differentials increased continuously and became significant again. Thus, financial markets started worrying about financial market risk long before the start of the financial crisis. The coefficient on the corporate bond yield spread rose from approximately five to 18 in Q1/2010, indicating that financial markets

¹⁴The estimation results of the extended set of regressors are available upon request from the authors.

¹⁵For some estimated coefficients we observe a widening of the confidence bounds at the end of the observed time period. This widening, however, starts way earlier than in the last two quarters and can therefore not be attributed to boundary effects, but to an increased volatility in the data after the outbreak of the financial crisis.

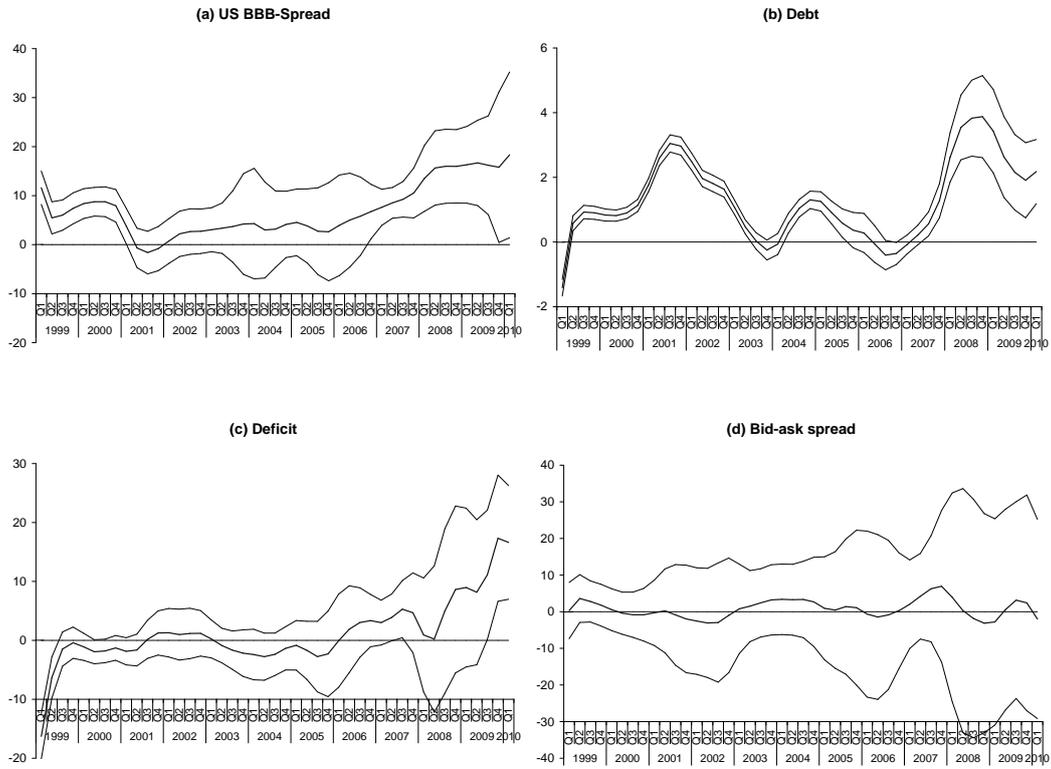


Figure 1: Nonparametric estimates of the time-varying coefficients

granted Germany safe haven status again, which is much more pronounced than at the beginning of EMU.

Figure 1(b) plots the estimated coefficient of the debt variable over time. Our results suggest that markets perceived and priced sovereign debt differentials significantly in general, with two exceptions: the periods of Q3/2003-Q1/2004 and Q4/2005-Q2/2007, when the coefficient of debt to GDP ratio is slightly insignificant.¹⁶ Mid-2001, a debt differential of 10 percent over Germany increased the yield spread by around 30 basis points. In the subsequent years, the impact of an increase in the debt ratio on the interest differential diminished somewhat and usually did not exceed more than 18 basis points

¹⁶The coefficient of the debt ratio is negative in 1999; however, this may be attributable to the interpolation of yearly debt data in this year to a quarterly frequency. The interpolated variable may be less accurate and this result should not be overstated.

for every 10 percent increase in the debt differential. With the outbreak of the financial crisis in Q3/2007, markets began to price fiscal indebtedness much more than they did before. The increase of the yield differentials in response to a debt to GDP differential of 10 percent rose from around 14 basis points at the end of 2007 to around 40 basis points at the end of 2008. In 2009, the coefficient of the debt variable decreased again somewhat, signaling easing financial market tensions. However, this easing did not last for long. With the beginning of the European fiscal crisis at the end of 2009, when financial markets started worrying about the sustainability of Greek, Irish and Portuguese debt, the coefficient on the debt differential increased again slightly. In Q1/2010, a 10 percent debt to GDP differential relative to Germany caused an increase in the yield differential of 22 basis points.

The estimated coefficients of the deficit variable are presented in Figure 1(c). We find that before the financial crisis, the coefficient of the deficit differential between the issuer country and Germany fluctuated around zero and was insignificant. Thus, while financial markets paid attention to the debt to GDP ratio, they ignored deficit differentials in this period. After the intensification of the financial crisis in Q3/2008, the coefficient of the deficit differential is continuously positive and shows an increasing trend. However, it is only with the beginning of the European fiscal crisis at the end of 2009 that the coefficient became significant. In Q1/2010, the interest differential increased by around 16 basis points for a one percentage point increase in deficit differential. Thus, only after the outbreak of the crisis did financial markets begin to perceive budget deficits and to punish financial loosening. This result coincides with the findings of Aßmann and Boysen-Hogrefe (2009), who also estimate a time-varying coefficient model.

Finally, the estimation results presented in Figure 1(d) suggest that liquidity premia never played a role in explaining bond yield differentials in EMU. The coefficient on the bid-ask spread mostly shows the expected positive sign, but it is always highly insignificant. We attribute this result to the fact that after the conversion of all existing government debt of the EMU countries into euro, the euro-denominated debt market increased for all countries substantially,

such that liquidity differences across member countries played not such an important role anymore. This result is consistent with a sound degree of financial market integration within the euro area.

Thus, comparing the estimation results of the time-varying coefficient model with that of the static fixed effects model in Table 1, it seems that a standard OLS fixed-effects panel model cannot adequately estimate the dynamics of the determinants of sovereign yield spreads. Additionally, when comparing the R^2 values of both models, the time varying coefficients framework clearly outperforms the standard OLS fixed-effects panel model. The former is able to explain about 95 percent of the variation in sovereign yield spreads, while the latter only around 70 percent.

6 Conclusions

This paper contributes to the literature on sovereign risk in the euro area by applying a time-varying coefficient fixed-effects panel model. We analyze the government bond yield spreads of 10 euro area countries between Q1/1999 and Q1/2010. By estimating time-varying coefficients we take into account that government bond yields may not only be affected by changes in macroeconomic fundamentals but also by shifts in the pricing of sovereign risks.

To model the time-varying determinants of EMU yield spreads, we use an additive nonparametric regression approach. Compared to parametric models, nonparametric regressions have the advantage that no information on the functional form of the model coefficient is needed, which is especially important for our data set, where we observe both periods of convergence and divergence in our yield spreads.

We find that the impact of fiscal variables and the global risk factor on EMU yield differentials varies considerably over time. Before the financial crisis, financial markets paid no attention to government deficit ratios, while they almost continuously monitored the debt to GDP ratio of the individual coun-

tries, which is also the more relevant variable to assess fiscal sustainability. Thus, financial markets play an important role in disciplining highly indebted countries.

We find that the observed strong increase of sovereign bond yield spreads during the financial crisis can be attributed to three factors: first, an increase in general investors' risk aversion; second, a deterioration of the fiscal position (both, in terms of debt but also in terms of deficits) of European governments; and third, by an increase in the price of risk. However, the pricing for the different risk components did not alter simultaneously. Financial markets began worrying about general market risk long before it worried about sovereign default risk. By the end of 2006, general investor risk aversion started having a growing impact on sovereign risk premia and to revitalize Germany's safe haven status. The strengthening of the fiscal discipline imposed by financial markets, however, was not pronounced until end of 2008, when the fiscal crisis intensified after the fall of Lehman Brothers.

The finding that financial market discipline has become stronger after the financial crisis underlines the importance of sound fiscal policies and the urgent need of fiscal consolidation of the highly indebted EMU countries in order to reduce their mounting interest rate burden.

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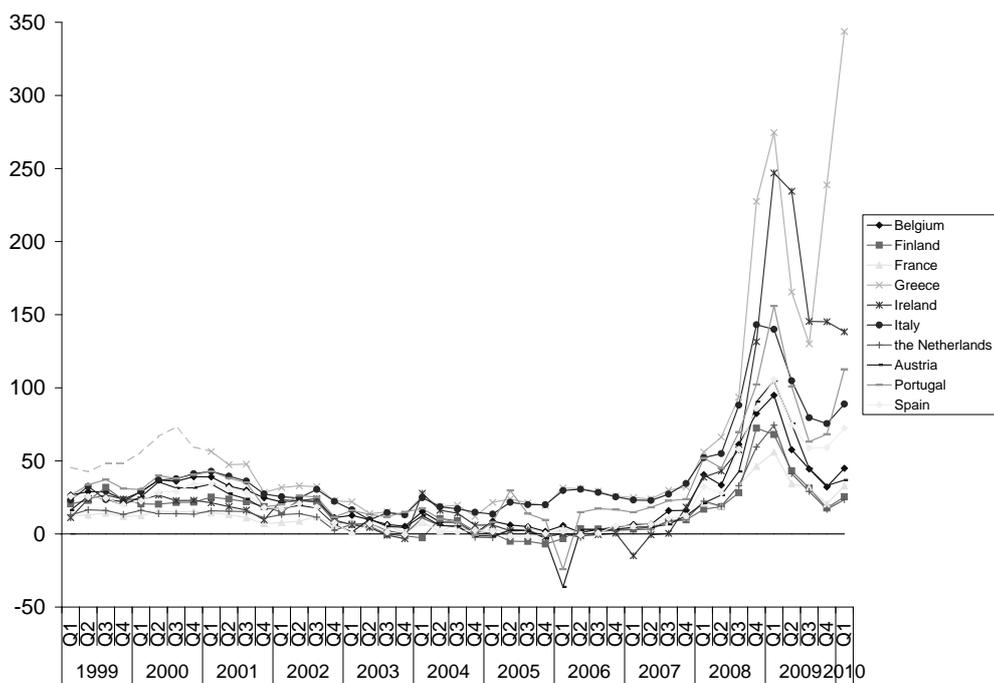
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Appendix

Figure 2: Ten-year sovereign bond yield spreads in basis points



Note: Yield spreads of Greece adjusted for exchange rate risk premia between Q1/1999-Q4/2000.

Figure 3: US and EU BBB corporate bond spreads in percent

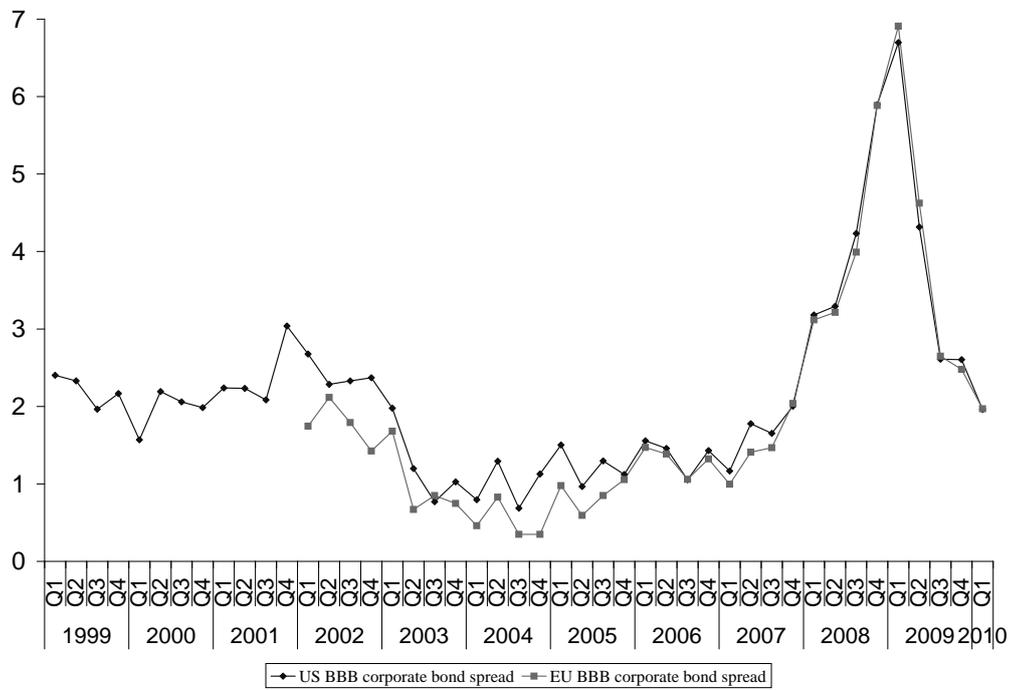


Table 2: Data Description

Variable	Description	Source	Mean	Std. Dev.	Min.	Max.
Yield spread	Spread between the ten-year sovereign bond of an EMU country and the ten-year German sovereign bond in basis points, quarterly frequency. ^{a)}	Bloomberg	27.86	38.85	-36.20	343.70
Debt differential	Difference of debt to GDP ratio of an EMU country over that of Germany in percent, quarterly frequency. ^{a)}	Eurostat	3.79	26.36	-42.80	59.65
Projected deficit differential	Difference of projected deficit (1-year ahead) to GDP ratio of an EMU country over that of Germany in percent, semi-annual frequency.	OECD Economic Outlook	0.89	2.36	-8.50	7.43
Bid-Ask Spread differential	Spread between bid and ask quotations for the relevant bond in basis points, quarterly frequency. ^{a)}	Bloomberg	0.12	0.75	-1.21	5.49
US BBB Spread	Spread between the yield on US seven to ten-year BBB corporate bonds and the yield on ten-year US treasury benchmark bonds in percent, quarterly frequency. ^{a)}	Merrill Lynch	2.15	1.22	0.69	6.70
EU BBB Spread	Spread between the yield on EMU seven to ten-year BBB corporate bonds and the yield on ten-year Euro benchmark bond (synthetic) in percent, quarterly frequency. ^{a)}	Merrill Lynch	1.89	1.54	0.35	6.91

^{a)} Measured as end-of-quarter values.