# Sovereign risk contagion in the Eurozone: a time-varying coefficient approach

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# Sovereign risk contagion in the Eurozone: a time-varying coefficient approach

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**Abstract**

This paper analyzes sovereign risk contagion in the Eurozone using an extension to the canonical model for contagion proposed by Pesaran and Pick (2007) and Metiu (2012) to allow for time-varying coefficients. This becomes necessary due to changes in the risk pricing of sovereign bonds since the onset of the recent crisis period and due to the presence of contagion in typically bounded time intervals. Controlling for changes in the risk pricing by investors, we detect several channels of pure contagion between 2008 and 2012. Further, we find that the bailout-programs for Greece, Ireland and Portugal led to a disruption in contagion of sovereign risk from these countries to Spain, Italy, France and Belgium as was desired by policymakers. For all countries considered, we observe an increase in the relevance of general risk aversion towards sovereign debt since May 2010. This development partially replaced the significance of country-specific credit risk factors in explaining bond yield spreads. Our model extension yields a device that is suitable to determine whether policy interventions are required and to judge their success ex-post.

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*Keywords:* contagion; credit events; Eurozone; financial crisis; sovereign risk; time-varying approach.

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# 1 Introduction

Until the outbreak of the global financial crisis, sovereign bond yield spreads in the Eurozone saw an unprecedented level of convergence. In 2007 the mean 10-year sovereign bond yield spread in relation to Germany amounted to a mere 24bp for Greece, 16bp for Portugal and 12bp for Ireland. At the end of 2008, however, spreads already rose to 201bp for Greece, 94bp for Portugal and 121bp for Ireland. Maximum values were reached at more than 5,500bp for Greece in March 2012, 1,400bp for Portugal in January 2012 and more than 900bp for Ireland in July 2011. What are the drivers of these sharp increases of bond yield spreads?

A large body of literature is devoted to the fundamental determinants of bond yield spreads. Bond yield spreads may contain a premium for the credit risk of the underlying debt (Codogno et al., 2003), for the liquidity risk arising in markets without sufficient depth or breadth (Barrios et al., 2009) as well as for the general risk aversion of market participants (Favero et al., 2010). Using time-varying coefficient and dummy variable approaches, respectively, Beirne and Fratzscher (2013), Bernoth and Erdogan (2012), Giordano et al. (2013) and von Hagen et al. (2011) find that changes in the pricing of sovereign risk by market participants ("wake-up-call contagion") contribute to explaining increasing yield spreads during the recent financial crisis. Especially, the general risk aversion towards Eurozone bonds in relation to German Bunds and the relevance of fiscal variables that represent credit risk increased considerably.

However, changes in the risk-pricing of these fundamentals may not suffice to fully explain the evolution of bond yields in times of increased market uncertainty. Negative shocks in a single country may directly lead to increasing spreads in other countries ("pure contagion"). Evidence for pure contagion in the Eurozone between 2008 and 2012 is provided by e.g. Afonso et al. (2012), Arezki et al. (2011), Wing Fong and Wong (2012), Kalbaska and Gatkowski (2012), Mink and de Haan (2013) and Missio and Watzka (2011). To identify pure contagion relationships, Pesaran and Pick (2007) propose a canonical model for contagion that directly accounts for the fundamentals of bond yields and identifies contagion effects under (observed and latent) interdependence (Metiu, 2012). In this model, the significance of positive contagion coefficients reveals the existence of shock-transmission from one to another country.

Applying the model proposed by Pesaran and Pick, Metiu (2012) provides evidence for contagion of sovereign risk in the Eurozone, e.g. from Greece to Belgium, France, Portugal and Spain. In his analysis, Metiu assumes the constancy of coefficients of fundamentals of bond yield spreads as well as of contagion coefficients in the entire sample (January 2008 - February 2012). A time-varying approach, however, is already inevitable due to the changes in risk pricing of yield spread determinants over time (Bernoth and Erdogan, 2012). Moreover, it is unlikely that contagion i) happens in the entire sample period and ii) occurs in the same intensity during the whole sample period. Instead, policy measures are likely to have an effect on contagion relations and may lead to their disruption. Moreover, an assessment of the effect of policy measures at a given time is not feasible when the time-variation of coefficients is not taken into account.

In this work we investigate the effects that arise from a combination of both the time-variation in coefficients of yield spread fundamentals as well as in pure contagion relationships. To measure credit risk, liquidity risk and the general risk aversion of market participants, we use these variables that the literature on bond yield fundamentals considers as the best proxies at high frequencies. We then propose the application of Pesaran and Pick's model in rolling windows of equal length which allows a simultaneous study of wake-up-call contagion and pure contagion.[1] The use of rolling windows allows for transitions of coefficient values without the need to pre-specify a stochastic process for the coefficients. Further, in every window only the information set is used that is observable until the end of the window. Therefore, at each point of time the contagion effects observed in the most recent sub-period can be extracted. Our approach can hence be used for prediction purposes and is a suitable means to judge the success of policy measures. It may also serve as a trigger to such measures as the presence of contagion could be effectively stricken down by policy interventions (Dornbusch et al., 2000, Pesaran and Pick, 2007). Moreover, as opposed to a large body of literature in which the distinction between pre-crisis period and possibly different crisis sub-periods is pre-defined exogenously by the

[1] By the time finalizing this paper, we became aware of a working paper prepared by Leschinski and Bertram (2013) who - independently from our work - also propose to apply the Pesaran and Pick model in rolling windows. Yet, their results are questionable as they imply e.g. a positive effect of Greek shocks on the spreads of other crisis and non-crisis countries (see Table 2 in their paper), stronger contagion effects in the subprime crisis period than during the actual sovereign debt crisis in the Eurozone (see Figure 4 in their paper) as well as similarly strong contagion effects to non-crisis countries in the tranquil period from 2003 till 2006 (see Figure 2 in their paper). Their results differ from the existing literature on contagion in the Eurozone debt crisis (see e.g. Missio and Watzka, 2011, Arghyrou and Kontonikas, 2012) and from ours for which we will give reasons in Section 2.4.

authors, our approach does not require such a sharp setting. Instead, we will determine structural breaks in the relation between yield spreads and their fundamentals as well as in contagion relations endogenously.

Our results include evidence for the presence of sovereign risk contagion in the Eurozone. We show that contagion effects are indeed time-dependent and should hence not be modeled assuming constant coefficients for the entire sample. Sovereign risk contagion from Greece, for example, can be detected only towards Portugal and Belgium in the constant coefficient model, whereas the generalized model with time-varying coefficients also shows contagion towards Ireland and Italy for sub-periods at the beginning of the sovereign debt crisis. On the other hand, the constant coefficient model yields evidence for contagion from France to Greece, which is not in accordance with the existing literature and cannot be confirmed by the time-varying coefficient model.

We find that contagion from the three countries mostly affected by the crisis (Ireland, Portugal and Greece) to the other four countries considered in this paper (France, Spain, Italy and Belgium) ceased after the bailout programs for the former set of countries were established. Despite the appearance of further credit events in Greece, Ireland and Portugal after their bailout, negative shocks were not transmitted to other countries anymore. Spain and France were not affected by credit events in Greece through contagion at any time. We find that their sovereign risk is rather influenced by changes in the pricing of sovereign risk, especially with regard to the increase in the general risk aversion of investors (interdependence).

For the fundamental factors explaining the levels of bond yield spreads, we argue that only a reduction of the general risk aversion can provide a systematic decrease of spreads in all Eurozone countries considered, cet. par. We find a substantial increase of the relevance of general risk aversion in the Eurozone for the yield spreads of all countries since May 2010. Country-specific credit risk measured by CDS spreads in relation to Germany was almost always significant for Spain, Italy, Portugal and Ireland, however with a decreasing trend in the coefficients for Portugal and Ireland. For Greece and Belgium this factor was only relevant till mid-2011 and is likely to have been substituted by the increased risk aversion of investors towards the Eurozone as a whole.

The paper is structured as follows. Section 2 explains the econometric framework of our analysis. Bond yield fundamentals are discussed in Section 3. In Section 4, we contribute a justification for the use of levels of bond yield spreads as opposed to first differences. Further, we run the model under the assumption of constant coefficients (as in Metiu, 2012) as well as with time-varying coefficients. We also report results from robustness checks to our model. Section 5 gives concluding remarks.

# 2 Econometric framework

We investigate the timing of contagion and its direction in a dynamic extension of the approach suggested by Metiu (2012). To test for contagion of sovereign risk in the Eurozone, Metiu applies Pesaran and Pick's (2007) canonical model for contagion (Section 2.1) and contributes a refined approach to identifying credit events that is based on violations of one-step-ahead Value-at-Risks derived from the distribution of innovations (Section 2.2). Further, Metiu (2012) recommends the use of new back-testing methods for Value-at-Risk recently developed by Candelon et al. (2011) to validate the selection of credit events (Section 2.3). As Metiu's approach assumes stability in the relationship between bond yield spreads and their fundamentals as well as stable shock-transmission channels throughout the whole sample, which is rather unlikely to happen, we apply a time-varying coefficient approach. Our strategy to extract time-varying parameter coefficients and point-wise confidence intervals is based on the application of Pesaran and Pick's canonical model for contagion in rolling windows of fixed size (Section 2.4).

The following description of our econometric setup follows a top-bottom approach, i.e. we start with the final equation to test for contagion and comment on steps taken to operationalize this equation.

## 2.1 Canonical model for contagion

Metiu (2012) extents the canonical model of Pesaran and Pick (2007) to assess the presence and direction of contagion (shock-transmission from country $j$ to country $i$) of sovereign risk by testing the significance of parameters $\delta_{i,j}$ in the following equation:[2]

$$y_{i,t} = \sum_{l=1}^{q} \alpha_{i,l} y_{i,t-l} + \beta_i' g_t + \gamma_i' cs_{i,t} + \sum_{\substack{j=1 \\ j \neq i}}^{N} \delta_{i,j} C_{j,t} + u_{i,t} \tag{1}$$

[2]For the ease of notation, we skip the constant term in this and the following regression equations.

Here, $y_{i,t}$ denotes the sovereign bond yield spread of country $i = 1, \ldots, N$ at time $t = 1, \ldots, T$. Yield spreads are explained by country-specific factors $cs_{i,t}$ as well as common global factors $g_t$. The inclusion of at least one country-specific variable is necessary to ensure the identifiability of the parameters $\delta_{i,j}$ (Pesaran and Pick, 2007). Using global as well as country-specific factors, the model accounts for the fundamentals of bond yields and identifies contagion effects controlling for (observed and latent) interdependence (Metiu, 2012). The first sum accounts for the autocorrelation of bond yield spreads. The indicator variable $C_{j,t}$ takes the value one if country $j$ is in a credit event at time $t$, otherwise zero (see Section 2.2). To estimate (1), we use an estimator that is robust against heteroscedasticity and autocorrelation in that we determine the variance of $u_{i,t}$ by a quadratic spectral kernel with data-dependent optimal bandwidth (Newey and West, 1994).

Further, Pesaran and Pick (2007) show that the endogeneity of credit event indicators would lead to inconsistent estimates of the coefficients $\delta_{i,j}$ when ordinary least squares (OLS) is used, i.e. the inference whether contagion occurred or not would be biased. Hence, the credit event indicator for country $j$ should be instrumented in a two-stage least squares (2SLS) regression. We apply the generalized instrumental variable estimation (GIVE) procedure as in Pesaran and Pick (2007) and Metiu (2012) who use lagged dependent variables $y_{i,t-l}$ of all countries $i \neq j$ to instrument $C_{j,t}$. As the credit event indicators are a non-linear function of the dependent variables, power series of the instruments are included in the set of instruments to increase their strength to approximate the indicators and to obtain consistent 2SLS estimates (Kelejian, 1971). This leads to the following first stage equation

$$C_{j,t} = \sum_{\substack{i=1 \\ i \neq j}}^{N} \sum_{l=1}^{q} \sum_{r=1}^{m} \mu_{i,l,r} (y_{i,t-l})^r + \epsilon_{j,t} \tag{2}$$

which is followed by the estimation of eq. (1) with the plug-in of the estimate $\hat{C}_{j,t}$ in place of the original credit event indicator $C_{j,t}$. The derivation of the credit event indicators will be explained in the following.

### 2.2 Determination of credit event indicators

In case of frequent (partial) credit defaults or restructuring of sovereign debt, the credit event indicators $C_{i,t}$ could be based on such real events. However, due to their scarcity, a broader definition of bond market distress has to be found. Pesaran and Pick (2007) derive such events from the time series of bond yield spreads and set the indicator to one for these times $t$ when the first difference of the yield spread exceeds twice its standard deviation.[3] This is a suitable approach for the case of a constant variance $\sigma_{i,t}^2 = \sigma_i^2$ of innovations $u_{i,t}$ of eq. (1). However, Forbes and Rigobon (2002) show that a change of their variance, which is likely to happen in times of financial distress, distorts the judgement of whether contagion took place or not. Metiu (2012) suggests a remedy to this issue. He infers credit events from violations of one-step-ahead Value-at-Risks $VaR_{i,t|t-1}$ of the probability distribution of innovations $u_{i,t}$ conditional on the information set $\mathcal{F}_{t-1}$ available at time $t-1$. Innovations are assumed to be $t(\nu)$-distributed such that

$$VaR_{i,t|t-1} = F^{-1}(p,\nu)\sigma_{i,t} \tag{3}$$

where $F(\cdot,\nu)$ denotes the cumulated distribution function of a $t(\nu)$-distributed random variable and $p$ equals the confidence level. The degrees of freedom of the $t$-distribution can be chosen according to the results of appropriate backtesting procedures for the calibration of Value-at-Risks (see Section 2.3). Hence, to test for contagion, credit events are defined as the points in time when the increase of a bond yield spread exceeds the upper bound of its prediction interval. In other words, credit events take place on trading days with significant deviations of the spread from the current risk pricing of market participants through global and country-specific factors.

To estimate the standard deviation $\sigma_{i,t}$ in eq. (3), Metiu (2012) proposes a GARCH(1,1) process for the conditional variance of the yield spread

$$\sigma_{i,t}^2 = \varphi_i + \theta_i u_{i,t-1}^2 + \tau_i \sigma_{i,t-1}^2 \tag{4}$$

whose parameters have to fulfill the conditions $\varphi_i > 0$, $\theta_i, \tau_i \geq 0$, and $\theta_i + \tau_i < 1$.[4] The length of the sub-intervals to estimate eq. (4) is kept equal to a fixed window with length $wl$, i.e. we consider rolling intervals $[t-wl+1,t]$. As old observations are thus continuously

[3] Alternatively, rating downgrades could be used to define credit events here. However, rating downgrades can also lag behind market developments: Afonso et al. (2012) provide evidence on two-way causality between ratings and sovereign bond yield spreads of 24 EU countries; Karmann and Maltritz (2012) show that bond market data implied increases of the probability of default of Greece several months before the downgrades of Greece by S&P and Fitch in December 2009.

[4] If these conditions were breached for a pair $(i,t)$, we set the credit event indicator to zero. We also ran a robustness check setting $\sigma_{i,t}$ in eq. (3) to $\sigma_{i,t-1}$ from the previous iteration instead. Our qualitative results regarding contagion of sovereign risk remained unchanged.

neglected, we account for the possibility of changing volatility-regimes.

To set up an equation for the conditional mean, disregarding the credit event indicators in eq. (1) (they are not known yet) gives a possible estimation approach (baseline model)

$$y_{i,r} = \sum_{l=1}^{q} \alpha_{i,l}^{t} y_{i,r-l} + (\beta_i^t)' g_r + (\gamma_i^t)' cs_{i,r} + u_{i,r}^t \tag{5}$$

with $r \in [t - wl + 1, t]$. Metiu suggests a modification of eq. (5) to incorporate findings by Lumsdaine and Ng (1999). To enhance the robustness against potential model misspecifications, additional terms will be included in eq. (5). Lumsdaine and Ng (1999) show that the integration of recursive residuals obtained from the baseline model (5) and their squared values helps control for slope and mean shifts in coefficients, ignoring an MA(1) error structure, omitting regressors in the conditional mean equation as well as over-differencing the data when they are actually trend-stationary. Further, cumulated sums of recursive residuals provide means to control for past additive outliers and can further improve the robustness against the misspecifications outlined above. Including these terms, eq. (5) becomes[5]

$$y_{i,r} = \sum_{l=1}^{q} \alpha_{i,l}^{t} y_{i,r-l} + (\beta_i^t)' g_r + (\gamma_i^t)' cs_{i,r} + \kappa_{i,1} \hat{\omega}_{i,r-1} + \kappa_{i,2} \hat{\omega}_{i,r-1}^2 + \kappa_{i,3} \sum_{s=t-wl+1}^{r-1} \hat{\omega}_{i,s} + \tilde{u}_{i,r}^t \tag{6}$$

with $r \in [t - wl + 1, t]$, where the recursive residual at time $r$ is determined from recursive coefficient estimates of eq. (5) that are based on the observations $[t - wl + 1, r - 1]$:

$$\hat{\omega}_{i,r} = y_{i,r} - \left[ \sum_{l=1}^{q} \hat{\alpha}_{i,l}^{r-1} y_{i,r-l} + (\hat{\beta}_i^{r-1})' g_r + (\hat{\gamma}_i^{r-1})' cs_{i,r} \right] \tag{7}$$

The rationale behind recursive residuals lies in the fact that they have an interpretation as one-step-ahead forecast errors (Kianifard and Swallow, 1996). Integrating recursive residuals in an auxiliary regression as in eq. (6), we incorporate information that has not been used before.

It remains to define a minimum number $k$ of observations that are needed to enter the determination of the estimators in eq. (7). Recursive residuals $\hat{\omega}_{i,r}$ are then obtained for

[5] Metiu (2012) did not include the partial sum of recursive residuals. We integrate this term in eq. (6) to further increase the robustness of the specification for the conditional mean equation, i.e. to mitigate the effects of past additive outliers which are likely to distort the contemporaneous conditional variance in eq. (4).

$r \in [t - wl + k + 1, t]$. It should be noted that in order to only use information included in $\mathcal{F}_{t-1}$ for estimating the conditional mean in eq. (6), recursive residuals must not enter this equation contemporaneously as they are derived from information at time $t$ as well. Therefore the recursive residual lagged by one period gives the most current term in eq. (6). The estimation of eq. (6) is hence based on $wl - k - 1$ instead of $wl - k$ observations.

Lastly, the credit event indicator for country $i$ can be derived comparing the residuals of eq. (6) with the estimates of the one-step-ahead Value-at-Risk according to eq. (3)

$$C_{i,t} = \begin{cases} 1 & \text{for } \tilde{u}^t_{i,t} > \widehat{VaR}_{i,t|t-1} \\ 0 & \text{otherwise} \end{cases} \tag{8}$$

for $t = wl + 1, \ldots, T$.

### 2.3 Backtesting the accuracy of the Value-at-Risk model

According to Christoffersen (1998), Value-at-Risks (VaR) determined as in eq. (3) have to fulfill certain criteria of goodness to be valid out-of-sample forecasts of a given time series. First, from the selected confidence level $p$ the so-called coverage rate $\alpha = 1 - p$ follows, which gives the frequency of VaR violations from eq. (8). Selecting a confidence level of e.g. 99% should lead to one VaR violation in 100 trading days on average (unconditional coverage hypothesis, UC). Second, violations should be distributed independently over the sample (independence hypothesis, IND). Otherwise the VaR forecast would not take into account changing volatility patterns (heteroscedasticity) and would provide clusters of VaR violations instead. A third hypothesis that should be tested for is the conditional coverage hypothesis (CC) which merges the UC and IND hypothesis.

The backtesting approach proposed by Candelon et al. (2011) is a duration-based approach that makes use of generalized moment conditions on orthonormal polynomials which are associated with the distribution of the durations under the null of a valid CC hypothesis. Durations are defined as

$$d_k = t_k - t_{k-1} \qquad k = 1, \ldots, K \tag{9}$$

with $t_k$ as the point of time of the $k$-th VaR violation according to eq. (8) and $t_0 = 0$. In contrast to existing backtesting methods, the approach proposed by Candelon et al. (2011) has better power properties and it is not necessary to define a distribution under the alternative. Under the null of a valid CC hypothesis, i.e. VaR violations are independent and occur at an expected frequency of $\alpha$, durations are distributed geometrically with parameter $\alpha$. A random variable $D$ that follows a geometric distribution can in turn be characterized by moment conditions on the orthonormal Meixner polynomials

$$E\left[M_j(D,\alpha)\right] = 0 \qquad \forall j \geq 1 \tag{10}$$

where the polynomials $M_j(d,\alpha)$ are defined recursively as

$$M_{j+1}(d,\alpha) = \frac{(1-\alpha)(2j+1)+\alpha(j-d+1)}{(j+1)\sqrt{1-\alpha}} M_j(d,\alpha) - \left(\frac{j}{j+1}\right) M_{j-1}(d,\alpha) \tag{11}$$

with $M_0(d,\alpha) = 1$ and $M_{-1}(d,\alpha) = 0$. Testing the CC hypothesis is equivalent to testing the moment conditions in (10). Given a finite sample of data, the moment conditions have to be reduced to a feasible order $J$, i.e.

$$E\left[M_j(D,\alpha)\right] = 0 \qquad j = 1,\ldots,J \tag{12}$$

Candelon et al. (2011) show that for a coverage rate of $\alpha = 1\%$ and $\alpha = 5\%$, respectively, an optimal selection for $J$ is 5 and 3, respectively. Due to the asymptotic independence and unit variance of the moments in (12), which follows from the orthonormality of the Meixner polynomials, they obtain the following test statistic for the CC hypothesis

$$J_{CC}(q) = \frac{1}{K}\sum_{j=1}^{J}\left(\sum_{k=1}^{K} M_j(d_k,\alpha)\right)^2 \xrightarrow[K\to\infty]{} \chi_J^2 \tag{13}$$

The UC hypothesis reflects the fact that the expected value of each duration variable $D$ equals $1/\alpha$. Inserting this expected value in eq. (11) in the case $j = 0$ yields an expression equivalent to the UC hypothesis

$$E\left[M_1(D,\alpha)\right] = 0 \tag{14}$$

This is a special case of eq. (12) such that the test statistic for the UC hypothesis results as a special case of eq. (13)

$$J_{UC}(q) = \frac{1}{K}\left(\sum_{k=1}^{K} M_1(d_k,\alpha)\right)^2 \xrightarrow[K\to\infty]{} \chi_1^2 \tag{15}$$

Lastly, the CC hypothesis also includes the IND hypothesis as a special case when the observed VaR violation frequency $\hat{\alpha}$ is used instead of the theoretical coverage rate $\alpha$. Hence, the moment conditions equivalent to the IND hypothesis are

$$E\left[M_j(D, \hat{\alpha})\right] = 0 \qquad j = 1, \ldots, J \tag{16}$$

and the test statistic reads as follows

$$J_{IND}(q) = \frac{1}{K} \sum_{j=1}^{J} \left( \sum_{k=1}^{K} M_j(d_k, \hat{\alpha}) \right)^2 \xrightarrow[K \to \infty]{} \chi^2_{J-1} \tag{17}$$

The adjustment of the degrees of freedom for the $J_{IND}$ test results from the fact that $E[M_1(d_k, \hat{\alpha})] = 0$ (Candelon et al., 2011).

In order to take into account the usually small number of VaR violations especially with respect to a coverage rate of 1%, Candelon et al. (2011) use the Monte Carlo simulation approach suggested by Dufour (2006) to calculate small sample size-adjusted $p$-values. Here, the test statistics are calculated under the null in $M$ (e.g. 9999) iterations and the $p$-value corresponds to the relative frequency of how many times the simulated test statistic exceeds the value of the test statistic originally obtained from the data. For the UC hypothesis, we draw geometrically distributed random variables with parameter $\alpha$ under the null until their sum exceeds the number of observations for which VaR were projected, i.e. $T - wl$. For the CC and IND hypothesis, respectively, we simulate $T - wl$ independent Bernoulli-distributed random variables with parameters $\alpha$ and $\hat{\alpha}$, respectively.[6]

### 2.4 Extension to a time-varying coefficient approach

The approach used by Metiu (2012) allows time-variation in the one-step-ahead Value-at-Risk values in order to tackle the issue of volatility clustering (see eq. (3)). Given the credit event indicators determined from (8), Metiu estimates the model (1) and tests for contagion from country $j$ (Greece, Ireland, Italy, Portugal, Spain) to country $i$ only once. His results can only be valid, however, if there is no time-variation within the coefficients of model (1). Yet, for the set of country-specific as well as global factors, Bernoth and Erdogan (2012)

[6] Candelon et al. (2011) also propose a remedy to estimation uncertainty using subsampling simulation methods to derive $p$-values of the test statistics. For the case relevant in our application ($\alpha = 1\%$, $J = 5$), however, size distortions of the $J_{CC}$ test increase when applying this simulation method. Therefore, we do not further follow the subsampling approach in this paper.

show instability regarding the significance and the value of corresponding coefficients over time, especially since the onset of the global financial crisis in 2007. Additionally, it is unlikely that the effect of sovereign risk contagion between two countries remains the same over the entire sample period (see Section 4 for empirical evidence).

Hence, we extent the approach by Metiu (2012) to allow for time-varying coefficients in eq. (1). To this end, we use rolling windows of a fixed length $wm$ and estimate eq. (1) in each window separately. The use of rolling windows allows for transitions of coefficient values without the need to pre-specify a process for the path of coefficients. Further, in every rolling window only such information is used that is observable until the end of the window. Therefore, at each point of time the contagion effects observed in the most recent sub-period can be extracted. Our approach can hence be used for prediction purposes and may serve as an early-warning-system. As a nice side-effect of the determination of credit events by one-step-ahead VaR conditional on $\mathcal{F}_{t-1}$, it suffices to derive credit event indicators only once before applying the rolling windows to eq. (1). The computation of the credit event indicators is the most costly operation in terms of computation time such that their estimation in each window would lead to a high overall time-consumption of the computations. Instead, in each window we only need to perform the 2SLS estimation of eq. (1).

Alternative methods such as the application of the Kalman filter (Aßmann and Boysen-Hogrefe, 2009) or a non-parametric kernel estimation approach (Bernoth and Erdogan, 2012) would also allow for time-varying coefficient values but have certain drawbacks. Kalman filtering is based on a pre-defined assumption regarding the process of state variables. The kernel estimation approach uses future observations to estimate coefficients at a single point of time which precludes its use for prediction purposes. For these reasons, we decided to use the rolling window approach.

As mentioned in the introduction, Leschinski and Bertram (2013) also apply the Pesaran and Pick (2007) model in rolling windows along the sample. We now state modeling issues which are likely to be the reason for the large differences of their results in comparison to ours. First, separately for each window, Leschinski and Bertram set the credit event indicator $C_{i,t}$ to 1 if the spread change at time $t$ exceeds the 80% quantile of the empirical distribution function in the respective window. This setting implies a frequency of credit events of one-in-five-days which is likely to average out substantial effects occurring after a one-in-100-days credit event as in Metiu (2012) and our work. Further, it reintroduces the issue of volatility clusters that was already resolved by Metiu (2012) and it leads to an inconsistent definition of credit events as they now depend on the specific window. Second, for non-crisis countries Leschinski and Bertram apply OLS instead of the GIVE procedure which implicitly assumes the absence of feedback effects from non-crisis to crisis countries. Moreover, the distinction between crisis- and non-crisis countries is rather subjective (selection bias). Third, they apply $F$-tests for the (joint) significance of contagion parameters $\delta_{i,j}$ but neglect an examination of the sign of these parameters. Negative shocks are only transmitted in the case of positive contagion parameters, whereas the $F$-test also assumes high values for significantly negative parameters. Fourth, the authors do not infer the explanatory variables for bond yield spreads from the existing literature on spread determinants, especially they do not consider premiums for country-specific credit risk and liquidity risk (see Section 3).

## 3 Observation period, variables and hypotheses

*Observation period*

We use data from Bloomberg (bid-ask-spreads of sovereign bonds) and Datastream (remaining variables) for the period 3 January 2005 until 31 December 2012. Hence, our dataset captures the global financial crisis of the years 2007 and 2008, and the European sovereign debt crisis of the years 2009 till 2012.[7]

Compared to the literature in which the distinction between pre-crisis and crisis period (possibly with different crisis sub-periods) is pre-defined exogenously by the authors, we determine structural breaks in the relation between yield spreads and their fundamentals as well as in shock transmission mechanism endogenously. Hence, we account for the dynamic structure of contagion effects such that our inference does not depend on the pre-definition of sub-periods.

*Variables and hypothesis*

We use time series of government bond yields at a constant maturity of 10 years and derive

[7]For Greece, we consider eq. (1) only until its quasi-default, i.e. when Greek debt was restructured under the second bailout program. This program was agreed upon in a meeting of the Eurogroup and the IMF on 20/21 February 2012. Greek credit-default swaps became obsolete since then which is reflected in flat CDS spreads since the beginning of March 2012.

yield spreads for Belgium, France, Greece, Ireland, Italy, Portugal and Spain in relation to German yields. The countries selected are the largest debtors and the most affected Eurozone countries in the recent crisis, respectively.

Our aim at this point is to correctly specify eq. (1) such that those determinants of sovereign bond yield spreads are included in the set of country-specific and global risk factors which are generally accepted in the literature. The size of yield spreads is typically explained by premiums that are related to credit risk, liquidity risk as well as the general risk aversion of market participants (see e.g. Bernoth and Erdogan, 2012, Favero et al., 2010, Gerlach et al., 2010, Maltritz, 2012, von Hagen et al., 2011).

The budget balance / GDP ratio or the debt / GDP ratio of a specific country are common measures for credit risk. The higher the debt ratio and the budget deficit, respectively, the larger is the probability that the country's debt reaches an unsustainable level and the more likely is the country's failure to pay back its loans. However, the use of these macroeconomic variables may be distorted in cases when it is difficult to distinguish between private and public debt (Barrios et al., 2009). In the recent financial crisis, governments issued guarantees for the debt of private banks (e.g. in mid-2007, Germany for the IKB and several Landesbanken; in September 2008, Ireland for its six largest banks; in October 2008, France and Belgium for Dexia). Agents on financial markets are likely to consider such information in their judgement on a country's creditworthiness.[8] Another reason for the non-suitability of macroeconomic variables for the use in this study is given by the frequency of these data. Macroeconomic variables are at best available at a monthly frequency. The assessment of contagion, however, requires data of a higher frequency to single out the effects of a rapid change in bond yield spreads. Therefore we decide to use credit default swap (CDS) spreads which have beneficial properties in measuring credit risk at a high frequency as opposed to other indicators (see e.g. Blanco et al., 2005, Longstaff et al., 2005).[9] Sovereign CDS provide insurance protection against the default of a country. The higher the CDS spread, which equals the price of the insurance, the more likely contractors consider the default of the underlying sovereign debt. Barrios et al. (2009) report a very high correlation between sovereign bond yield spreads and CDS spreads. As this may be the result of the endogeneity of CDS spreads in this case, we use lagged instead

[8] Karmann and Maltritz (2012) provide a general discussion of pros and cons of using market data instead of rating or (macro)economic data.

[9] Results of a robustness check that incorporates debt / GDP ratios are given in Section 4.

of contemporaneous values. We use CDS with a maturity of 5 years as they are the most liquid on the CDS market.

Additionally, we consider country-individual stock market returns which serve as a market-based proxy for the economic outlook of an entire economy (as in Afonso et al., 2012, Metiu, 2012). The smaller the returns, the weaker are the economic perspectives. A weak economic growth implies less tax revenues of the state, cet. par., and more difficulties to pay back its loans. Hence, a higher interest for a country's debt is likely in this case.

Liquidity risk is priced into sovereign yield spreads in the case of debt markets that do not offer a sufficient volume of buy and sell orders or exhibit sensitivity of bond prices on large-scale transactions (Barrios et al., 2009). This is reflected in wide bid-ask spreads of bond prices, which are commonly considered as the best measure for liquidity risk in the literature (Bernoth and Pick, 2012). We take country-individual bid-ask spreads of sovereign bonds from Bloomberg where the relevant time series is derived from the 10 year government bond most recently issued. To account for potential endogeneity, we use lagged bid-ask spreads.

To capture the general risk aversion of market participants, we use log-differenced values of the VSTOXX index (see e.g. Beber et al., 2009, Metiu, 2012). The VSTOXX belongs to the class of indices which measure the implied volatility of options on a given stock market and can be perceived as a measure of the market expectation of risk. In the case of VSTOXX, the underlying index is the EURO STOXX 50 which covers stocks from Eurozone countries only. Another prominent example is the VIX index which reflects the implied volatility of S&P 500 index options. Since our analysis focuses on yield spreads of Eurozone sovereign bonds, we decide to use the VSTOXX index. In Section 4, we provide results of a robustness check using the VIX index and another commonly used measure for global risk aversion instead of the VSTOXX index.

To be in line with the derivation of the dependent variable, we use spreads for the country-individual determinants over the respective German values (Codogno et al., 2003). This approach is used in the majority of research on bond spread determinants (Maltritz, 2012). As a robustness check, we will also calculate the model with explaining variables not being differenced with respect to Germany (see Section 4).

Appendix A contains descriptive statistics for the four country-specific variables, i.e. bond yield spreads, CDS spreads, bid-ask spreads and stock market returns in relation to German values, and for the log-differenced values of the VSTOXX index. For each country-specific variable, the largest absolute mean value is attained by Greece, the first country that was bailed-out and the only one with a haircut of its debt till the end of 2012. Second and third largest values for each variable are observable for Ireland and Portugal, the two other countries among the seven considered that were bailed-out till the end of 2012. There are relatively similar mean and maximum values for the corresponding Spanish and Italian variables which rank fourth and fifth. Smallest values are attained by France and Belgium. From these descriptive considerations, a fundamental link between explanatory country-specific variables and bond yield spreads already becomes apparent. In Section 4, we first assess these linkages as well as the contagion channels in a static way for the whole sample and then consider time-varying relationships.

## 4 Empirical results

*Levels vs. first differences*

In the majority of contributions, bond yield spreads are assumed to be stationary and the persistence of spread series is tackled by including lags of the dependent variable (see the first term in eq. (1)). However, some authors consider the possibility of spurious regression when eq. (1) is estimated in levels of bond yield spreads and use first differences of the yield spreads instead (e.g. Pesaran and Pick, 2007, Barrios et al., 2009). In the following, we test for unit roots in the bond yield spreads and contribute a justification of the use of levels of bond yield spreads.

The global financial crisis of 2007/2008 and the European sovereign debt crisis are events which are likely to have caused structural breaks in the bond yield series. Standard unit root tests such as the augmented Dickey-Fuller (ADF) test do not have power against trend-stationarity under structural breaks (Perron, 2006). Indeed, the ADF test does not reject the null of non-stationarity for all seven series of bond yield spreads at any conventional level. This result does not change in the case when a GLS-correction of the ADF test is included as suggested by Elliot et al. (1996).[10]

Therefore we also consider the unit root test proposed by Lee and Strazicich (2003). This test has power against trend-stationary alternatives with structural breaks and is derived

[10]Results for unit-root tests are available from the authors upon request.

with two breaks under both the null and the alternative.[11] We select the model with endogenous breaks in the level and trend and augment the test to take into account serial correlation of errors. Results confirm the trend-stationarity of all series at conventional levels. As a consequence, we do not difference bond yield spreads and continue our analysis with their levels as in eq. (1).[12]

*Determination of credit events and backtesting of Value-at-Risk*

As outlined in Section 2.2, a credit event takes place at time $t$ if the innovation $\tilde{u}_{i,t}^{t}$ in eq. (6) exceeds an ex-ante determined one-step-ahead Value-at-Risk which is based on the fundamental risk-pricing of yield spreads conditional on the information set $\mathcal{F}_{t-1}$. The Value-at-Risk is obtained from the conditional distribution of innovations at each point of time. To estimate the one-step-ahead VaR, we follow Metiu (2012) and select an in-sample window length of $wl = 500$ trading days. The first window thus ranges from 1 January 2005 till 1 December 2006 and gives a VaR for the 4 December 2006. The window thus includes only observations from a period well before the rise of sovereign bond yields. We then gradually include information of trading days closer to or within the crisis period and exclude information of past trading days. In sum, we obtain Value-at-Risks and credit event indicators $C_{i,t}$ for each trading day in the interval 4 December 2006 till 31 December 2012. This leads to a sample size of $T = 1586$ to estimate eq. (1).

To parameterize the Value-at-Risk according to eq. (3), we still have to define the degrees of freedom of the $t$-distribution of innovations and the confidence level $p$. As credit events we model VaR violations which happen in $\alpha = 1\%$ of all trading days. The confidence level thus equals 99%. Applying the backtesting methods described in Section 2.3, we find that the number of degrees of freedom should be optimally set to $\nu = 8$ as this leads to the acceptance of all three quality criteria for the calibration of Value-at-Risks (see Table 1).[13] Table 1 shows that there are no significant deviations from the assumed coverage rate of 1% and there is no spurious clustering of credit events which would result from

[11] If the null hypothesis of the test did not include structural breaks, the test might reject the null just because of the fact that there are structural breaks in the process. A (non-)rejection of the null by the test provides indeed evidence for (non-)stationarity instead.

[12] Moreover, a comparison of the results of the ADF test with the results of the test suggested by Lee and Strazicich (2003) confirms that there are structural breaks in the spread series. Thus, the appearance of time-varying coefficients in eq. (1) is likely. This underlines the need to apply an approach that allows time-variation in these coefficients.

[13] Using smaller (larger) values for $\nu$, the one-step-ahead Value-at-Risks become larger (smaller) which is more (less) conservative and leads to a decrease (increase) of the number of VaR violations.

Table 1: Backtesting of the Value-at-Risk parameterization

| Country | VaR violations | $\hat{\alpha}$ | $J_{UC}$ | $p$-value | $J_{CC}$ | $p$-value | $J_{IND}$ | $p$-value |
|---|---|---|---|---|---|---|---|---|
| Belgium | 14 | 0.0088 | 0.0905 | 0.7460 | 1.3573 | 0.6632 | 0.7081 | 0.8626 |
| France | 11 | 0.0069 | 0.2717 | 0.5887 | 1.6680 | 0.5797 | 0.3997 | 0.9504 |
| Greece | 18 | 0.0113 | 1.5057 | 0.2226 | 1.6423 | 0.5857 | 0.4942 | 0.9320 |
| Ireland | 14 | 0.0088 | 0.1580 | 0.6792 | 2.5834 | 0.3938 | 0.7292 | 0.8592 |
| Italy | 12 | 0.0076 | 0.3267 | 0.5589 | 2.4494 | 0.4169 | 0.9709 | 0.7622 |
| Portugal | 17 | 0.0107 | 0.7873 | 0.3640 | 1.4154 | 0.6476 | 0.5868 | 0.9035 |
| Spain | 17 | 0.0107 | 0.4725 | 0.4820 | 1.0939 | 0.7437 | 0.5490 | 0.9187 |

Note: the table contains the number of credit events, the empirical coverage rate $\hat{\alpha}$, the values of the test statistics to test for the unconditional (UC) and the conditional coverage (CC) and the independence (IND) hypothesis as described in Section 2.3. $p$-values are determined using the approach of Dufour (2006) and $M = 9999$ simulations.

mis-specified Value-at-Risks.

Table 3 in Appendix B shows the distribution of credit events for each country over time. The sum over all countries is the largest in 2008Q4 when 11 credit events took place. This is not surprising as a recession hit almost all European economies after the collapse of Lehman Brothers in September 2008. Credit events most often hit Greece (18), Portugal and Spain (each 17), whereas France (11) and Italy (12) were hit the least often which is in line with these countries' crisis exposure.

Further, we observe that the majority of credit events in Greece and Spain took place till mid-2010. For Greece, besides the recession starting in 2008 as a consequence of the dependence on pro-cyclical industries such as tourism and shipping, structural deficits and alarming debt and deficit levels were revealed at the beginning of 2010. In Spain, the global financial crisis led to the burst of the property bubble in 2007/2008.

The number of Irish credit events peaked in 2008Q4. Here, the burst of a property bubble caused major problems of Irish banks forcing the Irish state to guarantee bank deposits and bonds of the six main Irish banks in September 2008. For Portugal, credit events accumulated after the beginning of the recession in 2008Q4. Credit events became more concentrated for France since mid-2010. For Belgium and Italy, there is no concentration of credit events between 2007 and 2012.

*Whole sample coefficient estimates*

First, we estimate eq. (1) for the whole sample available, i.e. for the period 4 December 2006 till 31 December 2012, using 2SLS. It remains to define the number $q$ of lagged dependent variables in eq. (1). We choose $q = 5$ as this ensures the absence of autocorrelation and takes into account the within-week variation in trading patterns as well (Metiu, 2012). We use the same number of lags to instrument the credit event indicators in eq. (2) and include second and third powers of the instruments to improve their strength to approximate the credit event indicators.

Table 4 in Appendix B reports whole sample coefficient estimates for eq. (1) as well as robust standard errors. General risk aversion contributes most significantly to the bond yield spreads of each country (significance at the 1% level for all countries), whereas CDS spreads seem to not have explanatory power for Portugal, Ireland and Greece. Our time-varying coefficient approach, however, will reveal that this insignificance is likely due to the assumption of constant parameters in the static model.

Liquidity risk does not seem to be a relevant component in the risk-pricing of these Euro-zone bonds. Yet, we will later show that there are sub-periods in which the bid-ask spread is indeed significant. Coefficients with respect to stock returns are significant for Spain and Italy. All significant coefficients have the expected signs.

The last seven columns of Table 4 show estimates of the contagion coefficients $\delta_{i,j}$. The transmission of negative shocks, i.e. pure contagion, only takes place if this coefficient is positive. Results are similar to those obtained by Metiu (2012) who applies the canonical contagion model for the period 1 January 2008 till 1 February 2012. Contagion effects can be summarized as follows:

(i) Spain does not cause any contagion to other countries (the significant coefficient in the equation of Belgium is negative)

(ii) Italy influences France, Ireland and Belgium

(iii) France seems to have an impact on Greece (what we will show to be an unreliable result in the following) and Belgium

(iv) Portuguese shocks are transmitted to Greece (with a very large value of the coefficient) and Ireland

(v) Greece only influences Portugal and Belgium but none of the other countries (what will we show to be a result of neglecting time-variation in the contagion coefficients)

(vi) Belgian shocks hit the largest countries Spain, Italy and France

(vii) Ireland influences Italy, Portugal and Belgium

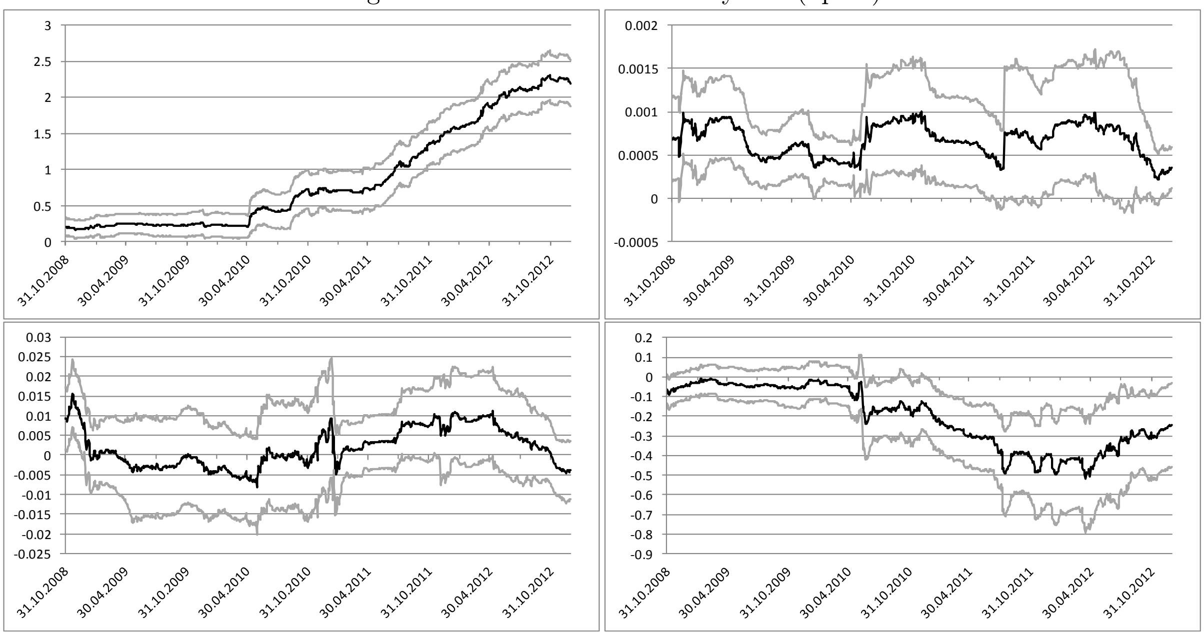
Yet, as argued by Bernoth and Erdogan (2012), an altering of risk pricing over time causes time-variation in the coefficients of fundamentals of bond yield spreads. We already indicated that several results stated above are not supported anymore when time-variation in contagion coefficients is considered as well. In the following, we present the main results of this paper.

*Time-varying coefficient estimates*

We now estimate eq. (1) in rolling windows of $wm = 500$ trading days. Compared to smaller window length, this selection guarantees that at least one credit event per country lies in each window. Further experiments showed that larger windows already average over sub-periods of significant contagion with sub-periods of no contagion.

Figures 1-7 display the time-varying coefficients of country-specific explanatory variables as well as of the VSTOXX index to proxy the general risk aversion of market participants. The graphs give the estimated coefficient values at the end date of each window as well as their 90% confidence intervals. It is easily seen that there are significant changes of parameter values over time as e.g. the coefficients related to VSTOXX in more recent years do not fall into the confidence interval belonging to points of time further away.

Figure 1: Determinants of bond yields (Spain)




Note: the panels in Figures 1-7 contain time-varying coefficients and 90% confidence intervals for the log-differenced VSTOXX index, the lagged CDS spreads, the lagged bid-ask spreads of a benchmark bond and the returns of the local stock market index (from top left to bottom right)

Figure 2: Determinants of bond yields (Italy)

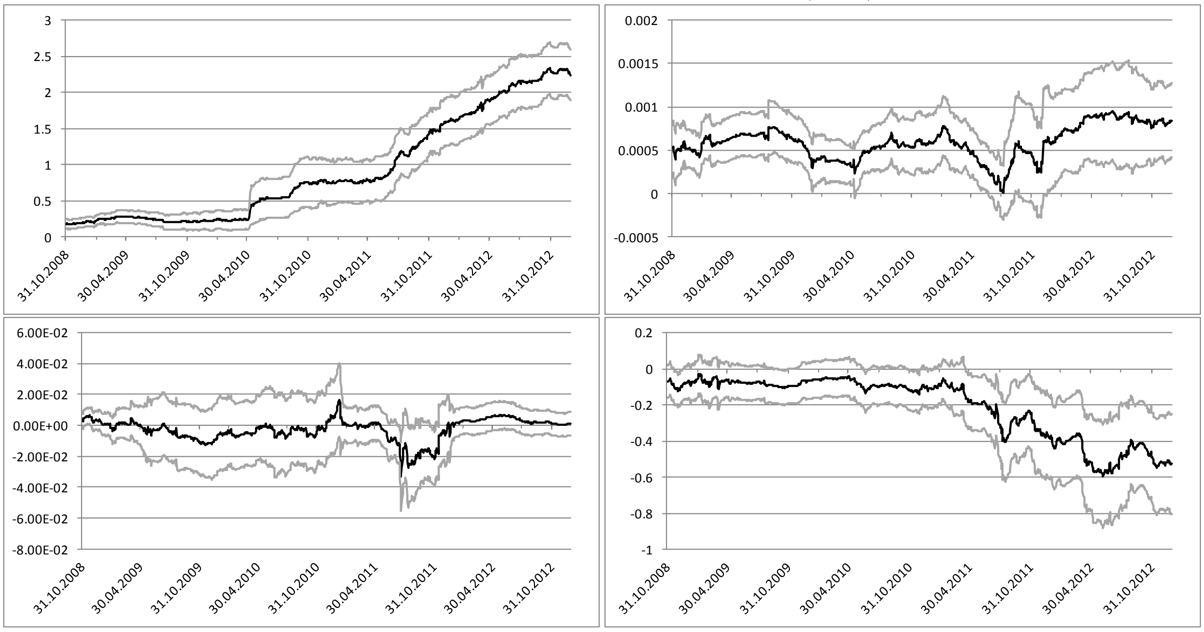
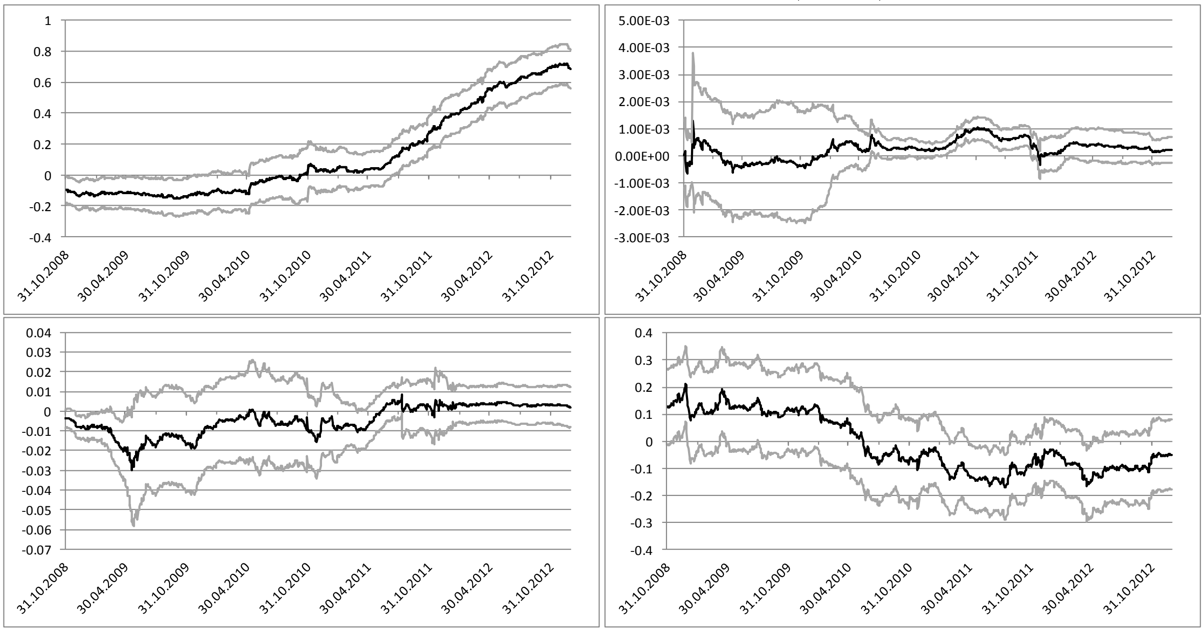



Figure 3: Determinants of bond yields (France)

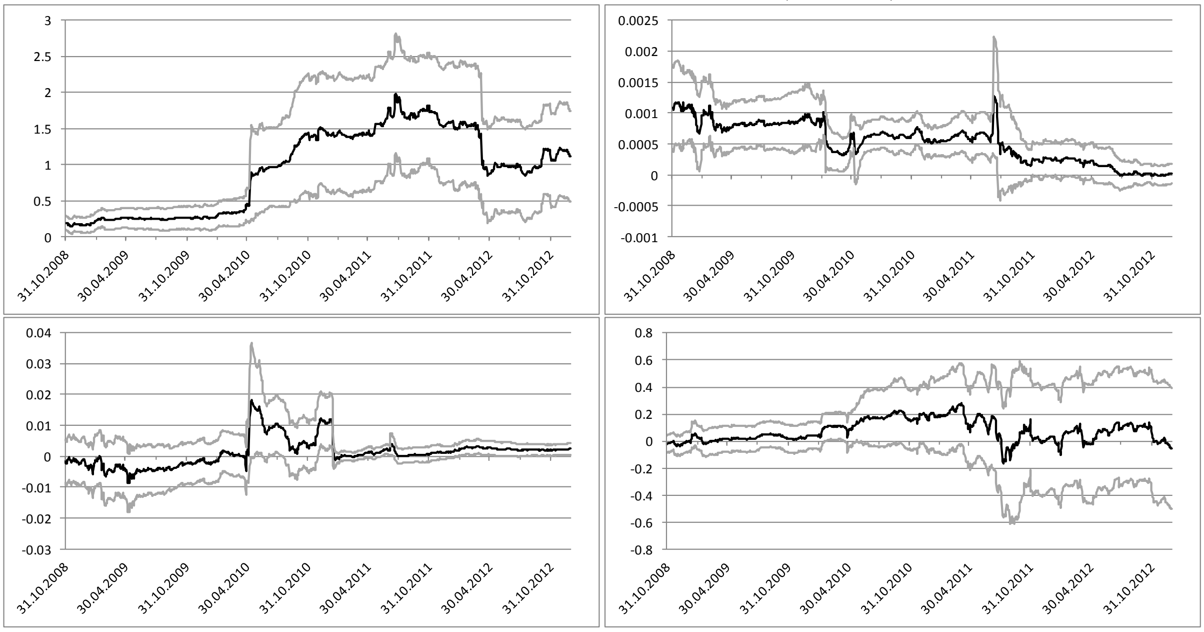



We find that the general risk aversion of market participants has an increasing influence on the risk pricing of bonds of all seven countries considered.[14] There is a large rise in the VSTOXX coefficient value for all countries since May 2010 which reflects rising premiums for general risk aversion ("wake-up-call contagion").[15] The timing of this development is

[14]Bernoth and Erdogan (2012) provide a similar finding using a spread series of BBB-rated US corporate bonds.

[15]This reasoning is possible since the log-differenced series of the VSTOXX index is stationary and does not contain a significant trend.

Figure 4: Determinants of bond yields (Portugal)



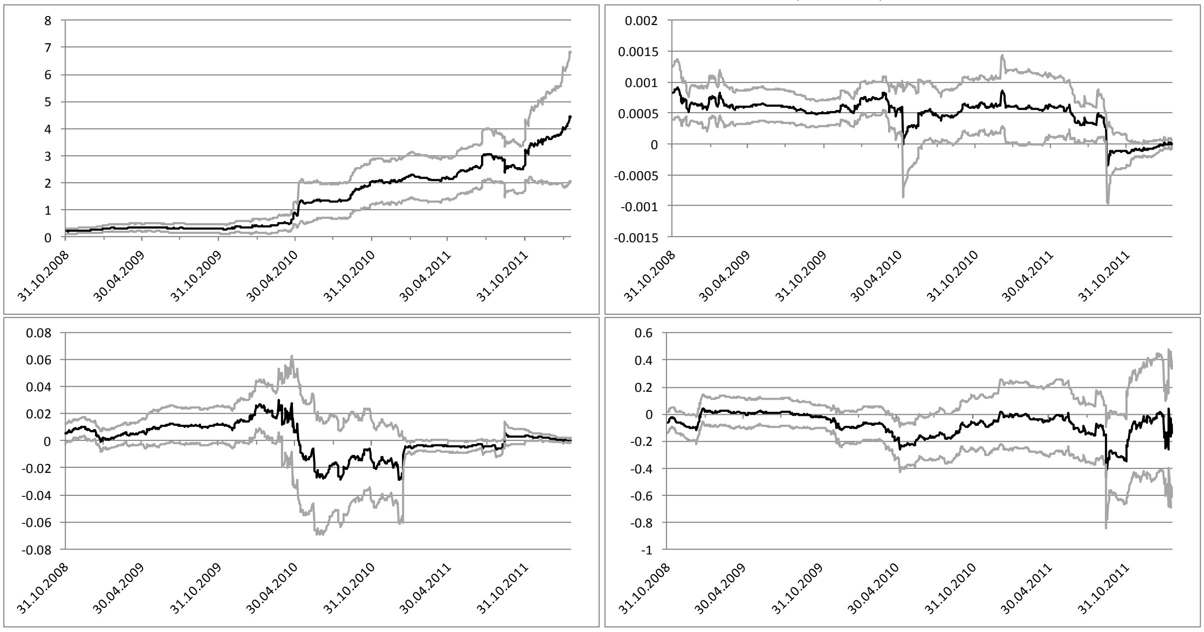

likely to result from the general distrust towards the Eurozone after uncovering the true economic and fiscal malaise in Greece at the beginning of 2010 and the fear of pro-cyclical austerity measures that could further foster negative developments in the entire Eurozone.

These premiums, however, are heterogeneous among countries as coefficient values differ significantly.[16] For Greece, we observe the largest premiums for general risk aversion and the smallest for France and Belgium. This corresponds to the general perception of the solvability of these countries. As opposed to the other countries considered in this analysis, the premium for general risk aversion has decreased for Portugal and Ireland after March 2012. For most countries, the width of confidence intervals remains relatively stable which underlines the robustness of the rising relevance of the risk premium for general risk aversion.

Country-specific credit risk measured by CDS spreads in relation to Germany was (almost) always significant for Spain, Italy and Ireland. For Greece, Portugal and Belgium this factor is only statistically significant till mid-2011 and is likely to have been substituted by the increased risk aversion in the Eurozone as a whole. For Ireland, a decreasing trend in the coefficients combined with decreasing CDS spreads led to a shrinking credit risk premium since mid-2011. CDS spreads of France are only significant for the period

[16] As we use the same "global factor", VSTOXX, in all country equations, we can directly infer a ranking of the relevance of this factor from individual coefficient values.

Figure 5: Determinants of bond yields (Greece)



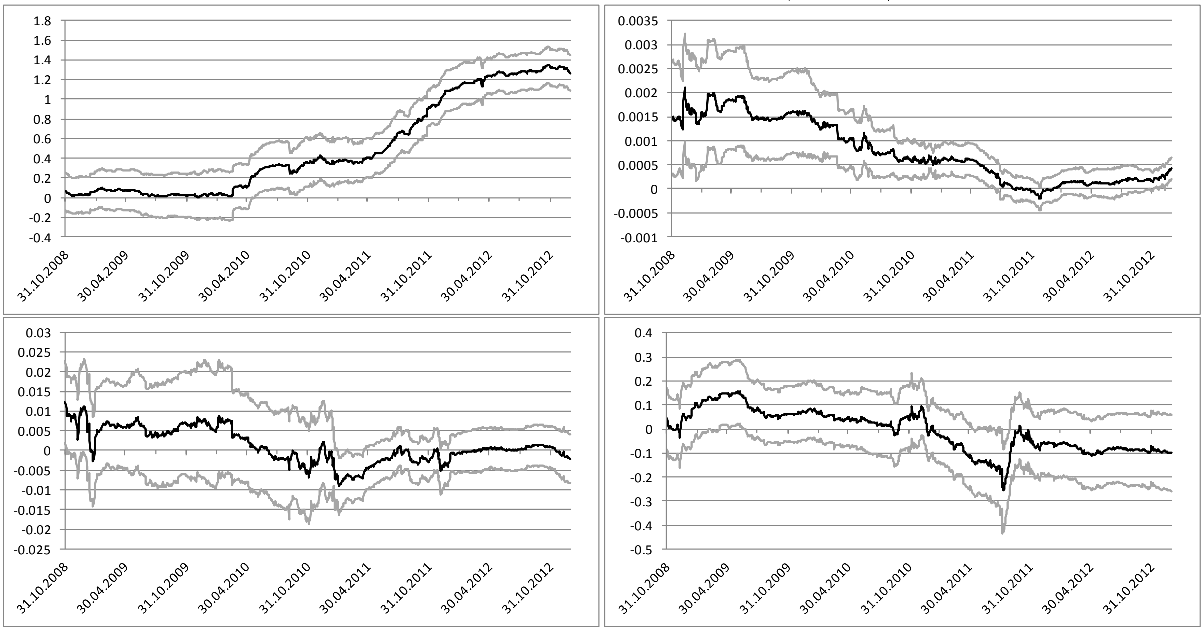

between December 2010 and October 2011.[17] These results are in contrast to the results for France, Portugal, Ireland and Greece in the constant coefficient approach applied before.

Bid-ask spreads were mostly insignificant for all countries what confirms the results of Bernoth and Erdogan (2012) with respect to liquidity risk. However, Greek bonds seem to have contained a liquidity premium in the first quarter of 2010 (Figure 5, lower left panel), a period of growing concerns about the economic and fiscal situation and the solvability of this country. Even more apparent is the relevance of the liquidity risk premium for Ireland in the period December 2008 till May 2010. This period began after the Irish government expressed guarantees for the six largest banks of the country and covers the (partial) nationalizations of three of these banks.

For Portugal, a significant liquidity premium can be observed for mid-2010, December 2010 as well as from February 2012 till the sample end (Figure 4, lower left panel). Moody's cut of Portugal's sovereign bond rating in the summer of 2010 reflected increasing concerns about Portugal's ability to cope with its high state debt and structural deficits and is likely to be the reason for the first peaks of the bid-ask spread coefficient in 2010. The last sub-period that exhibits a significant indicator of liquidity risk is likely to be connected with the downgrade of Portugal to a non-investment grade by S&P in January 2012.

[17]Barrios et al. (2009) also report that French CDS spreads do not have significant explanatory power for bond yields in their sample that covers the period until April 2009.

Figure 6: Determinants of bond yields (Belgium)



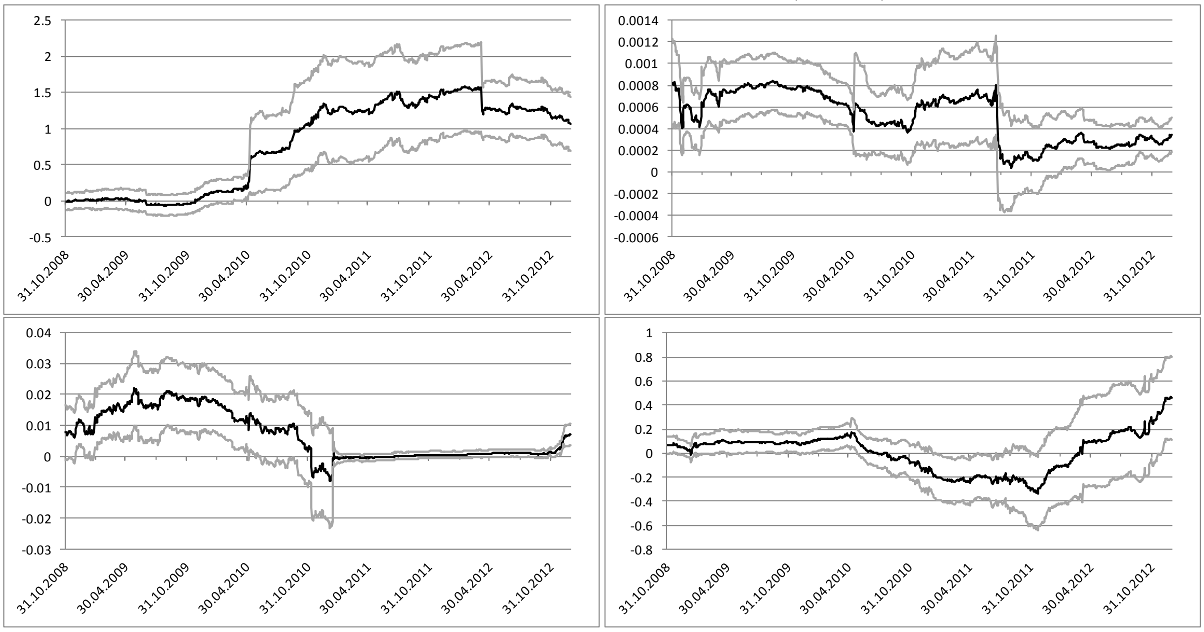

Our results are in line with the findings of Favero et al. (2010) who show that liquidity premiums are only significant in a subset of countries, e.g. in Portugal but not in Spain, Italy and France.[18] Beber et al. (2009) conclude that a liquidity premium is only relevant during times of heightened market uncertainty what matches our results as well.

To account for prospects for the entire economy, we included country-individual stock market returns in eq. (1). The results for time-varying coefficients of stock market returns confirm the findings of the constant coefficients model to a large extent. The coefficients for lagged stock market indices in the equations for France, Portugal, Greece, Belgium and Ireland are mostly insignificant. For Spain and Italy, respectively, we find periods with significant coefficients from June 2010 and April 2011, respectively, till the sample end (see the lower right panels in Figures 1 and 2).[19] This observation shows that investors perceive the worsening sitatuation of the economies of Italy and Spain, which are the largest among the peripheral Eurozone countries, as relevant for the deepening of the current government debt crisis. On the other hand, economic growth in other peripheral countries, indicated by rising stock market indices, is likely to not lead to a significant reduction of bond yield spreads in these countries if it is not accompanied by a smaller premium for investor risk aversion. Among the fundamental factors explaining the levels of bond yield spreads, only

[18]The authors did not consider Greece and Ireland in their sample.

[19]The increase of the coefficient for VSTOXX and the decrease of the coefficient for stock market returns in the equations for Italy and Spain are independent of each other, i.e. there are no disturbing effects from multicollinearity. We obtain the same results for these variables when they are used as the only variable besides the constant and lagged dependent variables.

Figure 7: Determinants of bond yields (Ireland)

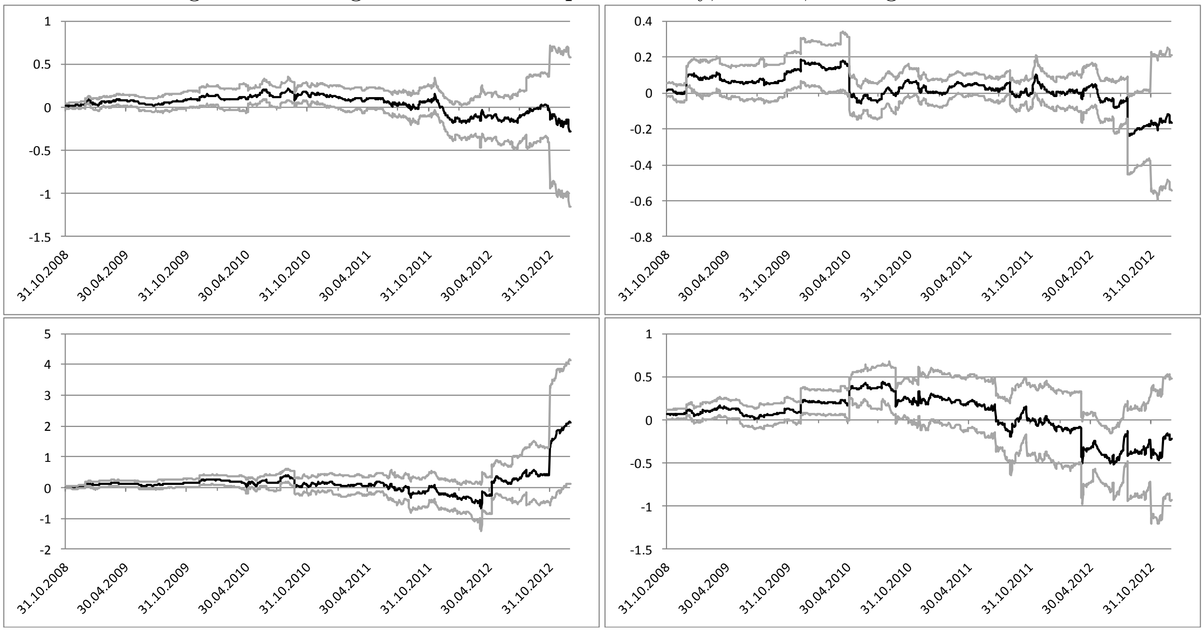



a reduction of the premium for general risk aversion can provide a systematic decrease of spreads in all Eurozone countries considered, *cet. par.*

Figures 8-14 provide evidence for the presence of pure sovereign risk contagion in the Eurozone. Moreover, they show that contagious effects are time-dependent and should hence not be modeled assuming constant contagion coefficients for the entire sample. Figures 8-14 only contain these graphs for the contagion coefficients between two countries when there is at least one sub-period with significant contagion coefficients. Remaining graphs can be obtained from the authors upon request. As in our discussion of the results of the constant coefficient approach, we state results for contagion channels from the point of view of the origination country.

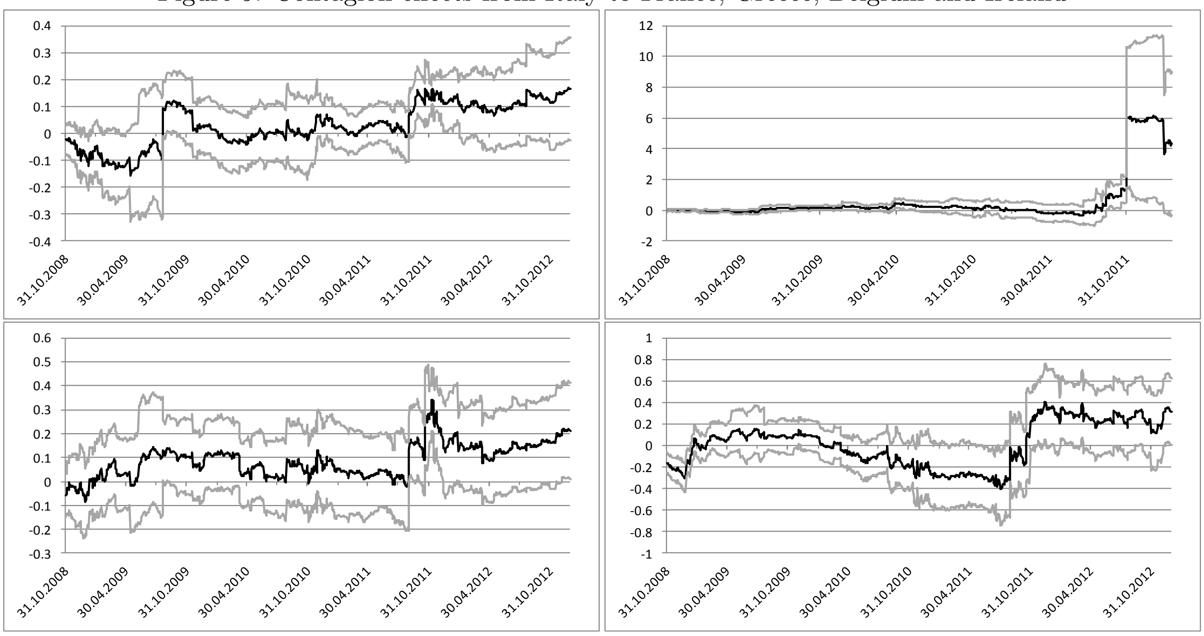
The results for contagious effects running from Spain to the other countries considered are depicted in Figure 8. In the constant coefficient approach, contagion coefficients are not significantly positive with respect to any other country. Now, we find sub-periods of significant contagion from Spain to Italy (May 2010 - January 2011), France (October 2009 - April 2010), Portugal (January 2009 - April 2010) and Ireland (December 2009 - October 2010). However, these sub-periods are relatively short in time and there is no contagion during 2011 and 2012. Moreover, there is no contagion to Greece and Belgium at all.

Figure 8: Contagion effects from Spain to Italy, France, Portugal and Ireland



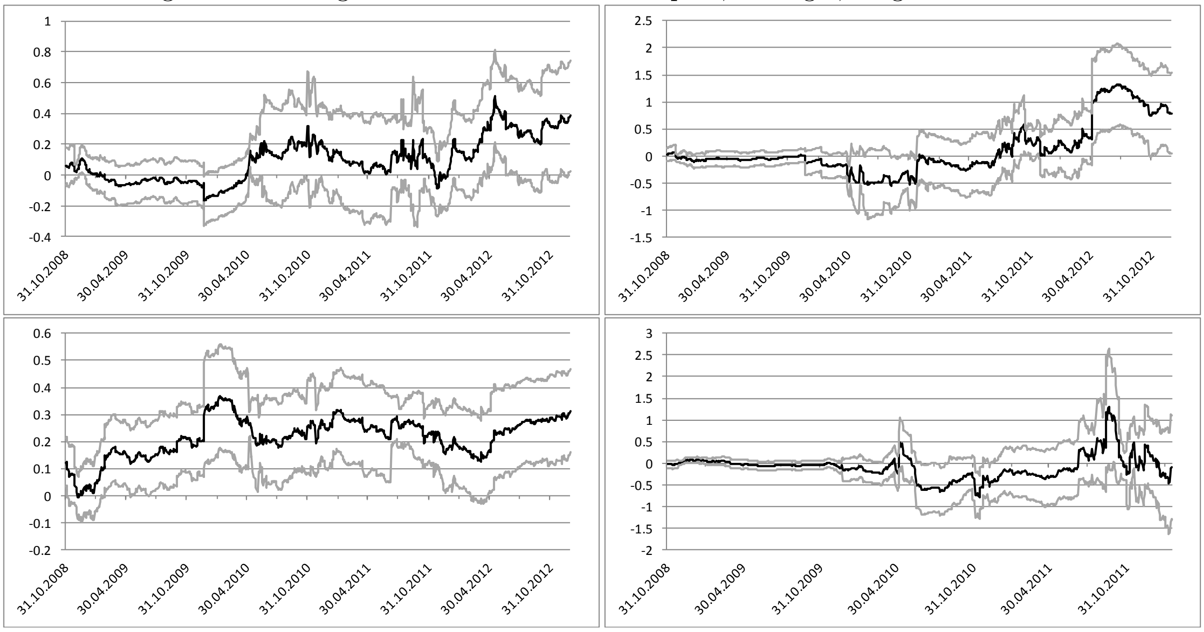

Note: the panels in Figures 8-14 contain time-varying coefficients and 90% confidence intervals for the contagion coefficients in eq. (1) for the countries mentioned in the respective figure head (from top left to bottom right)

Figure 9: Contagion effects from Italy to France, Greece, Belgium and Ireland




In 2012 Spanish banks required a bailout after incurring large losses from the burst of the property bubble (cf. the corresponding credit event in 2012/II in Table 3). Yet, solvability issues in Spain did not trigger an increase of bond yield spreads in the other Eurozone countries considered. Hence, investors either believe in the success of the recapitalization program for Spanish banks or bond-purchases of the ECB could counteract a possible

Figure 10: Contagion effects from France to Spain, Portugal, Belgium and Greece



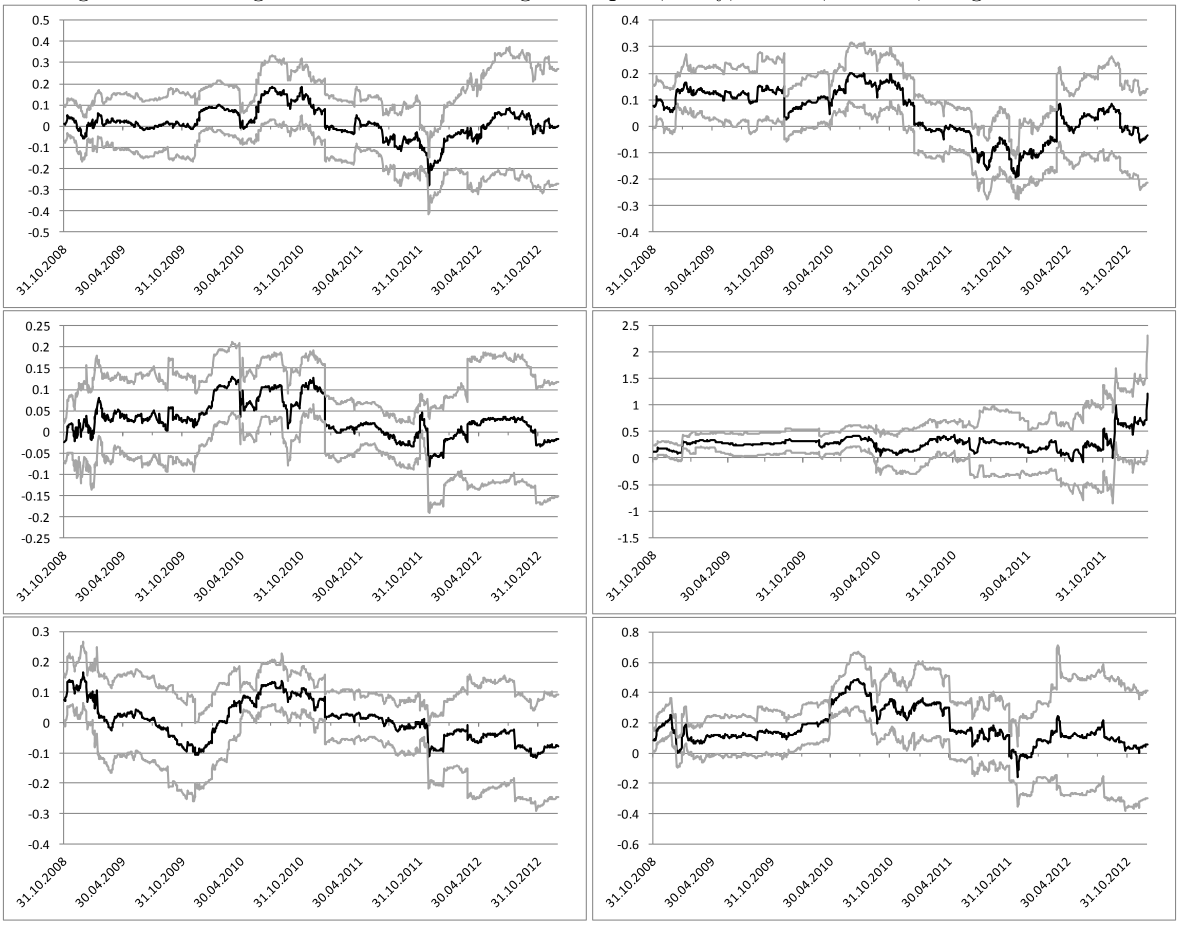

shock-transmission to other Eurozone countries.

As is the case for the other large peripheral country, Spain, contagious effects from Italy are relatively rare (see Figure 9). There is evidence for significant contagion to France (September 2011 - January 2012), Greece (September 2011 - January 2012), Belgium (September 2011 - December 2011) and Ireland (November 2011 - March 2012).[20] These periods are relatively short and began in September 2011 when the Italian government passed austerity measures to cope with increasing budget deficits and levels of state debt. Fears about a further worsening of economic prospects and political turmoil around the Berlusconi government led to increasing bond yield spreads at that time.

The relevance of shock transmissions from France to other Eurozone countries increased over time and led to significant contagion to Spain (April 2012 - July 2012) and Portugal (since May 2012), whereas shock transmission to Belgium is significant almost for the whole sample (since March 2009) (see Figure 10). The latter supposedly stems from close ties between the real and financial industries between these countries. In contrast to the

[20] As contagion to Greece, the country typically seen as the source of the Eurozone debt crisis (see e.g. Missio and Watzka, 2011), may look surprising, a comment is needed. According to our results, Portugal and Italy are the only countries from which significant shocks were transmitted to Greece. This observation is in line with the existing literature. Kalbaska and Gatkowski (2012) find evidence for the same causalities in their "crisis period" (i.e. after the credit crunch in August 2007). Among the countries considered in our analysis, Wing Fong and Wong (2012) detect contagion only from Italy to Greece. Their result that there is no contagion from Portugal to Greece may be due to differences in sample periods. We find contagion from Portugal to Greece to be most pronounced after February 2012 which is beyond the sample of Wing Fong and Wong (2012).

Figure 11: Contagion effects from Portugal to Spain, Italy, France, Greece, Belgium and Ireland

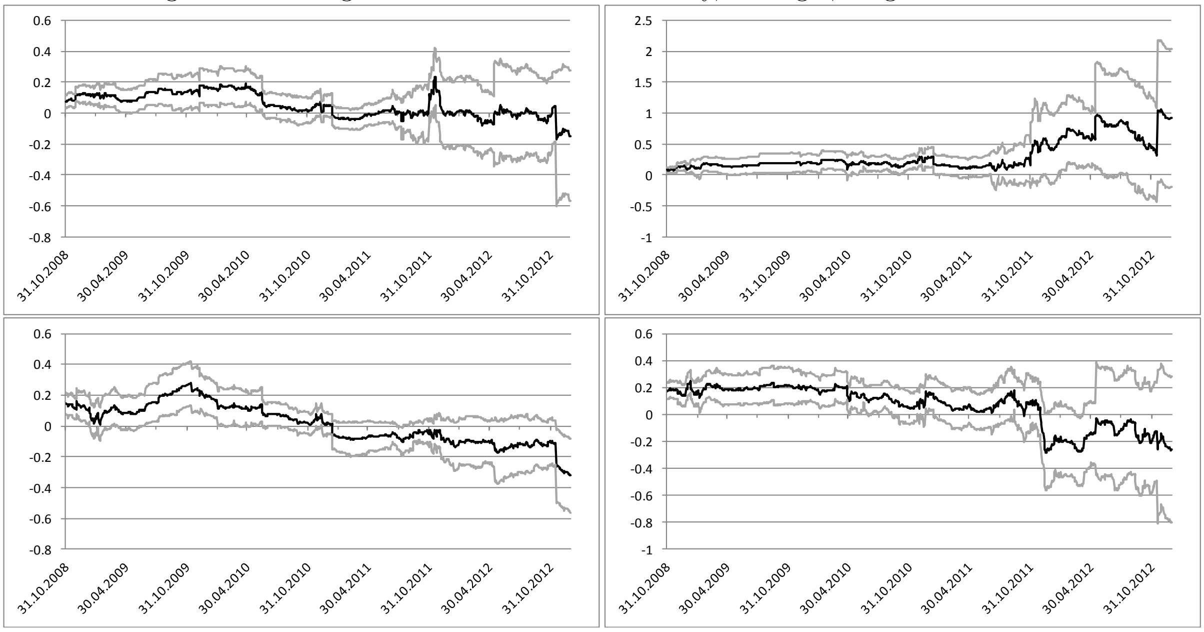



results of the model with constant coefficients, there is no shock-transmission from France to Greece in any sub-period. Instead, coefficients are negative for most windows and distinct peaks with positive values are not significant at the 10% level (see the lower right panel in Figure 10).

Figure 11 displays significant contagion of sovereign risk from Portugal to Spain (July 2010 - September 2010), Italy (January 2009 - December 2010), France (February 2010 - January 2011), Greece (January 2009 - April 2010), Belgium (May 2010 - December 2010) and Ireland (December 2009 - April 2011). Contagious effects came to an end until April 2011, when Portugal requested a bailout by the other Eurozone members and the International Monetary Fund.

For Greece, the application of the constant coefficient model yields significant contagion only towards Portugal and Belgium as documented above. In the model that allows time-varying coefficients, however, contagion can also be observed from Greece to Italy (November 2008 - June 2010) and Ireland (November 2008 - December 2010) (see Figure 12).

Figure 12: Contagion effects from Greece to Italy, Portugal, Belgium and Ireland

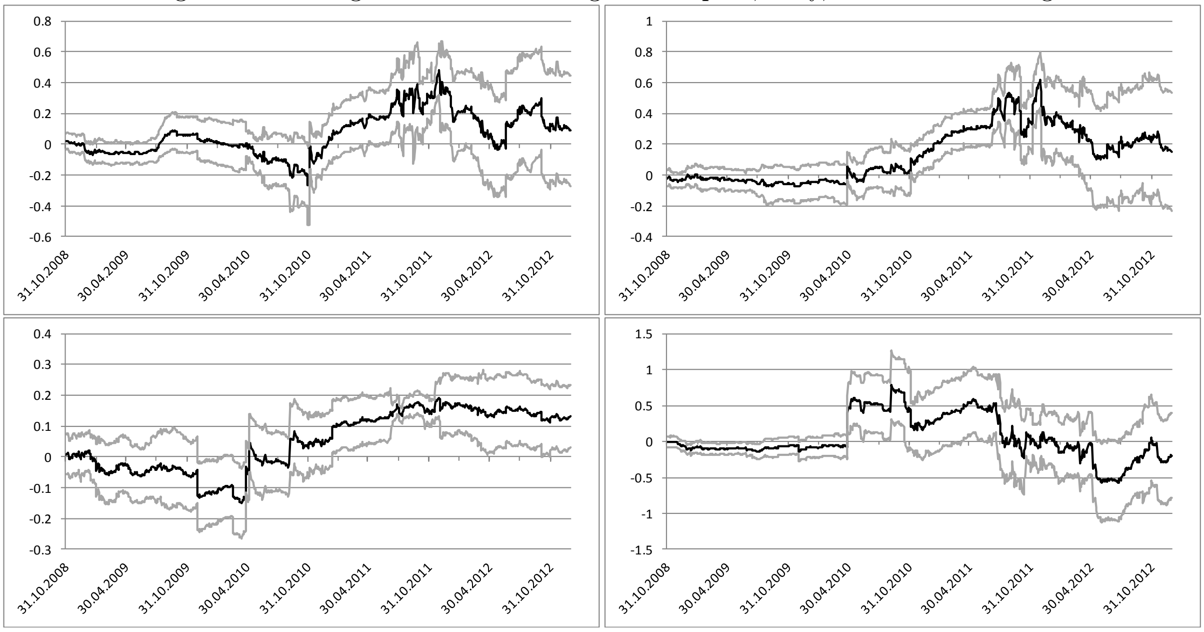



Shock transmissions to Portugal are significant from November 2008 till February 2011 as well as from January till June 2012, and to Belgium from June 2009 till August 2010. Spanish and French sovereign risk is not affected by credit events in Greece through contagion. As we could see before, their bond yield spreads are rather influenced by changes in the fundamental pricing of sovereign debt, especially with regard to the increase in the general risk aversion of investors ("wake-up-call contagion").

Sovereign risk contagion from Greece to Belgium, Italy and Ireland came to an end in mid-2010, after the first bailout program for Greece was designed in May 2010. Portugal is the only country that is still affected by Greek credit events after mid-2010. It seems that the Eurozone and IMF bailout programs for Greece and Portugal (as discussed above) led to a breakdown of contagion relations from Portugal and Greece to other Eurozone countries after their establishment.

The role of Belgium as an exporter of sovereign risk begins in mid-2010 when contagion coefficients started to rise (see Figure 13). Significant contagion effects are observable in relation to Spain (June 2011 - December 2011), Italy (November 2010 - March 2012), France (since January 2011) and Portugal (April 2010 - October 2010). The negative impact on France corresponds to the long-lasting significance of contagion effects from France to Belgium (see Figure 10 and the corresponding discussion). Financial challenges such as the high state debt / GDP ratio coincided with an absence of a ruling government

Figure 13: Contagion effects from Belgium to Spain, Italy, France and Portugal




from April 2010 till December 2011 which caused growing concern among investors and is likely to explain the timing of contagion that originated in Belgium.

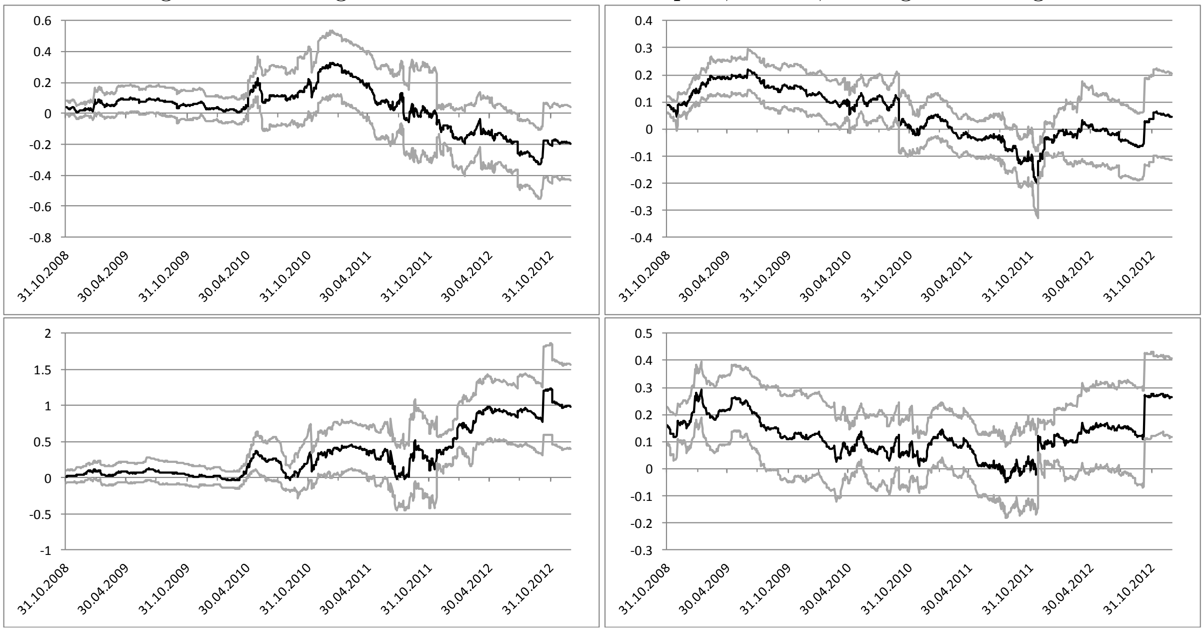
Lastly, contagion from Ireland to other Eurozone countries occurred relatively early (see Figure 14). Negative shocks from credit events were transmitted to France and Belgium from November 2008 till September 2010 and August 2009, respectively. Significant contagion took also place towards Spain (December 2010 - March 2011) and Portugal (December 2010 - April 2011 and since November 2011). As it was already the case for Portugal and Greece, contagion from Ireland to the sovereigns not being bailed-out ceased after the initiation of the rescue plan for the respective country. The Irish bailout program was set in power in November 2010. Although further credit events took place in Greece, Ireland and Portugal after their bailouts (see Table 3), negative shocks were not transmitted to the other countries anymore. This result allows the conclusion that the bailout programs for these countries successfully disconnected shock-transmission channels to healthier Eurozone countries.

*Robustness checks*

To assess the robustness of our results, we run different sensitivity checks. We summarize the results here and provide details upon request:

(i) As discussed by Maltritz (2012), the literature on fundamentals of bond yield spreads can be divided into two strands: in the first strand, fundamentals of bond yield

Figure 14: Contagion effects from Ireland to Spain, France, Portugal and Belgium




spreads are modeled in relation to a reference country (which is typically Germany or the United States); in the second strand, values are not differenced. As described in Section 3, the first approach is used in this paper. Applying the second approach does not lead to significant changes in the results for contagion coefficients. Among the fundamentals, results for the CDS spreads and bid-ask spreads of benchmark bonds remain the same, whereas the significant coefficients for Italian and Spanish stock market returns vanish.

(ii) Investor risk aversion was found to be an important explaining variable for bond yield spreads. Exchanging the log-differenced VSTOXX values for log-differenced VIX values (as used by e.g. Gerlach, 2010) or the yield spreads of BBB-rated US corporate bonds over benchmark US government bonds provided by Merrill Lynch (as used quite frequently, e.g. by Bernoth and Erdogan, 2012, von Hagen et al., 2011, Maltritz, 2012), does not change the results for the contagion effects observed in this paper. However, coefficients of the VIX index or BBB-rated corporate yield spreads are either insignificant or even negative which is exemplified in Figures 15 and 16. This finding creates doubt regarding the usefulness of these US based indicators to proxy investor risk aversion in the Eurozone. Arezki et al. (2011) support this conclusion in that they find that US stock market performance is unrelated with the Eurozone debt crisis. Beirne and Fratzscher (2013) do not find significant coefficients for the VIX index among their determinants for bond yield and CDS spreads either.

(iii) We also included quarterly GDP growth rates to account directly for the country-specific macroeconomic environment but did not find any changes in the number and location of significant contagion relations in time.

Figure 15: Coefficients of BBB-rated US corporate bond yield spreads in the equations for Italy and Ireland

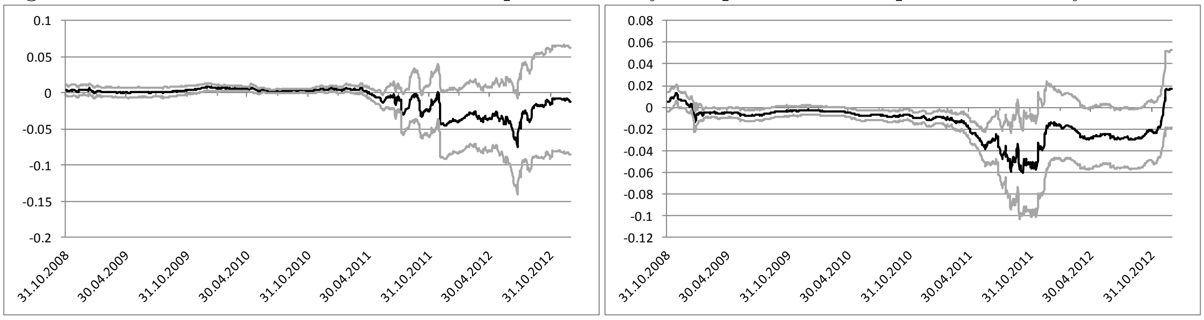
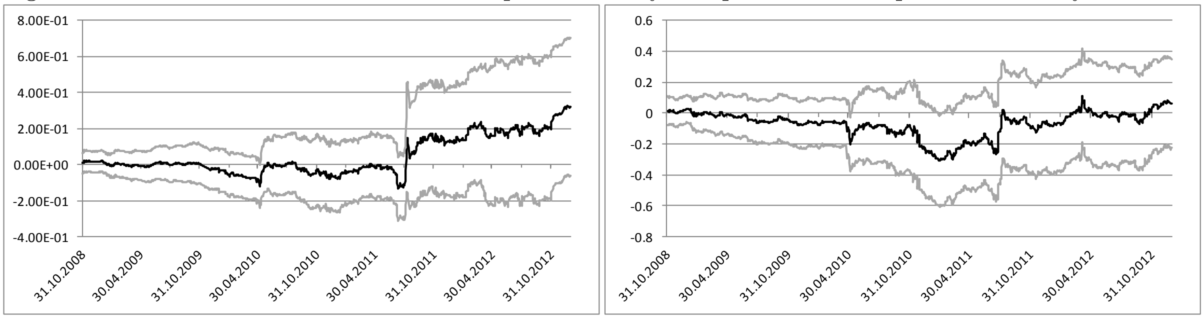



Figure 16: Coefficients of BBB-rated US corporate bond yield spreads in the equations for Italy and Ireland




# 5 Conclusions

This paper presented an extension of the canonical model for contagion by Pesaran and Pick (2007) and Metiu (2012) to assess the contagion of sovereign risk in the Eurozone. Contagion of sovereign risk is a necessary element to understand the development of bond yield spreads of Eurozone sovereigns. Contagion is examined by testing the significance of coefficients that represent shock-transmission after a credit event in a single country to another country. Credit are defined to take place on trading days with significant deviations of the bond yield spread from the current risk pricing of market participants through global and country-specific factors. To infer their location in time, we use the approach proposed by Metiu (2012) and increase its robustness with respect to past additive outliers.

The paper highlighted the necessity to allow for time-varying contagion coefficients in this model. Otherwise, certain contagion relations would not appear at all as they would be averaged out with these sub-periods when contagion did not take place. Moreover, contagion of sovereign risk is unlikely to happen in the same intensity over the whole sample period as it may have been intercepted by bailout programs. Therefore modeling contagion should allow for the time-variation of contagion coefficients. For this reason, we extended the modeling approach proposed by Pesaran and Pick (2007) and applied their basis equation in rolling windows of a fixed length. Thanks to this modification, the approach may serve more easily as an early-warning system to monitor possible contagion effects from crisis countries.

To be able to single out contagion effects, we controlled the evolution of yield spreads of sovereign bonds of seven Eurozone countries by country-specific indicators and an indicator for general risk aversion. We found an increasing relevance of general risk aversion towards Eurozone countries since May 2010 what partially replaced the country-specific credit risk factor. Moreover, liquidity risk seems to have played a significant role in the yield spreads of Ireland, Greece and Portugal in times of heightened uncertainty regarding the solvability of these countries. Lastly, as opposed to the smaller countries Portugal, Ireland and Greece, the (real) economic development in Spain and Italy has a significant impact on the development of bond yield spreads - even in the case of unchanged general risk aversion of market participants towards the entire Eurozone.

Our analysis confirms the presence of sovereign risk contagion in the Eurozone. The existence of contagion, however, is time-dependent. We find that contagion of sovereign risk originating in the three countries mostly hit by the crisis (Greece, Ireland and Portugal) to the other four countries considered (France, Italy, Spain and Belgium) terminated after the introduction of bailout programs for these countries. Hence, the policy measures taken by the Eurozone member states, the European Commission and the International Monetary Fund could successfully contain credit events within these three countries after their bailouts.

Spain is the least affected country by contagion from other Eurozone countries. Increasing Spanish bond yield spreads result mostly from homemade issues such as the financial distress of its banking system. As a consequence of culminating refinancing problems of Spanish banks in the first half of 2012, Spanish sovereign bond yields saw all-time highs. We showed that the corresponding credit event, however, did not cause contagion to the other Eurozone countries. Hence, investors either believe in the success of the recapitalization program for Spanish banks or bond-purchases of the ECB could counteract a possible shock-transmission to other Eurozone countries.

France and Belgium shared shock-transmission channels in both directions for a relatively long period. The relevance of France as a shock-exporting country rose during 2012. At the end of 2012, France caused significant contagion to Portugal and Belgium. Italy was responsible for contagion only in smaller sub-periods at the end of 2011 but was affected by contagion from Portugal and Greece for relatively long periods. Our analysis shows that among the three largest countries in our sample - France, Italy and Spain - during 2012 developments in France have received the greatest attention among market participants.

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# Appendix A: Descriptive statistics

Table 2: Descriptive statistics

| Variable | Country | Mean value | Median | Minimum | Maximum | Std. Dev. |
|---|---|---|---|---|---|---|
| Bond yield spreads in 100bp in relation to Germany | Belgium | 0.56 | 0.31 | -0.12 | 3.63 | 0.68 |
| | France | 0.28 | 0.14 | -0.12 | 2.04 | 0.40 |
| | Greece | 6.05 | 1.00 | 0.07 | 55.71 | 9.34 |
| | Ireland | 1.99 | 0.98 | -0.21 | 9.23 | 2.4 |
| | Italy | 1.23 | 0.55 | 0.04 | 5.75 | 1.46 |
| | Portugal | 2.59 | 0.45 | -0.10 | 14.59 | 3.69 |
| | Spain | 1.20 | 0.38 | -0.15 | 6.35 | 1.59 |
| CDS spreads in bp in relation to Germany | Belgium | 47.66 | 17.2 | -3.50 | 284.07 | 61.39 |
| | France | 22.58 | 3.40 | -4.50 | 137.91 | 33.51 |
| | Greece | 4542.48 | 97.60 | 1.40 | 37007.60 | 11224.2 |
| | Ireland | 193.39 | 103.88 | -4.20 | 1188.47 | 244.68 |
| | Italy | 103.4 | 51.79 | 0.50 | 490.82 | 124.93 |
| | Portugal | 242.18 | 34.26 | -4.10 | 1512.67 | 364.13 |
| | Spain | 108.54 | 45.14 | -3.33 | 553.00 | 133.51 |
| Stock market returns in relation to Germany | Belgium | -0.0070 | -0.0054 | -0.1319 | 0.124 | 0.0318 |
| | France | -0.0059 | -0.0056 | -0.0885 | 0.079 | 0.0206 |
| | Greece | -0.0164 | -0.0136 | -0.3138 | 0.3179 | 0.0774 |
| | Ireland | -0.0104 | -0.0089 | -0.176 | 0.1397 | 0.042 |
| | Italy | -0.0107 | -0.0104 | -0.128 | 0.1514 | 0.0323 |
| | Portugal | -0.0083 | -0.0064 | -0.1441 | 0.1408 | 0.0414 |
| | Spain | -0.0061 | -0.0052 | -0.1511 | 0.1534 | 0.0397 |
| Bid-ask spreads in bp in relation to Germany | Belgium | 0.78 | 0.40 | -1.10 | 9.80 | 1.47 |
| | France | 0.33 | 0.10 | -0.80 | 5.30 | 0.60 |
| | Greece | 37.41 | 0.80 | -1.10 | 708.00 | 107.22 |
| | Ireland | 9.42 | 0.20 | -1.00 | 58.3 | 16.54 |
| | Italy | 1.14 | 0.90 | 0.00 | 15.00 | 1.03 |
| | Portugal | 13.57 | 1.00 | -1.00 | 157.3 | 26.09 |
| | Spain | 1.15 | 0.30 | -1.00 | 12.3 | 2.19 |
| VSTOXX | | 0.0001 | -0.0023 | -0.1082 | 0.1423 | 0.0257 |

# Appendix B: Tables 3 and 4

Table 3: Timing of credit events

| Year | 2006 | 2007 | | | | 2008 | | | | 2009 | | | | 2010 | | | | 2011 | | | | 2012 | | | | Sum |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Quarter | IV | I | II | III | IV | I | II | III | IV | I | II | III | IV | I | II | III | IV | I | II | III | IV | I | II | III | IV | |
| Belgium | 0 | 0 | 0 | 2 | 0 | 1 | 1 | 0 | 2 | 0 | 0 | 0 | 1 | 0 | 2 | 0 | 1 | 0 | 0 | 2 | 2 | 0 | 0 | 0 | 0 | 14 |
| France | 0 | 0 | 0 | 0 | 0 | 1 | 1 | 0 | 1 | 0 | 0 | 0 | 1 | 0 | 2 | 0 | 1 | 0 | 0 | 2 | 2 | 0 | 0 | 0 | 0 | 11 |
| Greece | 0 | 1 | 0 | 3 | 0 | 2 | 0 | 1 | 2 | 2 | 0 | 0 | 2 | 1 | 1 | 0 | 1 | 0 | 0 | 1 | 1 | 0 | 0 | 0 | 0 | 18 |
| Ireland | 0 | 1 | 0 | 1 | 0 | 0 | 0 | 1 | 3 | 0 | 0 | 0 | 1 | 0 | 2 | 1 | 1 | 0 | 0 | 0 | 2 | 0 | 0 | 0 | 1 | 14 |
| Italy | 0 | 0 | 1 | 1 | 0 | 0 | 1 | 1 | 1 | 0 | 0 | 1 | 1 | 0 | 0 | 0 | 1 | 0 | 0 | 1 | 1 | 1 | 1 | 0 | 0 | 12 |
| Portugal | 0 | 0 | 0 | 1 | 1 | 0 | 0 | 1 | 1 | 2 | 0 | 0 | 2 | 0 | 2 | 1 | 1 | 0 | 2 | 0 | 2 | 1 | 0 | 0 | 0 | 17 |
| Spain | 1 | 2 | 1 | 1 | 1 | 1 | 1 | 2 | 1 | 0 | 0 | 0 | 1 | 0 | 1 | 2 | 1 | 0 | 0 | 0 | 0 | 0 | 1 | 0 | 0 | 17 |
| **Sum** | 1 | 4 | 2 | 9 | 2 | 5 | 4 | 6 | 11 | 4 | 0 | 1 | 9 | 1 | 10 | 4 | 7 | 0 | 2 | 6 | 10 | 2 | 2 | 0 | 1 | 103 |

Note: the table provides the numbers of credit events derived from one-steap-ahead Value-at-Risk violations in each quarter of the sample period considered.

Table 4: Whole sample coefficient estimates of the canonical contagion model

| Country | VSTOXX | CDS Spreads | Bid-Ask Spreads | Stock returns | Spain | Italy | France | Portugal | Greece | Belgium | Ireland |
|---|---|---|---|---|---|---|---|---|---|---|---|
| Spain | 1.0297*** [0.0991] | 0.0003*** [0.0001] | -0.0005 [0.0032] | -0.1887*** [0.0672] | | 0.0501 [0.1600] | -0.0587 [0.1788] | 0.0953 [0.1624] | -0.2206 [0.1377] | 0.4551** [0.1803] | -0.1464 [0.1239] |
| Italy | 1.0420*** [0.0993] | 0.0004*** [0.0001] | -0.0018 [0.0040] | -0.2478*** [0.0675] | -0.1268 [0.1395] | | -0.0311 [0.1750] | 0.0096 [0.1071] | -0.0245 [0.1579] | 0.4478** [0.1797] | 0.2366* [0.1263] |
| France | 0.1895*** [0.0467] | 0.0002** [0.0001] | 0.0014 [0.0035] | -0.0557 [0.0486] | -0.0472 [0.0729] | 0.2770*** [0.0967] | | 0.0439 [0.0953] | 0.0751 [0.0723] | 0.1970** [0.0774] | 0.1536 [0.0939] |
| Portugal | 0.8635*** [0.1953] | 0.0001 [0.0001] | 0.0007 [0.0007] | 0.0484 [0.1110] | -0.0446 [0.2595] | -0.2959 [0.3716] | 0.4601 [0.3562] | | 0.7628*** [0.2553] | -0.3663 [0.2948] | 0.8860*** [0.2861] |
| Greece | 2.8347*** [1.0626] | 0.0000 [0.0000] | -0.0022 [0.0020] | 0.2654 [0.3906] | 1.4729 [1.0216] | 4.1924 [2.7037] | 1.6347** [0.7565] | 2.6370** [1.0978] | | -0.8945 [1.2051] | -1.6352* [0.8375] |
| Belgium | 0.5475*** [0.0751] | 0.0003*** [0.0001] | -0.0001 [0.0024] | -0.0426 [0.0495] | -0.2292** [0.1070] | 0.2251* [0.1225] | 0.4128*** [0.1131] | 0.0480 [0.1010] | 0.2126** [0.0980] | | 0.2544** [0.1024] |
| Ireland | 0.7419*** [0.1506] | 0.0000 [0.0001] | 0.0002 [0.0004] | 0.0236 [0.0542] | 0.1754 [0.1665] | 0.4644*** [0.1767] | 0.2474 [0.1567] | 0.4374** [0.2085] | -0.1406 [0.1540] | -0.1581 [0.1679] | |

Note: the table contains coefficient values of eq. (1) and robust standard errors in brackets. The constant and coefficients of lagged dependent variables have been spared to save space. ***, ** and * denote significance at the 1%, 5% and 10% level, respectively.