Systemic regime switching in euro area government bond yield spreads: the impact of the financial crisis*

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Abstract

We investigate the presence of euro area-wide crisis regime periods in the risk pricing of euro area government bonds. We investigate how often and when the crisis regime occurs, how long it lasts, and how it affects the importance of the components that make up the risk premiums of euro area government bonds (i.e. the mean, the country-specific risk factor, and the common euro area-wide risk factor). To this end, a dynamic factor model with Markov switching parameters is estimated using monthly data for the 10 year government bond yield spreads of five euro area countries (Belgium, France, Italy, the Netherlands, and Spain) versus Germany over the period 1999 until 2012. We identify a single permanent regime shift in the risk components of the yield spreads during the first half of 2008, i.e. before the Lehman default (September 2008) and well before the outbreak of the government debt crisis in the euro area periphery. Following the regime shift, the impact on the spreads of both the country-specific risk factor and the area-wide risk factor is significantly higher in all countries considered. While all countries experienced qualitatively similar changes in the risk pricing of their bonds, the magnitude of the changes was different across countries and was most extreme for Italy and Spain.

JEL Classification: E43, G12

Keywords: Government bond yield spreads, financial crisis, euro area, state space, Markov switching

1 Introduction

In the run-up to the start of the European Monetary Union (EMU) in 1999 and during the years after the introduction of the euro, the 10 year government bond yield spreads of EMU member states versus Germany had declined significantly. By the end of 2006, they were quite small for many member states

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suggesting that the monetary union had increased government bond market integration in the euro area. It should be noted that full convergence was never entirely attained and positive spreads remained until 2007-2008 after which they increased exponentially. Many studies written before the onset of the financial crisis explored the reasons for the convergence and the sources of the lack of full convergence (Codogno et al., 2003; Bernoth et al., 2004; Gomez-Puig, 2008). Their findings suggest that default risk, defined as the probability that a country will not be able to fulfill its financial obligations or will do so only partially, liquidity differentials among issuing countries, and changes in general investors' risk aversion - usually captured by a common international risk factor - were significant in explaining these positive spreads. While spreads remained positive in the period between the introduction of the euro and the start of the crisis, their magnitude was low. During this period investors started to grant the same risk status to countries in the periphery of the euro area (Spain, Portugal, Greece, Ireland) as they did to countries belonging to the core of the euro area (Germany, France, the Netherlands). As such, euro area countries from the periphery were able to borrow at interest rates much lower than those available during the pre-EMU period. After the outbreak of the financial crisis in 2007-2008, the spreads of most euro area countries versus Germany started to rise again, and excessively so in Greece, Ireland, Italy, Portugal, and Spain ("GIIPS" countries). This reflected the higher risk premiums demanded by investors both in response to the worsening global economic climate and in response to country-specific evolutions like deteriorating government debts and deficits and problems in the banking sector (Favero et al., 2010; Bernoth and Erdogan, 2011; von Hagen et al., 2011).

The financial crisis has initiated a substantial literature that explores the channels through which financial turmoil affects euro area government bond spreads. Overall, most studies agree that investors' sensitivity with respect to country-specific factors increased considerably after the crisis and that there was also a sharp increase in international or euro area-wide risk (Haugh et al., 2009; Barrios et al., 2009; Attinasi et al., 2010). Haugh et al. (2009) also find evidence of interaction effects and non-linearities between domestic variables - captured by different measures of governments' indebtedness - and the international risk factor. Sgherri and Zoli (2009) do not use a proxy for international risk but filter out a time-varying measure for this factor. They find that sovereign default risk premiums mainly mirror global risk repricing, which is influenced by shifts in cyclical conditions and instability in financial markets. Allowing for the coefficients on the determinants of spreads to be time-varying, several studies confirm the importance of both country-specific macroeconomic fundamentals and changes in the international risk factor in post-crisis spreads (Aßmann and Boysen-Hogrefe, 2009; Bernoth and Erdogan, 2011; Mody, 2009). Other works explore the presence of flight to quality effects after the crisis (Borgy et al., 2011; von Hagen et al., 2011). They find evidence that countries such as France or the Netherlands benefited from markets' stronger sensitivity to fiscal variables during the crisis. By noting that the recent financial crisis is centered on the banking sector, Gerlach et al. (2010) find that when international risk increases, countries with large banking sectors are more severely penalized by market participants who then demand higher returns. Additionally, Attinasi et al. (2010) find that the announcement of banks rescue packages leads to an increase in spreads signalling the transfer of risk from the private financial sector to the public sector. Schwarz (2010) presents a novel approach by introducing newly constructed measures for liquidity and default risk based on microstructure datasets. She finds that liquidity risk explains much more of spreads movements than credit risk after the financial crisis. Dotz and Fisher (2011) offer support to country-specific factors as determinants of spreads. They find a significant role of the effect of financial market soundness and international competitiveness in determining spreads after the crisis. Afonso et al. (2011) explore the effect that rating announcements by credit agencies have on spreads. Both before and after the financial crisis they find a clear and quick reaction of euro area spreads to credit rating events with a stronger response to downgrade announcements. Exploring contagion effects during the crisis, some studies find that an increase in financial distress in a particular country can propagate to other relatively safer ones and that there is a non-linear effect of country-specific fiscal fundamentals on spreads which depends on other countries' fundamentals (Favero and Missale, 2012; Favero, 2012). Lastly, Maltritz (2012) deals with model uncertainty by conducting a Bayesian model averaging approach to test the robustness of a set of determinants of spreads included in previous works. He confirms previous results and finds that budget deficits or changes in debt are highly significant determinants of spreads during the financial crisis.

This paper adds to the literature by investigating the presence of euro area-wide crisis regime periods in the yield spreads of euro area government bonds (versus Germany) over the period after the introduction of the euro in January 1999 until April 2012. We investigate how often and when the crisis regime occurs, how long crisis episodes last, and to what extent and through which channels crisis episodes affect the pricing of government bond risk in euro area countries. We determine whether the impact of the crisis regime is qualitatively and quantitatively similar across countries or whether there are important qualitative or quantitative differences between countries. To answer these questions we estimate a regime switching dynamic factor model for government bond yield spreads for five euro area countries (Belgium, France, Italy, the Netherlands, and Spain) versus Germany. We decompose each spread into a country-specific intercept, a country-specific risk factor, and a common euro area-wide international risk factor. The country-specific intercept and the country-specific factor loadings on both factors are regime-dependent. We allow the latent regime variable to follow a first-order two-regime Markov switching process. We then use a multivariate linear state space approach with Markov regime switching (see Kim and Nelson (1999), chapter 5) to estimate the latent factor decomposition of the yield spreads and to endogenously determine in which of two possible regimes - a tranquil regime or a crisis regime - the system operates. Our empirical set-up has a number of advantages compared to what is done in the literature on government bond yield spreads. First, the unobserved country-specific risk premiums and the common international risk factor need not be obtained through conditioning variables or proxy's which imperfectly capture these premiums. Rather, a stochastic process is assumed for these factors after which they are filtered out of the spreads data with the Kalman filter. Second, the Markov switching approach used to model the factor loadings allows, *a priori*, for multiple endogenously determined shifts between the tranquil regime and the crisis regime.[1] Third, the combination of a decomposition of the spreads together with the possibility of shifts in the relative importance of the components of the spreads is both an informative and a theoretically compelling methodological approach. It is informative because it allows for an identification of the sources of the changes in spreads. It is theoretically compelling because the framework fits nicely into the contagion literature, i.e. the potential shift caused by the crisis in the country-specific impact of the

[1] Note that the Markov switching methodology assumes that, at least a priori, the crisis regime is re-occurring. Labeling the crisis regime according to specific historic crisis episodes (e.g., naming it "the euro debt crisis") makes no sense as these specific events are not re-occurring. With respect to a specific crisis episode, it also implies that, a priori, the crisis episodes should be thought of as an occurrence of the crisis regime or a succession of occurrences of the crisis regime.

country-specific risk factor captures "wake up call" contagion, the potential shift caused by the crisis in the country-specific impact of the common risk factor captures "shift" contagion, while the potential shift caused by the crisis in the country-specific intercept captures "pure" contagion (see e.g. Giordano et al. (2013) for this terminology).[2]

While our methodological approach allows for multiple regime shifts, our estimates identify only one shift, i.e. a permanent regime shift in the pricing of risk during the first half of 2008 which lasts until the end of the sample period. The shift occurs before the announcement of the Lehman default and well before the outbreak of the government debt crisis in the euro area periphery. It therefore seems that even before the announcement of the Lehman default euro area government bond market investors had already incorporated a high probability that the relatively tranquil regime had come to an end. Note that Acharya and Steffen (2015) observe a widening of yield spreads of a number of GIIPS countries as early as 2008 and attribute this to flight-to-quality of bank investors as they believed that rising sovereign yields threatened the solvency of European Banks. This massive unwinding of positions is explained by banks large exposures to the risk of divergence among Eurozone countries through their carry trade operations. Before the crisis banks were investing in risky long-term sovereign debt of the Eurozone periphery and financing themselves with short-term wholesale funding. As of 2008, this short-term funding dried up and banks were left holding large positions in GIIPS sovereign debt with rapidly rising yields. Following the regime shift, the average spreads and the factor loadings on both the idiosyncratic country-specific risk factor and the common euro area-wide risk factor are significantly higher in all countries considered. While all countries experience qualitatively similar changes in the pricing of their bonds, the magnitude of the changes is different across countries and is most extreme for Italy and Spain. The results imply that increases in the spreads are caused both by an increase in pure country-specific risk (i.e. both "pure" and "wake up call" contagion) and by an increase in the country-specific premium for euro area-wide international risk. The latter is the product of the common euro area-wide risk factor and the country-specific impact of this risk factor, both of which are shown to increase during the crisis regime. The higher exposure to common risk implies that the financial crisis has made the euro area countries considered more vulnerable to "shift" or "international" contagion. Additionally, the increased importance of both

[2]See Bekaert et al. (2012) for the same distinction of different types of contagion with a somewhat different terminology.

purely country-specific risk and common risk in the yield spreads makes the impact of the financial crisis on the degree of government bond market integration ambiguous a priori. Using a simple measure of country-specific market integration we find for instance that during the crisis regime Spain has become more and the Netherlands has become less integrated with the remaining countries.

The outline of the paper is as follows. Section 2 describes the empirical specification and the estimation method. A description of the data is also given. The results are presented in section 3. Section 4 concludes.

2 Empirical specification and methodology

2.1 A regime switching dynamic factor model for government bond yield spreads

We assume that the spread R_{it} of country i (where $i = 1, ..., N$) in period t versus the benchmark country can be expressed by the following latent factor model,

$$R_{it} = \mu_{it} + \alpha_{it} R_{it}^I + \beta_{it} R_t^W + \varepsilon_{it}$$

$$= \mu_i(S_t) + \alpha_i(S_t) R_{it}^I + \beta_i(S_t) R_t^W + \varepsilon_{it} \tag{1}$$

where R_{it}^I is the country-specific or idiosyncratic risk factor and where R_t^W is the common euro area-wide international risk factor. The country-specific error term ε_{it} is added to the specification to account for measurement error. The intercept $\mu_i(S_t)$ and the factor loadings $\alpha_i(S_t)$ and $\beta_i(S_t)$ are country-specific and are dependent on the regime variable S_t. The variable S_t is assumed to follow a first-order two-state Markov switching process so that it can take on two values. In particular, we specify

$$\mu_i(S_t) = \sum_{j=1}^{2} \mu_i^j S_t^j \tag{2}$$

$$\alpha_i(S_t) = \sum_{j=1}^{2} \alpha_i^j S_t^j \tag{3}$$

$$\beta_i(S_t) = \sum_{j=1}^{2} \beta_i^j S_t^j \tag{4}$$

where $S_t^j = 1$ if $S_t = j$ and $S_t^j = 0$ otherwise and where j is the regime index which can take on the values 1 and 2 (i.e. $j = 1, 2$). The transition probabilities are given by $P(S_t = j|S_{t-1} = j) \equiv p^j$ for $j = 1, 2$. In the paper we label regime 1 as the tranquil regime and regime 2 as the crisis regime (see below for identification issues). As such, if $\alpha_i^2 > \alpha_i^1$ this constitutes "wake up call" contagion induced by the euro area-wide crisis (i.e. a higher impact on the spreads of country-specific risk during the crisis regime), if $\beta_i^2 > \beta_i^1$ this constitutes "shift" contagion (i.e. a higher impact on the spreads of common euro area-wide risk during the crisis regime), and if $\mu_i^2 > \mu_i^1$ this constitutes "pure" contagion (see Giordano et al. (2013) for the terminology).

We model the unobserved risk factors R_{it}^I and R_t^W as $AR(1)$ processes,

$$R_{it}^I = \pi_i R_{it-1}^I + \eta_{it} \tag{5}$$

$$R_t^W = \pi_w R_{t-1}^W + \eta_{wt}. \tag{6}$$

where π_i and π_w are AR parameters for which $-1 < \pi_i < 1$ and $-1 < \pi_w < 1$ and where the error terms η_{it} and η_{wt} are white noise and follow $GARCH(1,1)$ processes,

$$\eta_{it} = [h_{it}]^{\frac{1}{2}} v_{it} \tag{7}$$

$$\eta_{wt} = [h_{wt}]^{\frac{1}{2}} v_{wt} \tag{8}$$

where $v_{it} \sim i.i.d(0,1)$ and $v_{wt} \sim i.i.d(0,1)$ and where the conditional variances h_{it} and h_{wt} are given by,

$$h_{it} = V_{t-1}(\eta_{it}) = c_i + \delta_i^a \eta_{it-1}^2 + \delta_i^b h_{it-1} \tag{9}$$

$$h_{wt} = V_{t-1}(\eta_{wt}) = c_w + \delta_w^a \eta_{wt-1}^2 + \delta_w^b h_{wt-1} \tag{10}$$

with parameter restrictions $c_i > 0$, $c_w > 0$, $0 < \delta_i^a < 1$, $0 < \delta_i^b < 1$, $0 < \delta_i^a + \delta_i^b < 1$, $0 < \delta_w^a < 1$, $0 < \delta_w^b < 1$, and $0 < \delta_w^a + \delta_w^b < 1$. The unconditional variance of η_{it} is given by $\sigma_{\eta i}^2 = c_i/(1 - \delta_i^a - \delta_i^b)$ and the unconditional variance of η_{wt} is given by $\sigma_{\eta w}^2 = c_w/(1 - \delta_w^a - \delta_w^b)$.

Two remarks about eq.(1) should be made at this point. First, we include an intercept μ_{it} in the specification for R_{it} rather than in the specifications for R_{it}^I and R_t^W. The reason is that, when estimating a factor model, we cannot attribute the mean of R_{it} to the unobserved components R_{it}^I and R_t^W in a non-arbitrary way. Second, the error term ε_{it} is measurement error so that we can assume that it is *i.i.d.* as there is no reason for it to be subject to $GARCH$ effects (see Harvey et al. (1992), p.138). As such, we assume that $\varepsilon_{it} \sim i.i.d(0, \sigma_{\varepsilon i}^2)$ where $\sigma_{\varepsilon i}^2$ is the unconditional variance of ε_{it} $(\forall i)$.

2.2 Estimation method and identification

2.2.1 Method

To obtain estimates for the unobserved factors R_{it}^I and R_t^W, for the conditional variance series h_{it} and h_{wt}, for the conditional probability series $P_t(S_t = j)$, and for the parameters in the model p^j, μ_i^j, α_i^j, β_i^j, $\sigma_{\varepsilon i}^2$, π_i, π_w, c_i, c_w, δ_i^k, and δ_w^k (where $i = 1, ..., N$, $j = 1, 2$, and $k = a, b$) we first put the model described by eqs.(1)-(10) in state space form. In particular, we estimate a *conditionally* Gaussian linear state space system with Markov switching parameters and including time-varying conditional variances for the error terms (see Harvey et al. (1992), Kim (1994), and Kim and Nelson (1999), chapters 5 and 6). In Appendix we report the state space representation of the model. The parameters in the system are estimated by maximum likelihood. Estimates of the regime probabilities $P_t(S_t = j)$ and of the factors R_{it}^I and R_t^W are obtained with the Kim filter (Kim (1994)) which combines the Hamilton filter that deals with the Markov switching part of the model (Hamilton (1989)) and the Kalman filter that estimates the unobserved factor part of the model.[3] The time-varying conditional variances further complicate the state space framework. To deal with this we follow the approach by Harvey et al. (1992) and augment the state vector with the shocks η_{it} and η_{wt}. The Kalman filter then provides estimates of the conditional variance of the shocks, i.e. estimates for h_{it} and h_{wt}.

To avoid potential computational difficulties caused by the multivariate nature of the state space system (i.e. $N > 1$) we follow the univariate approach to multivariate filtering as presented by Koopman and Durbin (2000) and Durbin and Koopman (2001) (chapter 6). A major advantage of this approach is that we can avoid taking the inverse of the variance matrix of the one-step-ahead prediction errors in the system. We refer to Koopman and Durbin (2000) for the filtering recursions and for the calculation of the likelihood.

[3]Note that we report filtered instead of smoothed estimates for both the states and and the regime probabilities. The reason is that we are interested in the real time ex ante calculation by the financial markets of potential regime shifts.

2.2.2 Identification

There are a number of identification issues in the empirical model, some of which require specific restrictions. First, note that we can multiply and divide the terms $\alpha_{it} R_{it}^I$ and/or $\beta_{it} R_t^W$ by a constant q and obtain a different decomposition of R_{it}, i.e. $R_{it} = (\alpha_{it} q)(R_{it}^I/q)+(\beta_{it} q)(R_t^W/q)+\varepsilon_{it} = \alpha_{it}^* R_{it}^{I*}+\beta_{it}^* R_t^{W*}+\varepsilon_{it}$. To obtain a unique decomposition of R_{it} and hence to uniquely identify the factor loadings α_{it} and β_{it} we impose an unconditional variance of unity on the shocks of the country-specific and common factors η_{it} and η_{wt}, i.e. $\sigma_{\eta i}^2 = 1$ and $\sigma_{\eta w}^2 = 1$. This amounts to setting $c_i = 1 - \delta_i^a - \delta_i^b$ $(\forall i)$ and $c_w = 1 - \delta_w^a - \delta_w^b$ in the $GARCH$ specification of the factor errors. Second, the signs of the factor loadings and of the factors are not identified since the likelihood remains the same if we multiply both R_{it}^I and α_{it} (or R_t^W and β_{it}) by -1. Therefore, we impose the restrictions $\alpha_{it} > 0$ and $\beta_{it} > 0$. Since $S_t^j \geq 0$ for $j = 1, 2$ and either S_t^1 or S_t^2 equals 1 in eqs.(2)-(4) we can impose positivity of α_{it} and β_{it} by imposing $\alpha_i^j > 0$ and $\beta_i^j > 0$ (for $j = 1, 2$) in the estimation of the system.[4] Third, to separately identify ε_{it} and R_{it}^I we need sufficient persistence in R_{it}^I, i.e. values for π_i that are not too close to 0. As will be reported below, our point estimates for π_i are well above 0.5. Finally, if one switches the labels for regime 1 and regime 2, the likelihood is unchanged. In the paper we label regime 1 as the tranquil regime and regime 2 as the crisis regime. An appropriate choice of parameter starting values in the likelihood optimization is sufficient to fix these labels on the regimes, i.e. we choose starting values for the parameters indexed by $j = 2$ that are somewhat higher than the starting values for the parameters indexed by $j = 1$.

2.3 Data

We estimate the empirical model with data for five countries ($N = 5$).[5] The countries are Belgium, France, Italy, the Netherlands, and Spain. The benchmark country in the 10 year government bond market segment is Germany. These six countries are the six largest economies in the euro area in terms

[4]It is conceivable that for some countries $\beta < 0$ so that increases in common risk decrease the yield spread - i.e. the risk premium - of a country. This could be the case if a country is considered a 'safe haven' so that demand for its bonds increases and the required risk premium decreases when the international economic climate deteriorates ('flight to quality'). Violation of the restriction $\beta > 0$ can easily be detected during estimation as it will lead to estimates for β which are very close to 0. We did not find such estimates however so that the restriction is justified.

[5]Our methodology is multivariate state space modeling using maximum likelihood for parameter estimation. This estimation approach necessitates a relatively small number of countries. With too many countries the number of parameters becomes very large which makes the optimization of the likelihood difficult and causes numerical problems. With $N = 5$ the total number of estimated parameters equals 55 and the optimization algorithm converges without problems in a reasonable time span.

of nominal GDP (2009). Their combined share in the total nominal GDP of the euro area equals almost 87%. The five countries for which the spreads are studied constitute an interesting sample taken from the 17 existing euro area economies. France and the Netherlands are countries with a relatively sound fiscal tradition belonging to the core of the euro area. Italy and Spain are peripheral countries considered to be "GIIPS" countries with problematic domestic fiscal conditions and, in the case of Spain, problems in the banking sector. Belgium is a core country but nevertheless struggles with its government debt and is characterized by periods of political instability.

We consider the period after the introduction of the euro in the countries considered, i.e. January 1999 (1/1999) until April 2012 (4/2012). This gives 160 monthly observations. Apart from computational considerations (i.e. limiting the computing time) we use monthly data for two reasons. First, when using monthly data it is sufficient to assume that the unobserved factors follow $AR(1)$ processes to get rid of autocorrelation in the system (while, for instance, with weekly data higher order autoregressive processes are necessary that involve the estimation of many more parameters). Second, structural regime shifts are easier to detect when the higher frequency movements in the data are limited by averaging.

Data for 10 year government bond yields are taken from Datastream/Thomson Financial (code: xxBRYLD where xx is the country code). To calculate the spread R_{it} for country i in period t we subtract the period t yield to maturity of a 10 year government bond issued by the benchmark country Germany from the period t yield to maturity of a 10 year government bond issued by country i (i.e. Belgium, France, Italy, the Netherlands, Spain). Data for the spreads of the five countries under consideration over the period $1/1999 - 4/2012$ are presented in Figure 1. From the figure it is clear that the spreads of all countries remained quite stable and quite close (though not equal) to zero for all countries before the onset of the financial crisis in 2008. After 2008 the spreads increased in all countries, first moderately, and from 2010 onward more dramatically (particularly in Italy and Spain). The descriptive statistics reported in Table $1(a)$ confirm these findings. When comparing the period prior to the Lehman default in September 2008 to the period after the default, we observe an increase in the means and the standard deviations of the spreads of all countries under consideration. These increases are largest for Italy and Spain, somewhat smaller for Belgium and France, and even more moderate for the Netherlands. In

Table 2(a) we also report the unconditional correlations of the spreads for all countries before and after September 2008. From the table we note that the correlations are generally high (always above 0.7 for the full sample). The impact of the crisis on the correlations is not the same for all countries however. Some correlations rise (e.g. between France and Italy) while others fall (e.g. between the Netherlands and all other countries) suggesting that the impact of the financial crisis on the degree of integration of the countries considered is ambiguous. We elaborate on this observation in the next section when analyzing the results of the estimated regime switching dynamic factor model.

3 Results

This section presents and discusses the results from the estimation of the model given by eqs.(1)-(10) for five euro area countries (Belgium, France, Italy, the Netherlands, and Spain) over the period 1/1999 – 4/2012. Table 2 presents the parameter estimates and the statistical tests conducted to determine the adequacy of the specification. Figures 2-7 present the estimated regime probabilities and the estimated factors and risk premiums and their conditional variances.

From Table 2 we report, first, that the estimates for the country-specific AR parameters and the common AR parameter are large (always above 0.5) and highly significant. The dependency structure is rather similar across countries. A Ljung-Box test for autocorrelation at different lag lengths (1, 4, and 12) conducted on the estimated one step ahead prediction errors of the system and reported in Table 2 further shows that the null hypothesis of no autocorrelation is never rejected at the 5% level of significance (except for the Netherlands at lag length 1). This supports our choice to model the unobserved factors in the state space system as AR(1) processes. Second, the δ^a parameters of the estimated GARCH processes are significant both for the country-specific factors and for the common factor. The δ^b parameters are significant for the common factor and for France and Italy but are close to zero for the remaining three countries. The latter result implies low persistence in the conditional variance series of the spreads of these countries and can be explained when noting that the regime shift introduced through the Markov switching variable S_t can capture much of the persistence that is present in the variance of the spreads. We test for remaining heteroscedasticity in the estimated one step ahead prediction errors of the system by conducting Ljung-Box tests for autocorrelation on the *squared* prediction errors (again at lag lengths 1, 4, and 12). We can never reject the null hypothesis of no heteroskedasticity in the prediction errors. Therefore, we conclude that the inclusion of GARCH(1,1) processes in the state errors is sufficient to capture the heteroscedasticity that is present in the data. Third, the estimated variances σ_ε^2 of the measurement error term are always of small magnitude and generally (almost) significant (except for the Netherlands).

In Figure 2 we present the estimated conditional probabilities of regime 1 and regime 2. From these graphs we observe an important permanent regime shift in the first half of 2008. Regime 1 is the regime prevalent before the shift and regime 2 is the regime prevalent after the shift, i.e. the crisis regime. Hence, a shift in the pricing of risk takes place during the first half of 2008, well before the announcement of the bankruptcy of Lehman Brothers in the US and more than one year before the outbreak of the sovereign debt crisis in the euro area periphery. This suggests that even before these key events took place euro area government bond market investors had already incorporated a high probability that a change from a tranquil to a crisis regime had taken place. While we allow for multiple regime shifts between tranquil and crisis periods in our empirical approach, we identify only this single shift to the crisis regime so that the crisis regime lasts until the end of the sample period. In line with these results, in Table 2 we report the estimated transition probabilities p^1 and p^2 of the Markov switching process followed by the regime variable S_t. Both are strongly significant and lie above 0.99 showing that both regimes are very persistent (i.e. once a regime is in place it is very hard to overturn it). In the table we also conduct a likelihood ratio test that compares our two-regime model with an alternative restricted model in which there is only a tranquil regime and no crisis regime. This restricted model is obtained from our model by setting the transition probabilities $p^1 = 1$ and $p^2 = 0$ so that $\mu_{it} = \mu_i^1$, $\alpha_{it} = \alpha_i^1$, and $\beta_{it} = \beta_i^1$.[6] We strongly reject the null hypothesis of a restricted model with only one regime and no regime shift.

The regime shift has affected the government bond yield spreads in a qualitatively identical way in all countries: through an increase of the country-specific means of the spreads μ, through an increase

[6]Setting $p^1 = 1$ and $p^2 = 0$ is sufficient for our model to collapse to a model with only one regime as it implies that the conditional regime probabilities in period 1 are given by $P_1(S_1 = 1) = 1$ and $P_1(S_1 = 2) = 0$ so that, given the extreme values of the transition probabilities, for all t we have $P_t(S_t = 1) = 1$ and $P_t(S_t = 2) = 0$ and regime 2 vanishes from the model.

of the country-specific factor loadings α on the idiosyncratic factor R^I, and through an increase of the country-specific factor loadings β on the common euro area-wide risk factor R^W. This implies that in all countries the financial crisis has caused three types of contagion: "pure" contagion (higher μ's), "wake up call" contagion (higher α's) and "shift" contagion (higher β's). In Table 2 we report the estimates for μ, α, and β under both regimes for all countries.[7] We note, first, that the estimated country-specific means μ are all positive and significant and much larger for all countries during the second regime. The increase is largest in absolute terms for Italy and Spain while the Netherlands show a rather modest increase. Second, the estimates of the loadings α on the idiosyncratic factor are positive and significant under both regimes. Again, the estimates show an increase during the second regime for all countries. This increase is particularly large for Italy, France, and Spain. Third, when looking at the estimated loadings β on the common risk factor the same conclusion holds. They are positive and significant for all countries under both regimes and larger during the second regime. The increase in the exposure to common risk is particularly large for Italy, Spain, and Belgium.

The estimation results therefore suggest that the spreads of all the countries in our sample are driven by both an idiosyncratic risk premium and a common risk premium and that the importance of both premiums has increased considerably after the financial crisis.

Figure 3 shows the estimated country-specific risk premium αR^I augmented with the intercept μ.[8,9] The figure shows an increase in all premiums starting in the first half of 2008. The magnitude of the increase for Italy and Spain is much larger than for France and the Netherlands however, which reflects the problematic magnitude of debt and deficit ratios in these countries and other country-specific problems. It should be noted that after the initial increase in the first half of 2008 the premiums continue to increase for Italy and Spain while they remain stable in 2009 for France and Belgium and decrease rather quickly and drastically in the Netherlands. From 2010 onwards Italy, Spain, France, and Belgium show further increases in the premiums which reach their highest points by the second half of 2011. The Netherlands,

[7]Since the unconditional variances of the factor errors are fixed to 1 for identification, the loadings α and β can be interpreted as the unconditional standard errors of the shocks to respectively the pure country-specific premium αR^I and to the country-specific common risk premium βR^W. Since the bond yield spreads are expressed in percentage terms, the α's and β's are therefore expressed in percentage points.

[8]The intercept is considered country-specific although it could also (partially) be part of the common factor. As mentioned in section 2.1 it is not possible to attribute the mean of the spreads R to the factors R^I and R^W in a non-arbitrary way.

[9]It is calculated as $P(S=1)(\mu^1 + \alpha^1 R^{I1}) + P(S=2)(\mu^2 + \alpha^2 R^{I2})$ where P denotes the regime probabilities and where R^{I1} and R^{I2} are the idiosyncratic factors in regime 1, respectively regime 2.

on the other hand, shows a further reduction in the country-specific risk premium until it reaches its lowest point by early 2012. Clearly, after the initial increase in 2008-2009 observed in all countries, investors have started to differentiate more strongly between the countries later on. In Figure 4 the estimated conditional variance series of the country-specific risk premium is presented, which confirms that in all countries from early 2008 onward there is increased financial turmoil.[10]

In Figure 5 we present the estimated common factor R^W and its estimated conditional variance. From these figures we observe that common euro area-wide risk, which most likely is strongly affected by events happening in the US, starts to rise in 2006 and increases sharply during 2007, i.e. at the time of the US subprime debt crisis. It stabilizes around the period 2009-2010 but then rises sharply again in 2011 possibly reflecting the grim economic outlook for both the US and the European countries. By then the latter have been severely hit by the sovereign debt crises. The conditional variance of the common factor shows a large peak around September 2008, at the time of the Lehman default, and again in 2011-2012.

By multiplying the common international risk factor R^W with the country-specific factor loadings β we obtain the country-specific common risk premium βR^W.[11] This estimated premium is presented in Figure 6 and its estimated conditional variance in Figure 7.[12] Since estimation results reported in Table 2 show that the β's are positive and significant *and* larger in the crisis regime, the impact of the common risk premium on the spreads during the crisis is twofold. First, the spreads have generally widened through the increase in common euro area-wide risk captured by the common factor R^W. This supports the findings of several recent studies such as Borgy et al. (2011) and von Hagen et al. (2011) who argue that the financial crisis caused a widening of spreads in the euro area through an increase in common international risk. Second, the exposure to international risk as measured by the country-specific factor loadings β is higher after the start of the crisis in all countries, in particular in Italy, Spain, and Belgium. From Figure 6 we indeed note that the magnitude of the increase in the common risk premium observed during the period 2008-2012 is highest for these three countries. It is most likely the problematic

[10] This variance is calculated as $P(S=1)((\alpha^1)^2 h^1) + P(S=2)((\alpha^2)^2 h^2)$ where P denotes the regime probabilities and where h^1 and h^2 denote the time-varying conditional variances of the idiosyncratic factor in regime 1, respectively regime 2.

[11] It is calculated as $P(S=1)(\beta^1 R^{W1}) + P(S=2)(\beta^2 R^{W2})$ where P denotes the regime probabilities and where R^{W1} and R^{W2} are the common factors in regime 1, respectively regime 2.

[12] This variance is calculated as $P(S=1)((\beta^1)^2 h_w^1) + P(S=2)((\beta^2)^2 h_w^2)$ where P denotes the regime probabilities and where h_w^1 and h_w^2 denote the time-varying conditional variances of the common risk factor in regime 1, respectively regime 2.

fiscal situation in these countries and their generally weaker fundamentals that have made them more vulnerable to the movements in common international risk (i.e. to "shift" contagion).

We end our discussion of the results with a note on the impact of the financial crisis on the integration of euro area government bond markets. The increased importance of both purely country-specific risk and common risk for the yield spreads makes the impact of the financial crisis on the degree of government bond market integration ambiguous a priori. The increased importance of purely country-specific risk suggests that integration may have decreased, while the increased importance of common risk suggests that integration may have increased. From an unconditional variance decomposition applied to eq.(1) under regimes 1 and 2 we obtain a simple measure of country-specific market integration under both regimes, i.e. $\frac{V(\beta_{it}R_t^W)}{V(\alpha_{it}R_{it}^I+\beta_{it}R_t^W)} = \frac{(\beta^j)^2(1-\pi_w^2)^{-1}}{(\alpha^j)^2(1-\pi^2)^{-1}+(\beta^j)^2(1-\pi_w^2)^{-1}}$ where j is the regime index, i.e. $j = 1, 2$, and where the derivation uses the restriction that fixes the unconditional variances of the errors of the country-specific and common factors to 1. The higher this ratio the higher the degree of integration of a country with the remaining countries in a specific regime. By calculating this ratio for all countries and comparing the change in the ratio from regime 1 to regime 2, we find no change for Belgium, a higher degree of integration for Spain and France, and a lower degree of integration for the Netherlands and Italy. While we do not find clear cut general results with respect to the impact of the financial crisis on the degree of financial market integration in euro area government bond markets, the decreased integration measured for the Netherlands is remarkable. Further research might shed light on whether this result reflects a decoupling from the other countries caused by a 'safe haven' or a 'flight to quality' effect.

4 Conclusions

The financial crisis has initiated a substantial literature that explores the channels through which financial turmoil affects the government bond yield spreads of euro area countries versus (usually) Germany. These spreads reflect the risk premium demanded by investors to be willing to hold government debt issued by euro area countries. This paper adds to the literature by investigating the presence of euro area-wide crisis regime periods in the yield spreads of euro area government bonds. We investigate how often and when the crisis regime occurs, how long crisis episodes last, and to what extent and through which channels they affect the pricing of government bond risk in euro area countries. To this end, a dynamic factor model with Markov switching parameters is estimated using monthly data for the 10 year government bond yield spreads of five euro area countries (Belgium, France, Italy, the Netherlands, and Spain) versus Germany over the period 1999/1 until 2012/4. We argue that our methodological approach has a number of advantages compared to the existing literature.

Our estimates identify a single permanent regime shift in the pricing of risk during the first half of 2008, i.e. the period before the Lehman default was announced (September 2008) and well before the outbreak of the government debt crisis in the euro area periphery. It therefore seems that even before the announcement of the Lehman default, euro area government bond market investors had already incorporated a high probability that the relatively tranquil regime had come to an end. Our results show that all countries experienced qualitatively similar changes in the pricing of their bonds and that increases in the spreads were caused both by an increase in pure country-specific risk and by an increase in the country-specific premium for common euro area-wide risk. The latter premium is the product of the common risk factor and the country-specific impact of common euro area-wide risk, both of which increased during the financial crisis. The magnitude of the changes is different across countries however and larger for Italy and Spain, countries plagued by fiscal problems and generally weaker fundamentals.

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Appendix. State space representation of the model

The state space system with state vector Ω_t and regime variable S_t is given by

$$y_t = Z_t(S_t)\Omega_t + \varepsilon_t \tag{A.1}$$

$$\Omega_t = T_t\Omega_{t-1} + G_t\eta_t \tag{A.2}$$

with,

$$\begin{aligned} \varepsilon_{t|t-1} &\sim N(0, H) \\ \eta_{t|t-1} &\sim N(0, Q_t) \end{aligned}$$

The state vector Ω_t is initialized with matrices M_1 and V_1 which, given that all states are covariance-stationary, contain respectively the unconditional means and the unconditional variances of the states included in Ω_t. The regime probabilities $P_t(S_t = j)$ for $j = 1, 2$ are initialized by the unconditional regime probabilities $P_1(S_1 = 1) = (1 - p^2)/(2 - p^1 - p^2)$ and $P_1(S_1 = 2) = (1 - p^1)/(2 - p^1 - p^2)$ where p^j (with $j = 1, 2$) are the transition probabilities of the Markov switching process for S_t (for a derivation see Kim and Nelson (1999), chapter 4).

Given $N = 5$ we have

$$y_t = [R_{1t}\ R_{2t}\ R_{3t}\ R_{4t}\ R_{5t}]'$$

$$\Omega_t = [1\ R^I_{1t}\ R^I_{2t}\ R^I_{3t}\ R^I_{4t}\ R^I_{5t}\ R^w_t\ \eta_{1t}\ \eta_{2t}\ \eta_{3t}\ \eta_{4t}\ \eta_{5t}\ \eta_{wt}]'$$

$$\varepsilon_t = [\varepsilon_{1t}\ \varepsilon_{2t}\ \varepsilon_{3t}\ \varepsilon_{4t}\ \varepsilon_{5t}]'$$

$$\eta_t = [\eta_{1t}\ \eta_{2t}\ \eta_{3t}\ \eta_{4t}\ \eta_{5t}\ \eta_{wt}]'$$

$$Z_t(S_t) = \begin{bmatrix} \mu_{1t} & \alpha_{1t} & 0 & 0 & 0 & 0 & \beta_{1t} & 0 & 0 & 0 & 0 & 0 & 0 \\ \mu_{2t} & 0 & \alpha_{2t} & 0 & 0 & 0 & \beta_{2t} & 0 & 0 & 0 & 0 & 0 & 0 \\ \mu_{3t} & 0 & 0 & \alpha_{3t} & 0 & 0 & \beta_{3t} & 0 & 0 & 0 & 0 & 0 & 0 \\ \mu_{4t} & 0 & 0 & 0 & \alpha_{4t} & 0 & \beta_{4t} & 0 & 0 & 0 & 0 & 0 & 0 \\ \mu_{5t} & 0 & 0 & 0 & 0 & \alpha_{5t} & \beta_{5t} & 0 & 0 & 0 & 0 & 0 & 0 \end{bmatrix}$$

where $\mu_{it} = \mu_i(S_t) = \mu^1_i S^1_t + \mu^2_t S^2_t$ (and similarly for α_{it} and β_{it}) for i=1,...,5 and S^j_t are the regimes (with j=1,2),

$$G_t = \begin{bmatrix} \mathbf{0}_6 \\ \mathbf{I}_6 \\ \mathbf{I}_6 \end{bmatrix}$$

where $\mathbf{0}_6$ is a 1×6 vector of 0's and $\mathbf{I}_6$ is an 6×6 identity matrix,

$$T = \begin{bmatrix}
1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & \pi_1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & \pi_2 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & \pi_3 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & \pi_4 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & 0 & \pi_5 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & 0 & 0 & \pi_w & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0
\end{bmatrix}$$

$$diag(H) = [\sigma^2_{\varepsilon_1}\, \sigma^2_{\varepsilon_2}\, \sigma^2_{\varepsilon_3}\, \sigma^2_{\varepsilon_4}\, \sigma^2_{\varepsilon_5}]'$$

$$diag(Q_t) = [h_{1t}\, h_{2t}\, h_{3t}\, h_{4t}\, h_{5t}\, h_{wt}]'$$

where $h_{it} = c_i + \delta^a_i\, \eta^2_{it-1} + \delta^b_i\, h_{it-1} = (1 - \delta^a_i - \delta^b_i) + \delta^a_i\, \eta^2_{it-1} + \delta^b_i\, h_{it-1}$ for i=1,...,5 and $h_{wt} = c_w + \delta^a_w\, \eta^2_{wt-1} + \delta^b_w\, h_{wt-1} = (1 - \delta^a_w - \delta^b_w) + \delta^a_w\, \eta^2_{wt-1} + \delta^b_w\, h_{wt-1}$

$$M_1 = [1\,0\,0\,0\,0\,0\,0\,0\,0\,0\,0\,0\,0]$$

$$diag(V_1) = [0\, \frac{\sigma^2_{\eta_1}}{1-\pi_1^2}\, \frac{\sigma^2_{\eta_2}}{1-\pi_2^2}\, \frac{\sigma^2_{\eta_3}}{1-\pi_3^2}\, \frac{\sigma^2_{\eta_4}}{1-\pi_4^2}\, \frac{\sigma^2_{\eta_5}}{1-\pi_5^2}\, \frac{\sigma^2_{\eta_w}}{1-\pi_w^2}\, \sigma^2_{\eta_1}\, \sigma^2_{\eta_2}\, \sigma^2_{\eta_3}\, \sigma^2_{\eta_4}\, \sigma^2_{\eta_5}\, \sigma^2_{\eta_w}]'$$

where $\sigma^2_{\eta_i} = 1$ $(\forall i)$ and $\sigma^2_{\eta_w} = 1$.

Tables and figures

Table 1: Descriptive statistics and unconditional correlations

(a) Descriptive statistics of 10 year government bond yield spreads versus Germany (in percentages, January 1999 - April 2012)

	Belgium	France	Italy	Netherlands	Spain
	Full sample 1/1999 - 4/2012				
Mean	0.3939	0.1902	0.6679	0.1537	0.5730
Maximum	2.9125	1.4672	4.8802	0.6667	4.2732
Minimum	-0.0125	-0.0103	0.0791	-0.0558	-0.0543
Stdv	0.5108	0.2597	0.9200	0.1504	0.9466
	First subsample 1/1999 - 8/2008				
Mean	0.1625	0.0725	0.2552	0.0811	0.1288
Maximum	0.4174	0.2122	0.6017	0.2138	0.3564
Minimum	-0.0126	-0.0104	0.0791	-0.0558	-0.0543
Stdv	0.1230	0.0572	0.0949	0.0640	0.1207
	Second subsample 9/2008 - 4/2012				
Mean	1.0041	0.5005	1.7560	0.3453	1.7442
Maximum	2.9125	1.4672	4.8802	0.6667	4.2732
Minimum	0.3313	0.2128	0.6343	0.1674	0.4620
Stdv	0.6315	0.3237	1.1977	0.1448	1.1571

(b) Unconditional correlations of 10 year government bond yield spreads versus Germany (January 1999 - April 2012)

	Belgium	France	Italy	Netherlands	Spain
	Full sample 1/1999 - 4/2012				
Belgium	1				
France	0.9633	1			
Italy	0.9779	0.9690	1		
Netherlands	0.7772	0.7860	0.7063	1	
Spain	0.9560	0.9245	0.9640	0.6801	1
	First subsample 1/1999 - 8/2008				
Belgium	1				
France	0.9112	1			
Italy	0.8171	0.7909	1		
Netherlands	0.8995	0.8619	0.7233	1	
Spain	0.9519	0.8760	0.7939	0.8676	1
	Second subsample 9/2008 - 4/2012				
Belgium	1				
France	0.9206	1			
Italy	0.9739	0.9494	1		
Netherlands	0.4041	0.4535	0.3204	1	
Spain	0.9071	0.8314	0.9247	0.1411	1

Table 2: Maximum likelihood estimation of the state space system eqs.(1)-(10),1999:1-2012:4

	Country-specific parameters$^{(a)}$					Common parameters$^{(a)}$
	Belgium	France	Italy	Netherlands	Spain	
μ^1	0.246	0.113	0.371	0.135	0.209	
	(0.055)	(0.024)	(0.061)	(0.029)	(0.047)	
μ^2	0.723	0.332	1.177	0.236	0.851	
	(0.052)	(0.028)	(0.087)	(0.012)	(0.066)	
α^1	0.017	0.052	0.025	0.026	0.024	
	(0.009)	(0.024)	(0.014)	(0.009)	(0.007)	
α^2	0.089	0.225	0.277	0.080	0.145	
	(0.053)	(0.139)	(0.212)	(0.033)	(0.045)	
β^1	0.034	0.015	0.032	0.018	0.017	
	(0.009)	(0.004)	(0.009)	(0.005)	(0.005)	
β^2	0.181	0.095	0.297	0.037	0.221	
	(0.052)	(0.028)	(0.087)	(0.012)	(0.066)	
p^1						0.994
						(0.006)
p^2						0.992
						(0.010)
π	0.938	0.617	0.961	0.656	0.967	0.962
	(0.030)	(0.081)	(0.031)	(0.059)	(0.016)	(0.016)
c	0.039	0.003	0.018	0.195	0.272	0.063
	(0.051)	(0.000)	(0.011)	(0.154)	(0.231)	(0.046)
δ^a	0.961	0.338	0.687	0.804	0.728	0.247
	(0.050)	(0.083)	(0.174)	(0.153)	(0.231)	(0.072)
δ^b	3.04e-13	0.659	0.294	2.20e-10	1.65e-10	0.690
	(2.94e-08)	(0.082)	(0.164)	(2.51e-06)	(1.36e-06)	(0.072)
σ^2_ε	3.68e-05	3.32e-05	1.71e-04	4.19e-11	6.75e-05	
	(2.51e-05)	(2.09e-05)	(4.58e-05)	(4.22e-08)	(4.25e-05)	
Country-specific Ljung-Box test for Autocorrelation$^{(b),(c)}$						
lag 1	0.202	0.206	0.122	5.738	0.147	
	[0.653]	[0.650]	[0.726]	[0.017]	[0.701]	
lag 4	3.175	2.001	0.891	6.287	0.636	
	[0.529]	[0.735]	[0.926]	[0.179]	[0.959]	
lag 12	12.521	9.370	9.143	14.550	9.589	
	[0.405]	[0.671]	[0.690]	[0.267]	[0.652]	
Country-specific Ljung-Box test for Heteroskedasticity$^{(b),(d)}$						
lag 1	0.278	0.103	0.773	1.026	3.67e-06	
	[0.598]	[0.748]	[0.379]	[0.311]	[0.998]	
lag 4	0.573	1.514	4.338	1.181	4.954	
	[0.965]	[0.824]	[0.362]	[0.881]	[0.291]	
lag 12	0.909	2.912	7.824	4.680	13.034	
	[0.999]	[0.996]	[0.798]	[0.967]	[0.366]	
Likelihood ratio test: H$_0$:$p^1 = 1, p^2 = 0, \mu_{it} = \mu_i^1, \alpha_{it} = \alpha_i^1, \beta_{it} = \beta_i^1 (\forall i, t)$						
test value $\sim \chi^2_{17}$	165.6					
p-value	2.08e-26					

Notes: (a) Standard errors are in parentheses. (b) p-values are in square brackets. (c) The null hypothesis is no autocorrelation in the one-step-ahead prediction errors. (d) The null hypothesis is homoscedasticity in the one-step-ahead prediction errors.

Figure 1: 10 year government bond yield spreads of Belgium, France, Italy, the Netherlands, and Spain versus Germany (January 1999 - April 2012)

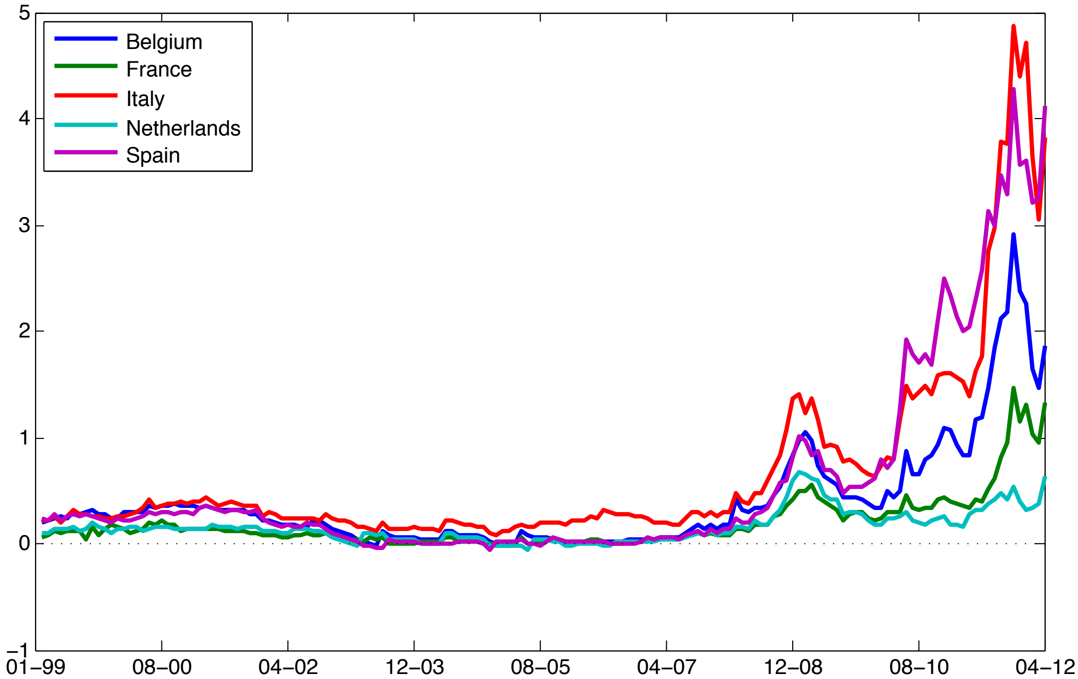
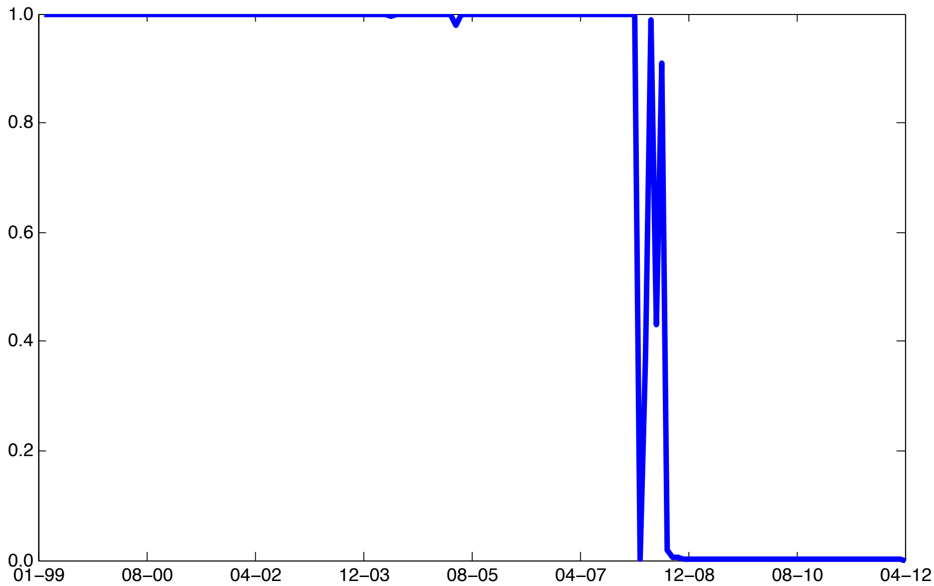
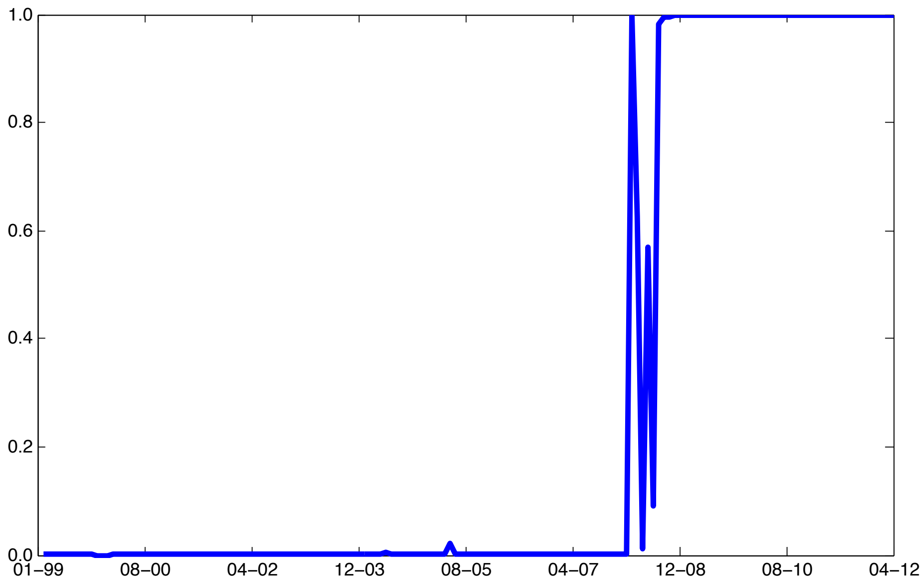


Figure 2: The estimated conditional probability of regime 1 and regime 2 (crisis regime)



PROBABILITY REGIME 1



PROBABILITY REGIME 2

Figure 3: The estimated country-specific risk premium $\alpha_{it} R^I_{it}$ augmented with the spread mean μ_{it}

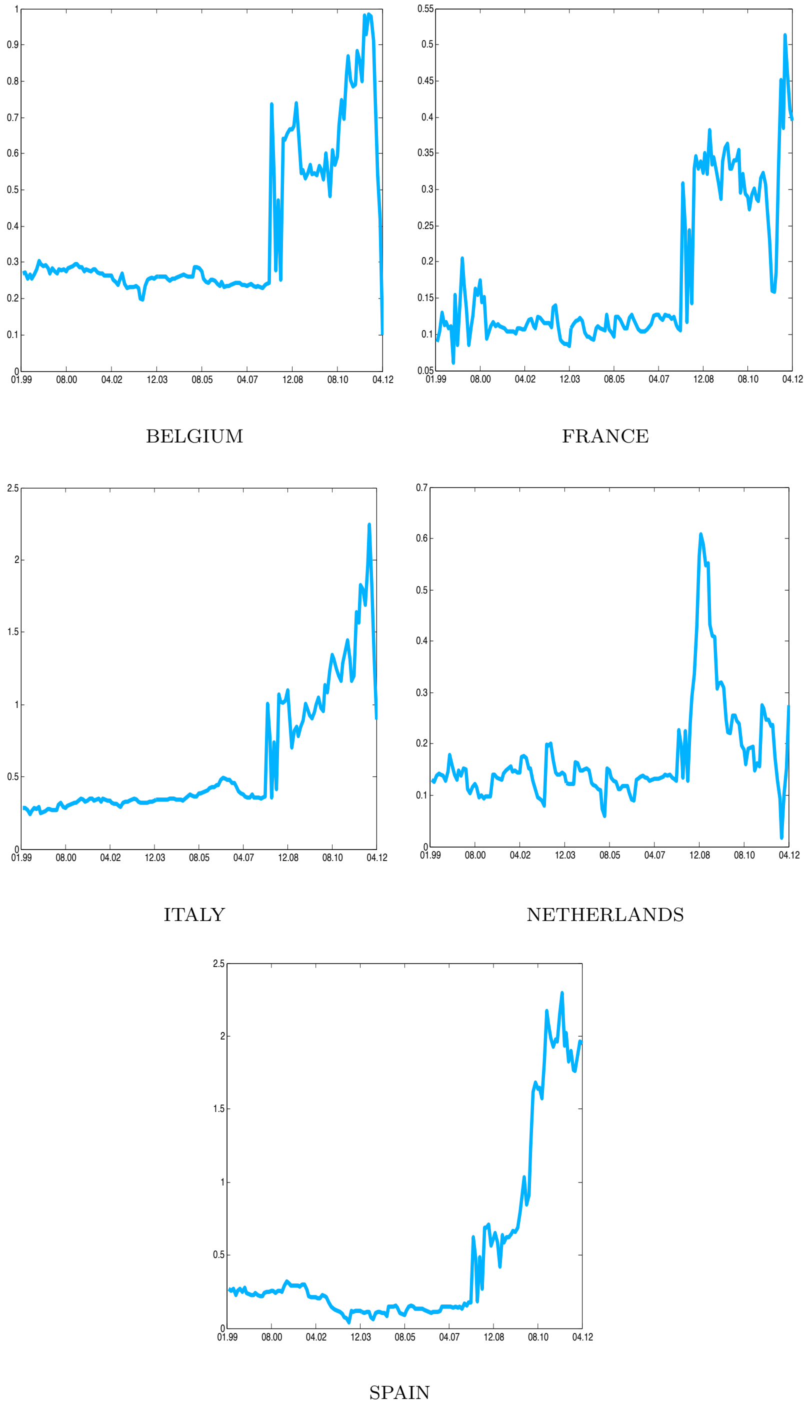
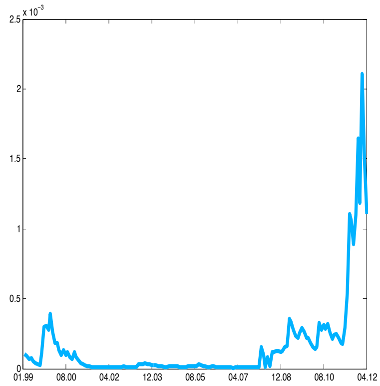
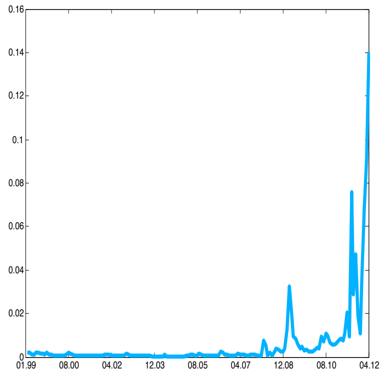


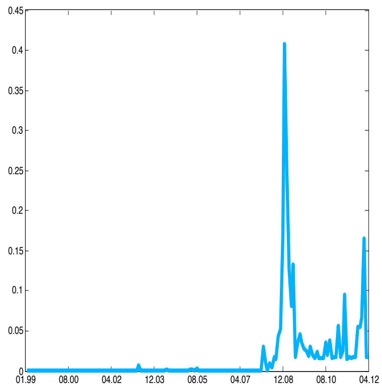
Figure 4: The estimated conditional variance of the country-specific risk premium



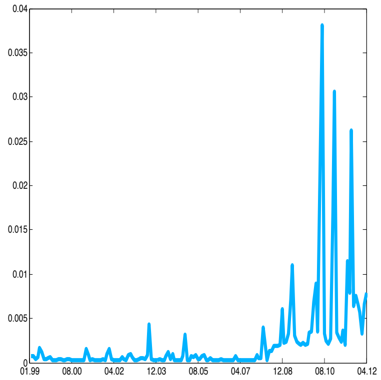
BELGIUM



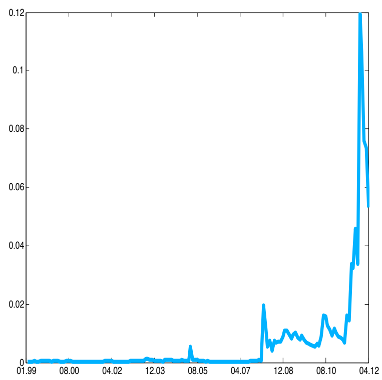
FRANCE



ITALY

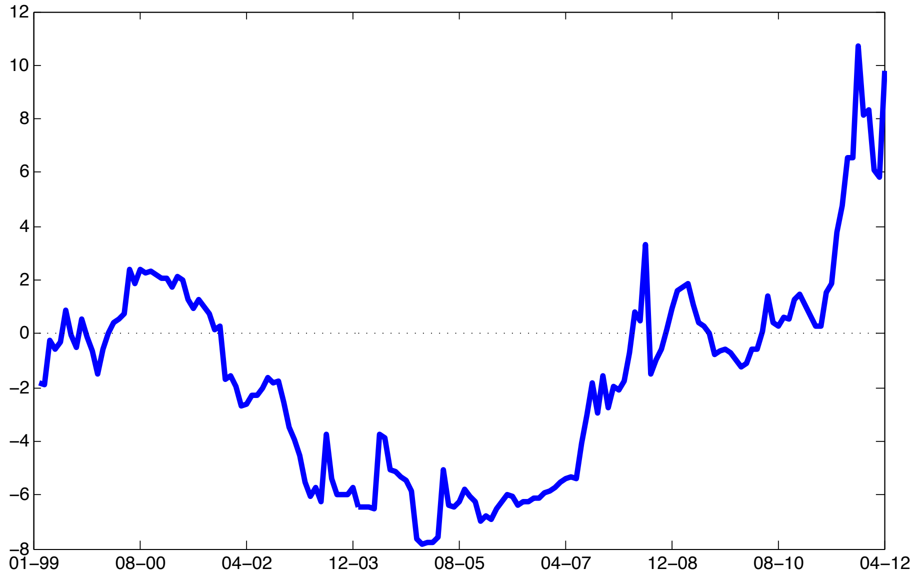


NETHERLANDS

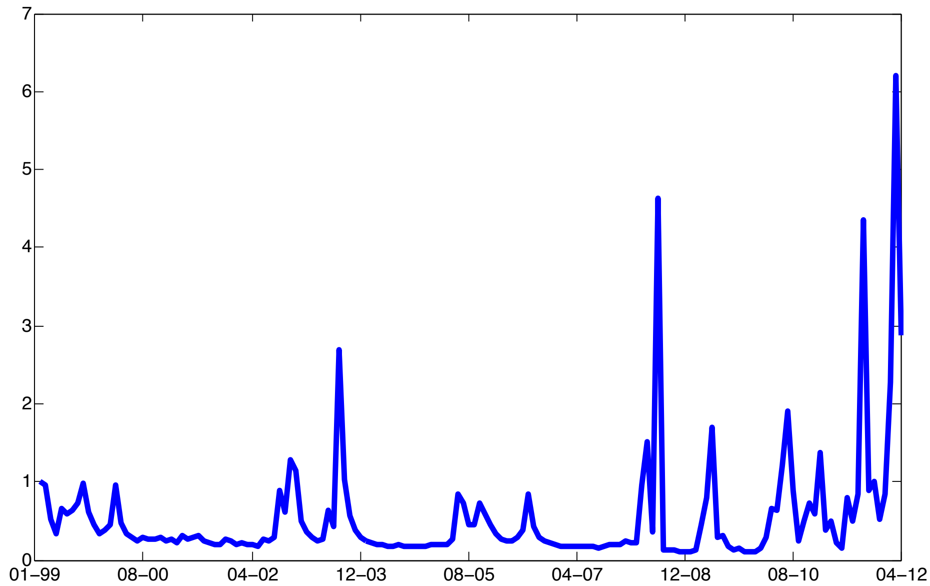


SPAIN

Figure 5: The estimated common factor R_t^W and its estimated conditional variance h_{wt}

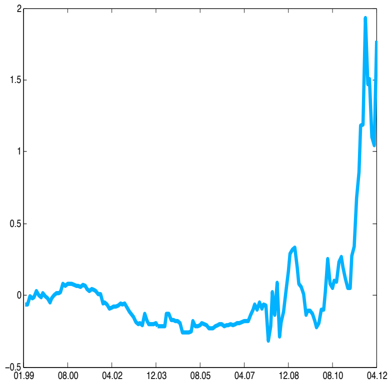


COMMON FACTOR

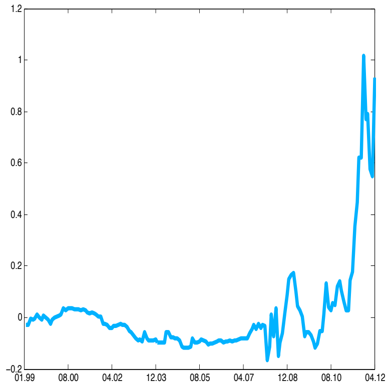


CONDITIONAL VARIANCE

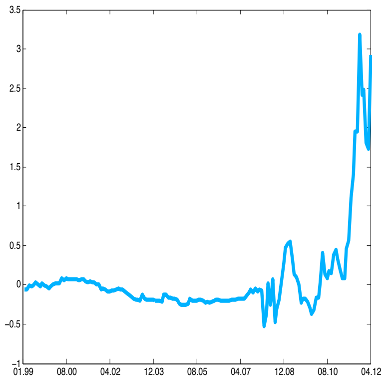
Figure 6: The estimated common risk premium $\beta_{it} R_t^W$



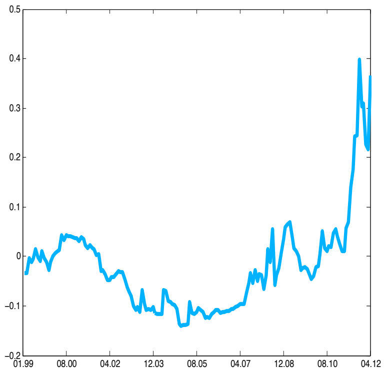
BELGIUM



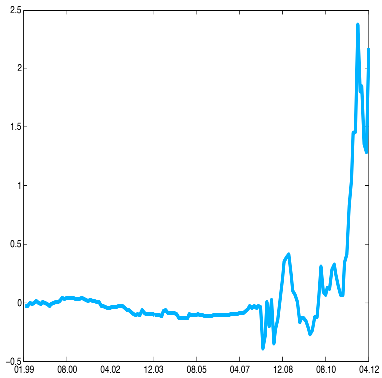
FRANCE



ITALY



NETHERLANDS



SPAIN

Figure 7: The estimated conditional variance of the common risk premium