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What can EMU Countries' Sovereign Bond Spreads Tell Us About Market Perceptions of Default Probabilities During the Recent Financial Crisis? *

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Abstract

This paper presents a new approach to analysing recent movements of EMU sovereign bond spreads. Based on a GARCH-in-mean model originally used in the exchange rate target zone literature, spreads are decomposed into a risk premium, an expected loss component and a liquidity premium. Time-varying probabilities of default are derived. The results suggest that the rise in sovereign spreads during the recent financial crisis mainly reflects an increased expected loss component. In addition, the rescue of Bear Stearns in March 2008 seems to mark a change in market perceptions of sovereign bond risk. The government bonds of some countries lost their former role as a safe haven. While price competitiveness always helps to explain sovereign spreads, it increasingly moved into investors' focus as financial sector soundness weakened.

JEL codes: E43, G15, C32, H63, F36

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1 Introduction

The spreads of euro-area government bonds over German Bunds widened substantially during the financial crisis. They peaked at 300 basis points and remained at their elevated level until April 2009, the end of our observation period. At the same time, considerable differences in yield spreads across countries have emerged since the second quarter of 2008. This paper aims, first, at explaining sovereign bond spread movements within the euro area during the crisis and, second, at providing high-frequency series of country-specific probabilities of default. A particular challenge consists in estimating the probability of an event that has not happened before (the default of an EMU member state) and, at the same time, separating this effect from liquidity concerns and premia which are due to the risk aversion of investors. The paper examines this issue using a GARCH-in-mean model, which was originally developed for the analysis of exchange rate target zones and which allows bond spreads to be decomposed into credit risk, liquidity premia and a component reflecting default expectations. The model is estimated for a calm period and a crisis episode. Following Mody (2009), the rescue of US investment bank Bear Stearns in mid-March 2008 is chosen as the turning point after which differentiation of sovereign bonds increased.

2 Related literature

The literature on credit spreads generally distinguishes between structural approaches derived from the Merton model (1974) and reduced form models such as those of Jarrow/Turnbull (1995).[1] In order to be able to use structural approaches to explain

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[1] According to the structural approach, an enterprise's liabilities constitute a put option held by the debtor on the enterprise's value. Wherever an enterprise's value falls below the nominal value of its liabilities, this leads to an – endogenously modelled – default and the option being exercised. By contrast, in the case of reduced-form models, the default is determined by an exogenously specified intensity process. This process can, in turn, depend on country-specific and macroeconomic factors.

sovereign spreads, it is necessary to define appropriate country-specific proxy variables for the level of indebtedness and the volatility of the firm's value – as, for instance, emphasised by Diaz Weigel/Gemmill (2006) and Oshiro/Saruwatari (2005). Such approaches have the disadvantage that the calculated measures of sovereign risk (distance to default) reflect not only country-specific factors but also risk premia – which vary according to investors' time-varying risk aversion.[2] Furthermore structural approaches are criticised as being unsuitable for the modelling of sovereign spreads. This argument is based on the premise that the state's incentives to default are much more complicated than those of enterprises, with the consequence that the option price theory offers insufficient modelling capability. Duffie et al (2003) reason that an enterprise effectively goes into default when it becomes unable to fulfil its payment obligations, whereas in the case of governments, matters largely hinge on a political decision by the government and its willingness to pay, which depends on a variety of considerations and where the default can take different forms.

Reduced-form approaches normally use a number of different macro variables as the determinants of country risk. The conventional literature, eg Reinhart et al (2003), Eichengreen et al (2003) or Goldstein/Turner (2004), analyses the country risks of emerging market economies, paying particular attention to debt sustainability, original sin and currency mismatches. Under these approaches, country risk is frequently measured on the basis of country ratings. However, the rating agencies have been slow to adapt their country ratings to the recent financial crisis triggered by events in the US real estate market. Moreover, such approaches are unable to provide any explicit information on the probability of default of an individual country. In the wake of the financial crisis and the resultant government rescue packages for financial institutions in many industrial countries, growing attention has been focused on the weaknesses of the financial sector as additional determinants of country risks (see, for instance, Mody, 2009). At the same time, according to Sgherri/Zoli (2009) it would seem that, in the wake of the financial crisis, the relative liquidity of markets has had a major impact on government bonds, a circumstance that is likely to have led to a temporary flight to safety and liquidity on the part of investors. With respect to the euro area, both Gomez-Puig (2006) and

[2] Remolona et al (2007) conclude that "…[the] notion that spreads might contain significant risk premia that are driven by investors' risk aversion is not seriously entertained."

Manganelli/Wolswijk (2009) identify indications that liquidity is an important explanatory factor for the yield spreads between government bonds.

Our GARCH-in-mean approach takes into account both macro variables and the soundness of the financial sector and simultaneously enables the decomposition of sovereign spread into three components (expected loss components, risk premia and liquidity premia). According to Flavin/Limosani (2007, p 105), who analyse the short-term yield differentials of a number of European countries prior to the introduction of the euro, an ARCH-in-mean approach is particularly well suited for this purpose "…as it captures the time variation in the premium while at the same time being consistent with many of the stylized facts of asset prices such as thick tails and volatility clustering." Kounitis (2007) applies the approach espoused by Flavin/Limosani to analyse corporate credit spreads and, in so doing, examines the empirical relevance of the determinants recommended by the Merton model (1974). Unlike the approaches put forward by Flavin/Limosani (2007) and Kounitis (2007), our approach explicitly considers the liquidity premia that are contained in sovereign spreads. What is more, by including financial sector soundness and international competitiveness in the scope of its analysis it focuses on determinants that could have played a major role in the financial crisis.

3 Sovereign yield spreads and the probability of default in a monetary union

The analysis of a relationship between yield differentials and perceived probabilities of default is based on uncovered interest parity augmented by a time-varying risk premium,

$$i_t - i_t^* = E_t(\Delta s_{t+k}) + \rho_t , \tag{1}$$

where i_t = yield on a domestic bond with a maturity k at time t, i_t^* = the yield on the equivalent foreign bond, s_t = logarithmic exchange rate between the currencies of the two countries under observation expressed in units of the domestic currency per unit of foreign currency, and ρ_t = time-varying risk premium for holding domestic bonds. The yield differential is equivalent to the rate at which the domestic currency is expected to have depreciated by the time the bond matures plus a risk premium to cover investors' risk aversion.

Equation (1) as it stands is, naturally, not suited to describing the situation for two countries participating in a currency union. However, the Bertola and Svensson (1993) approach to estimating a target-zone model can be used to expand equation (1) to include a regime change for cases where the chosen central parity does not appear entirely credible. To this end, the exchange rate in a target-zone regime is defined as consisting of two components, a central parity c_t and the current deviation of the exchange rate from this central parity d_t:

$$s_t \equiv c_t + d_t \,. \tag{2}$$

The expected depreciation rate is thus composed of the expected change in the central parity and the expected change in the deviation from the central parity. It is assumed that the central parity is constant apart from a possible discrete adjustment and that investors know neither the level nor the time of a future adjustment of the central parity. Denoting the probability of a regime change, ie an adjustment of the central parity, over the life k of the bond as π_{kt}, this yields the expected depreciation rate

$$E_t(\Delta s_{t+k}) = \pi_{kt} E_t(\Delta c_{t+k}) + (1 - \pi_{kt}) E_t(\Delta d_{t+k}) \,. \tag{3}$$

Hallwood et al (2000) is one of the papers on target zones which uses equations such as (3) in conjunction with (1) to determine the risk of an adjustment of the central parity.[3] If the above considerations are applied to the situation within a currency union, the second term on the right-hand side of equation (3) drops out; provided there is no regime change, no depreciation is expected in a currency union, and so $E_t(\Delta d_{t+k}) = 0$.

The first term on the right-hand side of equation (3), which describes a realignment of the central parity in the target-zone model, can, in the context of a currency union, be interpreted as expectations of an exit from the currency union. In this case, the exit would be associated with a discrete depreciation of the reintroduced national currency. Bond liabilities would be repaid in this national currency without the investor being compensated for the depreciation. In other words, the procedure would be as though the bond had been issued in national currency and not in euro. This would ultimately equate

[3] A similar equation is used by Weber (1992) to determine the risk of a realignment of the central parity in the EMS.

to a partial default. However, in a currency union, it is significantly less difficult to effect a partial default by repaying only part of the bond liabilities without abandoning the common currency. Assuming that the bond in the partner country is safe, $E_t(\Delta c_{t+k})$ can be interpreted as the percentage level of the default and π_{kt} as the probability of default, regardless of whether this type of regime change is associated with an exit from the currency union or with a default while maintaining the euro – the latter being regarded in the literature as more likely.[4,5]

The insertion of (3) into (1) taking account of $E_t(\Delta d_{t+k}) = 0$ yields

$$i_t - i_t^* = \pi_{kt} E_t(\Delta c_{t+k}) + \rho_t . \tag{4}$$

The yield spread within a currency union is composed of the expected default and the risk premium. Clearly, in the short history of the euro area, there has been no sovereign default yet. However, even rational investors may assign a positive value to the probability of default despite there having been no prior default event, for instance because the relevant observation period is not deemed long enough. Following the literature on exchange rates, this expected loss component can be termed "peso effect".

4 Econometric approach

In the decomposition of yield spreads into a peso effect, a risk premium and a liquidity premium according to equation (4) and the associated determination of time-varying probabilities of default, we basically adopt the approach of Hallwood et al (2000, hereinafter "HMM") yet modify it in some respects. Following Glosten et al (1993), they use a modified GARCH-in-mean model to describe the risk premium ρ_t. This model also proves appropriate for the present case. As described by Engle et al (1987), risk is positively correlated with the conditional variance of the residuals of an estimate of expected excess returns, h_t, the ARCH-in-mean term, if excess returns are normally

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[4] W Buiter, Sovereign default in the eurozone and the breakup of the eurozone: Sloppy Thinking 101, Financial Times, 14 January 2009, argues that the risk of a default or an existing default by a euro-area member is likely to reduce rather than increase the incentive to leave the euro area.

[5] Incidentally, the possibility of the first term of the right-hand side of (3) describing a traditional default without exchange rate change applies not only to a currency union, but also in the context of target zones. Such an interpretation is usually ignored in the literature on target zones, however, presumably mainly because, in the major target-zone systems in recent decades, such as the ERM, changes to the central parity have been much more frequent than defaults.

distributed. According to Glosten et al's (1993) asymmetrical GARCH(1,1)-in-mean specification, h_t is determined using an ARCH(1) term, a GARCH(1) term and a TARCH(1) term. The latter is equivalent to an ARCH(1) term that is multiplied by a dummy variable which assumes the value 1 if the residual of the previous period was negative. The TARCH(1) term takes account of the fact that the variance may asymmetrically depend on the residuals. This is based on the idea that rising spreads may cause greater volatility than falling spreads. To sum up, the risk premium is modelled as follows:

$$\rho_t = \delta h_t$$

$$h_t = \upsilon_0 + \upsilon_1 \varepsilon_{t-1}^2 + \upsilon_2 (\varepsilon_{t-1}^2 | \varepsilon_{t-1} < 0) + \upsilon_3 h_{t-1} \,, \tag{5}$$

where ε_t is the residual of an estimation of equation (4).

Yield spreads on sovereign bonds of alternative euro-area countries over corresponding German government bond yields, each with a maturity of ten years, have been used as endogenous variables $i_t - i_t^*$. The ten countries considered are Austria (AT), Belgium (BE), Spain (ES), Finland (FI), France (FR), Greece (GR), Ireland (IE), Italy (IT), the Netherlands (NL) and Portugal (PT). Data of daily frequency is used.

Uncovered interest parity as expressed in equation (1) assumes homogeneity of domestic and foreign bonds in terms of liquidity. As has been shown by Sgherri/Zoli (2009), however, liquidity concerns played a major role for investors during the recent crisis. We therefore extend the approach by adding a liquidity premium on the right hand side of equation (4). Unlike most of the literature, we thus allow for heterogeneity in the liquidity of bonds. Empirically, we consider two alternative measures of liquidity premia. First, in line with earlier studies for the United States and Germany (cf Longstaff, 2004), we use the difference between yields of 10-year government-guaranteed bonds issued by the German *Kreditanstalt für Wiederaufbau* (KfW) and German government bonds as an overall liquidity measure of EMU bond markets (λ_{1t}). In order to compute a country-specific liquidity premium, λ_1 is multiplied by a parameter γ_j which reflects the country's bonds' sensitivity to EMU liquidity preferences. Parameters γ_j are estimated. Our second measure (λ_{2jt}) is country-specific and based on the difference between bond and (relative) CDS spreads. It is computed as the country specific bond spread vis-à-vis Germany minus the difference between the country's CDS premium and the German CDS premium.[6] The idea behind this measure is that both bond and CDS spreads reflect the same credit risk, but for a number of reasons, only the bond spread includes a liquidity premium.[7]

The expected loss component, the peso effect, is modelled along the lines of HMM. For simplicity, the expected percentage level of the default $E_t(\Delta c_{t+k})$ in (4) is assumed to be constant (= α). Assuming that German government bonds are safe, π_{kt} represents the absolute probability of default for the relevant euro-area country. The probability of default is determined by exogenous variables. A probit transformation restricts the range of values which π_{kt} can assume to the interval [0; 1]. Let z_t be the vector of exogenous variables, β the associated coefficient vector and Φ the normal distribution function. The probability of default is then modelled as

$$\pi_{kt} = \Phi(\beta' z_t) \,. \tag{6}$$

The exogenous variables in z_{jt}, which are supposed to influence the default probability of country j's bonds, are the following: the spread between the yields of corporate bonds with a BBB credit rating and euro-area government bonds, each with a maturity of seven to ten years, x_t; a country-specific measure of financial sector soundness, y_{jt}; as well as an indicator of a country's price competitiveness, q_{jt}. As in Mody (2009), the variable y_{jt} is constructed as the ratio of the Thomson Financial equity index of the country's financial sector divided by Thomson Financial's overall equity index. Thus, a decrease in y_{jt} indicates a weakening of financial sector soundness. In a related manner, the corporate bond spread, x_t, is expected to serve as indicator of the severity of the crisis according to Gerlach et al (2010). The corporate bond spread reflects financing conditions for firms and the macroeconomic growth outlook, which should ultimately determine individual countries' sovereign risk assessment. Corporate bond spreads are only available since 4 February 2002, which limits the observation period to the subsequent period.

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[6] In a few cases, this liquidity measure yields negative values which, in the following, are set equal to zero.

[7] The CDS market is supposed to be much more liquid than the bond market, because the volume of CDS contracts is not fixed and it is easy to enter short positions. In addition, there seems to be a clear lead for CDS prices over credit spreads in the price discovery process; see Blanco et al (2005) and Dötz (2007).

An indicator of a country's price competitiveness, q_{jt}, is used as a third exogenous variable for determining probabilities of default. More specifically, the effective real exchange rate against 19 trading partners based on consumer price indices is normalised to its average since 1975. In order to obtain a relative, effective indicator, the (logarithmic) indicator value for Germany is deducted from the equivalent (logarithmic) real effective exchange rate of the country in question. The indicator based on consumer price indices has the advantage of being available on a monthly basis. As it is assumed that market players cannot forecast future indicator values, the monthly data are not interpolated, but assumed to be constant for all days within a month. Price competitiveness is included to take account of the argument put forward by Mody (2009) that countries' sensitivity to the financial crisis is more pronounced the greater the loss of competitiveness and growth potential. Mody's (2009) reasoning also suggests that interactions may exist between price competitiveness and financial distress. Therefore, an interaction term between the competitiveness indicator and the relative equity index of the financial sector, $q_{jt}y_{jt}$, is included in some specifications.

It is to be expected that the indicator of price competitiveness q_{jt} – given its relatively sticky development – has only a small effect on changes in yield differences over time but instead helps explain yield spreads across countries. In order to be able to take this into account, the model is estimated as a panel as it is done, for instance, in Chanda et al (2005). Two of the explanatory variables, the spread on corporate bonds in the euro area, x_t, as well as one measure of the liquidity premium, λ_{1t}, are identical across countries. They are multiplied by a country dummy (for countries $j = 2,...,10$, $D_j = 1$ for the currently considered country j and $D_j = 0$ otherwise; for the base country $j = 1$, Austria, D_j is always 0). This allows the sensitivity of the yield spreads to the corporate bond spread as well as the liquidity premium to be modelled in a country-specific way.[8] For a given country j, the vector of the explanatory variables for the probability of default is therefore $z_{jt} = (1 \;\; D_j x_t \;\; q_{jt} \;\; y_{jt})'$ or, if an interaction term between the competitiveness indicator and the relative equity index of the financial sector is included,

[8] Alternatively, country dummies were used as fixed effects in vector z_t. However, it emerged that the real exchange rate captures such fixed country effects relatively well, and so country dummies were subsequently omitted. In general, it should be noted that the variables that determine the probability of default enter into the model in a non-linear fashion as a result of the probit transformation. In an estimation of such a non-linear panel, fixed effects, for example, distort the results. However, as the bias is proportional to $1/T$ (cf Arellano/Hahn, 2006) and $T \geq 284$ in the present case, this distortion can be neglected.

z_{jt} = (1 $D_j x_t$ q_{jt} y_{jt} $q_{jt}y_{jt}$)'. Overall, using liquidity measure λ_{1t} and taking into account (5) and (6), equation (4) can be estimated using the system

$$
\begin{aligned}
i_{jt} - i_{DE,t} &= (\gamma + \sum_{j=2}^{10} \gamma_j D_j)\lambda_{1t} + \\
&\quad \{\alpha\Phi[\beta_0 + (\beta_1 + \sum_{j=2}^{10} \beta_{1,j} D_j)x_t + \beta_2 q_{jt} + \beta_3 y_{jt} + \beta_4 y_{jt} q_{jt}] + \delta h_{jt} + \varepsilon_{jt}\}/k \\
h_{jt} &= \upsilon_0 + \upsilon_1 \varepsilon_{j,t-1}^2 + \upsilon_2 (\varepsilon_{j,t-1}^2 \big| \varepsilon_{j,t-1} < 0) + \upsilon_3 h_{j,t-1}\,.
\end{aligned}
\tag{7}
$$

When using liquidity measure λ_{2jt}, $(\gamma + \sum_{j=2}^{10} \gamma_j D_j)\lambda_{1t}$ is replaced by $\gamma\lambda_{2jt}$. To avoid potential problems with endogeneity, all exogenous variables are lagged by one period in equation (7). The estimation method used is – as by HMM – FIML with the BFGS algorithm for non-linear maximisation. Because the heavily overlapping maturities of the endogenous variables mean autocorrelation has to be expected, Newey-West robust standard errors are applied.[9]

The default rate α is either estimated or, alternatively, set exogenously to 0.6. Imposing an exogenous value to the default rate serves two purposes. The default rate chosen by a government often depends more on the willingness to pay rather than the ability, and it is thus determined by domestic policy considerations. Furthermore, an exogenous default rate facilitates the maximisation of the likelihood function, which, in the present case, is difficult owing to multiple non-linearity (ARCH-in-mean term, probit transformation). The value of $\alpha = 0.6$ is taken from Bedford et al (2005), who determined average default rates of 50% and 70% respectively for the defaults of Russia in 2000 and Argentina in 2005.

[9] The number of lags used is set to six. The panel structure of the model may suggest applying instead a cluster robust variance estimator for two dimensions of clusters as proposed by Thompson (2010) and Cameron et al (2006) in order to deal with correlation across countries as well as autocorrelation. However, both Cameron et al (2006) and Thompson (2010) use Monte Carlo simulations to show that double clustering creates a size bias which results in considerable overrejections in small samples. Thus, Thompson (2010) recommends double clustering only for panels where both $N \geq 25$ and $T \geq 25$, while in the present case N = 10. In order to check for robustness nevertheless, we used a cluster robust variance estimator for one of the specifications and found that the estimated variances were only marginally affected.

5 Results

Because the financial crisis can be assumed to have a lasting impact on the coefficients of the estimate, the system (7) was estimated separately for the period prior to and the period since the onset of the financial crisis. We follow Mody (2009) in using the rescue of US investment bank Bear Stearns as the turning point between the two periods, which are thus defined as 17 March 2008 to 30 April 2009 and 4 February 2002 to 14 March 2008.

Table 1a presents the results for the period since the onset of the financial crisis using the country-specific liquidity premium measure λ_{2jt}, Table 1b the results using liquidity premium measure λ_{1t}. In specification (2), the interaction term is added which is not present in specification (1). Generally, plausible and significant coefficients are estimated for the GARCH equation (υ_0 - υ_3). Negative residuals have proven far less persistent than positive ones ($\upsilon_1 > \upsilon_1+\upsilon_2$). The GARCH-in-mean coefficient δ is significantly positive in line with the hypothesis that rising risk leads to larger interest rate spreads.

As expected, a higher liquidity premium in the mean equation raises the sovereign spread significantly ($\gamma > 0$ in table 1a and $\gamma+\gamma_j > 0$ in table 1b).[10] Table 1b also suggests that the sensitivity to liquidity concerns is higher in Belgium, Finland, Greece, Ireland and Italy compared to the other countries. The β_1 coefficients in tables 1a and 1b imply that the perceived probability of default and thus the yield spread over German government bonds rose as the virulence of the financial crisis increased (as measured by rising spreads on corporate bonds; cf $\beta_1+\beta_{1j}$).[11] However, the strength of the response varied. While the probability of default in France and Belgium increased only moderately, its rise was much more pronounced in countries like Greece and Italy.

A fall in the relative equity index of the financial sector indicates growing distress in the financial sector and thus raises sovereign spreads (cf $\beta_3 < 0$). According to β_2, a real appreciation is also associated with a mostly significantly higher probability of default. Lower price competitiveness leads investors to conclude that growth rates could be lower

[10] When using liquidity premium measure λ_{1t}, the base country Austria as well as Spain turn out to be exceptions to this rule (cf table 1b).

[11] As an exception, Austria in specification (1) as well as the Netherlands display a negative sign in table 1a. Such an outcome may be quite reasonable, as is explained in the section on the pre-Bear Stearns results.

and public debt higher in future. However, low competitiveness, which – in the currency union – is the result of high price and wage increases in the recent past, could also be associated with a lack of political will to expect the public to accept financial cutbacks. This would directly indicate a lack of willingness to consolidate and therefore increase the probability of default and thus interest rate spreads. The negative coefficient on the interaction term included in specifications (2), β_4, shows that the sensitivity to an increased virulence of the financial crisis has been more pronounced the lower the price competitiveness of the country considered.

In the estimation for the period prior to the financial crisis, the liquidity premium measure λ_{1t} has generally been used because, due to data limitations, λ_{2jt} is available for the post-Bear Stearns period only. The estimation results for the pre-crisis period are shown in Table 2. Only the results for the default rate, α, being exogenously set to 0.6 are shown because estimated default rates were implausibly small (see also the discussion on α in chapter 4). In terms of their sign, the results often do not differ much from those for the period since. One significant deviation, however, relates to the coefficient of the spreads for corporate bonds, which is negative for many countries in the period prior to the crisis, examples being Austria, Ireland, the Netherlands, Portugal and Spain (see eg β_1 or $\beta_1+\beta_{1,NL}$ respectively). In these countries, an increase in corporate spreads is likely to have been regarded less as a warning signal about the stability of the economy as a whole than as a company- or industry-specific issue at that time. In this case, investors are likely to have restructured their portfolios partly in favour of supposedly safe government bonds, thereby reducing their return. In that sense, β_1 is dominated by a substitution effect in some countries of the pre-crisis period.

As a second notable deviation from the post-Bear Stearns results, the coefficient for the real exchange rate in specification (2), β_2, is significantly negative in the pre-crisis period and that for the interaction term, β_4, is significantly positive. At this time, a real appreciation brought about by relatively high domestic inflation has obviously not been perceived as a sign of mounting problems but – in a more shortsighted view – rather as reflecting the often concomitant dynamic growth, which would facilitate the repayment of government debt. This result contrasts somewhat with Mody (2009), who found that the real exchange rate only had an impact on yield differentials in the euro area during the crisis.

The probabilities of default which can be calculated from the estimated model relate to a default event within the next ten years. They can be converted into probabilities of default within a one-year period using the formula

$$\pi_{1,t} = 1 - (1 - \pi_{10,t})^{1/10}. \qquad (8)$$

For the Netherlands, Ireland and Spain, this probability of default did not differ perceptibly from zero in the period prior to the crisis. For the other euro-area states, the likelihood of a default within a year was also very low in the pre-crisis period. For Italy, for example, the figure is less than 0.2% for most of the time.

During the financial crisis, the probability of default rose significantly in most countries (but not in Austria and the Netherlands), peaking in March 2009, and then dropped off again. This is exemplified by Italy in figure 2. The maximum probabilities of default reached within a one-year period (in %) are shown in Table 3 for all the countries in the sample. The probabilities of default in all euro-area countries, being close to zero previously, have risen considerably since the onset of the crisis. In addition, the table also demonstrates that the probabilities of default have fanned out significantly since the onset of the crisis, a result which is in line with Mody's (2009) findings, for example.

Nevertheless, the figures should be interpreted with caution. It should be stressed that the probabilities of default are ultimately calculated from observed yield differentials and therefore reflect the situation adequately only if one believes that the market is capable of doing so during the crisis, which was at times marked by panic.

Figure 3 illustrates to what extent the observed yield differentials during the crisis can be attributed to the peso effect of a default (the expected loss component) and to what extent they are the result of the liquidity premium or a risk premium, which reflects uncertainty about the expected return of the investment. Evidently, during the crisis the peso effect dominated interest-rate differentials, especially in countries where yield spreads were high. In Austria and the Netherlands, the risk premium makes a substantial contribution to the spread over German government bonds, particularly at the current end. This may hint at speculative pressure against these countries. The liquidity premium played an especially important role in Finland, France and Portugal.

6 Conclusions

This paper presents a new approach for analysing recent movements in EMU sovereign bond spreads. Based on a GARCH-in-mean model originally used in the target zone literature, spreads are decomposed into a risk premium, an expected loss component and a liquidity premium. Time-varying probabilities of default are derived. While the model could, in principle, also be applied to bonds with shorter maturities, we focus on long-term bonds with a maturity of 10 years due to the role of German Bunds as benchmark bonds. The structure of the model is general enough to be applicable to other countries or regions and other observation periods as well.

The results suggest that market perceptions of sovereign risk changed after the rescue of Bear Stearns in March 2008. As a result, the government bonds of some countries lost their previous role as a domestic safe haven. In the period prior to the Bear Stearns rescue, implied probabilities of default were negligible. The subsequent strong rise in several euro-area sovereign bond spreads mainly reflects an increased expected loss component. As an example, the implied probability of default for Irish sovereign bonds amounted to more than 6% at its peak. Important determinants of sovereign spreads, which are responsible for the rise in the expected loss, are a country's financial sector soundness and its price competitiveness. Interestingly, the combined effect of both variables has also proved important for spread developments during the crisis period considered. This suggests that price competitiveness moved into investors' focus as financial sector soundness weakened. Risk and liquidity premia generally played a minor part in spread widening of countries with high yield spreads, such as Greece or Italy. While there are signs that risk premia had an effect, particularly in Austria and the Netherlands, liquidity premia seem to have been most important in Finland, France and Portugal. The often dominant role of the expected loss component reflects the importance of fundamental country-specific factors as compared with global factors such as investors' general risk aversion.

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Table 1a: Estimation for the period since the rescue of Bear Stearns (17.03.2008 – 30.04.2009), liquidity measure: λ_{2jt} (difference between bond and relative CDS spreads)

Variable	Coefficient (standard error) $\alpha = 0.6$ no interaction term	Coefficient (standard error) $\alpha = 0.6$ interaction term present	Coefficient (standard error) α estimated no interaction term	Coefficient (standard error) α estimated interaction term present
α			0.27* (0.03)	0.36* (0.11)
β_0	0.78* (0.08)	-0.31* (0.01)	1.61* (0.35)	-0.11 (0.56)
β_1	-3.19* (0.59)	11.46* (0.31)	-1.70 (0.93)	13.19* (1.76)
$\beta_{1,BE}$	8.60* (0.48)	-2.19* (0.49)	9.98* (0.85)	-2.32* (1.14)
$\beta_{1,ES}$	13.21* (0.65)	-0.25 (0.32)	16.12* (1.00)	-0.73 (0.89)
$\beta_{1,FI}$	18.21* (0.57)	2.46* (0.56)	20.00* (1.47)	1.84 (1.77)
$\beta_{1,FR}$	6.84* (0.49)	-5.28* (0.45)	7.08* (0.67)	-5.99* (0.73)
$\beta_{1,GR}$	24.94* (0.67)	10.91* (0.39)	33.66* (2.80)	12.58* (1.80)
$\beta_{1,IE}$	10.00* (0.87)	-4.02* (0.57)	15.41* (1.41)	-6.35 (3.41)
$\beta_{1,IT}$	20.76* (0.60)	8.14* (0.45)	25.45* (1.54)	9.10* (1.57)
$\beta_{1,NL}$	-4.73* (0.75)	-13.94* (2.82)	-5.18* (1.40)	-14.90* (2.28)
$\beta_{1,PT}$	11.12* (0.68)	-2.08* (0.44)	14.62* (0.89)	-3.59 (2.03)
β_2	0.52* (0.15)	6.22* (0.27)	0.02 (0.17)	11.33 (7.47)
β_3	-0.62* (0.02)	-0.40* (0.003)	-0.73* (0.07)	-0.41* (0.12)
β_4		-1.14* (0.05)		-2.18 (1.49)
γ	0.64* (0.01)	0.58* (0.02)	0.59* (0.02)	0.60* (0.02)
δ	34.66* (1.91)	15.20* (2.21)	24.69* (2.12)	17.38* (1.99)
υ_0	0.00* (0.00)	0.00* (0.00)	0.00* (0.00)	0.00* (0.00)
υ_1	0.82* (0.05)	0.94* (0.08)	1.02* (0.09)	0.94* (0.08)
υ_2	-0.74* (0.05)	-0.58* (0.08)	-0.82* (0.07)	-0.62* (0.09)
υ_3	0.54* (0.01)	0.44* (0.05)	0.48* (0.02)	0.45* (0.03)

A star indicates significance at the 5% level.

Table 1b: Estimation for the period since the rescue of Bear Stearns (17.03.2008 – 30.04.2009), liquidity measure: λ_{1t} (difference between KfW and Bund yield)

Variable	Coefficient (standard error) $\alpha = 0.6$ no interaction term	Coefficient (standard error) $\alpha = 0.6$ interaction term present	Coefficient (standard error) α estimated no interaction term	Coefficient (standard error) α estimated interaction term present
α			0.43* (0.01)	0.16* (0.01)
β_0	0.76* (0.09)	0.97* (0.04)	1.53* (0.01)	1.89* (0.66)
β_1	18.40* (0.50)	18.35* (0.19)	20.09* (0.32)	30.79* (1.58)
$\beta_{1,BE}$	-9.93* (0.59)	-10.42* (0.54)	-11.88* (0.59)	-16.32* (2.34)
$\beta_{1,ES}$	-2.87* (0.60)	-2.71* (0.28)	2.70* (0.21)	-2.74 (1.45)
$\beta_{1,FI}$	2.53* (0.94)	2.93* (0.54)	3.19* (0.37)	-1.42 (2.53)
$\beta_{1,FR}$	-6.43* (0.49)	-6.50* (0.35)	-7.49* (0.29)	-12.28* (1.94)
$\beta_{1,GR}$	6.85* (1.01)	7.36* (0.38)	9.53* (0.37)	26.99* (3.97)
$\beta_{1,IE}$	-13.70* (1.00)	-14.32* (0.45)	-15.37* (0.31)	-25.92* (2.00)
$\beta_{1,IT}$	0.83* (0.42)	1.88* (0.44)	1.30* (0.37)	3.87* (1.73)
$\beta_{1,NL}$	-8.90* (0.43)	-9.25* (0.35)	-10.42* (0.22)	-17.24* (7.73)
$\beta_{1,PT}$	-7.09* (0.81)	-6.93* (0.46)	-6.87* (0.55)	-11.59* (1.53)
β_2	1.37* (0.65)	1.59* (0.12)	0.82* (0.01)	14.04* (3.80)
β_3	-0.65* (0.01)	-0.69* (0.01)	-0.79* (0.01)	-0.81* (0.14)
β_4		-0.14* (0.05)		-2.74* (0.83)
γ	-1.40* (0.35)	-1.30* (0.11)	-1.32* (0.23)	-0.97* (0.29)
γ_{BE}	5.22* (0.35)	5.20* (0.30)	5.29* (0.36)	5.33* (0.79)
γ_{ES}	0.63* (0.32)	0.79* (0.21)	0.82* (0.21)	0.41 (0.39)
γ_{FI}	5.18* (0.25)	4.98* (0.17)	5.07* (0.18)	4.82* (0.36)
γ_{FR}	2.59* (0.18)	2.48* (0.15)	2.57* (0.17)	2.74* (0.29)
γ_{GR}	4.93* (1.01)	5.54* (0.54)	5.36* (0.39)	1.58* (0.62)
γ_{IE}	4.69* (0.74)	5.28* (0.37)	5.49* (0.31)	5.71* (0.92)
γ_{IT}	5.24* (0.34)	5.20* (0.46)	5.28* (0.40)	4.70* (0.40)
γ_{NL}	1.41* (0.16)	1.39* (0.15)	1.44* (0.17)	1.84* (0.25)
γ_{PT}	2.88* (0.60)	3.08* (0.28)	3.03* (0.29)	2.94* (0.51)
δ	4.19* (1.01)	3.97* (0.65)	4.31* (0.91)	10.04* (1.63)
υ_0	0.00* (0.00)	0.00* (0.00)	0.00* (0.00)	0.00* (0.00)
υ_1	1.02* (0.09)	1.03* (0.07)	1.04* (0.09)	1.09* (0.07)
υ_2	-0.40* (0.12)	-0.40* (0.08)	-0.42* (0.11)	-0.69* (0.09)
υ_3	0.29* (0.05)	0.29* (0.04)	0.28* (0.04)	0.38* (0.04)

A star indicates significance at the 5% level.

Table 2: Estimation for the period until the rescue of Bear Stearns (04.02.2002 – 14.03.2008)

Variable	Coefficient (standard error) $\alpha = 0.6$ no interaction term	Coefficient (standard error) $\alpha = 0.6$ interaction term present
α		
β_0	-0.67* (0.02)	0.81 (0.50)
β_1	-32.08* (4.80)	-10.65* (3.44)
$\beta_{1,BE}$	52.73* (4.39)	31.59* (3.22)
$\beta_{1,ES}$	-67.66* (8.55)	-523.71* (17.12)
$\beta_{1,FI}$	57.28* (4.49)	32.64* (3.11)
$\beta_{1,FR}$	39.19* (5.02)	18.84* (3.53)
$\beta_{1,GR}$	37.84* (4.74)	35.14* (3.20)
$\beta_{1,IE}$	-72.21* (21.73)	-98.54* (9.89)
$\beta_{1,IT}$	49.72* (4.61)	36.63* (3.99)
$\beta_{1,NL}$	-40.80* (12.49)	-50.26* (12.70)
$\beta_{1,PT}$	-31.58* (5.80)	1.64 (3.56)
β_2	4.86* (0.17)	-34.94* (4.13)
β_3	-0.39* (0.01)	-0.71* (0.11)
β_4		8.15* (0.87)
γ	1.61* (0.20)	1.06* (0.18)
γ_{BE}	-0.33 (0.25)	-0.06 (0.24)
γ_{ES}	1.67* (0.30)	2.93* (0.24)
γ_{FI}	0.26 (0.26)	1.48* (0.21)
γ_{FR}	-0.26 (0.35)	0.29 (0.26)
γ_{GR}	-0.19 (0.44)	-0.81* (0.38)
γ_{IE}	-0.57 (0.84)	1.39* (0.22)
γ_{IT}	4.11* (0.76)	9.73* (0.66)
γ_{NL}	0.73* (0.26)	1.36* (0.25)
γ_{PT}	4.99* (0.31)	3.42* (0.36)
δ	27.01* (2.66)	5.72* (1.00)
υ_0	0.00* (0.00)	0.00* (0.00)
υ_1	0.66* (0.03)	0.53* (0.03)
υ_2	-0.32* (0.03)	-0.06* (0.02)
υ_3	0.49* (0.02)	0.49* (0.03)

A star indicates significance at the 5% level.

Table 3: Market perceptions of maximum probabilities of default (%) within a one-year period during the financial crisis according to the estimation results shown in the first column of table 1a (liquidity measure: λ_{2jt}, default rate: 60%, no interaction term)

AT	BE	ES	FI	FR	GR	IE	IT	NL	PT
0.2	1.1	1.4	0.5	0.5	3.5	6.4	2.2	0.2	1.4

Figure 1: Bond spreads on euro-area government bonds versus Germany

Countries: Greece, Ireland, France, Italy; maturity: 10 years

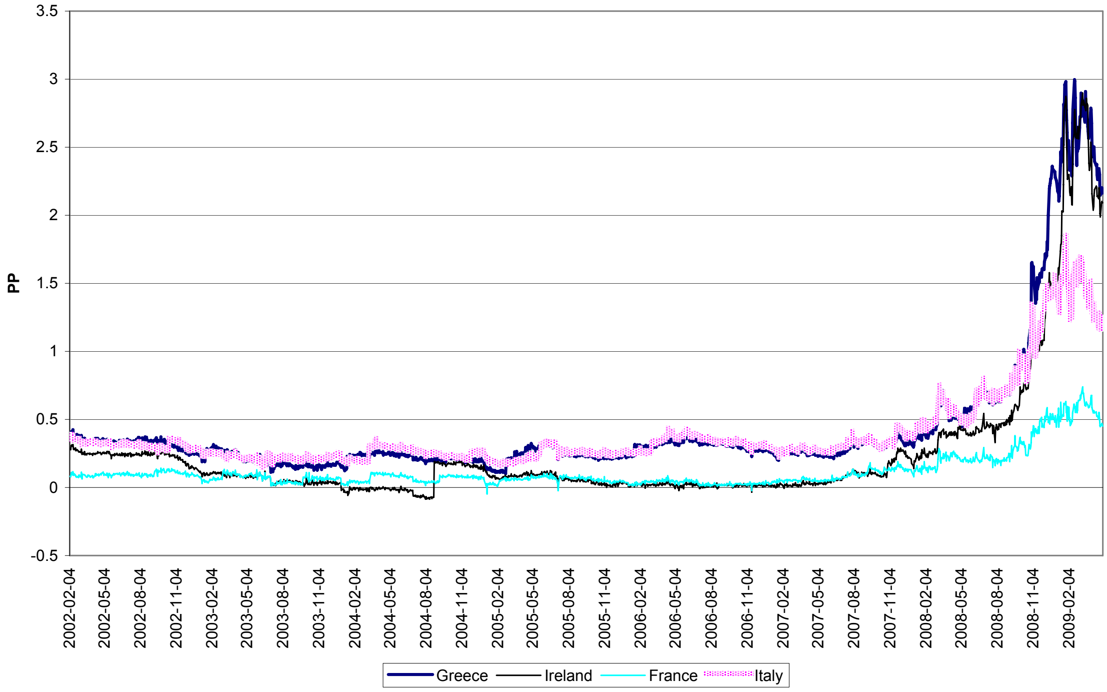


Figure 2: Probability of default for Italian sovereign bonds with a maturity of ten years based on the estimation results shown in the first column of table 1a

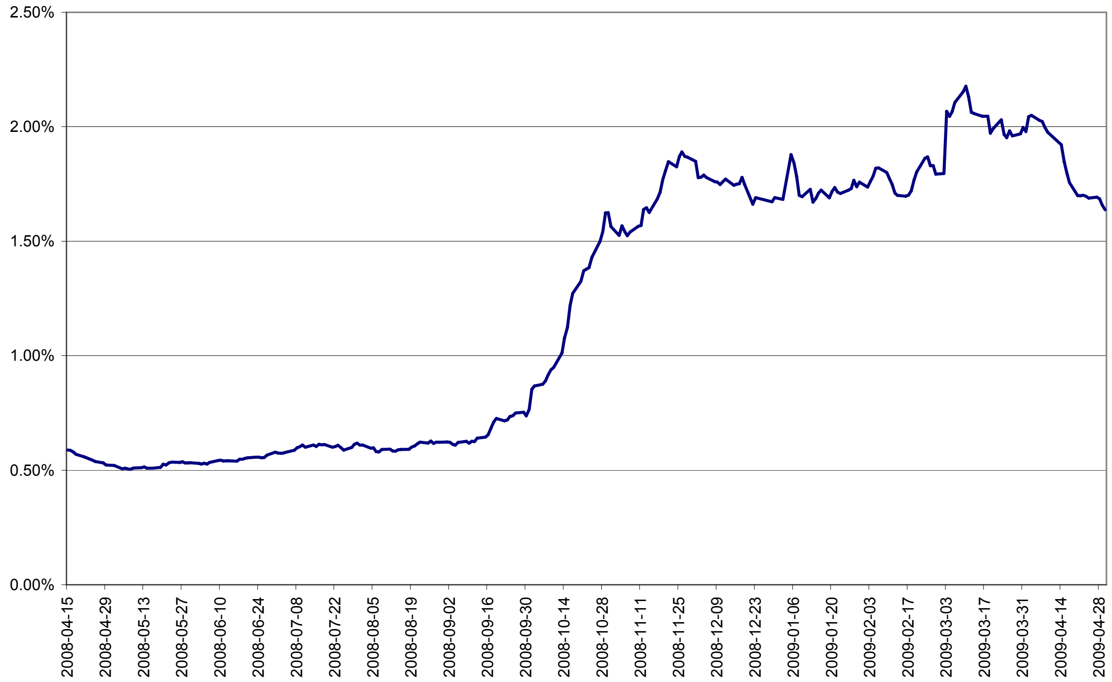
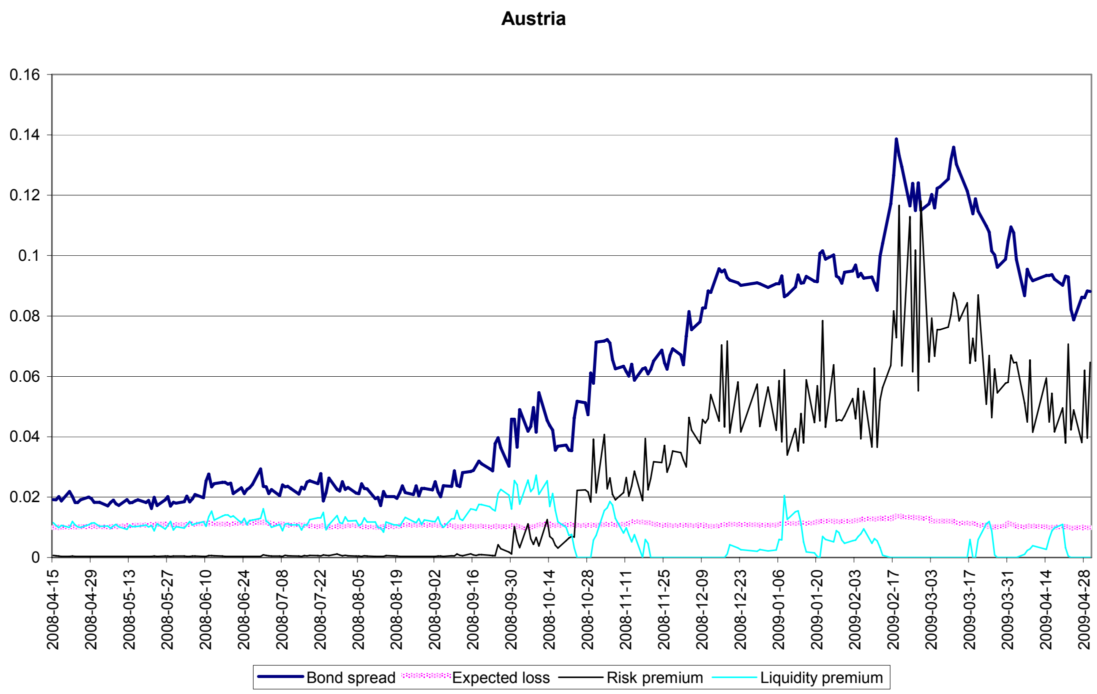
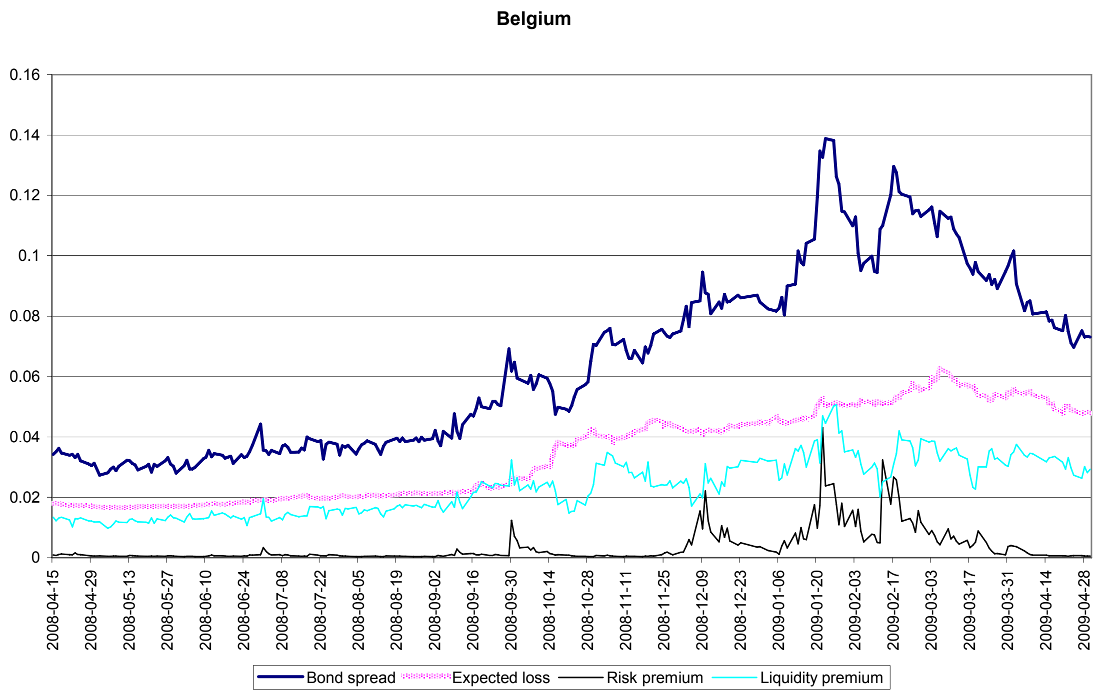
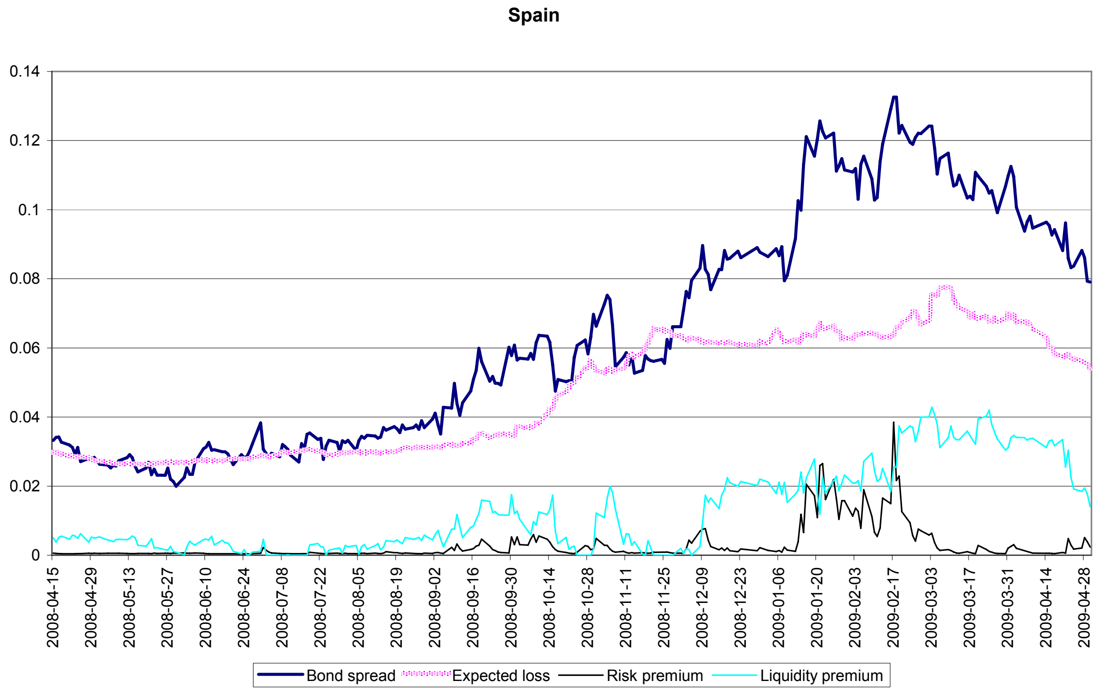


Figure 3: Decomposition of sovereign bond spreads (cumulated over 10 years) based on the estimation results shown in the first column of table 1a[12]





[12] Deviations between bond spread and the sum of expected loss, risk premium and liquidity premium are caused by residuals ε_{jt}.


Spain
0.14
0.12
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0.08
0.06
0.04
0.02
0
2008-04-15
2008-04-29
2008-05-13
2008-05-27
2008-06-10
2008-06-24
2008-07-08
2008-07-22
2008-08-05
2008-08-19
2008-09-02
2008-09-16
2008-09-30
2008-10-14
2008-10-28
2008-11-11
2008-11-25
2008-12-09
2008-12-23
2009-01-06
2009-01-20
2009-02-03
2009-02-17
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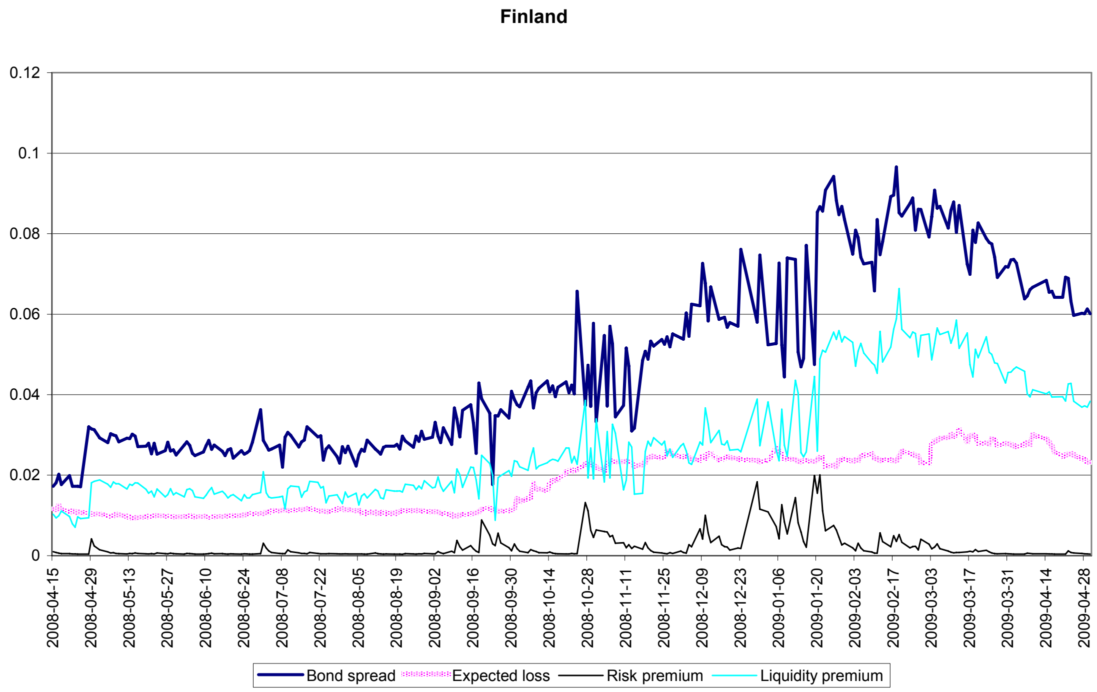
2009-03-17
2009-03-31
2009-04-14
2009-04-28
Bond spread
Expected loss
Risk premium
Liquidity premium


Finland
0.12
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2008-04-15
2008-04-29
2008-05-13
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2008-06-10
2008-06-24
2008-07-08
2008-07-22
2008-08-05
2008-08-19
2008-09-02
2008-09-16
2008-09-30
2008-10-14
2008-10-28
2008-11-11
2008-11-25
2008-12-09
2008-12-23
2009-01-06
2009-01-20
2009-02-03
2009-02-17
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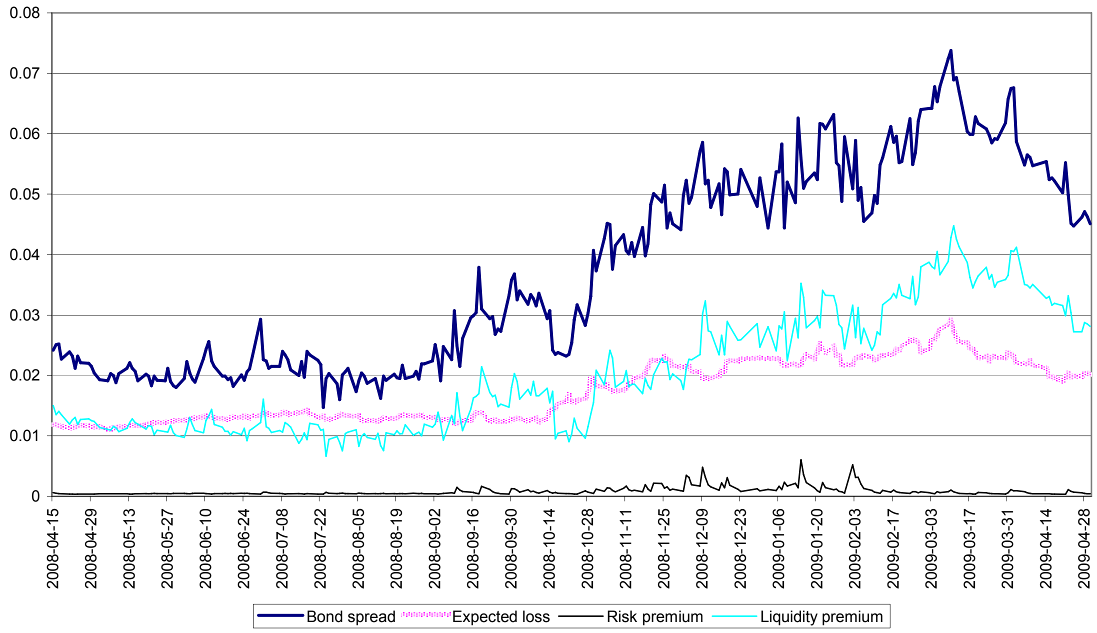
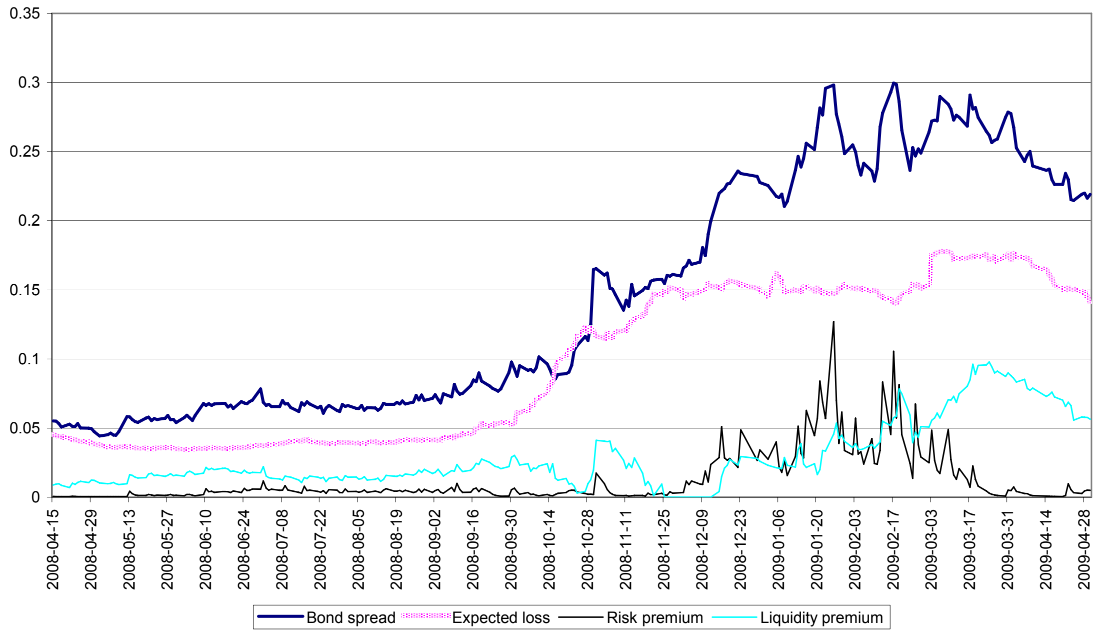
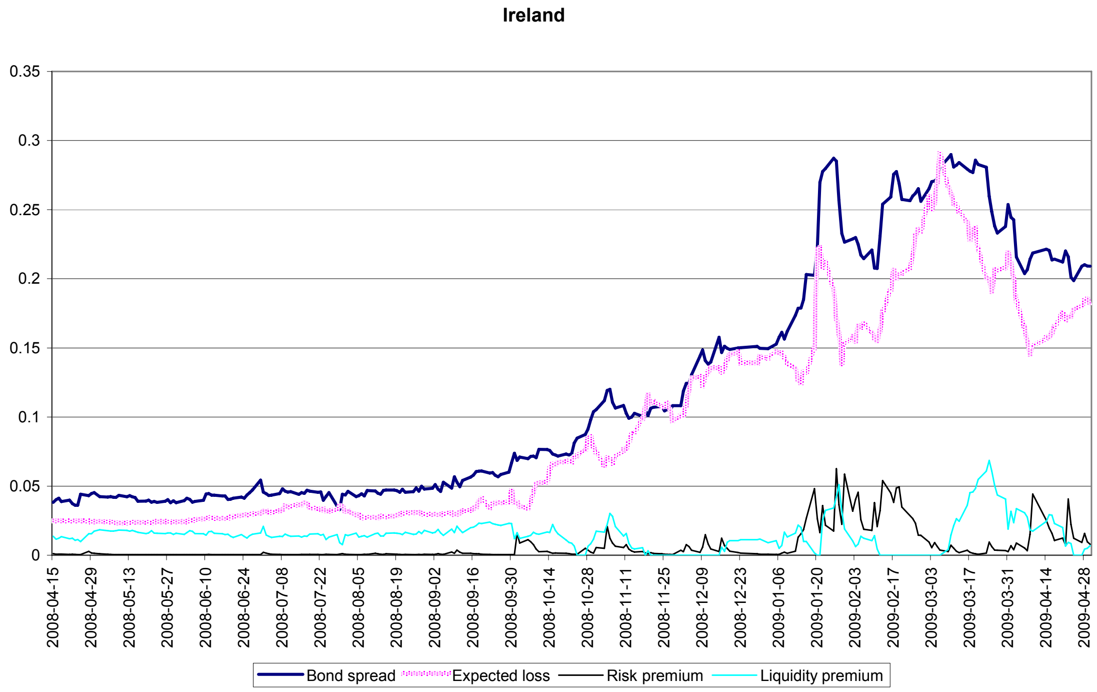
2009-04-14
2009-04-28
Bond spread
Expected loss
Risk premium
Liquidity premium

France



Greece




Ireland
0.35
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2008-05-27
2008-06-10
2008-06-24
2008-07-08
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2008-08-05
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2008-09-02
2008-09-16
2008-09-30
2008-10-14
2008-10-28
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2009-02-03
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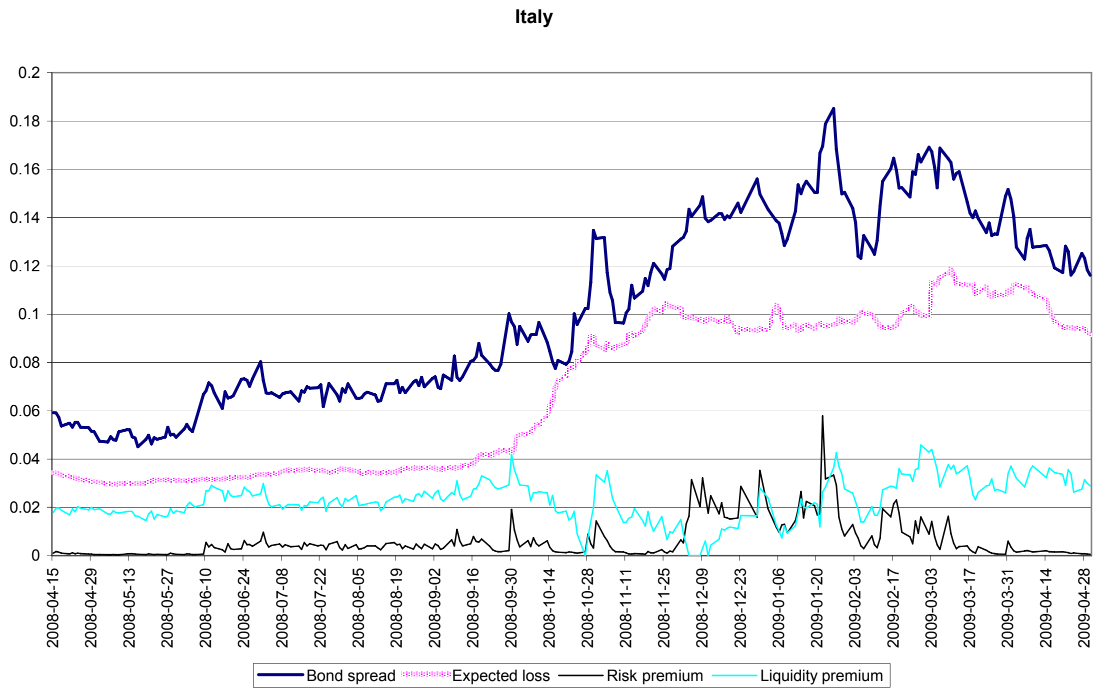
2009-04-28
Bond spread
Expected loss
Risk premium
Liquidity premium


Italy
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2008-04-15
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2009-01-06
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2009-02-03
2009-02-17
2009-03-03
2009-03-17
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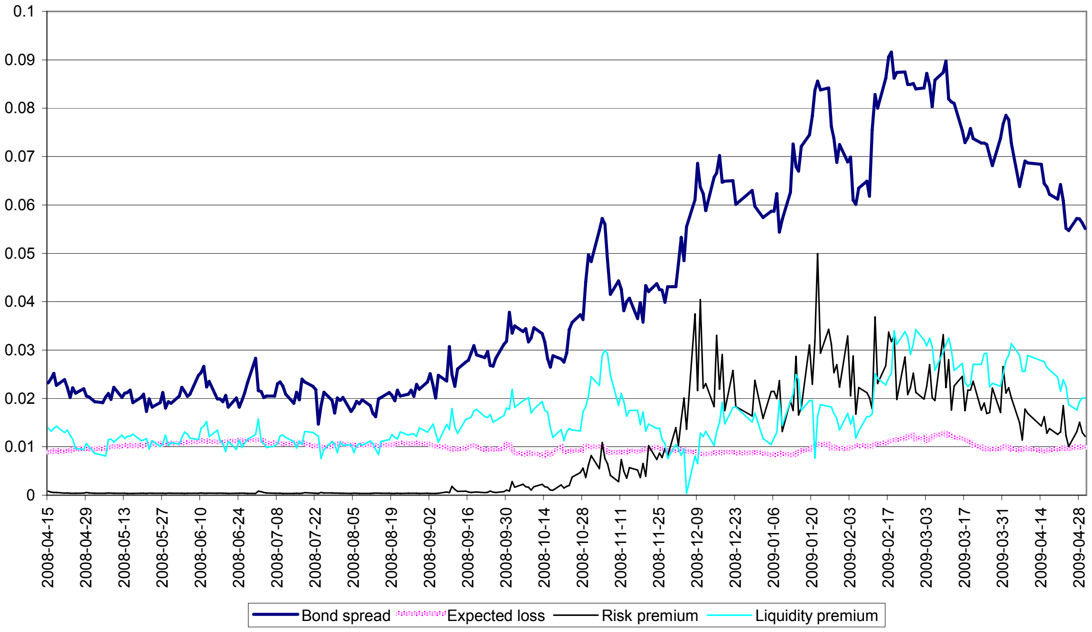
2009-04-14
2009-04-28
Bond spread
Expected loss
Risk premium
Liquidity premium

Netherlands



Portugal