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## Market perception of sovereign credit risk in the euro area during the financial crisis

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## WORKING PAPER SERIES

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# MARKET PERCEPTION OF SOVEREIGN CREDIT RISK IN THE EURO AREA DURING THE FINANCIAL CRISIS

Gonzalo Camba-Méndez and Dobromiř Serwa

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**NOTE:** This Working Paper should not be reported as representing the views of the European Central Bank (ECB). The views expressed are those of the authors and do not necessarily reflect those of the ECB.

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## **Abstract**

We study market perception of sovereign credit risk in the euro area during the financial crisis. In our analysis we use a parsimonious CDS pricing model to estimate the probability of default (PD) and the loss given default (LGD) as perceived by financial markets. We find that separate identification of PD and LGD appears empirically tractable for a number of euro area countries. In our empirical results the estimated LGDs perceived by financial markets stay comfortably below 40% in most of the samples. We also find that macroeconomic and institutional developments were only weakly correlated with the market perception of sovereign credit risk, whereas financial contagion appears to have exerted a non-negligible effect.

JEL-Classification: C11, C32, G01, G12, G15.

Keywords: sovereign credit risk, CDS spreads, euro area, probability of default, loss given default.

## Executive summary

The aim of this paper is to study market perception of sovereign credit risk in the euro area during the financial crisis. Understanding the factors that influence changes in market perception of sovereign credit risk is a question of special interest for economists, investors, and financial regulators. Similarly to a number of recent studies, we also monitor the market perception of sovereign risk by using information derived from CDS spreads.

In studying market perceptions of sovereign credit risk, one of the aims of this paper is to estimate measures of the probability of default (PD) as perceived by financial markets and the loss given default (LGD), also as perceived by financial markets, embedded in the CDS spreads. Earlier studies of the euro area sovereign debt crisis usually employed the CDS spreads alone and thus were not able to separately identify changes in the valuation of risk stemming from either increases of potential losses or a growing likelihood of default. Yet, in the pricing framework of ‘*fractional recovery of face value*’, the factors governing the dynamics of the PD and LGD play very distinct roles in the derivation of the price of the CDS contract, thus suggesting that PD and LGD can indeed, at least in theory, be separately identified. However, what is conceptually true may not always hold empirically, and to date the separate identification of PD and LGD remains challenging.

The second aim of this paper is to shed some light on the key drivers of the market perception of sovereign credit risk. In contrast with earlier studies, we aim not only to study the main economic drivers but also the potential impact of the numerous institutional changes introduced. For example, the European Financial Stability Facility (EFSF) was created to enable financing of the euro area member states that were in difficulty. Additionally, a new collection of measures to enforce fiscal discipline, referred to as the ‘*fiscal compact*’, was introduced. At the same time, and with a view to restoring the normal functioning of financial markets, the ECB introduced a number of non-standard monetary policy measures such as for example the Securities Markets Programme.

Beyond economic fundamentals and institutional factors, the phenomenon of contagion was also quoted as being at the centre of developments in the euro area sovereign debt markets. Correspondingly, much of the recent research on contagion during the financial crisis has studied ‘spillovers’ from the Greek sovereign debt crisis to other euro area sovereign debt markets. In this paper we adopt a more general approach and search for evidence of abnormal spillovers across euro area countries beyond those justified by economic fundamentals,

and not only financial spillovers from Greece.

In our empirical investigation we are able to identify PDs and LGDs for most countries and periods. Market implied measures of PD and LGD are computed for Germany, France, Greece, Ireland, Italy, Portugal and Spain. For a number of euro area countries, notably France, Ireland and Italy, the separate identification of PD and LGD appears empirically tractable. However, and at times of excessively high CDS spread levels, this appears slightly less so for Greece and Portugal. Identifiability appears more challenging altogether, however, for the cases of Germany and Spain.

With these caveats in mind, our reported estimates of PD and LGD reveal several interesting results. In contrast with the literature on corporate credit risk, the dynamics of the estimated PD and LGD are not always strongly positively correlated. Our estimated LGDs also show that high levels of LGD are often associated with large stocks of public debt. This is something broadly to be expected and aligned with the rationale of the haircuts derived from the simple static debt sustainability analysis proposed by the IMF. For the euro area countries under investigation here, the estimated LGD was for most of the sample comfortably below 40%. Only at a few points in time, and in the sample under study, was the estimate of LGD for Portugal recorded as surpassing 50%, which is the reported average LGD of all sovereign debt restructurings that took place in the world economy between 1970 and 2010.

Our estimates of PD and LGD are only weakly related to economic and institutional factors. We also find evidence of possible contagion from developments in the perceptions of credit risk in one euro area country to other euro area countries. The nature of contagion is, however, possibly not as unidirectional and simple as it is usually represented by the financial media, namely from Greece to the rest, because the coincidental shocks occurred on dates related to different crisis events.

# 1 Introduction

In the early stages of the crisis, the financial problems identified in a number of large banks triggered support recapitalization programs by various euro area governments. When the slowdown in the global economy became more apparent, and when the macroeconomic outlook for some euro area countries turned very pessimistic, international investors reacted nervously and started to seriously question the ability of many euro area governments to repay their debts. As a result, tensions in several euro area sovereign debt markets escalated to such a degree that the actual ability of some euro area governments to roll-over their debt was seriously hampered. Volatility levels, liquidity conditions and yield spreads reached historical peaks and reflected market malfunctioning. It was in this context that the European authorities decided to introduce a number of policy measures to deal with these tensions.

The aim of this paper is to study market perception of sovereign credit risk in the euro area during the financial crisis, and in particular, identifying the main factors that influence changes in market perception of sovereign credit risk. This is a question of special interest for economists, investors, and financial regulators. Similarly to a number of recent studies, we also monitor the market perception of sovereign risk by using information derived from CDS spreads (Longstaff, Pan, Pedersen, and Singleton, 2011; Bei and Wei, 2012; Ang and Longstaff, 2013). CDS spreads are a more accurate measure to gauge market perceptions of sovereign credit risk than are sovereign bond yield spreads. This is so because movements in sovereign bond yield spreads during the financial crisis also reflected liquidity distortions, limited arbitrage operations amid increasing risk aversion and the official interventions of the ECB, all of which were less apparent in the CDS market (e.g., Fontana and Scheicher, 2010).

In studying market perceptions of sovereign credit risk, one of the aims of this paper is to estimate measures of the probability of default (PD) and the loss given default (LGD) as perceived by financial markets embedded in the CDS spreads. Earlier studies of the euro area sovereign debt crisis usually employed the CDS spreads alone and thus were not able to separately identify changes in the valuation of risk stemming from either increases of potential losses or a growing likelihood of default (Blommestein, Eijffinger, and Qian, 2012; Beirne and Fratzscher, 2013). Yet, in the pricing framework of ‘*fractional recovery of face value*’, the factors governing the dynamics of the PD and LGD play very distinct roles in the derivation of the price of the CDS contract, thus suggesting that PD and LGD can indeed, at least in theory, be separately identified. However, as pointed by Pan and Singleton (2008), what is

conceptually true may not always hold empirically, and to date the separate identification of PD and LGD remains challenging. In this paper we show that, for a number of euro area countries, separate identification of PD and LGD appears empirically tractable.

The second aim of this paper is to shed some light on the key drivers of the market perception of sovereign credit risk. In contrast with earlier studies, we aim not only to study the main economic drivers, but also the potential impact of the numerous institutional changes introduced. Sovereign CDS spreads for the euro area countries were subject to large fluctuations during the financial crisis. This was understandable to a certain extent, as the recent financial crisis brought with it the largest recorded decline of economic activity and one of the fastest recorded increases in public debt since the second world war. However, economic fundamentals cannot be the sole explanation, as even in the euro area countries that had solid economic fundamentals, the sovereign CDS spreads were at historical record highs, and the functioning of the markets was affected by stress. This suggests that concerns over the institutional framework of the European Monetary Union, and in particular the lack of credibility of the EU rules to enforce fiscal discipline, were possibly being priced by the markets.<sup>1</sup> To respond to this, a number of important institutional changes were introduced. For example, the European Financial Stability Facility (EFSF) was created to enable financing of the euro area member states that were in difficulty. Additionally, a new collection of measures to enforce fiscal discipline, referred to as the '*fiscal compact*', was introduced. In contrast to the rules of the Stability and Growth Pact in force, the fiscal compact would have to be documented in primary legislation at the country level, thus rendering it more credible. At the same time, and with a view to restoring the normal functioning of financial markets, the ECB introduced a number of non-standard monetary policy measures such as the Covered Bond Purchase Programmes and the Securities Markets Programme.

Beyond economic fundamentals and institutional factors, the phenomenon of contagion was also quoted as being at the centre of developments in the euro area sovereign debt markets. Some of the European Central Bank (ECB) interventions were even said to have been motivated by the specific need to address contagion (Constancio, 2012). Correspondingly, much of the recent research on contagion during the financial crisis has studied 'spillovers' from the Greek sovereign debt crisis to other euro area sovereign debt markets (e.g., Mink and

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<sup>1</sup>Budget Balance Laws had been enshrined in the Treaty of the European Union signed in Maastricht in February 1992. In particular, article 104 of the Treaty stated that member states shall avoid excessive government deficits. Further to this, in June 1997, the European Council had passed a resolution on a Stability and Growth Pact, by which member states committed themselves to respect the medium-term budgetary objective of positions close to balance or in surplus as set out in their stability programme.



de Haan, 2013). In this paper we adopt a more general approach and search for evidence of abnormal spillovers across euro area countries beyond those justified by economic fundamentals, and not only financial spillovers from Greece.

In our empirical investigation we are able to identify PDs and LGDs for most countries and periods. Notably, we find that the estimated PDs vary considerably in time, and reach values close to one during the most turbulent periods. Meanwhile, the estimated LGD stays comfortably below 40% in most of the samples. We also find that the estimates of PD and LGD are only weakly related to macroeconomic and institutional factors. Our empirical results also show that financial spillovers beyond fundamentals, i.e. contagion, actually had a non-negligible impact on the market's perception of sovereign credit risk.

This paper is organised as follows. Section 2 describes our modeling strategy to estimate PD and LGD. Section 3 studies the main drivers underlying the market's perception of sovereign credit risk. Finally, section 4 concludes.

## 2 Measuring market's perception of credit risk

We adopt a modelling framework similar to that employed in Doshi (2011), where LGD is time varying. This time varying LGD is aligned with empirical observations and similarly with the current reporting of 'recovery ratings' for sovereigns by rating agencies (e.g., Das, Papaioannou, and Trebesch, 2012).<sup>2</sup>

### 2.1 Pricing CDS contracts

We follow the discrete time risk neutral valuation approach of Gouriou, Monfort, and Polimenis (2006) and Doshi (2011) to estimate the probability of default (PD) and the loss given default (LGD) perceived by financial markets. The default time  $\tau$  is modeled as a surprise event driven by a homogeneous Poisson process with the time-varying intensity parameter  $\lambda_t$ . The expected probability at time  $t$  of surviving at least  $h$  periods is defined by:

$$E_t^Q \{1_{(\tau > t+h)}\} = E_t^Q \left\{ \exp \left( \sum_{k=1}^h -\lambda_{t+k-1} \right) \right\}$$

where  $E_t^Q$  denotes the expected value computed under the risk-neutral measure, and  $1_{(\cdot)}$  is an indicator function equal to one when its argument is true, and equal to zero otherwise. The

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<sup>2</sup>Standard & Poor's first started providing recovery ratings for non-investment grade sovereigns in 2007.

PD is then simply defined as one minus the probability of surviving. It is further assumed that  $\lambda_t$  is a quadratic function of an unobservable factor  $x_t$ ,  $\lambda_t = x_t^2$ , thus enforcing positive survival and default probabilities. We further assume, as is commonly encountered in the literature, that the CDS contracts are priced under the assumption of ‘fractional recovery of face value’ of the debt. Thus, in the event of default, CDS holders will recover the fraction of the face value that bond issuers fail to pay. The LGD is allowed to vary over time and depends quadratically on an unobservable factor  $z_t$ . Its expected value under the risk neutral measure is assumed to be:

$$E_t^Q \{LGD_{t+j}^Q\} = E_t^Q \{\exp(-z_{t+j}^2)\}.$$

Once again, this functional setting guarantees that  $LGD_t$  will lie in the range  $[0, 1]$ . As in Doshi (2011), the factors  $x_t$  and  $z_t$  are assumed to follow autoregressive processes with correlated residuals:

$$\mathbf{s}_t = \boldsymbol{\mu} + \mathbf{\Gamma} \mathbf{s}_{t-1} + \boldsymbol{\epsilon}_t, \quad (1)$$

where:

$$\mathbf{s}_t = \begin{pmatrix} x_t \\ z_t \end{pmatrix}, \quad \boldsymbol{\mu} = \begin{pmatrix} \alpha_1 \\ \alpha_2 \end{pmatrix}, \quad \mathbf{\Gamma} = \begin{pmatrix} \beta_1 & 0 \\ 0 & \beta_2 \end{pmatrix}, \quad \boldsymbol{\epsilon}_t = \begin{pmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \end{pmatrix}$$

The residual vector  $\boldsymbol{\epsilon}_t$  has a bivariate normal distribution, i.e.  $\boldsymbol{\epsilon}_t \sim N(\mathbf{0}, \boldsymbol{\Sigma})$ , where

$$\boldsymbol{\Sigma} = \begin{pmatrix} \sigma_{xx} & \sigma_{xy} \\ \sigma_{xy} & \sigma_{yy} \end{pmatrix}$$

The factors  $x_t$  and  $z_t$  should resemble random walk processes, but we allow for a more general VAR process.

For the valuation of a CDS contract, the discounted payments of each of the two parties of the CDS contract, i.e. the protection buyer and the protection seller, should be considered. In our model, and in contrast to Doshi (2011), it is assumed that the discount factors, denoted by  $D(t+m)$  are exogenous and known at time  $t$ . We justify this choice primarily with the argument that adding more factors to our model would unnecessarily complicate the derivation, identification and estimation of the model, while not necessarily providing more flexibility to enhance the goodness of fit. Furthermore, it is sensible to assume *a priori* that the uncertainty related to the risk-free yield curve had a limited impact on the euro area CDS spreads during the crisis.

The protection buyer promises to pay the premium each quarter up to the termination of the CDS contract. However, when the credit event occurs (usually between the premium payment dates), the protection buyer pays the premium accrued and receives a payment,

$L$ , from the protection seller that is equal to the LGD. The expected value (under the risk neutral measure) of the discounted payments by the protection buyer at time  $t$  equals:

$$PB_t = E_t^Q \left\{ S_t \delta \sum_{i=1}^N D(t+i) \left[ 1_{(\tau > t+i)} + 1_{PA} 0.5 (1_{(\tau > t+i-1)} - 1_{(\tau > t+i)}) \right] \right\}. \quad (2)$$

where  $S_t$  is the annualized premium paid by the protection buyer, and  $\delta$  is the time between payment dates in annual terms (0.25 for quarterly payments). The function  $1_{PA}$  equals 1 if the contract specifies premium accrued, and 0 otherwise, as suggested by O’Kane and Turnbull (2003).  $N$  is the number of contractual payment dates until the contract matures.

The protection seller makes the payment  $LGD$  only in case of a credit event. The expected value (under the risk neutral measure) of the payment made is thus given by:

$$PS_t = E_t^Q \left\{ \sum_{j=1}^M D(t+j) LGD_{t+j-1}^Q 1_{(t+j-1 < \tau < t+j)} \right\} \quad (3)$$

where  $M$  is the number of periods to maturity for the CDS contract at time  $t$ . The spread of the CDS,  $S_t$  in formula (2), is set so that the expected value of the payments of the protection buyer equals the expected value of the payments of the protection seller, that is:

$$S_t = \frac{PS_t}{E_t^Q \left\{ \delta \sum_{i=1}^N D(t+i) \left[ 1_{(\tau > t+i)} + 1_{PA} 0.5 (1_{(\tau > t+i-1)} - 1_{(\tau > t+i)}) \right] \right\}} \quad (4)$$

In order to complete the CDS pricing model in equation (4), the expected values in formulas (2) and (3) need to be computed. Technical details on how we computed these expected values can be found in the appendix.

## 2.2 Parameter estimation

The combined formulas (4) and (1) for pricing sovereign CDS take the form of a non-linear state-space model:

$$\begin{aligned} \mathbf{y}_t &= \mathbf{f}(\mathbf{s}_t, \boldsymbol{\theta}) + \mathbf{u}_t & \mathbf{u}_t &\sim N(\mathbf{0}, \sigma_{uu} \mathbf{I}) \\ \mathbf{s}_t &= \boldsymbol{\mu} + \boldsymbol{\Gamma} \mathbf{s}_{t-1} + \boldsymbol{\epsilon}_t & \boldsymbol{\epsilon}_t &\sim N(\mathbf{0}, \boldsymbol{\Sigma}), \end{aligned} \quad (5)$$

where  $\mathbf{y}_t$  is the column vector of logged CDS spreads with maturities  $1Y, 2Y, \dots, 10Y$ , respectively. The parameter vector  $\boldsymbol{\theta}$  is defined as  $\boldsymbol{\theta} = (\alpha_1, \alpha_2, \beta_1, \beta_2, \sigma_{xx}, \sigma_{yy}, \sigma_{xy}, \sigma_{uu})$ . The function  $\mathbf{f}(\mathbf{s}_t, \boldsymbol{\theta})$  denotes the logged CDS pricing formula of equation (4). As the fit of the pricing formula is not perfect, a vector of residuals  $\mathbf{u}_t$  is added in order to account for pricing

errors.<sup>3</sup> The factors  $\mathbf{s}_t$  are unobserved and need to be estimated. A robust method to accomplish this, for a known parameter vector  $\boldsymbol{\theta}$ , is the Unscented Kalman filter (UKF). Given that there are parameters to be estimated, a least square estimation method in combination with the UKF needs to be used. We leave the technical details on the joint estimation of the state vector  $\mathbf{s}_t$  and the parameter vector  $\boldsymbol{\theta}$  for the appendix.

The most liquid sovereign CDS contracts for the countries in the euro area are denominated in US dollars. Therefore, we use dollar denominated CDS premia from the CMA Database in our analysis. We use annual maturities from one to ten years, and end-of-month observations, thus reflecting information disclosed throughout the month. The sample length depends on data availability and quality.<sup>4</sup> The zero coupon US yield curve from Thompson Reuters is used to compute the discount factors.

The estimated parameters are shown in Table 1. Notwithstanding the very parsimonious nature of our CDS pricing model, the high  $R^2$  coefficient values reported in the table suggest that the fit of the models is good. The largest errors are recorded for Greek and Portuguese data. The mean absolute value of the error term for Greece and Portugal amounted to 45 and 40 basis points respectively. These are large values in absolute terms. However, they can be regarded as small in relative terms if it is remembered that the Greek CDS premia stood well above 1000 basis points in 2011. The large errors primarily reflect the difficulties faced to successfully capture the large volatility of the CDS spreads in Greece and Portugal during the last months of the crisis. We also checked that keeping the LGD constant in time reduces the fit of the estimated models considerably, which suggests that a one factor model would be insufficient to replicate the term structure of CDS spreads effectively.

Similar to Pan and Singleton (2008), our results also point to the presence of some explosive (or random-walk) processes governing the changes in the PD and in the LGD for many sovereign CDS contracts. This can be seen in parameters  $\alpha_1$  and  $\alpha_2$  in the table. Pan and Singleton (2008) noted that whether the intensity was explosive or not appeared to be inconsequential for econometric identification. Our simulated results shown below also suggest that these estimated parameters, even when explosive, may be plausible.

The estimation results further suggest that the correlation between the process driving PD

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<sup>3</sup>We use logarithms in the observation equation to smooth the extreme changes in CDS premia in the time of crisis.

<sup>4</sup>Samples start in March 2004 (Greece), August 2005 (Portugal), June 2007 (Ireland), July 2007 (Spain), October 2007 (Italy) and November 2007 (France). All samples end in August 2012.

and the process driving LGD is close to zero, or very small, for most countries except Ireland and Spain, see parameter  $\rho_{xz}$  in Table 1. This is in contrast with the empirical literature on corporate CDS spreads which documents a positive correlation between PDs and LGDs, see Acharya, Bharath, and Srinivasan (2007) and Altman, Brady, Resti, and Sironi (2005).

## 2.3 The identification problem

We have analysed whether PD and LGD can be satisfactorily identified. This assessment was conducted by means of the graphical analysis employed in Christensen (2007), and also by means of the small-scale Monte Carlo exercise used in Pan and Singleton (2008).

Figure 1 shows different combinations of PD and LGD that would allow our model-implied CDS spreads to perfectly match the observed CDS spreads on a given date. The date chosen is December 2010, when tensions in sovereign debt markets were mounting but were still far from the critical peaks recorded in later months. Figure 1 shows that the CDS spreads of France, Ireland, Italy and Portugal react very differently, over different maturities, to changes in PD and LGD. Only one PD and LGD combination pair provides a perfect fit of the observed CDS spreads, suggesting that separate identification of PD and LGD is feasible. However, the CDS spreads at different maturities in Germany and Spain do not appear to react differently to alternative combinations of PD and LGD. Identifiability for Greece is also slightly more challenging for PD values larger than 60%. It appears that for large PDs, the range of LGD values that provide a perfect fit for all CDS contracts lies between 10% and 20%.

Despite this, it should be noted that our model-implied CDS spreads do not fit the observed spreads perfectly. The mean absolute value of the errors are relatively large, ranging from 4 to 50 basis points. Thus, when assessing identifiability of the parameters, it is sensible to report the combinations of PD and LGD which provide model-implied CDS spreads not departing from the true values by more than the average absolute estimation error (Christensen, 2007). Such combinations are shown in Figure 2. Once again, the possible set of combinations of the PD and LGD pairs that fit the CDS spreads within that margin of error is relatively narrow for France, Ireland, Italy and (slightly less so) Portugal. The results for Portugal are now slightly less satisfactory. The LGD values which provide CDS spreads within the margin of error of our model range from 40% to 60%. Nevertheless, the results for Germany, Greece and Spain are broadly aligned with those reported in Figure 1.

Naturally it should be remembered that these results are valid for a given date, whereas in our analysis the model is estimated over a sample period stretching more than 60 months.<sup>5</sup> This represents an additional constraint which should facilitate the identification of PD and LGD. As a final robustness check, and in line with Pan and Singleton (2008), we simulate data from the estimated models and check if the estimates based on this simulated data differ substantially from our estimation results. We find that the parameters estimated using simulated data are centered around those estimated using the original data for most countries, and that the dispersion of the simulated parameters is relatively narrow (cf., Table 2). The only notable exception is Portugal, for which the standard deviation of the estimated parameters is relatively large.

## 2.4 The market's perception of credit risk

It follows from the discussion above that our CDS model is not greatly robust with respect to separately identifying PD and LGD for Portugal and Greece from July 2011 onwards. To take account of this, we thus shorten the sample used for these two countries in the empirical analysis which follows.

The estimates of the PDs and LGDs within a two-year horizon are shown in Figure 3. One thing to note is that these estimates are computed under the risk neutral measure. Due to the presence of a positive risk premia, both the PD and the LGD are possibly best seen as upper bounds to actual PD and LGD (or those estimated under the 'physical' measure as usually referred to in the literature). Our reported estimates for LGD range from 2% for Spain at the beginning of 2008 to 50% for Portugal in mid 2010, (see Figure 3). However, for the majority of the countries studied here, the LGD fluctuated at levels comfortably below 20% for most of the sample. To put these numbers into perspective, it is worth noting that the actual average LGD of all sovereign debt restructurings in the world between 1970 and 2010 was 50%, and 50% was also the loss suffered by investors following the restructuring of Russian debt in 1998 (Cruces and Trebesch, 2013). The largest LGD recorded in our results (50%) would also be roughly aligned with a Standard and Poor's '*recovery rating*' of 3.<sup>6</sup> However, only two of the countries in our sample, Greece and Portugal, towards the end of 2012, were assigned recovery ratings slightly worse than 3. Indeed, by the end of 2012, no recovery rating had been assigned to any of the remaining euro area countries.

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<sup>5</sup>The chosen date is broadly representative of the problems encountered at different dates.

<sup>6</sup>Recovery rating of 2, 3 and 4 according to Standard and Poor's would be equivalent to expected recovery rates ranging from 70% to 90%, 50% to 70% and 30% to 50% respectively.

With the exception of Greece, the higher the ratio of the sovereign debt to GDP, the higher was the LGD. This appears to have been the case for Italy, Portugal and Ireland, the countries which from 2008 onwards have had the largest debt to GDP ratios and are usually associated with the highest LGD estimates. This accords well with economic theory. For example, prior to the sovereign debt crisis the costs of financing for the euro area governments were all broadly aligned, and commitments towards fulfilling the conditions of the Stability and Growth Pact dictated that the projected long-term paths of primary surpluses should be similarly broadly aligned across countries. Under these presumptions, disparities in the estimated haircuts to restore debt sustainability, computed from the simple static model used by IMF economists, would be exclusively driven by the magnitude of the debt to GDP ratio and GDP growth projections (Das, Papaioannou, and Trebesch, 2012).

In terms of scale of debt to GDP ratio, Greece has one of the largest ratios among the analysed countries from end-2008 onwards, and yet the level of the LGD is relatively low. It is only by mid-2011 that the LGD of Greece reaches values close to 30%. Nevertheless, this estimate is still somehow lower than that implied by the ‘*recovery rating*’ that was assigned to Greece by Standard and Poor’s in 2012. Our own last LGD estimate for Greece in 2011 is also lower than the losses that were associated with the 2012 Greek debt restructuring, which according to some recent studies, amounted to around 60% (Zettelmeyer, Trebesch, and Gulati, 2013).

The LGD estimates for Germany and Spain remain relatively unchanged for most of the sample under study. Furthermore, and in spite of the fact that the Spanish CDS spread continued to increase at a faster pace than that of countries like Germany and France, the estimated LGD for Spain is the lowest in our sample. This could be a reflection of the identification problems highlighted for Spain in the previous section.

The intensification of tensions in sovereign debt markets after May 2010 is more clearly illustrated in the estimates of PD shown in Figure 3. These estimates display a broad upward trend. The probability of default as perceived by financial markets surpassed the 50% level in three of the countries: Greece, Spain and Ireland.

### 3 Main drivers of the market’s perception of credit risk

We turn now to the issue of studying the main drivers underlying market perception of sovereign credit risk. The role of economic and financial developments, institutional developments and financial contagion will be studied by means of a regression analysis.

#### 3.1 Regression analysis

A multivariate regression model is estimated for every country:

$$\mathbf{y}_t = \boldsymbol{\alpha} + \mathbf{B}_1 \mathbf{e}_t + \mathbf{B}_2 \mathbf{i}_t + \mathbf{B}_3 \mathbf{n}_t + \boldsymbol{\varepsilon}_t \quad (6)$$

where  $\mathbf{y}_t = (y_t^2, y_t^5, y_t^7)'$  and  $y_t^k$  represents either the change in PD or the change in the LGD for a given country, where  $k$  denotes the (two-year, five-year, and seven-year) horizon over which the dependent variables are defined. Using changes in PD and LGD enables us to focus on the short-term effects of economic factors on sovereign credit risk, while at the same time we avoid problems associated with nonstationarity. The vector of regressors  $\mathbf{e}_t$  represents a set of economic and financial indicators;  $\mathbf{i}_t$  represents a set of dummy variables which serve as a proxy for institutional developments; and  $\mathbf{n}_t$  represents a set of dummy variables which serve as a proxy for ‘country news’.

Due to the large number of regressors employed, the model is over-parameterized. Consequently, we adopt a Bayesian model averaging method. In particular, every different combination of regressors represents different possible models, and the posterior probability of each such model being the true model can be computed for given assumptions. In our regression results we will report Bayesian Model Averaging (BMA) coefficients as this is standard in the literature. The BMA coefficient associated with a certain regressor is computed as the weighted sum of the coefficient associated with that regressor in every single model, with the use of the posterior probability associated with the model as weights. We leave the technical details on how to compute posterior model probabilities for the technical appendix. The list of economic variables to be included as part of  $\mathbf{e}_t$  is as follows:

- a. VIX. Implied volatility of Standard and Poor’s 500 index.
- b. VSP. Stock price volatility.
- c. RAT. Sovereign credit rating.
- d. iTraxx. Five-year iTraxx for European senior financials.
- e. BCDS. Median five-year CDS spread of main banks in the country.
- f. FUT-SER. Survey expectations of future demand in the services sector.
- g. FUT-GDP. Consensus Economics GDP growth one-year ahead.



- h. BC. Business cycle indicator.
- i. DEF. Monthly cash budget deficit.

The VIX serves as a proxy for global financial market conditions, while VSP accounts for conditions in domestic financial markets. We compute VSP as the standard deviation of the daily returns of the country equity price indexes using Datastream’s Global Equity Index database. To isolate ‘country specific effects’ from the world financial market volatility effects (proxied by the VIX), we regress the country volatility index on the VIX and use the residual from that regression as our measure for VSP. The sovereign rating indicator, RAT, is constructed using information from sovereign rating downgrades or upgrades, as well as revisions to the outlook for the sovereign rating. Information about sovereign ratings comes from Fitch, Standard and Poor’s and Moody’s. A value of 1 is assigned to a one-notch downgrade, 0.5 for a change to negative outlook, -0.5 for a change to positive outlook and -1 to a one notch upgrade. RAT is then defined as the average of the reported changes from the three rating agencies. The variable *iTraxx* serves as a proxy for tensions in the banking sector across euro area countries, while BCDS should identify country-specific tensions in the banking sector. FUT-SER is the indicator of the “*evolution of demand expected in the months ahead in the services sector*” published by the European Commission.<sup>7</sup> BC is constructed as a weighted average of five survey confidence indicators published by the European Commission: industry, services, consumers, construction and retail trade. The questions asked in these surveys relate to views on both current and future developments. Therefore, this weighted average index is regressed on the FUT-SER indicator above, and the computed residuals are used as a coincident business cycle indicator.<sup>8</sup> Finally, the series DEF is constructed as the ratio between monthly cash government budget deficit and the interpolated nominal GDP estimate. We use monthly budget balances of either the general government (when available), or the central government as published by the National Statistical Institutes. These monthly cash data provide valid information on the state of public finances as shown in Onorante, Pedregal, Perez, and Signorini (2010). Monthly seasonal factors have been used to correct for seasonality. These monthly seasonal factors have been computed as average percentage deviations from the estimated trend.

There are six ‘institutional’ dummy variables in  $\mathbf{z}_t$  which serve to proxy for a number of relevant changes to the institutional and policy framework introduced in the euro area dur-

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<sup>7</sup>For Ireland, due to the discontinuity of the data after April 2008, use is made instead of the “future business activity expectations” indicator published by Markit.

<sup>8</sup>For Ireland, data from the European Commission is discontinued, and use is made instead of the ‘economic climate indicator of the Irish Economic and Social Research Institute.

ing the financial crisis. These are:

- a. *Market Functioning*. Policies aimed at addressing malfunctioning in financial markets.
- b. *Ec & Fiscal Stimulus*. Economic policies aimed at boosting economic growth.
- c. *Bank Support*. Measures aimed at strengthening the financial position of banks.
- d. *Government Support*. Measures providing support for the financing of governments.
- e. *Standard MP*. Key changes in the standard monetary policy of the ECB.
- f. *Lehman Brothers*. Its collapse potentially signalled to investors that the ‘too big to fail’ argumentation may be misguided.

Finally, there are three ‘country news’ dummy variables in  $\mathbf{n}_t$ : *News Greece*, *News Ireland* and *News Portugal*.

The dummy variables have been constructed as follows. First, we created a list of important country news and institutional events, see Table 3. The events in the list are organised in chronological order and assigned either to the six previously defined types of ‘institutional’ changes, or to the three different types of ‘country news’. Then, a dummy variable for each group is defined by assigning a value of one to the dates of the events listed in Table 3 assigned to that specific group, and a value of zero otherwise. It should be noted that the dependent variable is modelled in first differences, and thus by this construction the impact of the institutional dummy is permanent in the level of the dependent variable. As a result, we will thus refer to this dummy variable as a **permanent-effect dummy**.<sup>9</sup> It seems sensible to assume that the impact of these institutional changes on the market’s perception of credit risk should be of a permanent nature. However, it cannot be *a priori* excluded that these announcements lacked credibility and thus had only a temporary effect. We thus assess the robustness of our results by running our regression analysis also using dummy variables of a second type. In particular, a **temporary-effect dummy** is constructed by assigning a value of one to the dates of the event, a value of minus one to the month following the event, and a value of zero otherwise.

Hereafter, similarly to most early warning systems, we choose to focus on the regression results for the risk of a credit event in the two-year horizon. Macroeconomic and financial factors are less likely to effectively explain the risk of default beyond this horizon. Regres-

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<sup>9</sup>There are two institutional changes which should be assigned a value of minus one to signal, for example, a non-supportive measure as opposed to a supportive measure. First, there are two announcements we associate with the *Ec. & Fiscal Stimulus* group which have been defined with a minus, namely realisation that a major recession lay ahead (Jan-2008), and the fiscal compact limiting structural deficits (Dec-2011). Second, increases in the ECB policy rates have been taken with a minus as opposed to cuts in rates which have been assigned a one.

sions results for the five and seven-year maturities have also been computed, although not reported, and are in line with our main conclusions. The Bayesian analysis shows that only a reduced set of variables is statistically relevant (Table 4). Furthermore, the  $R^2$  coefficients of the regressions are generally very low, suggesting that our set of economic and financial indicators and institutional dummies in fact fails to explain a large proportion of the movements in PDs and LGDs.

### 3.2 The role of economic and institutional factors

The empirical results suggest that macroeconomic indicators, and in particular the timely cash balances of the public sector (DEF) and the business cycle indicator (BC), are not of relevance to explain movements in the PD and LGD. Only the indicator on expectations on future economic activity (FUT) is marginally significant for Germany, France and Italy, and although significant, its impact is small in size, see Table 4. This result is broadly aligned with Blommestein, Eijffinger, and Qian (2012) who also identified some turbulent sub-periods during the euro area crisis where economic fundamentals did not affect sovereign risk measures.

Our empirical results also show that financial indicators do not covary with PD or LGD for most of the countries in our sample. This differs slightly from Longstaff, Pan, Pedersen, and Singleton (2011), who found there was a positive correlation between the risk premium component of sovereign CDS contracts and global risk factors such as the VIX for a number of developed and emerging market economies. Their sample, however, did not include euro area countries. Our results are also partly in contrast with the results reported in Ang and Longstaff (2013), who found a correlation between a systemic risk factor driving the PD of euro area countries and a number of financial indicators. The financial indicators which Ang and Longstaff (2013) found to be correlated with their systemic risk factor were the iTtraxx index for Europe and the returns on the German DAX index. In our analysis, the volatility of the country stock indexes (VSP) does not appear to be significant. It is only the VIX that helps to explain developments in the PD and LGD of Greece, the PD of Germany and the LGD of Portugal.<sup>10</sup> Interestingly, in the study by Ang and Longstaff (2013) the VIX was not found significant to explain the systemic risk factor.

Finally, our results show that the rating indicator (RAT) is only relevant to explain move-

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<sup>10</sup>As a robustness check, we also estimated principal components of changes in PDs and LGDs. Again, the VIX and national volatility indices appeared to be only weakly correlated with the common risk factors.

ments in the PD of Greece and Portugal. It appears significant to explain movements in Italy, but the coefficient reported in Table 4 comes with the wrong sign, as it suggests that a deterioration in the rating lowers the PD of Italy. This is counter-intuitive and may be driven by the fact that the RAT series for Italy remained unchanged for most of the sample, as downgrades of the Italian rating only materialised from mid 2011 onwards. It is thus sensible to ignore this significance and suggest that what the RAT variable in actual fact captures for Italy cannot be thus associated with what is to be expected of a rating announcement. It should also be noted that the credit rating of Germany and France remained unchanged for most of the sample, and it is thus not surprising that RAT is not significant to explain developments for the two countries.

As for institutional developments, only some of the group dummy variables included in the regressions to proxy for changes to the institutional framework appear relevant in explaining some of the dynamics of PD and LGD, see Table 4. The *Lehman Brothers* event dummy appears somewhat relevant in explaining developments in the French PD and the Irish LGD, and measures taken to provide *economic and fiscal stimulus* appear to covary negatively with the LGD in Portugal.

### 3.3 The role of contagion

In this paper we use the term contagion to refer to instabilities in a specific segment of the financial market that are transmitted to other financial markets, and which are beyond what might be expected by economic or institutional fundamentals. We take the following strategy to identify contagion. First, we test the possible impact that the release of country news may have had on the PD and LGD of other countries. Second, we make use of the correlations of the country residuals of our regression model (6) to identify shocks to PDs and LGDs that are not driven by economic fundamentals, nor by institutional factors, and nor by the previously identified country news. We refer to this phenomenon as ‘*dependence beyond fundamentals*’. Third, contagion may also be revealed by an analysis of the largest residuals from our PDs and LGDs country regressions. We describe this effect as ‘*dependence of extreme shocks*’. Large values are identified from the empirical distributions of the residuals (for each maturity and country separately), i.e. values beyond the sample 75th percentiles. Following this, the number of coincidences of these large residuals from PD (or LGD) regressions in different countries in each period is computed. The final coincidence index used is the share of countries with large PDs (or LGDs) in the total number of analysed countries at time  $t$ . Multiple coincidences of large country residuals are taken as evidence of contagion.

Results in Table 4 show that economic news from Greece, Ireland and Portugal are not associated with movements in PD and LGD with only one exception. The relevant exception is the small positive association reported between *News Portugal* and the PD of Greece and Ireland. Surprisingly, news about Portugal does not appear associated with movements in the PD and LGD of Portugal. Thus, the impact of country news, as captured in the regression analysis presented above, does not appear to reflect contagion.

However, the cross-country correlations of the residuals suggest otherwise. Table 5 shows evidence of the significant positive correlation of the PD residuals across countries. Results in the table point to significant inter-linkages in PD developments across countries. Interestingly, the PD residuals of Greece and Ireland failed to correlate with those from other countries, suggesting that the nature of contagion was possibly not as unidirectional and simple as the financial media tends to represent it, namely from Greece to the rest. A stronger correlation is reported for France vis-a-vis Germany, and a similar pattern is revealed in the LGD residuals, although the residuals of the LGD for Italy are not correlated with those in other countries.

Figure 4 shows the coincidences of large residual changes for both the PD residuals and the LGD residuals when using permanent dummy-effects.<sup>11</sup> If the residuals had been independent, instances of four large country residuals occurring at the same point in time would be unusual ( $prob < 4\%$ ). Some unobservable common factors among the markets can be identified here again. In particular, the peaks for these coincidences seem to cluster around three major periods. Firstly, in the second half of 2010 when the financial problems of Ireland were becoming more apparent and markets started to speculate on the possibility of a debt restructuring for one euro area country. Secondly, the Autumn of 2011, when some market participants speculated about the possibility that Greece might leave the euro area, and when tensions in the Italian sovereign debt markets escalated to peak levels. Thirdly, in the spring of 2012, when the PSI for Greece was finally agreed and tensions in Spain were at their peak.

Our results presented in Figure 4 thus suggest that common unobservable factors driving the sovereign risk in individual countries have a time-varying impact. This is also aligned with Alter and Schuler (2012) who found international effects of changes in sovereign CDS spreads in the euro area countries to vary in time (e.g. before and after government interventions in the financial sector) and also between countries. This changing impact may be attributed to

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<sup>11</sup>Similar results were obtained when using temporary dummy-effects, and are thus not shown.

the herding behaviour of investors as suggested by Beirne and Fratzscher (2013), or to changing market uncertainty affecting the cognitive biases of market participants (Blommestein, Eijffinger, and Qian, 2012).

### 3.4 Robustness to alternative specifications

It might be argued that our grouping of the events may be misguided, because the impact of certain individual events may be much larger than others. Consequently, as a check on the robustness of our empirical results we have also conducted a regression analysis without grouping the events, and from this we found similar results.<sup>12</sup> Of all the individual events studied, only *November 2008* is found to have a significant and positive impact across several countries. The positive sign would suggest that markets were regarding the additional fiscal boost of the European Economic Recovery Plan, and possibly the earlier launch of the bail-out plans, as potentially putting under pressure the state of the public finances. At that stage, markets possibly did not find it credible that the European Recovery Plan would promote economic activity.

## 4 Conclusions

The aim of this paper was to study market perception of the sovereign credit risk in euro area countries during the recent financial crisis. Measures of PD and LGD, as perceived by financial markets, were computed for Germany, France, Greece, Ireland, Italy, Portugal and Spain. For a number of euro area countries, notably France, Ireland and Italy, the separate identification of PD and LGD appears empirically tractable. However, and at times of excessively high CDS spread levels, this appears slightly less so for Greece and Portugal. Identifiability appears more challenging altogether however, for the cases of Germany and Spain.

With these caveats in mind, our reported estimates of PD and LGD reveal several interesting results. In contrast with the literature on corporate credit risk, the dynamics of the estimated PD and LGD are not always strongly positively correlated. Our estimated LGDs also show that high levels of LGD are often associated with large stocks of public debt. This is something broadly to be expected and aligned with the rationale of the haircuts derived from the simple static debt sustainability analysis proposed by the IMF, see e.g. IMF (2013). For the euro area countries under investigation here, the estimated LGD was for

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<sup>12</sup>For the sake of brevity, we do not report these results.

most of the sample comfortably below 40%. Only at a few points in time, and in the sample under study, was the estimate of LGD for Portugal recorded as surpassing 50%, which is the reported average LGD of all sovereign debt restructurings that took place in the world economy between 1970 and 2010. This finding potentially suggests that the credit events occurring in the eurozone were associated with investors' concerns about the liquidity of the sovereign state rather than concerns about the solvency of the sovereign state. Indeed, liquidity problems of sovereign states would in fact generate much lower losses for investors than full-blown debt crises.

Our estimates of PD and LGD were only weakly related to economic and institutional factors. We also found evidence of possible contagion from developments in the perceptions of credit risk in one euro area country to other euro area countries. The nature of contagion was, however, possibly not as unidirectional and simple as it is usually represented by the financial media, namely from Greece to the rest, because the coincidental shocks occurred on dates related to different crisis events.

The financial spillovers between countries reported in our results highlight the problem of how international investors may fail to price economic and institutional news during turbulent periods. Uncertainty, low quality of available information, and aggravated risk aversion of market participants played a crucial role in the transmission of risk between markets. Yet, our study suggests that extreme shocks were transferred to international sovereign markets through the channel of market perception rather than through economic fundamentals. We suggest that a key policy to counter the occurrence (and resultant destabilizing effect) of this would be to enhance the transparency of public accounts. Credible measures by policy makers and regulators to curb excessive sovereign indebtedness, such as the newly introduced fiscal compact, should also help to protect financial stability.

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# Technical Appendix

## A Pricing CDS contracts

As suggested by Doshi (2011), the Laplace transform of a quadratic form  $\Delta = \mathbf{A} + \mathbf{B}\epsilon + \epsilon'\mathbf{C}\epsilon$  can be computed to obtain the expectation of the exponential of this form:

$$E \{ \exp(\Delta) \} = \exp \left\{ -0.5 \ln(\det(\mathbf{I} - 2\mathbf{\Omega C})) + \mathbf{A} + 0.5\mathbf{B}(\mathbf{\Omega}^{-1} - 2\mathbf{C})^{-1}\mathbf{B}' \right\}, \quad (\text{A-1})$$

where  $\epsilon \sim N(\mathbf{0}, \mathbf{\Omega})$ ,  $\mathbf{I}$  is the identity matrix, and  $\det$  indicates determinant. Using (A-1) and the autoregressive formula (1) we define:

$$\begin{aligned} a(\mathbf{A}, \mathbf{B}, \mathbf{C}) &= \boldsymbol{\mu}'\mathbf{C}\boldsymbol{\mu} + \mathbf{B}\boldsymbol{\mu} + 2 * \boldsymbol{\mu}'(\mathbf{C}\mathbf{W}\mathbf{C}\boldsymbol{\mu} + \mathbf{C}\mathbf{W}\mathbf{B}') \\ &\quad + 0.5\mathbf{B}\mathbf{W}\mathbf{B}' - 0.5 \ln(\det(\mathbf{I} - 2\mathbf{\Sigma C})), \\ b(\mathbf{A}, \mathbf{B}, \mathbf{C}) &= 2\boldsymbol{\mu}'\mathbf{C}\boldsymbol{\Gamma} + \mathbf{B}\boldsymbol{\Gamma} + 4\boldsymbol{\mu}'\mathbf{C}\mathbf{W}\mathbf{C}\boldsymbol{\Gamma} + 2\mathbf{B}\mathbf{W}\mathbf{C}\boldsymbol{\Gamma}, \\ c(\mathbf{A}, \mathbf{B}, \mathbf{C}) &= \boldsymbol{\Gamma}'(\mathbf{C} + 2\mathbf{C}\mathbf{W}\mathbf{C})\boldsymbol{\Gamma}, \end{aligned} \quad (\text{A-2})$$

where  $\mathbf{W} = (\mathbf{\Sigma}^{-1} - 2\mathbf{C})^{-1}$  and  $\boldsymbol{\mu}$ ,  $\boldsymbol{\Gamma}$ , and  $\mathbf{\Sigma}$  contain parameters defined in (1). The above formulas are helpful in deriving the expectations in formulas (2) and (3). The protection leg of the contract presented in equation (2) may be well approximated by:

$$PS_t = E_t^Q \left\{ \sum_{j=1}^M D(t+j) \exp(-z_{t+j-1}^2) \left( \exp \left( \sum_{k=1}^j -x_{t+k-1}^2 \right) - \exp \left( \sum_{k=1}^{j-1} -x_{t+k-1}^2 \right) \right) \right\} \quad (\text{A-3})$$

Standard calculations reveal that the following expectations can be solved iteratively:

$$E_t^Q \{ 1_{(\tau > t+i)} \} = E_t^Q \left\{ \exp \left( \sum_{k=1}^i -x_{t+k-1}^2 \right) \right\} = \exp(\mathbf{D}_i + \mathbf{E}_i \mathbf{s}_t + \mathbf{s}_t' \mathbf{F}_i \mathbf{s}_t), \quad (\text{A-4})$$

$$E_t^Q \left\{ \exp(-z_{t+h-1}^2) \exp \left( \sum_{j=1}^h -x_{t+j-1}^2 \right) \right\} = \exp \{ \mathbf{G}_h + \mathbf{H}_h \mathbf{s}_t + \mathbf{s}_t' \mathbf{J}_h \mathbf{s}_t \}, \quad (\text{A-5})$$

$$E_t^Q \left\{ \exp(-z_{t+h-1}^2) \exp \left( \sum_{j=1}^{h-1} -x_{t+j-1}^2 \right) \right\} = \exp \{ \mathbf{K}_h + \mathbf{L}_h \mathbf{s}_t + \mathbf{s}_t' \mathbf{M}_h \mathbf{s}_t \}, \quad (\text{A-6})$$

where  $\mathbf{s}_t = (x_t, z_t)'$  is the vector of state variables in the state-space model (5). The formula (A-4) denoting the survival probability up to time  $t+i$  can be calculated by deriving the parameters  $\mathbf{D}_i$ ,  $\mathbf{E}_i$  and  $\mathbf{F}_i$  recursively:

$$\begin{aligned} \mathbf{D}_{i+1} &= a(\mathbf{D}_i, \mathbf{E}_i, \mathbf{F}_i) + \mathbf{D}_i, \\ \mathbf{E}_{i+1} &= b(\mathbf{D}_i, \mathbf{E}_i, \mathbf{F}_i), \\ \mathbf{F}_{i+1} &= c(\mathbf{D}_i, \mathbf{E}_i, \mathbf{F}_i) + \mathbf{F}_0, \end{aligned} \quad (\text{A-7})$$

with the starting values  $\mathbf{D}_0 = 0$ ,  $\mathbf{E}_0 = \begin{pmatrix} 0 & 0 \end{pmatrix}$ ,  $\mathbf{F}_0 = \begin{pmatrix} -1 & 0 \\ 0 & 0 \end{pmatrix}$ .

The formulas (A-5) and (A-6) can be calculated accordingly by deriving the parameters  $\mathbf{G}_i$ ,  $\mathbf{H}_i$ ,  $\mathbf{J}_i$ , and  $\mathbf{K}_i$ ,  $\mathbf{L}_i$ ,  $\mathbf{M}_i$ , recursively:

$$\begin{aligned} \mathbf{G}_{i+1} &= a(\mathbf{G}_i, \mathbf{H}_i, \mathbf{J}_i) + \mathbf{G}_i, \\ \mathbf{H}_{i+1} &= b(\mathbf{G}_i, \mathbf{H}_i, \mathbf{J}_i), \\ \mathbf{J}_{i+1} &= c(\mathbf{G}_i, \mathbf{H}_i, \mathbf{J}_i) + \mathbf{J}_0 \text{ for } i > 0, \text{ and } \mathbf{J}_1 = c(\mathbf{G}_i, \mathbf{H}_i, \mathbf{J}_0^*) + \mathbf{J}_0 \text{ for } i = 0, \\ \mathbf{K}_{i+1} &= a(\mathbf{K}_i, \mathbf{L}_i, \mathbf{M}_i) + \mathbf{K}_i, \\ \mathbf{L}_{i+1} &= b(\mathbf{K}_i, \mathbf{L}_i, \mathbf{M}_i), \\ \mathbf{M}_{i+1} &= c(\mathbf{K}_i, \mathbf{L}_i, \mathbf{M}_i) + \mathbf{M}_0 \text{ for } i > 0, \text{ and } \mathbf{M}_1 = c(\mathbf{K}_i, \mathbf{L}_i, \mathbf{M}_0^*) + \mathbf{M}_0 \text{ for } i = 0. \end{aligned} \tag{A-8}$$

The starting values of parameters are

$$\begin{aligned} \mathbf{G}_0 &= 0, \quad \mathbf{H}_0 = \begin{pmatrix} 0 & 0 \end{pmatrix}, \quad \mathbf{J}_0 = \begin{pmatrix} -1 & 0 \\ 0 & 0 \end{pmatrix}, \quad \mathbf{J}_0^* = \begin{pmatrix} 0 & 0 \\ 0 & -1 \end{pmatrix}, \text{ and} \\ \mathbf{K}_0 &= 0, \quad \mathbf{L}_0 = \begin{pmatrix} 0 & 0 \end{pmatrix}, \quad \mathbf{M}_0 = \begin{pmatrix} -1 & 0 \\ 0 & 0 \end{pmatrix}, \quad \mathbf{M}_0^* = \begin{pmatrix} -1 & 0 \\ 0 & -1 \end{pmatrix}. \end{aligned}$$

The equation (A-4) is used to solve the formula (3) and equations (A-5) and (A-6) are used in (2). The CDS spread in formula (4) is then derived from (2) and (3) for ten different maturities (1Y, 2Y, ... 10Y) of CDS contracts. In turn, the column vector of explained variables  $\mathbf{y}_t$  in the state-space model (5) consists of original (logged) CDS spreads for these ten maturities observed on the market at time  $t$ .

Finally, the formula for the (risk-neutral) expected LGD at time  $t + i$ , i.e.

$$E_t^Q \left\{ L_{t+i}^Q \right\} = E_t^Q \left\{ \exp(-z_{t+i}^2) \right\} = \exp(\mathbf{D}_i^* + \mathbf{E}_i^* \mathbf{s}_t + \mathbf{s}_t' \mathbf{F}_i^* \mathbf{s}_t), \tag{A-9}$$

may be derived analogously. The recursive equations for the parameters  $\mathbf{D}_i^*$ ,  $\mathbf{E}_i^*$  and  $\mathbf{F}_i^*$  take the following form:

$$\begin{aligned} \mathbf{D}_{i+1}^* &= a(\mathbf{D}_i^*, \mathbf{E}_i^*, \mathbf{F}_i^*) + \mathbf{D}_i^*, \\ \mathbf{E}_{i+1}^* &= b(\mathbf{D}_i^*, \mathbf{E}_i^*, \mathbf{F}_i^*), \\ \mathbf{F}_{i+1}^* &= c(\mathbf{D}_i^*, \mathbf{E}_i^*, \mathbf{F}_i^*) \end{aligned} \tag{A-10}$$

and the starting values are

$$\mathbf{D}_0^* = 0, \quad \mathbf{E}_0^* = \begin{pmatrix} 0 & 0 \end{pmatrix}, \quad \mathbf{F}_0^* = \begin{pmatrix} 0 & 0 \\ 0 & -1 \end{pmatrix}.$$

## B The Unscented Kalman Filter

The Unscented Kalman filter equations provide an estimator of the state  $\mathbf{s}_t$  conditional on information up to time  $t$ , this estimator is denoted as  $\mathbf{s}_{t|t} = E\{\mathbf{s}_t|y_1, \dots, y_t\}$  and its corresponding covariance matrix denoted as  $\mathbf{P}_{t|t} = E\{(\mathbf{s}_{t|t} - \mathbf{s}_t)(\mathbf{s}_{t|t} - \mathbf{s}_t)'|y_1, \dots, y_t\}$ . In our model the non-linearities are only present in the measurement equation, and the state vector is assumed to be Gaussian. Under this setting, the UKF formulation can be heavily simplified if the ‘prediction’ steps are ‘*marginalised*’. We follow here one simple ‘marginalised’ UKF procedure as reviewed in Briers, Maskell, and Wright (2003). The Unscented Kalman filter equations are thus given by:

$$\begin{aligned}
\mathbf{s}_{t+1|t} &= \boldsymbol{\mu} + \boldsymbol{\Gamma} \mathbf{s}_{t|t} \\
\mathbf{P}_{t+1|t} &= \boldsymbol{\Gamma} \mathbf{P}_{t|t} \boldsymbol{\Gamma}' + \boldsymbol{\Sigma} \\
\mathbf{y}_{t+1|t} &= E\{\mathbf{f}(\mathbf{s}_{t+1}, \boldsymbol{\theta})|y_1, \dots, y_t\} \\
\boldsymbol{\nu}_{t+1} &= \mathbf{y}_{t+1} - \mathbf{y}_{t+1|t} \\
\mathbf{P}_{t+1|t}^{yy} &= E\left\{(\mathbf{y}_{t+1|t} - \mathbf{y}_{t+1})(\mathbf{y}_{t+1|t} - \mathbf{y}_{t+1})'\right\} \\
\mathbf{P}_{t+1|t}^{xy} &= E\left\{(\mathbf{s}_{t+1|t} - \mathbf{s}_{t+1})(\mathbf{y}_{t+1|t} - \mathbf{y}_{t+1})'\right\} \\
\mathbf{K}_{t+1} &= \mathbf{P}_{t+1|t}^{xy} \left(\mathbf{P}_{t+1|t}^{yy}\right)^{-1} \\
\mathbf{s}_{t+1|t+1} &= \mathbf{s}_{t+1|t} + \mathbf{K}_{t+1} \boldsymbol{\nu}_{t+1} \\
\mathbf{P}_{t+1|t+1} &= \mathbf{P}_{t+1|t} - \mathbf{K}_{t+1} \mathbf{P}_{t+1|t}^{xy'}
\end{aligned}$$

These equations resemble those of the standard Kalman filter with the exception of the equations related to  $\mathbf{y}_{t+1|t}$ ,  $\mathbf{P}_{t+1|t}^{yy}$  and  $\mathbf{P}_{t+1|t}^{xy}$ . The expectations defined by those terms have no analytical solution and are thus approximated by means of an *unscented transformation*.

**Estimation.** The UKF filter relies on known values of the state variables at time zero, namely  $x_0$  and  $z_0$ , and additionally on the known parameter vector  $\boldsymbol{\theta}$ . Parameter estimation can be performed using least square fitting. Note that the prediction errors in the UKF iterations, namely  $\boldsymbol{\nu}_{t+1}$ , are a function of the parameters. The least square estimator for  $\boldsymbol{\theta}$  is then given by the solution to the following optimization problem:

$$\min_{\boldsymbol{\theta}} \sum_{t=1}^T \boldsymbol{\nu}_{t+1}' \boldsymbol{\nu}_{t+1}$$

The conditions under which this least squares estimator with the embedded UKF filter is consistent and asymptotically normal are provided in Ahn and Chan (2014).

## C Computation of posterior model probabilities

We follow the approach of Brown, Vannucci, and Fearn (1998). It is assumed that  $\mathbf{Y}$  is an  $n \times q$  random matrix, and that  $\mathbf{X}$  an  $n \times p$  matrix of fixed regressors. It is further assumed that  $p$  is large, and we are faced with the problem of selecting a subset of those regressors. Such subsets may be defined by using a  $p$ -dimensional vector variable  $\gamma$  with elements taken the value of zero or 1. If the  $j$ -th element of  $\gamma$  is one the  $j$ -th column regressor of  $\mathbf{X}$  is selected, if zero it is not selected. We further defined  $\mathbf{X}_\gamma$  as a matrix of dimension  $n \times p_\gamma$ , where  $p_\gamma$  is the number of ones in vector  $\gamma$ , which contains the regressors selected by the model defined by  $\gamma$ . We further assume

$$\mathbf{Y} - \mathbf{X}_\gamma \mathbf{B}_\gamma \sim N(\mathbf{I}_n, \Sigma)$$

where  $\mathbf{I}_n$  is an  $n$ -dimensional identity matrix. The unknown parameters are  $\mathbf{B}_\gamma$ , which is a  $p_\gamma \times q$  matrix, and the dispersion parameters  $\Sigma$  of dimension  $q \times q$ . These parameters are stochastic with well defined prior distributions. The vector  $\gamma$  is also random. The prior distribution of the parameters is assumed to be of the type:

$$p(\beta_\gamma, \Sigma, \gamma) = p(\beta_\gamma | \Sigma, \gamma) p(\Sigma) p(\gamma)$$

where  $\beta_\gamma = \text{vec}(\mathbf{B}_\gamma)$ . These priors are assumed to be of the type

$$\begin{aligned} \beta_\gamma | \Sigma, \gamma &\sim N(\beta_0, \Sigma \otimes g(\mathbf{X}'_\gamma \mathbf{X}_\gamma)^{-1}) \\ \Sigma &\sim iW(\Sigma_0, \nu_0) \end{aligned}$$

with  $g$  a constant parameter. Finally, the prior for  $\gamma$  is a multivariate Bernoulli with each element,  $\gamma_i$  being independent with individual probabilities  $\text{Prob}(\gamma_j = 1) = w_i$ , for  $w_i$  to be specified. This setting forms a multivariate equivalent to the so called Zellner (1986)  $g$  prior. Under this setting the posterior distribution of the parameters is thus simply given by:

$$p(\beta_\gamma, \Sigma, \gamma | \mathbf{Y}, \mathbf{X}) = p(\mathbf{Y} | \mathbf{X}, \beta_\gamma, \Sigma, \gamma) p(\beta_\gamma, \Sigma, \gamma)$$

where  $p(\mathbf{Y} | \mathbf{X}, \beta_\gamma, \Sigma, \gamma)$  is the likelihood term. The marginal posterior of  $\gamma$  is:

$$p(\gamma | \mathbf{Y}, \mathbf{X}) \propto \int p(\beta_\gamma, \Sigma, \gamma | \mathbf{Y}, \mathbf{X}) d\beta_\gamma d\Sigma$$

this probability is in effect proportional to the probability of that model defined by  $\gamma$ , and is shown in Brown, Vannucci, and Fearn (1998) to be proportional to:

$$p(\gamma | \mathbf{Y}, \mathbf{X}) \propto \left( \left| g(\mathbf{X}'_\gamma \mathbf{X}_\gamma)^{-1} \right| |\mathbf{K}_\gamma| \right)^{-q/2} |\mathbf{Q}_\gamma|^{-(n+\nu_0+q-1)/2} p(\gamma) \quad (\text{C-1})$$

with:

$$\begin{aligned} \mathbf{K}_\gamma &= \mathbf{X}'\mathbf{X} + \frac{1}{g}\mathbf{X}'_\gamma \mathbf{X}_\gamma \\ \mathbf{Q}_\gamma &= \Sigma_0 + \mathbf{Y}'\mathbf{Y} - \mathbf{Y}'\mathbf{X}\mathbf{K}_\gamma^{-1}\mathbf{X}'\mathbf{Y}. \end{aligned}$$

Table 1: Least Square estimation results of CDS pricing model.

	France	Germany	Greece	Ireland	Italy	Portugal	Spain
$\alpha_1$	-0.0001	0.0000	0.0010	0.0018	0.0001	-0.0001	0.0002
$\alpha_2$	0.0011	-0.0002	0.0000	0.0003	-0.0216	0.0000	-0.0006
$\beta_1$	1.0238	1.0083	1.0036	0.9925	0.9865	1.0215	0.9978
$\beta_2$	1.0058	0.9997	1.0047	1.0041	0.9645	1.0268	0.9978
$\sigma_{xx}$	4.4E-06	1.1E-06	7.4E-07	6.6E-07	1.3E-05	2.3E-05	2.2E-05
$\sigma_{zz}$	2.0E-05	8.4E-12	2.2E-06	4.4E-06	1.9E-05	3.5E-03	8.3E-07
$\sigma_{xz}$	1.6E-10	-1.2E-10	-2.6E-07	1.6E-06	-3.9E-08	2.1E-05	1.9E-06
$\sigma_{uu}$	4.0E-04	3.1E-06	1.2E-03	3.4E-06	2.4E-04	5.3E-01	6.6E-06
$x_0$	-0.005	-0.011	-0.002	-0.021	0.025	0.001	0.025
$z_0$	1.160	-1.436	1.238	-1.707	-2.150	0.748	-1.975
$\rho_{xz}$	0.000	-0.041	-0.200	0.943	-0.002	0.073	0.443
$R^2$	0.94	0.86	0.98	0.98	0.91	0.89	0.97
S.E.	0.25	0.37	0.40	0.24	0.31	0.64	0.22
$MAV$	4.1	6.5	46.9	17.8	15.4	38.7	11.5

NOTE:  $\alpha_1$ ,  $\alpha_2$ ,  $\beta_1$ , and  $\beta_2$  are the parameters of the state equation in model 5.  $\sigma_{ab}$  denotes covariance between  $a$  and  $b$ ,  $x_0$  and  $z_0$  are the respective starting values of  $x$  and  $z$ .  $\rho_{xz}$  is the correlation between the unobservable factors  $x$  and  $z$ .  $R^2$  is computed as one minus the ratio of the sum of squared residuals over the sum of squared dependent variables. S.E. denotes the residual standard error. Finally,  $MAV$  is the mean absolute value of the error term in basis points. Standard errors of the estimated parameters, computed by means of Monte Carlo simulations of the model, are shown in Table 2.

Table 2: Simulations of the CDS models

		$\alpha_1$	$\alpha_2$	$\beta_1$	$\beta_2$	$\sigma_{xx}$	$\sigma_{zz}$	$\sigma_{xz}$	$\sigma_{uu}$	$x_0$	$z_0$
Germany	true	0.00003	-0.00025	1.00830	0.99974	1.077E-06	8.390E-12	-1.245E-10	3.07522E-06	-0.0107	-1.4357
	mean	0.00003	-0.00025	1.00917	0.99703	1.070E-06	8.538E-12	-1.256E-10	3.1256E-06	-0.0108	-1.4486
	median	0.00003	-0.00025	1.00543	0.99824	1.077E-06	8.545E-12	-1.253E-10	3.14834E-06	-0.0108	-1.4465
	S.D.	0.00000	0.00001	0.01223	0.00512	4.820E-08	2.246E-13	3.266E-12	1.07125E-07	0.0002	0.0340
	relative error	2.4%	2.0%	1.2%	0.5%	4.5%	2.7%	2.6%	3.5%	2.0%	2.4%
France	true	-0.00006	0.00107	1.02376	1.00580	4.420E-06	2.032E-05	1.597E-10	0.00040	-0.0054	1.1602
	mean	-0.00005	0.00107	1.03375	1.00702	4.567E-06	2.107E-05	1.612E-10	0.00039	-0.0054	1.2143
	median	-0.00006	0.00108	1.03136	1.00292	4.455E-06	2.048E-05	1.607E-10	0.00040	-0.0054	1.1744
	S.D.	0.00000	0.00008	0.01956	0.00805	4.647E-07	2.436E-06	1.811E-11	6.15855E-05	0.0003	0.1024
	relative error	7.4%	7.6%	1.9%	0.8%	10.5%	12.0%	11.3%	15.5%	5.9%	8.8%
Greece	true	0.00099	0.00000	1.00358	1.00468	7.361E-07	2.230E-06	-2.565E-07	0.00120	-0.0024	1.2384
	mean	0.00100	0.00000	1.01748	1.00561	7.408E-07	2.256E-06	-2.596E-07	0.00121	-0.0024	1.2396
	median	0.00100	0.00000	1.01651	1.00570	7.385E-07	2.251E-06	-2.585E-07	0.00121	-0.0024	1.2432
	S.D.	0.00002	0.00000	0.00939	0.00209	3.793E-08	8.978E-08	9.426E-09	4.472E-05	0.0001	0.0304
	relative error	2.3%	3.0%	0.9%	0.2%	5.2%	4.0%	3.7%	3.7%	2.4%	2.5%
Ireland	true	0.00181	0.00035	0.99249	1.00409	6.554E-07	4.429E-06	1.606E-06	3.403E-06	-0.0207	-1.7069
	mean	0.00183	0.00034	1.00060	1.00511	6.385E-07	4.509E-06	1.584E-06	3.305E-06	-0.0219	-1.6868
	median	0.00184	0.00035	0.99174	1.00401	6.500E-07	4.495E-06	1.606E-06	3.370E-06	-0.0214	-1.7076
	S.D.	0.00012	0.00005	0.01914	0.00274	6.278E-08	5.333E-07	1.415E-07	5.852E-07	0.0019	0.0907
	relative error	6.6%	13.3%	1.9%	0.3%	9.6%	12.0%	8.8%	17.2%	9.0%	5.3%
Italy	true	0.00006	-0.02164	0.98650	0.96454	1.266E-05	1.937E-05	-3.881E-08	0.00024	0.0246	-2.1503
	mean	0.00006	-0.02184	0.97888	0.97491	1.275E-05	1.964E-05	-3.922E-08	0.00025	0.0248	-2.1501
	median	0.00006	-0.02181	0.98085	0.97643	1.272E-05	1.959E-05	-3.900E-08	0.00025	0.0248	-2.1548
	S.D.	0.00000	0.00070	0.02129	0.01001	4.933E-07	5.242E-07	1.476E-09	7.129E-06	0.0006	0.0473
	relative error	1.9%	3.3%	2.2%	1.0%	3.9%	2.7%	3.8%	2.9%	2.4%	2.2%
Portugal	true	-0.00007	0.00000	1.02148	1.02677	2.304E-05	3.465E-03	2.060E-05	0.53206	0.0015	0.7478
	mean	-0.00007	0.00000	1.03057	1.02253	2.054E-05	4.010E-03	1.878E-05	0.51163	0.0015	0.8619
	median	-0.00007	0.00000	1.02846	1.02206	2.288E-05	3.503E-03	2.052E-05	0.53208	0.0015	0.7579
	S.D.	0.00002	0.00000	0.01870	0.01156	7.058E-06	1.744E-03	5.089E-06	0.20584	0.0005	0.2136
	relative error	29.6%	19.3%	1.8%	1.1%	30.6%	50.3%	24.7%	38.7%	31.9%	28.6%
Spain	true	0.00019	-0.00059	0.99775	0.99780	2.209E-05	8.267E-07	1.891E-06	6.570E-06	0.0252	-1.9754
	mean	0.00019	-0.00060	1.00532	0.99823	2.282E-05	8.368E-07	1.903E-06	6.580E-06	0.0251	-1.9862
	median	0.00019	-0.00060	0.99980	0.99827	2.267E-05	8.251E-07	1.916E-06	6.649E-06	0.0253	-1.9815
	S.D.	0.00001	0.00003	0.01939	0.00370	1.544E-06	4.966E-08	8.079E-08	4.225E-07	0.0012	0.0482
	relative error	7.7%	4.9%	1.9%	0.4%	7.0%	6.0%	4.3%	6.4%	4.9%	2.4%

NOTE:  $\alpha_1$ ,  $\alpha_2$ ,  $\beta_1$ , and  $\beta_2$  are the parameters of the state equation in model 5.  $\sigma_{ab}$  denotes covariance between  $a$  and  $b$ ,  $x_0$  and  $z_0$  are the respective starting values of  $x$  and  $z$ . "True" are the estimated parameter values used in simulations. "Mean", "Median", and "S.D." are mean, median, standard deviations of estimated parameters from a number of simulations. "relative error" is the standard deviation divided by the true parameter value. 100 simulations were run for each model.

Table 3: Key events and news in the euro area during the financial crisis.

Date	Description	Grouping
	<i>Institutional Developments</i>	
Aug-2007	ECB provides liquidity assistance to calm tensions in interbank market	Market functioning
Jan-2008	US and EA economic indicators point to recession	Ec & Fiscal Stimulus
Mar-2008	ECB offers refinancing operations with longer maturities	Market Functioning
Jul-2008	ECB increases interest rates by 25 bps	Standard MP
Sep-2008	Lehman Brothers files for bankruptcy	Lehman Brothers
Oct-2008	Unlimited liquidity provision of ECB to banks	Market Functioning
	Launch of national bail-out plans for banks	Bank Support
	ECB cut rates by 50 bps	Standard MP
Nov-2008	Announcement of European Economic Recovery Plan	Ec & Fiscal Stimulus
	ECB cuts rates by 50 bps	Standard MP
Dec-2008	ECB cuts rates by 75 bps	Standard MP
Jan-2009	ECB cuts rates by 50 bps	Standard MP
Mar-2009	ECB cuts rates by 50 bps	Standard MP
Apr-2009	ECB cuts rates by 25 bps	Standard MP
May-2009	Announcement of European Financial Supervision Plan	Support Banks
Jun-2009	ECB launches covered bond programme	Market Functioning
May-2010	ECB announces Securities Markets Programme	Market Functioning
Jun-2010	European Financial Stability Facility launched	Government Support
Jul-2010	EU banks stress tests published	Bank Support
Jun-2010	Effective lending capacity of EFSF is increased	Government Support
Jul-2011	EU banks stress tests published	Bank Support
	ECB increases rates by 25 bps	Standard MP
Aug-2011	ECB's SMP programme is extended to Spain and Italy	Market Functioning
Dec-2011	ECB announces additional non-standard measures (3-year LTRO)	Market Functioning
	EU agrees on new fiscal compact limiting structural deficits	Ec & Fiscal Stimulus
Jun-2012	EU Council announces Compact for Growth and Jobs	Ec & Fiscal Stimulus
Jul-2012	President Draghi London speech in support of Euro	Market Functioning
	<i>Country News</i>	
Nov-2009	Greek announces public deficit would exceed 12%	News Greece
Apr-2010	announcement of EU/IMF financial support for Greece	News Greece
Oct-2010	Speculation on EU discussions on possible debt restructuring	News Ireland
Nov-2010	Irish government seeks financial support	News Ireland
Dec-2010	Irish package is agreed	News Ireland
Apr-2011	Portuguese government requests activation of aid mechanism	News Portugal
Jun-2011	DE and FR agree on need for PSI in Greek crisis	News Greece
Jul-2011	Eurogroup sets PSI as precondition for further aid to GR	News Greece
Oct-2011	Restructuring on GR debt deemed larger than previously envisaged	News Greece
Nov-2011	Announcement of referendum in Greece on aid assistance	News Greece
Mar-2012	Agreement on the (PSI) restructuring of Greek debt	News Greece

NOTE: bps are used to denote basis points.



Table 4: BMA regression results.

	Germany		France		Greece		Ireland		Italy		Portugal		Spain	
	PD	LGD	PD	LGD	PD	LGD	PD	LGD	PD	LGD	PD	LGD	PD	LGD
	With ‘ <i>permanent-effect dummies</i> ’													
VIX	-	0.03**	-	-	0.03**	0.17**	-	-	-	-	-	0.08**	-	0.04**
VSP	-	-	-	-	-	-	-	-	-	-	-	-	-	-
RAT	-	-	-	-	0.23**	-	-	-	-	-	0.39**	-	-	-
iTraxx	0.61**	-	0.42**	-	0.39**	0.21**	-	-	0.66**	-	-	0.07**	0.57**	-
BCDS	-	-	-	-0.20**	-	-0.03**	-	-	-	-	-	-	-	-
FUT-SER	-	-	-	-	-	-	-	-	-	0.04**	-	-	-	-
FUT-GDP	-	-	-	0.17**	-	-	-	-	-	-	-	-	-	-
DEF	-	-	-	-	-	-	-0.04**	-	-	-	-	-	-	-
BC	-	-	-	-	-	-	-	-	-	-	-	-	-	-
Market Functioning	-	-0.04**	-	-	-	-	-	-	-	-	-	-	-	-
Standard MP	-	-	-	-	-	-	-	-	-	-	-	-	-	-
Government Support	-	-	-	-	0.12**	-	0.03**	-	-	-	-	-	-	-
Bank Support	-	-	-	-	-	-	-	-	-	-	-	-	-	-
Ec & Fiscal Stimulus	-	0.03**	-	-	-	-	-	-	-	-	-	-0.11**	-	-
Lehman Brothers	-	-	0.21**	-	-	-	-	0.27**	-	-	-	-	-	0.07**
News Greece	-	-	-	-	-	-	-	-	-	-	-	-	-	0.05**
News Ireland	-	-	-	-	-	-	-	-	-	-	-	-	-	-
News Portugal	-	-	-	-	0.27**	-	0.03**	-	-	-	-	-	-	-
$R^2$	0.37	0.05	0.31	0.20	0.38	0.23	0.04	0.09	0.44	0.03	0.16	0.12	0.33	0.06
	With ‘ <i>temporary-effect dummies</i> ’													
VIX	-	-	0.13**	-	0.08**	0.17**	-	0.02**	-	-	-	0.16**	-	0.10**
VSP	-	-	-	-	-	-	-	-	-	-	-	-	-	-
RAT	-	-	-	-	0.28**	-	-	-	-	-	0.39**	-	-	-
iTraxx	0.60**	-	0.31**	-	0.18**	0.20**	-	-	0.65**	-	-	0.20**	0.57**	-
BCDS	-	-	-	-0.20**	-	-0.03**	-	-	-	-	-	-	-	-
FUT-SER	-	-	-	-	-	-	-	-	-	-	-	-	-	-
FUT-GDP	-	-	-	0.17**	-	-	-	-	-	-	-	-	-	-
DEF	-	-	-	-	-	-	-	-	-	-	-	-	-	-
BC	-	-	-	-	-	-	-	-	-	-	-	-	-	-
Market Functioning	-	-	-	-	-	-	-	-	-	-	-	-	-	-
Standard MP	-	-	-	-	-	-	-	-	-	-	-	-	-	-
Government Support	-	-	-	-	-	-	-	-	-	-	-	-	-	-
Bank Support	-	-	-	-	-	-	-	-	-	-	-	-	-	-
Ec & Fiscal Stimulus	0.14**	-	-	-	-	-	-	-	-	-	-	-	-	-
Lehman Brothers	-	-	0.13**	-	-	-	-	-	-	-0.16**	-	-	-	-
News Greece	-	-	-	-	-	-	-	-	-	-	-	-	-	-
News Ireland	-	-	-	-	-	-	-0.43**	0.44**	-	-	-	-	-	-
News Portugal	-	-	-	-	-	-	-	-	-	-	-	-	-	-
$R^2$	0.37	0.08	0.29	0.21	0.29	0.23	0.19	0.25	0.43	0.04	0.15	0.18	0.33	0.04

NOTE: Data has been normalised, results in the table relate to regressions conducted for the two-year PDs and LGDs. Values reported for the regression coefficients are only those significant and larger than 0.01. \* denotes significance at a 5% level of significance.

Table 5: Contagion across sovereign markets. Correlation of residuals.

	With ' <i>permanent-effect dummies</i> '											
	PD residuals						LGD residuals					
	France	Germany	Greece	Ireland	Italy	Portugal	France	Germany	Greece	Ireland	Italy	Portugal
Germany	0.57*						0.43*					
Greece	-0.09	0.08					0.35*	0.01				
Ireland	0.21	0.22	0.03				0.16	0.20	0.17			
Italy	0.20	0.37*	0.06	0.10			-0.08	-0.05	-0.08	-0.16		
Portugal	0.27*	0.15	-0.06	0.00	0.19		-0.08	-0.14	-0.11	0.31*	-0.04	
Spain	0.14	0.14	-0.24	0.21	0.34*	0.19	0.38*	0.19	0.30*	0.28*	-0.18	0.23
	With ' <i>temporary-effect dummies</i> '											
	PD residuals						LGD residuals					
	France	Germany	Greece	Ireland	Italy	Portugal	France	Germany	Greece	Ireland	Italy	Portugal
Germany	0.57*						0.46*					
Greece	-0.13	0.04					0.35*	0.02				
Ireland	0.23	0.19	-0.04				0.17	0.21	0.09			
Italy	0.20	0.38*	0.04	0.19			-0.08	-0.08	-0.09	-0.19		
Portugal	0.27*	0.15	-0.09	0.03	0.19		-0.07	-0.19	-0.12	0.28*	-0.05	
Spain	0.14	0.14	-0.21	0.33*	0.35*	0.19	0.39*	0.20	0.30*	0.35*	-0.18	0.20

NOTE: Correlations significant at the 5% level are denoted with an asterisk. Correlations have been computed with residuals of regressions of PD and LGD within 2 year maturity.

Figure 1: PD and LGD combinations providing model-implied CDS spreads which fit perfectly the observed CDS spreads over different maturities on December 2010.

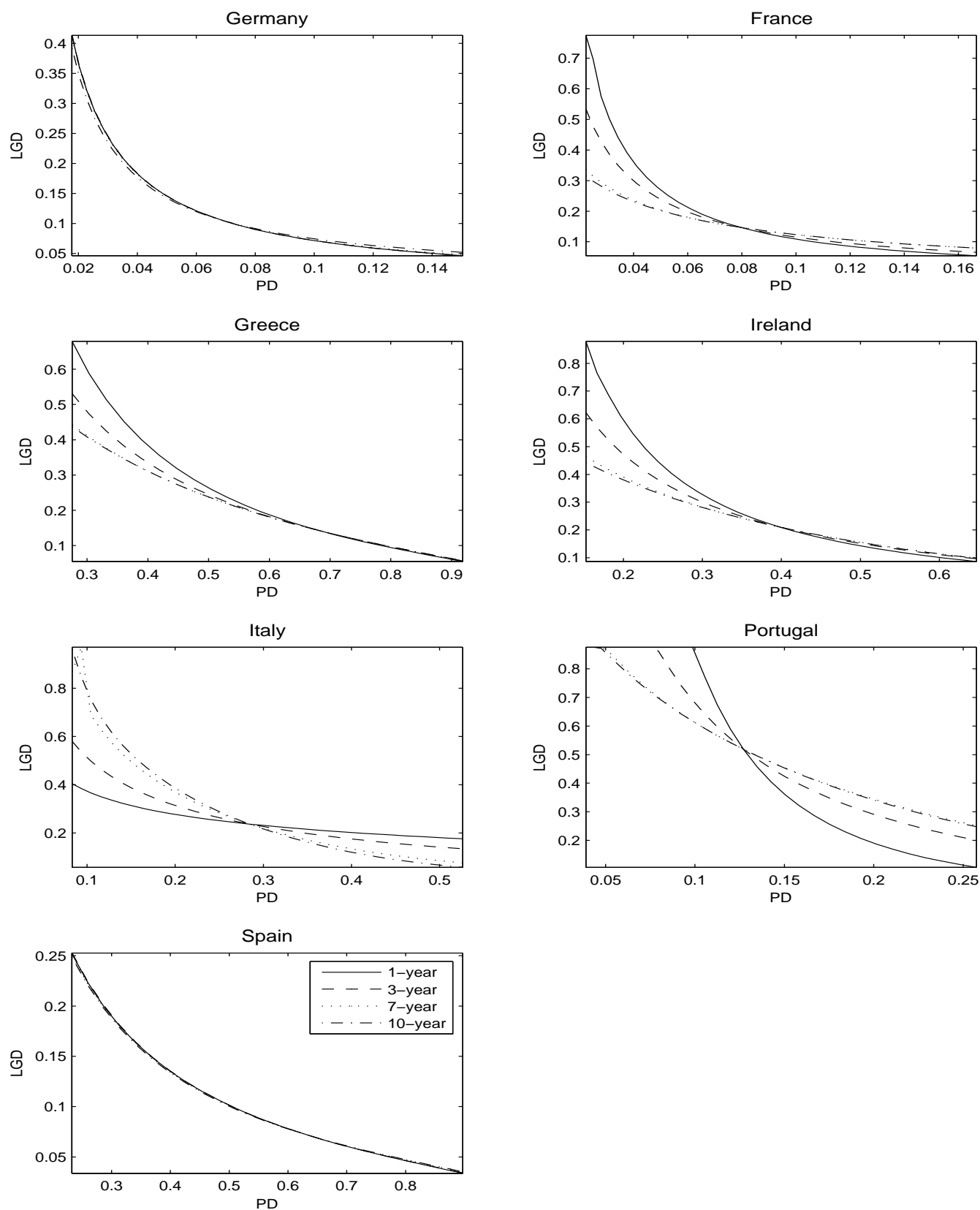


Figure 2: PD and LGD combinations providing model-implied CDS spreads within the estimated mean absolute error from observed CDS spreads over different maturities on December 2010.

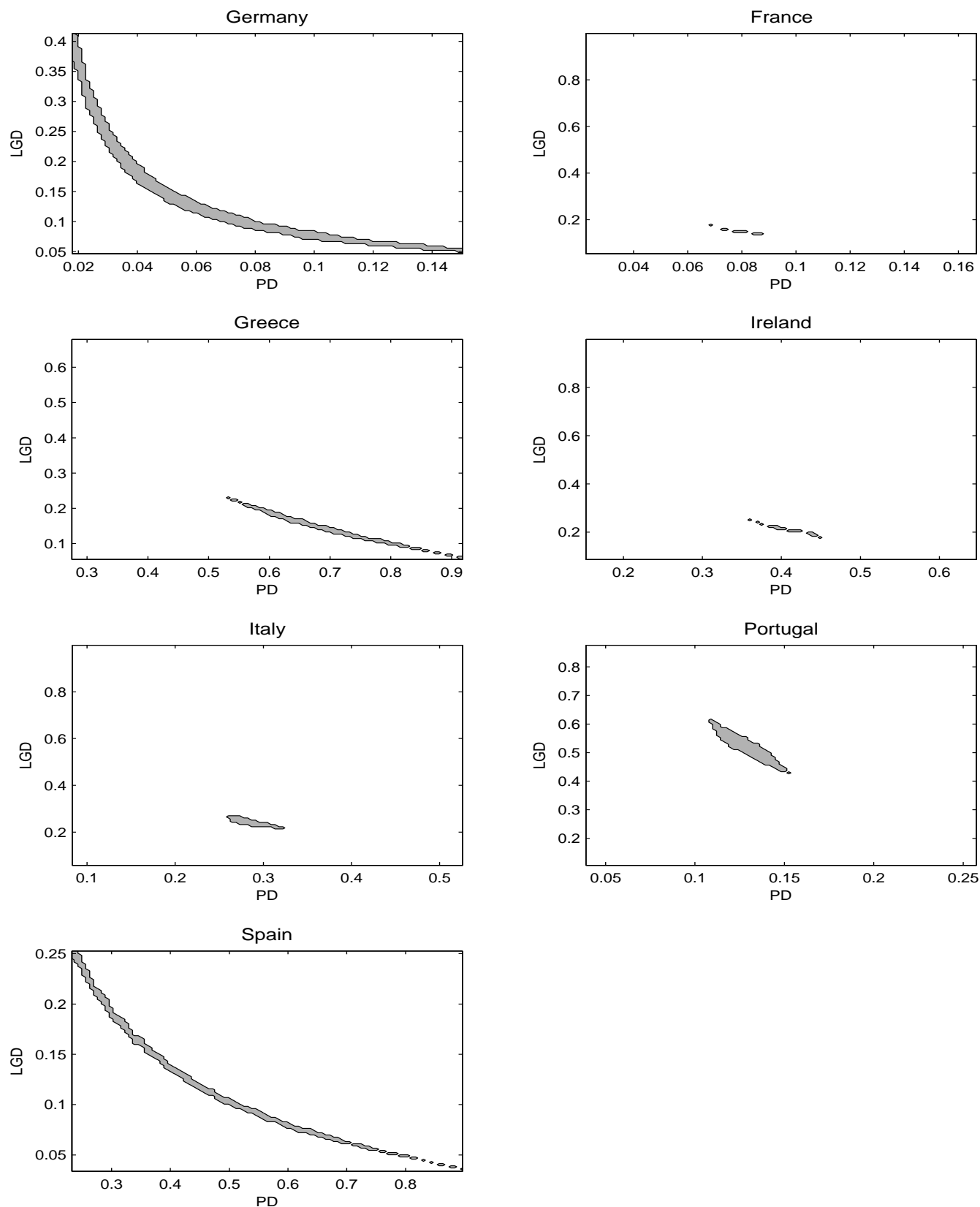


Figure 3: PD and LGD within 2 years.

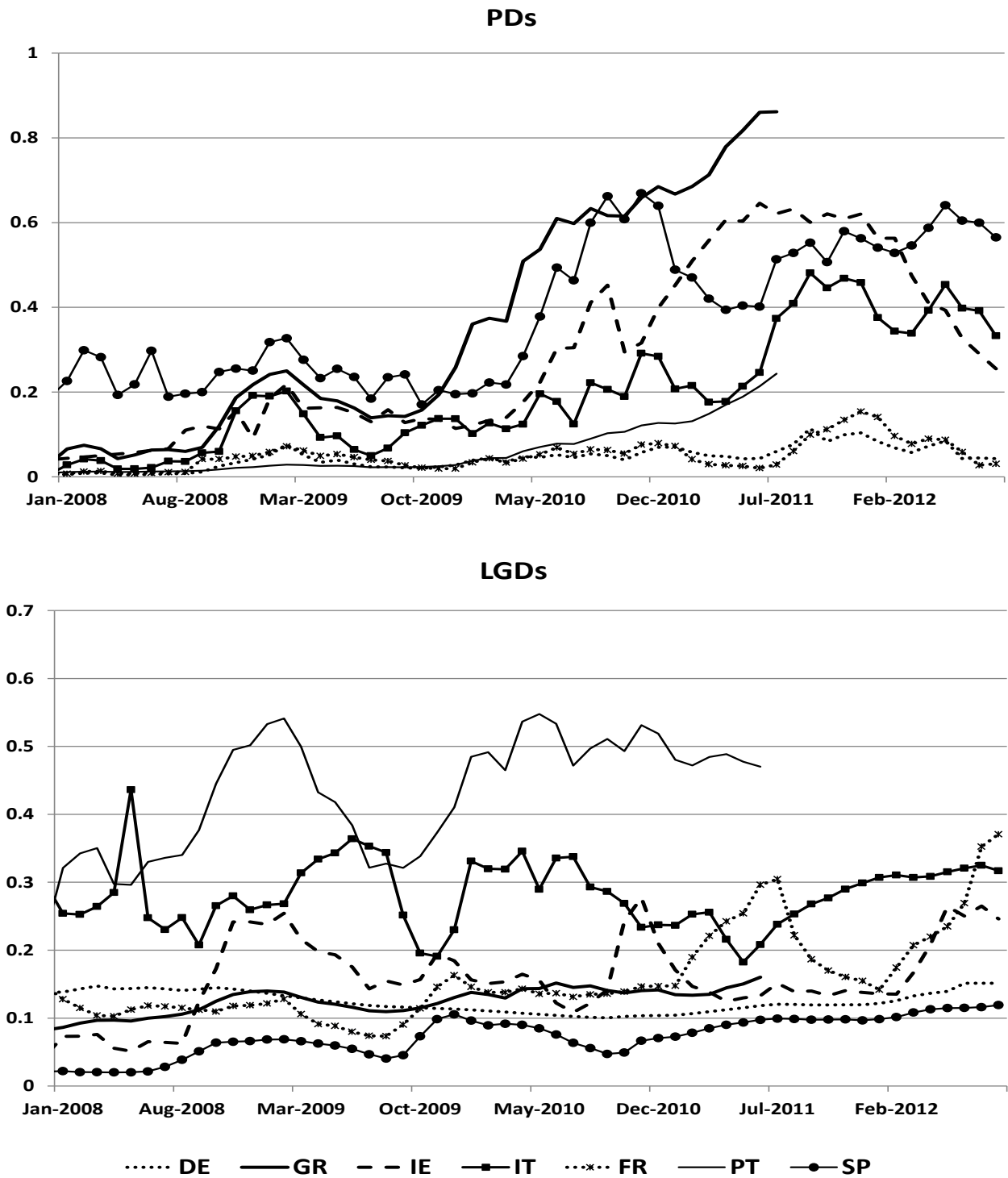
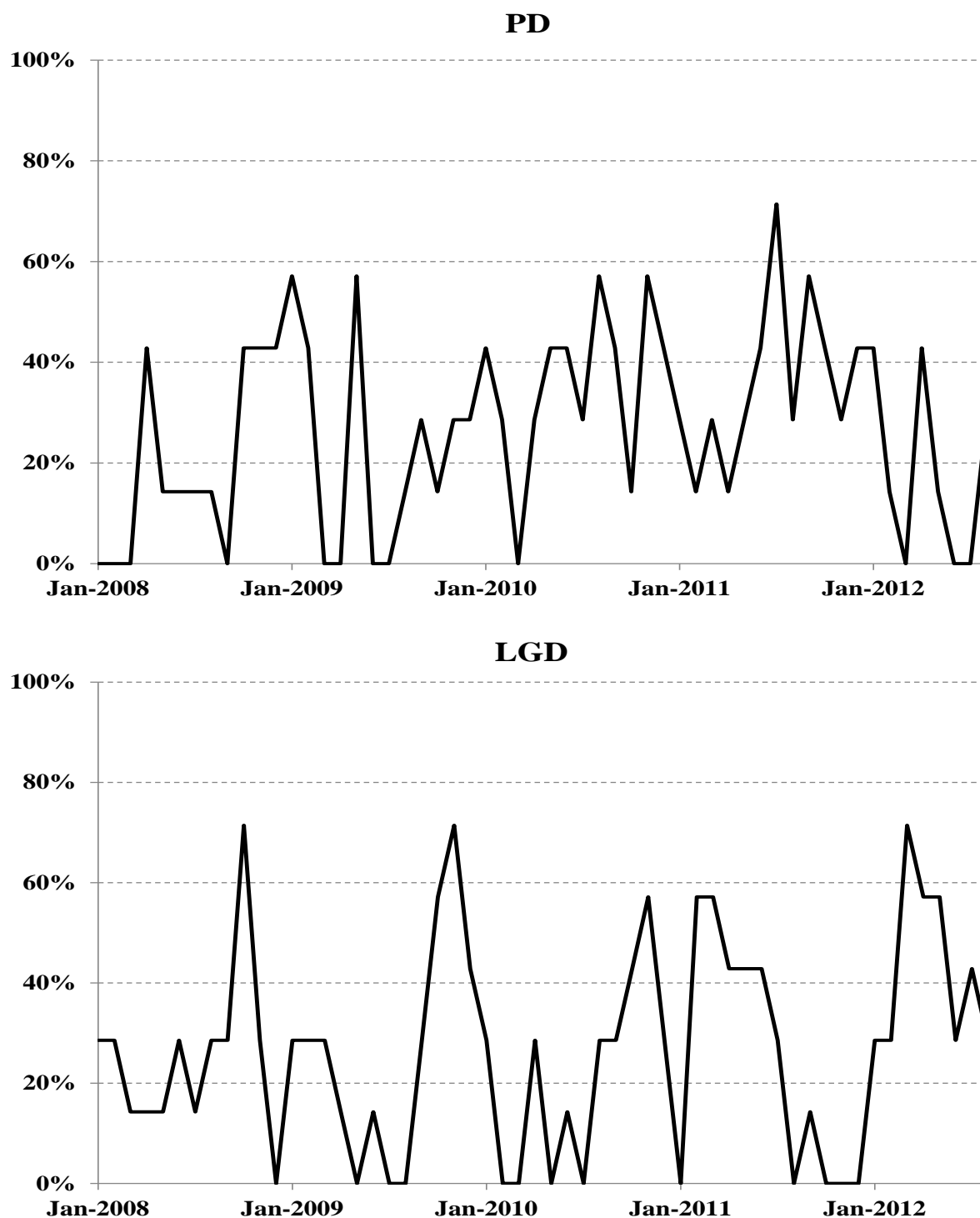


Figure 4: Coincidences of large residuals. With ‘*permanent-effect dummies*’.



NOTE: Coincidences are measured as the ratio of the number of residual values above their empirical 75th percentiles to the total number of the analysed countries in a given period. The residual values are estimated from the regressions of changes in PD (top chart) and LGD (bottom chart) on the set of economic fundamentals as explained in the main text.