Second Dimension Risk – A Reduced Form Analysis of European Sovereigns' Credit Spreads

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Abstract

This article analyzes a possible risk premium for uncertainties regarding the current and future default probabilities in the context of the European fiscal crisis. It is argued that this risk premium was an important driver of credit spreads on a singly country level and that it has catalyzed sovereign credit contagion effects in the past. The relevance of this risk premium in the context of the fiscal crisis is then empirically analyzed based on a doubly stochastic reduced form credit risk model.

Keywords: Sovereign Credit Risk, Variance Premium, European Fiscal Crisis

JEL Classification: G12, G13

1 Introduction

Sovereign credit spreads have become an important topic of debate and research in recent years. The strong increase in certain European countries' credit costs has triggered the current European financial crisis by exposing massive problems to refinance at costs that are affordable in the long run.

This article analyzes one possible driver of credit spreads during the European fiscal crisis: the risk premium that market participants expect because of uncertainties with respect to prospective and current default probabilities. This risk premium does not directly refer to the possibility of a default *per se* at a given default probability, but it refers to the possibility of unfavorable corrections regarding the default probability. We refer to this premium by the term "second dimension" risk premium.
It is argued in this article that such a risk premium was a very relevant driver of both sovereign credit spreads and correlations of sovereign credit spreads. In the centre of this argument are surprising insights into member countries' true fiscal situations. To test these hypotheses, we estimate a doubly stochastic reduced form credit risk model for several European countries under both a risk-neutral and the actual measure along the likes of Pan and Singleton (2008). In this context, it is tested whether second dimension risk premiums' correlation increased after the outbreak of the Greek crisis. The model estimation is based on European sovereign "credit default swap" (CDS) data for the years 2008-2012.

Previous work in this area was mostly devoted to other geographic areas or other periods and suggests that sovereign credit spreads are mainly driven by global financial market risk factors approximated by measures like the implied volatility index VIX (see e.g. Kaminsky and Reinhart (2002), Pan and Singleton (2008), respectively Longstaff et al. (2011), Favero et al. (2010), Zhou et al. (2013), Baek et al. (2005), Eichengreen and Mody (2000), Mauro et al. (2002), Remolona et al. (2008), Geyer et al. (2004)). Country specific economic data,

on the other hand, did not seem to be very important (see e.g. Alper et al. (2012)).
This research describes the correlation of sovereign credit spreads as rather strong (see e.g. Kaminsky and Reinhart (2002)), which is often assumed to be mainly caused by global financial market risk measures being important drivers of sovereign spreads. The described findings are also supported in the European sovereign context for the years before 2008 (De Santis (2012)). The explanatory power of variables like the VIX index with respect to European sovereign credit spreads decreased strongly during the past few years. The co-movement between spreads of specific countries stays high for this period (c.f. De Santis (2012)). De Santis (2012) suggests that in these years – in the cases of sovereigns like Portugal, Ireland or Spain – spreads are instead largely affected by contagion effects going back to the Greek crisis. This contagion may have been enforced by the bank rescue packages and the related risk transfer from banks to sovereigns (c.f. Ejsing and Lemke (2011)). A detailed understanding of how that contagion could have worked technically in the context of the European fiscal crisis is, however, still missing. The present article provides evidence on the relevance of the above risk premium and argues that this risk premium was an important driver of these contagion effects.

Longstaff et al. (2011) and Pan and Singleton (2008) analyze the relevance of the second dimension risk premium in a framework similar to the one established here. Their results suggest that the risk premium is highly relevant for the included sovereigns' credit spreads during the respective years. These articles are in opposition to the present one not based on sovereign credit data from the years of the European fiscal crisis and the possible interplay between the events during the European fiscal crisis and the examined risk premium is not analyzed. Moreover, the present article examines correlation between sovereign credit spreads in the context of the second dimension risk premium and whether general financial market nervousness is a relevant factor in this context.

2 The modelling framework

We consider a measure space $(\Omega, \mathcal{F}_1, P)$, an index set $S \neq \emptyset$[1] and a Poisson process

$$\mathcal{P}oi = (Poi_s, s \in S) \tag{1}$$

driven by the intensity λ_s. This Poisson process generates a filtration $\mathcal{F}_{1,s}$: $\mathcal{F}_{1,s} = \sigma\{Poi_t : 0 \leq t \leq s\}$ with $t \in S$. In this model, the default of a unit is in this model defined as a first jump of this Poisson process and the time of the first jump denoted as $\tau \in S$ is therefore the stopping time for this process as well. No-arbitrage pricing formulas for all kinds of credit risk related securities have been derived based on that and Lando (1998) presents for example pricing formulas for simple zero bonds.

For this example, define a zero bond with face value one, issued at time $s_0 \in S$, with a recovery rate $1-LR$ (denoting the fraction of the face value which is paid in the case of default right after the default occurred), maturity M (denoting the number of years until the principal is paid back) with $[s_0, s_0+M] \subset S$, and payoff Z_s for $s \in S$, with $Z_{s_0+M} = 1$ and $Z_{s'} = 0$ for all $s' \neq s_0 + M$ if $\tau \notin [s_0, s_0+M]$ as well as $Z_\tau = 1 - LR$ and $Z_{s''}$=0 for all $s'' \neq \tau$ if $\tau \in [s_0, s_0+M]$. Lando (1998) shows that for deterministic intensities the market price ZB_{s_0,s_0+M} of this bond in s_0 is for deterministic intensities given by:

$$ZB_{s_0,s_0+M} =$$
$$\mathbb{E}_{s_0}\left[e^{-\int_{s_0}^{s_0+M}\lambda_s^{\mathbb{Q}}+r_s^f ds}|\mathcal{F}_{1,s_0}\right] + (1-LR)\left[\int_{s_0}^{s_0+M}\mathbb{E}_{s_0}\left[\lambda_s^{\mathbb{Q}}e^{-\int_{s_0}^{s}\lambda_u^{\mathbb{Q}}+r_u^f du}|\mathcal{F}_{1,s_0}\right]ds\right], \tag{2}$$

where r_s^f denotes the risk free rate and the resulting discount factor complies with $ZB^f_{s_0,s_0+m}$ denoting the price of a risk free zero bond issued in s_0 with maturity m. r_s denotes the return expected by the investors in this zero bond and $\lambda_s^{\mathbb{Q}}$ denotes the risk neutral default intensity that allows to switch from r_s

[1] One time unit refers in the context of the estimation, which is discussed later on, to one year (and not one day).

to r_s^f. This framework is now extended for allowing stochastic intensities which follow a Cox-Ingersoll-Ross (CIR) diffusion:

$$d\lambda_s = (\mu_0 - \mu_1 \lambda_s) + \sigma_1 \sqrt{\lambda_s} dB_s \tag{3}$$

with B_s denoting a Brownian motion and μ_0, μ_1 and σ_1 being constant coefficients. The intensity process generates a filtration $\mathcal{F}_{2,s} = \sigma\{\lambda_t : 0 \leq t \leq s\}$ as well with $t, s \in S$. Finally, a filtration $\mathcal{F}_s$ is defined as

$$\mathcal{F}_s = \sigma\{\mathcal{F}_{1,s} \vee \mathcal{F}_{2,s}\}, \ \textit{for all } s \in S \tag{4}$$

with "$\vee$" in this context denoting the union of σ-fields respectively filtrations.

After the introduction of this second uncertainty dimension, equation 2 does not necessarily hold anymore. Accordingly, two equivalent probability measure with respect to $\lambda_s^{\mathbb{Q}}$ are introduced: $\widehat{\mathbb{Q}}$ and $\widehat{\mathbb{P}}$. The latter refers to the actual distribution law of $\lambda_s^{\mathbb{Q}}$ and $\widehat{\mathbb{Q}}$ refers to expectations with respect to (transforms of) $\lambda_s^{\mathbb{Q}}$ so that the pricing formula including a discount rate based on r_s^f holds despite of the possible existence of the respective "second dimension" risk premium. The expectations based on these distribution laws are denoted by $\mathbb{E}_s^{\widehat{\mathbb{P}}}$ respectively $\mathbb{E}_s^{\widehat{\mathbb{Q}}}$ in the following – following Pan and Singleton (2008) and Longstaff et al. (2011) – one rewrites for the zero-bond pricing formula 8:

$$\mathbb{E}_{s_0}^{\widehat{\mathbb{Q}}} \left[e^{-\int_{s_0}^{s_0+M} \lambda_s^{\mathbb{Q}} + r_s^f ds} | \mathcal{F}_{2,s_0} \right] + (1-LR) \left[\int_{s_0}^{s_0+M} \mathbb{E}_{s_0}^{\widehat{\mathbb{Q}}} \left[\lambda_s^{\mathbb{Q}} e^{-\int_{s_0}^{s} \lambda_u^{\mathbb{Q}} + r_u^f du} | \mathcal{F}_{2,s_0} \right] ds \right] \tag{5}$$

The distinction between the two distribution laws of $\lambda_s^{\mathbb{Q}}$ requires another notation of the diffusion equations driving $\lambda_s^{\mathbb{Q}}$ under both measures. Following Longstaff et al. (2005), Pan and Singleton (2008) and Longstaff et al. (2011)), one rewrites the underlying diffusion equations under the risk neutral measure $\widehat{\mathbb{Q}}$ as

$$d\lambda_s^{\mathbb{Q}} = \left(\mu_0^{\widehat{\mathbb{Q}}} - \mu_1^{\widehat{\mathbb{Q}}} \lambda_s^{\mathbb{Q}} \right) ds + \sigma_1 \sqrt{\lambda_s^{\mathbb{Q}}} dB_s^{\widehat{\mathbb{Q}}} \tag{6}$$

respectively under the actual measure $\widehat{\mathbb{P}}$

$$d\lambda_s^{\mathbb{Q}} = \left(\mu_0^{\widehat{\mathbb{P}}} - \mu_1^{\widehat{\mathbb{P}}}\lambda_s^{\mathbb{Q}}\right) ds + \sigma_1\sqrt{\lambda_s^{\mathbb{Q}}} dB_s^{\widehat{\mathbb{P}}}. \tag{7}$$

If market participants do not expect a specific remuneration for taking the uncertainty regarding $\lambda_s^{\mathbb{Q}}$, the expectations under $\widehat{\mathbb{P}}$ and $\widehat{\mathbb{Q}}$ with respect to this risk neutral intensities respectively the transforms included in these pricing formulas should not differ. The opposite is the case if the change in expected returns due to this uncertainty is high. The relevance of the "second dimension" risk premium can be analyzed accordingly based on the coefficients of these diffusion equations under both measures. In this context we calculate the Radon-Nikodym density based on the Girsanov theorem to describe the change from the actual measure $\widehat{\mathbb{P}}$ to the risk neutral measure $\widehat{\mathbb{Q}}$. The process, this Radon-Nikodym density depends on, is called "market price of risk" (c.f. Pan and Singleton (2008)).

Moreover, a specific functional form linking η_s and $\lambda_s^{\mathbb{Q}}$ is assumed. The specific form is chosen based on the plausible assumption that the increase in change in the respective intensity should be linear in this intensity (c.f. Cheridito et al. (2007) and Duffee (2002)). It is accordingly assumed that η_s depends on $\lambda_s^{\mathbb{Q}}$ in the following way:

$$\eta_s = \frac{\rho_0}{\sqrt{\lambda_s^{\mathbb{Q}}}} + \rho_1\sqrt{\lambda_s^{\mathbb{Q}}}. \tag{8}$$

This results in the actual difference in change of $\lambda_s^{\mathbb{Q}}$ being given by

$$\sigma_1\left(\rho_0 + \rho_1\lambda_s^{\mathbb{Q}}\right). \tag{9}$$

This implies the following link between ρ_0, ρ_1 and the CIR coefficients under both measures:

$$\rho_0 = \frac{\mu_0^{\widehat{\mathbb{Q}}} - \mu_0^{\widehat{\mathbb{P}}}}{\sigma_1}; \quad \rho_1 = \frac{\mu_1^{\widehat{\mathbb{P}}} - \mu_1^{\widehat{\mathbb{Q}}}}{\sigma_1}. \tag{10}$$

Accordingly, η_s refers to the change in the deterministic drift induced by a change from the historical to the risk neutral measure at a specific point in time.

For the present study, the estimation of such a doubly stochastic reduced form credit risk model has been based on historical spreads of "credit default swaps" (CDS). The pricing formula is, as described in Duffie (1999), given by

$$\begin{aligned} & SP_{s_0}(M)\sum_{n=1}^{2M}\left(\mathbb{E}_{s_0}^{\widehat{\mathbb{Q}}}\left[e^{-\int_{s_0}^{s_0+0.5n}\lambda_s^{\mathbb{Q}}ds}|\mathcal{F}_{2,s_0}\right]ZB_{s_0,s_0+0.5n}^{f}\right) \\ & = LR\left[\int_{s_0}^{s_0+M}ZB_{s_0,s}^{f}\mathbb{E}_{s_0}^{\widehat{\mathbb{Q}}}\left[\lambda_s^{\mathbb{Q}}e^{-\int_{s_0}^{s}\lambda_u^{\mathbb{Q}}du}|\mathcal{F}_{2,s_0}\right]ds\right]. \end{aligned} \tag{11}$$

For CDS, the loss rate complies with a face value. The index s_0 in this context refers not only to the point in time when a single contract is issued, but it is in turn index for the historical time series of CDS spreads used for the estimation. Following Duffie et al. (2003), it is assumed that the total default intensity λ_s combines the probabilities of different kinds of credit events like liquidation events or restructuring with λ_s being the sum of intensities referring to one particular default event. The loss rate is then correspondingly the average of the loss rates for all the different credit events, weighted by the particular probabilities.

The analysis is executed for several European countries: Spain, Ireland, Iceland, Estonia, Finland, Poland. These countries can be classified by different criteria: membership in the Euro area (this excludes Iceland and Poland) and countries that have been in acute stress during the crisis (this excludes Finland and Poland). The spreads for Estonia and Iceland have, moreover, decreased significantly from the first part of the sample to the second part, while the opposite can be said for Ireland and Spain. The sample period is from October 2008 to march 2012. The reason for not using earlier data is that historical CDS spread data does not reach back very far as CDS is rather a new security type. The historical CDS spread time series were supplied by Thomson-Reuters. The spreads for the particular period are depicted in figure 1. Both the spreads' strong increase during this period and the similarity in time series patterns is striking.

The prices for risk free zero bonds are approximated by using prices of zero bonds issued by AAA rated units. These prices are calculated based on the spot rate curve published by the ECB. The published data points (every three months with a range from three months to 30 years) are linked by linear interpolation.

3 The "second dimension risk premium" and the European fiscal crisis

The second dimension risk is highly relevant for spreads of both European countries actually struggling during the fiscal crisis and countries which have not been in acute distress. Revealed uncertainty regarding the current and future fiscal situations are an important aspect of the fiscal crisis. The Greek government significantly corrected previously published fiscal information[2]. This should have lead to a twofold increase in Greek credit spreads: On the one hand, the credit spreads increased due to an actual increase in the currently assumed default probability related to the actual deterioration of the observed Greek fiscal situation. The fact that the presumptions regarding the country's fiscal situation are based on information which turned out to be not very robust could have on the other hand lead to an increase in the second dimension risk and the related premium as well.

The strong corrections of fiscal information could also have lead to a twofold increase in other European sovereigns credit spreads. The default probabilities which are currently expected for other European sovereigns have increased since the real economic outlook for these countries had deteriorated due to the difficulties in Greek. In addition, the general sceptism toward fiscal information published by European sovereigns had increased and future default probabilities of other European countries were considered to be more uncertain than before and the second dimension risk premia might have increased accordingly.

A factor driving the uncertainty regarding the default probabilities of several

[2]In November 2009, "the Greek government revealed a revised budget deficit of -12.7% of GDP for 2009, which was the double of the previous estimate" (c.f. De Santis (2012))

countries at the same time could, for example, be the reputation of certain institutions. By accepting countries as members of the Euro zone, European institutions likewise implicitly rate both their fiscal information and their fiscal stability as sufficient. Being accepted as member in the Euro zone has however lost its characteristic as a positive signal in the course of the European fiscal crisis. Market participants' uncertainty regarding the assessment of the other countries' financial situations has since increased, even if the level of these other countries' default probabilities may not be impacted directly by a change in information with respect to the situation of the first country.
The second dimension risk premium might also have catalyzed correlation between these countries credit spreads during the European fiscal crisis. A comovement between two countries' credit spreads might be induced by the existence of a second dimension risk premium if these countries' risk premium components are driven by common factors. Such a factor might of course be the market participants' risk appetite itself, but it could as well be a common source driving the market participants' uncertainty regarding current and future default probabilities of two countries, like institutional quality.

Summing up, the second dimension risk premium might have been an important driver of sovereign credit spreads in Europe. Moreover, it might have been an important driver of the observed comovement between sovereign credit spreads respectively the contagion during the European fiscal crisis as well. In the following, the estimation of such a model is discussed and credit insurance securities are introduced in the presented framework.

4 Estimation and results

The key of the empirical analysis is the estimation and the comparison of the CIR coefficients under measure $\widehat{\mathbb{Q}}$ and measure $\widehat{\mathbb{P}}$.

4.1 Estimation under the risk neutral measure

For the applied estimation strategy, the set of coefficients $\{\widehat{\mu}_0^{\widehat{\mathbb{Q}}}, \widehat{\mu}_1^{\widehat{\mathbb{Q}}}, \widehat{\sigma}_1, \widehat{LR}\}$ is assumed ex-ante and the expectations in the pricing formula are substituted by the exponential linear functions depending on the realization of $\lambda_{s_0}^{\mathbb{Q}}$ and the horizon of the expectation as shown by Duffie and Singleton (1999). The coefficients of this exponential linear function are solutions to ordinary differential equations that only depend on the coefficients of the CIR-process. Based on that, an estimation $\widehat{\lambda}_{s_{0_i}}^{\mathbb{Q}}$ can then be obtained for each observation $s_{0_i} \in [s_{0_1}, s_{0_2}.., s_{0_N}]$. This is done based on the 5-year spreads. The extracted time series $\widehat{\lambda}_{s_{0_i}}^{\mathbb{Q}}$ is then however depending on the ex-ante determined coefficient set and it is therefore probably biased. This bias is, however, still going to be corrected:

spreads from contracts with other maturities (i.e. in the present case 1,3,7 and 10 years) are taken and the sum of squared distances between these observed spreads $SP_{s_{0_i}}(M)$ and the model spreads $\widehat{SP}_{s_{0_i}}(M)$ based on the time series of intensities estimated in our first step is minimized by choosing a new set of coefficients. The new set of coefficients is subsequently used for estimating a times series $\widehat{\lambda}_{s_{0_i}}^{\mathbb{Q}}$ which is again based on the time series of $SP_{s_{0_i}}(5)$. The estimated time series $\widehat{\lambda}_{s_{0_i}}^{\mathbb{Q}}$ is in turn used for the estimation of a new coefficient set by comparing model spreads $\widehat{SP}_{s_{0_i}}(M)$ with the actual spreads $SP_{s_{0_i}}(M)$ for $M \in [1, 3, 7, 10]$. Both steps are afterwards repeated until the estimates of the coefficients and the intensities converge. The final estimates of the intensity time series and the set of coefficients is then characterized by approximately equating the pricing formula for all maturities and all points in time.

4.2 Estimation under $\widehat{\mathbb{P}}$

The coefficients under the measure $\widehat{\mathbb{P}}$ can then be estimated based on the previously estimated intensity time series. In this context, it can be exploited that the transition distribution the CIR process is known in closed form. For this study, the average of the intensities has been chosen as non-parametric estimate for μ_0/μ_1. This is reasonable as μ_0/μ_1 complies with the mean reversion level of the particular CIR process. μ_1 is estimated afterwards via maximum-likelihood estimation (MLE) based on the previously obtained estimate for μ_0/μ_1.

4.3 Estimation Results

Table 1 presents the average model errors for all maturities respectively the "mean in relative difference" between model spreads and observed spreads. The low values for the 5-year maturity are caused by the estimation strategy. For Iceland, Ireland, and Finland the relative model error is modest (17% being the highest) for all maturities. In the Estonian and Polish cases, the errors are in a modest range for all maturities except 1-year. The fit for spreads with respect to maturities being higher than the 1-year case is only in the Spanish case rather disappointing.

It is remarkable that the results for the 1-year case are rather bad in three among six cases. In the Estonian case, the model even completely fails to replicate the 1-year spread. Summing up one can say that the model has a quite satisfactory fit for the 3-, 5-, 7- and 10-year maturities. Spain is the only country with rather disappointing average relative errors (more than 25%) for these maturities [3]. The model does, however, not work very well for the 1-year maturity in three cases. The relative error is finally in all six cases particularly small for the 5-year maturity[4]. The standard deviation of the model errors is moreover rather small. This indicates that the model spreads either systematically exceed the true spreads or that they are systematically below them, instead of fluctuating around them. This could again indicate that the model has difficulties to replicate the term structure of CDS spreads. The overall fit is however, as said before, satisfying.

[3] A reason, why the model fit is rather bad in the Spanish case compared to the other countries rather bad has not been detected. It may, however, be a sign for a structural break. The detection of such breaks is a topic for further research.

[4] This must, however, be interpreted cautiously as the intensities have been estimated based on this maturity.

The estimation results for all countries can be found in table 2. The number of iterations refers to the number of times the model had to be estimated until both intensities and coefficients converged. The estimated loss rates differ strongly from 0.75 which the typical assumption in the literature if the loss rate is not estimated itself. This result supports the suggestion by Pan and Singleton (2008) to estimate the loss rate within the model framework. The values of the objective functions for loss rates beyond one or below zero suggest however that the estimation results are the actual optimal in this model context.

The estimates of $\mu_1^{\widehat{\mathbb{Q}}}$ strongly differ from the estimates of $\mu_1^{\widehat{\mathbb{P}}}$: the estimated system is in all six cases mean reverting under $\widehat{\mathbb{P}}$ but it is only non-explosive under $\widehat{\mathbb{Q}}$ for Ireland. The estimate of $\mu_1^{\widehat{\mathbb{P}}}$ is in the latter case still higher than its counterpart under $\widehat{\mathbb{Q}}$. Moreover, $\frac{\mu_0^{\widehat{\mathbb{P}}}}{\mu_1^{\widehat{\mathbb{P}}}}$ is higher than $\frac{\mu_0^{\widehat{\mathbb{Q}}}}{\mu_1^{\widehat{\mathbb{Q}}}}$ in all cases besides the Irish one. For longer horizons, the values of the intensity which are expected under $\widehat{\mathbb{Q}}$ are accordingly higher than the ones expected under $\widehat{\mathbb{P}}$.
The coefficient estimates ρ_0 and ρ_1 (implied by the estimates for the CIR coefficients) can be found in table 2 as well. The estimate for ρ_0 is in some cases negative and in some cases positive, whereas the estimate for ρ_1 is always positive. Both coefficients being positive implies positive "market prices" of risk η_s respectively a positive change in deterministic drift for a change from measure $\widehat{\mathbb{P}}$ to $\widehat{\mathbb{Q}}$ for all values of $\lambda_s^{\mathbb{Q}}$ $(\sigma_1\sqrt{\lambda_s^{\mathbb{Q}}}\eta_s)$. In opposition to that, the market price of risk can be negative for small values of $\lambda_s^{\mathbb{Q}}$ when ρ_1 is negative. The market price average, which is calculated based on the whole sample period, respectively the average of the difference in the deterministic drift, which is calculated based on the whole sample period is positive in all six cases. This result is also reflected by the whole sample average of the difference in conditional expectations for 1-day and 1-year horizons under both measures.
In all cases, the average conditional expectations are cases higher under $\widehat{\mathbb{Q}}$ than under $\widehat{\mathbb{P}}$. Figure 2 plots the expected Spanish risk neutral intensities conditioned on the estimated current realization for the one year horizon. For each date, the expectations under $\widehat{\mathbb{Q}}$ are higher than the ones under $\widehat{\mathbb{P}}$.

The true model spreads $\widehat{SP}_{s_{0_i}}(M)$ are on average significantly higher than model spreads $\widehat{SP}^{\widehat{\mathbb{P}}}_{s_{0_i}}(M)$ with the expectations calculated based on $\widehat{\mathbb{P}}$. $\widehat{SP}^{\widehat{\mathbb{P}}}_{s_{0_i}}(M)$ is the hypothetical insurance model price, which would be valid as actual model price, if the uncertainty regarding future default probabilities had no impact on expected returns. In the following, this figure is denoted by "hypothetical model spread". Table 2 contains the based on the complete sample averaged values of the relative difference of the latter figure and the actual model spread. The average values of this figure are around 0.9 for four of six cases. The only country with a rather modest averaged relative deviation of the wrong model spreads from the true model spreads is Ireland. Ireland is also the only country for which the hypothetical model spread is at some dates smaller than the actual model spread. These results suggest accordingly that the second dimension risk premium has been positive for the other five sovereigns during the complete sample period. Figures 3 and 4 show the actual and the hypothetical 5-year model spreads for the Irish and the Polish case. Figures 5 and 6 show the relative difference between the actual and the hypothetical 5-year model spreads for the Irish and the Spanish case. Summing up the results referring to the complete sample, one can say, that the "second dimension" risk premium seems to be a very important driver of the included countries' CDS spreads. Based on these results, it can, however, not be concluded that the second dimension risk seems to be particularly important in the European currency union: for Poland and Iceland – i.e. the two non-member countries – the second dimension risk premium seems to be important as well.

Table 2 moreover includes results for the averaged difference in the deterministic drift $\sigma_1\sqrt{\lambda_s^{\mathbb{Q}}}\eta_s$ for two sub-samples. The sample is divided by the last day of November 2009. This was the day when significant corrections of Greek fiscal data were announced (c.f. De Santis (2012): in November 2009, "the Greek government revealed a revised budget deficit of -12.7% of GDP for 2009, which was double the previous estimate"). The results can be subdivided into three

cases: for Ireland, Spain and Finland, the average difference changed from being negative to being positive, for Iceland and Estonia, the opposite is the case and both values are positive but decreasing for Poland. This reflects the fact that the Spanish and Irish spreads are on average significantly higher in the second sub-sample compared to the values in the first sub-sample, whereas the opposite is the case for Iceland and Estonia.
The strongest relative change in averaged $\sigma_1\sqrt{\lambda_s^{\mathbb{Q}}}\eta_s$ is detected for Ireland and Spain, the strongest absolute change occurs for Spain, Ireland and Poland. This suggests that the changes of spreads, which led to the Spanish and Irish crisis, were strongly induced by changes in the market price of risk. This supports the hypothesis that the contagion from Greek to Spain and Ireland have indeed catalyzed by the second dimension risk premium. This may also explain the strong increase in the relative difference between actual and hypothetical spreads for these two countries (shown in figures 5 and 6), as well as the rather high standard deviations of the relative differences. The estimate for the latter can be found in table 2. In opposition to the Irish and Spanish cases, changes in Icelandic and Estonian spreads may have rather been driven by other factors, namely problems in the Icelandic banking sector and actual fiscal difficulties in Estonia.

In addition, correlations between 5-year spreads for all countries as well as correlations between all countries' $\sigma_1\sqrt{\lambda_s^{\mathbb{Q}}}\eta_s$ values are presented in tables 3 and 5. The correlations of the Euro sovereigns' spreads are not always positive. For example, the correlations between Irish and Estonian spreads are distinctly negative. The correlations between Finland and the non-Euro country Poland or between Estonia and non-Euro country Iceland are in opposition to that the highest positive ones. Two further pairs which show a distinct positive correlation are Estonia and Poland as well as Spain and Finland. These results do not suggest that membership in the Euro currency area leads to stronger correlations between spreads per se and comply with the correlations between the changes in drift $\sigma_1\sqrt{\lambda_s^{\mathbb{Q}}}\eta_s$.

The correlations of both figures have been calculated for both sub-samples. The difference between the respective correlations can be found in tables 4 and 6. The differences show that correlations between both spreads as well as the changes in drift decreased in all but two cases between the first and the second period. Only both figures' correlations between Ireland and Finland respectively Poland increased slightly. The strongest decreases in both figures' correlations were found for non-Euro country Iceland. The correlations between Spain and Ireland also decreased, but not as significantly as for country pairs including Iceland. The difference in correlation between changes in drift for the Spain and Ireland is particularly low.

The results for the change in spread correlations contradict the hypothesis that the outbreak of the Greek crisis lead to higher correlations between other European sovereigns' credit costs. The results regarding the change in the market price of second dimension risk contradict the hypothesis that the corrections of Greek fiscal balances lead to a stronger relation between the uncertainties regarding other European sovereigns' future default probabilities. For example, the correlation in Spanish and Irish changes in drift decreased slightly.

Figures 7 and 8 show the correlations between the Spanish and Irish $\sigma_1\sqrt{\lambda_s^{\mathbb{Q}}}\eta_s$ values for a rolling window with widths of 40 respectively 100 days. These plots do also not support the hypothesis of changes in correlations between sovereigns' second dimension risk premiums due to the Greek fiscal information correction. It is instead remarkable that these correlations vary strongly over time and that there is no stable linear dependency between these two countries' market prices of second dimension risk.

Moreover, the spread values $SP_s(5)$ are associated with data for the CBOE volatility index "VIX", measuring implied volatility for the S&P 500 stock index. The VIX index is often used as an approximation for global financial market "nervousness". Table 7 simply contains the adjusted R^2 values for the regressions of the 5 year CDS spread on the VIX index VIX_s:

$$SP_s(5) = \beta_0 + \beta_1 VIX_s + \epsilon_s^{SP,VIX}. \tag{12}$$

The adjusted R^2 value for Iceland decreases strongly from the first part of the sample to the second. In other words, the linear relation between the global financial market nervousness indicator and the spreads has been significantly stronger during the times of distress. This result seems to reflect that the fiscal crisis in Iceland has mainly been induced by problems of Icelandic banks. The adjusted R^2 values for Ireland and Spain are rather modest for both sub-samples compared to the Icelandic value for the first sample part, suggesting a relatively weak linear relation between the VIX index and the respective market price of risk. The change in this value from the first to the second sample is, moreover, relatively small. In combination with the finding that the average difference in drift changes more strongly between the two sub-samples for these two countries, this suggests that the global financial market nervousness may not have been a very important factor for the increases in Spanish and Irish spreads. These increases rather seem to be induced by an increase in the market price of second dimension risk. Moreover, the residuals from the regression of the difference of change in drift $\sigma_1\sqrt{\lambda_s^{\mathbb{Q}}}\eta_s$ on the VIX index are calculated:

$$\sigma_1\sqrt{\lambda_s^{\mathbb{Q}}}\eta_s = \beta_0 + \beta_1 VIX_s + \epsilon_s^{\sigma_1(\rho_0+\rho_1\lambda_s^{\mathbb{Q}}),VIX} \tag{13}$$

The adjusted R^2 values for that regression are presented in table 8. The values are similar to table 7. The residuals' correlations for the whole sample, respectively the difference in correlations between both sub-samples, are displayed in tables 9 and 10.

This correlation of the Irish and Spanish change in drift is still high after filtering the variation, which can be linearly explained by the VIX index. This suggests that the correlation between the market prices of risk is not induced by the simultaneous impact of the general global financial market nervousness. The correlation induced by changes in the market price of risk might instead be induced by simultaneous changes in the actual uncertainty regarding default intensities.

Figure 9 shows the correlations between the Spanish and Irish residuals for a rolling window with widths of 40 days. These graphs do also not support the hypothesis that the Greek fiscal information correction has lead to changes in the linear dependency of market participants' second dimension risk perception for all other European sovereigns after the impact of global market nervousness is filtered out. It is, however, eye catching that the variation of these correlations is much weaker than the variation of the correlations between changes in drift, which are plotted in figures 7 and 8. This suggests that there might be – independently from the Greek fiscal crisis – a stably strong linear dependency between the actual perception of these two countries' second dimension risk.

5 Concluding remarks

This article analyzes the relevance of the "second dimension" risk premium in the context of the European fiscal crisis. It is argued that second dimension risk may have been a crucial aspect for sovereign credit spreads in the context of this crisis and a reduced form credit risk model has been estimated to analyze the relevance of the second dimension risk premium in this context. The empirical results suggest that the second dimension risk premium is indeed an important driver for the credit spreads of the included Euro countries – this is however also the case for the countries, which are not members of the Euro currency area and are included in the sample. The results support moreover the hypothesis that – compared to the credit cost variations during the Icelandic and Estonian crises –the increase of the credit spreads of Spain and Ireland after the beginnings of the Greek crisis has been rather induced by the second dimension risk premium. A strong increase in the average market price of risk after the corrections of the Greek fiscal balances in both the Spanish and the Irish case suggests that the second dimension risk premium might have been in opposition to the other country pairs contagion catalysing for these two particularly troubled countries. The linear dependency between the uncertainty regarding both sovereigns' future default probability seems, moreover, to be strong. The empirical results do not support the hypothesis that the second dimension risk premium induced contagion among Euro countries in general or that the Greek fiscal balance corrections lead to stronger correlation among other European sovereigns' second dimension risk premia.

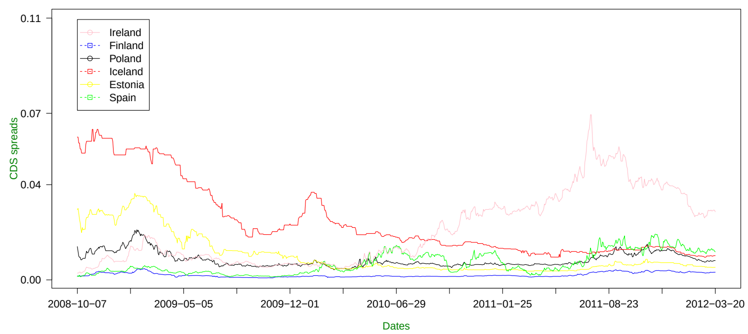

Figure 1: CDS spreads for all sovereigns

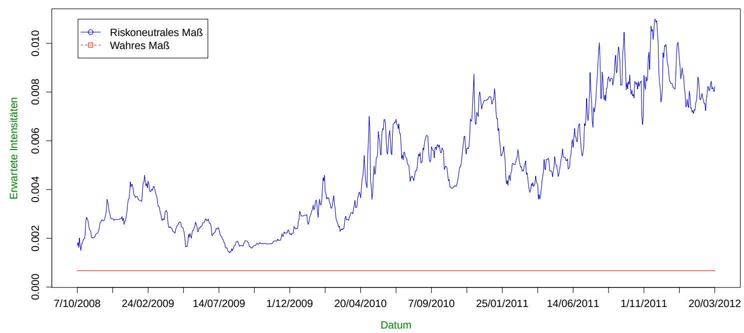

Figure 2: Expected intensities, conditioned on the actual estimated intensity realizations, Spain, horizon: 360 days

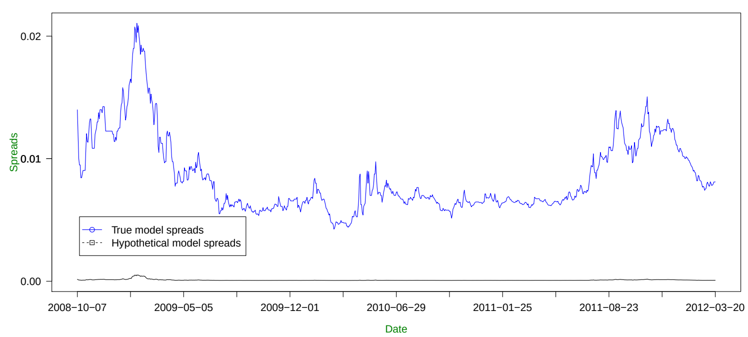

Figure 3: Actual and hypothetical 5-year model spreads, Poland

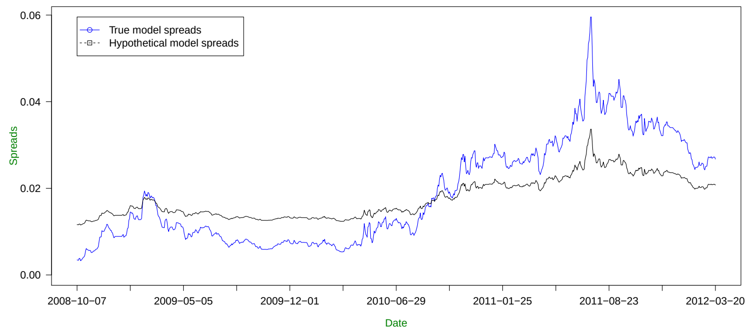

Figure 4: Actual and hypothetical 5-year model spreads, Ireland

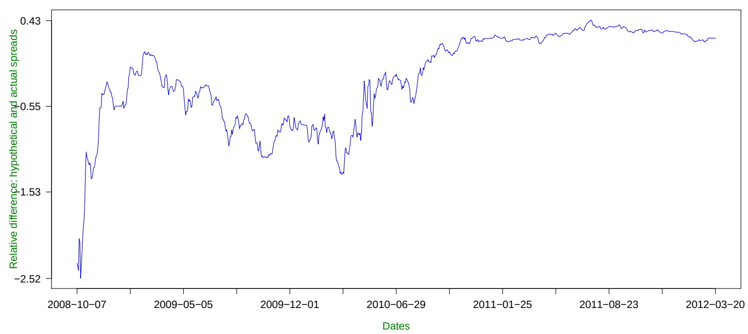

Figure 5: Relative difference in actual and hypothetical model spreads, Ireland

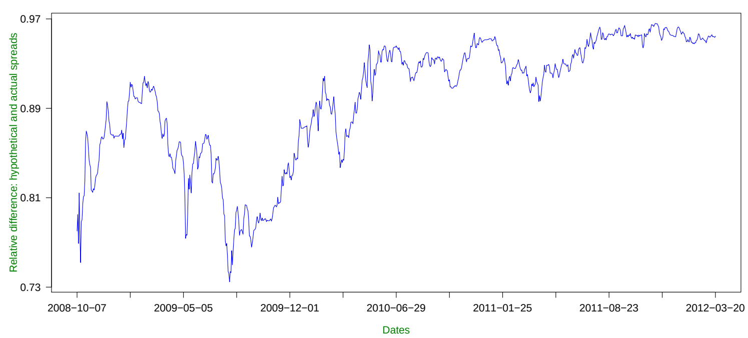

Figure 6: Relative difference in actual and hypothetical model spreads, Spain

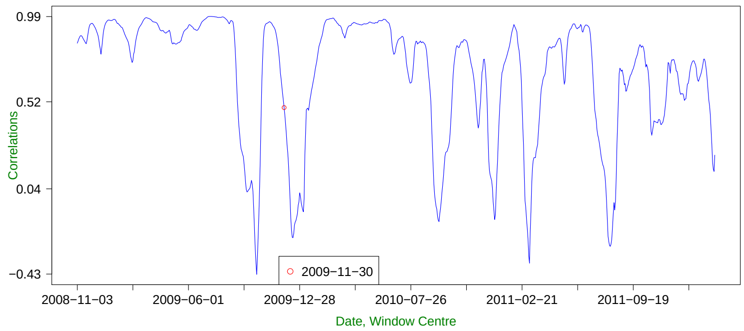

Figure 7: Correlations of Irish and Spanish changes in drift, rolling window, width: 40 days

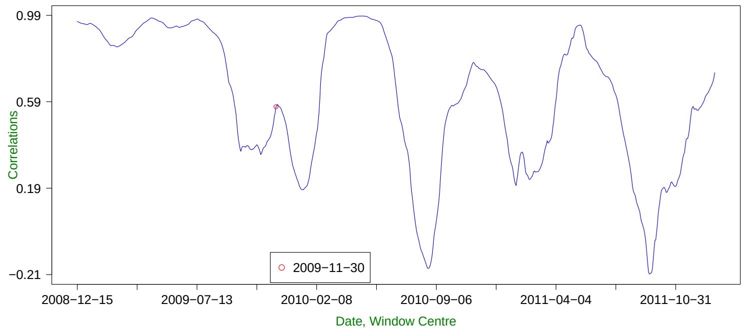

Figure 8: Correlations of Irish and Spanish changes in drift, rolling window, width: 100 days

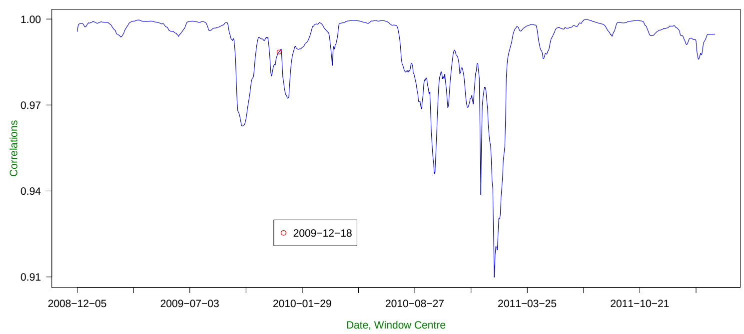

Figure 9: Correlations of Irish and Spanish residuals, regression: change of drift on VIX (equation 13, rolling window, width: 40 days

	1Y	3Y	5Y	7Y	10Y
Finland					
Mean in rel diff.	0.12	-0.03	-0.69×10^{-16}	0.07	-0.06
St.dev. difference	4.82×10^{-4}	2.86×10^{-4}	4.35×10^{-19}	1.53×10^{-4}	2.3×10^{-4}
Mean in difference	-2.28×10^{-4}	-1.29×10^{-4}	-1.46×10^{-19}	1.6×10^{-4}	-1.48×10^{-4}
Iceland					
Mean in rel diff.	-0.098	0.11	-1.1×10^{-17}	-0.07	-0.13
St.dev. difference	5.42×10^{-3}	3.32×10^{-3}	4.31×10^{-18}	2.72×10^{-3}	2.18×10^{-3}
Mean in difference	0.83×10^{-3}	0.54×10^{-3}	-1.93×10^{-19}	-3.17×10^{-4}	-1.61×10^{-3}
Poland					
Mean in rel diff.	-0.39	0.25	-0.99×10^{-17}	-0.09	-0.06
St.dev. difference	1.25×10^{-3}	0.71×10^{-3}	1.16×10^{-18}	0.92×10^{-3}	1.95×10^{-3}
Mean in difference	-1.65×10^{-3}	1.46×10^{-3}	-4.81×10^{-20}	-0.76×10^{-3}	-0.67×10^{-3}
—- Estonia					
Mean in rel diff.	-0.97	0.078	-1×10^{-16}	-0.21	-0.22
St.dev. difference	0.01	1.28×10^{-3}	2.31×10^{-6}	2.36×10^{-3}	1.3×10^{-3}
Mean in difference	-0.009	-1.75×10^{-4}	-1.05×10^{-18}	-2.56×10^{-5}	-1.54×10^{-3}
Spain					
Mean in rel diff.	-0.66	0.36	-3.46×10^{-17}	-0.25	-0.43
St.dev. difference	2.60×10^{-3}	0.95×10^{-3}	1.58×10^{-18}	0.96×10^{-3}	1.6×10^{-3}
Mean in difference	-1.34×10^{-3}	2.53×10^{-3}	-2.96×10^{-19}	-2.13×10^{-3}	-3.66×10^{-3}
Ireland					
Mean in rel diff.	-0.04	-0.05	-4.35×10^{-19}	0.03	0.06
St.dev. difference	2.28×10^{-3}	2.28×10^{-3}	3.2×10^{-18}	4.27×10^{-4}	0.94×10^{-3}
Mean in difference	-1.1×10^{-3}	-1.55×10^{-3}	-1.24×10^{-19}	3.23×10^{-4}	0.52×10^{-3}

Table 1: Model Errors - i.e. the average value of $\frac{1}{N}\sum_{i\in[1,..,N]}\frac{\widehat{SP}_{s_{0_i}}(M)-SP_{s_{0_i}}(M)}{SP_{s_{0_i}}(M)}$

Country	Spain	Ireland	Iceland
$\mu_0^{\hat{\mathcal{Q}}}$	-1.3×10^{-12}	1.46×10^{-3}	$-3.32\times^{-17}$
$\mu_1^{\hat{\mathcal{Q}}}$	-1.97	3.6×10^{-3}	-6.28
$\mu_0^{\hat{\mathcal{P}}}$	0.032	0.01	$5.87e^{-8}$
$\mu_1^{\hat{\mathcal{P}}}$	47.34	0.57	50.87
σ	0.25	0.21	0.00085
LR	1	1	0.99
ρ_0	-0.13	-0.04	-0.55×10^{-3}
ρ_1	198	2.68	53356
Avg. η_s	0.21	0.05	0.63
Avg. diff. in drift	1.32×10^{-3}	1.4×10^{-3}	0.73×10^{-7}
Pre 11/2009 avg. diff. in drift	-1.47×10^{-2}	-4.72×10^{-3}	0.54×10^{-6}
Post 11/2009 avg. diff. in drift	9.33×10^{-3}	4.45×10^{-3}	-1.62×10^{-7}
Avg. diff. in cond. expec. (1D)	$5.47e^{-3}$	$2.2e^{-4}$	$1.73e^{-3}$
Avg. diff. in cond. expec. (1Y)	0.86	0.07	0.98
Avg. rel. diff. in model spreads[5]	0.92	0.06	0.98
St. Dev. Avg. rel. diff. in model spreads	0.06	0.53	3×10^{-7}
Iterations	48	41	185

Country	Finland	Poland	Estonia
$\mu_0^{\hat{\mathcal{Q}}}$	-2.64×10^{-12}	-1.9×10^{-13}	$-7.52\times^{-15}$
$\mu_1^{\hat{\mathcal{Q}}}$	-0.48	-5.35	-5.35
$\mu_0^{\hat{\mathcal{P}}}$	0.015	0.0048	4.62×10^{-6}
$\mu_1^{\hat{\mathcal{P}}}$	20.98	0.42	6.79
σ	0.17	0.13	$3.03e^{-3}$
LR	0.99	0.03	0.91
ρ_0	-0.087	-0.04	-1.48×10^{-3}
ρ_1	126	45.32	3880
Avg. η_s	0.076	4.5	1.41
Avg. diff. in drift	3.43×10^{-4}	0.06	3.63×10^{-6}
Pre 11/2009 mean diff. in drift	-1.14×10^{-4}	0.1	1.34×10^{-5}
Post 11/2009 mean diff. in drift	5.71×10^{-4}	0.04	-1.26×10^{-6}
Avg. rel. diff. in cond. expec. (1D)	$1.34e^{-3}$	$1.48e^{-2}$	$1.47e^{-2}$
Avg. rel. diff. in cond. expec. (1Y)	0.38	0.98	0.97
Avg. rel. diff. in model spreads	0.65	0.97	0.9
St. Dev. Avg. rel. diff. in model spreads	0.16	0.002	4.5×10^{-6}
Iterations	18	201	31

Table 2: Estimation results under both measures

Notation:

- Avg. η_s: Refers to the average value for η_s over the complete sample.

- Avg. diff. in drift: Average of $\sigma_1\sqrt{\lambda^{\mathbb{Q}}_{s_{0_i}}}\eta_s$. This refers to the difference in the deterministic drift under $\widehat{\mathbb{P}}$ compared to $\widehat{\mathbb{Q}}$, i.e. a negative value characterizes a higher (i.e. more positive) deterministic drift under $\widehat{\mathbb{Q}}$.

- Avg. rel. diff. in cond. exp. refers to the average relative difference in expectations of the intensity conditioned on the respective current value with a one day (1D) respectively (1Y) horizon (i.e. $\frac{\mathbb{E}^{\widehat{\mathbb{Q}}}_{s_{0_i},\widehat{\mu}^{\widehat{\mathbb{Q}}}_0,\widehat{\mu}^{\widehat{\mathbb{Q}}}_1,\widehat{\sigma}_1}[\lambda^{\mathbb{Q}}_{s_{0_i}+1/360}]-\mathbb{E}^{\widehat{\mathbb{P}}}_{s_{0_i},\widehat{\mu}^{\widehat{\mathbb{P}}}_0,\widehat{\mu}^{\widehat{\mathbb{P}}}_1,\widehat{\sigma}_1}[\lambda^{\mathbb{Q}}_{s_{0_i}+1/360}]}{\mathbb{E}^{\widehat{\mathbb{P}}}_{s_{0_i},\widehat{\mu}^{\widehat{\mathbb{P}}}_0,\widehat{\mu}^{\widehat{\mathbb{P}}}_1,\widehat{\sigma}_1}[\lambda^{\mathbb{Q}}_{s_{0_i}+1/360}]}$ respectively $\frac{\mathbb{E}^{\widehat{\mathbb{Q}}}_{s_{0_i},\widehat{\mu}^{\widehat{\mathbb{Q}}}_0,\widehat{\mu}^{\widehat{\mathbb{Q}}}_1,\widehat{\sigma}_1}[\lambda^{\mathbb{Q}}_{s_{0_i}+1}]-\mathbb{E}^{\widehat{\mathbb{P}}}_{s_{0_i},\widehat{\mu}^{\widehat{\mathbb{P}}}_0,\widehat{\mu}^{\widehat{\mathbb{P}}}_1,\widehat{\sigma}_1}[\lambda^{\mathbb{Q}}_{s_{0_i}+1}]}{\mathbb{E}^{\widehat{\mathbb{P}}}_{s_{0_i},\widehat{\mu}^{\widehat{\mathbb{P}}}_0,\widehat{\mu}^{\widehat{\mathbb{P}}}_1,\widehat{\sigma}_1}[\lambda^{\mathbb{Q}}_{s_{0_i}+1}]}$).

- Rel. diff. in model spreads refers to the relative deviation of the 5-year model spread with expectations calculated based on $\widehat{\mathbb{Q}}$ (i.e.$\widehat{SP}_{s_{0_i}}(5)$) from the 5-year model spread with expectations calculated based on $\widehat{\mathbb{P}}$. This means: $\frac{average(\widehat{SP}_{s_{0_i}}(5))-average(\widehat{SP}^{\widehat{\mathbb{P}}}_{s_{0_i}}(M))}{average(\widehat{SP}_{s_{0_i}}(5))}$
with $\widehat{SP}^{\widehat{\mathbb{P}}}_{s_{0_i}}(M) == \frac{\widehat{LR}\left[\int_{s_{0_i}}^{s_{0_i}+M} ZB^f_{s_{0_i},s}\mathbb{E}^{\widehat{\mathbb{P}}}_{s_{0_i},\widehat{\mu}^{\widehat{\mathbb{P}}}_0,\widehat{\mu}^{\widehat{\mathbb{P}}}_1,\widehat{\sigma}_1}\left[\widehat{\lambda}^{\mathbb{Q}}_s e^{-\int_{s_{0_i}}^{s}\widehat{\lambda}^{\mathbb{Q}}_u du}|\mathcal{F}_{2,s_{0_i}}\right]ds\right]}{\sum_{n=1}^{2M}\left(\mathbb{E}^{\widehat{\mathbb{P}}}_{s_{0_i},\widehat{\mu}^{\widehat{\mathbb{P}}}_0,\widehat{\mu}^{\widehat{\mathbb{P}}}_1,\widehat{\sigma}_1}\left[e^{-\int_{s_{0_i}}^{s_{0_i}+0.5n}\widehat{\lambda}^{\mathbb{Q}}_s ds}|\mathcal{F}_{2,s_{0_i}}\right]ZB^f_{s_{0_i},s_{0_i}+0.5n}\right)}$

	Ireland	Finland	Poland	Iceland	Estonia	Spain
Ireland	1	0.55	0.18	-0.59	-0.42	0.68
Finland	0.55	1	0.83	0.13	0.32	0.58
Poland	0.18	0.83	1	0.54	0.72	0.24
Iceland	-0.59	0.13	0.54	1	0.93	-0.49
Estonia	-0.42	0.32	0.72	0.93	1	-0.39
Spain	0.68	0.58	0.24	-0.49	-0.39	1

Table 3: Correlations of spreads

	Ireland	Finland	Poland	Iceland	Estonia	Spain
Ireland	0	-0.11	-0.04	1.13	0.57	0.38
Finland	-0.11	0	0.04	1.23	0.43	0.05
Poland	-0.04	0.04	0	1.14	0.35	0.07
Iceland	1.13	1.23	1.14	0	0.51	1.06
Estonia	0.57	0.43	0.35	0.51	0	0.45
Spain	0.38	0.05	0.07	1.06	0.45	0

Table 4: Difference in correlations of spreads pre 11/2009 vs post 11/2009

	Ireland	Finland	Poland	Iceland	Estonia	Spain
Ireland	1	0.54	0.04	-0.58	-0.33	0.84
Finland	0.54	1	0.65	0.14	0.39	0.59
Poland	0.04	0.65	1	0.49	0.8	0.05
Iceland	-0.58	0.14	0.49	1	0.87	-0.59
Estonia	-0.33	0.39	0.8	0.87	1	-0.37
Spain	0.84	0.59	0.05	-0.59	-0.37	1

Table 5: Correlations $\sigma_1\sqrt{\lambda_s^{\mathbb{Q}}}\eta_s$

	Ireland	Finland	Poland	Iceland	Estonia	Spain
Ireland	0	-0.07	-0.01	1.17	0.68	0.13
Finland	-0.07	0	-0.06	1.22	0.45	-0.01
Poland	-0.01	-0.06	0	0.81	0.3	0.01
Iceland	1.17	1.22	0.81	0	0.39	1.25
Estonia	0.68	0.45	0.3	0.39	0	0.64
Spain	0.13	-0.01	0.01	1.25	0.64	0

Table 6: Difference in correlations of $\sigma_1\sqrt{\lambda_s^{\mathbb{Q}}}\eta_s$ pre 11/2009 vs post 11/2009

	complete sample	first sample	snd. sample
Ireland	0.04	-0.78×10^{-3}	0.08
Finland	0.16	0.4	0.19
Poland	0.49	0.49	0.33
Iceland	0.56	0.7	-1.64×10^{-3}
Estonia	0.62	0.63	0.18
Spain	0.01	0.15	0.21

Table 7: adjusted R^2 regression 12

	complete sample	first sample	snd. sample
Ireland	0.04	1.6×10^{-3}	0.06
Finland	0.16	0.4	0.19
Poland	0.3	0.22	0.29
Iceland	0.56	0.71	-1.66×10^{-3}
Estonia	0.56	0.5	0.17
Spain	0.05	0.17	0.12

Table 8: adjusted R^2 regression 13

	Ireland	Finland	Poland	Iceland	Estonia	Spain
Ireland	1	0.7	0.19	-0.65	-0.27	0.83
Finland	0.7	1	0.56	-0.28	0.15	0.75
Poland	0.19	0.56	1	0.15	0.71	0.2
Iceland	-0.65	-0.28	0.15	1	0.69	-0.66
Estonia	-0.27	0.15	0.71	0.69	1	-0.33
Spain	0.83	0.75	0.2	-0.66	-0.33	1

Table 9: Correlations $\epsilon_s^{\sigma_1(\rho_0+\rho_1\lambda_s^{\mathbb{Q}}),VIX}$

	Ireland	Finland	Poland	Iceland	Estonia	Spain
Ireland	0	0.01	0.13	1.14	0.95	0.19
Finland	0.01	0	0.06	1.04	0.99	-0.15
Poland	0.13	0.06	0	0.2	0.46	0.05
Iceland	1.14	1.04	0.2	0	-0.31	1.01
Estonia	0.95	0.99	0.46	-0.31	0	0.94
Spain	0.19	-0.15	0.05	1.01	0.94	0

Table 10: Correlations $\epsilon_s^{\sigma_1(\rho_0+\rho_1\lambda_s^{\mathbb{Q}}),VIX}$ Period 1 - Period 2

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