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ABSTRACT

Animal Spirits in the Euro Area Sovereign CDS Market

We study the determinants for the sovereign credit default swap (CDS) spreads of five Euro-area countries (Greece, Ireland, Italy, Portugal, Spain) in the post-Lehman-Brothers period. We find that there are regime switches in the process of sovereign CDS spread changes. We consider three alternative empirical hypotheses associated with regime switches. Under the first hypothesis, there are rational sunspot equilibria. Under the second hypothesis, there is a unique fundamental equilibrium and the regime switching is caused by changes in policy makers' preferences. The third hypothesis relaxes the rational expectations assumption. Under this hypothesis, indicators of the market fundamentals are not always precise. They are better indicators if cognitive biases are small and the rational expectations economy is a good approximation for reality. However, if market uncertainties enlarge the cognitive biases, the market-based indicators of fundamentals are no longer precise. In this case, they are not useful for CDS pricing. We find that the first two hypotheses are difficult to fit the data and the third hypothesis provides a good explanation for the sovereign CDS spread dynamics in our sample.

JEL Classification: F34, G15 and P34
Keywords: animal spirit, euro and sovereign cds

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1 Introduction

During the European sovereign debt crisis, sovereign credit default swap (CDS) spreads of the Euro countries drew a lot of public attention. The reason is that a country's CDS spread is usually taken as an indicator of that country's sovereign credit risk.[1] In this paper, we test the reliability of the sovereign CDS spread as an indicator of the sovereign credit risk. More specifically, we test whether changes in variables related to the sovereign credit risk are significant determinants for changes in the sovereign CDS spreads of five Euro-area countries (Greece, Ireland, Italy, Portugal and Spain) in the post-Lehman-Brothers period (from September 15, 2008 to December 19, 2011).

There are a number of empirical studies on the determinants of sovereign CDS spreads in developed countries. Longstaff et al. (2011) find that global financial market conditions significantly affect sovereign CDS spreads of 26 countries, including both developing and developed countries such as Japan and Korea. Dieckmann and Plank (2011) extend their analysis to Western European countries and find that global financial factors also play significant roles there. Moreover, they report that changes in the performance of the financial industry affect changes in the CDS spreads of Western European sovereigns. This finding is consistent with a private-to-public risk transfer hypothesis: prospective government debt necessary to help the distressed financial industry may increase a country's sovereign credit risk. Fontana and Scheicher (2010) focus on Euro area countries and also find changes in sovereign CDS spreads are related to global factors. They find that measures of investors' changing risk appetite play a prominent role in the sovereign CDS pricing.

All those previous empirical studies share two features. First, the empirical models are linear. More specifically, there is no regime switching in the models. Second, the covariates are assumed to be exogenous. In this paper, we show that those two features can bias the statistical inference.

Regime switching can arise from three different theories. Two of those theories are related to different concepts of "animal spirits". In the rational expectations framework, the animal spirits (henceforth we shall call this concept of animal spirits "animal spirits 1") are interpreted as sunspot shocks[2] to investors' expectations (Farmer, 2008). Those

[1] In this paper, we define the sovereign credit risk by the default probability of the sovereign bonds and the associated recovery rate after default.

[2] The sunspot shocks are defined as psychological changes which are not related to economic funda-

sunspot shocks cause multiple equilibria. The economy will be in a good equilibrium if people believe so while the economy will be in a bad equilibrium if people believe it to be bad. Such sunspot-driven multiple equilibria have been used to explain different economic phenomena. They are used to explain excessive volatility in macroeconomic variables such as output and inflation (Clarida et al., 2000; Lubik and Schorfheide, 2004; Davig and Leeper, 2007; Farmer et al., 2010). Diamond and Dybvig (1983) use them to explain bank runs. In international finance, they are used to explain self-fulfilling currency crises (Burnside et al., 2008; Jeanne, 2000). Jeanne and Masson (2000) propose an empirical test for the existence of rational sunspot equilibria in the currency crises context. They prove that the effects of the sunspot shocks are absorbed by discrete jumps in the intercept of a regression of the currency devaluation probability on fundamental variables. Therefore, a test for Markov regime switches in the intercept can be taken as a test for the existence of sunspot equilibria. We argue in Section 2 that this test can be applied to the sovereign CDS market *under the rational expectations assumption.*

While the theory of animal spirits 1 predicts regime switches in the intercept of the regression model, an alternative theoretical model under the rational expectations assumption predicts regime switches in the slopes of the regression model. Assuming that investors are rational and there is no sunspot equilibrium, the slopes change if governments change their preferences over different policy objectives. For example, when the financial crisis deepens, the weight attached to financial stability may become larger relative to economic growth in governments' objective functions. Anticipating this, rational investors will change their pricing behavior accordingly. Section 2 shows that this can lead to regime-dependent slope changes in the CDS spread determination equation.

Under the rational expectations assumption, investors are cognitively unlimited. Therefore, changes in market-based indicators of fundamentals always provide reliable information on the development of fundamental variables. Moreover, the information will be correctly incorporated into sovereign CDS spreads. Those results no longer hold if investors are not cognitively limited. When uncertainties overwhelm the market and there are time constraints for decision-making, the investors rely more on beliefs that are not necessarily based on rational calculations. We call those movements in beliefs of the cognitively limited investors "animal spirits 2" since they are different from the sunspot

mentals.

shocks ("animal spirits 1") in the rational expectations framework. The "animal spirits 2" concept is close to the definition of animal spirits in two recent theoretical papers by De Grauwe (2011a, 2012). In those two papers, agents are not fully rational, that is, they are cognitively limited, and use heuristics rather than rational calculations to make decisions. Agents' sentiments are self-fulfilling because they switch from an optimistic forecast rule to a pessimistic forecast rule if more other agents adopt the pessimistic rule. The widespread pessimistic psychology dampens aggregate demand and eventually leads to a bad outcome. De Grauwe (2011a, 2012) formalizes the concept of "confidence multiplier" of Akerlof and Shiller (2009). According to Akerlof and Shiller (2009), confidence is the belief of cognitively limited agents rather than a sunspot shock to the expectations of perfectly rational agents. De Grauwe (2011a, 2012) does not consider the possibility that agents can change their focus variables in their decision rules if market condition changes. Our definition of "animal spirits 2" allows this possibility.[3] Particularly, when market conditions become very uncertain, agents may drop decision rules based on observable fundamental variables. It is important to point out that ignoring the fundamentals does not mean that investors are irrational. It may be a *boundedly*[4] rational choice by the cognitively limited and imperfectly informed investors. It is because movements in the observable fundamental variables are driven by market participants whose cognitive abilities are also limited. The information content of those fundamentals is more seriously distorted by the cognitive biases in a more uncertain market. Therefore, observable fundamental variables which are useful when market uncertainty is low can become useless if market uncertainty becomes very high. The theory of animal spirits 2 is consistent with a two-state regime switching model. Under the regime with low market uncertainty, market-based indicators of fundamental variables have significant explanatory power for changes in sovereign CDS spreads. Under the regime with high market uncertainty, all market-based indicators of fundamentals are insignificant.

Above theoretical possibilities for regime switching motivate a form test for regime switching in regression models. Using the quasi-likelihood ratio test developed by Cho and White (2007), we show that the model linearity assumption in previous studies are

[3]Branch and Evans (2007) introduce a model in which agents can change their focus variable in their decision rules. Those agents are taken as econometricians. As pointed out by De Grauwe (2011a), such agents may have better cognitive skills than agents in the real world.

[4]We use the word "bounded" to suggest that agents are cognitively limited whereas perfect rationality means that agents are cognitively unlimited.

not valid.

Previous empirical studies on the determination of the sovereign CDS spreads assume that the covariates are exogenous. This assumption rules out the possibility that dynamics in the sovereign CDS spreads may affect fundamental variables. Ruling out such a possibility can be a source of bias. Particularly, it is possible that changes in the CDS spreads will feedback to governments' borrowing costs and affect domestic economic fundamentals.[5] Using a two-step estimation technique developed by Kim (2009), we estimate our regime switching model with instrumental variables and formally test for endogeneity based on the estimation results. Our test suggests that the domestic fundamentals are indeed endogenous in four sample countries (Ireland, Italy, Portugal and Spain). Therefore, compared to the previous studies using ordinary least squares (OLS), our results are more reliable; not only because we model the omitted nonlinearities caused by regime switches but also because we correct for reverse causality.

We find that there is no regime switch in the intercept of the regression equations. This indicates a failure of the joint hypothesis of rational expectations and sunspot equilibria. Therefore, the theory of animal spirits 1 is rejected. There are regime switches in the slopes of the regression equations. A possible explanation is that there is a unique rational expectations equilibrium in which the policy focus of the government changes. The difficulty with this explanation is that there is one regime under which we find that the sovereign CDS spreads are white noise. That is, they are completely disconnected from all the fundamental variables. The results are better explained by the theory of animal spirits 2. Observable indicators of fundamentals have little value to investors in the sovereign CDS market due to distortions caused by cognitive biases when market uncertainty is high. They are more valuable and used by investors to price the sovereign CDS contracts when market uncertainty is low.

The rest of the paper is organized as follows: Section 2 elaborates on three empirical hypotheses. Section 3 introduces the explanatory variables and describes the data. Section 4 provides estimates of OLS regression models for the determination of sovereign CDS spreads and tests for regime switching in the models. Section 5 shows estimated regime switching models with instrumental variables and tests for endogeneity. Section 6 concludes.

[5] See OECD (2012).

2 Empirical hypotheses

In the section, we elaborate on three alternative empirical hypothesis for the Euro-area sovereign CDS market.

Hypothesis 1 (animal spirits 1): agents are fully rational and there exist multiple sunspot equilibria.

According to Reinhart and Rogoff (2009), a country's default decision is the result of a cost-benefit analysis. Many countries default on their debts long before they run out of financial resources. Under the rational expectations assumption, Jeanne and Masson (2000) model a country's probability of currency devaluation as a result of its cost-benefit analysis. Due to the similarity, we can apply that model to our sovereign CDS context. More specifically, let us assume that the net benefit function of the government is $B(f_t, d_t)$, where f_t is an index of economic fundamentals, $d_t \equiv \int_0^1 d_t(i)di$ is the average estimate of the probability of default formed by a continuum of investors $i \in [0, 1]$.[6] The net benefit function is increasing in f_t, reflecting the idea that the better the fundamentals are, the higher will be the chance that the government will honor its debt. It is decreasing in d_t, suggesting that it is more costly to honor the debt if the investors have higher estimates for the default probability. More specifically, a higher expected default probability increases the interest rate for sovereign borrowing and induces the investors to divert their investment to safer assets, making rollover more difficult (De Grauwe, 2011b).

Investor i expects that the government will default if the net benefit of honoring its debt becomes negative. Therefore, $d_t(i) = Prob[B(f_{t+1}, d_{t+1}) < 0|f_t]$, where $Prob$ denotes probability. Following Jeanne and Masson (2000), we assume that the investors share common knowledge so that we can drop index i in the formula. Under some additional technical assumptions[7], there is a critical value of the fundamental index below which the government will default, given the market estimate of the default probability. Therefore, we can write the average estimate of the default probability as $d_t = Prob[f_{t+1} < f^{*e}|f_t] \equiv F(f_t, f^{*e})$, where f^* is the critical value defined by the equation $B(f^*, d_t) = 0$, the superscript e denotes expectation. Note that f^* is an implicit function of d_t and d_t is a function of f^{*e} , so f^* is a function of f^{*e}, which we denote by $g(f^{*e})$. Under the rational

[6]The two-step estimation approach of Kim (2009) is designed for time series not for panel data analysis. Therefore, we estimate empirical models separately for each country. That is why we only have the time subscript for variables.

[7]see Jeanne and Masson (2000) for details.

expectations assumption, $f^* = f^{*e}$, so $f^* = g(f^*)$. That is, f^* is a fixed point of the function g. Jeanne and Masson (2000) show that there can be more than one fixed point of g. Their proposition 1 further establishes that if there is more than one fixed point of g, there will be multiple sunspot equilibria. More specifically, there will be n states under which the threshold fundamental index value (denoted by f_s^*, where s is the state index) differs. The probability of default depends not only on the fundamental variables, but also on the transition probabilities from the current to the future states:

$$d_t = \Sigma_{s=1}^{n} q(s_t, s) F(f_t, f_s^*), \tag{1}$$

where $q(s_t, s)$ is the transition probability from the current state to state s in the next period, $F(f_t, f_s^*) \equiv Prob[f_{t+1} < f_s^* | f_t]$, where f_s^* is the critical value of the fundamental index under state s.

Following Jeanne and Masson (2000), we assume that the fundamental index is a linear function of the macroeconomic variables relevant for the policy maker's decision. More specifically, $f_t = \alpha' m_t$, where m_t is a vector of economic fundamentals, α is a vector of constant coefficients and $'$ is a transpose operator. Under this assumption, Jeanne and Masson (2000) show that equation (1) can be linearized to the following form:

$$d_t = \delta_{s_t} + \varphi' m_t, s_t = 1, ..., n, \tag{2}$$

where δ_{s_t} is a coefficient changing with the state, and φ is a vector of constant coefficients.

Under the rational expectation assumption, the sovereign CDS spread is determined by the default probability of the underlying bond (d_t) and other variables, such as the recovery rate of the defaulted bond and the investors' risk appetite. We write the linearized pricing equation for the sovereign CDS as follows:

$$CDS_t = l + \phi d_t + \chi' \mu_t, \tag{3}$$

where CDS_t is the sovereign CDS spread, l and ϕ are constants, χ is a vector of constant coefficients, and μ_t is a vector of determinants for the sovereign CDS spread other than the default probability. Substitute for d_t using equation (2), we get

$$CDS_t = \vartheta_{s_t} + \zeta m_t + \chi' \mu_t, \tag{4}$$

where $\vartheta_{s_t} \equiv l + \phi\delta_{s_t}$, $\zeta \equiv \phi\varphi'$. Equation (4) suggests that under Hypothesis 1, the sovereign CDS spread determination model can be approximated by a Markov regime switching model in which the intercept changes across states but the slopes are always constant.

Hypothesis 2 (changing policy focus): there is a unique *rational* expectations equilibrium, given a particular set of focus fundamental variables of the government. But the focus may change in the sample period.

Hypothesis 2 means that there is only one fixed point for the function g. In this case, $d_t = F(f_t, f^*)$. Its linearized version can be written as

$$d_t = a + bf_t, \tag{5}$$

where a and b are constants.[8] If the fundamental index f_t is a linear function of relevant fundamental variables, we will get a constant coefficient CDS pricing model. However, under hypothesis 2, f_t is not a linear function. Instead,

$$f_t = \alpha'_{j_t} m_t, j_t = 1, ..., J, \tag{6}$$

where α_{j_t} is a vector of coefficients which change with a discrete state variable j_t, J is the number of possible combinations of target fundamental variables in the objective function of the government, and m_t is the collection of all the potentially relevant variables. Equation (6) captures the idea that the set of fundamental variables in the government objective function can change during a turbulent period. More specifically, there is an unobservable latent state variable j whose value governs the changes in the preference of the government. Note that equation (6) not only allows changes in the relative weights of the same set of fundamental variables but also allows the set of relevant fundamental variables to change across states. Changes in the set of relevant fundamental variables can be modeled by setting different elements of α_{j_t} to zero under different states. Combining

[8]Following Jeanne and Masson (2000), we assume that $f_t = \bar{f} + cf_t$, where $\bar{f}$ and c are constants and cf_t is of the first order. Under this assumption, $a = F(\bar{f}, f^*)$ and $b = F_1(\bar{f}, f^*)$.

equations (3), (5), and (6), we get

$$CDS_t = \iota + \kappa_{s_t} m_t + \chi' \mu_t, \tag{7}$$

where $\iota \equiv l + a\phi$ and $\kappa_{s_t} \equiv b\phi\alpha'_{jt}$. Thus, under Hypothesis 2, it is the slope vector rather than the intercept that changes across different states. Note that it is possible that not only κ but also some elements of χ change with the state variable.[9] For example, the recovery rate also depends on the cost-benefit analysis of the defaulting government (Reinhart and Rogoff, 2009). Therefore, our reasoning for regime-dependent parameter changes in κ should also be applicable to the coefficients of determinants for the recovery rate.

Hypothesis 3 (animal spirits 2): agents are only boundedly rational. They rely on beliefs which are not related to the observable fundamentals if market uncertainty is high.

The derivation of equations (4) and (7) depends on the rational expectation assumption. More specifically, perfect rationality plays three important roles. First, it makes the information content of observable fundamental variables reliable to be used for forecasting the default probability of the sovereign bonds. Second, the forecast of the default probability will be unbiased because the investors are perfectly rational. Third, perfect rationality assures that the CDS spread will correctly incorporate all information on the unbiased forecast of the default probability. If agents are not perfectly rational, those three results will no longer be valid. If market uncertainty is low and cognitive biases are small, a CDS spread determination equation based on those three results may still be a good approximation of reality. In this case, the observable fundamental variables will have explanatory power for the dynamics in the sovereign CDS spreads. However, if market uncertainty is high and cognitive biases are large, it is no longer guaranteed that the observable fundamentals will have explanatory power for the dynamics in the sovereign CDS spreads. It is because the information content of the fundamentals can be highly distorted in a very uncertain environment, and there is no reason to use the incorrect information to price the CDS contract. Therefore, Hypothesis 3 is consistent with a two-state regime switching model.[10] In the less uncertain state, fundamental vari-

[9] In this case, we should add a subscript s to χ.

[10] In our empirical models, we restrict the number of states to two to save degrees of freedom. Consider only parameters to estimate in the transition matrix. Increasing the number of states from two to three will increase the number of parameters to estimate from 12 to 72 in the two-step regime switching model.

ables have nonzero coefficients in the CDS spread determination equation. In the more uncertain state, the coefficients of the fundamental variables are zero.

3 Variable and data description

3.1 The dependent variable: The sovereign CDS spread

The dependent variable in our empirical analysis is the sovereign CDS spread. A CDS contract can be taken as an insurance contract against the credit event specified in the contract.[11] Its spread, expressed in basis points, is the insurance premium the protection buyer has to pay. For example, a CDS spread of 20 basis points means the buyer of credit protection has to pay the seller an annual amount equal to 0.2 percent of the notional value of the reference debt obligation.[12] There are different credit events against which a sovereign CDS contract can insure. Following Dieckmann and Plank (2011), we consider only the CDS contracts on the credit event "complete restructuring", since it is the standard credit event in the European sovereign CDS contract. The contract maturity we consider is 10 years because the 10-year contract is the most liquid one for the European market. The spreads are quoted in US dollars, the standard currency for European sovereign CDS contracts. Our sample covers weekly data on 10-year government bond CDS spreads from September 15, 2008 to December 19, 2011. Importantly, our sample covers the period after April 2010, which is not covered in the previous studies surveyed in the introduction. Since sovereign debt problems in the sample countries become even more concerned by the public in this period, this extension is particularly interesting.[13] We start the sample from the collapse of Lehman Brothers since the study by Dieckmann and Plank (2011) suggests that European samples before and after the collapse of Lehman Brothers are very different. We include five Euro-area countries (Greece, Ireland, Italy, Portugal and Spain) into our sample. Those five countries are widely believed to have experienced a debt crisis in our sample period. Therefore, it is interesting to ask how reliable are sovereign CDS spreads of those countries as indicators for their sovereign

Since our sample is relatively small, it is better to restrict the number of states to two. Two is also the typical number of states specified in empirical regime switching models. For example, Jeanne and Masson (2000) use a two-state model.

[11]More precisely, it is a quasi-insurance instrument. See Pan and Singleton (2008) and Dieckmann and Plank (2011) for a more detailed description of the sovereign CDS contract.

[12]In our context, the reference debt is the sovereign bond.

[13]See OECD (2012).

credit risk during the crisis.

3.2 The covariates

Table 1 summarizes the covariates we use in the regression analysis. As we discussed, the probability of a government's default on its debt depends on the costs and benefits of honoring its debt. Thus, rational investors will use variables that can affect the government's cost-benefit analysis to conjecture the probability of a government default. In addition, they will use this probability of default to price the sovereign CDS contract, an insurance for the sovereign credit risk. Hence, we include variables that are commonly perceived to affect the country's willingness to pay its debt as covariates in the regression analysis. Note that we do not impose the rational expectations assumption for the regression. Rather, we take statistical insignificance of the variables that should have explanatory power to changes in the sovereign CDS spreads under the rational expectations assumption as a failure of the assumption.

Theoretically, the state and volatility of the economy may affect a country's willingness to pay its debt. Fiscal reforms necessary to honor the government's debt obligation can impose additional pressure on the already distressed economy. Therefore, when the domestic economy is weak and unstable, the policy maker will be less willing to implement the reforms. Following the literature, we use the domestic stock market return and volatility to proxy the economic state and volatility, respectively. In the rational expectations framework, one should expect the lower the stock market return or the more volatile the return, the higher the sovereign CDS spread, reflecting the unwillingness of the government to take fiscal reforms in an already weak and unstable economy. While Dieckmann and Plank (2011) use the domestic stock price index return, we use the gross return which also includes dividends. This choice is because changes in dividends also contains information on the performance of firms, which affect the performance of the economy. Another domestic variable we consider is the stock market performance of domestic financial firms, the Dow Jones Total Market(DJTM) Financials index. Dieckmann and Plank (2011) argue that this variable measures the private-to-public risk transfer due to the costs of helping the distressed financial industry. That means we should expect a higher sovereign CDS spread when the DJTM financials index is low.

Longstaff et al. (2011) suggest that changes in the global stock and bond markets can explain a large part of the variation in an individual country's sovereign CDS spread. Empirical studies on the European sovereign CDS market (Fontana and Scheicher, 2010; Dieckmann and Plank, 2011) find the same result. For this reason, we also include indicators of developments in the global stock and bond markets as covariates. More specifically, we follow Dieckmann and Plank (2011) to use the EuroStoxx 50 return and MSCI World Financials index as indicators for global stock market developments. We use 10-year German Bund rate and iTraxx Europe corporate CDS spread as indicators for global bond market developments. Dieckmann and Plank (2011) use corporate bond spreads rather than the iTraxx index to proxy European corporate credit spread. The corporate credit spread is not significant in their time series analysis. By contrast, Fontana and Scheicher (2010) find that the iTraxx index has strong explanatory power in the equation for European sovereign CDS spreads. For this reason, we use the iTraxx Europe index as the proxy for European corporate credit spread.

Theoretically, including global variables into the analysis captures the international spillover effect. The European Monetary Union(EMU)-wide stock market performance, EuroStoxx 50 return, is a proxy for the state of the Euro-area economy. Through trade linkages, the economic conditions in the other member countries can affect the home country's economy. This spillover effect need not to be fully captured by the current domestic stock market return due to the fact that a bad union-wide economic condition may affect the home economy with lags. More importantly, in a monetary union, a sovereign country's probability of default is partly affected by the willingness of the other member countries to bail it out, and the other member countries' willingness to pay will depend on their own economic conditions. In this case, a decline in the union-wide economy, proxied by the EuroStoxx 50 return, will increase the sovereign CDS spread. Similarly, a bad state of the world financial industry may affect the willingness of the international community to help an individual sovereign nation out of its debt problem.[14] Therefore, a decline in the World Financials index may increase the home country's sovereign CDS spread. A higher German Bund rate signals a higher rate of economic growth in Germany. This favorable outcome can in turn help improve the economic conditions of the other EMU countries and increase their willingness to help the member countries which have debt problems. Even if Germany's economic growth does not affect other member countries'

[14] We use a worldwide proxy for the performance of the financial sector rather than a Euro-area one because the later is not available.

economic performance, an improvement in its own economy alone can significantly affect the market expectation of defaults by the Euro-area periphery countries. This spillover effect is because Germany plays a leading role in negotiations on the bailout plans. Thus, we expect that an increase in the German Bund rate may reduce the sovereign CDS spreads of the periphery countries. The European corporate CDS spread index, iTraxx, measures the corporate credit spread in Europe. It contains a proxy for the overall state of the European economy since the recovery rates of defaulted corporate bonds increase as the overall business climate improves (Collin-Dufresne et al., 2001). Because lower recovery rates lead to higher corporate CDS spreads, an increase in the iTraxx index implies a deteriorating macroeconomic condition. In this sense, we expect sovereign CDS spreads to be positively related to the iTraxx index. The iTraxx index also contains a proxy for investors' risk appetite. When investors become more risk averse, they will ask for higher credit spread for both corporate bonds and sovereign bonds. This again suggests a positive relationship between iTraxx and the sovereign CDS spreads.

If changes in the iTraxx index fully capture changes in investors' risk appetite, there is no need to include an additional proxy for the risk appetite into the analysis. Fontana and Scheicher (2010) find that the risk appetite proxy constructed from the Chicago Board Options Exchange Market Volatility Index(VIX) is not significant when the iTraxx index is included in the regression. Nevertheless, we add an additional proxy for investors' risk appetite for robustness. More specifically, we use the difference between the implied and realized volatility of EuroStoxx 50 return as the proxy for the global risk premium. This variable captures the pricing of the volatility risk, and therefore contains information on the investors' risk appetite (Longstaff et al., 2011). The implied volatility is the VSTOXX index directly available from Datastream while the realized volatility is estimated by the Garman and Klass (1980) estimator using a rolling 20-day window.

Finally, we include the nominal Euro-US Dollar exchange rate as a covariate. It is measured by the amount of Euros per 100 US dollars. Thus, a higher value means a depreciation of the Euro against the US dollar. We expect a positive sign of this variable. In other words, a depreciation of the Euro increases the sovereign CDS spread. The exchange rate is taken as a global variable since the exchange rate is determined by the macroeconomic fundamentals of the EMU rather than a single member state.

3.3 Orthogonalization

Financial asset returns are highly correlated to each other (see Table 2). That means including different asset returns into the regression can cause a multicollineararity problem which affects identification. Therefore, it is better to orthogonalize the variables before using them as covariates in the regression. We follow Dieckmann and Plank (2011) to construct the orthogonalized value of a variable as the sum of the estimated intercept and residuals of a regression of that variable on other covariates correlated to it. More specifically, domestic Financials index returns are regressed on the domestic stock market returns and the World financials index return; the World Financials index return is regressed on the global stock market return. Dieckmann and Plank (2011) do not orthogonalize the domestic stock market returns and the European corporate credit spread. Fontana and Scheicher (2010) suggest that orthogonalizing the domestic stock market returns also helps improve identification. Therefore, we orthogonalize the domestic stock market returns by regressing them on the global stock market return and construct the domestic stock market volatility indicators using the orthogonalized series. Alexander and Kaeck (2008) find that changes in the iTraxx index can be explained by changes in VSTOXX and changes in global stock and bond market conditions. Thus, to facilitate identification, we orthogonalize the change in the iTraxx index by regressing it on the change in the VSTOXX index, the global stock market return, the World Financials index and the 10-year German Bund rate.

4 OLS regression analysis

Table 4 summarizes the estimation results of the following linear OLS regression model.

$$\Delta CDS_t = \Delta x_t^{'} \beta + \epsilon_t, \tag{8}$$

where CDS_t is the sovereign CDS spread, x_t is the vector of covariates listed in Table 1, ϵ_t is the i.i.d. error term and Δ is a first difference operator. The OLS regressions assume that ϵ_t is independent of x_t. We follow the previous studies to run the regression with first differenced data.[15] This approach facilitates comparison of the results. Con-

[15]See Table 3 for descriptive statistics of first differenced data.

sistent with previous studies, our OLS results suggest that changes in the global bond market conditions have strong explanatory power to changes in sovereign CDS spreads. More specifically, increases in the 10-year German Bund rate significantly reduce the sovereign CDS spreads of Ireland, Italy and Spain; increases in the European corporate credit spreads significantly increase the sovereign CDS spreads of Greece, Ireland, Italy and Spain; better Euro-area economic performance (a higher EuroStoxx 50 return) significantly reduces the sovereign CDS spreads of Italy and Spain. These results are also consistent with the theoretical expectation *under the rational expectations assumption*, as we discussed in the last section. Consistent with the private-to-public risk transfer hypothesis, improvement in local financial firms' performance can reduce the sovereign CDS spread. This reduction effect is statistically significant in Italy and Portugal. Signs of the estimated coefficients of the World Financials index are positive, which is not only different from the finding of Dieckmann and Plank (2011), but also different from the theoretically expected sign we discussed in the last section. However, due to the econometric deficiency of equation (8), both the point estimates and the inference based on it are not reliable. Serial independence test results in Table 4 suggest that even if there is just one regime, inference based on standard errors reported in Table 4 will be distorted. If the single-regime assumption holds, the serial correlation problem can be corrected by using the serial-correlation robust standard errors for inference. However, if the single-regime assumption fails, even the serial independence test results in Table 4 will be unreliable.

Testing for regime switching is quite tricky because there are nuisance parameters that are only identifiable under the alternative hypothesis of two regimes but not under the null hypothesis of one regime. More specifically, a single-regime model can be represented in three different ways. First, it can be taken as a model with two regimes with the same regression coefficients. In this case, the probability associated with each regime is not identifiable. In the other two ways, the single-regime model can be taken as a model with two regimes under which the regression coefficients differ but one of the regime happens with zero probability. In such ways of representation, the regression coefficients of the regime which happen with zero probability are not identifiable. In addition, because probabilities cannot be larger than one, there is a boundary condition imposed in the estimation of the regime-switching model. Due to those facts, the typical likelihood ratio test statistics do not follow the usual χ^2 limiting distribution. Cho and

White (2007) propose a quasi-likelihood ratio test for regime switching and tabulated critical values at the 5 percent level. Carter and Steigerwald (2011) point out that critical values reported by Cho and White (2007) are based on 10,000 replications, but fewer than 100,000 replications do not produce stable critical values. They provide 5 percent critical values based on 100,000 replications. Table 5 reports the quasi-likelihood ratio test statistics for the null of one regime against an alternative of two regimes. Those values are far larger than the critical values tabulated in Carter and Steigerwald (2011). Therefore, the null hypothesis of a single regime is clearly rejected, and we should not make inference based on the OLS model.

5 Regime switching model analysis with instrumental variables

Like the OLS model, the standard regime switching models also assume that the error term is independent of the covariates. However, in our context, this assumption may not be plausible. It is possible that the insurance premium of sovereign borrowing affects the borrowing cost and therefore affect the domestic economy. In this case, the local variables are not exogenous and the standard maximum likelihood estimation of a regime switching model will give us biased results. Kim (2009) proposes a two-step maximum likelihood estimator with instrumental variables to solve this problem. Formally, the model can be written as follows:

$$\Delta CDS_t = \Delta x_t' \beta_{S_{1t}} + e_t, S_{1t} = 1, 2, ..., J_1, \tag{9}$$

$$\Delta x_t = Z_t' \gamma_{S_{2t}} + \Sigma_{v,S_{2t}}^{1/2} v_t, S_{2t} = 1, 2, ..., J_2, \tag{10}$$

where S_{1t} and S_{2t} are unobservable state variables; $Z_t = I_k \otimes z_t$, I_k is a $k \times k$ identity matrix with k being the dimension of x_t, $\otimes$ denotes the Kronecker product[16], and z_t is a $q \times 1$ vector of instrumental variables; $\Sigma_{v,S_{2t}}$ is a $k \times k$ matrix; J_1 and J_2 denote the

[16]Let a_{ij} be the element on the ith row and the jth column of a $m \times n$ matrix A. A$\otimes$B is defined as

$$\begin{pmatrix} a_{11}B & ... & a_{1n}B \\ \vdots & & \vdots \\ a_{m1}B & ... & a_{mn}B \end{pmatrix}$$

.

number of states; the joint distribution of e_t and v_t is

$$\begin{pmatrix} v_t \\ x_t \end{pmatrix} \sim i.i.d.N \left(\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} I_q & \rho_{S_{1t}} \sigma_{e,S_{1t}} \\ \rho'_{S_{1t}} \sigma_{e,S_{1t}} & \sigma^2_{e,S_{1t}} \end{pmatrix} \right),$$

$\rho_{S_{1t}}$ is a vector of correlation coefficients, and $\sigma_{e,S_{1t}}$ is the standard deviation of e_t. Equation (9) is similar to equation (8) but now the parameters in β change with the unobservable state variable S_{1t}. The Lucas critique suggests that a regime shift in the policy process governing equation (9) can lead to a regime shift in the dynamics of the CDS spread determinants. Therefore, we allow regime shifts in equation (10) as well. The unobservable state variable S_{2t} is correlated to S_{1t} according to the Lucas critique. One way to estimate the system composed of equations (9) and (10) is to specify the joint process of S_{1t} and S_{2t} and estimate the model by a joint maximum likelihood method. However, as pointed out by Kim (2009), such a joint estimation typically has too many parameters to estimate and suffers from the "curse of dimensionality". Furthermore, S_{2t} will be correlated to but different from S_{1t} if there is no perfect policy credibility and the agents have to learn to respond to the policy. Kim (2009) suggests that a two-step estimation approach which ignores the correlation between the state variables suffers less from the "curse of dimensionality". It has better finite sample performance than the joint maximum likelihood estimation when the correlation between S_{1t} and S_{2t} is not perfect. Moreover, it is more robust when the instrument variables are weak. The two-step approach of Kim (2009) first estimates equation (10) as a standard regime switching model. This procedure will give consistent estimates for $\gamma_{S_{2t}}$ and $\Sigma_{v,S_{2t}}$ since there are no endogenous covariates in equation (10). The elements of the residual vector $\hat{v}_t$ are used as control variables in the second-step estimation of equation (9).[17] Kim (2009) proves that this two-step approach will give us consistent estimates for the parameters in equation (9).[18]

To save degrees of freedom, we restrict the number of possible states for both S_{1t} and S_{2t} to two. We instrument the local determinants of the CDS spread ($\Delta sdri_t$, $\Delta svol_t$, and $\Delta fdri_t$) by the second and third lags of those local variables and the lagged dependent variable ΔCDS_{t-2} and ΔCDS_{t-3}. Table 6 summarizes our two-step estimation results

[17]See the appendix for a brief description of major steps of the second-step estimation.

[18]The second-step standard errors are biased due to the generated regressor problem. The standard errors in the tables are corrected using the method provided by Kim (2009).

of equation (9). Changes in the global bond market conditions (*gbi* and/or *itraxx*) remain to be significant explanatory variables for changes in country-specific sovereign CDS spreads under at least one regime. Moreover, the estimated signs of *gbi* and *itraxx* are consistent with the theory *under the rational expectations assumption*. More specifically, the 10-year German Bund rate (*gbi*) has a negative sign when significant, suggesting that investors expect a lower sovereign credit risk when Germany has a better economic performance. The iTraxx index has a positive sign when significant. As we discussed above, both a worse business climate in the European countries and a higher degree of risk aversion can lead to a higher iTraxx index. Therefore, both a worse economic state of EU and a higher degree of risk aversion can increase the prices of insurances on the sovereign bonds. Similar to the finding by Fontana and Scheicher (2010), the other proxy for investors' risk appetite, *vp*, is not significant when the iTraxx index is included as a regressor. The World Financials index is significantly negative under one regime in Italy. This suggests that there is a private-to-public risk transfer in Italy. Under the specific regime, a worse performance of the global financial sector increases the possibility that foreign countries have to spend money to bail out their own financial firms and hence less willing to help the home country. As a result, the sovereign CDS spread increases. Note that it is the performance of the global rather than local financial industry that matters. This finding suggests that compared to the possibility that the Italian government has to bail out its domestic financial firms, the market is more concerned about whether there will be international financial assistance if Italy is in trouble. Under regime 2, $\Delta fgro_t$ turns insignificant while the proxy for domestic economic performance turns significant in Italy. This suggests that under this regime, investors care more about the Italian economy than contingent government debt for bailing out the financial sector. Note that the signs of the estimated coefficients of the World Financials index are positive in some sample countries in some regimes. However, those coefficients are not statistically significant. Hence, it is better to be interpreted as no effect rather than a positive effect. Our results do not support Hypothesis 1, the rational sunspot-equilibria interpretation of the animal spirits. The slopes rather than intercepts are regime-dependent in the sample countries. Hypothesis 2 is also not plausible, though it allows regime-dependent slope changes. Under one of the two regimes, none of the indicators for economic performance and financial health are significant. In this case, Hypothesis 2 will lead to the conclusion that both economic growth and financial stability become unimportant for the government, which is very unlikely to be true. However, our results are consistent with Hypothesis 3. Increasing market uncertainty can enlarge cognitive biases in the market-based indicators for economic and financial health, making them useless for sovereign CDS pricing. Note that the iTraxx index contains a market-based proxy for risk aversion *under the rational expectations assumption*. It becomes insignificant when all observable fundamentals become insignificant. This does not mean that the degree of risk aversion does not matter. Instead, it means that the market-based measure of investors' risk appetite becomes very imprecise when market uncertainty is high and the rational expectations approximation is far from reality.

5.1 Tests for endogeneity and serial independence

Kim (2009) suggests that endogeneity of the explanatory variables can be tested by the standard Wald test using the second-step estimation outputs. More specifically, in the two-step estimation, endogeneity is captured by the first-step regression residuals of the endogenous variables on the instrumental variables. These residuals are used in the second-step regression as control variables to eliminate the endogeneity. Therefore, we can test for endogeneity by testing the statistical significance of the first-step residuals in the second-step regression. Formally, the second-step estimation equation can be written as

$$\Delta CDS_t = \Delta x_t^{'}\beta_{S_{1t}} + \hat{v}_t^{'}\theta_{S_{1t}} + \omega_t, S_{1t} = 1, 2, ..., J_1, \tag{11}$$

where $\theta_{S_{1t}}$ is a vector of regime-dependent coefficients, $\hat{v}_t$ is the first-step estimate for v_t, and ω_t is an i.i.d. normal random variable given a specific value of S_{1t}. The variance of ω_t changes across regimes. We denote it by $\sigma_{\omega,S_{1t}}$.[19] No endogeneity means $\theta_1 = \theta_2 = ... = \theta_{J_1} = 0$. Under the null hypothesis of no endogeneity, the asymptotic distribution of the Wald statistics $\hat{\theta}' c\hat{o}v(\hat{\theta})^{-1}\hat{\theta}$ is $\chi^2(h)$, where cov denotes the covariance; $\hat{\theta} = [\hat{\theta}_1' = \hat{\theta}_2' = ... = \hat{\theta}_{J_1}']'$ is the vector of estimated values for $\theta_{S_{1t}}$, $S_{1t} = 1, 2, ..., J_1$; h is the dimension of $\hat{\theta}$. Table 7 summarizes the Wald test results. The null hypothesis of variable exogeneity is rejected in all sample countries, except Greece. This verifies the importance of controlling for potential endogeneity.

Since we cannot directly apply the Hamilton (1996) test for autoregression to our

[19]See Kim (2009) for details.

regime-switching model with endogenous variables, we test for autoregression by adding the lagged dependent variable, ΔCDS_{t-1}, to the second-step equation and test the statistical significance of the autoregressive term. In order to avoid correlation between higher-order lags of ΔCDS_t and ΔCDS_{t-1}, we exclude them from the original instrument variable set. That is, we only use lags of the local variables as instrument variables. Table 8 summarizes the estimated coefficients of ΔCDS_{t-1} and their standard errors. The lagged dependent variable is not significant in any sample country under either regime, which suggests no serial correlation in the original model.

5.2 The endogeneity of the performance of global financial sector

In the econometric analysis above, we considered only the potential endogeneity of the local variables. Now we consider the potential endogeneity of a global variable: the change in the performance of the global financial sector, $\Delta fgro_t$. Such endogeneity can arise if financial firms outside the home country are highly involved in the trading of the specific country's sovereign CDS contracts.[20] Taking $fgro$ as an additional endogenous variable, we re-estimate the regime switching model. We use the second and third lags of $\Delta sdri_t, \Delta svol_t, \Delta fdri_t, \Delta fgro_t$ and the lagged dependent variable ΔCDS_{t-2} and ΔCDS_{t-3} to instrument the potentially endogenous variables ($\Delta sdri_t, \Delta svol_t, \Delta fdri_t, \Delta fgro_t$). We test the endogeneity of $fgro$ based on the new estimation results. As we mentioned in the last subsection, the test for endogeneity is equivalent to the test for the statistical significance of the corresponding first-stage residuals. Table 9 summarizes the test results. Those results suggest that changes in the Irish and Portuguese sovereign CDS spreads have significantly affected changes in the performance of financial firms outside those two countries at least under one regime. Table 10 reports the estimation results for Ireland the Portugal, taking $fgro$ as an endogenous variable. The previous result that changes in the fundamental variables do not explain changes in the Irish or Portuguese sovereign CDS spreads under regime 2 is unchanged. This means that the type-2 animal spirits of investors are indeed the driver of changes in the Irish and Portuguese sovereign CDS spreads under regime 2. Changes in the Euro-Dollar rate and the iTraxx index significantly affect changes in the Irish sovereign CDS spread under regime 1. More

[20]See OECD (2012).

specifically, a depreciation of the Euro relative to the US Dollar and an increase in the European corporate CDS spread lead to an increase in the Irish sovereign CDS spread. The significant positive sign of the iTraxx index suggests that either a worse business climate increases the sovereign credit risk or a higher degree of risk aversion increases the insurance premium for the sovereign borrowing. In Portugal, under regime 1, the 10-year German Bund rate appears to be the only significant fundamental driver of the sovereign CDS spread. The negative sign of *gbi* suggests that a larger increase in the German growth rate implies a higher increase in the probability that the EMU will provide financial support to the Portugal government if it is in trouble.

6 Conclusion

We have studied the determinants of changes in the sovereign CDS spreads of five Euro-area countries (Greece, Ireland, Italy, Portugal and Spain) after the failure of Lehman Brothers. Two distinct regimes under which the coefficients of the determinants differ are identified.

On the one hand, under regime 2, the usual determinants of changes in the sovereign CDS spreads of Greece, Ireland, and Portugal lose their explanatory power.[21] We argue that the animal spirits; that is, the psychological movements of cognitively limited investors are the key drivers of the sovereign CDS spreads in such situations. This has important implications for both policy makers and academic researchers. As a widely-used indicator of the sovereign credit risk, the sovereign CDS spread can be highly distorted in the sense that it can be completely disconnected from the country's fundamental economic movements. In the rational expectations framework, the existence of non-fundamental determinants of sovereign CDS spreads does not necessarily mean that the CDS spreads cannot predict sovereign defaults. That is because non-fundamental sunspot shocks to investors' expectation can lead to self-fulfilling sovereign debt crises. If the market believes that a debt crisis is under way, it will happen. And if the market participants are perfectly rational, the sovereign default probabilities will be correctly included in the pricing of the corresponding sovereign CDS spreads. However, our empirical results do not support the story of rational self-fulfilling debt crises. Rather, there are periods in which

[21] Under this regime, the market is more turbulent in the sense that the conditional variance of changes in the sovereign CDS spreads is larger than under regime 2. See Tables 6 and 10.

the boundedly rational market participants fail to price the sovereign credit risk correctly. In this case, the sovereign CDS spread is not very useful for evaluating the sovereign credit risk. Hart and Zingales (2011) suggest using CDS spreads on the long-term debt of large financial institutions to regulate them. The purpose is to contain systemic risk caused by the failure of those too-big-to-fail financial institutions. The idea is that a high CDS spread signals a high default risk of debt issued by the institution. In this case, the financial institution should be required to issue more equity or the regulator should intervene. There may be a tendency to extend this idea to public debt management. That is, one may suggest using sovereign CDS spread to monitor the sovereign credit risk and guide actions to prevent or resolve a sovereign debt crisis. However, a precondition of such a policy proposal is that sovereign CDS spreads should be reliable indicators of the sovereign credit risk. Unfortunately, our results suggest that the reliability of sovereign CDS spreads as indicators of the sovereign credit risk is questionable.

On the other hand, our results also suggest that there are periods in which the fundamental variables matter. In addition, the estimated effects of those fundamental variables reflect at least bounded rationality of the market participants. More specifically, we find that in the periods when the investors behave more rationally, the global bond market conditions are particularly important for the pricing of the sovereign CDS spreads. A better economic prospect of Germany or a better European-wide business climate implies a higher chance that other union members will be willing to provide financial support for the fiscally distressed countries. Therefore, the market confidence on sovereign borrowing will be enhanced and sovereign CDS spreads decrease. A depreciation of the Euro against the US dollar significantly increases Italian and Spanish sovereign CDS spreads. This result also suggests that the Euro-area economic prospect can affect an individual member country's sovereign credit risk. It is because the depreciation of the Euro can signal a weaker Euro-area economy. Another interesting finding is that there are periods in which traders of the Italian sovereign CDS contracts are concerned by the performance of the global financial industry. A worse performance of the global financial industry increases the probability that foreign countries will have to bail out their own financial sectors and become less willing to provide financial aids for Italian fiscal reforms. By contrast, in some other periods, investors in the market of Italian sovereign CDS contracts only cares about the domestic economic performance of Italy, and pay little attention to the performance of the global financial sector. The reason for this change in investors' concern can be an interesting topic for future research.

Table 1: Variable definitions

Variable	Definition
forex	Nominal Euro to US Dollar exchange rate, the amount of Euros per 100 US Dollars
stoxx	EuroStoxx 50 return (orthogonalized), percentage point
gbi	10-year benchmark German Bund interest rate, basis point
itraxx	iTraxx Europe 10-year CDS spread (orthogonalized), basis point
vp	Volatility risk premium, percentage point
fgro	MSCI World Financials index return (orthogonalized), percentage point
sdri	DJTM domestic stock market return (orthogonalized), percentage point
svol	GARCH(1,1) Domestic stock market volatility, percentage point
fdri	DJTM Financials index return (orthogonalized), percentage point

Notes: All data are from Datastream.
See the texts for detailed description on the orthogonalization of variables.
The volatility risk premium is proxied by the difference between the implied volatility and the Garman-Klass realized volatility of EuroStoxx 50.

Table 2: Correlation between stock market returns

Greece					Ireland				Italy			
	stoxx	fgro	sdri	fdri	stoxx	fgro	sdri	fdri	stoxx	fgro	sdri	fdr
stoxx	1.00				1.00				1.00			
fgro	0.84	1.00			0.84	1.00			0.84	1.00		
sdri	0.68	0.59	1.00		0.77	0.77	1.00		0.95	0.81	1.00	
fdri	0.61	0.53	0.95	1.00	0.59	0.63	0.74	1.00	0.89	0.78	0.96	1.00
Portugal					Spain							
	stoxx	fgro	sdri	fdri	stoxx	fgro	sdri	fdri				
stoxx	1.00				1.00							
fgro	0.84	1.00			0.84	1.00						
sdri	0.78	0.64	1.00		0.92	0.77	1.00					
fdri	0.58	0.51	0.73	1.00	0.88	0.78	0.97	1.00				

Notes: Correlation coefficients are calculated using non-orthogonalized data.

7 Appendix

In this appendix, we show the major steps of the second-step estimation for our two-state model. Our purpose is to estimate $\beta_{S_{1t}}$, $\theta_{S_{1t}}$, $\sigma_{e,S_{1t}}$ and p_{ij}, the transition probability from state i to state j. From equation (10), we have

$$\hat{v}_t = inv(\hat{\Sigma}^{1/2}_{v,S_{2t}})(\Delta x_t - Z_t'\hat{\gamma}_{S_{2t}}), \tag{12}$$

where $inv(\cdot)$ denotes the inverse, and $\hat{\Sigma}^{1/2}_{v,S_{2t}}$ and $\hat{\gamma}_{S_{2t}}$ denote the first-step estimates for $\Sigma^{1/2}_{v,S_{2t}}$ and $\gamma_{S_{2t}}$, respectively.

Using equations (11) and (12), we can derive the conditional density function of ΔCDS_t for given values of S_{1t} and S_{2t}. More specifically, for $j_1 = 1, 2$ and $j_2 = 1, 2$, the density functions can be represented as: $f(\Delta CDS_t|\Delta Z_t, \Delta x_t, S_{1t} = j_1, S_{2t} = j_2; \lambda_1, \hat{\lambda}_2) = \frac{1}{\sqrt{2\pi\sigma^2_{\omega,j_1}}} exp\{-\frac{1}{2\sigma^2_{\omega,j_1}}\{\Delta CDS_t - x_t'\beta_{j_1} - [inv(\hat{\Sigma}^{1/2}_{v,j_2})(\Delta x_t - Z_t'\hat{\gamma}_{j_2})]'\theta_{j_1}\}^2\}$, where λ_1 denotes the vector of parameters to be estimated in the second step, and $\hat{\lambda}_2$ denotes the vector of estimated parameters in the first step.

Using the standard smoother for the regime switching model, we can get, from the first-step estimation, $Prob(S_{2t} = 1|\Delta\tilde{x}_T)$ and $Prob(S_{2t} = 2|\Delta\tilde{x}_T)$, where $\Delta\tilde{x}_t$ denotes the historical information on Δx until time t, T is the end of the sample period.[22] We can calculate the conditional densities for $j_1 = 1, 2$: $f(\Delta CDS_t|\Delta Z_t, \Delta x_t, S_{1t} = j_1; \lambda_1, \hat{\lambda}_2) = f(\Delta CDS_t|\Delta Z_t, \Delta x_t, S_{1t} = j_1, S_{2t} = 1; \lambda_1, \hat{\lambda}_2) \times Prob(S_{2t} = 1|\Delta\tilde{x}_T) + f(\Delta CDS_t|\Delta Z_t, \Delta x_t, S_{1t} = j_1, S_{2t} = 2; \lambda_1, \hat{\lambda}_2) \times Prob(S_{2t} = 2|\Delta\tilde{x}_T)$.

Denote the historical information on ΔCDS_t until period $t-1$ by $\Delta\widetilde{CDS}_{t-1}$. If $Prob(S_{1t} = j_1|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T)$ is known, we can calculate the predictive density of ΔCDS_t by the following equation:

$f(\Delta CDS_t|\Delta\widetilde{CDS}_{t-1}, \Delta x_t; \lambda_1, \hat{\lambda}_2) = f(\Delta CDS_t|\Delta Z_t, \Delta x_t, S_{1t} = 1; \lambda_1, \hat{\lambda}_2) \times Prob(S_{1t} = 1|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T) + f(\Delta CDS_t|\Delta Z_t, \Delta x_t, S_{1t} = 2; \lambda_1, \hat{\lambda}_2) \times Prob(S_{1t} = 2|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T)$.[23]

However, we do not know $Prob(S_{1t} = j_1|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T)$. Given initial values $Prob(S_{10} = j_1|\Delta\widetilde{CDS}_0, \Delta\tilde{x}_T)$, we can calculate the filtered probabilities as follows: $Prob(S_{1t} = 1|\Delta C\tilde{D}S_{t-1}, \Delta\tilde{x}_T) = p_{11}Prob(S_{1,t-1} = 1|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T) + p_{21}Prob(S_{1,t-1} = 2|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T)$.

Similarly, $Prob(S_{1t} = 2|\Delta C\tilde{D}S_{t-1}, \Delta\tilde{x}_T) = p_{12}Prob(S_{1,t-1} = 1|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T) + p_{22}Prob(S_{1,t-1} = 2|\Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T)$.

[22] See Hamilton (1994) for details on the standard regime switching model.

[23] Note that in our model, Z_t includes past values of CDS_t and x_t.

The probabilities can be updated using the following equation:

$Prob(S_{1t} = j_1 | \Delta\widetilde{CDS}_t, \Delta\tilde{x}_T) = \frac{f(\Delta CDS_t | \Delta Z_t, \Delta x_t, S_{1t}=j_1; \lambda_1, \hat{\lambda}_2) \times Prob(S_{1t}=j_1 | \Delta\widetilde{CDS}_{t-1}, \Delta\tilde{x}_T)}{f(\Delta CDS_t | \Delta\widetilde{CDS}_{t-1}, \Delta x_t; \lambda_1, \hat{\lambda}_2)}$, where $j_1 = 1, 2$.

Iterating the procedure listed above, we can get the log likelihood function to be maximized.

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Table 3: Descriptive statistics

	CDS	forex	stoxx	gbi	itraxx	vp	fgro	sdri	svol	fdri
Greece										
Mean	42.81	0.05	-0.02	-1.39	0.12	0.00	-0.01	-0.03	-0.31	-0.04
Median	8.00	-0.01	-0.68	-1.5	0.22	0.18	0.03	-0.32	3.87	0.19
Maximum	3648.5	5.18	26.16	35.20	19.67	20.43	14.85	16.50	294.61	22.43
Minimum	-1622.96	-4.20	-13.44	-36.20	-30.00	-22.04	-11.72	-16.09	-393.21	-15.00
Standard deviation	397.60	1.40	5.88	12.59	7.19	5.34	3.83	6.89	47.40	3.88
Ireland										
Mean	3.47	0.05	-0.02	-1.39	0.12	0.00	-0.01	0.06	-4.28	-0.10
Median	2.94	-0.01	-0.68	-1.5	0.22	0.18	0.03	-0.05	-0.67	-0.35
Maximum	291.44	5.18	26.16	35.20	19.67	20.43	14.85	27.36	6.35	74.67
Minimum	-305.68	-4.20	-13.44	-36.20	-30.00	-22.04	-11.72	-26.38	-54.16	-76.52
Standard deviation	51.99	1.40	5.88	12.59	7.19	5.34	3.83	4.93	10.41	17.04
Italy										
Mean	2.66	0.05	-0.02	-1.39	0.12	0.00	-0.01	0.01	0.05	0.00
Median	1.25	-0.01	-0.68	-1.5	0.22	0.18	0.03	-0.02	-2.71	0.25
Maximum	132.22	5.18	26.16	35.20	19.67	20.43	14.85	7.35	156.34	8.48
Minimum	-101.55	-4.20	-13.44	-36.20	-30.00	-22.04	-11.72	-7.43	-19.96	-6.49
Standard deviation	25.85	1.40	5.88	12.59	7.19	5.34	3.83	1.99	16.34	2.64
Portugal										
Mean	5.48	0.05	-0.02	-1.39	0.12	0.00	-0.01	-0.01	-1.46	-0.05
Median	2.69	-0.01	-0.68	-1.5	0.22	0.18	0.03	-0.05	-1.45	0.20
Maximum	354.84	5.18	26.16	35.20	19.67	20.43	14.85	13.81	25.08	22.95
Minimum	-287.39	-4.20	-13.44	-36.20	-30.00	-22.04	-11.72	-10.21	-22.73	-22.36
Standard deviation	56.81	1.40	5.88	12.59	7.19	5.34	3.83	3.38	5.78	5.79
Spain										
Mean	2.04	0.05	-0.02	-1.39	0.12	0.00	-0.01	-0.01	-0.74	0.01
Median	1.50	-0.01	-0.68	-1.5	0.22	0.18	0.03	-0.13	-1.83	-0.01
Maximum	91.06	5.18	26.16	35.20	19.67	20.43	14.85	9.35	41.26	6.33
Minimum	-107.15	-4.20	-13.44	-36.20	-30.00	-22.04	-11.72	-9.95	-16.17	-7.60
Standard deviation	24.52	1.40	5.88	12.59	7.19	5.34	3.83	2.48	7.16	2.06

Notes: All data are first differenced.

Table 4: OLS results

	Greece	Ireland	Italy	Portugal	Spain
constant	0.40	0.02	0.02	0.07	0.02
	(0.30)	(0.04)	(0.02)	(0.04)	(0.01)
forex	-2.70	3.99	2.98**	1.23	2.43**
	(26.43)	(3.29)	(1.32)	(3.16)	(1.28)
stoxx	-10.07	-0.35	-0.69**	-1.28	-0.80**
	(7.26)	(0.87)	(0.36)	(0.87)	(0.34)
gbi	-0.46	-0.75**	-0.55***	-0.63	-0.40***
	(2.81)	(0.35)	(0.14)	(0.34)	(0.13)
itraxx	10.01**	1.18**	0.77***	0.55	1.01***
	(4.49)	(0.55)	(0.24)	(0.55)	(0.21)
vp	2.05	1.75	0.60	1.56	0.57
	(8.16)	(0.99)	(0.40)	(0.97)	(0.39)
fgro	4.24	1.43	0.48	2.71***	0.74
	(8.27)	(1.07)	(0.42)	(0.99)	(0.40)
sdri	-8.04	-0.27	-2.41***	-5.76***	-2.45***
	(4.52)	(0.83)	(0.92)	(1.20)	(0.67)
svol	-0.00	0.06	0.18	1.74***	0.53**
	(0.66)	(0.36))	(0.10)	(0.66)	(0.23)
fdri	-4.19	-0.24	-1.64***	-2.12***	-0.85
	(8.06)	(0.23)	(0.61)	(0.63)	(0.76)
Adjusted R-squared	0.29	0.16	0.43	0.32	0.42
Serial independence	0.56	0.00	0.00	0.13	0.03

Notes: Standard errors in parentheses. ***,** denotes significance at one and five percent level, respectively.
Serial independence is the Lagrange Multiplier (LM) test p value for serial correlation up to two orders.

Table 5: Tests for regime switching

	Greece	Ireland	Italy	Portugal	Spain
test statistics	462.6	115.1	75.96	157.8	75.48

Notes: Test statistics are Cho and White (2007) Quasi-Likelihood Ratio test statistics.
Null hypothesis: one regime.

Table 6: Regime switching model results-local variables instrumented

	Greece		Ireland		Italy		Portugal		Spain	
	1	2	1	2	1	2	1	2	1	2
constant	0.06	2.01	0.01	0.38	-0.01	0.03	0.04	0.22	0.01	0.00
	(0.04)	(3.64)	(0.03)	(6.05)	(0.01)	(0.06)	(0.02)	(1.31)	(0.02)	(0.03)
forex	5.47	-38.25	3.93	3.82	2.47***	2.19	1.63	12.94	2.67**	2.26
	(4.25)	(395.50)	(2.20)	(653.70)	(0.88)	(7.08)	(2.04)	(99.84)	(1.31)	(3.19)
stoxx	-1.23	-58.30	-0.56	(-6.67)	-0.17	-0.05	-0.58	-10.46	0.02	-1.10
	(1.20)	(89.51)	(0.51)	(195.20)	(0.35)	(1.51)	(0.61)	(25.06)	(0.50)	(0.83)
gbi	-0.84**	13.08	-0.34	-2.39	-0.27**	-0.95	-0.62***	(3.27)	-0.28	-0.74**
	(0.41)	(36.99)	(0.23)	(68.41)	(0.13)	(0.53)	(0.19)	(10.20)	(0.25)	(0.34)
itraxx	0.94	30.43	1.23***	-3.70	0.64***	1.79	0.72**	4.51	0.39	1.94***
	(0.79)	(33.88)	(0.32)	(199.50)	(0.23)	(0.96)	(0.36)	(10.29)	(0.27)	(0.47)
vp	-0.30	16.76	0.05	2.14	-0.16	1.67	0.29	2.36	0.10	0.87
	(1.20)	(140.90)	(0.54)	(267.00)	(0.44)	(1.13)	(0.58)	(36.34)	(0.59)	(0.87)
fgro	0.25	6.18	0.18	5.10	-0.97***	1.30	0.84	-1.61	-0.45	1.34
	(1.25)	(104.10)	(0.65)	(192.00)	(0.32)	(2.15)	(0.70)	(28.37)	(0.54)	(1.12)
sdri	0.02	-36.01	-1.89	21.13	1.04	-16.78**	-2.17	2.76	1.06	1.49
	(1.93)	(74.62)	(1.08)	(746.50)	(1.48)	(8.61)	(1.64)	(84.37)	(1.61)	(3.36)
svol	0.24	-0.69	-0.22	101.60	-0.06	1.21	0.65	-11.70	0.15	0.02
	(0.25)	(6.99)	(0.28)	(781.00)	(0.09)	(1.11)	(0.70)	(39.40)	(0.51)	(0.67)
fdri	0.76	-6.64	0.26	-16.33	0.18	2.98	-0.99	-3.00	-0.21	-0.98
	(4.82)	(64.09)	(0.34)	(102.80)	(1.02)	(5.22)	(0.76)	(45.02)	(1.44)	(5.74)
p_{ii}	0.97	0.91	0.97	0.89	0.93	0.94	0.97	0.87	0.98	1.00
σ_{ω}	0.39	7.00	0.22	0.55	0.08	0.16	0.20	0.49	0.08	0.18

Notes: Standard errors in parentheses. ***,** denotes significance at one and five percent level respectively.
p_{ii} denotes the probability of staying under regime i next period if i is the current regime.

Table 7: Endogeneity tests (local variables only)

	Greece	Ireland	Italy	Portugal	Spain
Wald statistics	1.36	20.39	14.25	13.98	16.97
p value	0.97	0.00	0.03	0.03	0.01

Notes: Testing for endogeneity of the local variables, taking the global variables as exogenous.

Table 8: Serial correlation tests for the regime switching model

	Greece	Ireland	Italy	Portugal	Spain
regime 1	0.01	-0.03	0.08	0.07	0.09
	(0.05)	(0.09)	(0.10)	(0.06)	(0.12)
regime 2	-0.21	1.40	-0.33	0.35	-0.02
	(0.73)	(27.4)	(0.27)	(0.22)	(0.34)

Notes: Estimated coefficients of ΔCDS_{t-1} with standard errors in parentheses. ***,** denotes significance at one and five percent level, respectively.

Table 9: Tests for the endogeneity of $fgro$

	Greece	Ireland	Italy	Portugal	Spain
regime 1	-0.11	-0.01	0.00	0.07**	-0.01
	(0.14)	(0.04)	(0.02)	(0.03)	(0.02)
regime 2	-2.14	1.18**	0.00	0.09	-0.06
	(8.08)	(0.06)	(0.06)	(0.34)	(0.05)

Notes: Standard errors in parentheses. ***,** denotes significance at one and five percent level, respectively.

Table 10: Regime switching model results-local variables and $fgro$ instrumented

	Ireland		Portugal	
	regime 1	regime 2	regime 1	regime 2
constant	0.01	0.02	0.02	0.11
	(0.02)	(1.11)	(0.02)	(0.43)
forex	4.38***	-0.86	3.22	-5.26
	(1.68)	(122.8)	(1.79)	(35.73)
stoxx	-0.71	1.40	-0.37	-4.36
	(0.48)	(34.61)	(0.54)	(12.61)
gbi	-0.34	-1.67	-0.49**	-0.91
	(0.18)	(10.85)	(0.20)	(4.02)
itraxx	1.07***	0.74	0.59	4.85
	(0.29)	(21.62)	(0.32)	(6.75)
vp	-0.10	4.51	0.24	3.57
	(0.48)	(27.53)	(0.56)	(13.63)
fgro	-0.87	25.38	-1.14	7.25
	(0.79)	(201.3)	(1.51)	(25.01)
sdri	-0.16	-1.00	0.57	-3.35
	(0.20)	(132.1)	(0.61)	(11.37)
svol	-0.30	5.39	-0.91	-1.37
	(0.21)	(51.48)	(0.96)	(33.96)
fdri	0.16	-40.80	-0.60	1.55
	(0.94)	(188.7)	(1.02)	(39.63)
p_{ii}	0.98	0.95	0.96	0.92
σ_ω	0.20	0.65	0.15	0.55

Notes: Standard errors in parentheses. ***,** denotes significance at one and five percent level respectively.
p_{ii} denotes the probability of staying under regime i in the next period if i is the current regime.