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Was the European sovereign crisis self-fulfilling? Empirical evidence about the drivers of market sentiments ☆

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ABSTRACT

We investigate the presence of self-fulfilling dynamics during the European sovereign crisis in the light of a theoretical model that we bring to the data. Our empirical framework allows us to empirically test the presence of self-fulfilling dynamics and to identify what may have driven the market sentiment during this crisis. To do so we estimate the probability of default of five European "peripheral" countries during January 2006 to September 2011 with a panel smooth threshold regression. Our estimation results suggest that (1) both the fundamentals and "animal spirit" ignited the European sovereign crisis; (2) we isolate the risk indicator through which investors' belief coordinate.

1. Introduction

The fiscal crisis in Greece that began in the autumn of 2009 has turned into a full-fledged sovereign crisis across Europe. The ten-year process of interest rate convergence has been wiped out and two distinct categories have emerged, the peripheral and the core European economies. Some economists put forward the presence of self-fulfilling speculation, or more precisely a situation where the fear of default is precisely what leads to default.[1] Was speculation self-fulfilling? Did it make sovereign states vulnerable to erratic speculative movements? If it has been the case, we would like to know the channels of coordination of market expectations. What drove market sentiments? The answers to these questions are important because they will determine subsequent regulation responses to address self-fulfilling herd behaviors.

The academic answer to these topical questions is still being debated. A few theoretical papers have argued in favor of the presence of self-fulfilling speculation. The fact that investors have put peripheral countries under scrutiny and the subsequent risk of speculative attacks have been emphasized by several works indeed. The reasons invoked are policy coordination failures among domestic executives (De Grauwe and Ji, 2012), the lack of a lender of last resort until the ECB's announcement of September 2012 to buy debt in unlimited amounts (Pâris and Wyplosz, 2013), the lack of a credible backstop at the European level such as a fiscal union (Chamley, 2012). In the existing papers however the surge in the spreads is

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[1] See e.g. Krugman in "A Self-Fulfilling Euro Crisis?" (the New York Times, August 7, 2011).

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due to a shift from optimistic to pessimistic market sentiments (Arghyrou and Kontonikas, 2012; Conesa and Kehoe, 2011; De Grauwe, 2011), while a more precise idea about what drives market sentiment would be needed.

Similar questions motivated the development of the "second generation" approach to currency crises.[2] In the second-generation model, the economic fundamentals are not sufficient to explain the sudden eruption of a crisis. The credibility of the government's commitment to maintaining a fixed-exchange rate regime becomes a subject of speculation by rational investors. The expectation of devaluation increases the cost of maintaining a peg and therefore the policy-maker will move to devalue. Such interaction between investors' beliefs and the government's objectives gives rise to self-fulfilling dynamics and multiple equilibria.

In this paper, we draw on these theoretical elements to give a functional form to the European sovereign crisis. More precisely, we refer to Jeanne and Masson's (2000) escape clause model that analyzes the benefits and costs to policymakers of exiting from a peg and specifies the probability of devaluation as applied to the European Monetary System crisis of 1993. We transpose their approach to model the probability of default in the context of the European sovereign crisis. Their framework has the advantage of proposing a linearized reduced form of the self-fulfilling speculation model, which is amenable to the data using econometric techniques. We extend Jeanne and Masson's (2000) model to obtain a linearized form of the econometric specification where not only the constant but also the coefficients of the fundamentals are allowed to vary depending on observable variables. In sum, we rely on their framework to assess the plausibility of self-fulfilling dynamics and multiple equilibria empirically during the European sovereign crisis and we adapt their model in order to associate observable variables with the underlying equilibria.

Indeed, a limit of Jeanne and Masson's approach is that the variable that coordinates investors with optimistic or pessimistic expectations is not observable. In other words, the model is tuned on the dynamics of the beliefs of market participants. Yet, it is key to better understanding the crisis and designing proper regulations.

To address this issue, we estimate the model within a threshold regression model. This specification has the advantage of offering a parametric solution to explain the nonlinearity. Indeed, it allows the parameters to change as a function of a threshold variable. We test different market signals that may have coordinated the expectations of market participants during the crisis and induced nonlinearity. We select six candidates among the financial variables that convey public information both about the economy as well as the mood of the market participants.[3] In sum, we use the panel smooth threshold regression approach (PSTR), initially proposed by González et al. (2005)), to estimate the sovereign spreads of five European "peripheral" countries: Spain, Ireland, Italy, Portugal and Greece during January 2006 to September 2011. This modeling strategy allows us to test the hypothesis that the elasticities in the spread determination model changed smoothly over time according to market signals, a nonlinear pattern that we interpret as evidence of multiple equilibria.

The contributions of this paper are threefold. First, we adapt and extend an existing model of self-fulfilling speculation to obtain a structural approach to assess the nature of the European sovereign crisis. Second, we bring the model to the data. Our estimation results suggest that both the fundamentals and "animal spirit" ignited the European sovereign crisis. Third, we adopt an empirical strategy to capture the dynamics of investors' beliefs during the crisis. We find that the Credit Default Swap (CDS) market has played a dominant role in driving market sentiments, a concerning finding given the opaqueness and concentration of this market. This finding leads us to suggest regulation implications.

The remainder of this paper is organized as followed. In the first section, we present our theoretical framework. In Section 2, we justify our empirical strategy and, in Section 3, we present the estimation procedure and data. Our empirical results are detailed in Sections 4 and 5. We conclude in Section 7.

2. The escape clause model and sovereign crises

The basic logic of self-fulfilling multiple equilibria derives from the circularity between market expectations and the policy-maker's decision. In the seminal model, the policy-maker's decision is about maintaining the fixed exchange rate or devaluing. In this Section, we transpose the reasoning to a situation in which the government decides to default or not. We rely on Jeanne and Masson (2000) (JM hereafter) and clarify which modifications we introduce to extend their model with the objective of reducing constraints.

The benefit of defaulting arises from the reduction of the interest burden on the outstanding debt. The authorities' optimal policy may validate market expectations *ex post*; that is, default if investors expect a default. This is due to the fact that default expectations increase the policymaker's benefit from defaulting. In fact, if investors become pessimistic, they sell government bonds, which increases the interest rate and interest rate payments and thus leads to the burden of public debt and the subsequent required austerity efforts. The benefit from defaulting then becomes higher. In sum, whether or not a default occurs depends on market expectations.

Default expectations depress output by rising the interest rate, which makes fiscal austerity more costly. In consequence on the one hand the benefit function of default ($B(\cdot)$) is higher than the cost (the loss of credibility in the capital market)

[2] Seminal papers include Obstfeld (1986), Eichengreen and Wyplosz (1993), Krugman (1996), Flood and Marion (1996, 1999). Jeanne (2000) proposed a taxonomy of second-generation models. In turn, in the context of a single currency area, the 3rd generation crisis model implying balance sheet effects due to currency mismatches seems less relevant to analyze the sovereign bond pricing during the crisis.

[3] For example, the *Euribor-OIS spread*, the difference between the Euro Interbank Offered Rate and the overnight indexed swap rate, which reflects both the cost of lending as well as the perception of risk by banks in lending to each other.

when fundamentals,[4] ϕ_t, fall short of a certain threshold, ϕ^*. On the other hand, this threshold results from a *strategic complementarity* between market expectations and the government's decision rule. To clarify this circularity, JM's model defines both investors' expectations and the government's benefit.

The expectations of identical rational investors are forward looking.[5] They not only depend on the investors' beliefs about future fundamentals but also on their own beliefs about the future beliefs of other investors. Rational investors know that the expectations of other investors will influence the benefits of defaulting in the next period as well as the objective probability of default, π_t[6]:

$$\pi_t = Prob[B(\phi_{t+1,}\pi_{t+1}) > 0|\phi_t] \tag{1}$$

Denoting ϕ^{*e} as the level of the fundamental under which investors expect the policymaker to default, the default probability is precisely the probability that fundamentals of the next period $t+1,\ \phi_{t+1}$, will be lower than ϕ^{*e}, conditionally on the current value of the fundamentals ϕ_t:

$$\pi_t = Prob[\phi_{t+1} < \phi^{*e}|\phi_t] = F(\phi_t, \phi^{*e}), \tag{2}$$

where the conditional cumulative distribution function $F(.,.)$ is supposed to have a negative first partial derivative F_1.[7]

In turn, the government chooses the optimal triggering level of the fundamental, ϕ^*, which makes its net benefit equal to zero, given investors' expectations:

$$\phi \mapsto B(\phi, F(\phi, \phi^{*e})).$$

As we suppose that the benefit function is a strictly decreasing function of the fundamental, ϕ^* is the unique level of the fundamental at which the net benefit is equal to zero. In sum, there is a unique equilibrium for each level of investors' expectations.[8]

Solutions with multiple equilibria, which are the key feature of JM's model, are due to shifts in investors' expectations. More precisely, if expectations shift from being optimistic to pessimistic, investors sell government bonds, which increases the interest rate and thus the benefit to the policymaker from defaulting. The self-fulfilling character of the default expectations comes from the fact that a high default probability tends to validate itself by increasing the net benefit of default.

To formalize this idea, JM (2000) suppose n different states, $s = 1, \ldots, n$, each one corresponding to a different level of the fundamental triggering default, in our case, ϕ^*_s. If the state at date t is s, the policymaker defaults if and only if $\phi_t < \phi^*_s$. At time t, there are as many critical thresholds ϕ^*_s as there are possible states of the economy[9] as perceived by the agents. The selection of the state depends on investors expectations. Therefore, the probability of default is the sum of the default probabilities, $F(\phi_t, \phi^*_s)$, weighted by the probability to be in one of the n different states of the economy in the future given the current state:

$$\pi_t = \sum_{s=1}^{n} Prob(s_{t+1} = s|s_t)F(\phi_t, \phi^*_s) \tag{4}$$

From here forward, we extend JM's model (2000) by assuming that the government refers to a different fundamental process, ϕ^s_t, at each state, s. More precisely, at each state, the government refers to a combination of different fundamentals, such as debt to GDP and unemployment. We assume that the weights of the fundamentals in this combination vary with the state. For example, the deeper the recession (bad state s), the higher the debt-to-GDP ratio and the closer to default. Hence, the government is more sensitive to the level of the debt-to-GDP ratio in a bad state of the economy than in a good state. We therefore have different associated critical thresholds ϕ^*_s.

Accordingly, we introduce the probabilities $F^{(s_t,j)}(\phi^{s_t}_t, \phi^{*e}_j)$ that fundamental ϕ^j_{t+1} in $t+1$ will be lower than the expected critical threshold, ϕ^{*e}_j, conditionally on the current fundamental, $\phi^{s_t}_t$, for each couple of states, (s_t, j).[10] Eq. (4) becomes:

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[4] In Jeanne and Masson (2000), $B(\cdot)$ is a function of economic fundamentals influencing the policymaker's decision and the optimal benefit function is not explicitly specified. We adopt the same approach.

[5] Backward looking rather than forward-looking expectations, or more precisely the presence of adaptive expectations, would not affect our main conclusion of multiple equilibria, as shown by Branch and Mc Gough (2004). We thank an anonymous referee for his suggestion.

[6] Such specification allows a representation of intertemporal strategies without changing the main results. For example, a default to domestic creditors can be an inter-temporal trade-off between capital repaid today and the level of taxes to rise in the future, which would imply solving a dynamic stochastic program. At the end it is possible to show that the benefit function still depends only on fundamentals as in the seminal setting. We thank an anonymous referee for this comment.

[7]

$$F_1(\phi_t, \phi^{*e}) = \frac{\partial F}{\partial \phi_t}(\phi_t, \phi^{*e}) \leq 0 \tag{3}$$

This property means that the fundamental process is not negatively autocorrelated, or, in other words, that an increase in the current value of the fundamental shifts the conditional cumulative distribution function of the next period fundamental in the same direction.

[8] See the justification in Appendix A.

[9] As in Jeanne and Masson (2000), we suppose a (strict) ordering of the different thresholds. But, in our case, we suppose that $\phi^*_1 > \ldots > \phi^*_n$ if state $s = 1$ is better than state $s = 2$ and so on.

[10] We suppose that each of these functions has the specific properties required in Jeanne and Masson (2000). See the Appendix A for details.

$$\pi_t(s_t) = \sum_{j=1}^{n} Prob(s_{t+1} = j|s_t) F^{(s_t,j)}(\phi_t^{s_t}, \phi_j^{*e}) \tag{5}$$

The circularity between market expectations and the policy-maker 's decision is represented here precisely: at any date, t, the government takes into account not only the state, s_t, and the corresponding fundamental process, $\phi_t^{s_t}$, but also the expectations of the investors through the probability $\pi_t(s_t)$ specified in Eq. (5). Accordingly, at each state $s_t(= 1, \ldots, n)$, the net benefit function of the government becomes a function of $\phi_t^{s_t}$ only, as specified as follows:

$$\phi_t^{s_t} \to B[\phi^{s_t}, \sum_{j=1}^{n} Prob(s_{t+1} = j|s_t) F^{(s_t,j)}(\phi_t^{s_t}, \phi_j^{*e})]$$

As previously, the government chooses the optimal triggering level of fundamental $\phi_{s_t}^*$, which makes its net benefit equal to zero:

$$\phi_{s_t}^* = H_{(s_t)}(\phi_1^{*e}, \ldots, \phi_n^{*e})$$

In a rational expectations equilibrium, each $\phi_{s_t}^*$ should satisfy the fixed point equations:

$$\forall s = 1, \ldots, n, \phi_s^* = H(s)(\phi_1^*, \ldots, \phi_n^*)$$

The vector of solutions $(\phi_1^* \ldots \phi_n^*)$ corresponds to the *sunspot* equilibria. There are at least n equilibria, but JM (2000) prove that this result implies an infinite number of equilibria. In addition, each equilibrium results from self-fulfilling dynamics. In fact, the level of the fundamental under which investors expect the policy maker to default, ϕ_s^{*e}, is validated, $\phi_s^{*e} = \phi_s^*$. Due to the properties of the different F-type functions and of the benefit function, these solutions exist and are unique (see details in the Appendix A).

In our last step, to bring the model to the data, we need to linearize (Eq. (5)). We specify the fundamental processes:

$$\phi_t^s = \alpha_{0,s} + \alpha_s' X_t + u_{t,s},$$

where α_s is the vector of coefficients. X_t is a vector of relevant economic fundamentals, and $u_{t,s}$ is an i.i.d. stochastic term reflecting other exogenous determinants of the policy maker's behavior. As in JM, we suppose that the fluctuations of the fundamentals are of limited magnitude at each state. Thus linearizing the default probability around the mean value $\bar{\phi}^{st}$ of ϕ_t^{st} yields:

$$\pi_t(s_t) \approx \rho_{0,s_t} + \rho_{s_t}' X_t + u_{t,s_t} \tag{6}$$

In the particular case where $n = 2$, one finds that ρ_{0,s_t} and ρ_{s_t} are given by:

$$\rho_{0,s_t} = \sum_{j=1}^{2} Prob(s_{t+1} = j/s_t)[F^{(s_t,j)}(\overline{\phi^{(s_t)}}, \phi_j^*) + F_1^{(s_t,j)}(\overline{\phi^{(s_t)}}, \phi_j^*)(\alpha_{0,s_t} - \overline{\phi^{s_t}})]$$

$$\rho_{s_t} = \sum_{j=1}^{2} Prob(s_{t+1} = j|s_t = i) F_1^{(s_t,j)}(\overline{\phi^{s_t}}, \phi_j^*)\alpha_{s_t}$$

where F_1 is the first partial derivative of F (Details are given in Appendix A).

If there is only one equilibrium, the probability of default is a linear function of the fundamentals; otherwise, it is a nonlinear function, according to:

$$\pi_t \approx \sum_{s=1,\ldots,n} [\rho_{0,s} + \rho_s' X_t + u_{t,s}] * 1_{s_t=s} \tag{7}$$

Note that, unlike in JM (2000), in our model, not only the constant but also the coefficients of the (observable) fundamentals X_t vary with the state of the economy. The self-fulfilling speculation model with multiple equilibria can now be tested empirically by testing the hypothesis of linearity. In the following, we explain our empirical strategy.

3. Empirical strategy: specification and estimation

The theoretical model involves non-linearity in the specification of the probability of default, a result that leads us to adopt a regime-switching approach in the estimation. Instead of adopting a Markov Switching Regime (MSR) approach á la Hamilton (1994) as JM (2000) did, we estimate the model using a threshold regression (TR) model. In fact, the MSR does not reveal the sources of nonlinearity: the determination model of default probability changes because of a shift in investors'expectations and these regime shift are due to a latent variable, a sunspot, that suddenly modifies the state of default expectations. In turn it is interesting to allow the expectations to change according to an observable signal that reveals market sentiments. This is precisely the advantage of a TR model that allows us to characterize nonlinearity as a function of an observable transition variable. More precisely, the default probability is estimated according to the following specification:

$$\pi_t = \left[\rho_{0,1} + \rho_1' X_t\right](1 - g(q_t, c)) + \left[\rho_{0,2} + \rho_2' X_t\right] g(q_t, c) + u_t, \tag{8}$$

where $g(\cdot)$ is an indicator function:

$$g(q_t; c) = \begin{cases} 1 & \text{if } q_t \leqslant c \\ 0 & \text{otherwise} \end{cases}$$

At each date, the observable variable, q_t, which is supposed to account for coordination in investors' expectations is compared to an estimated value called the location parameter, c. For illustration, q_t can be the sovereign grade of the country by rating agencies. If the sovereign grade is higher than c, the market is optimistic, which means that the estimated default probability equals $\hat{\pi}_t = \rho_{0,1} + \rho_1' X_t$ (regime 1). In turn, if the sovereign grade is downgraded below the location parameter, the market's expectations shift to pessimistic and the estimated default probability is equal to $\hat{\pi}_t = \rho_{0,2} + \rho_2' X_t$ (regime 2). However, this specification allows only a sharp transition. To circumvent this limit, a solution is to use a smooth transition function – a logistic function of order 1:

$$g(q_t; \gamma, c) = \frac{1}{1 + \exp\left[-\gamma(q_t - c)\right]} \quad , \gamma > 0. \tag{9}$$

This continuous function, bounded between 0 and 1, has an S-shape. The γ parameter determines the smoothness, i.e., the speed of the transition from one regime to the other. The higher the value of the γ parameter, the faster (i.e., sharper) the transition. Note that the model allows for an infinite number of intermediate regimes between regime 1 and regime 2 as defined above, which is well adapted to account for the infinite number of equilibria that are found by JM.

In sum, our empirical strategy has two main advantages. First, the introduction of an observable variable explaining the nonlinearity sheds light on what may coordinate investors' beliefs. Second, the infinite number of intermediate regimes is consistent with the theoretical result of an infinite number of equilibria.

From now on, we present the STR specification applied to panel data (PSTR model initially proposed by González et al. (2005)). The choice of panel data is motivated by the low time dimension of macroeconomic data and justified by the hypothesis that the countries of our panel are governed by the same type of economic forces. In addition, the PSTR model is a solution to account for individual heterogeneity (Fouquau et al., 2008).

The PSTR specification of Eq. (10) is the following:

$$\begin{aligned} \pi_{it} &= \mu_i + \rho_1' X_{it}(1 - g(q_{it}; \gamma, c)) + \rho_2' X_{it} g(q_{it}; \gamma, c) + u_{it} \\ &= \mu_i + \rho_1' X_{it} + (\rho_2' - \rho_1') X_{it} g(q_{it}; \gamma, c) + u_{it} \\ &= \mu_i + \beta_1' X_{it} + \beta_2' X_{it} g(q_{it}; \gamma, c) + u_{it} \end{aligned} \tag{10}$$

for $i = 1, \ldots, n$, with $\beta_1' = \rho_1'$ and $\beta_2' = (\rho_2' - \rho_1')$. The terms u_{it} are i.i.d. errors, μ_i represent individual fixed effects and q_{it} are the threshold variables introduced above.

The estimation of the PSTR model consists of several stages. In the first step, a null hypothesis of linearity is tested against the alternative hypothesis of a threshold specification. Then, if the linear specification is rejected, the estimation of the parameters of the PSTR model requires eliminating the individual effects, μ_i, by removing individual-specific means and then applying nonlinear least squares to the transformed model (see González et al., 2005).

In González et al. (2005) procedure, testing the linearity in a PSTR model (Eq. 10) can be done by testing $H_0: \ \gamma = 0$ or $H_0: \ \beta_2 = 0$. In both cases, the test is non-standard since the PSTR model contains unidentified nuisance parameters under H_0 (Davies, 1987). The solution is to replace the transition function, $g(q_{it}; \ \gamma, c)$, with its first-order Taylor expansion around $\gamma = 0$ and to test an equivalent hypothesis in an auxiliary regression. We then obtain:

$$\pi_{it} = \mu_i + \theta_0 \ X_{it} + \theta_1 \ X_{it} q_{it} + \epsilon_{it}^*. \tag{11}$$

In these auxiliary regressions, parameter θ_1 is proportional to the slope parameter, γ, of the transition function. Thus, testing the linearity against the PSTR simply consists of testing $H_0: \ \theta_1 = 0$ in (11) for a logistic function with an usual LM test.

4. Data

The estimation of the model of Eq. (10) is subject to two major data constraints. On the one hand, the macroeconomic variables included to measure economic fundamentals have a low frequency (quarterly or monthly) and some are available with a lag of two quarters. On the other hand, the sovereign crisis started in 2009, representing three years of crisis at the time of this analysis. Therefore, to obtain a critical number of observations, our estimation is based on an unbalanced panel of the five peripheral European countries in which the sovereign yield has been most under pressure, Greece, Ireland, Italy, Spain and Portugal, between January 2006 and September 2011.

As a dependent variable, we retain the sovereign bond spread (in percentage) which is supposed to be a proxy for the sovereign default probability.[11] This spread is defined as the difference between the sovereign bond yield and the risk-free rate

[11] In the light of the recent sovereign debt crisis affecting euro-area countries, a number of papers have focused on the determinants of spreads in euro-area countries.For example, Dötz and Fischer (2010) decompose sovereign spreads to generate country-specific default probabilities.

of the same maturity. For each country in the sample, we use the long-term German yield, which is the benchmark risk-free rate for the Euro area, and the government yield of this country at the same maturity. We rely on monthly observations of Maastricht criterion bond yields provided by the Eurostat database.

A key choice is the set of explanatory variables included in X_t in Eq. (10). We test the following variables: debt-to-GDP ratio, unemployment, unit labor cost, risk, liquidity.

First, the country's credit risk is traditionally related to fiscal sustainability. We therefore include the debt-to-GDP ratio from Eurostat.[12] To address the issue of poor quality of Greek statistical data, we use only revised data.

Other variables relevant in forming default expectations are those variables that may appear in the authorities' objective function. The economic activity and the country's competitiveness are potential candidates because the deterioration of these fundamentals increases the social cost of austerity efforts and thus the benefit from defaulting. We proxy the economic activity using the unemployment rate rather than GDP to avoid redundance with the debt-to-GDP ratio. The unit labor cost is included to proxy the country's competitiveness. These data are taken from Eurostat. The trade balance (a proxy for competitiveness) is excluded from the vector of determinants because of the specific behavior of Ireland, which ran a trade surplus (the variable is positive), contrary to the other countries in the sample. This variable was found to be not significant in other studies (De Grauwe and Ji, 2012). An issue with our macroeconomic data is that they are available only at a quarterly frequency (debt, unemployment and unit labor cost). To transform them to monthly frequencies, we used a local quadratic with the average matched to the source data.[13]

In line with the literature, we include a variable of liquidity risk and a measure of international risk aversion. Our proxy for liquidity is the size of the government's bond markets. For each country in the sample, liquidity is measured as the share of total outstanding Euro-denominated long-term government securities issued in the Euro zone. Data are available on a monthly basis from the European Central Bank. Following Borgy et al. (2011), our measure for international risk aversion is computed as the spread between US AAA corporate bonds and US 10-year sovereign bonds.[14] Data are available on a daily basis from Bloomberg. We compute the average of daily data to obtain monthly frequencies. In the following, we proceed to the estimation of Eq. (10) in two steps.

5. TV-PSTR estimation results

We start the empirical estimation of Eq. (10) using a TV-PSTR and then proceed to the PSTR in the next section. In this case, the threshold variable is imposed to be time. The primary objective is to check the rejection of linearity, which will be interpreted as a first indication for multiple equilibria. In fact, if the linearity hypothesis in the test presented below is rejected, this will indicate that the determination of default probability (proxied by the spread) was modified during the period of the estimation.

The TV-PSTR equation is the following:

$$\pi_{it} = \mu_i + \beta_1' X_{it} + \beta_2' X_{it} g(T; \gamma, c) + u_{it} \tag{12}$$

for $i = 1, \ldots, n$ and $t = 1, \ldots, T$, μ_i represent individual fixed effects and u_{it} are i.i.d. errors. X_{it} include: debt-to-GDP, squared debt-to-GDP, unemployment, unit labor cost, risk, liquidity. As the effect of debt is usually found to be nonlinear and this effect is captured through the introduction of the squared debt-to-GDP ratio (De Grauwe and Ji, 2012), we include it to avoid the rejection of linearity due only to this effect.

Table 1 reports the estimated parameters of the TV-PSTR and the linearity tests. The result of the parameter constancy test rejects the null hypothesis of a linear relationship at the 1% significance level ($LM = 87,\ 3$). It corroborates the hypothesis that the determination of the probability of default changed during the period. Other papers have also shown that the spread determination was not constant during the same period using break models or regime-switching features (Borgy et al., 2011; Mody, 2009). However it is not realistic to consider a sharp transition given the progressive increase in the spreads. Our approach has the advantage of allowing a smooth transition process (see Fig. 1). The threshold value, c, representing the inflexion point of the transition process, is located in March 2010. The complete modification of the spread determination occurred within one year between October 2009 and October 2010 (in October 2009, the spread determination was defined at 97% by regime 1 and in October 2010 at 97% by regime 2). Our TV-PSTR model thus correctly captures the increase in market tensions about the European sovereign starting with the announcement of the revision of the fiscal deficit in Greece by Prime Minister Papandreou in November 2009. The determination model of default probability for the European sovereign appears to have radically changed in Fall 2010 in respect to Fall 2009, a result that we interpret as evidence of a shift in investors' expectations.

In fact, Table 1 indicates that most coefficients increased: debt (from $\hat{\beta}_1 = 0.05$ to $\hat{\beta}_1 + \hat{\beta}_2 = 0.26$), risk (from 0.48 to 1.33), and unemployment (from -0.05 to 0.25). We mention that the increase in the weight of debt is slightly reduced by the

[12] We exclude deficit data to avoid collinearity with the rest of the economic variables. The correlation between the primary deficit and unemployment is 0.46 and that between the primary deficit and the unit labor cost is -0.37.

[13] We used Eviews software for this transformation. To check the robustness, we compared our results with a transformation based on a cubic spline with the last observation matched to the source data. We present the results in Table 5.

[14] In addition to international risk aversion, this spread can also be interpreted as a liquidity measure as outlined by Longstaff (2004) (we thank the anonymous referee for this comment).

Table 1
Linearity tests and estimation of the probability of default with a time-varying PSTR model.

Determinants	β_1	β_2
Debt	0.055***	0.209***
	(4.74)	(5.43)
Squared debt	0.000	−0.001***
	(0.33)	(−3.51)
Unemployement	−0.048*	0.297***
	(-3.19)	(7.51)
Unit labor cost	−0.011	−0.167***
	(−0.79)	(−6.34)
Liquidity	1.543*	−14.698***
	(1.78)	(−6.14)
Risk	0.480***	0.851*
	(9.85)	(1.77)
Smooth parameter γ		0.529
Loc Parameter		51.5
Linearity test		87.26***
RSS		76.28
Information crit. BIC		−1.22

Notes: The T-stat in parentheses are corrected for heteroskedasticity. β_1 and β_2 correspond to the coefficient in Eq. (11). β_1 is the coefficient in the first extreme regime. The coefficient in the second extreme regime is $\beta_1 + \beta_2$.
** Significant at the 5% level.
* Significant at the 10% level.
*** Significant at the 1% level.

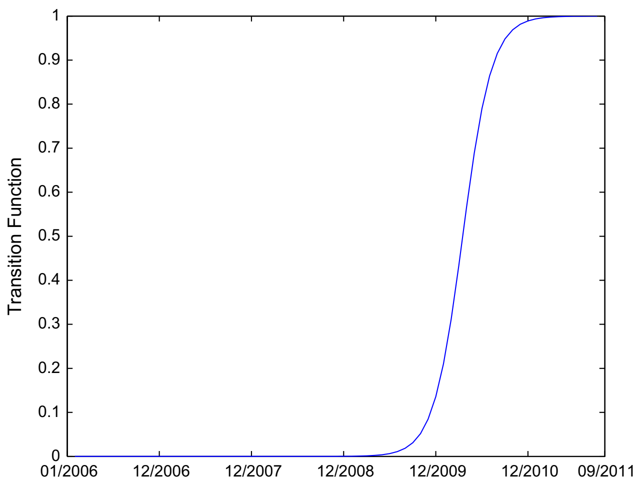


Fig. 1. Transition function with a TV-PSTR.

negative coefficient of the squared debt in the second regime (from 0 to −0.001). The effect of liquidity also increases significantly. While it has a sign contrary to expectations in the first regime ($\hat{\beta}_1 = 1.54$), it becomes highly negative in the second extreme regime ($\hat{\beta}_1 + \hat{\beta}_2 = -13.15$), implying that the lack of liquidity increases the probability of default (consistent with the linear findings in Beber et al. (2009)). In addition, the coefficient of our competitiveness indicator (ULC) goes from −0.04 to −0.19, contrary to the expected effect. However, eliminating ULC does not modify the value of the other estimated coefficients.[15] In total, the estimation reveals the increasingly important constraint on fiscal policy played by financial markets. At the same time, investors also became sensitive to the business cycle, a result that shows the potential counter-effective impact of fiscal austerity. The estimation results illustrate the dilemma faced by European policy makers between fiscal austerity and stimulating growth policies.

This first step gives an indication for multiple equilibria and identifies precisely the period of transition and its specific dynamics. Now, we would like to go one step forward and identify the drivers that instantaneously coordinated the expectations of all investors. To do so, in the following section, we proceed with the estimation of a PSTR model that allows the nonlinearity to depend on an observable variable.

[15] Results available upon request to the authors.

6. Sunspots or observable drivers of investors expectations?

We test different market signals that may have coordinated the expectations of market participants. We recall that the PSTR specification of the spread is as follows:

$$\pi_{it} = \mu_i + \beta_1' X_{it} + \beta_2' X_{it} g(q_{it}; \gamma, c) + u_{it} \tag{13}$$

for $i = 1, \ldots, n$ and $t = 1, \ldots, T$, μ_i represent individual fixed effects and u_{it} are i.i.d. errors. In order to estimate the PSTR model, we need threshold variables q_{it}. We select six candidates among financial variables that convey public information both about the economy as well as the mood of the market participants. The candidate threshold variables q_{it} are: rating, *Sovereign CDS*, *Bank Risk*, *i-traxx Europe*, *i-traxx Crossover*, *Euribor-OIS spread*.

First, *rating* is the average of the sovereign grades published by the three main international rating agencies, Standard and Poors, Moodys and Fitch (taken from Reuters). In fact, the sovereign crisis brought credit ratings agencies to the front. Rating agencies help investors overcome their lack of information about the variables that will determine whether a borrower will service debt. These agencies use qualitative letter ratings in descending order.[16] We use the linear transformation of Afonso (2007) to obtain a continuous numerical scale from the letter ratings.

Second, *sovereign CDS* is the premium of sovereign credit default swaps, which are bilateral contracts between a buyer and seller under which the seller sells protection against the credit risk of the reference country. The CDS premium, the insurance cost, is used here to measure market assessments of the health of borrowers and the likelihood of default. We select the 5-year maturity, which is the most traded contract in the CDS market, taken from Bloomberg.

Third, *Bank Risk* denotes the premium of CDS on the main banks in the country where the default probability is estimated. The nexus of the financial sector, sovereign credit risk, is a feature of financial crises in general (Reinhart and Rogoff, 2009) and the European sovereign crisis in particular (De Grauwe, 2010; Acharya et al., 2011). To avoid a credit crunch and loss of real sector output, governments engaged in large-scale financial-sector bailouts. Such bailouts are costly because they require immediate issuance of additional debt by the sovereign. This leads to an increase in the sovereign's credit risk. We use the average of the CDS premia of major banks weighted by the CDS market volume, taken from Reuters.

Fourth and fifth, we consider two broader indicators of the health of the corporate sector in Europe: i-Traxx Europe and i-Traxx Crossover are credit default swap index products. i-Traxx Europe comprises the most liquid 125 CDS referencing European investment grade credits while Crossover comprises the most risky 40 constituents at the time the index is constructed.

Last, *Euribor-OIS spread* captures the difference between the Euro Interbank Offered Rate and the overnight indexed swap rate. It reflects the risk banks perceive in lending to each other (the higher the spread, the more reluctant the banks are to lend to each other). The three last variables are taken from Reuters.

Before proceeding to the estimation we need to be cautious about a potential risk of simultaneity and more generally endogeneity between the dependent variable and three threshold variables, *Sovereign CDS*, *Bank Risk* and *Rating*. To address this issue a solution is to lag the variables to reduce endogeneity bias due to simultaneity. As a month lag may imply a significant loss of financial information about agents' expectations though, we implement two estimations, one with lagged threshold variables and another with contemporaneous variables. In addition our TV-PSTR estimations serves as a benchmark since time is an exogenous variable. We will conclude that our findings are robust if we obtain similar coefficient results in the three estimations.

For each model, the first step is to test the linear specification of the spread against a specification with threshold effects. The results of these tests are reported in Table 2. The linearity tests clearly reject the null hypothesis of a linear relationship regardless of which threshold variable is included in the specification. The remarkably high level of rejection provides a positive indication for multiple equilibria.[17] This is consistent with our preliminary result from the time-varying specification. The second step consists of selecting the best threshold variables, with the objective of identifying the drivers that mostly coordinate investors expectations. As suggested by González et al. (2005), the "optimal" threshold variable corresponds to the variable that leads to the strongest rejection of the linearity hypothesis. Among the six variables tested, the *sovereign CDS* is unambiguously the market variable that drives investors' expectations as it yields the strongest rejection statistics of the null hypothesis (LM = 282).[18] This first result illustrates the important role that the sovereign CDS market has played during the crisis. It is consistent with the findings of Delatte et al. (2012) pointing to the amplification role played by the credit derivative market in times of market distress. According to the estimation, the CDS market plays a more important role in coordinating investors' expectations than do the rating agencies, which rank second, also with very high rejection statistics (LM = 231). Bank Risk rank third, also with high rejection statistics (LM = 186). In comparison, the European corporate CDS indices (i-Traxx Europe and i-Traxx Crossover) and the Euribor-OIS spread have much lower rejection statistics (LM = 51.8, 77.9 and 39.9), which suggests that they are not good candidates for threshold variables. In total, the PSTR specification identifies three market variables that coordinate investors' expectations, with the sovereign CDS market clearly issuing the leading signal.

[16] S & P and Fitch use similar ratings from AAA to CCC-, while Moody' system goes from Aaa to Caa3. Although they do not use the same qualitative codes, there is a correspondence between each rating level.

[17] Given the quality issue of Greek statistics and to make sure that our results do not depend on the case of Greece, we re-estimated a sub-panel excluding Greece. Our results confirm our main conclusion about the indication of multiple equilibria. Results are available upon request.

[18] The order is not affected in the estimations using lagged threshold variables (results reported in Table 4).

Table 2
Linearity tests with a PSTR model.

	Sovereign CDS	Bank risk	Rating	ItraX Europe	Itrax	EURIBOR OIS
LM	282.2	186.3	231.7	77.9	51.81	39.87
p-value	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
RSS	19.83	50.9	57.9	140.2	142.1	148.9
BIC	−2.57	−1.63	−1.50	−0.61	−0.61	−0.56

Notes: The corresponding *LM* statistic has an asymptotic $\chi^2(p)$ distribution under H_0. The corresponding p-values are reported in parentheses.

Table 3
Estimation of the probability of default with a PSTR model (quadratic transformation).

	Model 1 Sovereign CDS		Model 2 Bank Risk		Model 3 Rating	
Determinants	β_1	β_2	β_1	β_2	β_1	β_2
Debt	−0.030**	0.211***	0.032	−0.097	0.313***	−0.291***
	(-2.61)	(4.51)	(0.79)	(-1.44)	(6.39)	(-6.23)
Squared debt	0.000***	−0.001***	−0.001**	0.002***	−0.001***	0.001***
	(3.68)	(-4.48)	(-2.14)	(3.73)	(-6.38)	(5.61)
Unemployement	−0.099***	0.335***	−0.253***	0.561***	0.804***	−0.791***
	(-2.97)	(3.63)	(-3.13)	(4.45)	(4.92)	(-4.6)
Unit labor cost	0.045**	−0.062**	0.056*	−0.087***	−0.237***	0.25***
	(2.16)	(-2.02)	(1.83)	(-2.41)	(-7.98)	(8.03)
Liquidity	1.694*	−4.31	19.314***	−35.68***	−0.801	−1.000
	(1.76)	(-0.92)	(5.80)	(-5.01)	(-0.14)	(-0.14)
Risk	−0.2	1.447*	−2.184***	4.455***	2.242***	−2.11***
	(-0.94)	(1.71)	(-4.05)	(4.55)	(7.39)	(-5.47)
Smooth parameter γ		0.002		0.003		0.554
Loc parameter		466.1		9.06		15.7
RSS		19.8		50.9		57.8
Information crit. BIC		−2.57		−1.63		−1.66

Notes: Strut notes: The T-stat in parentheses are corrected for heteroskedasticity. β_1 and β_2 correspond to the coefficient in Eq. (11). β_1 is the coefficient in the first extreme regime. The coefficient in the second extreme regime is $\beta_1 + \beta_2$.
* Significant at the 10% level.
** Significant at the 5% level.
*** Significant at the 1% level.

We examine more precisely the impact of these variables on the determination of default probabilities by investors. We consider which determinants have their weight changed most when the sovereign CDS premia increase. We also consider which determinants matter most to investors when their expectations based on these indicators become strongly pessimistic.

Table 3 reports the value of the estimated coefficients in the three models that best reject linearity with a contemporaneous transition variable. The coefficients are defined at each date and for each country as weighted averages of the values obtained in the two extreme regimes. The coefficients in the PSTR model can therefore be different from the estimated parameters defined in the extreme regimes, i.e., the parameters $\hat{\beta}'_1$ and $\hat{\beta}'_1 + \hat{\beta}'_2$ in Eq. 11. For each model, we first need to interpret the sign of parameter $\hat{\beta}'_2$, which indicates an increase ($\hat{\beta}'_2 > 0$) or a decrease ($\hat{\beta}'_2 < 0$) in the estimator as the threshold variable increases.

As announced before, to deal with the endogeneity issue, we re-estimate the PSTR models by introducing the lagged transition variables instead of the contemporaneous ones.[19] Table 4 reports the estimation results. We observe that the estimated coefficients in Tables 1 (TV-PSTR) 3 and 4 are very similar, a fact that suggests that the simultaneity bias does not influence the results. The estimated coefficient of the determinant variables risk and unemployment unambiguously increase in the second regime. The way investors price the fiscal situation is captured by the interaction of debt and squared debt, which makes a direct interpretation of the coefficients impossible. We plot it below. The coefficient of ULC becomes negative in the second regime, which is contrary to the expected sign. Only the evolution of liquidity is ambiguous as it is not consistent across the three models. Removing ULC and liquidity does not change our results.[20] Last we find similar patterns for a majority of the coefficients in the three selected models, which suggests that our estimations are robust.

We would like to examine the variation in the impact of each determinant during the period. However, as mentioned above, the coefficients could be different from the estimated coefficients in the extreme regimes. The only information is that

[19] This is the approach that is generally adopted as a first step to deal with endogeneity problems. Of course a formal test should be implemented; but, to our knowledge, such a test has not been described in the literature in the PSTR framework.
[20] Results available upon request.

Table 4
Estimation of the probability of default with a PSTR model and lagged variables (quadratic transformation).

	Model 1 $SovereignCDS_{t-1}$		Model 2 $BankRisk_{t-1}$		Model 3 $Rating_{t-1}$	
Debt	-0.04***	0.25***	0.01	0.05**	0.35***	-0.31***
	(-3.07)	(4.58)	(0.71)	(2.21)	(6.96)	(-6.62)
Squared debt	0.0002***	-0.0001***	0.0003***	0.001	-0.001***	0.001***
	(3.86)	(-4.53)	(3.96)	(-0.38)	(-6.82)	(5.78)
Unemployement	-0.09**	0.31***	-0.0025	0.04	$0.72_{(4.53)}$***	-0.70***
	(-2.44)	(3.08)	(-0.15)	(1.62)	(-4.30)	
ULC real	0.05**	-0.08**	-0.04***	-0.07***	-0.26***	0.26***
	(1.96)	(-2.20)	(-3.36)	(-4.37)	(-8.24)	(8.63)
Liquidity	2.41**	-7.47	0.13	-14.34***	-4.40	3.43
	(2.00)	(-1.34)	(0.22)	(-5.07)	(-0.96)	(0.63)
Risk	-0.27	1.64*	0.48***	2.47***	1.85***	-1.62***
	(-1.08)	(1.75)	(10.3)	(4.67)	(6.90)	(-4.70)
Smooth parameter γ	0.003		0.177		0.669	
Loc parameter	437.6		219.8		16.17	
Linearity test	277***		163***		224***	
RSS	22.88		89.31		58.56	
Information crit. BIC	-2.43		-1.05		-1.47	

Notes: Strut notes: The T-stat in parentheses are corrected for heteroskedasticity. β_1 and β_2 correspond to the coefficient in Eq. (11). β_1 is the coefficient in the first extreme regime . The coefficient in the second extreme regime is $\beta_1 + \beta_2$.
* Significant at the 10% level.
** significant at the 5% level.
*** Significant at the 1% level.

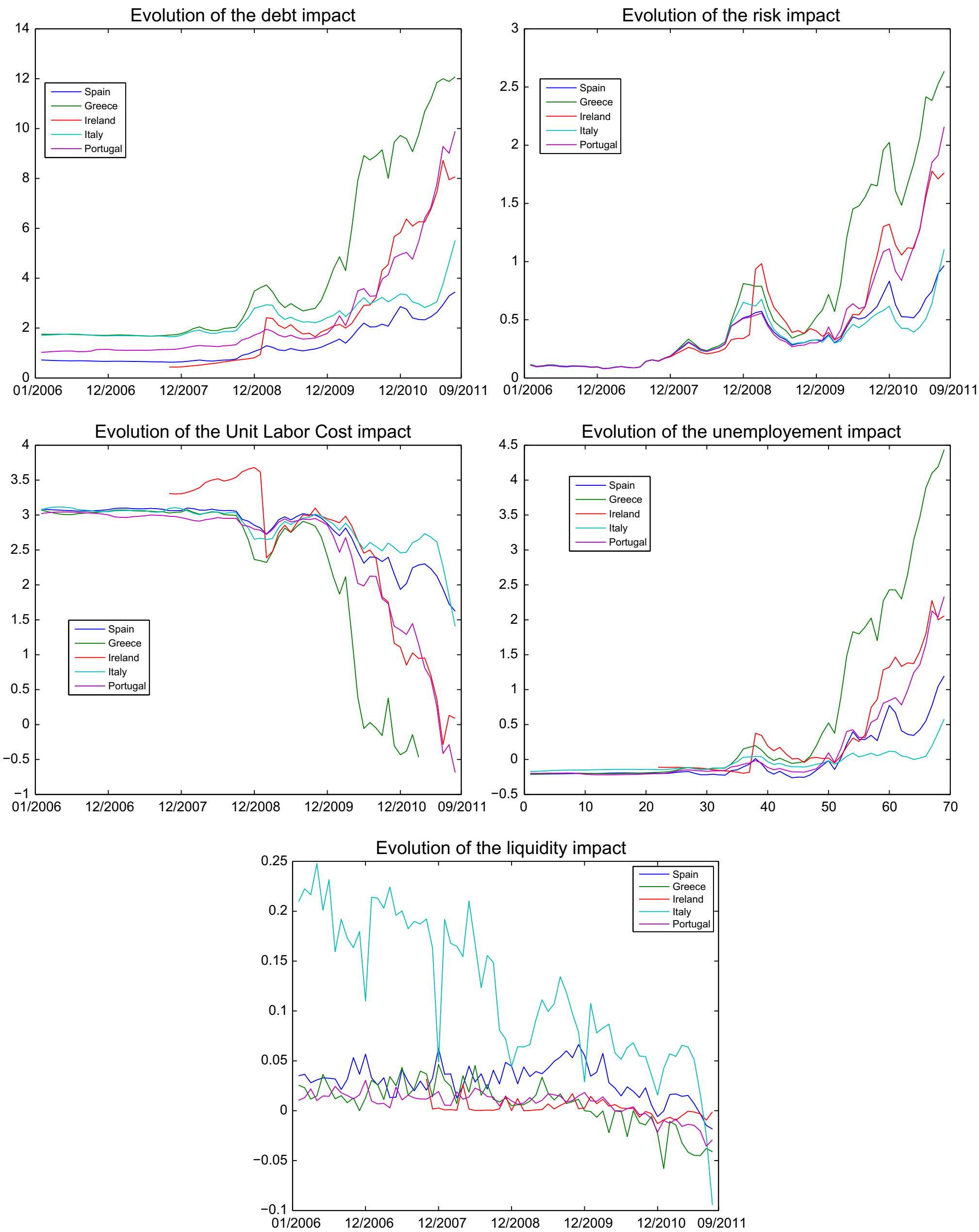
Table 5
Estimation of the probability of default with a PSTR model (cubic transformation).

	Model TV-PSTR		Model 2 Sovereign CDS	
Determinants	β_1	β_2	β_1	β_2
Debt	0.06***	0.22***	-0.03**	0.20***
	(4.97)	(5.55)	(-2.28)	(4.5)
Squared debt	0.00	0.00***	0.00***	0.00***
	(0.35)	(-3.61)	(3.18)	(-4.35)
Unemployement	-0.05***	0.31***	-0.10***	0.34***
	(-3.49)	(7.54)	(-2.91)	(3.55)
Unit labor cost	-0.01	-0.17***	0.05**	-0.06**
	(-0.52)	(-6.41)	(2.31)	(-1.98)
Liquidity	1.88**	-15.47***	1.70*	-4.08
	(2.08)	(-6.66)	(1.84)	(-0.9)
Risk	0.52***	0.56	-0.18	1.40*
	(9.63)	(1.15)	(-0.88)	(1.69)
Smooth parameter γ	0.567		0.003	
Loc parameter	51.7		446.1	
Linearity test	82.0***		275.5***	
RSS	74.3		19.8	
Information crit. BIC	-1.25		-2.57	

Notes: Strut notes: The T-stat in parentheses are corrected for heteroskedasticity. β_1 and β_2 correspond to the coefficient in Eq. (11). β_1 is the coefficient in the first extreme regime. The coefficient in the second extreme regime is $\beta_1 + \beta_2$. The variable debt is with a cublic spline.
* Significant at the 10% level.
** Significant at the 5% level.
*** Significant at the 1% level.

the transition seems rather smooth ($\gamma = 0.002$ with Sovereign CDS). Therefore, to get more accurate picture, we plot the evolution of each estimator multiplied by the variable using the historical values of the threshold variable (for example, $\hat{\beta}'_1 risk + \hat{\beta}'_2 risk * g(q_{it}; \gamma, c)$) (Fig. 2). To interpret the proper evolution of the fiscal situation, we plot the sum of debt and squared debt multiplied by their respective coefficients. For the sake of synthesis and for statistical argument, we do this exercise for the sovereign CDS model only. In fact, this model performs better in rejecting linearity and minimizes the sum of the squared residuals. In sum, this specification appears to best capture the determination model used by investors to price the spread of a country.

We note that sovereign CDS continuously increased during the period. Fig. 2 indicates that the fiscal situation has become more and more influential in the determination of European spreads during the period, a finding that confirms



Note: we plot the evolution of each estimator multiplied by the variable along the historical values of the threshold variable (for example, $\hat{\beta}_1' x_t + \hat{\beta}_2' x_t g(q_{it}; \gamma, c)$.) with x_t is an explanatory variable defined in the text.

Fig. 2. Impact of the determinant factors with a PSTR model.

our time-varying results and the existing results in the recent literature (Haugh et al., 2009; Borgy et al., 2011). In addition, this influence becomes primary at the end of the period. For example, in September 2011, the estimated fiscal situation alone implied a spread equal to 796 bp in Portugal, while it was 951 bp in reality. Fig. 1 also plots the evolution of the coefficients of risk and unemployment. The graphical representation indicates that the influences of unemployment and risk are almost null in the optimistic state and they become very important in the pessimistic state. In particular, the level of unemployment was not priced in the spread before the crisis but it became a significant driver afterwards, which confirms the argument that the business cycle matters to investors. In sum, unemployment adds to the fiscal situation in the macroeconomic variables monitored by investors, a pattern that implies no simple economic resolution of the crisis. In the conclusion we draw regulation implications.

7. Concluding remarks

Here, we have assessed the nature of the European sovereign crisis in the light of a model borrowed from the second generation of currency crises. We estimated the probability of default using panel non-linear estimation methods, the TV-PSTR and PSTR models. Two important objectives were to empirically test the presence of self-fulfilling dynamics and to identify what may have driven the market sentiment during this crisis. In total, our PSTR estimation confirms that the determination model of default probability is not linear, a result that we interpret as evidence of multiple equilibria and self-fulfilling mechanisms during the European crisis. The progressive deterioration of the market sentiment about peripheral sovereigns has been validated by an increase in these countries' spreads. Our main finding is that the sovereign credit derivatives prices affect market sentiment and serve as a coordinating device for speculation.

Delatte et al. (2012) show that the sovereign CDS drive the price discovery process when uncertainty is high, which means that the credit derivative market reveal information first, explaining why market participants focus on CDS to price sovereign risk. The main reason why CDS play this role is due to a growing liquidity of this market in the time of a crisis as suggested by anecdotal evidence that investors take positions on the credit derivative market rather than the underlying market during an episode of stress. In At this light, it is interesting to observe that unconventional monetary measures after September 2012 contributed to balance the liquidity differential and reduce tensions.

These effects are not uncorrelated from the political context as put in the model (fundamentals depend on the beliefs on policy decisions). Coordination failures, the lack of credible macroeconomic rules, the weaknesses of European financial and institutional arrangements have unambiguously paved the way for speculative attacks. Given the difference between political and financial timing, financial markets certainly play an important role in revealing the price of risk. In this sense, our paper emphasizes the disciplining role of markets on fiscal policy while it warns against the destabilizing effects due to multiple equilibria. An important policy conclusion of our paper is that regulation needs to make sure that information conveyed in the CDS is reliable, i.e. not manipulated by too few actors.[21]

Two different regulations of the CDS market have been recently adopted in the European Union. Although they constitute an important step towards safer markets, both regulations suffer from important limits. First the European Union has implemented a ban of uncovered CDS on sovereign entity.[22] From November 2012 investors willing to trade sovereign CDS in a European Union country must hold the underlying bond or a portfolio of assets correlated to the value of the sovereign debt. However two exemptions in the European regulation constitute severe loopholes that seriously mitigate its impact.[23]

The second initiative is a strengthening of the derivatives markets regulatory framework. In June 2012 the European Market Infrastructures Regulation (EMIR) has been adopted with the objective of increasing transparency in the OTC market along similar moves in the United States through the Dodd-Franck act. On the one hand, the implementation of trade repositories provides an effective tool for mitigating the opacity of OTC derivatives markets. On the other hand, in order to improve the resilience of operations, the objective is that 80 percent of all CDS be cleared through a central counterparty. However central clearing obligation affects only new contracts, a fact implying that the transition from the books of the large banks to central counterparty will be dramatically slow. The pace of the reform seems clearly at odds with the emergency situation experienced by the peripheral sovereign cash markets in Europe.

Finally, we hope that the framework presented in this paper opens opportunities for new research. In particular, it would be insightful to relate the volumes traded in the sovereign and banking CDS markets with the nonlinear effects evidenced here. This would constitute a step forward in assessing the plausibility of speculative attacks against sovereigns.

Appendix A. Existence of multiple Sunspot equilibria

At each date t, the probability the investors attribute to default for next period is the sum of the (conditional) default probabilities $F^{(s_t,j)}(\phi_t^{s_t}, \phi_j^{*e})$ in the different states j at date $t+1$, weighted by the corresponding transition probabilities, i.e.:

$$\pi_t(s_t) = \sum_{j=1}^{n} Prob(s_{t+1} = j/s_t) F^{(s_t,j)}(\phi_t^{s_t}, \phi_j^{*e})$$

where ϕ_j^{*e} denote the expected value of the critical threshold in state j. As in JM, we suppose that the partial derivative of each functions $F^{(s_t,j)}$ with respect to $\phi_t^{s_t}$ is negative. This property means that an increase in the current value of the fundamental shifts the conditional cumulative distribution function of the next period fundamental in the same direction. Given these expectations, in each state s_t at date t, the net benefit function of the policymaker is a function of the current value ϕ of $\phi_t^{(s_t)}$:

$$\phi \rightarrow \mathrm{B}(\phi, \pi_t(s_t)) = B(\phi, \sum_{j=1}^{n} Prob(s_{t+1} = j/s_t) F^{(s_t,j)}(\phi, \phi_j^{*e})) \quad (1)$$

[21] We thank an anonymous referee for her comment.

[22] On arguments in favor of a ban see R. Portes, "CDS: useful, misleading, dangerous?" in Vox, April 30, 2012).

[23] For a discussion of the regulation see A.L. Delatte, "The European ban on naked CDS: a fake good idea" in Vox, July 23, 2012

We suppose that the function:

$$(\phi, \pi) \to B(\phi, \pi)$$

is respectively decreasing and increasing with respect to ϕ and π. First, the fundamental phi reflects the sustainability level of the country's economy. If it is high, the state is rather good and the benefit from default is low; second, when the default probability increases, the benefit from default also increases, because the interest rates increase as explained in the text. Thus the function defined in (1) is decreasing in ϕ; indeed, its partial derivative with respect to ϕ has for expression:

$$B_1(\phi, \pi_t(s_t)) + \sum_{j=1}^{n} Prob(s_{t+1} = j/s_t) B_2(\phi, \pi_t(s_t)) F_1^{(s_t,j)}(\phi, \phi_j^{*e})$$

and is strictly negative because $B_1 < 0,\ B_2 > 0$ and $F_1^{(s,j)} < 0$.

Thus the government chooses the unique level of ϕ for which the net benefit is equal to zero. We denote this value by $\phi_{s_t}^* = H_{(s_t)}(\phi_1^{*e}, \ldots, \phi_n^{*e})$.

In this way we define n values ϕ_s^* for the n possible values of s. In a rational expectations equilibrium, each ϕ_s^* should be equal to the expected corresponding threshold ϕ_s^{*e} and the set of these thresholds should therefore satisfy the fixed point equations:

$$\forall s, \phi_s^* = H_{(s)}(\phi_1^*, \ldots, \phi_n^*)$$

We suppose that:

$$\phi_1^* > \ldots > \phi_n^*$$

if state $s = 1$ is better than state $s = 2$ and so on.

Now, the arguments of Jeanne and Masson (2000) apply. The fundamental-based equilibria can be viewed as degenerate cases of the sunspot ones, when the transition probabilities $\text{Prob}(s_{t+1} = j/s_t)$are equal to 1 if $s_{t+1} = s_t$ and 0 otherwise and the F-type functions $F^{(i,j)}$ reduce to one unique function F. In that case, the economy never jumps and always remains in its initial state; thus, JM prove that there exists at least one equilibrium and there may be multiple fundamental-based equilibria associated with different thresholds, provided that the function F and the benefit function B have the good properties mentioned above.

Now, let us turn to the sunspot equilibria and remark that the probability that economy shifts to higher states than state 1 in the next period increases investors' default expectations and decreases the corresponding fundamental threshold chosen by the policymaker to a level $\phi_1^* = H_{(1)}(\phi_1^*, \ldots, \phi_n^*) < H(\phi_1^*)$, because the benefit function decreases with the level of the fundamental process. Similarly, the threshold $\phi_n^* = H_{(n)}(\phi_1^*, \ldots, \phi_n^{*n})$ associated with the worst state n has to be higher than $H(\phi_n^*)$. These inequalities can be consistent with the inequality $\phi_1^* > \phi_n^*$ if and only if there are multiple solutions in the case of fundamental-based equilibria with the shape of function H as the one depicted in JM (p. 334) and with $\phi_n^* \in [0, \phi_I]$ and $\phi_1^* \in [\phi_{II}^*, \phi_{III}^*]$.

So provided that the $F^{(i,j)}$ functions on one hand and the functions F and B on the other hand have the good properties expressed before, one can claim that there exist multiple sunspot equilibria.

Appendix B. Linearization of the default probability

First, we specify the fundamental variable as a linear combination of macroeconomic indicators, depending on the underlying state:

$$\forall t, \phi_t^{s_t} = \alpha_{0,s_t} + \alpha'_{s_t} X_t + u_{t,s_t} \tag{2}$$

with X_t denoting a vector of different economic indicators.

Moreover, in the lines of Jeanne and Masson (2000), we suppose that the fundamental processes ϕ_t^s do not deviate too much from their mean values $\overline{\phi^s}$:

$$\forall t, \forall s = 1, 2 \phi_t^s = \overline{\phi^s} + \delta\phi_t^s$$

where $\delta\phi_t^s$ is supposed to be of limited magnitude.

Thus, the default probability specified as previously:

$$\pi_t(s_t) = \sum_{j=1}^{2} Prob(s_{t+1} = j/s_t) F^{(s_t,j)}(\phi_t^{s_t}, \phi_j^*) \tag{3}$$

can be linearized around $\overline{\phi^{(s_t)}}$ as follows:

$$\pi_t(s_t) \approx \sum_{j=1}^{2} Prob(s_{t+1} = j/s_t)[F^{(s_t,j)}(\overline{\phi^{s_t}}, \phi_j^*) + F_1^{(s_t,j)}(\overline{\phi^{s_t}}, \phi_j^*)(\phi_t^{s_t} - \overline{\phi^{s_t}})] + u_{t,s_t}$$

Accordingly, the previous equation can be rewritten as:

$$\pi_t(s_t) \approx \rho_{0,s_t} + \rho'_{s_t} X_t + u_{t,s_t} \tag{4}$$

with c_{s_t} and θ_{s_t} given by:

$$\rho_{0,s_t} = \sum_{j=1}^{2} Prob(s_{t+1} = j/s_t)[F^{(s_t,j)}(\overline{\phi^{(s_t)}}, \phi_j^*) + F_1^{(s_t,j)}(\overline{\phi^{(s_t)}}, \phi_j^*)(\alpha_{0,s_t} - \overline{\phi^{s_t}})] \tag{5}$$

$$\rho_{s_t} = \sum_{j=1}^{2} Prob(s_{t+1} = j/s_t) F_1^{(s_t,j)}(\overline{\phi^{(s_t)}}, \phi_j^*)\alpha_{s_t}$$

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