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Regime-Dependent Sovereign Risk Pricing During the Euro Crisis*

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Abstract

Previous work has documented a greater sensitivity of long-term government bond yields to fundamentals in euro area peripheral countries during the euro crisis, but we know little about the driver(s) of regime switches. Our estimates based on a panel smooth threshold regression model quantify and explain them: (1) investors have penalized a deterioration of fundamentals more strongly from 2010 to 2012; (2) the higher the bank credit risk, measured with the premium on credit derivatives, the higher the extra premium on fundamentals; (3) after ECB President Draghi's speech in July 2012, it took 1 year to restore the noncrisis regime and suppress the extra premium.

JEL classification: E44, F34, G12, H63, C23

Keywords: European sovereign crisis, Panel Smooth Transition Regression Models, CDS indices

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1. Introduction

Financial market participants have a particular taste for locutions that describe the dynamics of asset prices. In 2011, when sovereign spreads for European peripheral countries successively soared, bond market participants asserted the presence of a cliff risk, the point at which a small shift in a bond's value can have a big impact on its price.[1] A similar pattern was emphasized by policymakers (with different terminology) when they complained about growing mistrust on the part of investors, a fact that drove self-reinforcing dynamics.[2] A way to picture these comments is to say that sovereign risk pricing is regime-dependent and subject to threshold effects. It is clear from Figure 1, which plots spreads between 10-year peripheral and German sovereign bonds, that the trend breaks after 2010, a break that is hard to reconcile with the gradual deterioration of economic conditions.[3]

There is an extensive body of research examining sovereign bond prices in the context of the euro crisis, and we have learned several important lessons. First, the massive holding of peripheral sovereign bonds by the European banking sector created a dangerous nexus between sovereigns and banks. It made banks' balance sheets sensitive to sovereign shocks, and this in turn increased pressure on sovereigns, because they were expected to bail out the banks. These feedback loops have been put forward by Gennaioli, Martin, and Rossi (2010), Huizinga and Demirguc-Kunt (2010), Acharya and Steffen (2013), Acharya, Drechsler, and Schnabl (2013), Coimbra (2014), and Gaballo and Zetlin-Jones (2016). Second, there have been liquidity spirals such as the sell-off in Irish bonds in November 2010, driven by an attempt by market participants to regain liquidity after being unable to meet collateral requirements.[4] Liquidity conditions in the euro area did not recover after the subprime crisis, but rather showed a clear drop in liquidity after 2011. But so far, however, we do not know the details: it is unclear by how much these two effects, the sovereign–bank nexus and the liquidity spirals, have affected the peripheral sovereign bond markets and which one has dominated.

The last lesson we have learned: previous empirical work documents a regime switch in the spread determination model for euro area peripheral sovereigns during the crisis. Two different regimes have been described, a crisis and a noncrisis regime, with a higher sensitivity of yields to fundamentals in the crisis regime (Aizenman, Hutchison, and Jinjarak, 2013; Costantini, Fragetta, and Melina, 2014; Afonso, Arghyrou, and Kontonikas, 2015). But this work does not tell us what drove the change in regime.

In this article, we integrate these different pieces by exploring the possibility that the switch to the crisis regime was triggered by the deterioration of the banks' risk, the liquidity spirals, or both: two endogenous mechanisms potentially implying self-amplifying

1 See, for example, "Bond Investors Fear Cliff Risks", *Financial Times* (November 7, 2011).

2 "The Greek Financial Crisis: From Grexit to Grecovery", Speech by Mr George A Provopoulos, Governor of the Bank of Greece, for the Golden Series lecture at the Official Monetary and Financial Institutions Forum (OMFIF), London, February 7, 2014.

3 In Spain, for example, the public debt amounted to less than 60% of GDP even by end of 2009. The Italian primary budget surplus implied that if interest rates had stayed low, only modest fiscal adjustment would have been necessary to service the debt. Unemployment and the trade deficit had been increasing gradually. And Ireland's trade balance had been improving at the time of the crisis.

4 "Irish Bond Yields Leap after Selling Wave", *Financial Times* (November 10, 2010).

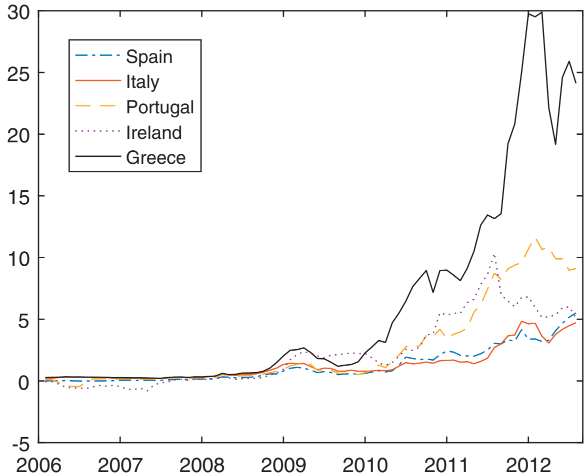

Figure 1. Sovereign spreads. The figure presents the evolution of sovereign spreads variable.

dynamics. We also control for alternative mechanisms, such as the rise of systemic risk in the market and the rise of volatility on several market segments.[5]

These questions require testing for regime-switching dynamics in bond spread determination and investigating the triggers. To do so, we use the smooth transition regression model extended in panel by González, Terasvirta, and van Dijk (2005). Contrary to the alternative family of nonlinear models employed in previous works, the smooth transition regression (STR) model offers a parametric solution to account for nonlinearity by allowing the parameters to change smoothly as a function of an observable variable. We exploit this advantage by taking an off-the-shelf model estimating the impact of economic fundamentals on the spreads of sovereign bonds. We allow the coefficients to change as a function of several measures of risk that might induce regime change. Linearity tests establish a ranking among those hypothetical drivers of regime switch following González, Terasvirta, and van Dijk (2005). We compute our own indicators of risk in the banking sector and of liquidity risk in the euro area by decomposing indicators of systemic risk recently designed by Federal Reserve and European Central Bank researchers (Hakkio and Keeton, 2009; Hollo, Kremer, and Lo Duca, 2012).

In order to work on a homogeneous sample of countries, we focus on the five peripheral member countries which have faced most financial stress during the crisis: Spain, Ireland, Italy, Portugal, and Greece. We start the estimation in 2006 to examine the transition from the noncrisis to the crisis regime and stop right before the spreads decline drastically in July 2012 to document the dynamics specific to the crisis period. We then investigate the reversion mechanisms by extending our period of estimation until March 2014.[6]

A preview of our results is the following. First, sovereign yield spreads became more sensitive to fundamentals between 2010 and 2012; interestingly we do not confirm the finding of Aizenman, Hutchison, and Jinjarak (2013) and Afonso, Arghyrou, and Kontonikas

5 We thank an anonymous referee for this valuable suggestion.

6 Again, we thank the referee for this suggestion.

(2015) of an extra premium on fiscal imbalances for Italy, Spain, and Portugal. In these countries, we find that an extra premium was instead attached to a deterioration of domestic competitiveness on the one hand and of rising uncertainty and risk aversion in global financial markets on the other. Second, the bank–sovereign nexus is the leading driver of nonlinearities, well beyond liquidity spirals and systemic risk. The deterioration of banks' credit risk changed the way investors' price risk of the sovereigns. It exacerbated the effect of initial shocks to the fundamentals. We find that the threshold value of bank credit risk that triggers amplification effects is relatively low. Lastly, we find that the spreads switched back to the noncrisis determination regime during the year following ECB President Mario Draghi's speech in July 2012. In that speech, he asserted the lender of last resort role of the ECB, saying it would do "whatever it takes" to safeguard the monetary union.

Our work complements earlier research on sovereign credit risk during the euro crisis (Attinasi, Checherita, and Nickel, 2009; Dieckman and Planck, 2012; Ang and Longstaff, 2013; Acharya, Drechsler, and Schnabl, 2013; Avino and Cotter, 2014). Technically our work imposes fewer constraints than previous work on the functional form of nonlinearities and allows parameters to change smoothly as a function of an observable variable. The innovations here are therefore the identification of the amplification mechanisms; pinpointing the bank–sovereign nexus working through aggregate credit risk for financial names; quantifying the resulting change in the relative weight of the determinants; and documenting the reversion process after the crisis. More generally, documenting nonlinear dynamics in asset pricing during a crisis episode should contribute to a better understanding of drivers of financial instability.

The remainder of this article is organized as follows. Section 2 reviews the abundant literature on sovereign bond pricing during the euro crisis in order to specify our contribution. Section 3 introduces the panel smooth transition regression (PSTR) specification methodology. Section 4 summarizes our dataset, and Section 5 discusses the estimation results. Section 6 concludes.

2. Sovereign Risk Pricing: What Have We Learned?

Substantial research has examined the sovereign bond price in the context of the euro crisis. On the one hand, there is a consensus that a sovereign–bank nexus generated feedback loops in the dynamics of government bond spreads during the crisis: the deterioration of the sovereign's creditworthiness fed back onto the financial sector, reducing the value of its guarantees and existing bond holdings and increasing its sensitivity to future sovereign shocks. On the other hand, bank risk affects the sovereigns, which are expected to bail out systemically important institutions that represents a significant risk given the size of banks compared to the size of the public backstop (Acharya, Drechsler, and Schnabl, 2013). A theoretical paper suggestive for our empirical investigation is by Coimbra (2014), who shows how the initial shock is exacerbated and feeds back to credit conditions. After a rise in sovereign risk, the banks' VaR constraint binds, which reduces their demand for sovereign bonds, thereby raising the sovereign risk premium. This in turn leads to adverse sovereign debt dynamics, which raise sovereign risk.

Attinasi, Checherita, and Nickel (2009) empirically confirm the effect of the bank–sovereign nexus in a model of government bond yield spreads (over Germany) of ten European countries. They find that government bond yield spreads are significantly affected by the announcements of bank rescue packages in addition to standard measures of government creditworthiness. Acharya and Steffen (2013) find that credit default swap (CDS) spreads of banks and those of governments tend to move more closely together after the announcement of financial sector bailouts.[7] But these papers assume a linear relationship between bank credit risk and government yield. We find it more realistic to relax the linearity assumption to account for self-reinforcing dynamics in the feedback loop.

Liquidity spirals during the euro crisis may have amplified the effect of initial shocks. More precisely, liquidity spirals occur when an initial shock on sovereign bonds degrades the quality of collateral. This forces banks to sell off bonds to regain liquidity or restore their capital ratio, reinforcing the initial downgrading. In addition to the example of the Irish bond sell-off mentioned in the introduction, we have the spiral on the Italian sovereign bond market documented by Pelizzon *et al.* (2015). They find threshold effects in the dynamic relationship between changes in Italian sovereign credit risk and liquidity: there is a structural change in this relationship above 500 basis points (bp) in the sovereign Italian CDS spread, because of changes in collateral and margins for Italian bonds. Brunnermeier and Pedersen (2009) have theoretically modeled liquidity spirals[8]: debt pricing becomes more "information sensitive" and safe assets become less safe, so investors are more selective about the quality of assets they accept as collateral. Their demand for the sovereign bonds perceived to be more risky declines, thereby raising the sovereign risk premium. So there is a liquidity spiral: a falling sovereign bond market leads financial intermediaries to fly to liquidity, and this amplifies the effects of the initial price reduction. Relatively small shocks can cause liquidity suddenly to dry up, leading to a major correction of asset prices.

We have learned, therefore, that banking credit risk and liquidity deterioration affected sovereign credit risk during the euro crisis. In addition, theoretical models point to endogenous amplification effects. Consequently, handling these variables as extra regressors in the sovereign risk pricing model is misleading. Our work tests the hypotheses that the deterioration of banking risk and liquidity shocks has had self-reinforcing effects on sovereign pricing. Before proceeding, we conclude the literature review by examining existing evidence of nonlinearities in the euro area sovereign bond spread.

Several empirical papers find a regime switch in the spread determination model for euro area peripheral sovereigns during the crisis (Gerlach, Schulz, and Wolff, 2010; Borgy *et al.*, 2011; Favero and Missale, 2012; Aizenman, Hutchison, and Jinjarak, 2013; Costantini, Fragetta, and Melina, 2014; Montfort and Renne, 2014; Afonso, Arghyrou, and Kontonikas, 2015). But these papers are silent on what triggered these changes. We go beyond them by testing two channels that may have changed the relative influence of variables determining spreads and thereby triggered the amplification mechanism we describe. Our empirical strategy allows the estimated coefficient of the spread determinants to change as a function of an observable variable.

7 Several papers have focused on the opposite direction of the feedback loop: Acharya and Steffen (2013) find that the Eurozone banks actively engaged in a "carry trade" in the crisis period, increasing their exposure to risky sovereign debt. Gennaioli, Martin, and Rossi (2010) argue that the sovereign risk affects the banks through their exposure to sovereign bonds. Huizinga and Demirguc-Kunt (2010) provide evidence in a large cross-country sample that bank CDS spreads responded negatively to the deterioration of government finances in 2007–08.

8 Stiglitz (1982) and Geanakoplos and Polemarchakis (1986) initially pointed out this externality.

3. Empirical Strategy

Previous work neither explains nor quantifies the mechanism driving the regime change in the sovereign bond pricing. We use a STR model, in order to model the transition process with an observable variable. The PSTR model can be thought of as a regime-switching model that allows a continuum of regimes bounded by two extreme regimes (Fouquau, Hurlin, and Rabaud, 2008). Each intermediate regime is characterized by a different value of the threshold variable and the shape of the transition function. We compare the effect of different potential channels of amplification. With linearity tests we identify the predominant driver of regime shift. We quantify this shift by estimating the coefficients in both extreme regimes.

To model the regime switching and provide an economic interpretation, we use a parametric specification. More precisely, we employ a PSTR model developed by González, Terasvirta, and van Dijk (2005). The choice of panel data is motivated by the low temporal dimension of macroeconomic data. The PSTR model allows us to characterize nonlinearity as a function of an observable variable. The sovereign spread S_{it} is estimated as follows:

$$S_{it} = \mu_i + \beta_1' X_{it} + \beta_2' X_{it} g(q_{it}; \gamma, c) + u_{it} \tag{1}$$

for countries $i = 1, \ldots, N$ and $t = 1, \ldots, T$. Here μ_i represents individual fixed effects, X_{it} is a set of variables that capture credit risk, liquidity risk, and international risk aversion, γ the smooth parameter, c the location parameter defined below, and u_{it} are i.i.d. errors. The transition function $g(.)$ is continuous and bounded between 0 and 1. This specification requires making an assumption on the functional form of $g(.)$. González, Terasvirta, and van Dijk (2005) design their empirical framework with a logistic function of order 1 or 2. We use a logistic function of order 1 that has an S shape and is used in most empirical work[9]:

$$g(q_{it}; \gamma, c) = \frac{1}{1 + \exp\left[-\gamma(q_{it} - c)\right]}, \gamma > 0. \tag{2}$$

where q_{it} is the observable threshold variable. The parameter γ determines the smoothness, that is, the speed at which the vector of coefficients goes from ${\beta_1}'$ to ${\beta_1}' + {\beta_2}'$; the higher the value of the parameter, the faster the transition. The location parameter c shows the inflection point of the transition, that is, the threshold value at which the regime shifts. Thus, the regime switching depends not only on the choice of the transition form but also on the estimated parameters. In order to get an accurate grasp of the pricing evolution during the crisis period, we will plot $g(.)$, the combination of q_{it}, γ, and c to show for every date in which regime applies, this regime being potentially an intermediate regime.

The estimation procedure is reported in Appendix A.

4. Data Description

The estimation of the model in Equation (1) is subject to two major data constraints. Frequency mismatch is the first constraint: while macroeconomic fundamentals have a low frequency (annual, quarterly, or monthly), financial data have a high frequency. We

9 In our estimates, the information criteria indicate that a logistic function of order 1 fits the data better than a function of order 2. The results of Table 4 with logistic of order 2 are available upon request.

therefore transform all series to monthly data. The number of observations of the transition process is the second constraint. In fact, the sovereign crisis started in late 2009, and the Outright Monetary Transactions (OMT) program implemented in September 2012 successfully narrowed the spreads when it was announced in July 2012. So we have only 3 years during which the hypothesized transition might have occurred. Therefore, to obtain a sufficient number of observations, our estimation is based on a balanced panel of the five peripheral Eurozone countries in which the sovereign yield was under pressure between January 2006 and July 2012: Greece, Ireland, Italy, Spain, and Portugal. Subsequently, in order to test the robustness of our findings, we extend our estimates up until March 2014.

4.1 Determinants of the Sovereign Bond Spread

Our dependent variable is the long-term government bond spread, defined as the difference between country *i*'s government bond yield and the risk-free rate of the same maturity. For each country in the sample, we use the long-term German yield as the risk-free rate for the euro area (Dunne, Moore, and Portes, 2007), and the government yield of this country at the same maturity as the German yields. We use daily observations of 10-year bond yields provided by Bloomberg, from which we compute a monthly average.[10] The descriptive statistics of our variables are presented in Table I.

A key choice is the set of explanatory variables included in X_t in Equation (1). The government bond yield spread represents the risk premium paid by governments relative to the benchmark government bond.[11] From a theoretical perspective, these instruments can be priced by decomposing the risk premium into credit risk and liquidity risk.[12] Credit risk is influenced by variables that affect the sustainability of the debt and the ability and willingness to repay. For a sovereign entity, these are macroeconomic variables determining internal and external balances, that is, the budget deficit and the current account. The empirical evidence in the euro area context suggests that significant determinants include fiscal variables, activity-related and competitiveness-related variables (see Attinasi, Checherita, and Nickel, 2009; Haugh, Ollivaud, and Turner, 2009; De Grauwe and Ji, 2013). Liquidity risk is related to the size of the issuer, with an expected negative relationship due to larger transaction costs in small markets. In contrast with findings on credit risk, empirical evidence is mixed about the pricing of a liquidity premium in the sovereign bond spread.[13] Beyond these two theoretical risk premia, Longstaff *et al.* (2011) find that a large component of sovereign credit risk is linked to global factors, while Ang and Longstaff (2013) find that the systemic default risk of European countries is highly correlated with financial market variables. In total, we draw on the previous research mentioned above to test a large range of macroeconomic and financial determinants.

10 For Ireland, only 8-year bond yields are available, so we computed the spread using the 8-year German yield.

11 Early and influential empirical papers include Edwards (1986), Eichengreen and Portes (1989), Cantor and Packer (1996).

12 For countries in the euro area, most of the government bonds are held by euro area investors, so we can ignore foreign exchange risk. Recall also that our spread variable is the spread over the euro-denominated bund.

13 For example, Geyer, Kossmeier, and Pichler (2004) find that liquidity plays a minor role for the pricing of EMU government yield spreads. Favero, Pagano, and von Thadden (2010) find that investors value liquidity, but they value it less when risk increases.

Table I. Descriptive statistics

This table presents the descriptive statistics of the sovereign spreads and explanatory variables. Out. Issues: Outstanding euro-denominated long-term government securities issued in the eurozone. Fisc. Balance: Fiscal Balance. R. Eff. Exch. Rate: Real effective Exchange Rate. Unconv Monetary: Unconventional Monetary Policy. Manuf. Prod.: Manufacturing production index. Hous. Permits: New housing permits. Industry: Industrial production index.

	Spread	Debt	Fisc. Balance	R. Eff. Exch rate	VIX	Bid–ask
Mean	2.678	85.591	−6.780	100.98	23.086	0.143
Median	0.870	94.626	−5.518	100.59	20.723	0.013
Maximum	29.886	174.882	3.134	115.22	62.254	5.886
Minimum	−0.801	23.159	−34.081	92.95	10.787	0.001
Std. Dev.	4.686	35.237	6.477	3.456	10.624	0.539
Skewness	3.364	−0.042	−1.779	1.277	1.708	7.273
Kurtosis	16.384	2.235	8.495	6.436	6.346	64.823
	Out. Issues	Unconv. Monetary	Manuf. Prod.	Hous. Permits	Industry	Unemployment
Mean	0.082	64.051	105.46	205.72	105.08	11.141
Median	0.047	2.571	100.80	138.43	103.24	9.200
Maximum	0.262	283.61	132.47	907.47	140.99	25.300
Minimum	0.007	0.000	81.110	25.294	51.500	4.300
Std. Dev.	0.089	93.716	11.788	179.72	14.714	4.864
Skewness	1.244	1.275	0.588	2.028	0.060	1.024
Kurtosis	2.893	3.290	2.285	7.231	3.192	3.357

To capture fiscal factors, we include the debt-to-GDP ratio and fiscal balance from Eurostat. The expected signs are positive for debt-to-GDP and negative for fiscal balance because a deterioration of fiscal sustainability increases the sovereign risk; we add the squared value of the debt-to-GDP ratio to capture nonlinear dynamics due to threshold effects of sovereign debt on real growth. These fiscal data are revised data, necessary due to the presence of Greece in the sample, although these are not the data initially observed by market participants. Other relevant variables are economic activity and the country's competitiveness. We proxy economic activity with four variables: Unemployment, the manufacturing index and the new housing permits from Eurostat, and the industrial production index from the International Financial Statistics (IFS). Unemployment has an expected positive sign while the other activity variables are expected to reduce the spread when they increase. The country's competitiveness is proxied with the real effective exchange rate defined as the relative price of domestic to foreign consumer price index from IFS. An increase is an appreciation, hence a deterioration of competitiveness, implying that the expected coefficient is positive. In addition we use the trade balance from Eurostat, which is expected to have a negative coefficient.[14] Second, we include a variable for liquidity risk, proxied by the bid–ask spread measured in the bond yields from Bloomberg; it is expected to have a positive coefficient, because an increase of the bid–ask spread is a deterioration of

14 All data are available at a quarterly frequency, except for unemployment and real exchange rate (monthly) and fiscal deficit (annual).

liquidity. Because the liquidity effects were mixed in previous studies, we also use the country's share of total outstanding euro-denominated long-term government securities issued in the eurozone, from ECB, and expected to have a negative coefficient. We include the CBOE Volatility Index (VIX) from Bloomberg as a measure of international risk aversion, because it is often considered to be the world's premier barometer of investor sentiment and market volatility (e.g., Rey, 2013). The coefficient is expected to be positive.

Lastly, we control for the effect of nonstandard monetary measures adopted by the ECB during the crisis. In May 2010, the ECB decided to start the Securities Markets Program (SMP) with large securities' purchases in order to address tensions in certain market segments.[15] We use the amount of securities held for monetary purposes (divided by 100), reported in the ECB's weekly financial statements.[16]

4.2 Endogenous Drivers of Nonlinearities—Two Hypotheses

We use a set of financial data to capture our two hypotheses: bank–sovereign nexus and liquidity spirals. They represent the set of threshold variables that we will include alternatively in our nonlinear estimations. They are composed of indicators of uncertainty and stress in the banking sector and liquidity risk. In addition to including usual well-known measures of such risk, we decompose the indicator of systemic risk designed by the Kansas City Fed which aggregates risk of different market segments, and we recalculate the individual components measuring banking and liquidity risk with European data (Hakkio and Keeton, 2009). This allows us to obtain twenty-two measures tested in alternative specifications to obtain robust findings. All threshold variables are described in Appendix B, Table B1.

5. Estimation Results

5.1 The Changing Composition of the Yield Spreads over Time

In order to test the linearity assumption and select the optimal threshold variable, we need a single specification for the whole set of threshold variables. Selecting explanatory variables by linear models might not be appropriate, since some variables could be important in a nonlinear way.[17] So we select the common specification using a time-varying PSTR (TV-PSTR) which allows the coefficients to vary with time. It has both advantages of allowing nonlinearity and not imposing a particular observable threshold variable. To proceed, we estimate a TV-PSTR on alternative specifications and select the optimal specification according to information criteria.[18]

15 The SMP was terminated in September 2012 in favor of OMTs in sovereign secondary bond markets.

16 The ECB provided, in December 2011 and March 2012, more than 1 trillion Euros of additional liquidity to the financial system with the very LTROs. Unfortunately, publicly available data are not broken down by country so they are not relevant in our panel estimates.

17 We thank the anonymous referee for this comment.

18 We test the largest possible vector of determinants by simultaneously including several proxies of the same effect (for example we include the real exchange rate and the trade balance together). The only exception is the four alternative proxies for economic activity because of their strong correlation.

Table II. Selection of the optimal specification with a TVPSTR model

This table presents estimations of TV-PSTR model on alternative specifications and the optimal specification is selected according to information criteria. The *t*-statistics in parentheses are corrected for heteroskedasticity. (*) significant at the 10% level; (**) significant at the 5% level; and (***) significant at the 1% level. β_1 and β_2 correspond to the coefficients in Equation (1) . β_1 is the coefficient in the first extreme regime. The coefficient in the second extreme regime is $\beta_1 + \beta_2$.

	Specification 1		Specification 2		Specification 3		Specification 4	
	β_1	β_2	β_1	β_2	β_1	β_2	β_1	β_2
Debt-to-GDP	0.137***	0.353***	0.034**	0.666***	0.030**	0.645***	0.021	0.686***
	(7.78)	(2.61)	(2.21)	(5.29)	(2.04)	(4.78)	(1.32)	(5.09)
Debt-to-GDP^2	−0.001***	0.000	0.000	−0.002***	0.000***	−0.002***	0.000***	−0.002***
	(−6.16)	(−0.52)	(1.44)	(−3.80)	(2.89)	(−3.48)	(3.24)	(−3.71)
Fiscal balance	0.162***	−0.845***	0.096***	−1.649***	0.094***	−1.992***	0.073***	−2.059***
	(4.69)	(−5.10)	(7.13)	(−9.80)	(6.69)	(−10.3)	(4.26)	(−10.21)
Real Effect. Exch Rate	0.230***	−0.395***	0.127***	–0.156	0.121***	−0.573***	0.126***	−0.615***
	(9.28)	(−4.22)	(8.46)	(−1.19)	(8.65)	(−6.40)	(8.34)	(−6.02)
Trade balance	28.39***	−46.93***	24.05***	−47.51***	25.49***	−64.05***	26.58***	−60.03***
	(6.54)	(−9.84)	(7.94)	(−8.34)	(7.98)	(−17.05)	(7.73)	(−13.36)
VIX	0.022***	0.034	0.019***	0.087**	0.018***	0.079	0.019***	0.055
	(5.87)	(0.98)	(7.17)	(2.11)	(7.09)	(1.61)	(6.93)	(1.06)
Bid–ask	4.059*	–2.491	–0.505	5.561***	–0.106	5.075***	0.083	4.584***
	(1.76)	(−0.53)	(−0.56)	(2.99)	(−0.13)	(2.89)	(0.1)	(2.62)
Outstanding issues	−75.76***	−12.52***	0.333	10.78***	–8.300	18.88***	–2.082	14.93**
of LT govt sec	(−3.49)	(−2.84)	(0.02)	(2.46)	(−0.58)	(3.63)	(−0.15)	(2.42)
Unconventional	0.017***	−0.033***	0.013***	−0.031***	0.012***	−0.021**	0.012***	−0.024***
Monetary Policy	(4.72)	(−3.97)	(7.00)	(−3.66)	(6.32)	(−2.51)	(5.41)	(−2.81)
Unemployment	0.054	0.541***	–	–	–	–	–	–
	(1.02)	(3.42)	(–)	(–)	(–)	(–)	(–)	(–)
Manufacturing	–	–	0.005	−0.393***	–	–	–	–
production index	(–)	(–)	(1.1)	(−3.17)	(–)	(–)	(–)	(–)
Industry production	–	–	–	–	0.004*	–0.006	–	–
index	(–)	(–)	(–)	(–)	(1.67)	(−0.21)	(–)	(–)
New housing permits	–	–	–	–	–	–	0.001***	0.031
	(–)	(–)	(–)	(–)	(–)	(–)	(2.67)	(0.8)
Smooth parameter γ	0.072		0.179		0.211		0.221	
Loc parameter c	72.0		72.0		72.0		72.0	
Linearity stat.	222.7***		227.4***		207.7***		208.2***	
RSS	153.9		139.1		149.7		148.0	
Schwarz crit.	−0.549		−0.651		−0.577		−0.588	

The linearity test results reported at the bottom of Table II lead to a strong rejection of the null hypothesis of a linear relationship (estimated Lagrange Multiplier (LM) statistics go from 207 to 227 across the different specifications and p-values are inferior to 1%). It is, therefore, clear that linear models of sovereign spreads are misspecified during this period of estimation. Our specifications yield similar slope parameters (γ is estimated between 0.07 and 0.22), the same inflection date ($c = 72$ corresponds to December 2011 when the

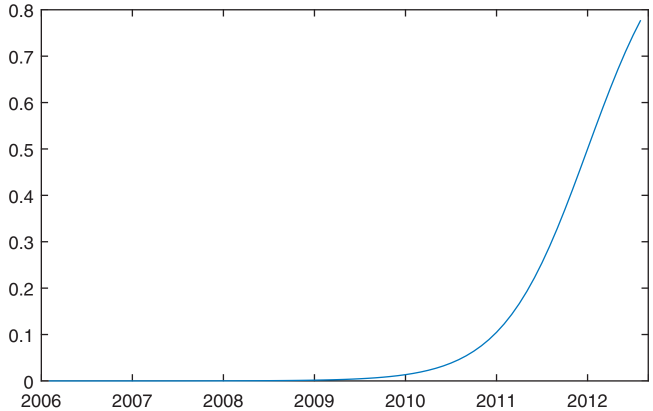

Figure 2. Transition function in the TV-PSTR model.

long-term refinancing operation (LTRO) was launched) and consistent estimated values and signs across different specifications.

Figure 2 which plots the estimated transition function indicates that investors have priced sovereign risk differently during the crisis, and the transition from the noncrisis to the crisis regime has taken 2 years. The information criterion suggests that the second specification including the manufacturing production index is optimal (Schwarz = −0.65). In the following, we focus on this specification to comment on the changing composition of the spread determinants over time.

First, investors price fiscal risk, throughout the period under examination, through the debt-to-GDP ratio and the fiscal balance. In the crisis regime, however, they penalize fiscal imbalances more strongly, attaching an extra premium on the stock of debt ($\hat{\beta}_2 = 0.66$) and the fiscal balance ($\hat{\beta}_2 = -1.65$).[19] Before the crisis, the effect of competitiveness was ambiguous because of the unexpected positive sign of the estimated coefficient. Since the crisis, however, the relationship has become unambiguous: the sign is negative implying that the deterioration of the trade balance is now associated with a higher yield ($\hat{\beta}_2 = -47.51$). Since the crisis, yield spreads increase as a response to a slowdown in economic activity, proxied by the manufacturing production index ($\hat{\beta}_2 = -0.39$). The international risk aversion is statistically significant in explaining spreads before the crisis, but its role becomes critical during the crisis when the relationship between the two variables is multiplied by 5. Liquidity becomes significant only during the crisis, as a higher bid–ask spread is associated with a higher yield spread only during the crisis ($\hat{\beta}_2 = 5.56$).[20] Lastly, as expected, the yield spreads decrease as a response to the OMT program during the crisis.

Overall, we confirm a key finding of previous work: the change of the fundamentals is not sufficient to explain yields over the crisis period, and an increase in the sensitivity to fundamentals and the pricing of new risks are also relevant (Aizenman, Hutchison, and Jinjarak, 2013; Afonso, Arghyrou, and Kontonikas, 2015). So far we have allowed the coefficients to vary over time, but we argue that the regime shift may be endogenous due to

19 The increase is attenuated by the negative coefficient of squared debt $\hat{\beta}_2 = -0.002$. The aggregate sign is, however, positive.

20 This effect is confirmed in two out of four specifications reported in Table II.

Table III. Linearity tests with a PSTR model (specification 2)

The variable that gives rise to the strongest rejection of linearity is chosen as the transition variable. The corresponding LM statistic has an asymptotic $\chi^2(p)$ distribution under H_0. (**) significant at the 5% level; and (***) significant at the 1% level. We have used the specification 2 of Table II.

	H1: Fire-sale liquidation			H2: Bank–sovereign loop	Control
	Flight to liquidity	Flight to quality	Asymmetry information		
Aaa/10-year bond spread	147.3***				
10-year swap spread	110.2***	110.2***			
A/10-year treasury spread	92.10***	92.10***			
High-yield bond/Baa spread	77.6***	77.6***	77.6***		
StockbondsCorr		80.4***			
Cross-section dispersion banks			63.2***		
IVOL bank				144.8***	
Cmax Fi				140.6***	
Euribor-OIS				123.2***	
CDS Snr-Fin				210.8***	
CDS Sub-Fin				177.2***	
I-traxx Europe					120.4***
X-over					84.10***
Hivol					79.3***
Vstoxx					63.2***
RVOL Germ					24.1***
RVOL Nonfin					78.7***
RVOL Pound					54.6***
RVOL Doll					20.3**
RVOL Yen					45.2***
FTSE 300					70.4***
S& P 350					69.6***
Domestic indices returns					26.8***
CISS					79.7***

self-reinforcing dynamics. What are the drivers of regime shift? In the following, we answer by relaxing the linearity assumption again and we allow the coefficients to vary with the different observable variables that capture the bank–sovereign nexus, liquidity risk, and the controls.

5.2 The Prominent Role of the Bank–Sovereign Nexus

The results reported in Table III indicate that the null hypothesis of a linear relationship is strongly rejected regardless of which threshold variable is included in the specification. As mentioned by van Dijk, Terasvirta, and Franses (2002) and González, Terasvirta, and van Dijk (2005), the linearity test can be used to test the null hypothesis of a linear or homogeneous relationship when the threshold variable is known. However, when this variable is theoretically unknown, the linearity test allows to select the best threshold variable among a set of candidate variables. More precisely, González, Terasvirta, and van Dijk (2005, p. 6) indicate that "the test is carried out for a set of candidate transition variables and the variable that gives rise to the strongest rejection of linearity (if any) is chosen as the transition variable".

In Table III, the ranking of the test statistics reveals that four out of the five proxies of the bank–sovereign nexus rank in the top five highest rejection statistics: CDS Snr-Fin, CDS Sub-Fin, IVolBank, and Cmaxi Fi reject linearity with 210.8, 177.2, 144.8, and 140.6, respectively. Only one indicator of liquidity risk ranks in the top five: Aaa/10-year Bund spread gets a rejection statistic of 147.3, while the five alternatives get a significant lower statistics mostly below 100. Similarly, the remaining candidate in the set of threshold variables get much lower rejection statistics (e.g., CISS, the indicator of systemic risk, gets a rejection statistics of 79.7, almost three times lower than the banking CDS index).

We find, therefore, that investors are sensitive to the risk in the banking sector, and this triggers nonlinear dynamics. While the bank–sovereign loop has been documented before, we are the first to give a functional form to the subsequent amplification effects in the government bond pricing. More precisely, the pricing model is a nonlinear function of fundamentals, where the weight of these fundamentals varies with the risk of banks. The deterioration of market conditions for banks changes the way investors' price risk of the sovereigns. We examine the evolution of the estimated coefficient below.

Given the high rejection statistics obtained in every model, we check the robustness of our selection choice using BIC information criteria. While the model with the banking CDS index rejects linearity with the highest statistics, the BIC criterion indicates that the model with the banking stress indicator Cmax Fi is more efficient (Table IV). So in the last step of our empirical investigation, we estimate the two specifications to examine the variation of coefficient loads.

5.3 Heterogeneity in the Sample

The threshold variable Cmax Fi has an individual dimension (i.e., it takes different values across countries, see Figure 3) contrary to the homogeneous CDS Snr-Fin, a feature allowing us to spot heterogeneity in our sample and suggesting two different dynamics across countries. Indeed, the threshold value of Cmax Fi that triggers the regime shift, $c = 0.86$, was never crossed in Italy, Spain, and Portugal, while Ireland and Greece went from the first to the second regime (Figure 3). Therefore, our estimates suggest that their spreads have different dynamics. Gonzalez-Hermosillo and Johnson (2014) point out similar heterogeneous dynamics in the sovereign CDS of the five stressed countries. This finding leads us to split our sample into two subsamples, one including Italy, Spain, and Portugal, the other Greece and Ireland. The smaller subsample still has 162 observations, which is sufficient for reasonably precise and stable estimates.

We reestimate the model in each subsample (Tables V and VI). We obtain a parsimonious specification by adopting a general-to-specific modeling approach, where we eliminate variables based on their statistical significance and the Schwartz information criterion.

5.3.a Italy, Portugal, and Spain

Results in Table V report the transition speed γ, the location parameter c, and the estimated coefficients in regime 1 and regime 2 ($\hat{\beta}_1$ and $\hat{\beta}_1 + \hat{\beta}_2$) in two estimations, one using the

Table IV. Comparing two nonlinear models

This table presents the PSTR estimation for two different threshold variables, CDS Snr-Fin and Cmax Fi. The BIC criterion indicates that the model with the banking stress indicator Cmax Fi is more efficient. We have used the specification 2 of Table II. (***) significant at the 1% level.

Threshold	CDS Snr-Fin	Cmax Fi
Linearity stat	210.7***	140.6***
Smooth parameter	0.928	549.9
Loc parameter	259.1	0.859
RSS	238.9	145.92
Schwarz crit.	−0.110	−0.603

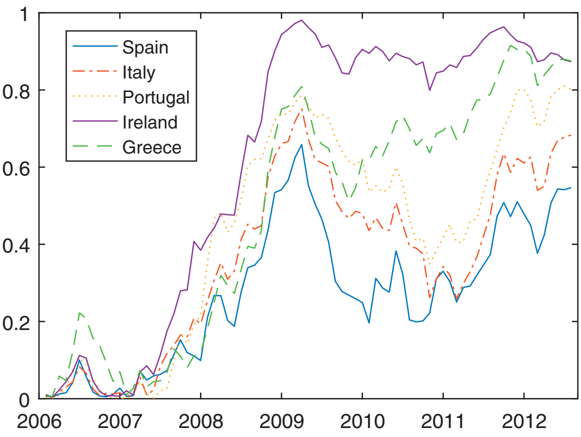

Figure 3. Threshold variable Cmax Fi. The figure presents the evolution of *C*max Fi variable. $\text{Cmax}_t = 1 - \frac{P_t}{\max[P_{t-24} \ldots P_t]}$ with P_t, the domestic banking stock index. The more bearish the market, the closer to 1 the indicator.

optimal threshold variable CDS Snr-Fin and the other using Cmax Fi for robustness check. We comment only on the first estimate. The transition from the first to the second regime is sharp ($\gamma = 95.4$) and the threshold value, c is 130.7 bp. Our model predicts that investors price the sovereign risk differently when the banking CDS index is over 130.7 bp, a value which was crossed in Autumn 2010 shortly after the Greek crisis broke. When we focus on the crisis period, the transition is sharp, which may illustrate the sudden contagion effects. The plot of a sharp transition function does not carry much information, so we focus instead on the numerical evolution of the coefficients.

Estimates confirm the time-varying PSTR result of an increase in the sensitivity to fundamentals. Investors apply an extra premium to competitiveness and international risk aversion ($\hat{\beta}_2 = 0.03$ for real exchange rate and $\hat{\beta}_2 = 0.03$ for the VIX). In turn, the extra premium on fiscal imbalances uncovered in the large sample is much less pronounced in this sub-sample: when we plot the evolution of the weight, we observe that the increase is

Table V. Estimates of the sovereign bond model for Italy, Spain, and Portugal

This table presents the PSTR estimation for two different threshold variables, CDS Snr-Fin and Cmax Fi. The specification 2 of Table II has been used. β_1 and β_2 correspond to the coefficient in Equation (1). 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; and (***) significant at the 1% level.

	CDS Snr-Fin		CMax Fi	
	β_1	β_2	β_1	β_2
Debt – to – GDP	0.064***	−0.068**	0.097***	−0.015***
	(2.56)	(−2.48)	(12.26)	(−4.27)
Debt – to – GDP^2	0.000	0.001***	–	–
	(−1.14)	(3.17)	(–)	(–)
Fiscal balance	0.035***	0.011	−0.042***	0.130***
	(3.72)	(0.16)	(−2.84)	(4.50)
Real Exchange Rate	0.044***	0.034*	0.050***	0.060***
	(2.82)	(1.82)	(3.69)	(4.20)
Trade balance	–	–	−7.444***	10.03***
	(–)	(–)	(−3.57)	(4.19)
VIX	0.014***	0.028***	0.022***	0.006
	(6.73)	(3.67)	(7.52)	(0.96)
Bid–ask	17.72***	−13.19***	4.872***	–0.119
	(3.62)	(−2.68)	(7.29)	(−0.2)
Outstanding stock	–7.045	−9.766***	–	–
	(−0.59)	(−5.38)	(–)	(–)
Unconventional Monetary Policy	–0.003	0.014***	0.004***	0.007***
	(−1.27)	(6.29)	(6.25)	(6.61)
Manufacturing prod. index	−0.008**	–0.015	0.042***	−0.036***
	(−1.97)	(−1.00)	(7.85)	(−3.05)
Smooth parameter γ	95.4		42.2	
Loc parameter c	130.7		0.530	
Linearity stat.	94.6***		79.7***	
RSS	21.6		18.5	
Schwarz crit.	-1.843		-2.053	

very limited.[21] In sum, the market discipline effect works through a higher sensitivity to the countries' perceived competitiveness rather than the fiscal situation. Lastly, the SMP program does not have the expected negative effect on the yield spread.

5.3.b Greece and Ireland

The results of the second subsample including Greece and Ireland reported in Table VI also indicate that the yield spreads have become more sensitive to fundamentals since 2010. Figure 4 plots the smooth transition to the crisis regime. The fact that the transition is smooth and not sharp in this sample may be due to the presence of Greece, the epicenter of the crisis from which contagion effects then spread.

21 The graph is available upon request.

Table VI. Estimates of the sovereign bond model for Greece and Ireland

This table presents the PSTR estimation for two different threshold variables, CDS Snr-Fin and Cmax Fi. The specification 2 of Table II has been used. β_1 and β_2 correspond to the coefficient in Equation (1). 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; and (***) significant at the 1% level.

	CDS Snr-Fin		CMax Fi	
	β_1	β_2	β_1	β_2
Debt-to-GDP	−0.222**	1.08***	−0.123***	0.376***
	(−2.39)	(4.02)	(−4.47)	(5.52)
$Debt-to-GDP^2$	0.000	−0.004***	0.001***	−0.001***
	(0.90)	(−2.75)	(6.86)	(−3.49)
Fiscal balance	0.336***	−0.895***	−0.088**	0.108*
	(3.74)	(−4.32)	(−2.13)	(1.80)
R Effect. Exch Rate	−0.179**	1.304***	0.060**	–0.021
	(−2.14)	(7.88)	(2.06)	(−0.81)
Trade balance	44.63***	−67.13***	28.51***	−43.48***
	(4.53)	(−3.79)	(5.13)	(−6.04)
VIX	0.104	−0.356**	0.016***	–0.032
	(1.50)	(−2.00)	(2.66)	(−1.49)
Bid−Ask	–	–	4.054***	–0.657
	(−)	(−)	(10.63)	(−1.17)
Outstanding stock	–	–	–11.696	−416.2***
	(−)	(−)	(−0.31)	(−4.59)
Unconventional Monetary Policy	–	–	0.029***	−0.048***
	(−)	(−)	(6.41)	(−6.51)
Manufacturing prod. index	0.473***	−1.693***	–	–
	(6.12)	(−6.80)	(−)	(−)
Smooth parameter γ	0.007		438.4	
Loc parameter c	176.5		0.861	
Linearity stat.	132.6***		59.7***	
RSS	241.2		86.9	
Schwarz crit.	1.049		0.186	

Contrary to the previous sample, we find that an extra premium is applied to fiscal imbalances: the coefficient of debt-to-GDP increases in the second regime as well as the absolute value of the coefficient of fiscal balance ($\hat{\beta}_2 \neq 0$). So the higher sensitivity to fiscal imbalances seen in the larger sample was driven by the presence of Greece and Ireland, two counties that have faced fiscal deterioration to a much larger extent than Italy, Spain, and Portugal. In addition, a higher sensitivity is detected for competitiveness (the real effective exchange rate and trade balance have both a higher absolute coefficient in the second regime) and economic activity (manufacturing production index). We do not detect a significant effect of the SMP program in this subsample either.

In total, splitting the sample highlights that an extra premium on fiscal deterioration is applied in Greece and Ireland only. Robustness tests are reported in Appendix C.

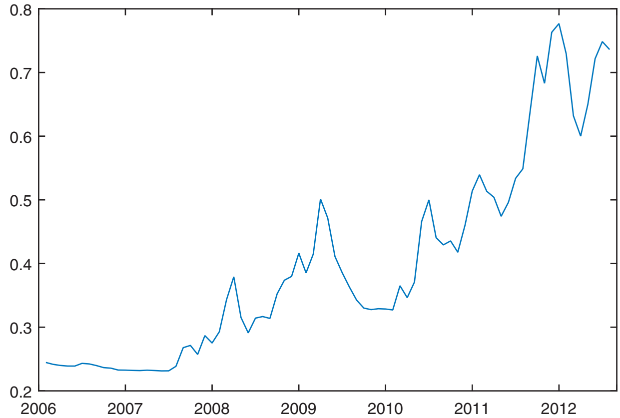

Figure 4. Transition function in Greece and Ireland from 2006 to 2012.

5.4 Dynamics after Draghi's Speech and Macro-Prudential Implications

Our objective in this article was to shed light on the regime shift during the crisis. We start the estimation in 2006 to examine the transition toward the crisis regime and stop right before the spreads decline drastically in July 2012. It is interesting, however, to examine whether our model captures the drastic decline afterward, a sudden decline that cannot be due to the evolution of fundamentals. It may be that the ECB President's commitment to do "whatever it takes" blurred market signals, so that "spreads no longer show us what investors think about debt sustainability" (Paris and Wyplosz, 2013). In our analysis, that would introduce a third regime in government yield pricing after Draghi's speech, where the vector of determinants and their sensitivity change again. Alternatively, one may argue that Draghi's speech tamed market tensions and restored the pricing regime prevailing before the crisis. In that case, we would find that the same endogenous mechanisms operated in reverse. To check, we extend the estimates of our optimal specification in both subsamples up until March 2014, the maximum date with available data.[22]

The takeaway is that the evolution of the coefficient load is very similar to the previous estimation period and the same regime-shifting mechanism operates in reverse. Indeed, Figure 5 of the new transition functions indicates that the model shifts back progressively to the first regime after July 2012. By the end of 2013, the shift was complete with the coefficients back to their precrisis level. The financial CDS index is still a key driver of regime shift (LM statistics is 155 and 137 in each subsample, respectively). The fact that it gets progressively back to its precrisis value drives the shift back to the first regime of coefficients. Our estimates, therefore, show that the reversion to the noncrisis regime was driven by a break of the feedback loop between the sovereign and the banks. It is interesting to observe that it occurred well before macro-prudential measures were

22 There are missing data for the Irish yield after 2012 because liquidity was scarce during the assistance program. In order to bridge the missing data, we mix three different maturities, the 7-, 8-, and 9-year maturity (the longer maturity yield data include all bonds with lower maturity).

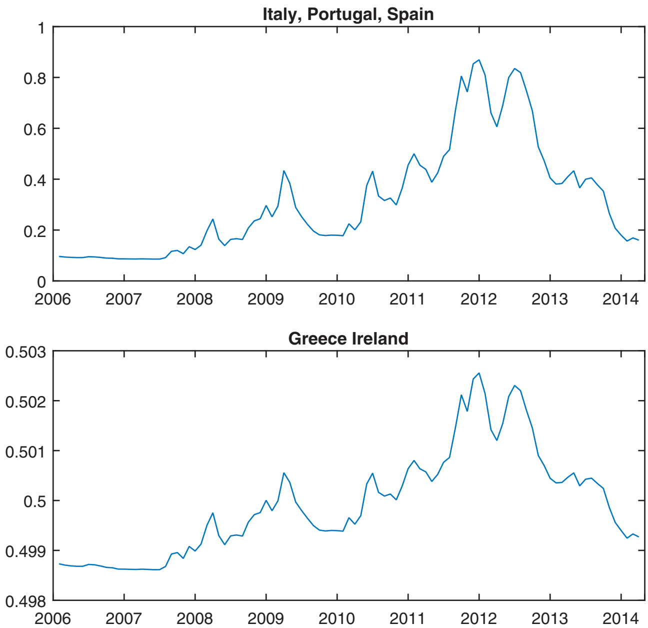

Figure 5. Transition functions from 2006 to 2014.

enforced to address the fragility of the banking sector's balance sheet, including the banking union and the stress tests in Fall 2014. The ECB broke the sovereign–bank nexus and interrupted the feedback loop. This bought time while macro-prudential measures were being implemented.

6. Concluding Remarks

We estimate the sovereign spread of five peripheral members of the euro area using panel nonlinear estimation methods. Our objectives were three-fold: (1) test for nonlinear sovereign bond pricing, (2) discriminate between two potential drivers of nonlinearity, the sovereign–bank nexus and liquidity spirals, and (3) quantify the threshold effects and coefficient regime shifts in order to draw lessons for economic policy.

Our PSTR estimations confirm the previous finding that the changing sensitivity of bond yields to fundamentals is necessary to explain yields during the crisis period (Aizenman, Hutchison, and Jinjarak, 2013; Afonso, Arghyrou, and Kontonikas, 2015). We find that investors then attached an extra premium to competitiveness, international risk, and to a lesser extent liquidity. Contrary to previous studies, we find an extra premium on fiscal imbalances only in Greece and Ireland, not in Italy, Spain, and Portugal. We show that the increasing risk in the banking sector was not only a significant determinant of sovereign risk, but it also amplified the effects of movements in fundamentals. This was a key link in the bank–sovereign nexus. Finally, we find that bond yields returned to their precrisis spread determination regime during the year after Draghi's speech, demonstrating the power of the lender of last resort to stabilize markets. These findings of regime switch and switch back are new, revealed by our estimation method.

There are significant lessons for European regulators and policymakers here: (1) Domestic fiscal discipline and structural reforms could not bring yields down as long as the bank–sovereign feedback loop was not fully addressed. (2) Regime shift was better explained by risk in the banking sector than a general systemic risk indicator. So tracking the financial CDS index would effectively complement the macro-prudential toolkit of policymakers. (3) The individual dynamics were driven by the aggregate banking risk, a risk that the ECB intervention has successfully tamed. So, a more speculative conclusion: (4) Limiting the risk sharing of the ECB operations in the sovereign bond markets as in the asset purchase program announced in January 2015 carries the risk of reigniting tensions.

Beyond the specific eurozone crisis event, our findings may contribute to a better understanding of financial instability, with macro-prudential lessons. The financial price determination models prevailing in normal times may be invalid during crises; the risk pricing of financial assets is fundamentally state-dependent. Our empirical framework gives a simply implementable method to track regime changes and identify the trigger. It is a key to act on it quickly. When the risk trigger is systemic, the central bank can change the state to restore the pricing dynamics, by virtue of its unique role as lender of last resort.

Appendix A: PSTR Estimation

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.

In the González, Terasvirta, and van Dijk (2005) procedure, testing linearity in a PSTR model (Equation (1)) can be done by testing $H_0 : \gamma = 0$ or $H_0 : \beta_0 = \beta_1$. In both cases, the test is nonstandard, 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:

$$S_{it} = \mu_i + \theta_0 X_{it} + \theta_1 X_{it} q_{it} + \epsilon^*_{it}. \tag{A.1}$$

In these auxiliary regressions, parameter θ_1 is proportional to the slope parameter γ of the transition function. Thus, testing linearity against the PSTR simply consists of testing $H_0 : \theta_1 = 0$ in Equation (A.1) for a logistic function with the usual LM test. The corresponding LM statistic has an asymptotic $\chi^2(p)$ distribution under H_0.

Appendix B: Endogenous Drivers of Nonlinearities

Table B1. Definition of threshold variables

Variables	Definition
Liquidity spirals	
Aaa/10-year bund spread	Spread between European corporate bonds rated Aaa and the 10-year German bund. All corporate bond indices are Markit i-boxx European corporate bonds
10-year swap spread	Difference between the fixed rate component of swap and the yield on a 10-year treasury.
High-yield bond/Baa spread	Spread between "junk bonds" and Baa-rated corporate bonds
StockbondsCorr	Three-month rolling correlation between the domestic stock index of each country of our panel and the 10-year bund index. We use the negative values of the correlations.
Cross-section dispersion banks	CAPM regression of the daily return on each bank's stock index against the daily return on the S&P Europe 350 index, using data for the previous 12 months. The estimated coefficients are then used to calculate the forecast errors of the current month. Last we calculate the interquartile range for these residuals in order to keep the central 50%. The lower the interquartile value, the smaller the dispersion across banks. We use daily data on the S&P Europe 350 and the stock prices of the 82 largest commercial banks in terms of market value. The larger the cross-section dispersion, the larger the information asymmetry.
Banking–sovereign nexus	
IVOL bank	Standard deviation of residual returns from a CAPM regression using an aggregate European banking sector price index and the S&P Europe 350. Equivalent of the VIX for the banking industry.
Cmax Fi	$\text{Cmax}_t = 1 - \frac{P_t}{\max[P_{t-24} \ldots P_t]}$ with P the five domestic banking stock indices. The more bearish the market, the closer to 1 the indicator.
Euribor-OIS	The difference between the Euro Interbank Offered rate and the overnight indexed swap rate. This indicator must be taken with some caution because of the alleged manipulation of the Euribor rate.
CDS Snr-Fin	Basket of 25 single CDS covering 25 senior subordination European banks
CDS Sub-Fin	Basket of 25 single CDS covering 25 junior subordination European banks
I-traxx Europe	Most liquid 125 CDS referencing European investment grade credits
X-over	Sub-investment grades names
Hivol	Highest spread non-financial names from iTraxx Europe
Vstoxx	European equivalent of the VIX
RVOL Germ	Realized volatility using the 10-year German government bond index computed as the monthly average of absolute daily rate changes
RVOL Nonfin	Realized volatility of domestic non-financial sector stock market indices

(continued)

Table B1. Continued

Variables	Definition
RVOL Pound	Realized volatility of euro exchange rate against British pound
RVOL Doll	Realized volatility of euro exchange rate against US Dollar
RVOL Yen	Realized volatility of euro exchange rate against Japanese Yen
FTSE 300	Returns of the FTSE 300 stock market indices
S&P 350	Returns of the S&P 350 stock market index
Domestic indices returns	Matrix of the domestic stock returns indices of the five countries in our panel (PSI, IBEX, ATHEX, FTSEMIB, ISEQ)
CISS	Composite Indicator of Systemic Stress of the ECB which aggregates five market-specific subindices (Hollo, Kremer, and Lo Duca, 2012)

Appendix C: Robustness

To check the robustness of our results, we proceed to alternative estimates:

- In the first subsample (including Italy, Spain, and Portugal), overall amplification effects are confirmed when Cmax Fi is used as a threshold variable in an alternative specification reported in Table V. In particular, these estimates confirm that fiscal imbalances are not priced more severely in the crisis.
- Banking CDS and sovereign bonds may price the same information, which would raise an endogeneity bias due to simultaneity. To address this, we reestimate our optimal model by lagging the threshold variable. Linearity is strongly rejected (LM Mj179.9), and amplification effects are confirmed.
- Last, we check that our nonlinearity finding does not result from omitting the financial CDS index as an explanatory variable so that a linear regression would be enough.[23] Our results are not affected by the introduction of the financial CDS index in the vector of determinants (X_{it} in Equation (1)), and its coefficient is not significant. That indicates that this variable drives nonlinear effects in the sovereign bond pricing (LM= 216.8).

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