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Sovereign - Bank Default Risk Linkages During the Greek Financial Crisis and the Role of the Italian Debt

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Abstract

The Greek crisis has brought to light the strong nexus between the credit risks of European banks and their sovereign. We study this phenomenon in Germany, France, Italy and Spain by estimating the conditional correlations between sovereign and bank CDS bond spreads over the period 2006-2015. Trivariate time-varying regime switching correlation analyses, such as the STCC-GARCH and DSTCC-GARCH, are implemented to associate causally the state shifts to the dynamics of the so-called "transition variables". We find evidence of significant changes in the correlation structures due to the evolution of both the Greek and Italian crises.

Keywords: CDS spreads, Greek financial crisis, STCC- and DSTCC-GARCH correlation analysis, Contagion

Jel Classification: E43, E52, F36, C32

1. Introduction

We focus on the time varying correlation between the credit default swap (CDS) spreads of the bonds of major international banks and of sovereign issuers over the period 2006-2015. From here onwards, we use the term nexus to define the link between the default risk of a sovereign issuer and the default risk of banks or the reverse.

Acharya et al. (2015) and De Bruyckere et al. (2013) analyze the feedback loops between sovereign risk and bank risk. The direction of causality can run from bank to sovereign risk in countries with sound public finances and weak banking sector or the other way around, i.e. from sovereign to banks, when an over-indebted public sector jeopardizes the solvency of domestic banks. Gennaioli et al. (2014) investigate the repercussions of a sovereign debt crisis on the banking system and on the real economy via the banks' holdings of sovereign debt (assets side).

Indeed, the unsustainability of public debt affects sovereign creditworthiness and, depending on the exposure of banks' portfolios to government loans, bank's balance sheets. The overall effects, i.e. the strength of the nexus, will depend on the degree of portfolio diversification. (Note 1) It is noteworthy that, in response to the Lehman crisis, governments often provided explicit guarantees to bank bond issuers in order to restore transactions on the wholesale funding market (liabilities side). Recently, Leonello (2018) shows that guarantees link banks and sovereign stability even in absence of banks sovereign's exposure and that under certain conditions a larger size of the guarantee can be beneficial to the nexus as it enhances financial stability.

As is well known a sovereign downgrade will increase the cost of funding, which in turn will affect credit availability and economic growth, determining spillovers on credit quality and thus raising the default probability of banks (Davies and Ng, 2011, Panetta, 2011).

Both sides can be at work at the same time. Podstawski and Velinov (2018) find heterogenous and time varying effects of bank exposure on sovereign credit risk in the Euro area. (Note 2) A destabilizing impact - running from bank exposure to sovereign default risk - characterizes Spain, Italy and Portugal, especially during phases of financial turmoil, whereas a stabilizing effect characterizes the EMU core countries. Gomez-Puig et al. (2018), using three different interconnection measures, find evidence of bidirectional linkages between country-level banking and sovereign risk indicators for Spain and Italy during the European debt crisis. Similarly, Buchholz and Tonzer (2016), study the EU sovereign credit risks interconnections and find that the correlation/contagion structure varies across time and countries, requiring therefore a dynamic approach.

Since the inception of the Greek crisis in 2010, several papers have attempted to address the issue of contagion from the Greek sovereign bonds to the European sovereign markets and national financial systems. This kind of contagion is far from undisputed, however, as the mechanism behind is hazy. Mink and De Haan (2016), for example, find that news about Greek public finance, per se, do not generate abnormal bank stock returns, with the exception of Portuguese, Irish and Spanish banks, while news related to the likelihood of a bail-out affect both the European bank abnormal returns and the sovereign bond prices of Portugal, Ireland and Spain. Similarly, Buchel (2013), analyzing the impact of communications on rescue of indebted countries on PIIGS' CDS and bond yield spreads, find that statements from German, French, EU officials and ECB governing members exert an immediate (asymmetric) influence.

An insignificant economic effect of Greek CDS spread changes on the stock returns of banks of other countries is also detected by Beltratti and Stulz (2017), who find that shocks from larger countries or multiple peripheral countries, instead, have a substantial impact.

According to their analysis the relation between contagion and the holdings of peripheral country bonds by banks from other countries is weak. Similarly, Pradigis et al. (2015), using a corrected Dynamic Conditional Correlation model, and Philippas and Siriopoulos (2013), using a regime switching model and a time varying copula, do not find an overall contagion effect from the Greek crisis to other countries.

Recently, Koutmos (2018) analyzes CDS spreads in the EU to test for contagion and finds not only a heterogenous and time-varying pattern of interdependence but also that the role of Greece, as catalyst for the shocks cannot be statistically proved. (Note 3)

About the nature of the Greek contagion, Gomez-Puig and Sosvilla-Rivero (2016) detect both pure and fundamentals-based bond yield spread co-movements. Arghyrou and Kontonikas (2012), instead, find that evidence of contagion during the European sovereign crisis is restricted mostly to the peripheral countries.

Using a conditional value-at-risk (CoVaR) measure based on copulas and vine copulas, Roboredo and Ugolini (2015) investigate the systemic risk implications of a potential Greek debt default before and after the onset of the financial turmoil. Sovereign debt was found to imply a homogeneous positive systemic risk for domestic financial systems across Europe before the crisis. With the onset of the Greek crisis, however, the systemic impact of sovereign debt increased for countries like Greece, Italy and Portugal, and remained stable or even decreased in other countries.

R-vine copulas have been used by Zhang et al. (2018) to explore the tail dependence between financial stress indicators (including an index of vulnerability of the public and financial sector) of 11 European countries. They find that Spain, Italy, Belgium and France are the most interconnected.

Our investigation deals with the strength and the nature of the nexus between domestic sovereign and bank bond CDS spreads over the decade 2006-2015. The within country analysis is complemented by cross country (contagion) considerations via the transition variables. The domestic nexus correlation structure varies, at first, according to the dynamics of the Greek-German sovereign bond spread, and later on – in the case of Germany and France - also in reaction to the behavior of the Italian-German sovereign bond spread. Analyses of the nexus based on standard techniques, which posit a priori causality linkages, are likely to be distorted during crises since the direction of causality may vary.

We implement thus a correlation analysis procedure and - in order to avoid biases due to volatility shifts (see Forbes and Rigobon, 2002, among many others) - we use a dynamic conditional correlation approach: the smooth transmission constant correlation STCC-GARCH of Berben and Jansen (2005) and Silvennoinen and Teräsvirta (2005) and its two transition functions extension, the DSTCC-GARCH by Silvennoinen and Teräsvirta (2009). These techniques - along with a preliminary probit analysis of the factors that bring about shifts in DCC-GARCH conditional correlations - allow us to identify two transition variables that determine changes in the correlation regime status.

Our paper contributes to, and improves upon, the extant literature on contagion, defined here as a significant increase in the co-movement of the rates of return of the sovereign and bank bonds CDS spreads, in the following ways. First, by focusing on correlation analysis we avoid the indeterminacy of ad hoc causality assumptions. Second, the use of a complex trivariate STCC-GARCH methodology allows to by-pass most of the shortcomings that affect the effectiveness of previous empirical investigations. Indeed, we upgrade the heteroskedasticity consistent procedure of Forbes and Rigobon (2002) by avoiding any ad hoc assumption on the timing of the crisis and the subsequent sub-sample selection bias. Third, we date when correlations among returns increase and, in this way, identify contagion events more precisely. Fourth, the transition variables help us to detect which factor brings about contagion. We avoid in this way the causal indeterminacy of the extreme events/copula contagion analyses. Finally, we assess both the minimal dimension of the shocks required to generate a reaction of CDS investors (and thus a shift in the nexus) and the speed of their reaction, which reflects the relative heterogeneity of their expectations.

The paper is structured as follows. Section 2 presents an overview of the empirical methodology. In Section 3 we focus on the relation between the Greek financial turmoil and the domestic nexuses whereas in Section 4, considering a larger perspective, we include in the analysis the concerns about the sustainability of the Italian public debt as a risk magnifying factor. Finally, Section 5 concludes the paper.

2. Trivariate Parameterizations of Time-Varying Correlations via STCC-GARCH (1, 1) and DSTCC-GARCH (1, 1)

In a time-varying context, conditional correlations are usually estimated with the help of the DCC-GARCH of Engle (1992). The STCC-GARCH implemented in this paper extends this procedure by linking the shifts of the correlations to specific explanatory transition variables.

Consider a 3x1 vector of CDS daily rates of change, with the following conditional mean dynamics

$$\begin{aligned} DsvC_t &= a_{01} + \textstyle\sum_{z=1}^{l^\circ} a_{z1} DsvC_{t-z} + u_{1t} \\ DbkB_{1t} &= a_{02} + \textstyle\sum_{z=1}^{h^\circ} a_{z2} DbkB_{1t-z} + u_{2t} \\ DbkB_{2t} &= a_{03} + \textstyle\sum_{z=1}^{q^\circ} a_{z3} DbkB_{2t-z} + u_{3t} \end{aligned} \tag{1}$$

$DsvC_t$ is the rate of change of a sovereign bond CDS, where C is a country index, and $DbkB_{it}, i = 1, 2,$ is the rate of change of a bank bond CDS, where B_i denotes a domestic bank. u_t is a 3x1 vector of residuals $(u_{1t} \;\; u_{2t} \;\; u_{3t})'$ such that

$$u_t | \Psi_{t-1} \sim iid(0, H_t) \tag{2}$$

where Ψ_{t-1} is the relevant information set.

The conditional variance matrix of the residuals has the following time-varying structure

$$H_t = E(u_t u_t' | \Psi_{t-1}) \tag{3}$$

Bollerslev (1990) posits in the CCC-GARCH parameterization that the conditional variance of each residual time series $u_{it}, i = 1, \dots, 3,$ follows a GARCH(1,1) process and that the correlations are constant. The conditional second moments are thus modeled as

$$h_{iit} = \omega_i + \alpha_i u_{it-1}^2 + \beta_i h_{iit-1},\ \ i = 1, \dots, 3 \tag{4}$$

$$h_{ijt} = \rho_{ij}(h_{iit}, h_{jjt})^{0.5},\ \ 1 \le i < j \le 3 \tag{5}$$

Denoting D_t as a 3x3 diagonal matrix with diagonal elements given by $(h_{iit})^{0.5}$ and Γ as a constant 3x3 correlation matrix, the conditional covariance matrix H_t reads as

$H_t = D_t \Gamma D_t$ and can be rewritten in extended form as

$$\begin{bmatrix} h_{11t} & h_{12t} & h_{13t} \\ h_{21t} & h_{22t} & h_{23t} \\ h_{31t} & h_{32t} & h_{33t} \end{bmatrix} = \begin{bmatrix} h_{11t}^{0.5} & 0 & 0 \\ 0 & h_{22t}^{0.5} & 0 \\ 0 & 0 & h_{33t}^{0.5} \end{bmatrix} \begin{bmatrix} 1 & \rho_{12t} & \rho_{13t} \\ \rho_{21t} & 1 & \rho_{23t} \\ \rho_{31t} & \rho_{32t} & 1 \end{bmatrix} \begin{bmatrix} h_{11t}^{0.5} & 0 & 0 \\ 0 & h_{22t}^{0.5} & 0 \\ 0 & 0 & h_{33t}^{0.5} \end{bmatrix} \tag{6}$$

Berben and Jansen (2005) and Silvennoinen and Teräsvirta (2005) modify the CCC-GARCH model and introduce smoothly time-varying conditional correlations. The latter are assumed to switch over time from one (extreme) constant correlation regime to the other according to the distance from a threshold value of a transition variable. The shifts in turn depend on the dynamics of a continuous logistic function.

In this case, at time t the 3x3 conditional correlation matrix P_t can be written as

$$P_t = (1 - G_t)P_1 + G_t P_2 = \rho_{ijt} = (1 - G_t)\rho_{ij}^1 + G_t \rho_{ij}^2,\ \ 1 \le i < j \le 3 \tag{7}$$

where P_1 and P_2 are assumed to be constant 3x3 positive definite correlation matrices. The logistic function G_t is defined as

$$G_t(x_t; \gamma, c) = \frac{1}{1+exp\{-\gamma(x_{t-d} - c)\}},\ \ \gamma > 0 \tag{8}$$

x_{t-d} is a transition variable with delay *d*. The coefficient γ and the threshold *c* determine, respectively, the speed of adjustment and the location of the transition between the two regimes. P_t is, indeed, a mixture of the two correlation matrices P_1 and P_2. When $(x_{t-d} - c)$ is large and positive, G_t is close to 1 and P_t nears P_2, and when $(x_{t-d} - c)$ is large and negative, G_t is close to 0, and P_t nears P_1.

In the DSTCC-GARCH(1,1) model the conditional correlations vary according to two transition variables. They are parameterized as follows

$$P_t = (1 - G_{1t})P_{1t} + G_{1t}P_{2t},\ \ P_{kt} = (1 - G_{2t})P_{k1} + G_{2t}P_{k2},\ \ k = 1,2 \tag{9}$$

$$\rho_{ijt} = (1\text{-}G_{2t})[(1\text{-}G_{1t})\,\rho_{ij}^{11} + G_{1t}\,\rho_{ij}^{21}] + G_{2t}[(1\text{-}G_{1t})\,\rho_{ij}^{12} + G_{1t}\,\rho_{ij}^{22}],\ \ 1 \le i < j \le 3 \tag{10}$$

with, as transition functions, the logistic functions

$$G_{kt}(x_{kt}; \gamma_k, c_k) = \frac{1}{1+exp\{-\gamma_k(x_{kt-d_k} - c_k)\}},\ \ \gamma_k > 0,\ \ k = 1,2 \tag{11}$$

For each k (k = 1, 2), x_{kt-d_k} are transition variables with delay d_k,. The coefficients γ_k and the thresholds c_k determine, respectively, the speed of adjustment and the location of the transitions between regimes. P_t is thus a (convex) positive definite mixture of four 3x3 positive definite symmetric extreme state correlation matrices P_{11}, P_{12}, P_{21} and P_{22}, with entries ρ_{ij}^{11}, ρ_{ij}^{12}, ρ_{ij}^{21} and ρ_{ij}^{22} , $1 \le i < j \le 3$.

3. The Impact of the Greek Financial Turmoil on Domestic Nexuses

We use a data set of daily observations on sovereign and banks CDS 5 year spreads (the corresponding contract being the most liquid of the CDS market) and on the Greek and Italian sovereign spread, i.e. the differential between the yields of the Greek and Italian 10 year sovereign bonds and of the German 10 year bund. The latter are our selected transition variables.

The panel consists of four large European countries, Germany, France, Italy and Spain, which accounted, in recent years, for 75% of the GDP of the EMU. Their net sovereign debt to GDP ratios differ significantly and range – in 2015 - from 119% in the case of Italy, to 89% for France, 65% for Spain and to 48% in the case of Germany. As we shall see the order of these ratios coincides with the ranking of the severity of the estimated impact of the Greek financial crisis on the national banks - sovereign nexuses.

The graphs of Figure 1 are highly informative. CDS premia vary substantially over the sample period and reflect the shifts in the probabilities of bond potential defaults that are priced by the market. The series in level are not stationary. (Note 4) In France and Germany sovereigns are perceived as substantially less risky than banks whereas in Spain and, especially in Italy, the CDS levels are alike. (Note 5) The sheer dimension of the financial disequilibria hinders, in these countries, any public intervention in favor of distressed banks. The co-movements between CDS spreads on sovereign and bank bonds too change over time, and justify the stochastic correlation approach adopted hereafter.

The statistics set out in Table 1 deal with the rates of change of the CDS spreads on sovereign bonds and on bonds issued by eight major banks, two each for Germany, France, Italy and Spain. They reflect the turbulence of the sample period, since all the time series are strongly serially correlated and are affected by nonlinearities.

Indeed, the BDS test statistics of Brock et al. (1987) strongly reject, with embedding dimension 2, the null hypothesis that the rates of return, filtered for first order serial dependence, are iid. (Analogous results are obtained for the unfiltered returns, and with embedding dimensions varying from 2 to 6.) The standard tests, moreover, suggest that their distributions are non-normal (mostly leptokurtic) and conditionally heteroskedastic. An analysis of the co-movement of these time series requires, therefore, the use of a multivariate GARCH procedure such as the STCC-GARCH(1,1).

In order to corroborate the selection of the transition variables, we perform a probit analysis of first step estimates of DCC-GARCH parameterizations of the conditional correlations of each trivariate system. In every relationship, the dependent variable takes value 1, if the magnitude of the conditional correlation at time t is larger than its average value, and 0 otherwise. The probit specification includes, besides a Lehman dummy, two additional dummies obtained from abnormal shifts of the selected transition variables. The empirical evidence suggests that they have a clear-cut and highly significant impact. (Note 6)

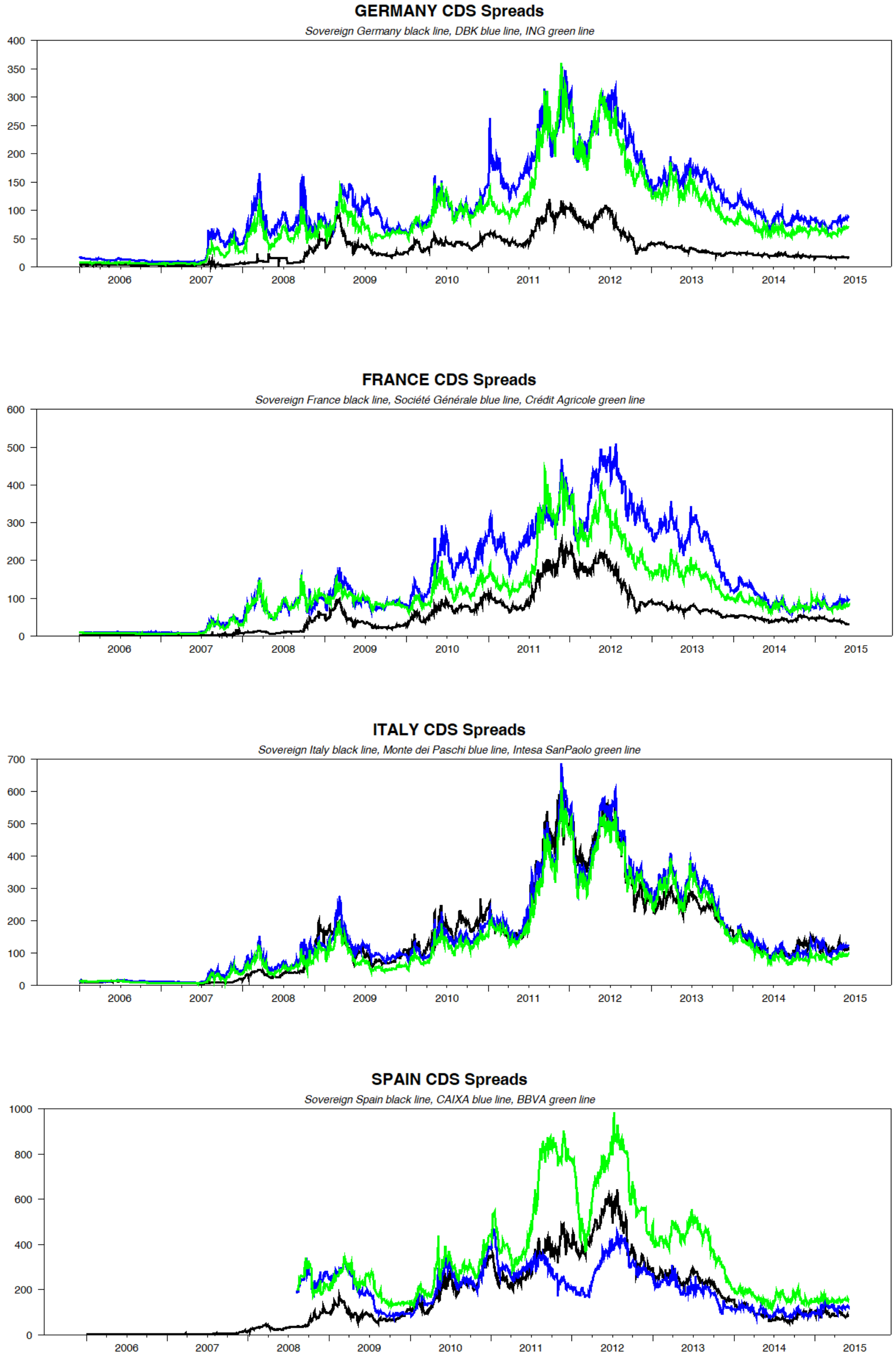

Figure 1. CDS spreads

Table 1. Summary statistics

Variab.	Mean	Std. Dev.	Skew.	Kurt.	JB	AR (1)	AR (5)	ARCH (1)	ARCH (5)	BDS (2)
*Dsv*BD	0.0044	0.1037	10.79	253.54	6473642.0 [0.00]	112.97 [0.00]	122.92 [0.00]	31.24 [0.00]	31.36 [0.00]	12.82 [0.00]
*Dsv*FR	0.0035	0.0847	2.61	39.67	140432.20 [0.00]	166.45 [0.00]	174.01 [0.00]	80.15 [0.00]	363.51 [0.00]	19.474 [0.00]
*Dsv*IT	0.0020	0.0456	1.30	20.41	31715.16 [0.00]	2.33 [0.00]	4.75 [0.00]	53.47 [0.00]	174.21 [0.00]	10.912 [0.00]
*Dsv*SP	0.0024	0.0506	0.39	17.08	20360.29 [0.00]	0.92 [0.00]	12.93 [0.00]	21.93 [0.00]	253.94 [0.00]	8.371 [0.00]
*Dbk*DBK	0.0016	0.0466	2.47	37.60	125102.0 [0.00]	123.19 [0.00]	126.03 [0.00]	243.93 [0.00]	358.93 [0.00]	14.501 [0.00]
*Dbk*ING	0.0019	0.0471	0.65	9.93	5097.04 [0.00]	84.62 [0.00]	90.17 [0.00]	138.82 [0.00]	549.68 [0.00]	9.356 [0.00]
*Dbk*CAG	0.0018	0.0446	0.95	11.74	8191.50 [0.00]	92.46 [0.00]	106.03 [0.00]	133.83 [0.00]	411.47 [0.00]	13.114 [0.00]
*Dbk*SGA	0.0019	0.0451	0.52	9.99	5122.38 [0.00]	136.98 [0.00]	159.17 [0.00]	241.67 [0.00]	669.69 [0.00]	13.842 [0.00]
*Dbk*MPS	0.0020	0.0469	0.98	13.24	11134.27 [0.00]	406.89 [0.00]	420.96 [0.00]	207.76 [0.00]	452.41 [0.00]	15.616 [0.00]
*Dbk*ISP	0.0018	0.0525	1.40	20.05	30564.91 [0.00]	53.06 [0.00]	61.34 [0.00]	91.59 [0.00]	267.48 [0.00]	15.385 [0.00]
*Dbk*CAIXA	0.0010	0.0489	0.82	11.63	5728.03 [0.00]	29.19 [0.00]	34.72 [0.00]	140.61 [0.00]	157.03 [0.00]	8.750 [0.00]
*Dbk*BBVA	0.0011	0.0445	0.23	8.47	2234.47 [0.00]	73.046 [0.00]	113.45 [0.00]	120.48 [0.00]	653.81 [0.00]	14.579 [0.00]
VIX	0.0596	7.2255	1.27	9.35	4791.80 [0.00]	8.43 [0.00]	13.56 [0.00]	34.99 [0.00]	122.61 [0.00]	7.833 [0.00]
GGsp	0.0118	0.2930	0.54	26.72	57734.69 [0.00]	66.58 [0.00]	91.54 [0.00]	171.48 [0.00]	601.38 [0.00]	14.602 [0.00]
ΔGG_{sp}^{+}	0.0754	0.2051	6.04	54.82	322766.51 [0.00]	71.45 [0.00]	120.27 [0.00]	46.12 [0.00]	135.58 [0.00]	5.276 [0.00]
ΔIG_{sp}^{+}	0.0229	0.0522	4.51	29.32	96333.13 [0.00]	3.69 [0.05]	12.61 [0.03]	135.58 [0.00]	3.09 [0.69]	1.485 [0.14]

Notes: *Dsv*C = daily rate of change of the CDS premium on sovereign bonds issued by country C, C = BD, FR, IT and SP; *Dbk*B = daily rate of change of the CDS premium on

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bonds issued by bank B, B = DBK, ING, SGA, CAG, MPS, ISP, CAIXA and BBVA; ΔVIX_t = daily change of the VIX; □□$GGsp$: daily change in the spread between the yields of Greek and German 10 years bonds; ΔGG_{sp}^{+}: positive daily change in the spread between the yields of Greek and German 10 years bonds; ΔIG_{sp}^{+}: positive daily change in the spread between the yields of Italian and German 10 years bonds; Probability values in square brackets; Skew: Skewness; Kurt: Excess Kurtosis; JB: Jarque-Bera normality test; AR(n): Ljung-Box test statistic for n-th order serial correlation of the time series; ARCH(n): Ljung-Box test statistic for n-th order serial correlation of the squared time series; BDS(k): z-test statistic, with embedding dimension k and Є value =.9, of the null that the time series, filtered for a first order autoregressive structure, is independently and identically distributed.

Table 2. System 1

$$P_t = (1 - G_t)P_1 + G_tP_2 = \rho_{ijt} = (1 - G_t)\rho_{ij}^1 + G_t\rho_{ij}^2,\ 1 \le i < j \le 3 \quad (12)$$

$$G_t(x_t; \gamma, c) = \frac{1}{1+exp\{-\gamma(x_{t-d}-c)\}},\ \gamma > 0 \quad (13)$$

	GERMANY	FRANCE	ITALY	SPAIN
Transition Variable	$\Delta GGsp_{t-3}$	$\Delta GGsp_{t-6}$	$\Delta GGsp_{t-7}$	$\Delta GGsp_{t-4}$
Usable data	2006:01:10 - 2015:06:03	2006:01:10 - 2015:06:03	2006:01:10 - 2015:06:03	2008:08:11 - 2015:06:03
	SOV. DBK ING	SOV. CAG SGA	SOV. ISP MPS	SOV. CAIXA BBVA
ρ_{12}^1	0.3911 (36.0656)	0.3014 (25.6189)	0.4398 (15.7684)	0.5478 (14.6712)
ρ_{13}^1	0.41482 (37.9990)	0.3139 (26.7486)	0.41894 (16.8921)	0.1677 (2.2909)
ρ_{32}^1	0.7230 (125.0255)	0.7144 (118.2438)	0.8025 (93.3619)	0.3284 (5.0037)
ρ_{12}^2	0.6866 (30.8745)	0.6420 (22.8153)	0.7053 (27.0543)	0.7350 (7.3253)
ρ_{13}^2	0.7334	0.6349	0.7102	0.8332

	(38.4523)			(26.0420)			(33.3659)			(6.8294)		
ρ_{32}^2	0.8215			0.7463			0.9076			0.9040		
	(66.1891)			(41.4232)			(97.3057)			(23.0761)		
ϒ	30.4523			21.9397			92.6383			2.2127		
	(3.1898)			(3.7756)			(4.2748)			(4.5038)		
C	0.1523			0.1918			0.0531			0.5989		
	(12.8655)			(11.4399)			(4.2246)			(3.1471)		
LLF	15078.8882			14327.5278			15237.4094			10067.8219		
	ε_{1t}	ε_{2t}	ε_{3t}	ε_{1t}	ε_{2t}	ε_{3t}	ε_{1t}	ε_{2t}	ε_{3t}	ε_{1t}	ε_{2t}	ε_{3t}
E(ε_{lt})*	0.017	0.029	0.032	0.032	0.041	0.019	0.021	0.026	0.023	0.014	0.006	-0.021
E(ε_{lt}^2)	1.000	0.996	0.996	1.004	0.994	0.998	0.999	0.996	0.998	1.005	1.010	1.012
ARCH(1)	0.000	0.774	0.373	1.809	0.057	1.081	0.182	0.113	1.088	0.253	1.586	0.031
	[0.988]	[0.379]	[0.541]	[0.179]	[0.811]	[0.298]	[0.670]	[0.737]	[0.297]	[0.615]	[0.208]	[0.859]
ARCH(2)	0.156	0.781	0.582	2.334	0.078	1.125	0.294	0.265	1.089	1.572	1.898	0.457
	[0.925]	[0.677]	[0.747]	[0.311]	[0.962]	[0.570]	[0.863]	[0.876]	[0.580]	[0.456]	[0.387]	[0.796]
ARCH(5)	0.894	4.962	9.832	4.363	7.788	5.407	1.382	2.922	6.974	4.033	4.153	1.234
	[0.971]	[0.421]	[0.080]	[0.498]	[0.168]	[0.368]	[0.926]	[0.712]	[0.223]	[0.545]	[0.528]	[0.942]
JB	1999.3	863.0	1559.4	10409.2	1385.7	392.7	3824.5	1194.4	592.5	418.7	215.5	3830.6
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
BDS(2)	0.587	1.262	1.039	4.225	0.799	1.353	1.796	0.827	-0.585	-0.368	-0.711	-0.095
	[0.557]	[0.207]	[0.299]	[0.000]	[0.424]	[0.176]	[0.072]	[0.408]	[0.559]	[0.712]	[0.477]	[0.924]
BDS(3)	0.985	0.754	1.285	4.138	0.852	1.235	2.351	0.263	-0.567	-0.908	-0.807[	-0.426
	[0.325]	[0.451]	[0.199]	[0.000]	[0.394]	[0.217]	[0.019]	[0.793]	[0.570]	[0.364]	0.419]	[0.670]

Notes. *: $\varepsilon_{it} = u_{it}/h_{iit}^{0.5}, i = 1, 2, 3$; Prob. values in square brackets; JB: Jarque-Bera normality test; ARCH(n): Ljung-Box test statistic for n-th order serial correlation of the squared time series; BDS(k): z-test statistic, with embedding dimension k and and Є value =.9, of the null that the standardized residuals are independently and identically distributed.

Table 2 presents the conditional correlations and the smooth transition parameters of equations (7) and (8), where the transition variable is the Greek-German sovereign 10 year bond spread first difference. Strongly significant from a statistical point of view, they have the appropriate size and the expected sign. (Note 7) The usual misspecification tests performed using the standardized residuals, suggest that the quality of fit is adequate (E(ε_{lt}) = 0, E(ε_{lt}^2) = 1 and ε_{lt} conditionally homoskedastic and serially uncorrelated, for $l = 1, \ldots, 3$).

Indeed, the BDS(2) and BDS(3) test statistics, resulting from BDS tests with embedding dimensions 2 and 3, fail to reject (with one exception only) the null that the standardized residuals are iid. The nonlinearities detected in the return time series of Table 1 are filtered away by the model. However, since the Jarque-Bera statistics systematically reject the null of normality, we compute the estimates using the robust QMLE procedure developed by Bollerslev and Wooldridge (1992).

Table 3. Average dimension and persistence of the conditional correlations over the two regimes

	Number of days in regime 1 (no contagion)	Average value of the cond. correlations between sovereign and domestic banks bonds CDS spreads rates of change in regime 1 (no contagion)	Number of days in regime 2 (contagion)	Average value of the cond. correlations between sovereign and domestic banks bonds CDS spreads rates of change in regime 2 (contagion)	Increase in regime 2 cond. correlations vs. regime 1 cond. correlations. (pct.)	Number of days in regime 1/ number of days in regime 2
Full Sample 2006:01:10 - 2015:06:03						
Germany	2125	0.4127	333	0.6802	64.7983	6.3813
France	2189	0.3197	269	0.6060	89.5981	8.1375
Italy	1836	0.4362	622	0.6956	59.5675	2.9518

Spain	1710	0.4065	69	0.6436	93.6009	24.7826
Pre-Greek crisis period 2006:01:10 - 2010:01:04						
Germany	1030	0.4082	16	0.64785	58.6679	64.3751
France	1037	0.3149	9	0.5484	74.1971	115.2222
Italy	975	0.4351	71	0.6823	56.8788	13.7324
Spain	367	0.4485	0			
Greek crisis period 2010:01:05 - 2015:06:03						
Germany	1095	0.4170	317	0.6811	63.2989	3.4542
France	1152	0.3370	260	0.6066	80.5878	4.4307
Italy	861	0.4850	551	0.6957	44.2452	1.5626
Spain	1343	0.4495	69	0.6436	55.9896	19.4637

Note: For each country, the sovereign-bank bonds CDS correlations are simple averages of the correlations between the rates of change of the spreads of sovereign bonds CDS and the rates of change of the spreads of the CDS of the bonds issued by the corresponding national banks.

To extract additional useful insights, in Table 3 we label as "contagious" the regime in which the transition function is larger than 0.5 and the conditional correlation - the nexus - is closer to its high extreme value P_2 (regime 2) than to its low extreme value P_1(regime 1). The relative number of days spent in each regime and the relative dimension of the corresponding conditional correlations differ among countries.

In Germany and France, the number of days in the "no contagion" regime is from 6 to 7 times larger than the number of days in the "contagion" regime. In the latter, the size of the nexus rises by 65 percent in Germany and by 90 percent in France. In the peripheral countries of the sample, the results are less homogeneous.

The number of days in the "no contagion" regime is only 3 times larger than the number of days in the "contagion" one in Italy and 24 times larger in Spain. In the same way, the nexus in Italy increases by 60 percent in the "contagion" regime, in line with the increases in the core countries of the sample. This is not the case in Spain, where the rise in the nexus is huge (larger than 90 percent). It is noteworthy that the estimates repeated over the Greek-crisis subsample are qualitatively similar to the full sample ones, as the contagion phenomena turn out to occur mostly during the Greek financial turmoil.

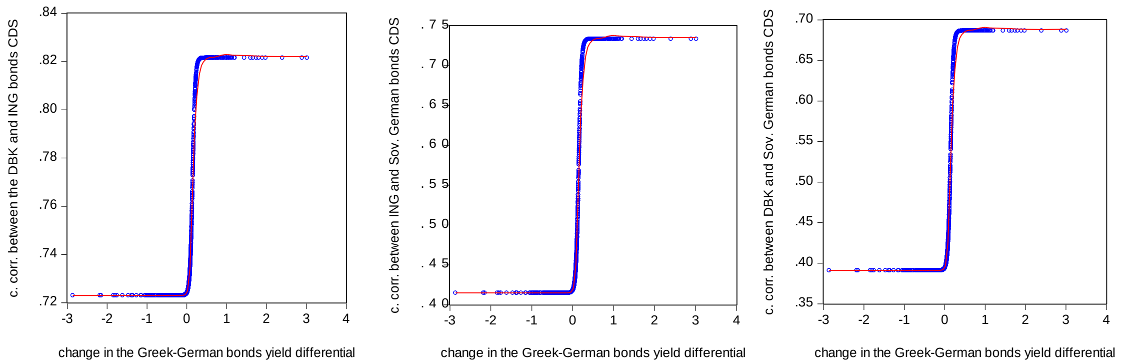
Additional information is provided by the graphs of Figure 2, where the nexuses are related to changes in the differences between Greek and German bond yields (the values of the selected transition variable). The shape of the curves depends upon the values of the coefficient gamma (speed of adjustment variable) and c (threshold parameter) estimates in Table 2. (Note 8) In Germany and France we detect a similar market psychology, since the dynamics of the nexuses (i.e. the conditional correlations between banks and sovereign bond CDS) displays a strong similarity.

This is not the case for Spain and Italy. Agents' reactions are strong and homogenous in Italy whereas they are slow and highly heterogeneous in Spain. In Italy, the nexuses fluctuate frequently and abruptly from one regime to the other, while in Spain they change more slowly and tend to be persistent.

In Italy traders have a common risk perception and react to small variations of the Greek German yield spread, in Spain an opposite behavior holds; the dimension of the public (rather than of banks') debt seems to be the discriminating factor in the risk assessment of bond traders.

The larger the stock of sovereign debt the faster and more homogeneous is the positive shift in the pricing of risk and the smaller the absolute value of the Greek-German sovereign risk differential that triggers it. The differing patterns of reaction, detected in Figure 2, reveal that the focus of the markets is on sovereign financial equilibrium, in line with the major policy recommendations of the European institutions.

Germany



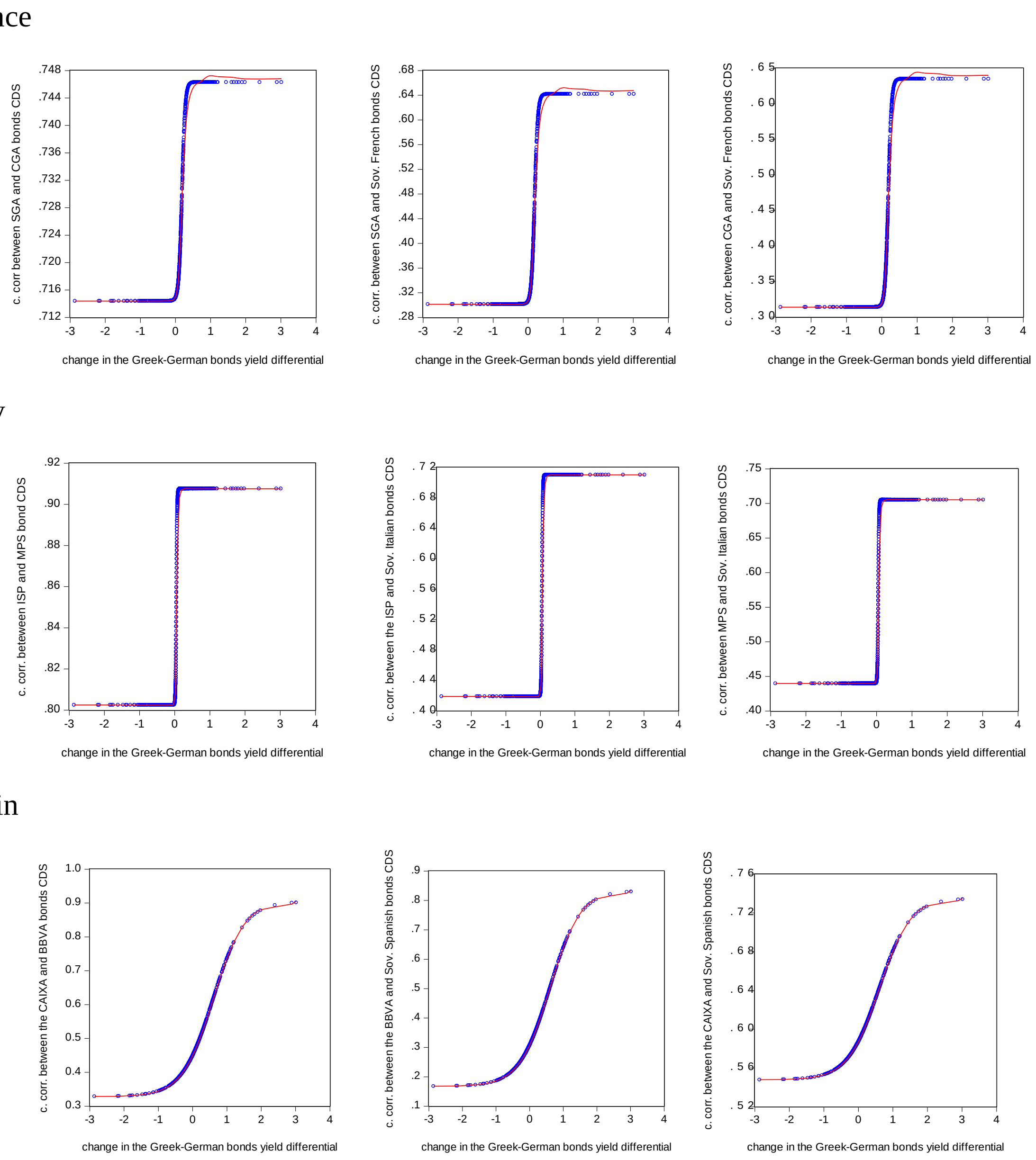


Figure 2. Speed/homogeneity of the reaction of the nexus to shifts of the Greek-German yield differential

4. The Role of the Italian Public Debt

On the basis of the above considerations and taking into account the relevance of the systemic risk channel pointed out by Beltratti and Stultz (2017), who notice that holdings of peripheral country bonds by core banks may not be a statistically and economically significant contagion channel, we extend our analysis introducing a second transition variable, the positive changes of the BTP Bund sovereign bond spread. (Note 9) To give substance to the Italian channel hypothesis, we compute - using data from the BIS quarterly review - the share of the outstanding claims on Greek and Italian official sectors by German and French banks with respect to their total claims on the foreign official sector.

The findings, set out in Table 4, support the view that, after 2011, Greece should not be viewed as the unique source of contagion. The share of the Italian official sector claims reported by French and German banks seems to have magnified the effect of the Greek turmoil; the sheer size of the Italian public debt being able to transform tensions in the Italian sovereign sector into a threat to the survival of the euro area.

Table 4. Claims of German and French banks on the Greek and Italian official sectors as percentage of their respective claims on the overall foreign official sector

GREECE						
	2010	*2011*	*2012*	*2013*	*2014*	*2015*
Germany	5.32	2.60	0.02	0.06	0.07	0.01
France	4.53	1.77	0.01	0.02	0.01	0.00
ITALY						
	2010	*2011*	*2012*	*2013*	*2014*	*2015*
Germany	18.18	16.11	14.73	15.58	14.95	12.88
France	29.96	17.98	27.21	31.33	22.64	22.99

Note: Raw data are obtained from the BIS quarterly report statistics.

In order to account for the role of an additional highly indebted peripheral country, such as Italy, the STCC-GARCH(1,1) model has been extended by adding a second transition variable, the daily positive changes in the difference between the yields of the Italian 10 year BTP and of the German 10 year Bund. The DSTCC-GARCH(1,1) model parameterized by equations (9), (10) and (11) of Section 2 is therefore estimated, where x_{1t} and x_{2t} are, respectively, the positive changes in the differentials between the Greek-German and Italian-German 10 year sovereign bonds yields, ΔGG_t^+ and ΔIG_t^+.

The conditional correlations and the smooth transition parameters are set out in Table 5. The estimates are significant from a statistical point of view and have the appropriate size and the expected sign. (Note 10) The usual misspecification tests, performed using the standardized residuals, suggest that the quality of fit is adequate and the BDS(2) and BDS(3) tests fail to reject (with one exception only) the null that the standardized residuals are iid. The nonlinearities of the return time series of Table 1 are filtered away by the DSTCC-GARCH(1,1) model. The estimates are performed using the robust QMLE procedure developed by Bollerslev and Wooldridge (1992), since the Jarque-Bera statistics systematically reject the null of normality.

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Table 5. System 2

$$P_t = (1 - G_{1t})P_{1t} + G_{1t}P_{2t}, \; P_{kt} = (1 - G_{2t})P_{k1} + G_{2t}P_{k2} \;\; k = 1,2 \tag{14}$$

$$G_t(x_{kt}; \gamma_k, c_k) = \frac{1}{1+exp\{-\gamma_k(x_{kt-d}-c_k)\}}, \; y_k > 0$$

$$\rho_{ijt} = (1 - G_{2t})\left[(1 - G_{1t})\,\rho_{ij}^{11} + G_{1t}\,\rho_{ij}^{21}\right] + G_{2t}\left[(1 - G_{1t})\,\rho_{ij}^{12} + G_{1t}\,\rho_{ij}^{22}\right], \; 1 \le i < j \le 3 \tag{15}$$

	GERMANY	FRANCE
Transition Variables	$\Delta GG_{sp}^{+}{}_{t-3}$ $\Delta IG_{sp}^{+}{}_{t-3}$	$\Delta GG_{sp}^{+}{}_{t-5}$ $\Delta\Delta IG_{sp}^{+}{}_{t-5}$
Usable data	2006:01:10 - 2015:06:03	2006:01:10 - 2015:06:03
	SOV. DBK ING	SOV.CAG SGA
ρ_{12}^{11}	0.3202 (23.3204)	0.0716 (2.9856)
ρ_{13}^{11}	0.3233 (22.0359)	0.0838 (3.5583)
ρ_{32}^{11}	0.6978 (93.3591)	0.6821 (49.0902)
ρ_{12}^{12}	0.5566 (21.4249)	1.0344 (16.8893)
ρ_{13}^{12}	0.6416 (40.0859)	0.7734 (9.0947)
ρ_{32}^{12}	0.7803 (74.0478)	0.8132 (12.5108)
ρ_{12}^{21}	0.6988 (0.0475)	0.4844 (26.6381)
ρ_{13}^{21}	0.7795	0.5056

	(17.2453)			(25.4176)		
ρ_{32}^{21}	0.8399 (33.5343)			0.7464 (84.0399)		
ρ_{12}^{22}	0.7501 (21.4721)			0.6968 (18.2619)		
ρ_{13}^{22}	0.7691 (29.5915)			0.7391 (21.8697)		
ρ_{32}^{22}	0.8167 (39.9912)			0.7062 (25.4551)		
γ_1	10.8468 (7.6162)			25.5783 (4.4065)		
c_1	0.2153 (10.3284)			0.01218 (2.8935)		
γ_2	233.6876 (1.9886)			38.8457 (4.5767)		
c_2	0.0227 (7.6424)			0.1103 (12.7656)		
LLF	15102.597			14342.492		
	ε_{1t}	ε_{2t}	ε_{3t}	ε_{1t}	ε_{2t}	ε_{3t}
E(ε_{lt})*	0.018	0.024	0.026	0.047	0.061	0.047
E(ε_{lt}^2)	0.993	1.005	0.999	1.013	0.988	0.996
ARCH(1)	0.005 [0.942]	0.791 [0.374]	0.282 [0.595]	1.671 [0.196]	0.077 [0.782]	1.138 [0.286]
ARCH(2)	0.108	0.802	0.428	2.168	0.110	1.183

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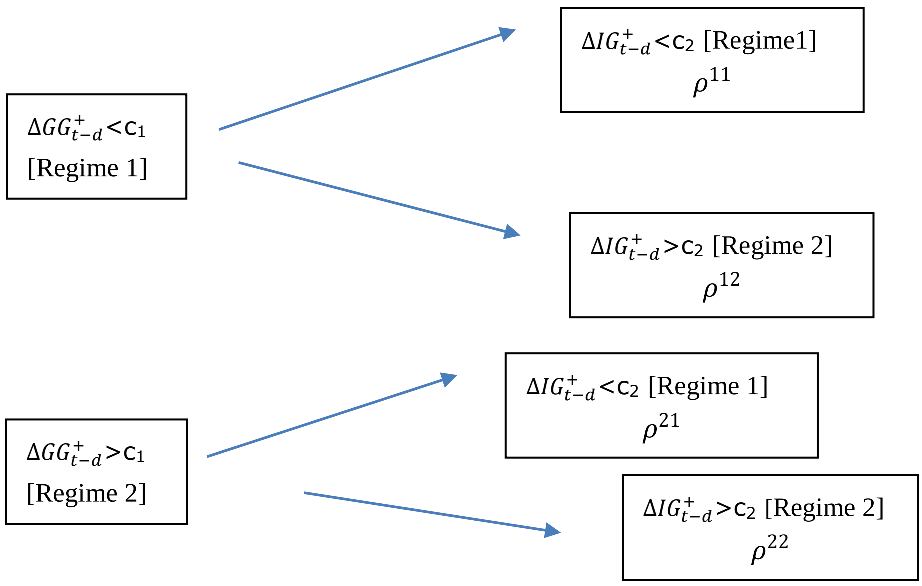
	[0.947]	[0.669]	[0.807]	[0.338]	[0.947]	[0.553]
ARCH(5)	0.908 [0.969]	4.898 [0.428]	8.538 [0.128]	4.185 [0.523]	8.132 [0.149]	5.386 [0.371]
JB	2072.12 [0.000]	1042.08 [0.000]	1137.71 [0.000]	11667.8 [0.000]	1516.66 [0.000]	360.25 [0.000]
BDS(2)	0.483 [0.629]	1.265 [0.206]	0.933 [0.351]	4.049 [0.000]	0.854 [0.393]	1.384 [0.166]
BDS(3)	0.835 [0.404]	0.738 [0.460]	1.153 [0.249]	4.068 [0.000]	0.921 [0.357]	1.273 [0.203]

Notes. *: $\varepsilon_{it} = u_{it}/h_{iit}^{0.5}, i = 1,2,3$; Probability values in square brackets; JB: Jarque-Bera normality test; ARCH(n): Ljung-Box test statistic for n-th order serial correlation of the squared time series; BDS(k): z-test statistic, with embedding dimension k and and Є value =.9, of the null that the standardized residuals are independently and identically distributed.

The interaction of regimes 1 and 2 for the Greek-German and Italian-German transition variables produces the correlation tree sketched below.



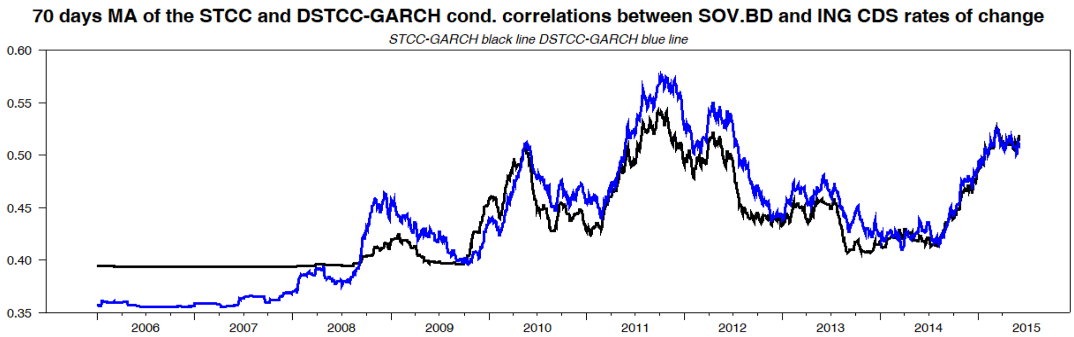
The positive shifts of the Greek-German yield differential determine two regimes according to the positive changes of the sovereign spread being below or above a threshold value c_1, each of which, in turn, can be associated with two regimes generated by positive shifts of the Italian-German yield differentials. We obtain in this way the four regime paths above, where the transitions from one regime to the other are modeled by smooth transmission mechanisms.

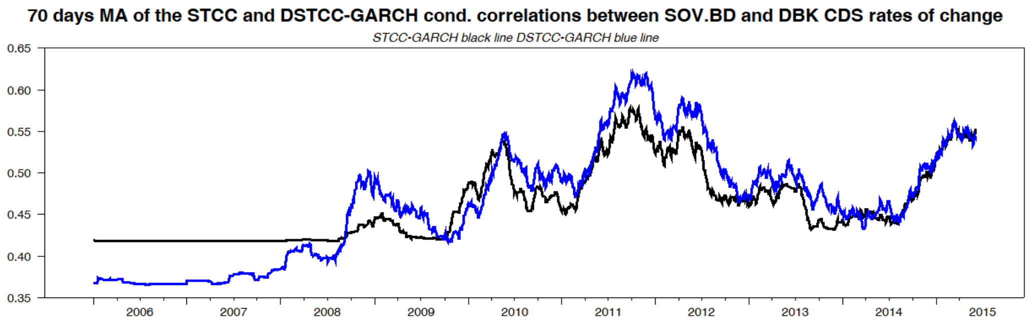
A perusal of the estimates of Table 5 shows that when both transition variables are in regime 1, the nexuses $(\rho_{12}^{11}, \rho_{13}^{11})$ are quite small. They increase substantially when the transition variable associated with the Italian debt is in regime 2 ($\rho_{12}^{12}, \rho_{13}^{12}$). The correlations of Table 2 ($\rho_{12}^{1}, \rho_{13}^{1}$), for Germany and France, thus, are likely to be the by-product of a combination of shocks, which have to be disentangled. We find an analogous result when the Greek-German bonds yield differential is in regime 2. On average, the association with a contagion regime in Italy determines a significant increase in the nexuses ($\rho_{12}^{22}, \rho_{13}^{22}$) both in Germany and France with respect to ρ_{12}^{21} and ρ_{13}^{21}. Here too, the size of the single transition variable correlations estimates of Table 2 ($\rho_{12}^{2}, \rho_{13}^{2}$) seems to be due to multiple causes.

The seventy days centered moving averages of the conditional correlations obtained from historical simulations of the estimates of the STCC and DSTCC -GARCH models of Tables 2 and 5, set out in Figure 3, support the hypothesis of an amplifying effect of the Italian financial stress. In each graph, the blue line – associated with the DSTCC-GARCH model - lies above the black STCC-GARCH line whenever the Italian-German yield differential is large and below it whenever this differential is small even if the absolute value of the shifts of the latter are much smaller than those of the correspond Greek-German bond yield differential. (Note 11)

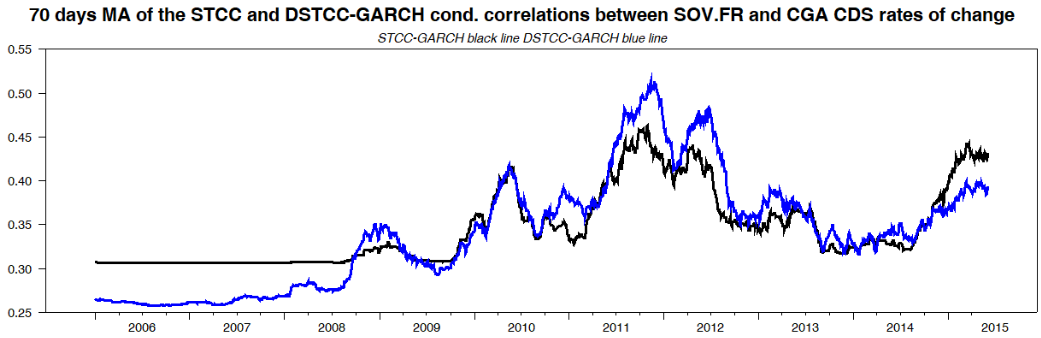
As for the estimated speed of convergence, γ_1 is always lower than γ_2, which suggests that the Italian default risk is likely to trigger a much faster reaction of market agents. As for the threshold values c, they are not comparable since in Table 5 they refer only to positive changes of the Greek and Italian sovereign bond yields spread whereas in Table 2 both positive and negative changes are considered.

Germany





France



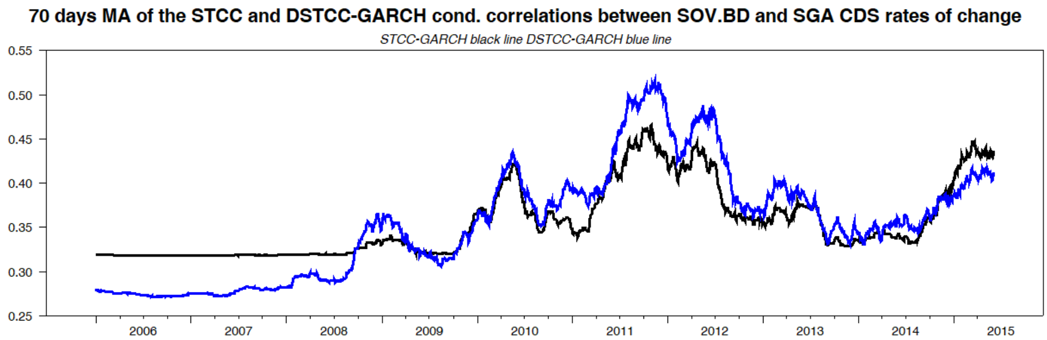


Figure 3. 70 days MA of the STCC and DSTCC GARCH conditional correlations

The results suggest that the two transition variables matter in both core countries, since concomitant positive changes of sovereign bond spreads determine a significant rise in the conditional correlations. In France, however, the behavior of the estimated correlation coefficient through regimes suggests that the role of the Italian debt be dominant, a result in line with the size of the outstanding claims of French banks towards Italian official counterparties (see Table 4).

5. Conclusions

The interconnections between sovereign and banks CDS spreads are highly informative and provide new insights on the financial contagion triggered by the Greek crises. Using the STCC-GARCH methodology we find similar patterns of behavior in core countries (Germany and France) and strong dissimilarities in the so called peripheral ones (Italy and Spain). In the peripheral countries, the main driver of contagion is the perceived default risk of the sovereign issuer, which is linked to the size of the outstanding public debt for a well know debt sustainability issue. Actually, the Italian banks are hit by the Greek turmoil more severely than the Spanish ones.

We extend, therefore, the model introducing a second transition variable related to the Italian public debt. We find that core countries nexuses are affected and increase in a significant way whenever the Italian-German 10 year bond spread change rises above a regime threshold. This highlights the key role of the Italian sovereign debt on the tenability of the EMU project. However, it is the concomitant occurrence of tensions on the Italian and Greek sovereign bond markets that matters. Indeed, positive changes of the Italian spread exert a magnifying effect on the nexuses of core countries, especially in the case of France.

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References

Acharya, V., Drechsler, I., & Schnabl, P. (2015). A Pyrrhic victory? Bank bailouts and sovereign credit risk. *Journal of Finance, 69*, 2689-2739.

Arghyrou, M. G., & Kontonikas, A. (2012). The EMU sovereign-debt crisis: fundamentals, expectations and contagion. *Journal of International Financial Markets, Institutions and Money*, *22*, 658-677.

Beltratti, A., & Stultz, R. M. (2017). How important was contagion through banks during the European sovereign crisis?. *Dice Center WP n. 2017-15.*

Berben, R. P., & Jansen, W. J. (2005). Comovement in international equity markets: a sectoral view. *Journal of International Money and Finance*, *24*, 832-857.

Bollerslev, T. (1990). Modelling the coherence in short-run nominal exchange rates: a multivariate generalized ARCH model. *Review of Economics and Statistics*, *72*, 498-505.

Bollerslev, T., & Wooldridge, J. M. (1992). Quasi-maximum likelihood estimation and inference in dynamic models with time-varying covariances. *Econometric Reviews*, *11*, 143-172.

Brock, W. A., Dechert, W. D., Scheinkman, J. A., & LeBaron, B. (1987). *A test for independence based on the correlation dimension.* Retrieved from www.princeton.edu/~joses/wp/BDS.pdf

Buchel, K. (2013). Do words matter? The impact of communication on the PIIGS CDS and bond yield spreads during Europe's sovereign debt crisis. *European Journal of Political Economy, 32*, 412-431.

Buchholz, M., & Tonzer, L. (2016). Sovereign credit risk co-movements in the Eurozone: simple interdependence or contagion?. *International Finance*, *19*, 246-268.

Cleveland, W. S. (1993). *Visualizing Data.* Hobart Press: Summit, NJ.

Davies, M., Ng, T. (2011, September). The rise of sovereign credit risk: implications for financial stability. *BIS Quarterly Review*, 59-70.

De Bruyckere, V., Gerhardt, M., Schepens, G., & Van der Vennet, R. (2013). Bank/sovereign risk spillovers during the European debt crisis. *Journal of Banking and Finance, 37*, 4793-4809.

De Santis, R., & Stein, M. (2016). Correlation changes between the risk- free rate and sovereign yields of euro area countries. *Working Paper Series #1979*, European Central Bank, Eurosystem.

Engle, R. (2002). Dynamic conditional correlation: a simple class of multivariate generalized autoregressive conditional heteroskedasticity models. *Journal of Business and Economic Statistics*, *20*, 339-350.

Forbes, K. J., & Rigobon, R. (2002). No contagion, only interdependence: measuring stock market comovements. *Journal of Finance*, *57*, 2223-2261.

Gennaioli, N., Martin, A., & Rossi, S. (2014). Sovereign default, domestic banks, and financial institutions. *Journal of Finance, 69*, 819-866.

Gomez-Puig, M., & Sosvilla-Rivero, S. (2016). Causes and hazards of the euro area sovereign debt crisis: pure and fundamentals-based contagion. *Economic Modelling, 56,* 133-147.

Gomez-Puig, M., Singh, M., & Sosvilla-Rivero, S. (2018). The robustness of the sovereign-bank interconnection: evidence from contingent claim analysis. *Research Institute of Applied Economics Working Paper 2018/04*.

Kalbaska, A., & Gatkowski, M. (2012). Eurozone sovereign contagion: evidence from the CDS market (2005-2010). *Journal of Economic Behavior and Organization*, *83*, 657-673.

Koutmos, D. (2018). Interdependence between CDS spreads in the European Union: is Greece the black sheep or black swan?. *Annals of Operations Research*, *266*, 441-498.

Lee, J. L., & Strazicich, M. C. (2003). Minimum LM unit root test with two structural breaks. *Review of Economics and Statistics*, *85*, 1082-1089.

Lee, J. L., & Strazicich, M. C. (2004). Minimum LM unit root test with one structural break. *Appalachian State University Working Paper*.

Leonello, A. (2018). Government guarantees and the two-way feedback between banking and sovereign debt crises. *Journal of Financial Economics*.

Mink, M., & de Haan, J. (2013). Contagion during the Greek sovereign debt crisis. *Journal of International Money and Finance*, *34*, 102-113.

Panetta, F. (2011). The impact of sovereign credit risk on bank funding conditions. *CGFS Papers n. 43*.

Philippas, D., & Siriopoulos, C. (2013). Putting the C into the crisis: contagion, correlation and copulas and EMU bond markets. *Journal of International Financial Markets, Institutions and Money, 27,* 161-176.

Podstawski, M., & Velinov, A. (2018). The state dependent impact of bank exposure on sovereign risk. *Journal of Banking and Finance*, *88*, 63-75.

Pradigis, I. C., Aielli, G. P., Chionis, D., & Schizas, P. (2015). Contagion effects during the financial crisis: evidence from the Greek sovereign bonds markets. *Journal of Financial Stability*, *18*, 127-138.

Reboredo, C., & Ugolini, A. (2015). A vine-copula conditional value-at-risk approach to systemic sovereign debt risk for the financial sector. *North American Journal of Economics and Finance*, *32*, 98-123.

Silvennoinen, A., & Teräsvirta, T. (2005). Multivariate autoregressive conditional heteroskedasticity with smooth transition in conditional correlations. *SSE/EFI Working Paper Series in Economics and Finance No. 577.*

Silvennoinen, A., & Teräsvirta, T. (2009). Modelling multivariate autoregressive conditional heteroskedasticity with the double smooth transition conditional correlation GARCH model. *Journal of Financial Econometrics*, *7*, 373-411.

Zhang, D., Yan, M, & Tsopanakis A. (2018). Financial stress relationship among Euro area countries: A R-Vine copula approach. *The European Journal of Finance.*

Notes

Note 1. It should be noticed that current regulation, which provides for a preferential treatment to euro sovereign securities, has probably reinforced this correlation.

Note 2. They use a Markov switching structural vector autoregressive in a heteroscedastic framework.

Note 3. Supporting the time varying results, Kalbaska and Gatkowski (2012) - who analyze the EU CDS markets before 2011 - find that Greece, Spain and Italy have a lower power to trigger contagion than the core EU countries.

Note 4. We analyze in this paper the CDS on bonds issued by the following banks: Deutsche Bank (DBK) and Ing (ING) for Germany (the latter is Dutch, but no alternative data were available), Société Générale (SGA) and Credit Agricole (CAG) for France, Monte dei Paschi (MPS) and Intesa SanPaolo (ISP) for Italy and Caixa (CAIXA) and Banco Bilbao Vizcaya Argentaria (BBVA) for Spain. The empirical analysis is carried out in terms of rates of change of the spreads since CDS, in levels, are nonstationary. The unit root tests – performed both with the standard ADF tests and the LM tests with unknown structural breaks of Lee and Strazicich (2003, 2004) are available from the authors upon request.

Note 5. It is well known that, beside a pure credit risk, the CDS premia includes a liquidity risk and a systemic/macroeconomic risk (see De Santis and Stein, 2016 page 6). These components explain the large simultaneous volatility shifts and the differences among the premia of Figure1.

Note 6. The estimates are not reported here for lack of space and are available from the authors upon request.

Note 7. The full set of mean and variance equations parameters of the GARCH estimates are not reported here for the sake of parsimony and are available from the authors upon request along with the econometric routines written in RATS.

Note 8. Each graph contains a scatter plot of the conditional correlations between the CDS rates of change and the deviations of the difference between the transition variable (viz. the changes in the Greek German bond yields spread) and the threshold c from zero. We report the former on the vertical axis and the latter on the horizontal one. For the sake of clarity, we have interpolated the scatter plots using local first order polynomial regressions with bandwidth based on the nearest neighbor approach. The local regressions are performed on a sub sample selected according to the Cleveland (1993) procedure and involves about 100 evaluation points. Tricube weights are used in the weighted regressions aimed at minimizing the weighted sum of squared residuals. The bandwidth span of each local regression is set to 0.3.

Note 9. Beltratti and Stultz (Table 2, 2017), find that core country banks' net holding of bonds issued by Greece accounted for 4.99% of banks' market capitalization in 2010 (5.88% if normalized by banks' tangible equity). A year later these percentages were respectively 5.89% and 2.97%. As for bonds issued by Italy, the percentages were 18.45% and 19.69% in 2010, whereas in 2011 the figures rose to 27.15% in terms of market capitalization and fell to 9.19% in terms of tangible equity. The figures for Greece are quite small if compared to the severe turmoil generated by the so called "Greek crisis".

Note 10. The full set of mean and variance equations parameters of the GARCH estimates are not reported here for the sake of parsimony and are available from the authors upon request.

Note 11. The first graph in Figure 3 shows the centered seventy days moving averages of the Greek-German and of the Italian-German log-term bond yield differentials. They provide a visual chronology of the timing of the respective sovereign bond crises along with their size.