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European Government Bond Market Contagion in Turbulent Times*

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Abstract
In this paper we investigate the dynamics of European government bond market contagion during the financial crisis and, subsequently, during the European sovereign debt crisis. Following Bae et al. (2003), we use the coexceedance variable—joint occurrences of extreme negative and positive returns in different countries on a given day—to measure contagion. We also analyze the underlying determinants of the dynamics of contagion using an ordered logistic regression. Our results reveal that interest rates, stock market returns and market volatility help explain contagion in European government bond markets; however, their individual relevance varies from crisis to crisis. We also find that past contagion significantly increases the probability of more episodes of contagion today. Finally, we find statistically significant evidence of contagion from the "old" European Monetary Union (EMU) members to the new members during the sovereign debt crisis and to the non-EMU EU-15 members during both crises. Interestingly, our results show that the new members are those that behave most differently in our analysis.

1. Introduction

Contagion effects across European government bond markets have become a major area of concern for both policymakers and market participants. Policymakers are particularly keen to understand the mechanisms that link these markets in order to be able to maintain financial stability and make effective monetary policy decisions. Likewise, an understanding of bond market linkages can help market participants formulate appropriate risk management strategies and investment decisions. This interest becomes even greater in years of turmoil, when financial markets are hit by extreme shocks.

In the financial literature, there is no general consensus on the exact meaning to be attributed to contagion, while a wide variety of frameworks have been employed in the empirical examination of contagion across markets.[1] European government bond markets are no exception here and, since the onset of the sovereign debt crises, a growing body of literature has focused on analyzing contagion across these markets. For example, in the framework of a structural vector error correction model, de Santis (2012) finds that contagion effects in the euro area are linked to ratings. Defining

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[1] Pericoli and Sbracia's (2003) survey of financial contagion reports the five most common definitions of "financial contagion" and Dungey *et al.* (2005) review alternative methods of testing for the presence of contagion during financial market crises and show how such methods are related in a unified frame--work.

contagion as the change in propagation mechanisms when large shocks occur and using a Bayesian quantile regression model, Caporin *et al.* (2013) report no change in the intensity of the transmission of shocks among European countries during the sovereign debt crisis. A different perspective is given by Metiu (2012), who extends the canonical econometric model of contagion proposed by Pesaran and Pick (2007) and finds evidence for significant contagion effects among long-term bond yield premia after 2008. Beirne and Fratzscher (2013) also report that regional contagion was not particularly evident during the 2008–2011 sovereign debt crisis in Europe and that the cross-country spillovers of sovereign risk were stronger prior to the crisis than during the crisis. Finally, applying the Granger-causality approach and the endogenous breakpoint test, Gómez-Puig and Sosvilla-Rivero (2014) provide clear evidence of contagion in the aftermath of the current euro debt crisis in eleven EMU countries.

This study assesses contagion in European government bond markets during periods of turbulence, when investors and policymakers have a particularly strong interest in knowing whether and, if so, how shocks propagate to other countries. Following Bae *et al.* (2003), we use the coexceedance variable—joint occurrences of extreme negative and positive returns in different countries on a given day—to measure contagion. Bae *et al.* (2003) capture the coincidence of extreme return shocks across countries within a broader region and across regions. They define contagion within regions as the fraction of the coexceedance events that cannot be explained by fundamentals, and contagion across regions as the fraction of coexceedance events that is left unexplained by fundamentals but can be explained by exceedances from other countries.[2] This approach is used by Christiansen and Ranaldo (2009) to analyze the financial integration of the stock markets in the ten new EU member states comprising former communist countries of Eastern and Central Europe as well as the integration of two groups of countries, namely new and old member states. In this paper, we are interested in analyzing contagion in European government bond markets and in determining the underlying determinants of contagion. We test for contagion (i) across all countries in the sample, (ii) within groups of countries and (iii) across groups of countries. To this end, we use an ordered logistic regression model to determine the underlying determinants of the observed dynamics of contagion.

Specifically, we address the following questions: First, which factors are associated with an increase (decrease) in the probability of observing extreme returns across markets? Second, did the effects of these factors change during the financial crisis and, subsequently, during the European sovereign debt crisis? Third, is contagion in European government bond markets driven by global (US) or regional (EMU) factors? And, are there differences across countries?

The main results of this paper can be summarized as follows: First, we find that contagion in European government bond markets can be explained by interest rates, stock market returns and market volatility in both regional and international

[2] Their approach has two advantages. First, contrary to standard correlation measures, it is robust to time-varying volatility and departure from normality. Second, the correlation coefficient is a linear measure, which is inappropriate for analysing nonlinear phenomena, as financial market integration potentially might be (see Baur and Schulze, 2005; Dungey and Martin, 2007).

stock markets. However, the relevance of these factors is heterogeneous and differs between crises. Second, our analysis reveals that contagion has a cluster effect and past contagion significantly increases the probability of more episodes of contagion today. Third, we find statistically significant evidence of contagion from the "old" European Monetary Union members to the new members during the sovereign debt crisis and to the non-EMU EU-15 members during both crises. Finally, we find that the new European Union members are those that behave most differently.

The rest of this paper is organized as follows: In Section 2 we present our data, while in Section 3 we describe the ordered logit model. In Section 4 we examine the determining factors of European government bond market contagion. Finally, we conclude the paper in Section 5.

2. Data

The data consist of the ten-year JPMorgan Government Global Bond Index (JPMGBI), expressed in terms of a common currency, the euro, and the sample includes 16 European countries. Our study focuses on ten EMU EU-15 countries (Austria, Belgium, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal and Spain)[3] and six non-EMU countries (Denmark, the Czech Republic, Hungary, Poland, Sweden and the UK). These bond market indices are transformed into returns by taking the first difference of the natural log of each bond price index. All data were collected from Thomson Datastream.

We use daily data for the period from 7 August 2007 to 15 December 2013; thus our sample covers the recent years of turmoil (initially the financial crisis and, subsequently, the European sovereign debt crisis). We define the starting point of the financial crisis as August 2007 (see Mishkin, 2010), when equity markets initially fell and central banks started intervening to provide liquidity to financial markets. For our analysis, we match the end of the financial crisis with the beginning of the European sovereign debt crisis. As pointed out by Christiansen (2014), dating the European sovereign debt crisis is not a straightforward task, as no official dates are available. Generally, it is considered to have begun in late 2009 (see, for example, Alter and Beyer, 2014; Caporin *et al.*, 2013). Therefore, we define the starting point of the sovereign debt crisis as January 2010 and it persists until the end of our dataset in December 2013.

Following Bae *et al.* (2003), we define an extreme return, or exceedance, as one that lies either below (above) the 5th (95th) quantile of the marginal return distribution. We consider separately the left side of the distribution, or the negative return exceedances ("bottom tail"), and the right side, or the positive return exceedances ("top tail"). Similarly, episodes of coexceedance refer to occurrences of extreme negative and positive returns in different countries on the same day. We define a coexceedance count of i units for negative returns as the joint occurrence of i exceedances of negative returns on a particular day.

As is standard in the literature, we have divided the European countries into three groups: (1) EMU EU-15 old countries, (2) non-EMU new EU countries and (3) non-EMU EU-15 countries. In our study, these groups are composed of the fol-

[3] Finland is not included in the study due to a lack of available data.

Table 1 Summary Statistics of Exceedances

	TOP TAIL				BOTTOM TAIL			
	Financial crisis	Sov debt crisis	*z*-statistic	*P*-value	Financial crisis	Sov debt crisis	*z*-statistic	*P*-value
Panel A: EMU EU-15 old countries								
Austria	39	72	0.66	(0.51)	40	67	0.15	(0.88)
Belgium	49	60	-1.52	(0.13)	45	63	-0.79	(0.43)
France	41	69	0.18	(0.86)	38	70	0.64	(0.52)
Germany	44	69	-0.19	(0.85)	35	72	1.18	(0.24)
Greece	34	82	2.01*	(0.05)	24	95	4.05*	(0.00)
Ireland	27	90	3.40*	(0.00)	24	91	3.83*	(0.00)
Italy	27	84	3.04*	(0.00)	27	81	2.85*	(0.00)
Netherlands	42	67	-0.10	(0.92)	38	69	0.57	(0.57)
Portugal	14	105	5.58*	(0.00)	17	101	5.14*	(0.00)
Spain	30	89	2.96*	(0.00)	29	85	2.84*	(0.00)
Panel B: Non-EMU new EU countries								
Czech. Rep.	55	42	-3.81*	(0.00)	58	44	-3.94*	(0.00)
Hungary	54	55	-2.53*	(0.01)	54	52	-2.79*	(0.01)
Poland	49	38	-3.52*	(0.00)	44	47	-2.06*	(0.04)
Panel C: Non-EMU EU-15 countries								
Denmark	32	87	2.58*	(0.01)	32	84	2.39*	(0.02)
Sweden	34	79	1.80*	(0.07)	43	71	0.08	(0.93)
UK	46	68	-0.52	(0.61)	52	64	-1.55	(0.12)

Notes: The financial crisis extends from 7 August 2007 to 31 December 2009 and the sovereign debt crisis extends from 1 January 2010 to 15 December 2013. We estimate a logit model where the occurrence of exceedances is explained by a constant and a dummy variable indicating the beginning of the sovereign debt crisis. The *z*-statistic tests where the coefficient on the dummy variable is equal to zero. * denotes significance at the 10% level.

lowing countries: Austria, Belgium, France, Germany, Greece, Ireland, Italy, Portugal, Spain and the Netherlands (EMU EU-15 old); the Czech Republic, Hungary and Poland (non-EMU new EU); and Denmark, Sweden and the UK (non-EMU EU-15).

Table 1 shows the number of extreme returns in each European government bond market on a particular day. We compute the number of positive and negative exceedances for the financial crisis (from 7 August 2007 to 31 December 2009) and for the European sovereign crisis (from 1 January 2010 to 15 December 2013) separately. To test for differences in the number of exceedances during both periods, we estimate a logit model where the occurrence of exceedances is explained by a constant and a dummy variable indicating the beginning of the sovereign debt crisis.[4] The distribution of negative and positive exceedances is largely symmetrical. As expected, the logit results show that the number of exceedances (negative and positive) is significantly higher during the sovereign debt crisis than during the financial crisis for all peripheral bond markets (Greece, Ireland, Italy, Portugal

[4] We thank an anonymous reviewer for this suggestion.

and Spain). The logit results also show that the number of exceedances in the central bond markets (Austria, Belgium, France, Germany and the Netherlands) does not differ significantly between the two crises and in the non-EMU EU-15 countries the results are heterogeneous, depending on the country and the tail of the distribution. Finally, the number of exceedances in the non-EMU new EU countries is, in general, significantly higher during the financial crisis than during the sovereign debt crisis. As Abad and Chuliá (2013) suggest, these markets appear to be more strongly influenced by US news than they are by European news and, therefore, were more exposed to the financial crisis than to the sovereign debt crisis.

As for differences across countries, the bond markets of the peripheral countries show more exceedances than the other bond markets during the sovereign debt crisis and fewer during the financial crisis, though this is not surprising given that these countries were most affected by the euro-area sovereign debt crisis. On the whole, there are no great differences in the number of exceedances across the central countries, non-EMU new EU countries and non-EMU EU-15 countries, regardless of the particular crisis.

2.1 Explanatory Variables

We examine five main hypotheses relating market conditions to the likelihood of coexceedances and, to this end, we use a large set of explanatory variables. First, several arguments such as the flight-to-quality proposed by Caballero and Krishnamurthy (2008) and the liquidity spirals proposed by Brunnermeier and Pedersen (2009) suggest a relation between equities, money market instruments and bonds in turbulent periods. In line with these arguments, we would expect contagion in bond markets to be connected to the stock and money markets. To test this hypothesis, we include daily returns of the Euro Stoxx 50 and the three-month interbank interest rate (Euribor).

Second, according to Cristiansen and Ranaldo (2009), the propagation of shocks is more likely in a highly volatile environment. Thus, the hypothesis to be tested is whether contagion is strengthened when volatility is pervasively high in the financial markets. As a proxy of European financial market volatility, we use the VSTOXX index (which measures implied volatility in the Euro Stoxx 50 index options).

Third, given the pivotal role of monetary policy in determining bond returns, it is likely that monetary policy impacts the bond returns in each country (Bredin *et al.*, 2010). Thus, we examine whether contagion can be attributed to the release of monetary policy announcements. To test this hypothesis, we include the "surprise" or the unexpected component of the news announcements[5] released by the ECB.

Fourth, we define a cluster effect when past contagion significantly increases the probability of more contagion today.[6] To test this hypothesis, we include a previous coexceedance in the model as an explanatory variable. A positive value for this variable indicates higher probabilities of observing a coexceedance today if a coexceedance occurred the previous day.

[5] An important common finding in the extant literature is that only the surprise component of monetary policy has a significant effect on asset returns, whereas the effect of expected policy actions is statistically insignificant (see Bomfim, 2003, and Bernanke and Kuttner, 2005, among others).

[6] We thank an anonymous reviewer for this suggestion.

Finally, with the aim of distinguishing regional factors from global factors, the fourth group of variables is associated with the US. These variables are the return of the US stock market (S&P500 Composite index), the three-month Treasury bill rate, the US stock return volatility proxied by the VIX index (which measures implied volatility in Standard and Poor's 500 index options) and the "surprise" or unexpected component of the news announcements released by the Federal Reserve Bank (Fed). The hypotheses to be tested are (i) whether there is a relation between European and US assets similar to that between European asset classes, (ii) whether higher volatility in international financial markets increases contagion and, (iii) whether US monetary policy surprises are a source of contagion in European government bond markets.

To obtain a measure of the surprise in the Fed announcements we use the methodology proposed by Kuttner (2001). For an event taking place on day d, the unexpected or "surprise" target rate change can be calculated as the change in the rate implied by the current-month futures contract, scaled up by a factor related to the number of days in the month affected by the change. In sum, we compute the unexpected target rate change or the "surprise" as

$$S = \left[D/(D-d)\right]\cdot\left(f_d - f_{d-1}\right) \tag{1}$$

where f_d is the current-month futures rate at the end of the announcement day d and D is the number of days in the month. Kuttner (2001) uses a scaled version of the one-day change in the current-month federal funds future rate because in the US the futures contract's payoff depends on the monthly average federal funds rate, and the scaled factor is included to reflect the number of days remaining in the month that are affected by the change. This scaled factor is not required to obtain a measure of the surprise in the ECB announcement and, following Bredin *et al.* (2007), we proxy surprises in ECB policy rates using the one-day change in the three-month Euribor futures rate.[7] The data for the monetary policy-related variables are provided by Bloomberg.

3. The Model

Our aim is to identify the underlying determinants of contagion in European government bond markets and to determine whether their importance varies across and within groups of countries and across time. Here, we define a coexceedance as a variable that counts the number of large negative (large positive) returns on a given day across countries. To capture the probabilities associated with this polychotomous variable, Bae *et al.* (2003) use a multinomial logistic regression model that allows them to condition on attributes and characteristics of the exceedance events using control variables (or covariates). However, in so doing, they discard information, since multinomial logit ignores the ordered aspect of the outcome. To avoid this problem, we use the ordered logit model, which provides a means to exploit the ordering information.

[7] Bernoth and von Hagen (2004) find that the three-month Euribor futures rate is an unbiased predictor of euro-area policy rate changes.

Defining y_t^* as a latent variable (the level of contagion), $\tilde{x}_t^{'}$ as a vector of the explanatory variables discussed in Section 2, $\mathbf{\beta}$ as a vector of unknown parameters and epsilon as an error term, the ordered logit model is given by:

$$y_t^* = \tilde{x}_t^{'}\mathbf{\beta} + \varepsilon_t \tag{2}$$

The observation y is assumed to be related to the continuous latent variable y^* by:

$y = 0 \quad \text{if} \quad y^* \leq \mu_0$

$y = 1 \quad \text{if} \quad \mu_0 < y^* \leq \mu_1$

$y = 2 \quad \text{if} \quad \mu_1 < y^* \leq \mu_2 \ldots$

where y is the ordered measure of the level of contagion (coexceedance) and $\mathbf{\mu}$ the vector of unknown threshold parameters. We assume that ε has a standard logistic distribution and then:

$$\Pr\left(y_t = i \middle| \tilde{x}_t\right) = \Phi\left(\mu_i - \tilde{x}_t\mathbf{\beta}\right) - \Phi\left(\mu_{i-1} - \tilde{x}_t\mathbf{\beta}\right) \tag{3}$$

where Φ is the cumulative standard normal. The vector of the slopes $\mathbf{\beta}$ is not indexed by the category index c, thus the effects of the covariates are constant across response categories. A positive coefficient implies an increase in the log of the probability ratio, indicating that higher values of the explanatory variables imply a higher probability of observing a higher number of coexceedances (contagion). Similarly, a negative coefficient implies a lower probability of observing a higher number of coexceedances. The model is estimated using maximum likelihood and goodness-of-fit is measured using McFadden's (1974) pseudo-R^2 approach and testing the overall significance.

We estimate the model separately for positive and negative coexceedances to allow the factors to have different effects on each tail. As we are also interested in determining whether the effects of the factors differ during the financial and the European sovereign debt crises, we estimate the model separately for both periods. We used the regression parameters and each mean variable to generate predicted probabilities of possible coexceedances. This was done twice, first fixing the financial crisis and then fixing the sovereign debt crisis. We test for contagion (i) across all countries in the sample, (ii) within groups of countries and (iii) across groups of countries.

4. Determinants of European Government Bond Market Contagion

4.1 Contagion across All Countries

Table 2 shows the estimates of the ordered logit model for the top and bottom tails for all the countries considered together. To test the relation between markets, we include both European Union money and stock market variables. Our findings show that the Euribor is a statistically important covariate for both positive and negative coexceedances during the sovereign debt crisis and that the European stock market only appears to be useful in explaining positive coexceedances during the sovereign debt crisis.

Table 2 Contagion Across Countries

	TOP TAIL				BOTTOM TAIL			
	Financial crisis		Sov debt crisis		Financial crisis		Sov debt crisis	
CO_{t-1}	0.12*	(0.00)	0.17*	(0.00)	0.11*	(0.00)	0.13*	(0.00)
SR_t^{ECB}	-14.42	(0.98)	-0.43	(0.56)	-0.65	(0.56)	-1.78	(0.10)
R_t^{EMU}	0.00	(0.97)	0.38*	(0.02)	0.08	(0.36)	0.50*	(0.00)
S_t^{EMU}	0.67	(0.89)	9.85*	(0.05)	2.08	(0.68)	6.30	(0.20)
$V2X_t$	-0.01	(0.75)	0.19*	(0.00)	0.10*	(0.00)	0.15*	(0.00)
SR_{t-1}^{FED}	-0.42	(0.53)			0.93	(0.14)		
R_{t-1}^{US}	-0.40*	(0.01)	-0.59	(0.67)	-0.43*	(0.01)	-0.86	(0.53)
S_{t-1}^{US}	-10.12*	(0.01)	5.70	(0.35)	3.18	(0.45)	-12.79*	(0.03)
VIX_{t-1}	0.05*	(0.08)	-0.15*	(0.00)	-0.05	(0.12)	-0.11*	(0.00)
χ^2	143.88	(0.00)	185.62	(0.00)	152.74	(0.00)	163.14	(0.00)
Likelihood	-689.92		-1362.65		672.49		-1369.63	
R^2	0.09		0.06		0.10		0.06	
Prob y = 0	70%		55%		72%		57%	
Prob y = 1	17%		23%		17%		20%	
Prob y > 1	13%		22%		12%		22%	

Notes: The number of coexceedances of daily returns (*CO*) is modeled as an ordered logit. SR_t^{ECB} and SR_{t-1}^{FED} refer to monetary policy surprises announced by the ECB and the Fed, respectively; R_t^{EA} and R_{t-1}^{US} refer to the three-month interbank interest rate (Euribor) and the three-month Treasury bill rate, respectively; S_t^{EA} and S_{t-1}^{US} refer to the Euro Stoxx 50 index returns and the S&P500 index returns, respectively; and *V2X* and *VIX* refer to the VSTOXX index and VIX index, respectively. χ^2 refers to the statistic to test that the model considered provided a better fit than the null model with no independent variables. R^2 refers to McFadden's pseudo-R^2. During the sovereign debt crisis, no surprises were announced by the Fed. Prob denotes the adjusted probabilities (from the ordered logit model) that the number of exceedances is equal to 0, 1 or more than 1. We consider that no surprises were announced by the ECB or the FED and hold the other variables at their means. * indicates significance at the 10% level.

In general, our results confirm the hypothesis that shock propagation increases in a highly volatile European environment during both crises. We also find strong evidence for the presence of a cluster effect during both crises and in both tails. The positive and significant coefficients of the previous coexceedance show that there is a higher probability of observing a coexceedance today if a coexceedance occurred the previous day.

In addition, we examine whether some fraction of the coexceedance events in European government bond markets can be explained by the explanatory variables associated with the US, i.e. whether global factors have an effect on contagion.[8] Our results show that, in general, the likelihood of observing coexceedances is negatively

[8] Owing to timing conventions (European markets close before their US counterpart), US explanatory variables enter the model lagged one period. Similar to Bae *et al.* (2003), we interpret these results as evidence of the predictability of coexceedances.

related to the US interest rate and stock market. US volatility is helpful for predicting contagion across European government bond markets. However, the effect of the volatility during the sovereign debt crisis is negative, suggesting that the higher the US volatility yesterday, the less likely we are to observe episodes of coexceedance in European government markets today.[9]

An examination of the impact of unexpected news announcements released by the ECB and the Fed shows that they are not useful in explaining contagion in any tail. This result is consistent with that of Connolly and Wang (2003), who show that the bulk of the observed comovement in international equity markets cannot be attributed to public information about economic fundamentals.

As expected, our results also show that the probability of contagion (the probability of exceedances being more than 1) is higher during the sovereign debt crisis than during the financial crisis. Finally, the chi-square test indicates that all the models considered provided a better fit than the null model with no independent variables when predicting cumulative probabilities of contagion.

4.2 Contagion within Groups of Countries

Table 3 shows the estimates of the ordered logit model for the top and bottom tails when we group the countries as is standard in the literature: EMU EU-15 old countries (Panel A), non-EMU new EU countries (Panel B), and non-EMU EU-15 countries (Panel C). If we consider the old members, the results are similar[10] to those in *Table 2*, but now the European stock market appears to be useful in explaining both positive and negative coexceedances during both crises.

The results for the non-EMU new EU countries reveal three important differences with respect to the old members. First, there appears to be no relation between European interest rates and contagion within these countries. Second, in general, US assets fail to provide an adequate explanation of the probability of coexceedances within these bond markets. Third, the probability of contagion is higher during the financial crisis than during the sovereign debt crisis.

Finally, our results for the non-EMU EU-15 countries (Panel C, *Table 3*) show that what distinguishes them most obviously from the old members is that shock propagation only increases in a highly volatile European environment during the sovereign debt crisis.

4.3 Contagion across Groups of Countries

We are also interested in testing whether contagion within the group of old EMU members can help predict contagion within the other groups of countries. We define contagion across groups of countries as the fraction of the coexceedance events in each group of countries that is left unexplained by its own covariates but is explained by coexceedances within the group of old members.

Our results in *Table 4* show that our findings on contagion within groups of countries (see *Table 3*) are robust to the inclusion of the coexceedances within

[9] Bae *et al*. (2003) also find that the effect of US conditional volatility on the probability of contagion in Latin America is negative.

[10] Ten of the 16 countries in the sample belong to this category of "old countries".

Table 3 Contagion within Groups of Countries

	TOP TAIL				BOTTOM TAIL			
	Financial crisis		Sov debt crisis		Financial crisis		Sov debt crisis	
Panel A: EMU EU-15 old countries								
$CO_{t\text{-}1}$	-0.29*	(0.06)	-0.25*	(0.00)	-0.41	(0.16)	-0.37*	(0.00)
SR_t^{ECB}	-11.38	(0.98)	-0.71	(0.39)	-0.01	(1.00)	-1.37	(0.21)
R_t^{EMU}	0.09	(0.49)	0.63*	(0.00)	0.05	(0.67)	0.80*	(0.00)
S_t^{EMU}	-34.62*	(0.00)	9.04*	(0.09)	29.74*	(0.00)	9.46*	(0.08)
$V2X_t$	-0.09*	(0.08)	0.19*	(0.00)	0.10*	(0.05)	0.15*	(0.00)
SR_{t-1}^{FED}	-0.44	(0.63)			1.56*	(0.03)		
R_{t-1}^{US}	-0.61*	(0.02)	-1.22	(0.41)	-0.27	(0.21)	-1.04	(0.48)
S_{t-1}^{US}	-13.04*	(0.02)	1.20	(0.85)	5.76	(0.32)	-11.19*	(0.07)
$VIX_{t\text{-}1}$	0.12*	(0.01)	-0.15*	(0.00)	-0.06	(0.19)	-0.11*	(0.00)
χ^2	68.06	(0.00)	128.23	(0.00)	51.40	(0.00)	139.57	(0.00)
Likelihood	-346.17		-1124.83		-358.99		-1130.40	
R^2	0.09		0.05		0.07		0.06	
Prob y = 0	92%		66%		90%		67%	
Prob y = 1	2%		17%		3%		17%	
Prob y > 1	6%		17%		6%		17%	
Panel B: Non-EMU new EU countries								
$CO_{t\text{-}1}$	0.32*	(0.06)	0.13	(0.55)	0.25	(0.14)	0.32*	(0.07)
SR_t^{ECB}	-14.07	(0.99)	-0.10	(0.93)	-12.93	(0.99)	-12.07	(0.98)
R_t^{EMU}	-0.08	(0.43)	-0.10	(0.76)	0.04	(0.75)	-0.30	(0.35)
S_t^{EMU}	40.87*	(0.00)	59.50*	(0.00)	-35.88*	(0.00)	-39.38*	(0.00)
$V2X_t$	0.04	(0.34)	0.17*	(0.00)	0.06	(0.15)	0.15*	(0.00)
SR_{t-1}^{FED}	-0.54	(0.63)			-0.51	(0.58)		
R_{t-1}^{US}	-0.03	(0.88)	0.58	(0.80)	-0.38*	(0.07)	-2.18	(0.38)
S_{t-1}^{US}	5.87	(0.28)	-7.43	(0.44)	1.67	(0.76)	-8.30	(0.42)
$VIX_{t\text{-}1}$	0.01	(0.79)	-0.12*	(0.01)	-0.01	(0.74)	-0.11*	(0.02)
χ^2	118.63	(0.00)	83.11	(0.00)	140.47	(0.00)	113.08	(0.00)
Likelihood	-324.63		-367.34		-308.73		-355.89	
R^2	0.15		0.10		0.19		0.14	
Prob y = 0	86%		92%		87%		93%	
Prob y = 1	11%		7%		10%		5%	
Prob y > 1	4%		1%		3%		2%	
Panel C: Non-EMU EU-15 countries								
$CO_{t\text{-}1}$	0.10	(0.64)	0.29*	(0.06)	0.46*	(0.01)	0.33*	(0.02)
SR_t^{ECB}	-13.29	(0.99)	0.10	(0.93)	0.03	(0.98)	-13.12	(0.98)

R_t^{EMU}	-0.07	(0.60)	0.46*	(0.09)	0.15	(0.23)	-0.15	(0.56)
S_t^{EMU}	-21.20*	(0.00)	-55.96*	(0.00)	6.39	(0.29)	63.08*	(0.00)
$V2X_t$	-0.05	(0.21)	0.17*	(0.00)	0.03	(0.46)	0.09*	(0.01)
SR_{t-1}^{FED}	-0.40	(0.61)			1.48*	(0.02)		
R_{t-1}^{US}	-0.65*	(0.02)	1.67	(0.43)	-0.89*	(0.00)	-1.79	(0.38)
S_{t-1}^{US}	5.57	(0.28)	14.88	(0.10)	8.35*	(0.09)	-7.22	(0.42)
VIX_{t-1}	0.09*	(0.03)	-0.15*	(0.00)	0.00	(1.00)	-0.08*	(0.07)
χ^2	84.65	(0.00)	179.70	(0.00)	99.17	(0.00)	108.10	(0.00)
Likelihood	-270.73		-476.77		-289.00		-501.81	
R^2	0.14		0.16		0.15		0.10	
Prob y = 0	90%		89%		90%		88%	
Prob y = 1	8%		7%		8%		8%	
Prob y > 1	2%		4%		2%		4%	

Note: See notes to *Table 2*.

the group of old members as an explanatory variable. The contagion test for the new members is presented in Panel A of *Table 4.* The regression coefficient on the number of coexceedances in the old countries is significant during the sovereign debt crisis. This means that during that crisis, there are higher probabilities of observing contagion within the new members today if contagion has occurred in the old countries. Nevertheless, as discussed above, the probability of contagion within the group of new members is higher during the financial crisis than during the sovereign debt crisis.

If we consider contagion from the old members to the non-EMU EU-15 countries (Panel B, *Table 4*), we find that the regression coefficient on the number of coexceedances in the old countries is positive and significant during both crises and in both tails. This positive relationship indicates that both groups of countries are connected even though they do not share a common currency.

Finally, although the use of McFadden's pseudo-R^2 for comparison across models has some limitations (Greene, 2000), in general we observe that McFadden's pseudo- R^2 is higher during the financial crisis than during the sovereign debt crisis, especially in the models that analyze contagion across groups of countries.

5. Conclusions

Using the coexceedance measure proposed by Bae *et al.* (2003), we have analyzed contagion in European government bond markets. Specifically, we have tested for contagion (i) across all countries in the sample, (ii) within groups of countries and (iii) across groups of countries. To this end, we have used an ordered logistic regression model to analyze the underlying determinants of the dynamics of contagion.

We report evidence that interest rates, stock market returns and market volatility in both regional and international stock markets are statistically important covariates that help explain contagion in European government bond markets.

Table 4 Contagion across Groups of Countries

	TOP TAIL				BOTTOM TAIL			
	Financial crisis		Sov debt crisis		Financial crisis		Sov debt crisis	
Panel A: Non-EMU new EU countries								
CO_{t-1}	0.30*	(0.08)	0.02	(0.93)	0.24	(0.15)	0.24	(0.18)
SR_t^{ECB}	-13.73	(0.98)	0.11	(0.92)	-12.95	(0.99)	-13.12	(0.99)
R_t^{EMU}	-0.08	(0.48)	-0.19	(0.54)	0.03	(0.76)	-0.41	(0.21)
S_t^{EMU}	39.08*	(0.00)	57.41*	(0.00)	-36.14*	(0.00)	-39.52*	(0.00)
$V2X_t$	0.03	(0.51)	0.14*	(0.00)	0.06	(0.17)	0.14*	(0.00)
SR_{t-1}^{FED}	-0.66	(0.56)			-0.56	(0.56)		
R_{t-1}^{US}	-0.05	(0.78)	0.74	(0.75)	-0.38*	(0.07)	-1.85	(0.46)
S_{t-1}^{US}	6.99	(0.20)	-7.16	(0.46)	1.60	(0.77)	-7.21	(0.47)
VIX_{t-1}	0.03	(0.51)	-0.10*	(0.04)	-0.01	(0.76)	-0.09*	(0.05)
CO_Old_t	-0.13	(0.10)	0.23*	(0.00)	0.02	(0.76)	0.25*	(0.00)
χ^2	121.96	(0.00)	93.64	(0.00)	140.56	(0.00)	126.01	(0.00)
Likelihood	-322.96		-362.08		-308.69		-349.43	
R^2	0.16		0.11		0.19		0.15	
Prob $y = 0$	86%		92%		87%		93%	
Prob $y = 1$	11%		7%		10%		5%	
Prob $y > 1$	4%		1%		3%		2%	
Panel B: Non-EMU EU-15 countries								
CO_{t-1}	-0.01	(0.95)	0.35*	(0.03)	0.36*	(0.07)	0.42*	(0.01)
SR_t^{ECB}	-12.24	(0.98)	0.55	(0.62)	-1.33	(0.38)	-13.30	(0.98)
R_t^{EMU}	-0.17	(0.25)	0.15	(0.58)	0.08	(0.59)	-0.43	(0.13)
S_t^{EMU}	-5.50	(0.45)	-58.34*	(0.00)	-14.69*	(0.03)	70.70*	(0.00)
$V2X_t$	-0.01	(0.90)	0.11*	(0.01)	-0.07	(0.11)	0.08*	(0.04)
SR_{t-1}^{FED}	0.07	(0.93)			0.50	(0.50)		
R_{t-1}^{US}	-0.57*	(0.06)	2.44	(0.27)	-1.25*	(0.00)	-0.70	(0.75)
S_{t-1}^{US}	-0.96	(0.86)	12.03	(0.23)	9.87*	(0.07)	-1.16	(0.91)
VIX_{t-1}	0.04	(0.39)	-0.10*	(0.03)	0.09*	(0.03)	-0.11*	(0.02)
CO_Old_t	0.44*	(0.00)	0.54*	(0.00)	0.75*	(0.00)	0.73*	(0.00)
χ^2	151.90	(0.00)	263.08	(0.00)	242.48	(0.00)	269.11	(0.00)
Likelihood	-237.11		-435.07		-217.35		-421.30	
R^2	0.24		0.23		0.36		0.24	
Prob $y = 0$	91%		90%		93%		90%	
Prob $y = 1$	8%		7%		6%		8%	
Prob $y > 1$	1%		3%		1%		2%	

Notes: See note to Table 2. CO_Old_t refers to the number of coexceedances in old countries

Nevertheless, the relevance of these variables is heterogeneous and differs between crises. In addition, we find evidence indicating that past contagion significantly increases the probability of more contagion today. Third, we find that the probability of contagion among the non-EMU new EU members has a statistically significant relationship with coexceedances occurring among the old members during the sovereign debt crisis and in the case of non-EMU EU-15 members during both crises. Finally, we find that the new European Union members are those that behave most differently.

Our results should enable market participants to make effective investment decisions, given that they need to have an understanding of the way in which extreme shocks propagate across European government bond markets and of the factors underlying contagion. Additionally, our findings have two policy implications. First, our results show that the probability of contagion was higher for the new EMU members during the financial crisis and for the rest of members during the sovereign debt crisis. This result indicates the need for stabilization mechanisms to avoid contagion, though individual countries ought to consider what their main source of contagion is. Second, since we have found that both money markets and stock markets are important determinants of sovereign debt market contagion, policy regulators should make an effort to control all financial markets simultaneously. The link between money markets and contagion is especially important since interest rates were the most frequently employed monetary policy tool in response to the crises.

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