The Walking Debt Crisis

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Abstract

We use recent methodologies suggested by Phillips, Wu, and Yu (2011) and Phillips and Yu (2011) to shed light on the question whether the EMU crisis was triggered by the US subprime crunch. We define crisis regimes by explosive behavior of interest rates and government bond yield spreads. As expected, we find clear evidence for explosive behavior during the EMU crisis and coincident with the bankruptcy of Lehman Brothers in the spreads. We estimate the time-varying persistence for a US house price proxy. Furthermore, we employ an explosiveness migration test and obtain a migration from the house price proxy to the spreads. EMU interest rates are considered to investigate whether the migration process is persistent during the EMU crisis. The results reported in this paper indicate that there is explosiveness migration caused by the Lehman Brother bankruptcy, but no migration during the EMU crisis from US mortgage markets to EMU interest rates. These findings suggest that the EMU debt crisis is a homemade problem. Our results remain unchanged after performing some robustness checks.

JEL classification: H63, C58.

Keywords: Sovereign Debt Crisis, Contagion, Explosiveness.

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1. Introduction

U.S. house prices peaked in the years 2005 and 2006. As a consequence, the availability of credit for potential buyers of real estate decreased and borrowers experienced more and more problems to refinance their loans. This new environment increased the pressure on real estate prices. The resulting dramatic fall of U.S. house prices, of course, had massive negative effects on the prices of U.S. subprime mortgage-backed securities. These collateralized bonds had also been bought by financial institutions in Europe and Asia. Therefore, not only U.S. banks (e.g., Lehman Brothers and Washington Mutual) all of a sudden found themselves to be in deep trouble. The collapse of Lehman Brothers intensified the problems. At this point one of the central questions (see Eichengreen, Mody, Nedeljkovic, and Sarno, 2012) seems to be how the Subprime Crisis, a problem in a rather small segment of U.S. financial markets, was able to have such serious negative consequences for the global economy. One of the key answers to this important question clearly is the global banking system. In fact, international banks have played a critical role in the transmission of the crisis from the U.S. to Europe and other parts of the world. Most importantly, banks have been responsible for causing some additional fiscal problems in some European countries. Basse, Friedrich, and Kleffner (2012), for example, have identified structural change in the relationship between German and Italian government bond yields. They have found two structural breaks that can be explained by changes to sovereign credit risk. While the first structural break identified coincides with the U.S. Subprime Crisis and the resulting bank rescue programmes in Europe, the second structural break might be a consequence of a phenomenon that could be called the European sovereign debt crisis. Three countries (namely Greece, Portugal and Ireland) played a special role in this second part of the crisis. The sudden increase of the importance of sovereign credit risk in Europe has had major consequences for the pricing of fixed income securities in one of the biggest bond markets of the world. Sibbertsen, Wegener, and Basse (2014), for example, have argued that the crisis at least for the moment has ended the process of interest rate convergence in the European Monetary Union (EMU). Only in Ireland the fiscal problems of the government can be explained by the costs resulting from measures to stabilize the financial system of the country. Moreover, there also have been dangerous imbalances within the Eurozone between surplus nations with higher exports than imports and deficit nations that imported more goods and services than they exported. Given that the existence of the common currency made it impossible for deficit nations to devalue and thereby improve their competitiveness Varoufakis (2013), for example, has argued that even without the credit crunch in the U.S. and the subsequent events in 2008 something bad simply had to happen. Therefore, it could also be argued that this second part of the crisis actually was no second part but a crisis of its own. Ludwig (2014) already has presented an interesting discussion of this issue. We try to find new relevant empirical evidence. More specifically, we use a methodology that recently has been suggested by Phillips, Wu, and Yu (2011) and Phillips and Yu (2011) to shed new light on this question.

The paper is structured as follows: Section 2 gives a short review of the relevant literature focusing on interest rate convergence in the EMU. The 3rd section discusses methodological issues and introduces the data examined. The empirical evidence is presented in section 4. Section 5 summarizes the results of several robustness checks and the last section concludes.

2. Literature Review

The introduction of the Euro in 1999 has been very important for the bond markets of the EMU countries because the new common currency has eliminated exchange rate risk among the member states. Therefore, it is no surprise that Kim, Moshirian, and Wu (2006) have been able to documented that the Euro has caused structural change in the bond market. Lund (1999) has argued that a binding time table for the introduction of the common currency existed before 1999. Consequently, the prospects of monetary union should already have fixed income markets before the introduction of the Euro. Laopodis (2008) has reported an increase in the correlation of the returns on Euro government bonds after the introduction of the new currency. Using techniques of cointegration analysis this empirical study also has identified the existence of two groups of EMU countries – a core group (including Germany and France) and some peripheral countries (including Italy and Ireland). Also employing methods of cointegration analysis Jenkins and Madzharova (2008) have been able to find cointegration among nominal government bond yields in the Euro area after the introduction of the Euro. Thus, they have argued convincingly that interest rates in EMU countries have converged. Meanwhile, the European debt crisis has caused some concerns about sovereign credit risk and possibly even redenomination risk (which means the return of currency risk due to the breakdown of the EMU) in the market for fixed income securities. This is a relatively new strand of literature. Gruppe and Lange (2014) have shown that higher sovereign credit risk has caused structural change among government bond yields in Germany and Spain. Moreover, using a similar approach Basse (2014) has reported that government bond yields in Austria, Belgium, Finland and the Netherlands seem to be cointegrated with German government bond yields and that there has been no sign for structural change caused by the crisis. Gómez-Puig and Sosvilla-Rivero (2014) also have searched for structural change in EMU government bond markets and have argued that more than half of the breakpoints identified seem to be connected to the Euro sovereign debt crisis. Moreover, Sibbertsen, Wegener, and Basse (2014) have tested for a break in the persistence of EMU government bond yield spreads examining data from France, Italy and Spain using German sovereign bonds as benchmark. Their results seem that there are structural breaks. The persistence of the examined time series has increased significantly during the crisis. This could be a sign of higher sovereign credit risk and of redenomination risk.

3. Data and Methodology

3.1. Data

We use daily data from 8/31/2001 until 7/31/2015 taken from the Bloomberg Database. All series are normalized to 1 at the beginning of the sample. Spreads – market by $\tilde{\Delta}$ – are created as the simple difference between Germany and Greece (GR_t), Italy (IT_t), Spain (ES_t) and Portugal (PT_t) in levels. The daily house price proxy is indicated by REI_t. This is the Dow Jones Equity REIT price index. Equity REITs invest in properties. Therefore, a broad index consiting of U.S. Equity REITs should be regarded as a useful measure of economic activity in the North American real estate sector. Data on this share price index is available on a daily basis.[1]

3.2. Testing explosiveness

We are interested in whether our time series show explosive behavior. This paper deals with a mildly form of explosiveness – as analyzed by Phillips and Magdalinos (2007) – defined by the model

$$y_t = \rho y_{t-1} + \varepsilon_t \tag{1}$$

with y_t as a stochastic process in discrete time, $t = 1, ..., T$, ε_t as the innovation sequence and $\rho = 1 + \frac{c}{k_T}$ with $c > 0$. $(k_T)_{T \geq 1}$ is a sequence which increases to infinity such that $k_T = o(T)$ when $T \to \infty$. ε_t is an independent and identically random variable or weakly dependent with $E(\varepsilon_t) = 0$ and $E(\varepsilon_t^2) < \infty$.

In order to identify explosiveness, we apply the procedure by Phillips, Wu, and Yu (2011)

[1] To underline our assumption, that REITs are adequate as a house price market proxy, we investigated the number of cointegration relations between REITs and S&P Case-Shiller Home Price Index using the procedure suggested by Johansen (1988, 1991). The hypothesis of zero relations has been rejected while the hypothesis of one relation can not be rejected on a 1% significance level. The S&P Case-Shiller Home Price Index examined here is a very popular measure of residential real estate prices in 20 U.S. metropolitan areas.

to each of our time series. The regression of the augmented Dickey-Fuller (ADF) test

$$y_t = \mu + \rho y_{t-1} + \sum_{j=1}^{J} \theta_j \Delta y_{t-j} + \varepsilon_t, \varepsilon_t \sim NID\left(0, \sigma^2\right) \tag{2}$$

is estimated by Ordinary Least Squares (OLS) for some lag length J. Δ indicates first differences and NID denotes independent and normal distribution. We are interested in the hypothesis of an unit root process $H_0 : \rho = 1$ against the right tailed alternative $H_1 : \rho > 1$.

We assume a temporary limited mildly explosive model of the form, as analyzed by Phillips and Yu (2009),

$$y_t = y_{t-1} 1\left\{t < \tau_e\right\} + \rho y_{t-1} 1\left\{\tau_e \leq t \leq \tau_f\right\} + \left(\sum_{k=\tau_f+1}^{t} \varepsilon_k + y^*_{\tau_f}\right) 1\left\{t > \tau_f\right\} + \varepsilon_t 1\left\{t \leq \tau_f\right\}, \tag{3}$$

with $\rho = 1 + \frac{c}{T^\alpha}$, $c > 0$ and $\alpha \in (0, 1)$. This model switches from an unit root to an explosive regime at τ_e and back to unit root behavior at τ_f. The model comes to a new level $y^*_{\tau_f}$ with a re-initialization at τ_f. Furthermore, a short transitional period is allowed when it comes from explosive to unit root behavior in which the process is mean reverting.

Thus, we use a forward recursive approach to test against explosiveness. This procedure, proposed by Phillips, Wu, and Yu (2011), deals with the estimation of model 2 involving subsamples of the data by the expansion of one observation at each run. The first estimation includes $\tau_0 = [Tr_0]$ observations. r_0 is some fraction of the whole sample and $[x]$ indicates the integer part of x. Thus, the regression involves $\tau = [Tr]$ observations for $r_0 \leq r \leq 1$. Denoting the t-statistic by ADF_r, Phillips, Wu, and Yu (2011) specify the limiting distribution under the null as

$$ADF_r \Rightarrow \frac{\int_0^r \widetilde{W} dW}{\left(\int_0^r \widetilde{W}^2\right)^{1/2}} \quad \text{and} \quad \sup_{r \in [r_0, 1]} ADF_r \Rightarrow \sup_{r \in [r_0, 1]} \frac{\int_0^r \widetilde{W} dW}{\left(\int_0^r \widetilde{W}^2\right)^{1/2}},$$

with W as the standard Brownian motion and $\widetilde{W}(r) = W(r) - \frac{1}{\tau} \int_0^1 W$. To stamp the origination $\hat{r}_e$ and the collapse date $\hat{r}_f$ of the explosive behavior, Phillips and Yu (2011) construct the estimates as

$$\hat{r}_e = \inf_{s \geq r_0} \left\{ s : ADF_s > cv^{adf}_{\beta_T}(s) \right\}, \hat{r}_f = \inf_{s \geq \hat{r}_e + \gamma \ln(T)/T} \left\{ s : ADF_s < cv^{adf}_{\beta_T}(s) \right\}. \tag{4}$$

$\gamma \ln(T)$ ensures that a short episode after the origination is not considered for a collapse date stamping and $cv^{adf}_{\beta_T}(s)$ is the right-sided critical value with a significance level of β_T.

For practical implementation, the authors suggest to set the critical value to $cv_{\beta_T}^{adf}(s) = -0.08 + \ln([Tr])/C$. This helps to ensure the consistent estimation of both parameters by a slowly varying rate of $cv_{\beta_T}^{adf}(s)$. We set C to 1000, thus for large sample sizes, we are close to the ADF_r 5% critical values.

3.3. Testing the migration of explosiveness

The following exposition draws heavily from Phillips and Yu (2011). The authors propose a test procedure – which makes use of the recursive estimation of ρ as introduced in the foregone section – to test against migration of explosive behavior from one variable to another. We have two time series y_t and x_t with mildly and timely limited explosive autoregressive regimes as in equation (3). Suppose that the start of explosiveness is denoted by $\tau_{ey} = [Tr_{ey}]$ and $\tau_{ex} = [Tr_{ex}]$ respectively. Furthermore, the estimated autocorrelation coefficient $\hat{\rho}_y$ peaks at $\tau_{py} = [Tr_{py}]$ and $\hat{\rho}_x$ peaks at $\tau_{px} = [Tr_{px}]$. Additionally, it is assumed that $r_{py} > r_{px}$.

We obtain for ρ_y under the null

$$\rho_y(\tau) = \begin{cases} 1, & \tau < \tau_{ey} = [Tr_{ey}] \\ 1 + \frac{c_y}{T^\alpha}, & \tau > \tau_{ey} = [Tr_{ey}] \end{cases}, \tag{5}$$

and under the alternative

$$\rho_y(\tau) = \begin{cases} 1, & \tau < \tau_{ey} = [Tr_{ey}] \\ 1 + \frac{c_y}{T^\alpha} + d\frac{c_x}{T^\alpha}\left(\frac{\tau - \tau_{px}}{m}\right)^2, & \tau > \tau_{ey} = [Tr_{ey}] \end{cases}, \tag{6}$$

with $m = \tau_{py} - \tau_{px} = [Tr_{py}] - [Tr_{px}]$. ρ_x is defined under both hypothesis as

$$\rho_x(\tau) = \begin{cases} 1, & \tau < \tau_{ex} = [Tr_{ex}] \\ 1 + \frac{c_{ex}}{T^\alpha}, & \tau > \tau_{ex} = [Tr_{ex}] \\ 1 + \frac{c_x}{T}\left(\frac{\tau - \tau_{px}}{m}\right), & \tau > \tau_{px} = [Tr_{px}] \end{cases}. \tag{7}$$

We assume $c_{ex} > 0$ and a negative localizing coefficient function $c_x(\cdot) < 0$. Thus, ρ_x is local to unity upon the explosive regime which influences the behavior of ρ_y. Phillips and Yu (2011) assume a linear relation for ρ_x with a constant $c_x < 0$. This leads to $dc_x = 0$ under the null and to $dc_x > 0$ under the alternative and enables us to test the hypothesis

$$H_0 : \beta_1 = 0 \text{ vs. } H_1 : \beta_1 < 0 \tag{8}$$

with β_1 from the regression model

$$\hat{\rho}_y(\tau) - 1 = \beta_0 + \beta_1 (\hat{\rho}_x(\tau) - 1) \frac{\tau - \tau_{px}}{m} + \epsilon(\tau), \tag{9}$$

where β_0 and β_1 are OLS estimates and $\epsilon(\tau)$ is the error sequence for $\tau = [Tr_{px}]+1, ..., [Tr_{py}]$. Figure 1 and figure 2 show the trajectory of ρ_x and ρ_y under both hypothesis.

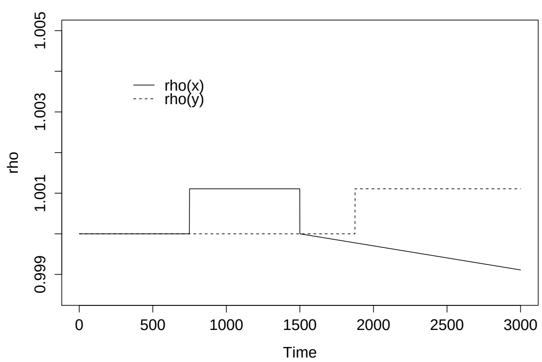

Fig. 1. ρ_x and ρ_y under H_0.

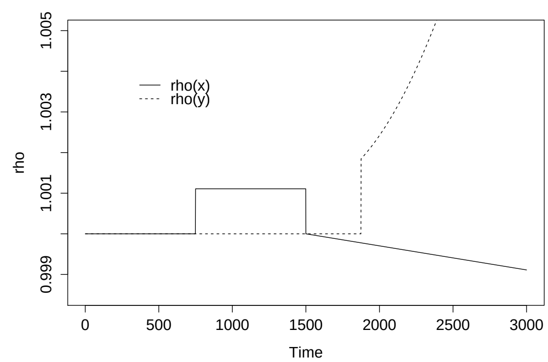

Fig. 2. ρ_x and ρ_y under H_1.

See Phillips and Yu (2011) for the limit theory of $\hat{\beta}_1$. They constructed an asymptotically conservative and consistent test of the hypothesis in equation (8) based on the statistic

$$Z_\beta = \frac{\hat{\beta}_1}{L(m)}, \text{ with } \frac{1}{L(m)} + \frac{L(m)}{n^e} \to 0 \text{ as } n \to \infty \text{ for any } e > 0. \tag{10}$$

The test has asymptotically zero size because $\hat{\beta}_1/L(m) \to_p 0$ under the null and unit power because $Z_\beta = O_p(T^{1-\alpha}/L(m))$ under the alternative for some slowly varying function $L(m)$. Z_β is compared to critical values from the standard normal distribution $cv_{N,\alpha}$ and rejects the H_0 if $|Z_\beta| > cv_{N,\alpha}$. The authors suggest to set $L(m) = a\log(m)$ with $1/3 \leq a \leq 3$ to control the size of the test. Figure 3 shows the density of Z_β with $L(m) = 3\log(m)$, $m = 100$ under the null. The dotted line is the density of the $N(0,1)$ distribution and the vertical line indicates the 0.95 quantile. This result is obtained by Monte Carlo Simulation with $mc = 10000$ iterations. We have $n = 3631$ observation for our empirical application. Thus, we use another Monte Carlo Simulation with $n = 3000$, $m = 600$ and the same settings as in the case before. The result indicates that the test holds it's size with $a = 1/3$[2].

[2] We used $mc = 300$ iterations. It will be necessary to increase mc to get clearer evidence about the empirical size and power of this procedure. However, due to the recursive estimation of ρ_x and ρ_y, we are faced with extensive computing time.

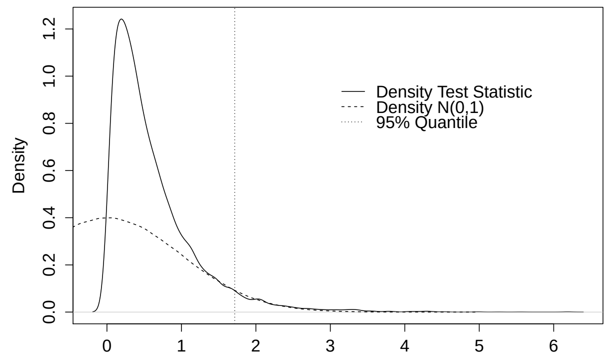

Fig. 3. Density of $|Z_\beta|$

4. Empirical Results

4.1. Testing against explosiveness and migration effects

At first, we apply the test procedure by Phillips, Wu, and Yu (2011) to REI_t. Figure 4 indicates arising and collapsing explosiveness from the beginning of our sample until 2007. The recursive estimated $\hat{\rho}_{REI_t}$ peaks at 2/8/2007, when HSBC announced higher provisions for bad mortgage loans. Now, we apply this procedure to the spreads of Greece, Spain, Italy

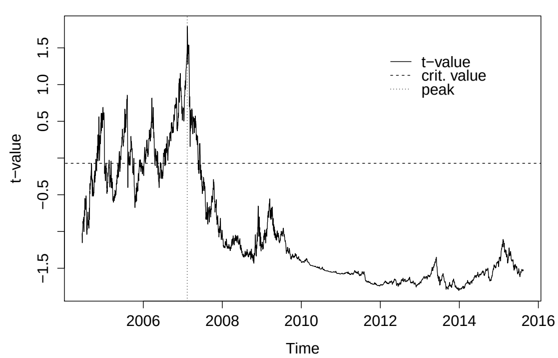

Fig. 4. Trajectory of $t_{\rho_{REI}}$.

and Portugal against Germany as reported by figure 5 to figure 8. The values of the maximal autocorrelation coefficient are reported by table 1.

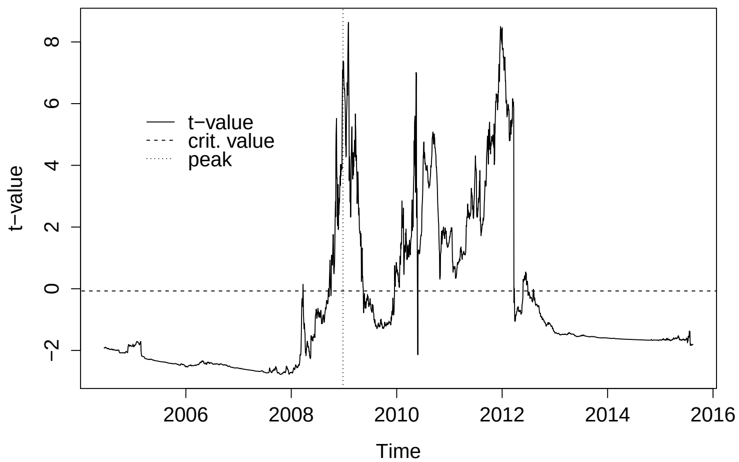

Fig. 5. Trajectory of $t_{\hat{\rho}_{\tilde{\Delta} GR_t}}$.

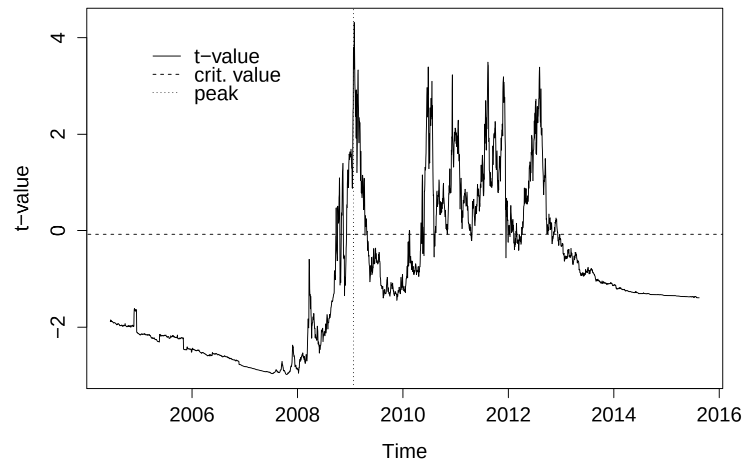

Fig. 6. Trajectory of $t_{\hat{\rho}_{\tilde{\Delta} ES_t}}$.

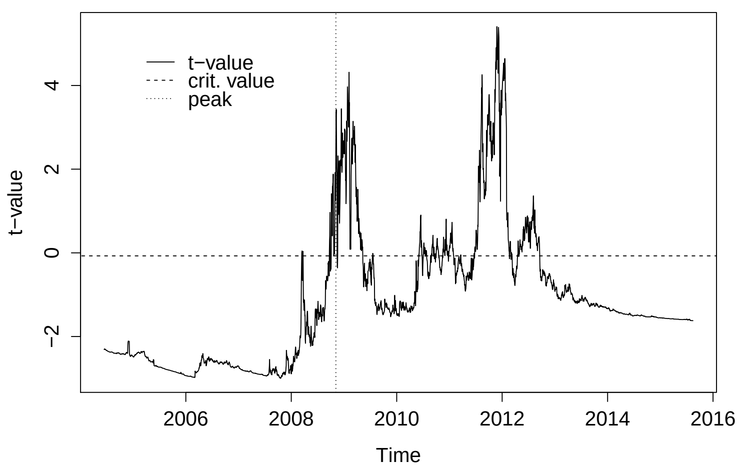

Fig. 7. Trajectory of $t_{\hat{\rho}_{\tilde{\Delta} IT_t}}$.

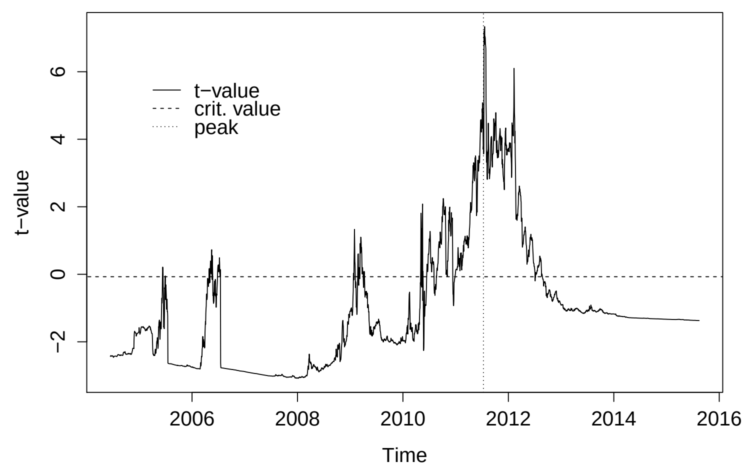

Fig. 8. Trajectory of $t_{\hat{\rho}_{\tilde{\Delta} PT_t}}$.

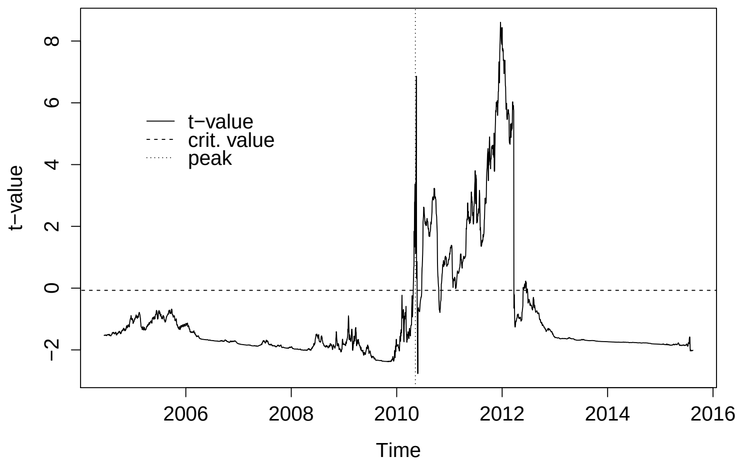

Fig. 9. Trajectory of $t_{\hat{\rho}_{GR_t}}$.

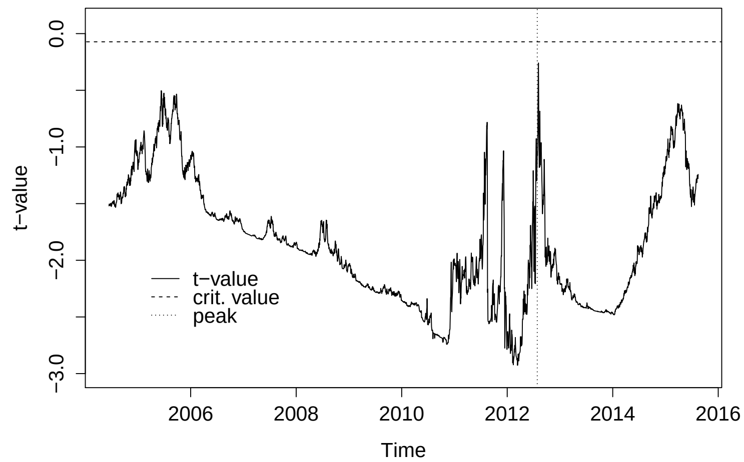

Fig. 10. Trajectory of $t_{\hat{\rho}_{SP_t}}$.

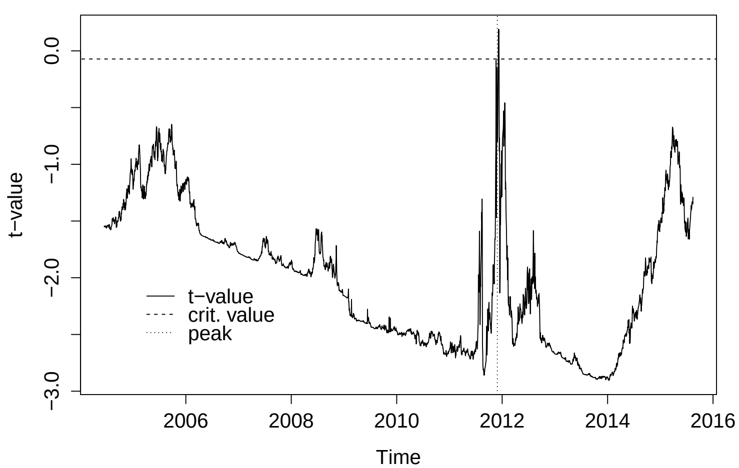

Fig. 11. Trajectory of $t_{\hat{\rho}_{IT_t}}$.

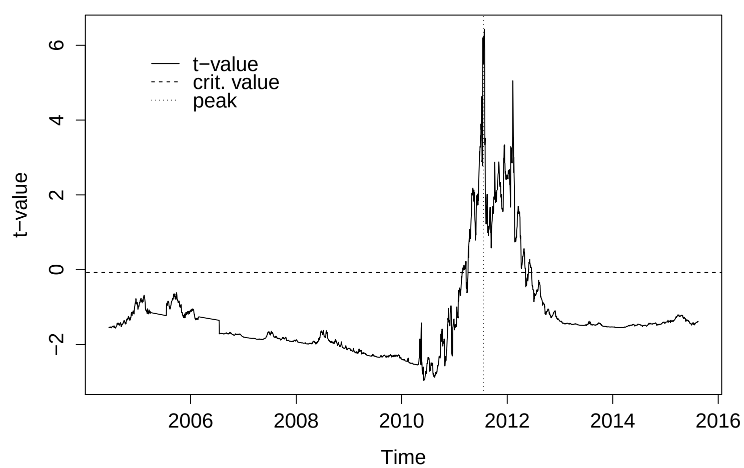

Fig. 12. Trajectory of $t_{\hat{\rho}_{PT_t}}$.

Table 1: Value of the maximal autocorrelation coefficient.

Country	Date	$\max_t\{\hat{\rho}_t\}$
$\tilde{\Delta}GR_t$	12/22/2008	1.014
$\tilde{\Delta}IT_t$	11/3/2008	1.008
$\tilde{\Delta}ES_t$	1/23/2009	1.007
$\tilde{\Delta}PT_t$	7/13/2011	1.008

Table 2: Value of the maximal autocorrelation coefficient.

Country	Date	$\max_t\{\hat{\rho}_t\}$
GR_t	5/10/2010	1.015
IT_t	11/29/2011	1.000
ES_t	7/26/2012	1.000
PT_t	7/19/2011	1.009

The results indicate explosive behavior of the spreads between 2008 and 2009 apart from Portugal. Here, we have explosiveness in 2011. Furthermore, we apply the procedure to the interest rate time series as reported by table 2 and figure 9 to 12 to isolate the regimes after the subprime crisis. The t-value is not considerable larger than the critical value in the case of Italy and Spain. However, we see explosiveness for Greece and Portugal. Now, we apply the migration test as described in the foregone section to the REI_t and the spreads against Germany. Table 3 and table 4 show the results of the test.

Table 3: Migration test applied to spreads.

Country	Test statistic	$\hat{\beta}_1$
GR_t	10.338	-9.261
IT_t	6.971	-6.172
ES_t	5.199	-4.693
PT_t	6.807	-6.948

Table 4: Migration test applied to interest rates.

Country	Test statistic	$\hat{\beta}_1$
GR_t	0.622	-0.607
PT_t	0.934	-0.954

We see that the spread seems to be strongly affected by the explosiveness of REI_t. All results are significant on a level of 0.01. However, the test applied to the interest rates indicates no explosiveness migration. Thus, theses findings underline the hypothesis that the EMU crisis is a homemade problem by the countries while the first explosive regime was triggered by the bankruptcy of Lehman Brothers.

4.2. Robustness checks

We evaluate the robustness of our results in three ways: Firstly, we use an alternative way to stamp the dates of explosiveness by Harvey, Leybourne, and Sollis (2015). This is motivated by the fact, that the authors receive better power properties of this procedure compared to the recursive approach in order to stamp explosive regimes. Secondly, we use a procedure by Kruse and Wegener (2016) to test against strong dependent innovations in the explosive regime. The authors show that the right tailed unit root test has severe size distortions under strong dependent innovations. They propose adjusted critical values to overcome this problem and a procedure to test against fractional integrated residuals. Thirdly, we use an indirect inference estimator to estimate the autoregressive coefficient of model (2) as proposed by Phillips, Wu, and Yu (2011). This is motivated by the well known bias of the conventional OLS estimator in the vicinity of unity in small samples. Kruse and Kaufmann (2015) show by simulation, that the indirect inference estimator is a valuable alternative to other procedures.

Harvey, Leybourne, and Sollis (2015) consider the following data generating process for y_t with $t = 1, ..., T$,

$$y_t = \mu + u_t \tag{11}$$

$$u_t = \begin{cases} y_{t-1} + \epsilon_t \text{ for } t = 2, ..., [r_1 T] \\ (1+\delta_1) y_{t-1} + \epsilon_t \text{ for } t = [r_1 T] + 1, ..., [r_2 T] \\ (1-\delta_2) y_{t-1} + \epsilon_t \text{ for } t = [r_2 T] + 1, ..., [r_3 T] \\ y_{t-1} + \epsilon_t \text{ for } t = [r_3 T] + 1, ..., T \end{cases} \tag{12}$$

with $\delta_1 \geq 0$ and $\delta_2 \geq 0$, ϵ_t as an error term. This process has a unit root up to $\tau_1 = [r_1 T]$, followed by explosive behavior for $\delta_1 > 0$ up to $\tau_2 = [r_2 T]$, collapse of the explosiveness up to $\tau_3 = [r_3 T]$ for $\delta_2 > 0$ and finally unit root behavior until the end of the sample. The authors use a Bayesian Information Criterium to chose the optimal OLS estimated model from the following data generating processes:

1. $0 < r_1 < 1$, $r_2 = 1$: unit root, explosiveness to sample end
2. $0 < r_1 < r_2 < 1$,$r_2 = r_3$: unit root, explosiveness, unit root to sample end
3. $0 < r_1 < r_2 < 1$,$r = 1$: unit root, explosiveness, collapse to sample end
4. $0 < r_1 < r_2 < r_3 < 1$: unit root, explosiveness, collapse, unit root to sample end

As a slight modification, we allow stationary local-to-unity behavior for all regimes, because interest rates might be mean reverting in moderate economic times (see Sibbertsen, Wegener, and Basse, 2014). Furthermore, we use the estimated breakpoints to test against explosiveness using the t_ρ-statistic

$$t_\rho = \frac{\rho - 1}{\sigma_\rho} \tag{13}$$

of the regression model

$$y_t 1\left\{\tau_i < t < \tau_{i+1}\right\} = \mu + \rho y_{t-1} 1\left\{\tau_i < t < \tau_{i+1}\right\} + \epsilon_t 1\left\{\tau_i < t < \tau_{i+1}\right\} \tag{14}$$

where σ_ρ is the standard deviation of ρ. Kruse and Wegener (2016) account for short-run dynamics of the error term ϵ_t using the prewhitening procedure by Qu (2011). To test against strong dependent innovations, they suggest to use the test by Demetrescu, Kuzin, and Hassler (2008) with adjusted critical values depending on the estimate of ρ. If the results of this test show indications for strong dependent residuals, we employ a local Whittle estimator to estimate the degree of integration $I(d)$ with $d \in [0, 0.5)$. Kruse and Wegener (2016) suggest response curves depending on d to adjust the critical values of the right tailed unit root test.

Furthermore, it is well known that the OLS estimator has a bias in the region of unity in finite samples (see Phillips, Wu, and Yu, 2011; Kruse and Kaufmann, 2015). The authors suggest to use the indirect inference estimator

$$\hat{\rho}_H^{II} = \underset{\rho \in \Theta}{\operatorname{argmin}} \|\hat{\rho} - \frac{1}{H} \sum_{h=1}^{H} \hat{\rho}^h(\rho)\| \tag{15}$$

by Phillips, Wu, and Yu (2011). Here, $\hat{\rho}^h(\rho)$ is the OLS estimator from a simulated series with $AR(1)$ coefficient ρ. H is the number of available simulation paths, Θ is a compact parameter space and $\|\cdot\|$ is a distance metric. For $H \to \infty$ Phillips, Wu, and Yu (2011) obtain

$$\hat{\rho}_H^{II} = \underset{\rho \in \Theta}{\operatorname{argmin}} \|\hat{\rho} - q(\rho)\| \tag{16}$$

where $q(\rho) = E\left(\hat{\rho}^h(\rho)\right)$ is the binding function. Thus, the idea is to compare the estimates $\hat{\rho}$ from a grid of true values for ρ with it's average OLS estimates. The indirect inference estimator leads to the minimal distance between $\hat{\rho}$ and the average OLS estimator. See Kruse and Kaufmann (2015) for details and simulation studies for different bias correction procedures.

Table 5: Results of the robustness check.

	$\hat{\rho}$	$\hat{\rho}^{II}$	Breakpoints		ALM	t_ρ	prewhitening
REI_t	1.0188	1.0192	10/19/2006	11/20/2008	0.1889**	2.4700***	$ARFIMA(1, 0.07, 1)$
$\tilde{\Delta}GR_t$	1.0003	1.0054	3/23/2007	4/22/2009	-1.0385	0.1297**	$ARFIMA(1, 0, 1)$
$\tilde{\Delta}IT_t$	1.0048	1.0075	5/16/2006	6/13/2008	-0.9015	13.9398***	$ARFIMA(1, 0, 1)$
$\tilde{\Delta}ES_t$	1.0004	1.0055	8/25/2006	9/25/2008	-0.5721	0.0961**	$ARFIMA(1, 0, 1)$
$\tilde{\Delta}PT_t$	0.9983	1.0047	5/11/2010	7/25/2012	2.8503***	-2.3465	$ARFIMA(1, 0.30, 1)$
GR_t	1.0024	1.0063	1/8/2010	3/8/2012	-0.7464	1.1105***	$ARFIMA(1, 0, 1)$
IT_t	0.9977	0.9997	11/8/2012	7/31/2015	-0.7479	-1.9408	$ARFIMA(1, 0, 1)$
ES_t	0.9892	0.9996	12/31/2012	7/31/2015	-0.9774	-1.9102	$ARFIMA(1, 0, 1)$
PT_t	0.9878	0.9983	11/19/2012	7/31/2015	-0.8616	-2.0848	$ARFIMA(1, 0, 1)$

Table 5 reports the results of the robustness checks. Firstly, all the beginnings and the ends of all explosive periods are confirmed by the procedure proposed by Harvey, Leybourne, and Sollis (2015). Furthermore, the second explosive regimes, as indicated by the recursive right tailed unit root test for the spreads, seem to be not explosive using this alternative stamping procedure. In order to conserve space, we reported the respective regimes with the highest autoregressive coefficient. Secondly, in the case of REI_t and $\tilde{\Delta}PT_t$ we consider indications for strong dependent innovations. The result that REI_t shows explosive behavior is not affected while the result of the right tailed unit root test for $\tilde{\Delta}PT_t$ does not longer indicate explosiveness. However, the consistent estimation of ρ and the results of the migration test are not impaired by strong dependent innovations (see Magdalinos, 2012). Thirdly, the indirect inference estimator indicates explosiveness for all spreads and the government bond yield from Greece. Thus, the required condition of migrations of explosiveness from REI_t is fulfilled for the all spreads and the yield from Greece. To summarize the results of the robustness checks: The results of the procedure by Harvey, Leybourne, and Sollis (2015) and the indirect inference estimator underline the findings of the recursive right tailed unit root test. Strong dependent innovations are only relevant in the case of the spread between Portugal and Germany. However, also the results of the recursive approach indicate, that Portugal is a part of another story.

5. Conclusion

Employing a test procedure recently introduced by Phillips, Wu, and Yu (2011) we searched for explosive behavior in the US housing market and the EMU government bond market. With regard to EMU bonds, we considered the interest rates of a number of countries and the government bond yield spreads (with German government bond yields as benchmark) in order to distinguish between two explosive regimes: Firstly, explosiveness initiated by the bankruptcy of Lehman Brothers as a result of the bursting of the US housing market bubble and secondly, explosive behavior provoked by the EMU sovereign debt crisis. Further, we used a procedure of Phillips and Yu (2011) to investigate migration effects from the US house price bubble to EMU government bond yield spreads and to EMU interest rates. Examining the spreads our results seem to indicate that there are two crisis. The first crisis reflected in the yield differentials – as expected – is a result of the collapsing US housing market. Phrased somewhat differently, the results reported here indicate the existence of a migration process. Therefore, the first problems encountered in Europe most probably should indeed be regarded as a result of the US Subprime Crisis. However, the second crisis seems to be a consequence of the explosiveness of yields caused by the sudden appearance of sovereign credit risk and redenomination risk in EMU government bond markets. These observations indeed do suggest that the EMU debt crisis is a homemade problem. Consequently, the crisis originating in the US housing market really has moved from US mortgage baked securities to European banks and (most probably via bank rescue programmes) to EMU government bonds. It therefore really is some sort of a walking debt crisis. But this crisis does not seem to be a long distance runner because it most probably is not the cause of the EMU sovereign debt crisis. Thus, the results of our empirical investigations do support the point of view that the second part of the crisis was no second part but a major crisis of its own.

References

Basse, T., 2014. Searching for the EMU core member countries. European Journal of Political Economy 34, S32–S39.

Basse, T., Friedrich, M., Kleffner, A., 2012. Italian government debt and sovereign credit risk: An empirical exploration and some thoughts about consequences for European insurers. Zeitschrift für die gesamte Versicherungswissenschaft 101, 571–579.

Demetrescu, M., Kuzin, V., Hassler, U., 2008. Long memory testing in the time domain. Econometric Theory 24, 176–215.

Eichengreen, B., Mody, A., Nedeljkovic, M., Sarno, L., 2012. How the subprime crisis went global: Evidence from bank credit default swap spreads. Journal of International Money and Finance 31, 1299–1318.

Gómez-Puig, M., Sosvilla-Rivero, S., 2014. Causality and contagion in EMU sovereign debt markets. International Review of Economics & Finance 33, 12–27.

Gruppe, M., Lange, C., 2014. Spain and the European sovereign debt crisis. European Journal of Political Economy 34, S3–S8.

Harvey, D. I., Leybourne, S. J., Sollis, R., 2015. Improving the accuracy of asset price bubble start and end date estimators. In: *School of Economics, University of Nottingham Discussion Paper*.

Jenkins, M. A., Madzharova, P., 2008. Real interest rate convergence under the Euro. Applied Economics Letters 15, 473–476.

Johansen, S., 1988. Statistical analysis of cointegration vectors. Journal of economic dynamics and control 12, 231–254.

Johansen, S., 1991. Estimation and hypothesis testing of cointegration vectors in gaussian vector autoregressive models. Econometrica: Journal of the Econometric Society pp. 1551–1580.

Kim, S.-J., Moshirian, F., Wu, E., 2006. Evolution of international stock and bond market integration: Influence of the European Monetary Union. Journal of Banking & Finance 30, 1507–1534.

Kruse, R., Wegener, C., 2016. Testing for neglected strong dependence in explosive models. Tech. rep., Leibniz University Hannover.

Kruse, Y. R., Kaufmann, H., 2015. Bias-corrected estimation in mildly explosive autoregressions. Tech. Rep. 2013-10, CREATES Working Paper.

Laopodis, N. T., 2008. Government bond market integration within European Union. International Research Journal of Finance and Economics 19, 56–76.

Ludwig, A., 2014. A unified approach to investigate pure and wake-up-call contagion: Evidence from the Eurozone's first financial crisis. Journal of International Money and Finance 48, 125–146.

Lund, J., 1999. A model for studying the effect of EMU on European yield curves. European Finance Review 2, 321–363.

Magdalinos, T., 2012. Mildly explosive autoregression under weak and strong dependence. Journal of Econometrics 169, 179–187.

Phillips, P. C., Magdalinos, T., 2007. Limit theory for moderate deviations from a unit root. Journal of Econometrics 136, 115–130.

Phillips, P. C., Wu, Y., Yu, J., 2011. Explosive behavior in th 1990s NASDAQ: When did exuberance escalate asset values? International Economic Review 52, 201–226.

Phillips, P. C., Yu, J., 2009. Limit theory for dating the origination and collapse of mildly explosive periods in time series data. Singapore Management University Working Paper .

Phillips, P. C., Yu, J., 2011. Dating the timeline of financial bubbles during the subprime crisis. Quantitative Economics 2, 455–491.

Qu, Z., 2011. A test against spurious long memory. Journal of Business & Economic Statistics 29, 423–438.

Sibbertsen, P., Wegener, C., Basse, T., 2014. Testing for a break in the persistence in yield spreads of EMU government bonds. Journal of Banking & Finance 41, 109–118.

Varoufakis, Y., 2013. From contagion to incoherence towards a model of the unfolding eurozone crisis. Contributions to Political Economy 32, 51–71.