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Abstract

This article sheds light on the question whether arising sovereign credit risk in the EMU has been triggered by the US subprime crunch. By adapting recent econometric methodologies suggested in the related field of speculative bubbles, we find clear evidence for fast diverging (and even explosive) behavior of EMU government bond yields of peripheral countries relative to Germany during the financial *and* the European debt crisis. This might be caused by flight-to-quality effects to German government bonds coincident with the collapse of Lehman Brothers and by a loss of confidence in the fiscal stability of Greece, Ireland, Italy, Portugal and Spain during the European debt crisis. First, we find compelling evidence for bubbles in the Dow Jones Equity Real Estate Investment Trust (REITs) index which serves as a weekly measure of economic activity in the North American real estate sector. Second, in our main analysis, we test whether the collapsing bubble in the housing market triggered the diverging government bond yields during two crisis regimes. Our findings indicate that this was the case in the course of the financial, but not during the EMU sovereign debt crisis. These results suggest that the severe fiscal problems in peripheral countries are homemade rather than imported from the US.

JEL classification: C15, G12, H63.

Keywords: Sovereign Debt Crisis, Sovereign Credit Risk, Subprime Crisis, Bubbles, Explosive Behavior, Bubble Migration.

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1. Introduction

At the end of a multi-year rally, US house prices peaked in the year 2006. Back then US real estate assets were quite expensive. Kivedal (2013), for example, presented empirical evidence for the existence of a speculative bubble in US housing markets before 2007. Its collapse marked the beginning of a downward house price spiral. The availability of credit for potential buyers of real estate in the US decreased and borrowers experienced more and more problems to refinance their loans (see Emmons and Noeth, 2013; Bullard, Neely, and Wheelock, 2009). The resulting dramatic fall of US house prices had massive negative effects on the prices of US subprime mortgage-backed securities. These collateralized bonds had also been bought by financial institutions in Europe and Asia (Noeth and Sengupta, 2012). Therefore, not only US 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 this problem. At this point a central question (see Bullard, Neely, and Wheelock, 2009; Eichengreen, Mody, Nedeljkovic, and Sarno, 2012) seems to be how the Subprime Crisis – an issue in a rather small segment of overall US 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 played a critical role in the transmission of the crisis from the US to Europe and other parts in the world. Most importantly, banks were responsible for causing some additional fiscal problems in a number of European countries. In Ireland, for example, the fiscal problems of the government can be explained, inter alia, by the costs resulting from measures to stabilize the financial system of the country.

The abrupt increase of the importance of sovereign credit risk in Europe had major ramifications for the pricing of fixed income securities in one of the biggest bond markets of the world and for the financial system in Europe. Ludwig (2014b) identified structural breaks in the long-run cointegration relation between German government bond yields and interest rates from other European countries that might be explained by changes in sovereign credit risk. The author noted that the government bond yields of Greece, Ireland, Italy, and Spain started to converge to German interest rates and diverged later considering time series data from January 1995 to December 2012. Furthermore, he found empirical evidence for structural breaks in the case of Greece and Ireland in 2007 and 2010. The first structural change in 2007 might coincide with the US Subprime Crisis and the resulting bank rescue programs in Europe, the second break in 2010 might be a consequence of a phenomenon that could be related to the European sovereign debt crisis. Both structural breaks were characterized by faster diverging interest rates.

Sovereign credit risk and financial stability are closely related. In fact, Noeth and Sengupta (2012) noted that depositors are more likely to flee from countries with less financial resources to provide funding for deposit insurance schemes in times of crisis. Interestingly, this environment caused outflows from banks in Greece, Italy and Spain – countries in which the banking sector was less heavily exposed to the US real estate market – to financial institutions in Germany and France that invested more funds in US mortgage backed securities in general.

Moreover, there also were dangerous imbalances within the euro area between surplus nations like Germany with higher exports than imports and deficit nations like France, Greece and Spain with more goods and services imported than exported (see Holinski, Kool, and Muysken, 2012). These imbalances clearly have to be classified as an European problem. 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, argued that even without the credit crunch in the US and the subsequent events in 2008 something bad simply had to happen concerning the fiscal situation of some European countries. However, Lane (2012) argues that the financial crisis caused the re-evaluation of asset prices and prospects for growth. From this point of view the financial crisis in the US might not be the cause but a trigger for the EMU debt crisis.

In our work, we provide new empirical evidence regarding the question whether the financial crisis in the US sparked the arising sovereign credit risk in the European Union (EU). In particular, we ask whether the bursting US home price bubble sparked flight-to-quality effects to Germany and loss of investors confidence in the fiscal situation of the peripheral countries. We adapt recent time series econometric methodologies for explosive processes originating in the field of speculative bubbles. In particular, we rely on the popular right-tailed recursive unit root test suggested by Phillips, Shi, and Yu (2015a) and the bubble migration test by Phillips and Yu (2011).

In the related literature on bubbles, explosive prices decouple from their fundamental value during periods of exuberance. In our applications, we study (i) the Dow Jones Equity Real Estate Investment Trust (REITs) index as a measure of the US housing market and (ii) European Monetary Union (EMU) government bond yield spreads containing sovereign credit risk. The latter cannot be understood as an asset price in the sense of Blanchard and Watson (1982) (see Section 4). But they are closely related to Credit Default Swaps. Thus, explosiveness of government bond yield spreads indicates rapidly increasing sovereign credit (and redenomination) risk. As argued below in detail, the test by Phillips, Shi, and Yu (2015a) is also appropriate in this case. Finally, we consider migration effects of explosiveness from the US housing market to government bond yield spreads in Europe using a modification of the procedure by Phillips and Yu (2011) to the case of two explosive regimes.

Contrary to the current relevant literature (see Section 2) this study takes the existence of arising and collapsing bubbles in prices and explosiveness in sovereign bond yield spreads into account and tests directly for migration effects between both phenomena. For the best of our knowledge this is a completely new approach to identify linkages between crisis-related events.

The article is structured as follows: Section 2 gives a short review of the relevant literature. Section 3 presents a timeline of the financial and the EMU sovereign debt crisis. Section 4 gives an overview of the underlying economic theory in order to motivate the empirical modeling strategy. Section 5 describes the data and introduces methodologies to identify bubbles in the US housing market and explosiveness of EMU government bonds yield spreads. Furthermore, we present tests for migration effects of collapsing bubbles in the US to rapidly arising sovereign credit risk in Europe. Section 6 presents the empirical results of our testing approaches for explosiveness and migration effects and Section 7 concludes.

2. Literature Review

In this section we focus on four strands of the literature relevant for our research questions. Subsection 2.1 deals with the convergence process of government bond yields in the euro area as the consequence of the introduction of the Euro. Next, Subsection 2.2 summarizes studies which consider the divergence process of interest rates in the EMU in the course of the financial crises. These two strands of literature are considered to motivate the test for fast diverging government bond yields (see Subsection 5.2). In order to justify the migration test (see Subsection 5.3) and to put our results into the global context we focus on the literature (Subsection 2.3 and Subsection 2.4) that deals with contagion and global financial shocks.

2.1. On the Way to the European Monetary Union

The introduction of the Euro in 1999 was very important for the bond markets of the EMU countries since the new common currency eliminated exchange rate risk among the member states. Therefore, it is no surprise that Kim, Moshirian, and Wu (2006) showed that the Euro caused structural changes in bond and equity markets by strengthening market integration. More specifically, they modeled the relationship between financial market segmentation and three main economic channels and documented structural change in the estimated parameters. Lund (1999) argued that a binding time table for the introduction of the common currency existed before 1999. Consequently, the prospects of creation of a monetary union in Europe should already have affected fixed income markets before the introduction of the Euro.

The empirical evidence reported by Kim, Moshirian, and Wu (2006) seems to be compatible with this assumption. In this context it is worthwhile to note that Laopodis (2008) reported an increase in the correlation of the returns on Euro government bond indices after the introduction of the new currency. Using cointegration techniques, this study also identified the existence of two groups of EMU countries – a core group (including Germany and France) and some peripheral countries (including Italy and Ireland). In addition, Jenkins and Madzharova (2008) applied a similar approach and identified a long-run relationship between nominal government bond yields in the euro area after the introduction of the Euro. Trying to explain this empirical finding they argued that interest rates in EMU countries converged after the new currency regime had been established. Additionally, Frömmel and Kruse (2015) analyzed European government bond yield spreads (Belgium, France, Italy, the Netherlands versus Germany) applying an econometric methodology which allows to estimate the dates of structural breaks endogenously. The authors found convergence behavior of European interest rates in the course of the introduction of the Euro.

2.2. The Crises and Diverging Interest Rates

Meanwhile, the European debt crisis caused some concerns about sovereign credit risk and possibly even redenomination risk (which means the return of currency risk due to a possible breakdown of the EMU) in the market for fixed income securities. While crises of such a type were quite common in less developed economies this is a relatively new strand of literature examining industrialized countries. At this point two important areas of the literature have to be discussed.

First of all, there is a body of literature that focuses on government bond yield spreads which normally use interest rates from Germany as the measure of the risk-free rate. A notable example is Gruppe and Lange (2014). They argued that higher sovereign credit risk might have caused structural changes among government bond yields in Germany and Spain (which have been found to be cointegrated). In particular, they used a testing procedure to detect and date structural changes in the parameters of a Vector Error Correction Model (VECM). Basse, Friedrich, and Kleffner (2012) used the same methodology and documented structural changes in the relationship between government bond yields in Germany and Italy. They identified two breakpoints – the first structural break coincides with the US Subprime Crisis while the second might be related to the EMU debt crisis. Furthermore, Basse (2014) documented that government bond yields in Austria, Belgium, Finland and the Netherlands – countries that usually are considered not to have significant fiscal problems – are cointegrated with German government bond yields in a stable way, even during the crisis.

Gómez-Puig and Sosvilla-Rivero (2014) also searched for structural changes in EMU government bond markets. They used Granger-causality and endogenous breakpoint tests to examine the relationship between interest rates in different member countries of the currency union and they argued that more than half of the breakpoints identified can be linked to the Euro sovereign debt crisis. Moreover, Sibbertsen, Wegener, and Basse (2014) 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 also indicate that there are structural breaks. The persistence of the examined time series might have increased significantly during the crisis. This could be a sign of higher sovereign credit and probably even of redenomination risk.

2.3. The Crises and Financial Contagion

The second important strand of literature focuses on contagion.[1] Studies belonging to this group of studies originally also focused strongly on emerging market economies (see Forbes and Rigobon, 2002; Aloui, Aïssa, and Nguyen, 2011). Kaminsky and Reinhart (2000), for example, argued that the 1980s will be remembered as a period of systemic crisis in the emerging market countries and that especially Latin American economies fell like domino stones having to cope with high debt burdens, devaluations, banking crises and recessions.

The importance of international financial crises clearly rose in the age of globalization (see Summers, 2000; Moser, 2003). After the US Subprime crisis and the European sovereign debt crisis the concept of contagion now is also used to analyze financial and fiscal problems in developed countries (see Kalbaska and Gatkowski, 2012; Gómez-Puig and Sosvilla-Rivero, 2014). One focus of studies belonging to this literature is the default risk of countries. Generally speaking, contagion describes a situation where a crisis in one country spreads across borders and affects financial markets and the economy in other countries (see Kenourgios, Samitas, and Paltalidis, 2011).

The literature distinguishes between two important relevant types of contagion – namely "pure contagion" and "wake-up-call contagion". With regard to the question examined here, pure contagion could be defined as an environment where asset prices in two or more countries fall because of a crisis in one of the countries and where there is no good funda-

[1] See, for example, for empirical models of contagion Pesaran and Pick (2007) and Ludwig (2014a).

mental explanation for this movement of asset prices. This type of contagion could be the consequence of herding (see Beirne and Fratzscher, 2013).

Wake-up-call contagion describes a situation where new information from one country makes investors reassess the default risk of other similar countries (see Giordano, Pericoli, and Tommasino, 2013; Bekaert, Ehrmann, Fratzscher, and Mehl, 2014). In this case, the new information does not necessarily have direct implications for other countries. Italian banks were not very strongly invested in US mortgage backed securities but there was an outflow of deposits (see Noeth and Sengupta, 2012) because of fears with regard to the ability of the government to help the banking system in an emergency. Thus, the empirical evidence documented by Basse, Friedrich, and Kleffner (2012) could be interpreted as a consequence of wake-up call contagion.

2.4. Global Financial Shocks and Sovereign Debt in the EMU

In any case, the financial sector plays a major role trying to explain the European sovereign debt crisis. Empirical evidence reported by Ludwig and Sobański (2014) is compatible with an interpretation of the events that focuses on the banking industry in the euro area. Using rolling Granger causality tests, they showed that the fragility linkages in the banking sector of the euro area increased markedly with the outbreak of the US subprime crisis in the year 2007 and that with the problems in Greece the epicenter of spillover risk shifted from financial services firms in the periphery towards the banks in the core countries of the euro area.

In fact, it now seems to be generally accepted in the literature that the global financial system does matter trying to explain sovereign credit risk and the European sovereign debt crisis (see Ang and Longstaff, 2013; Kräussl, Lehnert, and Stefanova, 2016). Taking the perspective of economic historians Reinhart and Rogoff (2011), for example, noted that banking crises often precede or accompany sovereign debt crises. Moreover, examining data from the Credit Default Swap (CDS) market Acharya, Drechsler, and Schnabl (2014) documented that there is a two-way feedback between banking sector and sovereign credit risk in the euro area. They argued convincingly that the announcements of government sponsored bank rescue packages in the euro area caused an immediate widening of sovereign CDS spreads combined with a narrowing of bank CDS spreads and that then (in the post-bailout era) there emerged a significant co-movement between bank CDS and sovereign CDS. Similar results were also presented by Ejsing and Lemke (2011) who detected clear empirical evidence for a structural break in the relationship between bank and sovereign CDS premia after the bailouts in the European financial services sector had been announced. In this context, it is of interest that Ang and Longstaff (2013) reported empirical evidence indicating that systemic sovereign credit risk is highly correlated with financial market variables and does not appear to be directly caused by macroeconomic integration between the US and the euro area.

3. A Timeline of the Crises

In order to relate our empirical results (see Section 6) to the recent historical context of the US financial and the European sovereign debt crisis we propose the following summary and a tabular overview of some key events (see Table 1). While the classification seems to be quite clear in most cases, there are some global events caused contagion (or migration) effects from the US to Europe (for example the bank run of Northern Rock).

Date	Event	Related to
February 2007	Freddie Mac declares to withdraw from risky mortgage deals	US
July 2007	Bear Stearns liquidates two funds which were heavily invested in Mortgage Backed Securities	US
August 2007	BNP Paribas liquidates two funds which were heavily invested in Mortgage Backed Securities	US
September 2007	Northern Rock asks for emergency credits from the Bank of England	Global
August 2008	Lehman Brothers collapses	Global
September 2008	Fed supports American International Group	Global
January 2009	Bank of America takes over Merrill Lynch	Global
November 2009	Greece revises the public deficit from 6% to 12.7%	EU
December 2009	Fitch rates Greece from A- to BBB	EU
April 2010	Greece asks for help from the EU	EU
	Portugal asks for help from the EU	EU
May 2010	European Financial Stability Facility becomes enacted with 440 billions Euro	EU
November 2010	Ireland asks for help from the EU	EU
February 2012	Greece gets the second rescue package	EU
June 2012	Spain asks the EU to rescue domestic banks	EU
September 2012	Outright Monetary Transaction Program comes on stream	EU

Table 1: Summary of some key events related to European and US crises.

A central event during the outset of the Subprime Crisis is the announcement of Freddie Mac in February 2007 to withdraw from risky real estate business. Furthermore, in August 2007 BNP Paribas liquidated two funds which were heavily invested in Mortgage Backed Securities and on March 14, 2008 Bear Stearns – also invested in mortgage markets – announced that the liquidity situation deteriorated significantly during the last 24 hours. At the same time the Federal Reserve Bank (Fed) and JPMorgan Chase started a rescue program. On March 16, 2008 JPMorgan Chase bought Bear Stearns with funding of the Federal

Reserve Bank.

"Given the current exceptional pressures on the global economy and financial system, the damage caused by a default by Bear Stearns could have been severe and extremely difficult to contain. Moreover, the adverse effects would not have been confined to the financial system but would have been felt broadly in the real economy through its effects on asset values and credit availability." (Ben S. Bernanke, The economic outlook, Before the Joint Economic Committee, US Congress, April 2, 2008)

However, not only in the US but also in Europe banks got in trouble. One important event here is related to Northern Rock. This financial institution got problems in September 2007 as a consequence of the liquidity shortage during the Subprime Crisis which culminated in a bank run. In order to reassure scared investors the British finance minister declared that the Bank of England and the government will guarantee for deposits at Northern Rock. From this point in time decisions of an European administration were directly triggered by consequences of the Subprime Crisis.

Back in the US, in April 2008 the US government and the Federal Reserve provided rescue loans to Fannie Mae and Freddie Mac in July 2008 – both companies government sponsored enterprises. In the aftermath of these rescue programs the political pressure on the US government became too high. Therefore, the government and the Federal Reserve were unable to avert the Chapter 11 bankruptcy of Lehman Brothers in August 2008. In September 2008, Washington Mutual – a financial service company with the core business of mortgage loans – collapsed. From that date at the latest, the global financial system got into deep trouble.

Furthermore, in September 2008 the Fed gave a rescue credit to American International Group (AIG) in order to avoid a downgrade of AIG's rating and in November 2008 the US government supported AIG with 150 billion US Dollar. All these and the following measures were regarded as necessary to maintain the functionality of Credit Default Swaps all over the world. In order to avoid even more worse effects for the gobal financial system Bank of America took over Merrill Lynch in January 2009. Governments and central banks all around the world tried to cope with the major problems of the global banking system that resulted from a loss of confidence in the stability and soundness of financial institutions. In this environment banks implemented new and more stringent risk management strategies and as a consequence more or less stopped lending funds to each other. This situation caused a de facto freezing of money markets not only in the US.

Back in Europe, the financial stability of governments began to be questioned. In November 2009 Greece revised the public deficit from 6% to more than 12.7%. As a consequence Fitch rated Greece with BBB (former A-) in December 2009. This caused the first doubts about the credibility of Greece or even about the fiscal situation of all European peripheral countries. Due to higher capital costs Greece and Portugal asked for help from the EU in April 2010. Subsequently in May the European Financial Stability Facility became operational and Ireland asked for help in November. Afterwards, Greece received the second rescue package in February 2012 and Spain made a request for help from the EU in order to save the domestic banking system. In September 2012 the heretofore critical stage of this crisis ended with Mario Draghi's words "*whatever it takes*".

4. Bursting Real Estate Bubbles in the US and Sovereign Credit Risk in the EMU

In this section we review two fundamental economic theories and their empirical implications in order to identify potential connecting points between the bursting US house price bubble and the arising sovereign credit risk in the EMU. They form the basis of our empirical modeling and testing strategy.

4.1. Rational Bubbles

The fundamental equation of asset pricing (see Blanchard and Watson, 1982) is given by

$$P_t = E_t\left[\left(D_{t+1} + P_{t+1}\right) / \left(1 + R_{t+1}\right)\right]. \tag{1}$$

From this equation, the current price P_t is determined as the conditional expectation (based on the information set available in t) of future discounted price P_{t+1} and dividend payment D_{t+1}. Equation (1) is the standard model of asset price determination and is quite flexible to use. Most importantly, D_{t+1} can also be considered as coupon payment in the case of bonds and as rent in the case of real estate assets (see, for example, Engsted, Hviid, and Pedersen, 2016).

Phillips and Yu (2011) provide an extensive discussion of assuming a constant discount rate R and the effects of time-varying discount rates. As the time series properties of the price and dividend process do not change with a time-varying, stationary discount rate, we follow Phillips and Yu (2011) and proceed with a constant rate.

Thus, equation (1) simplifies to

$$P_t = \frac{1}{(1+R)} E_t \left[(D_{t+1} + P_{t+1}) \right]. \tag{2}$$

Recursive substitution until period k yields to

$$P_t = \underbrace{\sum_{i=1}^{k} \frac{1}{(1+R)^i} E_t \left(D_{t+i} \right)}_{F_t} + \underbrace{\frac{1}{(1+R)^k} E_t \left(P_{t+k} \right)}_{B_t} \tag{3}$$

whereby a unique solution is obtained if the transversality condition $\lim_{k \to \infty} \frac{1}{(1+R)^k} E_t \left(P_{t+k} \right) = 0$ holds. In this case, the price equals the fundamental part F_t and is a martingale process. Otherwise, the price contains a bubble component B_t which renders P_t being a submartingale process.

Note that in this framework of rational bubbles the price contains a bubble component if investors pay more for the asset than they know is justified by discounted expected future dividends because they assume that they can sell the asset for an even higher price in the future. With regard to the US real estate market it has to be noted in the context of rational bubbles that financial institution had no major problems with borrowers in financial difficulties as long as house prices were rising because real estate assets in this market environment could be sold with a profit (see Bullard, Neely, and Wheelock, 2009). Moreover, mortgage risk was "offloaded" to capital markets using mortgage backed securities (see Wachter, 2015).

This framework allows to discriminate between a time series integrated of order one (i.e. a martingale process) and an explosive time series (i.e. a submartingale process) by applying right-tailed unit root tests. Another common empirical strategy relies on the fact that the dividend and the price process should share a common stochastic trend in the absence of a bubble component. However, the recursive type of the right-tailed unit root tests is necessary to test for migration effects as described in Sections 5.3 and it is superior in the case of periodically collapsing bubbles (see Evans, 1991).

4.2. *Interest Rate Spreads*

Furthermore, we examine the fiscal problems in some European countries in more detail. The uncovered interest rate parity condition can be seen as the theoretical foundation to analyze the relationship between interest rates in different countries. Assuming that the bonds issued by the two sovereign states examined are homogenous with regard to credit risk and liquidity the uncovered interest rate parity condition demands that

$$1 + i_t = \left(\frac{E_t\left(s_{t+k}\right)}{s_t} \right) (1 + i_t^*) \tag{4}$$

with i_t as the domestic and i_t^* as the foreign interest rate. $E_t(s_{t+k})$ is the expected future spot exchange rate for tie $t+1$, given information up to time t. For small i_t^* we obtain

$$i_t - i_t^* \approx E_t\left(\Delta s_{t+k}\right). \tag{5}$$

Thus, the yield differential should approximately equal the expected exchange rate change $E_t(s_{t+k})$ when investors are risk neutral. Given the assumptions expected exchange rate movements and possibly a risk premium compensating investors that are risk averse for exchange rate risk are the only determinants of the yield spread.

We examine interest rates in EMU member states. Consequently, there is no exchange rate and the spread should be equal to zero when the bonds examined are considered to be homogeneous. Yield differentials therefore are a consequence from risk premia compensating investors for liquidity risk, sovereign credit risk and possibly even redenomination risk.

In the crisis the last two types of risk obviously are of major importance. Before the European government debt crisis certain differences with regard to the level of liquidity should have existed. Thus, even in this environment the spread was not zero all the time. But the yield differential should be at least stationary around zero and the interest rates (as long as both are integrated of order one) therefore under certain conditions ought to be cointegrated with the constant vector of $(1, -1)$ (see Ludwig, 2014b). This behavior of converging government bond yields in the EMU is well documented in the course of the implementation of the Euro (e.g. Zhou, 2003; Frömmel and Kruse, 2015).

However, Sibbertsen, Wegener, and Basse (2014) reported a break in the persistence of the spreads from mean reverting to unit root behavior coincident with the start of the financial crisis. In addition, the approach by Ludwig (2014b) – which applies cointegration methodology – indicates divergence behavior.

This might be caused be arising sovereign risk and redenomination credit risk. Thus, equation (5) comes to

$$i_t - i_t^* \approx E_t\left(\Delta s_{t+k}\right) + RP_t \tag{6}$$

with RP_t compensating investors for sovereign credit risk and potentially for redenomination risk. Generally, if market participants realize the existence of sovereign credit risk suddenly, this risk premium might imply explosive behavior of the spread. Thus, it is appropriate to test the unit root hypothesis against an explosive alternative.

5. Data and Methodology

5.1. Data

For all series we use data at a weekly frequency ranging from January 05, 2001 to June 24, 2016 taken from the Bloomberg Database. The resulting number of observations for each single time series is therefore $T = 808$.

We investigate generic government bond yields with a maturity of 10 years[2] for Germany, Greece, Ireland, Italy, Portugal and Spain. Our sample starts in January 2001 in order to ensure that neither the Euro introduction nor the dot-com bubble infer with our research objective. Spreads are computed as the simple difference of bond yields (in levels) between Germany and the European peripheral countries.

There is no useful aggregate index reflecting the movements of US house prices that is published on a weekly basis. Thus, we examine REITs. The Dow Jones Equity REIT Price Index is a popular measure for the performance of real estate investment trusts in the US. The index is based on stock price data from all publicly traded equity REITs on the three most important US stock exchanges NYSE, AMEX and NASDAQ. REITs are companies that invest in real estate assets (e.g., office buildings, apartments and shopping centers) or that lend money to owners of real estate assets. In the US there are three types of REITs (see Lee and Chiang, 2004).

While mortgage REITs directly lend money to owners of real estate assets or buy mortgage backed securities, equity REITs only invest in properties (see Colwell and Park, 1990). Hybrid REITs own real estate assets and also lend funds to owners of real estate assets. Equity REITs seem to be related to the property market in a very close way. Therefore, the time series examined here, which is a broad index consisting of US Equity REITs should be a very useful measure of economic activity in the North American real estate sector.[3]

Given that house price indices do not consider current rent payments but – at least from the perspective of corporate finance – are the discounted value of future expected rent income we examine the Dow Jones Equity REIT Price Index and not the Dow Jones Equity REIT Total Return Index. Thus, we ignore the effect of dividend payments. This does make sense

[2]We follow Ludwig (2014b) who noted that government bond yields with 10 years maturity are used to evaluate whether the Maastricht Treaty criterion of interest rate convergence in the EU is fulfilled. This is also appropriate in the context of this study.

[3]The literature demonstrates that cointegration between the prices of REITs and unsecurized real estate is a phenomenon of economic relevance (see Hoesli and Oikarinen, 2012). Therefore, private and public real estate markets in many countries are closely related to each other and REITs prices thus ought to be a useful measure of real estate market activity in general. With regard to the US (where explosiveness of house prices is a problem) we examine whether the Case-Shiller index and the REITs price index show co-explosive behavior. This hypothesis can not be rejected on a 5% level of significance.

because equity REITs in the US derive the majority of their revenues from rents and by law are required to distribute at least 90% of their net income as dividends to shareholders. Thus, the REIT price index should be a better proxy for house prices that the equivalent REIT total return index.

5.2. Testing Explosiveness

In the following, we present the econometric methodology for testing for explosiveness in a series y_t and migration of explosiveness from y_t (e.g., the house price series) to another time series x_t (e.g., government bond yield spreads). Before we can test for migration, we need to test for explosiveness in the single series y_t and x_t. In particular, we are interested in testing whether our time series show temporarily explosive behavior in the following form. We start for simplicity with the case of a single explosive regime which starts at time point t_e after a "regular" unit root regime during $t = 1, 2, ..., t_e - 1$. For notational convenience we suppress the intercept, but our empirical testing strategy accounts for possible drifts. In its simplest form, the data generating process is given as

$$y_t = \begin{cases} y_{t-1} + \varepsilon_t, & t < t_e = [Tr_e] \ , \quad \text{"non-explosive"} \\ \rho y_{t-1} + \varepsilon_t, & t \geq t_e = [Tr_e] \ , \quad \text{"explosive"} \\ y^*_{t_c} + \sum_{i=[r_e T]+1}^{t} \varepsilon_i, & t \geq t_p = [Tr_p] \ , \quad \text{"non-explosive with re-intialization"}. \end{cases} \tag{7}$$

After time t_p, the explosive regime collapses and the process y_t jumps to a different new initial value $y^*_{t_c}$. This model specification also allows for short mean reverting transitional dynamics to level $y^*_{t_c}$ and continuous as a unit root process (see Phillips and Yu, 2009; Phillips, Wu, and Yu, 2011). The autoregressive parameter ρ controls the persistence of the series. For the value of unity, the series is non-stationary but not explosive. On the contrary, in case of a slightly bigger coefficient than one, the process is mildly explosive defined via $\rho = 1 + \frac{c}{k_T}$ with the so-called "localization coefficient" $c > 0$; k_T is a sequence such that $k_T \to \infty$ as $T \to \infty$ which ensures only slight upward deviations from unity.[4] For the innovations ε_t, we impose the standard assumption of weak dependence.

In order to test the null hypothesis of non-explosiveness throughout the whole sample ($t = 1, 2, ..., T$) as an alternative to the described data generating process above, the corresponding

[4]Mild explosiveness resembles the local-to-unity framework for bridging highly persistent but stationary series to non-stationary series. It has been heavily used in finance, see for example Campbell and Yogo (2006) for an application in a predictive regression setting.

Dickey-Fuller test regression (see Phillips, Wu, and Yu, 2011) for $t = 1, 2, ..., T$ reads

$$y_t = \mu + \rho y_{t-1} + \sum_{j=1}^{J} \theta_j \Delta y_{t-j} + u_t. \tag{8}$$

This test regression contains J extra lags of the differenced series ($\Delta y_{t-j}; j \geq 1$) to capture additional serial correlation beyond the first lag. In practice, this test regression is estimated by Ordinary Least Squares (OLS) and J is selected via the Bayesian information criterion (BIC). We are interested in the hypothesis of a unit root process $H_0 : \rho = 1$ against the right-sided alternative $H_1 : \rho > 1$. For this purpose, the simple t-ratio $ADF = \frac{\hat{\rho}-1}{\hat{\sigma}_{\hat{\rho}}}$ can be used. As for the classic Dickey-Fuller test, critical values are non-standard and obtained via simulation (see Phillips, Wu, and Yu, 2011).

In this exposition, the time series y_t would be mildly explosive under the alternative throughout the whole sample. In order to allow for temporarily explosive behavior, we apply the method by Phillips, Shi, and Yu (2015a) to each of our time series. This test is an enhancement of the test by Phillips, Wu, and Yu (2011) which is based on the limit theory for moderate deviations from a unit root by Phillips and Magdalinos (2007).

First, in the setting of Phillips, Wu, and Yu (2011), the test regression (see equation (8)) is estimated in a recursive fashion leading to a sequence of ADF statistics. The ADF statistic computed on the full sample ($t = 1, 2, ..., T$) is denoted as ADF_0^1. The recursion parameter is commonly specified as $r_0 = 0.01 + 1.8/\sqrt{T}$ and the corresponding ADF statistic is denoted as $ADF_0^{r_0}$. During the recursion, the sample is extended by using the fraction $r_2 \in [r_0, 1]$ of observations. Thus, each regression involves a window of $[Tr_2]$. The supremum test statistic ($SADF$) is obtained as the supremum of the sequence of recursive ADF statistics:

$$SADF(r_0) = \sup_{r_2 \in [r_0, 1]} ADF_0^{r_2}. \tag{9}$$

Suppose, the null hypothesis of a non-explosive unit root behavior throughout the whole sample is rejected in favor of a (temporary) explosive regime. For determination of the unknown start and the end date of the explosive episode, Phillips, Wu, and Yu (2011) construct the following estimators r_e and r_f, respectively:

$$r_e = \inf_{r_2 \geq r_0} \left\{ r_2 : ADF_0^{r_2} > cv_{r_2}^{\beta_T} \right\} \text{ and } r_f = \inf_{r_2 \geq r_e + \gamma \log(T)/T} \left\{ r_2 : ADF_0^{r_2} < cv_{r_2}^{\beta_T} \right\}. \tag{10}$$

The user sets the parameter γ in order to specify a minimal duration condition for the length of the explosive regime. In particular, $\gamma \log(T)/T$ ensures that a very short episode of explosiveness after the origination is not necessarily considered for a collapse date stamping.

For example, in the case of 15 years of weekly data (780 observations), γ is set to the value of 2 if the duration should exceed a period of six weeks. Furthermore, $cv_{r_2}^{\beta_T}$ is the right-sided critical value with a significance level of β_T of the Dickey-Fuller statistic with $[Tr_2]$ observations.

The procedure by Phillips, Shi, and Yu (2015a) has been shown to have better power properties in the presence of two (or more) explosive regimes compared to the test by Phillips, Wu, and Yu (2011). This test deals with a double recursion of regression (see equation (8)). In contrast to the methodology of Phillips, Wu, and Yu (2011) the starting point of the recursion r_1 varies until the end of the sample. Thus, each regression involves $[T\,(r_2 - r_1)]$ observations. The test statistic is given by

$$GSADF(r_0) = \sup_{\substack{r_2 \in [r_0, 1] \\ r_1 \in [0, r_2 - r_0]}} BSADF_{r_1}^{r_2} \tag{11}$$

where $BSADF_{r_1}^{r_2}$ is backward supremum ADF statistic sequence (see Phillips, Shi, and Yu, 2015a, for details). Here, the recursion parameter is also specified as $r_0 = 0.01 + 1.8/\sqrt{T}$ and the used critical values are reported by Phillips, Shi, and Yu (2015a).

For stamping the start and the end of the explosive episodes, Phillips, Shi, and Yu (2015a) use the estimators

$$r_e = \inf_{r_2 \geq r_0} \left\{ r_2 : BSADF_0^{r_2} > scv_{r_2}^{\beta_T} \right\} \text{ and } r_f = \inf_{r_2 \geq r_e + \gamma \log(T)/T} \left\{ r_2 : BSADF_0^{r_2} < scv_{r_2}^{\beta_T} \right\} \tag{12}$$

where $scv_{r_2}^{\beta_T}$ is the right-sided critical value with a significance level of β_T of the backward supremum ADF statistic with $[T(r_2 - r_1)]$ observations.

Note that the date stamping technique by Phillips, Wu, and Yu (2011) generally underestimates the number of explosive periods in the presence of more than one mildly explosive regime while the procedure by Phillips, Shi, and Yu (2015a) is consistent even for more than two explosive periods (see Phillips, Shi, and Yu, 2015b). Hence, we focus on the application of the latter procedure by Phillips, Shi, and Yu (2015a) as we find at least two explosive regimes for each series, even when applying the less appropriate test by Phillips, Wu, and Yu (2011). By using the date stamping technique by Phillips, Wu, and Yu (2011), we apply reliable and consistent estimators for the start and ending periods of explosive regimes.

5.3. Testing the Migration of Explosiveness

Phillips and Yu (2011) propose a test statistic that makes use of the (implicit) recursive estimation of ρ as introduced in the previous section to test for migration of explosive behavior from one series to another. Hence, we have two time series y_t and x_t with mildly and timely limited explosive autoregressive regimes

$$x_t = \rho_x x_{t-1} + \varepsilon_{x,t} \quad \text{and} \quad y_t = \rho_y y_{t-1} + \varepsilon_{y,t}. \tag{13}$$

For $\varepsilon_{x,t}$ and $\varepsilon_{y,t}$ we impose also the assumption of weak dependence like for ε_t defined before.

In the following, we assume that explosiveness exists for a certain period in x_t and that this regime collapses before the end of the sample. This implies in essence a time-varying behavior of the autoregressive coefficients ρ_x and ρ_y; therefore, we use the notation $\rho_x(t)$ and $\rho_y(t)$ in the following. Formally, we have

$$\rho_x(t) = \begin{cases} 1, & t < t_{ex} = [Tr_{ex}] \ , & \text{"non-explosive"} \\ 1 + \frac{c_{ex}}{T^\delta}, & t \geq t_{ex} = [Tr_{ex}] \ , & \text{"explosive"} \\ 1 + \frac{c_x}{T} \cdot \widetilde{t}_m, & t \geq t_{px} = [Tr_{px}] \ , & \text{"collapse"} \end{cases} \tag{14}$$

with $\widetilde{t}_m = \frac{t - t_{px}}{m}$, $m = t_{py} - t_{px}$ and $0 < \delta < 1$. This explosive regime starts at t_{ex} and ends where the autoregressive parameter $\rho_x(t)$ peaks at t_{px} (with $t_{px} > t_{ex}$ as an identifying restriction). After the peak, the autoregressive parameter drops to unity and decreases thereafter. It is assumed that $c_{ex} > 0$ and that $c_x(\cdot) < 0$ is a negative localizing coefficient function which leads to the "explosive" and the "collapse" regimes.

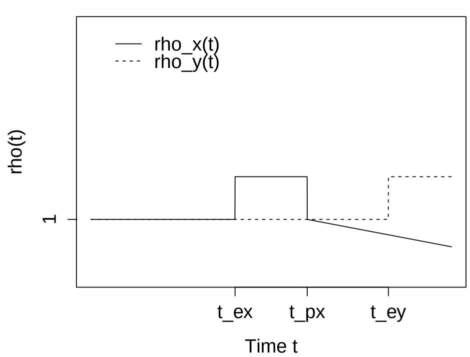


Fig. 1. The graph shows the stylized behavior of $\rho_x(t)$ (equation (14)) and $\rho_y(t)$ (equation (15) with $d = 0$) over time t in case of no migration from x to y.

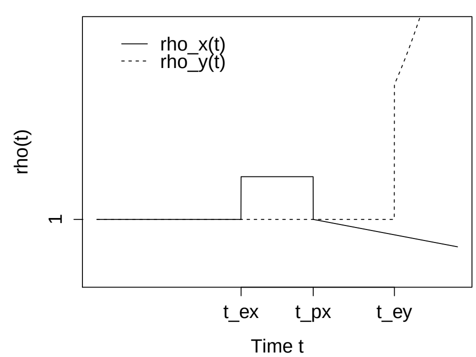


Fig. 2. The graph shows the stylized behavior of $\rho_x(t)$ (equation (14)) and $\rho_y(t)$ (equation (15) with $d < 0$) over time t in case of migration from x to y.

For y_t, an explosive regime starts at $t_{ey} > t_{px}$, that is after the beginning of the collapse in x_t. From t_{ey} onward, the autoregressive coefficient of y_t, i.e. $\rho_y(t)$, is composed of possibly two different explosive sources. In particular, we have the following structure for the autoregressive coefficient of y_t:

$$\rho_y(t) = \begin{cases} 1, & t < t_{ey} = [Tr_{ey}] \ , \quad \text{"non-explosive"} \\ 1 + \frac{c_y}{T^\delta} + d \cdot \frac{c_x}{T^\delta} \cdot \widetilde{t}_m^2, & t \geq t_{ey} = [Tr_{ey}] \ , \quad \text{"explosive"} \end{cases} . \tag{15}$$

After a "regular" unit root regime lasting until $t_{ey} - 1$, the series becomes explosive. In the explosive regime, the first component $\frac{c_y}{T^\delta}$ (with $c_y > 0$) is an "intrinsic" part which drives the explosiveness of y_t by itself. If the collapsing explosive regime in x_t contributes positively to the explosiveness in y_t via the term $d \cdot \frac{c_x}{T^\alpha} \cdot \widetilde{t}_m^2$ with negative localizing coefficient $c_x < 0$ (due to collapsing) and $d < 0$, we say that the explosiveness migrates from x_t to y_t. Importantly, the explosiveness in x_t fades away while the explosiveness in y_t arises. In the context of speculative bubbles, such a behavior is called bubble migration, while we adapt the formulation of migration of explosiveness. The parameter d measures the strength of the migration impact. In case of $d = 0$, there is no migration of explosiveness from x_t to y_t. When setting up a bubble migration test, Phillips and Yu (2011) test the null hypothesis $d = 0$ (no migration) against the alternative $d < 0$ (migration). Figures 1 and 2 visualize the trajectory of $\rho_x(t)$ and $\rho_y(t)$ under both hypotheses.[5]

In practice, the autoregressive coefficients are unknown and estimated recursively as outlined above in Section 5.2. Thereafter, the unknown breakpoints t_{px} and t_{py} are determined from the previously obtained series of recursively estimated autoregressive coefficients.[6]. Based on these quantities, a t-type statistic can be calculated from the following (infeasible) auxiliary test regression

$$[\rho_y(t) - 1] = \beta_0 + \beta_1 [\rho_x(t) - 1] \cdot \widetilde{t}_m + \epsilon(t), \tag{16}$$

for $t_{px} + 1, ..., t_{py}$. The test statistic for $H_0 : \beta_1 = 0$ is given by

$$Z_{\beta_1} = \frac{\widehat{\beta}_1}{L(m)}. \tag{17}$$

[5]It is not necessary to assume that the explosive regime in y lasts until the end of the sample. This is simply done for presentation purposes. Otherwise, one can add two more regimes ("collapse" and "non-explosive") at the end of the process.

[6]In detail, $t_{px} = \arg\max_t \rho_x(t)$ and $t_{py} = \arg\max_{t \in (t_{px}+1,...,t_{px}+\lambda T)} \rho_y(t)$ with λ being set equal to 1.2 defining a neighbourhood of observations similar to the bubble duration condition in equations (10) and (12).

Here, $L(m)$ is a slowly varying function depending on m.[7] As shown by Phillips and Yu (2011) this test is asymptotically conservative under the null hypothesis and consistent under the alternative hypothesis. We follow Phillips and Yu (2011) who suggest to set $L(m) = a \log(m)$ with $a = \{1/3, 1, 3\}$ to control the size-power tradeoff of the test. The parameter a can be seen as a tuning parameter, similar to many other time series testing situations. The t-type test statistic Z_{β_1} is compared to critical values from the standard normal distribution and rejects the H_0 if $|Z_{\beta_1}| > cv_{T,1-\alpha}$. In practice, unknown quantities are replaced by their estimated counterparts.

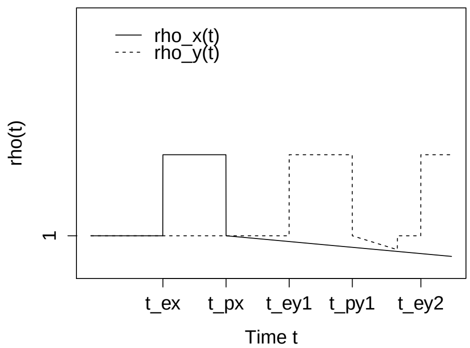


Fig. 3. The graph shows the stylized behavior of $\rho_x(t)$ (equation (14)) and $\rho_y(t)$ (equation (18) with $d_1 = 0$ and $d_2 = 0$) over time t in case of no migration from x to y at all.

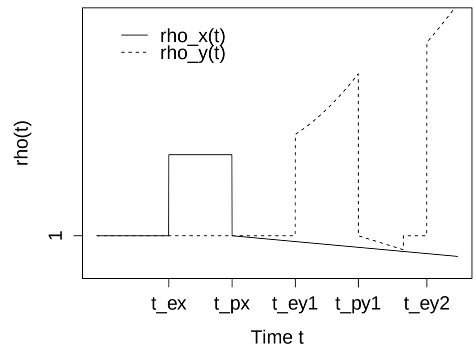
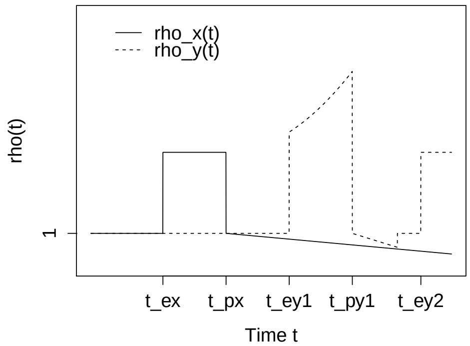


Fig. 4. The graph shows the stylized behavior of $\rho_x(t)$ (equation (14)) and $\rho_y(t)$ (equation (18) with $d_1 < 0$ and $d_2 < 0$) over time t in case of migration to both explosive regimes.

This test by Phillips and Yu (2011) allows the migration of one explosive regime from one series to another. However, we are interested whether a collapsing explosive regime in x may support *two* further explosive regimes in y. In this setting, $\rho_x(t)$ has the same properties under both hypotheses as in equation (14), but $\rho_y(t)$ now changes to

$$\rho_y(t) = \begin{cases} 1, & t < t_{ey,1} = [Tr_{ey,1}] \ , & \text{"non-explosive"} \\ 1 + \frac{c_{y,1}}{T^\delta} + d_1 \cdot \frac{c_x}{T^\delta} \cdot \widetilde{t}^2_{m_1}, & t \geq t_{ey,1} = [Tr_{ey,1}] \ , & \text{"explosive"} \\ 1 + \frac{c_y^*}{T} \cdot \left(\frac{t - t_{py,1}}{m^*}\right), & t \geq t_{py,1} = [Tr_{py,1}] \ , & \text{"collapsing"} \\ 1, & t \geq t_{fy} = [Tr_{fy}] \ , & \text{"non-explosive"} \\ 1 + \frac{c_{y,2}}{T^\delta} + d_2 \cdot \frac{c_x}{T^\delta} \cdot \widetilde{t}^2_{m_2} & t \geq t_{ey,2} = [Tr_{ey,2}] \ , & \text{"explosive"} \end{cases} , \tag{18}$$

[7]The slowly varying function needs to satisfy the condition $\frac{1}{L(m)} + \frac{L(m)}{T^e} \to 0$ as $T \to \infty$ for any $e > 0$.



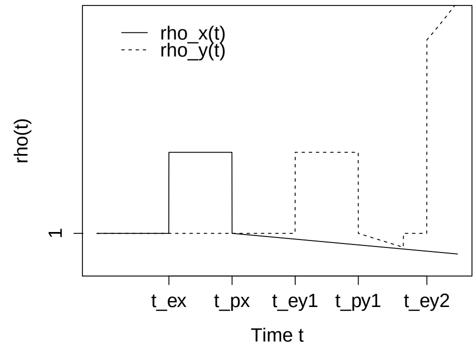


Fig. 5. The graph shows the stylized behavior of $\rho_x(t)$ (equation (14)) and $\rho_y(t)$ (equation (18) with $d_1 < 0$ and $d_2 = 0$) over time t in case of migration the first explosive regime only.

Fig. 6. The graph shows the stylized behavior of $\rho_x(t)$ (equation (14)) and $\rho_y(t)$ (equation (18) with $d_1 = 0$ and $d_2 < 0$) over time t in case of migration the second explosive regime only.

with $m_1 = t_{py,1} - t_{px}$, $m_2 = t_{py,2} - t_{fy}$ und $m = m_1 + m_2$. Here, the process starts with a non-explosive unit root. Then, it becomes explosive with its "intrinsic" part $\frac{c_{y,1}}{T^\delta}$ and a first migration of explosiveness $\frac{c_x}{T^\delta} \cdot \widetilde{t}^2_{m_1}$ from x_t to y_t if $d_1 < 0$. In case of $d_1 = 0$, there is no migration of explosive behavior from x_t to y_t in the first explosive regime of y_t.

After that, the process y_t collapses with c_y^* as a negative localization coefficient. Upon $t_{ey,2}$ the time series y_t contains a unit root and becomes finally explosive again with the "intrinsic" part $\frac{c_{y,2}}{T^\delta}$ and a second migration process $\frac{c_x}{T^\delta} \cdot \widetilde{t}^2_{m_2}$ as long as $d_2 < 0$. Again, for $d_2 = 0$, there is no migration during the second explosive regime.[8]

Thus, we have four possible situations:

1. No migration from x_t to y_t at all ($d_1 = 0, d_2 = 0$), see Figure 3.

2. Migration from x_t to both explosive regimes in y_t ($d_1 < 0, d_2 < 0$), see Figure 4.

3. Migration from x_t to the first explosive regime in y_t only ($d_1 < 0, d_2 = 0$), see Figure 5.

4. Migration from x_t to the second explosive regime in y_t only ($d_1 = 0, d_2 < 0$), see Figure 6.

[8]Similarly to the case of a single explosive regime in equation (14), it is not necessary to assume that the second explosive regime lasts until the end of the sample.

The auxiliary test regression accounting for two possible migrations is given by

$$[\rho_y(t)-1] = \beta_0 + \beta_1 \cdot \mathcal{I}_1 \cdot [\rho_x(t)-1] \cdot \widetilde{t}_{m_1} + \beta_2 \cdot \mathcal{I}_2 \cdot [\rho_x(t)-1] \cdot \widetilde{t}_{m_2} + \epsilon(t). \tag{19}$$

Here, $\mathcal{I}_1$ and $\mathcal{I}_2$ are dummy variables defined as follows. For $t = t_{px}+1, ..., t_{py,1}$, the first dummy variable $\mathcal{I}_1$ equals 1, otherwise it takes the value 0. Similarly, from $[Tr_{fy}]$ to $[Tr_{py,2}]$, the second dummy variable $\mathcal{I}_2$ equals 1 and 0, otherwise. Thus, these dummies indicate periods of explosiveness in y_t. In practice, we test the pairs of hypotheses

$$\begin{aligned} &H_0^1 : \beta_1 = 0 \text{ vs. } H_1^1 : \beta_1 < 0, \\ &H_0^2 : \beta_2 = 0 \text{ vs. } H_1^2 : \beta_2 < 0 \end{aligned} \tag{20}$$

using t-type test statistics from the test regression (19). The test statistics are adapted from the single migration case:

$$\begin{aligned} Z_{\beta_1} &= \frac{\widehat{\beta}_1}{L(m_1)}, \\ Z_{\beta_2} &= \frac{\widehat{\beta}_2}{L(m_2)} \end{aligned} \tag{21}$$

where $L(m_1)$ and $L(m_2)$ are slowly varying functions similar to $L(m)$ define above. Following Phillips and Yu (2011), we set $L(m_1) = a\log(m_1)$ and $L(m_2) = a\log(m_2)$ with $a \in \{1/3, 1, 3\}$. The null hypothesis is rejected if $|Z_{\beta_1}| > cv_{T,1-\alpha}$ respectively $|Z_{\beta_2}| > cv_{T,1-\alpha}$.

6. Empirical Results

In this section we present empirical results of the $GSADF$ (see Phillips, Shi, and Yu, 2015a) test applied to the logarithmic REITs index and interest rate spreads from Europe.[9] Furthermore, we stamp the start and the end of explosive regimes as described in Section 5.2. Main results of the migration tests are reported thereafter in Section 6.2.

6.1. Testing for Explosiveness

First of all, we apply the test procedure for explosiveness to the logarithmic REITs and the government bond yield spreads. As our findings indicate, see Table 2, there is compelling evidence for explosive behavior for all time series under consideration. As we have found

[9]As the application of the $SADF$ statistic does not provide any additional insights, we decide to save space and refrain from reporting those.

strong evidence in favor of explosive regimes in the data series (see $GSADF$ in Table 2), we turn our attention to the determination of their starting and ending dates. All results are summarized by Table 2 and are discussed in more detail in the following subsections.

6.1.1. House Prices

In addition to Table 2, Figure 7 shows the result of the $BSADF$ (see equation (12)) date stamping method applied to logarithmic REITs. It indicates periods of bubbles in US house prices between Oct 10, 2003 and Apr 10, 2009.

Most important for the migration analysis following in Subsection 6.2 is the "trigger" collapse of the house price bubble because it is the starting point of the migration analysis to explosive interest rate spreads in the EMU. We stamp this date to Mar 9, 2007. Earlier bubble regimes (Oct 10, 2003 to Apr 16, 2004 and Nov 19, 2004 to Jan 7, 2005) might be related to the same price rally which "final" collapse is located in 2007. Later explosive regimes (from Oct 24, 2008 to Dec 12, 2008 and from Feb 20, 2009 to Apr 10, 2009) might be related to speculative behavior during the financial crisis and thus, these bubble periods are part of another story.

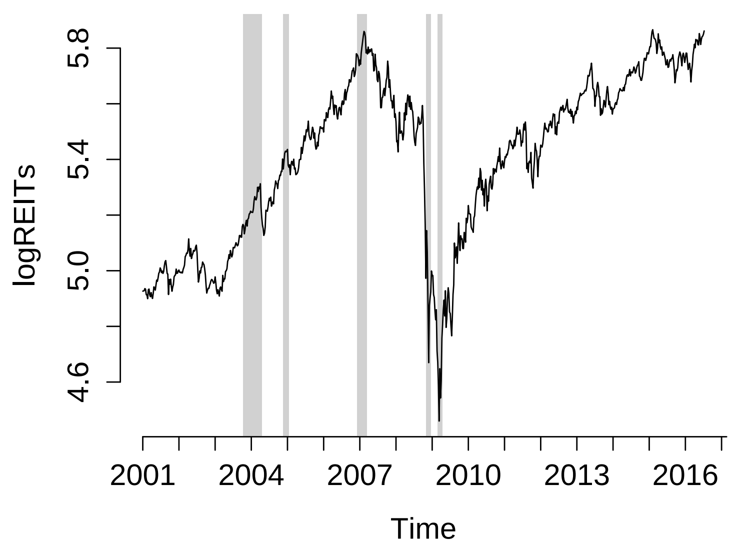


Fig. 7. The graph shows the logarithmic REITs from January 05, 2001 to June 24, 2016. Gray shaded areas indicate explosive regimes according to the results of the $BSADF$ date stamping (see equation (12)), see Table 3.

Series	REITs		Greece		Ireland	
GSADF	3.26^r		9.88^r		8.63^r	
Regime	Start	End	Start	End	Start	End
I	Oct 10, 2003	Apr 16, 2004	Sep 26, 2008	Sep 28, 2007	Apr 18, 2003	Aug 8, 2003
II	Nov 19, 2004	Jan 7, 2005	Jan 29, 2010	Mar 23, 2012	Apr 6, 2007	Dec 7, 2007
III	Dec 1, 2006	Mar 9, 2007	Apr 27, 2012	Jun 29, 2012	Dec 12, 2008	Jul 24, 2009
IV	Oct 24, 2008	Dec 12, 2008	Feb 6, 2015	May 15, 2015	Sep 24, 2010	Sep 2, 2011
V	Feb 20, 2009	Apr 10, 2009				
VI						
Series	Italy		Portugal		Spain	
GSADF	6.09^r		7.49^r		5.92^r	
Regime	Start	End	Start	End	Start	End
I	Mar 8, 2002	Apr 5, 2002	Dec 3, 2004	Feb 18, 2005	Aug 3, 2007	Sep 28, 2007
II	Sep 26, 2008	Apr 24, 2009	Mar 14, 2008	Apr 11, 2008	Nov 30, 2007	Jan 4, 2008
III	Jul 15, 2011	Feb 10, 2012	Oct 17, 2008	May 8, 2009	Mar 14, 2008	Apr 25, 2008
IV	May 25, 2012	Sep 21, 2012	May 14, 2010	Jul 30, 2010	Sep 12, 2008	May 8, 2009
V			Sep 3, 2010	Aug 31, 2012	May 14, 2010	Feb 3, 2012
VI			Apr 11, 2014	May 23, 2014	Apr 6, 2012	Oct 5, 2012

Table 2: This table shows the results of the recursive $GSADF$ (equation (11)) tests. The lag length is selected via BIC. The null hypothesis of non-explosiveness (at an unspecified point in time) is rejected whenever the test statistic exceeds its critical value. The initial recursion parameter is set to $r_0 = 0.01 + 1.8/\sqrt{T}$. The superscript ^{r} indicates a rejection of the null hypothesis at nominal significance level of five percent. Furthermore, results of the $BSADF$ date stamping procedure by Phillips, Shi, and Yu (2015a) (see equation (12)) are reported. For each series, the estimated start and end points of explosive regimes are given.

Therefore, we focus on the arising and collapsing house price bubbles between Oct 10, 2003 and Mar 9, 2007 which might have been caused by the same reasons. The most important might be the credit availability at this point time. It was very easy to get a property loan – even for low-income households. On the one hand, the Fed followed a policy of low interest rates. On the other hand, Fannie Mae and Freddy Mac bought mortgage loans from commercial banks. Thus, these banks were able to offer favorable terms and granted additional real estate loans with this released liquidity. However, in 2006 the Fed raised interest rates because of the impending inflation. Thus, the real estate financing became dramatically expensive. This might be one reason, inter alia, for the burst of the bubble in 2007. From this point in time, market participants realized that debt burdens for private households in the USA were too high.

"'It appears quite clear at this juncture,' said Joseph Brusuelas of IDEAglobal, 'that the consumer has reached a psychological point where expectations of future price declines have become entrenched. We consider this to be eminently rational behavior on the part of potential homeowners and until the new homes market observes a decline in the median price of homes and falling rates, there will be little incentive to step up purchasing activity.'" ("US November new home sales plunge 9 pct to 12-year low", Forbes, December 28, 2007)

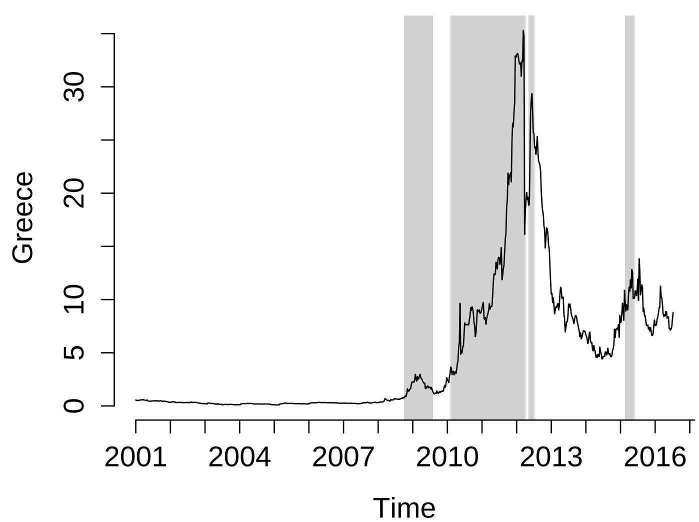


Fig. 8. The graph shows the spread between Greek and German government bond with 10 years maturity from January 05, 2001 to June 24, 2016. Gray shaded areas indicate explosive regimes according to the results of the $BSADF$ date stamping (see equation (12)), see Table 3.

6.1.2. *Greece*

While we focused on the bubble in the US housing market so far, we now consider explosiveness in government bond yield spreads in the EMU. As already mentioned, the $GSADF$ test indicates explosive behavior which might be related to fast growing sovereign credit risk for Greece.

The $BSADF$ break point estimator – see equation (12) – indicates essentially three explosive regimes as suggested by Figure 8. The first regime lasts from Sep 26, 2008 to Jul 24, 2009, while the second regime starts Jan 29, 2010 and ends Jun 29, 2012. Thus, the first explosive regime might be related to the bankruptcy of Lehman Brothers. The second phase of explosiveness might be related to the European sovereign debt crisis. A key event in this respect is Fitch's downgrade of Greece from A- to BBB as a consequence of the revision of the public deficit from 6% to 12.7% (see Section 3). In addition, the $BSADF$ statistic indicates also a third explosive regime from Feb 6, 2015 to May 15, 2015 which might be a consequence of the uncertainty of the Greek legislative election, January 2015.

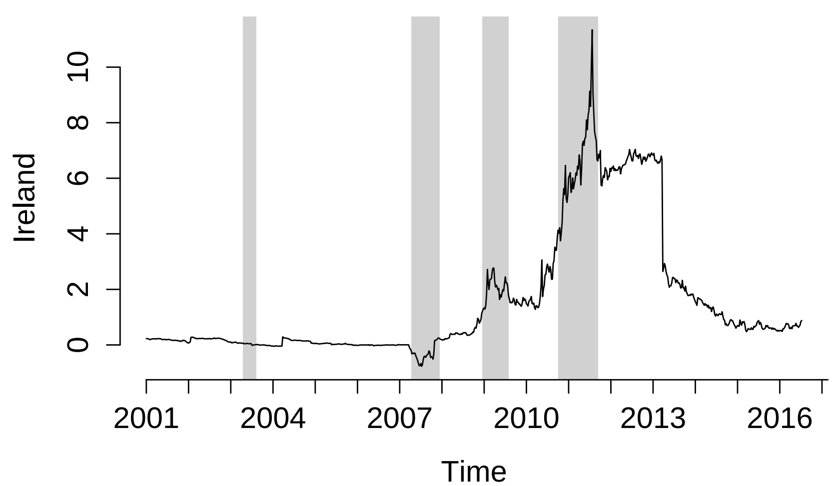


Fig. 9. The graph shows the spread between Irish and German government bond with 10 years maturity from January 05, 2001 to June 24, 2016. Gray shaded areas indicate explosive regimes according to the results of the $BSADF$ date stamping (see equation (12)), see Table 3.

6.1.3. Ireland

The results of the date stamping procedure are presented by Figure 9 for the spreads between Ireland and Germany.

Ireland is a special case (though there are some similarities to Spain). The country had to cope with an own housing bubble (see Wachter, 2015) that resulted in a banking crisis. The measures taken by the government in Dublin then caused severe fiscal problems. These financial difficulties then forced Ireland to ask for support from the EU. However, we obtain four explosive regimes using the date stamping procedure by Phillips, Shi, and Yu (2015a): The first starting Apr 18, 2003 and ending Aug 8, 2003, and the second begins Apr 6, 2007 and ends Dec 7, 2007 and the third starts Dec 12, 2008 and ends Jul 24, 2009 and the fourth start is stamped on Sep 24, 2010 and the end is indicated by Sep 2, 2011.

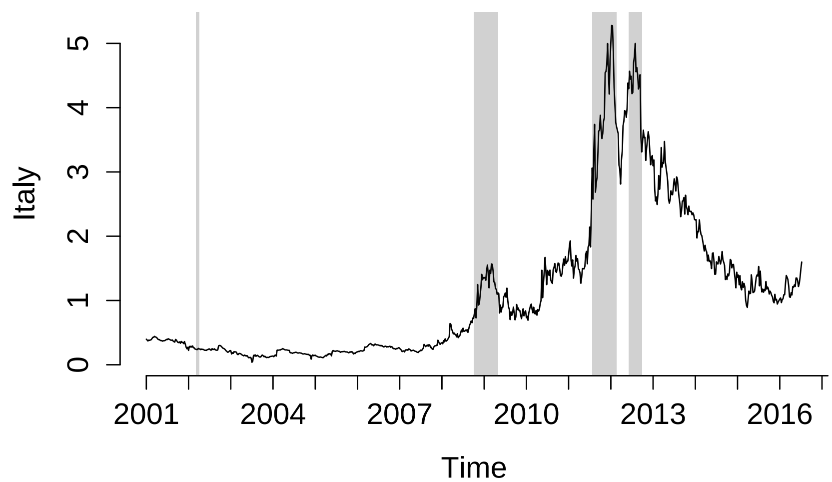


Fig. 10. The graph shows the spread between Italian and German government bond with 10 years maturity from January 05, 2001 to June 24, 2016. Gray shaded areas indicate explosive regimes according to the results of the $BSADF$ date stamping (see equation (12)), see Table 3.

6.1.4. Italy

Figure 10 reports the results of $BSADF$ break date estimation approach for spreads of the government bond yields between Italy and Germany.

While Italy did not need financial support from the EU there is (as discussed in Section 2) empirical evidence that could be interpreted as a sign for contagion. Our findings are in line with the results reported so far and indicate explosiveness of this spread between Sep 26, 2008 and Apr 24, 2009. Furthermore, we find two explosive regimes between Jul 15, 2011 and Sep 21, 2012 which might be related to the sovereign debt crisis. A third phase of explosiveness is indicated between Mar 8, 2002 and Apr 5, 2002.

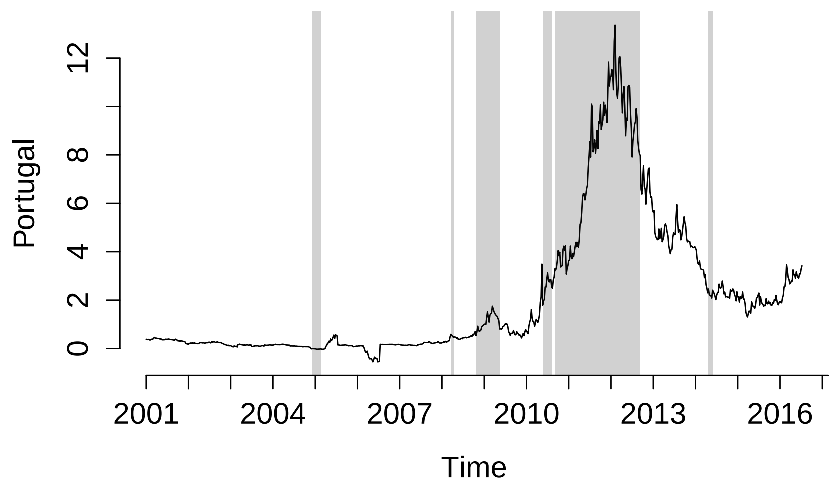


Fig. 11. The graph shows the spread between Portuguese and German government bond with 10 years maturity from January 05, 2001 to June 24, 2016. Gray shaded areas indicate explosive regimes according to the results of the $BSADF$ date stamping (see equation (12)), see Table 3.

6.1.5. Portugal

The results for spreads between Portugal and Germany are visualized by Figure 11.

The $BSADF$ approach stamps explosiveness for six regimes. We aggregate these six regimes to four phases of explosiveness: The first from Dec 3, 2004 to Feb 18, 2005, the second from Mar 14, 2008 to May 8, 2009, the third from May 14, 2010 to Aug 31, 2012 and the fourth from Apr 11, 2014 to May 23, 2014. The second might be related to the financial crisis while the third coincides with the sovereign debt crisis.

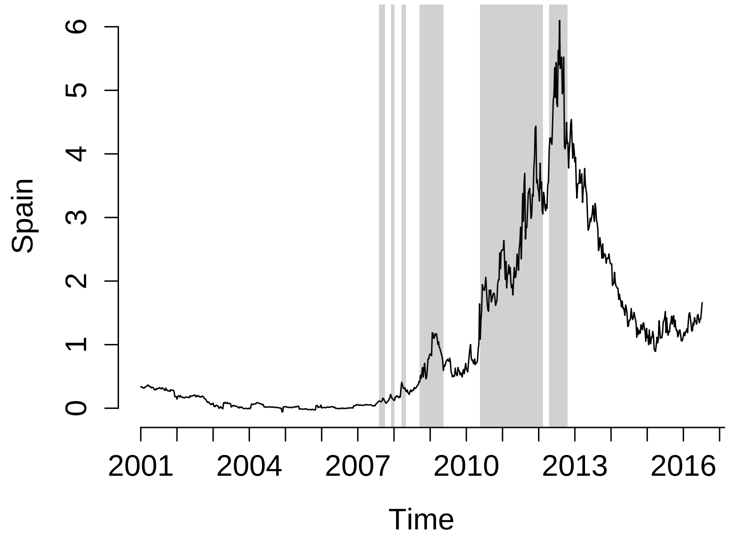


Fig. 12. The graph shows the spread between Spanish and German government bond with 10 years maturity from January 05, 2001 to June 24, 2016. Gray shaded areas indicate explosive regimes according to the results of the $BSADF$ date stamping (see equation (12)), see Table 3.

6.1.6. Spain

Figure 12 shows the results for the spread between Spanish and German government bond yields. Here, we also obtain essentially two phases of explosive regimes: From Aug 3, 2007 to May 8, 2009 – which might be related to the financial crisis – and from May 14, 2010 to Oct 5, 2012 – which is coincident with the sovereign debt crisis – using the date stamping procedure in equation (12). The "slowdown" in 2012 might be a consequence of the principal refinancing operations by the European Central Bank in favor of European Banks.

With the end of 2012, we do not consider further sustainable explosive regimes in the spreads apart from Greece and Portgal. Thus, the Outright Monetary Transaction Program and the famous words of Mario Draghi *"whatever it takes"* might have fulfilled its specific aim to strengthen investor's trust in the EMU and the improvement of the fiscal situation in European peripheral countries for our observation period.

6.2. Migration Analysis

This subsection presents the results of the main analysis of this study – the findings of the migration tests (see Subsection 5.3). Key ingredients to the analysis are the recursively estimated persistence parameters $\rho_x(t)$ and $\rho_y(t)$ with house prices $x = \{\text{REITs}\}$ and interest rate spreads $y = \{\text{Greece, Ireland, Italy, Portugal, Spain}\}$. As an initial analysis, we consider the plots (see Figures 13 to 17) of the persistence of the REITs ($\rho_x(t)$) in relation to the persistence of the respective EMU government bond spread ($\rho_y(t)$) for a particular country over time.

The graphs reveal that the persistence of the REITs decreases just before the persistence of the spreads increases – this might be related to a first regime of arising sovereign credit risk in the EMU in the aftermath of the bankruptcy of Lehman Brothers. In particular, we stamp the date of the "final" decline of the persistence of the REITs – which is the start date of the migration analysis – on February 16, 2007 because this is the local peak of the persistence of the REITs (see *Peak-to-peak* in Table 3). In addition, we stamp the date also by the *BSADF* date stamping procedure (see *BSADF* in Table 3).

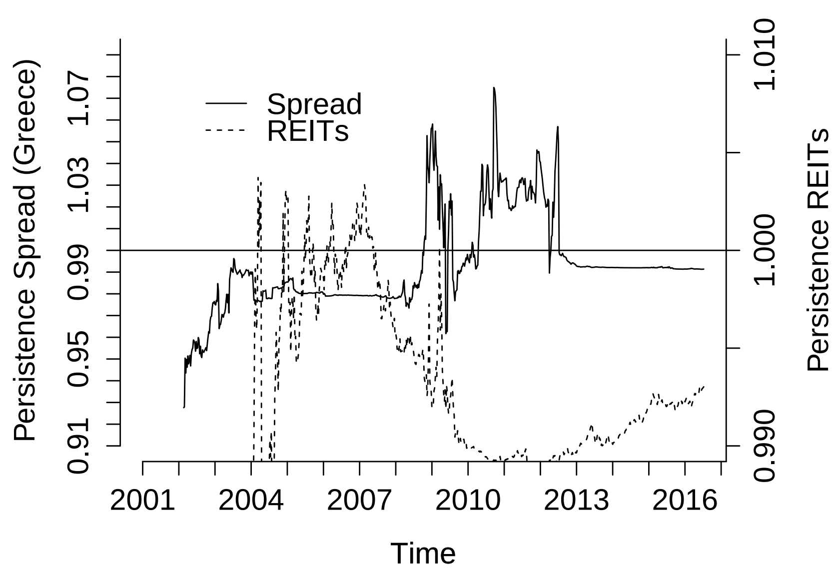


Fig. 13. The graph shows the recursively estimated persistence of the spread between Greek and German government bond yields (left scale) and of REITs (right scale) from January 05, 2001 to June 24, 2016.

Global financial crisis

Peak-to-peak

Series	$\widehat{\beta}_1$	Z_{β_1}	m	t_{px}	t_{py}
Greece	-7.04	-3.53^r	99	Feb 16, 2007	Jan 2, 2009
Ireland	-1.36	-0.67	106	Feb 16, 2007	Feb 20, 2009
Italy	-6.66	-3.36^r	96	Feb 16, 2007	Dec 12, 2008
Portugal	-4.46	-2.19^r	109	Feb 16, 2007	Mar 13, 2009
Spain	-4.36	-2.15^r	106	Feb 16, 2007	Feb 20, 2009

BSADF stamping

Series	$\widehat{\beta}_1$	Z_{β_1}	m	t_{px}	t_{py}
Greece	-6.52	-3.11^r	125	Mar 9, 2007	Jul 24, 2009
Ireland	-1.69	-0.81	125	Mar 9, 2007	Jul 24, 2009
Italy	-8.54	-4.17^r	112	Mar 9, 2007	Apr 24, 2009
Portugal	-5.44	-2.65^r	114	Mar 9, 2007	May 8, 2009
Spain	-4.46	-2.17^r	114	Mar 9, 2007	May 8, 2009

Table 3: This table reports the results of the single migration test statistic $Z_{\beta_1} = \widehat{\beta}_1 / \log_{10}(m)$ for $H_0 : \beta_1 = 0$ in the auxiliary regression $[\rho_y(t) - 1] = \beta_0 + \beta_1 [\rho_x(t) - 1] \cdot \widetilde{t}_m + \epsilon(t)$, for $t_{px} + 1, ..., t_{py}$ and $m = t_{py} - t_{px}$. The null hypothesis is "no migration of explosiveness" from REITs to interest rate spreads in Greece, Spain, Portugal, Italy and Ireland (separately) during the global financial crisis. The superscript ^{r} indicates a rejection of the null hypothesis at nominal significance level of five percent. In the upper panel, we report the results for the peak-to-peak determination of t_{px} and t_{py}, while the lower panel contains results where t_{px} and t_{py} are determined by the *BSADF* date stamping procedure, see Table 2.

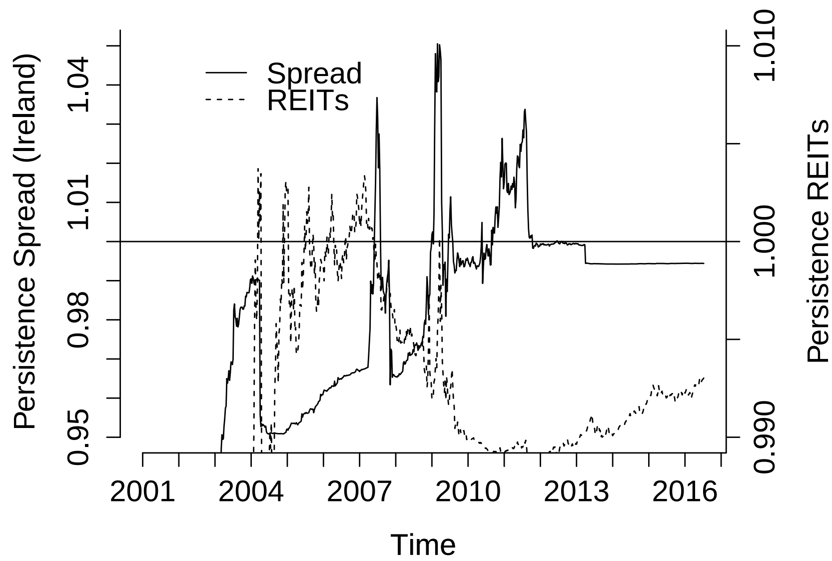


Fig. 14. The graph shows the recursively estimated persistence of the spread between Irish and German government bond yields (left scale) and of REITs (right scale) from January 05, 2001 to June 24, 2016.

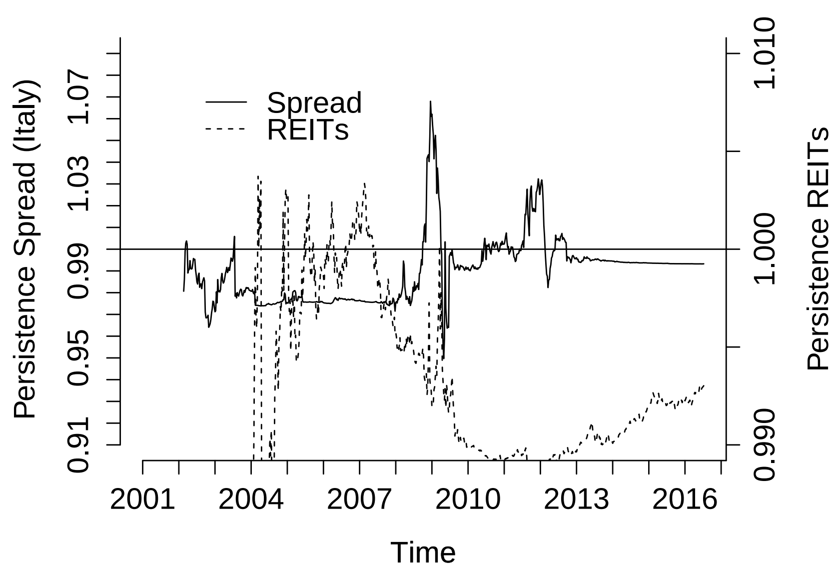


Fig. 15. The graph shows the recursively estimated persistence of the spread between Italian and German government bond yields (left scale) and of REITs (right scale) from January 05, 2001 to June 24, 2016.

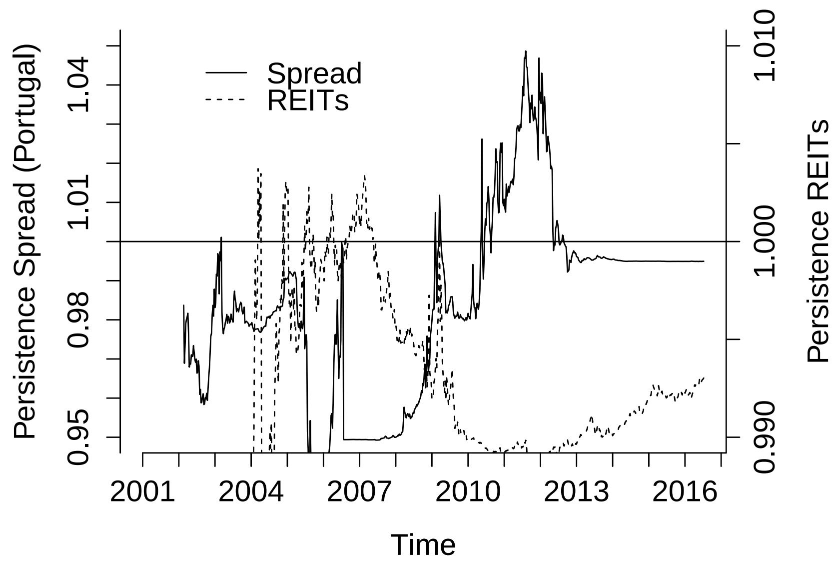


Fig. 16. The graph shows the recursively estimated persistence of the spread between Portuguese and German government bond yields (left scale) and of REITs (right scale) from January 05, 2001 to June 24, 2016.

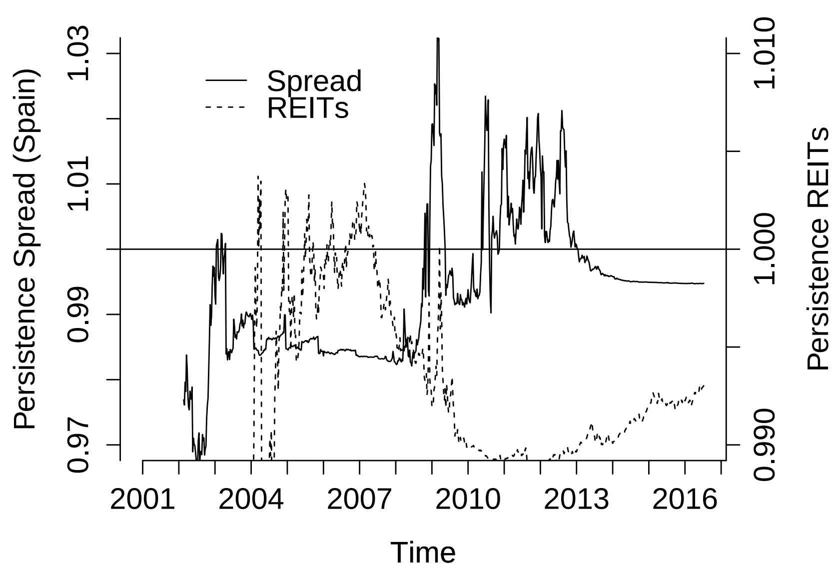


Fig. 17. The graph shows the recursively estimated persistence of the spread between Spanish and German government bond yields (left scale) and of REITs (right scale) from January 05, 2001 to June 24, 2016.

	Global financial crisis					Sovereign debt crisis				

Peak-to-peak

Series	$\widehat{\beta}_1$	Z_{β_1}	m_1	t_{px}	$t_{py,1}$	$\widehat{\beta}_2$	Z_{β_2}	m_2	$t_{py,1}+1$	$t_{py,2}$
Greece	-4.90	-2.45^r	99	Feb 16, 2007	Jan 2, 2009	-1.77	-0.91	88	Jan 9, 2009	Sep 10, 2010
Ireland	-0.39	-0.19	106	Feb 16, 2007	Feb 20, 2009	-1.46	-0.69	126	Feb 27, 2009	Jul 22, 2011
Italy	-3.82	-1.93^r	96	Feb 16, 2007	Dec 12, 2008	-1.16	-0.57	109	Dec 19, 2008	Jan 14, 2011
Portugal	-3.38	-1.66^r	109	Feb 16, 2007	Mar 13, 2009	-1.44	-0.81	61	Mar 20, 2009	May 14, 2010
Spain	-3.20	-1.58	106	Feb 16, 2007	Feb 20, 2009	-0.64	-0.35	69	Feb 27, 2009	Jun 18, 2010

BSADF stamping

Series	$\widehat{\beta}_1$	Z_{β_1}	m_1	t_{px}	$t_{py,1}$	$\widehat{\beta}_2$	Z_{β_2}	m_2	$t_{py,1}+1$	$t_{py,2}$
Greece	-6.71	-3.20^r	125	Mar 9, 2007	Jul 24, 2009	-3.02	-1.41	139	Jul 31, 2009	Mar 23, 2012
Ireland	-2.08	-0.99	125	Mar 9, 2007	Jul 24, 2009	-1.41	-0.69	110	Jul 31, 2009	Sep 2, 2011
Italy	-8.81	-4.30^r	112	Mar 9, 2007	Apr 24, 2009	-2.47	-1.14	146	May 1, 2009	Feb 10, 2012
Portugal	-5.13	-2.49^r	114	Mar 9, 2007	May 8, 2009	-1.71	-0.94	64	May 15, 2009	Jul 30, 2010
Spain	-4.47	-2.17^r	114	Mar 9, 2007	May 8, 2009	-1.69	-0.79	143	May 15, 2009	Feb 3, 2012

Table 4: This table reports the results of the double migration test statistics $Z_{\beta_1} = \widehat{\beta}_1/\log_{10}(m_1)$ for $H_0 : \beta_1 = 0$ and $Z_{\beta_2} = \widehat{\beta}_2/\log_{10}(m_2)$ for $H_0 : \beta_2 = 0$ in the auxiliary regression $[\rho_y(t) - 1] = \beta_0 + \beta_1 \cdot \mathcal{I}_1 \cdot [\rho_x(t) - 1] \cdot \widetilde{t}_{m_1} + \beta_2 \cdot \mathcal{I}_2 \cdot [\rho_x(t) - 1] \cdot \widetilde{t}_{m_2} + \epsilon(t)$. The null hypothesis is "no migration of explosiveness" from REITs to interest rate spreads in Greece, Spain, Portugal, Italy and Ireland (separately) during the global financial crisis and the sovereign debt crisis. The superscript ^{r} indicates a rejection of the null hypothesis at nominal significance level of five percent. In the upper panel, we report the results for the peak-to-peak determination of t_{px} and t_{py}, while the lower panel contains results where t_{px} and t_{py} are determined by the *BSADF* date stamping procedure, see Table 2.

Then, after a short decline of the persistence of the spreads – which marks the country specific first endpoint of the migration analysis – it increases again. This might be related to a second regime of arising sovereign credit risk in the EMU as a result of the sovereign debt crisis. Depending on the respective country the persistence decreases one more time. This marks the second endpoint of the migration analysis. We use the two tests as presented in Subsection 5.3 in order to carry out the migration analysis from decreasing persistence of REITs to both regimes of increasing persistence of the spreads.

Table 3 reports the findings of the test for a single migration process in equation (16). The null hypothesis of no migration from the collapse in REITs to explosive regimes of the EMU government bond yield spreads coincident with the bankruptcy of Lehman Brothers is rejected for a significance level of 5% and $a = 1$ in the cases of Greece, Italy, Portugal and Spain for both date stamping procedure (*Peak-to-peak* and *BSADF* in Table 3).

Only in the case of Ireland we find no indications for migration processes. This might be caused by the fact that Ireland had its own housing bubble as mentioned in Subsection 6.1.3. However, this results indicate that the first explosive regime in EMU government bond spreads has been triggered by the collapse of the US house price bubble and therefore, there might have been a "walking debt crisis" from US private households to European sovereigns.

However, the findings presented in Subsection 6.1 indicate a second explosive regime. Thus, we test for migration effects to both regimes using the auxiliary regression in equation (19). All results are presented by Table 4.

Using the same values for the tuning parameter a and both date stamping procedures we confirm the results from above. We consider migration effects from the collapsing bubble in the US to the first explosive regime of EMU government bond yields (Ireland is still an exception) for $a = 1$ and a level of significance of at least 5%. However, the hypothesis of no migration to the second regime is rejected for $a = 1$ in every case and for both date stamping procedures (*Peak-to-peak* and *BSADF*). Thus, the migration effects might have stopped after the first regime of arising sovereign credit risk in the EMU.

In the above discussion of test results, we have set the value of a equal to one as an intermediary choice when computing the Z-statistics. Variation of this tuning parameter affects the value of test statistic and potentially the test decision. In most cases, the parameter can be varied in certain limits without affecting the results. For instance, in the case of Greece with statistics $\widehat{\beta}_1 = -4.90$ and $\widehat{\beta}_2 = -1.77$ (peak-to-peak), the result is robust for the ranges $a \in [1/3, 3/2]$ and $a \in [1/3, 9/5]$, respectively. These intervals are relatively wide showing that the results are robust. In other cases, like Spain with $\widehat{\beta}_1 = -3.20$ and a statistic of $Z_{\beta_1} = -1.58$, a slight variation of a may lead to a rejection of the null hypothesis which we would expect.

Our study provides an analysis of the exact timing of potentially relevant migration effects. Therefore, it could be argued that our empirical research strategy might be problematic because it may suffer from the post hoc ergo propter hoc fallacy (see Hoover and Perez, 1994). This is a problem that already has been discussed in the contagion literature (see Moser, 2003). Causality indeed seems to be a philosophical concept that creates some difficulties in applied econometric work. Hoover and Perez (1994) argued convincingly that a narrative approach can be quite helpful for empirical economists trying to cope with "true" causality.

In fact, we believe that the interpretation of our empirical results does not suffer from the post hoc ergo propter hoc fallacy because there is some reason to assume that the US economy seems to be quite independent from the crisis events in Europe. Most importantly, economic growth in the US started to recover in the 2009 and was positive in the 3rd and 4th quarter of this year.

Moreover, numerous different measures of US house prices already show increasing values before the sovereign debt crisis in Europe gained momentum. Figure 18, for example, shows annualized real economic growth rates and real residential property prices in the US.

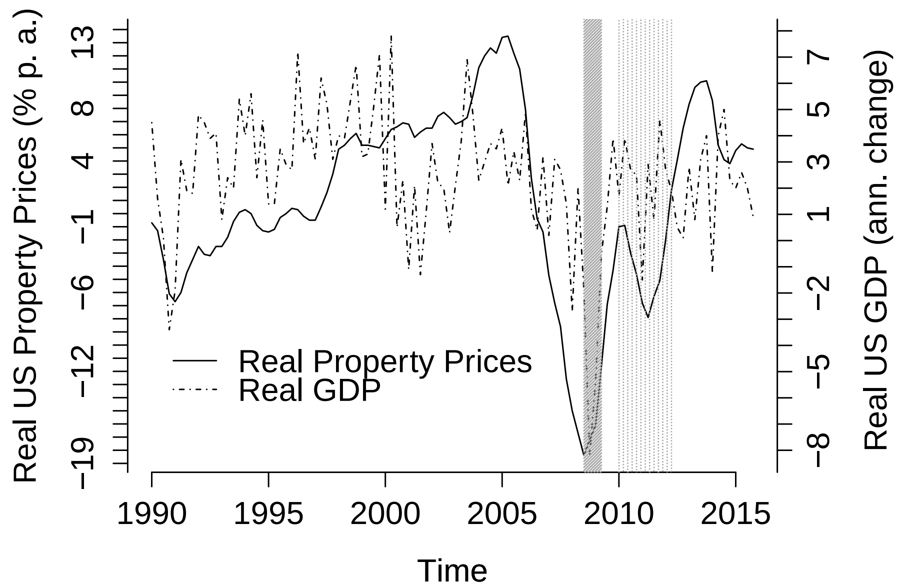


Fig. 18. The graph shows that economic growth and house prices in the US (here in real terms – data taken from the FRB of St. Louis) started to rise again before the soverign debt crisis in Europe gained momentum.

Thus, based on a narrative approach it can be argued that economic activity and house prices in the US seem to be exogenous variables and that (as a consequence) there should be no reverse causation among REITs and the European government bond market. This important observation ought to imply that our approach is not based on a problematic post hoc ergo propter hoc reasoning.

7. Conclusions

Employing a popular test introduced by Phillips, Shi, and Yu (2015a) we searched for explosive behavior in the US housing and the EMU government bond market. With regard to EMU bonds, we considered spreads between government bond yields from European peripheral countries (Greece, Spain, Portugal, Italy and Ireland) and Germany. Concerning US housing market we applied this test also to REITs. The results of this test indicated that there was a house price bubble collapsing in the second half of 2007. Furthermore, we found two explosive regimes in EMU government bond yield spreads. One started in September 2008 and collapsed at the beginning of 2009 and one arose in the beginning of 2010 and collapsed during 2012. Furthermore – as the essential innovation of this study – we used a recent testing procedure by Phillips and Yu (2011) to investigate migration effects of explosiveness from the US house price series to EMU government bond yield spreads.

The results indicate the existence of a migration process with regard to the first explosive regime of EMU government bond yield spreads. Therefore, the first problems encountered in Europe most probably should indeed be regarded as a result of the US Subprime Crisis. However, we did not consider a statistically significant migration effect to the second explosive regime.

These observations indeed suggest that the EMU debt crisis is a homemade problem. The crisis originated in the US housing market moved from US mortgage backed securities to European banks and (most probably via bank rescue programs and flight-to-quality effects) to EMU government bonds. The first period of arising sovereign credit risk in the EMU therefore really is some sort of a walking debt crisis but was not a long distance runner because it most probably was not the trigger of the EMU sovereign debt crisis. Thus, the results of our empirical investigations support the point of view that the European sovereign debt crisis has not been sparked by the collapsing US house price bubble.

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