The Walking Debt Crisis

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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 Trusts (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 crisis, 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 problems refinancing 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 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 (see Bullard, Neely, and Wheelock, 2009; Eichengreen, Mody, Nedeljkovic, and Sarno, 2012). 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 of 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 consequences for the pricing of fixed income securities in one of the biggest bond markets in 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 were also 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. 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 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 triggered the arising sovereign credit risk in the European Union (EU). In particular, we ask whether the bursting US home price bubble lead to flight-to-quality effects to Germany and a 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 Trusts (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 reflects rapidly increasing sovereign credit (and redenomination) risk. As argued below in detail, the test by Phillips, Shi, and Yu (2015a) is 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. To the best of our knowledge, this is a 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 over 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 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 of tremendous importance for the bond markets of the EMU countries since the new common currency eliminated exchange rate risk among member states. Therefore, it is much in line that Kim, Moshirian, and Wu (2006) found that the Euro caused structural changes in bond and equity markets by strengthening market integration across the countries. More specifically, they modeled the relationship between financial market segmentation and three main economic channels and documented structural change in the estimated parameters. In addition, Lund (1999) argued that a binding time table for the introduction of the common currency existed before 1999. Consequently, the prospects of creating a monetary union in Europe should already have affected fixed income
markets before the introduction of the Euro via expectation formation (see Frömmel and Kruse, 2015). In this respect, the 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.

There is a body of literature that focuses on government bond yield spreads which normally use interest rates from Germany as a measure of the risk-free rate. A notable example is Gruppe and Lange (2014), who 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. 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

Another important strand of literature focuses on contagion. Studies belonging to this
body of research originally also focused strongly on emerging market economies (see Forbes
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 such studies is dedicated to the default risks of countries. Generally
speaking, contagion describes a situation where a crisis in one country spreads across borders
and affects financial markets as well as the economies of other countries (see Kenourgios,
Samitas, and Paltalidis 2011).

The literature distinguishes between two 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 fundamental explanation
for this movement of asset prices. This type of contagion could be the consequence of herding
(see Beirne and Fratzscher 2013).

For empirical models of contagion see Pesaran and Pick (2007), Ludwig (2014a) and Leschinski and
Bertram (2017).
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 Tommasin, 2013; Bekaert, Ehrmann, Fratzscher, and Mehl, 2014). In this case, the new information does not necessarily have direct implications for other countries. Italian banks did not have substantial investments 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 vital role trying to explain the European sovereign debt crisis. Empirical evidence reported by Ludwig and Sobañski (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 problems in Greece the epicenter of spillover risk shifted from financial services firms in the periphery towards banks in the core countries of the euro area.

In fact, it is by now a generally accepted view in the literature that the global financial system matters when 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 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 during the post-bailout era a significant co-movement between bank CDS and sovereign CDS emerged. 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 following the announcement of the bailouts in the European financial services sector. In this context, it is of interest that Ang and Longstaff (2013) found 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 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 that caused contagion (or migration) effects from the US to Europe (for example the bank run of Northern Rock).

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

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 had 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 all deposits at Northern Rock. As a consequence, policies of European administrations had been triggered by the subprime crisis. This example underlines how US related problems were directly connected to the economic situation in Europe.

In the first half of 2008, the US government and the Federal Reserve provided rescue loans to Fannie Mae and Freddie Mac – both companies are 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 time onwards, 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 US$150 billion. 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 prevent even worse effects on the global 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%. In 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 oper-
ational 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

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

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

\[ P_t = \frac{1}{1 + R} E_t (D_{t+1} + P_{t+1}). \]
Recursive substitution until period $k$ yields to

$$P_t = \sum_{i=1}^{k} \left\{ \frac{1}{(1 + R)^i} E_t(D_{t+i}) + \frac{1}{(1 + R)^k} E_t(P_{t+k}) \right\}$$

(3)

whereby a unique solution is obtained if the transversality condition $\lim_{k \to \infty} \frac{1}{(1 + R)^k} E_t(P_{t+k}) = 0$ 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 institutions 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 us 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

$$1 + i_t = \left[ \frac{E_t(s_{t+k})}{s_t} \right] (1 + i^*_t)$$

(4)
with \(\Delta s_t\) as the domestic and \(\Delta s_t^*\) as the foreign interest rate. \(E_t(s_{t+k})\) is the expected future spot exchange rate for time \(t+k\) with \(k \geq 1\), given information up to time \(t\). For small interest rates we obtain

\[
i_t - i_t^* \approx E_t (\Delta s_{t+k}) .
\]

Thus, the yield differential should approximately equal the expected exchange rate change \(E_t(\Delta s_{t+k})\) when investors are risk neutral. Given the assumptions, expected exchange rate movements and possibly a risk premium compensating investors who 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 examined bonds are considered to be homogeneous. Therefore, yield differentials 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 are obviously of major importance. Before the European government debt crisis started, certain differences with regard to the level of liquidity should have existed. Even in this environment, the spread was not equal to zero at all time points. But the yield differential should be at least stationary and the interest rates (as long as both are integrated of order one) 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 reversion 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 by rising sovereign risk and redenomination credit risk. Thus, equation (5) becomes

\[
i_t - i_t^* \approx E_t (\Delta s_{t+k}) + RP_t
\]

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 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 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. However, we work with the Dow Jones Equity REIT Price Index which 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.

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

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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 As shown by Breitung and Kruse (2013) the maximum likelihood estimator yields March 2000 as the date for the burst of the dot-com bubble in the NASDAQ 100 index.

4 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.
REIT Total Return Index. Thus, we ignore the effect of dividend payments. This does
make sense 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. Hence, the REIT price index is considered as a better proxy for house prices
than the equivalent REIT Total Return Index.

5.2. Testing Explosiveness

In the following, we present the econometric methodologies for testing against 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 such migration, we
need to test against 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.
For simplicity, we first look at 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

$$
y_t = \begin{cases} 
y_{t-1} + \varepsilon_t, & t < t_e, \quad "non-explosive" \\
r y_{t-1} + \varepsilon_t, & t \geq t_e, \quad "explosive" \\
y^*_t + \sum_{i=[r_eT]+1}^{t} \varepsilon_i, & t \geq t_p, \quad "non-explosive with re-initialization". 
\end{cases}
$$

After time $t_p$, the explosive regime collapses and the process $y_t$ jumps to a different new
initial value $y^*_t$. This model specification also allows for short-live mean reverting transitional
dynamics to level $y^*_t$ and continues as a unit root process (see Phillips and Yu, 2009; Phillips,
Wu, and Yu, 2011).

The autoregressive parameter $\rho$ 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\footnote{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.}. For the innovations $\varepsilon_t$, 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
Dickey-Fuller test regression (see Phillips, Wu, and Yu, 2011) for \( t = 1, 2, \ldots, T \) 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 Augmented Dickey Fuller \( t \)-ratio \( ADF = \frac{\hat{\rho} - 1}{\hat{\sigma}_\rho} \) 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 \( (t = 1, 2, \ldots, T) \) is denoted as \( ADF_{10} \). The recursion parameter is commonly specified as \( r_0 = 0.01 + 1.8/\sqrt{T} \) and the corresponding \( ADF \) statistic is denoted as \( ADF_{0r_0} \). During the recursion, the sample is extended by using the fraction \( r_2 \in [r_0, 1] \) of observations. Thus, each regression uses a window of \( [Tr_2] \) observations. The Supremum Augmented Dickey Fuller (SADF) test statistic is obtained as the supremum of the sequence of recursive \( ADF \) statistics:

\[
SADF(r_0) = \sup_{r_2 \in [r_0, 1]} ADF_{0r_2} \tag{9}
\]

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 end date of the explosive episode, Phillips, Wu, and Yu (2011) construct the following estimators \( \hat{r}_e \) and \( \hat{r}_p \) for the relative time points:

\[
\hat{r}_e = \inf_{r_2 \geq r_0} \left\{ r_2 : ADF_{0r_2}^{r_2} > cv_{r_2}^{\beta_T} \right\} \quad \text{and} \quad \hat{r}_p = \inf_{r_2 \geq r_e + \gamma \log(T)/T} \left\{ r_2 : ADF_{0r_2}^{r_2} < cv_{r_2}^{\beta_T} \right\}. \tag{10}
\]

The user sets the parameter \( \gamma \) 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), $\gamma$ is set to the value of two 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 $\beta_T$ of the Dickey-Fuller statistic with $[T r_2]$ observations.

The Generalized Supremum Augmented Dickey Fuller (GSADF) test 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

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

where $BSADF_{r_1}^{r_2}$ is Backward Supremum Augmented Dickey Fuller 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

$$\hat{r}_e = \inf_{r_2 \geq r_0} \{ r_2 : BSADF_0^{r_2} > scv_{r_2}^{\beta_T} \} \quad \text{and} \quad \hat{r}_p = \inf_{r_2 \geq r_e + \gamma \log(T)/T} \{ r_2 : BSADF_0^{r_2} < scv_{r_2}^{\beta_T} \}$$

where $scv_{r_2}^{\beta_T}$ is the right-sided critical value with a significance level of $\beta_T$ 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 $\rho$ 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

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

For $\varepsilon_{x,t}$ and $\varepsilon_{y,t}$ we impose also the assumption of weak dependence like for $\varepsilon_t$ 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. In essence, this implies a time-varying behavior of the autoregressive coefficients $\rho_x$ and $\rho_y$; therefore, we use the notation $\rho_x(t)$ and $\rho_y(t)$ in the following. Formally, we have

$$\rho_x(t) = \begin{cases} 
1, & t < t_{ex} = \lceil Tr_{ex} \rceil, \quad \text{"non-explosive"} \\
1 + \frac{c_{ex}}{T}, & t \geq t_{ex} = \lceil Tr_{ex} \rceil, \quad \text{"explosive"} \\
1 + \frac{c_x}{T} \cdot \tilde{t}_m, & t \geq t_{px} = \lceil Tr_{px} \rceil, \quad \text{"collapse"}
\end{cases} \hfill (14)$$

with $\tilde{t}_m = \frac{t_t - t_{px}}{m}$, $m = t_{py} - t_{px}$ 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.

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 $y_t$:

$$\rho_y(t) = \begin{cases} 
1, & t < t_{ey} = \lceil Tr_{ey} \rceil, \quad \text{"non-explosive"} \\
1 + \frac{c_y}{T} + d \cdot \frac{c_x}{T} \cdot \tilde{t}_m^2, & t \geq t_{ey} = \lceil Tr_{ey} \rceil, \quad \text{"explosive"}
\end{cases} \hfill (15)$$

After a "regular" unit root regime lasting until $t_{ey} - 1$, the series becomes explosive. In the explosive regime, the first component $\frac{c_y}{T}$ (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 $d \cdot \frac{c_x}{T} \cdot \tilde{t}_m^2$ with negative localizing coefficient $c_x < 0$ (due to collapsing) and $d < 0$, we say that the explosiveness migrates from $x_t$ to $y_t$. 

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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 $y$.

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.

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. Based on these quantities, a $t$-type statistic can be calculated from the following (infeasible) auxiliary test regression

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

(16)

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.

In detail, $t_{px} = \arg \max_t \rho_x(t)$ and $t_{py} = \arg \max_{t \in (t_{px} + 1, \ldots, t_{px} + \lambda T)} \rho_y(t)$ with $\lambda$ being set equal to 1.2 defining a neighbourhood of observations similar to the bubble duration condition in equations (10) and (12).
for \( t_{px} + 1, ..., t_{py} \). The test statistic for \( H_0 : \beta_1 = 0 \) is given by

\[
Z_{\beta_1} = \frac{\hat{\beta}_1}{L(m)}. \tag{17}
\]

Here, \( L(m) \) is a slowly varying function depending on \( m \). 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 \( Z_{\beta_1} \) 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

\footnote{The slowly varying function needs to satisfy the condition \( \frac{1}{L(m)} + \frac{L(m)}{T} \to 0 \) as \( T \to \infty \) for any \( e > 0 \).}
under both hypotheses as in equation (14), but $\rho_y(t)$ now changes to

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

with $m_1 = t_{py,1} - t_{px}, m_2 = t_{py,2} - t_{fy}$ and $m = m_1 + m_2$. Here, the process starts with a non-explosive unit root. Then, it becomes explosive with its "intrinsic" part $\frac{c_y^1}{T}$ and a first migration of explosiveness $\frac{c_y^2}{T} \cdot \tilde{t}_{m_1}^2$ 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 $\frac{c_y^2}{T}$ and a second migration process $\frac{c_x}{T^2} \cdot \tilde{t}_{m_2}^2$ as long as $d_2 < 0$. Again, for $d_2 = 0$, there is no migration during the second explosive regime.

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.

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.
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.

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

$$[\rho_y(t) - 1] = \beta_0 + \beta_1 \cdot I_1 \cdot [\rho_x(t) - 1] \cdot \tilde{t}_{m_1} + \beta_2 \cdot I_2 \cdot [\rho_x(t) - 1] \cdot \tilde{t}_{m_2} + \epsilon(t).$$ \hspace{1cm} (19)

Here, $I_1$ and $I_2$ are dummy variables defined as follows. For $t = t_{px} + 1, ..., t_{py}$, the first dummy variable $I_1$ equals 1, otherwise it takes the value 0. Similarly, from $[Tr_{fy}]$ to $[Tr_{py,2}]$, the second dummy variable $I_2$ equals 1 and 0, otherwise. Thus, these dummies indicate periods of explosiveness in $y_t$. In practice, we test the pairs of hypotheses

$$H^1_0: \beta_1 = 0 \text{ vs. } H^1_1: \beta_1 < 0,$$

$$H^2_0: \beta_2 = 0 \text{ vs. } H^2_1: \beta_2 < 0$$ \hspace{1cm} (20)

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

$$Z_{\beta_1} = \frac{\hat{\beta}_1}{L(m_1)},$$

$$Z_{\beta_2} = \frac{\hat{\beta}_2}{L(m_2)}$$ \hspace{1cm} (21)

where $L(m_1)$ and $L(m_2)$ are slowly varying functions similar to $L(m)$ defined 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 $|Z_{\beta_2}| > cv_{T,1-\alpha}$.

Here, a neglected second explosive regime does not cause size or power distortions of the $t$-type test statistics from the test regression as long as the second regime is located after the first regime in the considered sample. The intuition behind this is that we consider observations from $[Tr_{px}]$ to $[Tr_{py}]$ in the single migration case respectively from $[Tr_{px}]$ to $[Tr_{py,1}]$ in the double migration case. The later explosive regime up to observation $[Tr_{py,2}]$ is not part of this subsample and thus, it could not affect the limiting distribution of $Z_{\beta_1}$. In general, size and power of $Z_{\beta_1}$ are not affected if we apply the single migration or the double migration test. As long as the second regime follows the first and identified regime with the
same arguments as before, \( Z_{\beta_2} = \beta_2/L(m_2) \) is a conservative test statistic for \( L(m_2) \to \infty \).

In this context, it is important to note that the bubble migration test by Phillips and Yu (2011) in its classic form is motivated by the idea that a bursting bubble in one asset class \((x)\) is connected to a bubble in another asset class \((y)\) by migration mechanism (see for the motivation of this test Caballero, Farhi, and Gourinchas, 2008b,a). In these articles, multiple bubble migration processes from \(x\) to \(y\) are not considered.

However, in the context of this study, we do not consider migration of bubbles, but migration of explosiveness, from one time series \((x;\) in our case house prices\) to another time series \((y;\) in our case government bond yield spreads\). Here, \(y\) is composed of sovereign credit risk, redenomination and/or liquidity risk. Hence, we motivate the potential migration process by the concept of financial contagion – in particular by a form of wake-up-call contagion – and flight-to-quality (or liquidity) effects (see, for a multiple asset rational expectations model to explain contagion, Kodres and Pritsker, 2002).

Flight-to-quality (or liquidity) refers to a situation where investors move their capital to assets which are expected to retain their value even in times of financial crises or turbulent market conditions. The concept of wake-up-call contagion characterizes a situation in which new information from one country lead to a reassessment of the default risk of other countries (see, for example, Giordano, Pericoli, and Tommasino 2013 Bekaert, Ehrmann, Fratzscher, and Mehl, 2014). Phrased somewhat differently, the changed environment leads to a situation where investors just start to worry more strongly about risk. The slight difference to our case is as follows rather than new information about the fundamentals of one country, a particular market – the US housing market – potentially triggered the reassessment of the default risk of European countries relative to Germany.

Both, flight-to-quality (or liquidity) and wake-up-call contagion effects lead to divergence of the interest rates and thus, possibly to explosive behavior of the spreads (see Section 4.2). First, the interest rates decreased for German government bonds – this refers to flight-to-quality (or liquidity) effects. Second, the interest rates increased for government bonds from Greece, Ireland, Italy, Portugal and Spain – this refers to a form of wake-up-call contagion.

Safe haven and wake-up-call contagion effects are likely to be present during the financial crisis and the sovereign debt crisis and might have been triggered in both cases by the bursting bubble in the US housing market. For this reason, we test against migration effects from the REITs to government bond yield spreads during both crises. Thus, we are interested in the question whether a collapse of the bubble in \(x\) supported two further explosive regimes in \(y\).

Please note that in the case of evidence in favor of migration behavior, we would not summarize that the crises have been caused by the collapse of the housing market in the US. However, in this situation, the divergence of interest rates from peripheral countries and
Germany would have been triggered in the sense that the spreads diverged even faster due to the burst of the US house price bubble.

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. 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.

![Graph showing 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 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

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10 As the application of the SADF statistic does not provide any additional insights, we decide to save space and refrain from reporting those.
<table>
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<th>Ireland</th>
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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.
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 2.

As we have found 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.

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 October 10, 2003 and April 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. The resulting date is March 9, 2007. Earlier bubble regimes (October 10, 2003 to April 16, 2004 and November 19, 2004 to January 7, 2005) might be related to the same price rally which finally collapsed in 2007. Later explosive regimes (from October 24, 2008 to December 12, 2008 and from February 20, 2009 to April 10, 2009) might be related to speculative behavior during the financial crisis and thus, these bubble periods are part of another story.

Therefore, we focus on the arising and collapsing house price bubbles between October
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 2.

10, 2003 and March 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", Moore, December 28, 2007)
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 2.

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 indicates essentially three explosive regimes as suggested by Figure 8. The first regime lasts from September 26, 2008 to July 24, 2009, while the second regime starts January 29, 2010 and ends June 29, 2012. Thus, the first explosive regime might be related to the bankruptcy of Lehman Brothers. The second phase of explosiveness is coincident with 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 February 6, 2015 to May 15, 2015 which might be a consequence of the uncertainty of the Greek legislative election, Jan 2015.

In Figure 9, the results for the spread between Ireland and Germany are presented. Ireland is a special case (though there are some similarities to Spain). The country had to cope with its 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 April 18, 2003 and ending August 8, 2003, and the second begins April 6, 2007 and lasts until December 7, 2007 and the third starts December 12, 2008 and ends July 24, 2009 and the fourth start is stamped on September 24, 2010 and the end is indicated by September 2, 2011.

Figure 10 contains the results for the Italian-German government bond spread. 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 September 26, 2008 and April 24, 2009. Furthermore, we find two explosive regimes between July 15, 2011 and September 21, 2012 which might be related to the sovereign debt crisis. A third phase of explosiveness is indicated between March 8, 2002 and April 5, 2002.

For Portugal, the BSADF approach stamps explosiveness for six regimes, see Figure 11. We aggregate them to four phases of explosiveness: The first from December 3, 2004 to February 18, 2005, the second from March 14, 2008 to May 8, 2009, the third from May 14,
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 2.

2010 to August 31, 2012 and the fourth from April 11, 2014 to May 23, 2014. The second might be related to the financial crisis while the third coincides with the sovereign debt crisis.

The Spanish case is presented in Figure 12. Here, we essentially obtain two phases of explosive regimes: From August 3, 2007 to May 8, 2009 – which might be related to the financial crisis – and from May 14, 2010 to October 5, 2012 – which is coincident with the sovereign debt crisis. 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 Portugal. 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).

![Graph showing persistence analysis](image)

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

<table>
<thead>
<tr>
<th>Series</th>
<th>$\hat{\beta}_1$</th>
<th>$Z_{\hat{\beta}_1}$</th>
<th>$m$</th>
<th>$t_{px}$</th>
<th>$t_{py}$</th>
</tr>
</thead>
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<tr>
<td>Ireland</td>
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<td>-0.67</td>
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<td>Feb 20, 2009</td>
</tr>
<tr>
<td>Italy</td>
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<td>-3.36</td>
<td>96</td>
<td>Feb 16, 2007</td>
<td>Dec 12, 2008</td>
</tr>
<tr>
<td>Spain</td>
<td>-4.36</td>
<td>-2.15</td>
<td>106</td>
<td>Feb 16, 2007</td>
<td>Feb 20, 2009</td>
</tr>
</tbody>
</table>

#### BSADF stamping

<table>
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<th>Series</th>
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<th>$Z_{\hat{\beta}_1}$</th>
<th>$m$</th>
<th>$t_{px}$</th>
<th>$t_{py}$</th>
</tr>
</thead>
<tbody>
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<td>Jul 24, 2009</td>
</tr>
<tr>
<td>Ireland</td>
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<td>125</td>
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<td>Jul 24, 2009</td>
</tr>
<tr>
<td>Italy</td>
<td>-8.54</td>
<td>-4.17</td>
<td>112</td>
<td>Mar 9, 2007</td>
<td>Apr 24, 2009</td>
</tr>
<tr>
<td>Portugal</td>
<td>-5.44</td>
<td>-2.65</td>
<td>114</td>
<td>Mar 9, 2007</td>
<td>May 8, 2009</td>
</tr>
<tr>
<td>Spain</td>
<td>-4.46</td>
<td>-2.17</td>
<td>114</td>
<td>Mar 9, 2007</td>
<td>May 8, 2009</td>
</tr>
</tbody>
</table>

Table 3: This table reports the results of the single migration test statistic $Z_{\hat{\beta}_1} = \hat{\beta}_1 / \log_{10}(m)$ for $H_0 : \beta_1 = 0$ in the auxiliary regression $[\rho_y(t) - 1] = \beta_0 + \beta_1 [\rho_x(t) - 1] \cdot \tilde{t}_m + \epsilon(t)$, 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.
Table 4: This table reports the results of the double migration test statistics $Z_{β_1} = \tilde{β}_1 / \log_{10}(m_1)$ for $H_0 : β_1 = 0$ and $Z_{β_2} = \tilde{β}_2 / \log_{10}(m_2)$ for $H_0 : β_2 = 0$ in the auxiliary regression $[ρ_y(t) - 1] = β_0 + β_1 \cdot I_1 \cdot [ρ_x(t) - 1] \cdot \tilde{t}_{m_1} + β_2 \cdot I_2 \cdot [ρ_x(t) - 1] \cdot \tilde{t}_{m_2} + ε(t)$. 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, marking the second endpoint of the migration analysis. We use the two tests presented in Subsection 5.3 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 at 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. 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.

In Section 6.1 we interpreted later phases of explosiveness (from October 24, 2008 to December 12, 2008 and from February 20, 2009 to April 10, 2009) as ”part of another story”. However, these explosive regimes are not removed from the sample when it comes to the bubble migration analysis to ensure that further transmission effects are not neglected. Furthermore, we also tested against migration effects starting March 13, 2009 as this is the peak in later explosive phases. We could not reject the null hypothesis of no migration effects at nominal significance level of five percent. Hence, the explosive regimes from October 24, 2008 to December 12, 2008 and from February 20, 2009 to April 10, 2009 are not the main driver of the results for the sovereign debt crisis.

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 not rejected for \( a = 1 \) in all cases and for both date stamping procedures (Peak-to-peak and BSADF). Thus, the migration effects might have

\[\text{Results are not reported in order to conserve space and are available on request from the authors.}\]
stopped after the first regime of arising sovereign credit risk in the EMU.

In the discussion of test results mentioned earlier, 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 eventually 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 $\hat{\beta}_1 = -4.90$ and $\hat{\beta}_2 = -1.77$ (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 $\hat{\beta}_1 = -3.20$ 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 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.

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 sovereign debt crisis in Europe gained momentum.
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 a 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 the same 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.

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 the 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 indicated 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.

References


