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A yield spread perspective on the great financial crisis: Break-point test evidence☆

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ABSTRACT

We use a simple partial adjustment econometric framework to investigate the effects of financial crises on the dynamic properties of yield spreads. We find that crises manifest themselves in the form of substantial disruptions revealed by changes in the persistence of the shocks to spreads as much as by in their unconditional mean levels. Formal breakpoint tests confirm that in the U.S. the Great Financial Crisis has been over approximately since the Spring of 2009 and provide a conservative dating centered around the August 2007–June 2009 dates. However, some yield spread series point to an end of the most serious disruptions as early as in December 2008. Some symptoms of an impending crisis re-appear instead in the second half of 2011. We also uncover evidence that the LSAP program implemented by the Fed in the U.S. residential mortgage market has been effective, in the sense that the risk premia in this market have been uniquely shielded from the disruptive effects of the crisis.

1. Introduction

The financial crisis that has allegedly taken place between 2007 and 2009 in the United States has been viewed as the worst financial disruption since the Great Depression of 1929–1933. Many commentators have in fact taken the habit of referring to it as the Great Financial Crisis (henceforth, GFC). The banking crises of the Great Depression involved runs on banks by depositors, whereas the GFC reflected widespread panic in wholesale funding markets that left banks unable to roll over short-term debt. That has deteriorated to engulf most fixed income (FI) markets, both in the US and internationally where persistently high, often historically abnormal yields and yield spreads between different instruments have been observed. The reaction to the crisis by central banks and governments around the world has been massive. It has involved large-scale interventions in both short- and long-term, in private as well as public segments of international bond markets. Although by the end of 2009, a majority of analysts became willing to admit that the worst of the GFC was over, throughout 2010 and 2011 lingering doubts have persisted as to whether the GFC could be dated as a closed, and yet painful event of the recent financial history.

Because a number of such interventions have directly involved the segments of the fixed income (FI) markets more severely affected by the GFC, in this paper we take a perspective that is based on *yield spread* data. A yield spread is the difference between the yield to maturity of a riskier bond and the yield of a comparatively less risky (or riskless) bond.[1] The dimensions of risk that are measured by yield spreads may be many, but they can be grouped as originating from either their default "intensity" (i.e., probability of default and loss given default) or their liquidity risk. In this paper, we ask four related questions:

- How can we date a financial crisis, at least on the basis of the yield spread perspective adopted in this paper? This relates to the general question of what properties of yield spreads are affected by a crisis.
- In particular, can we date the GFC? Most researchers have been referring to the crisis as a 2007–2009 phenomenon: is this dating as correct as commonly held and/or can we be more precise about its dating as it is usually required of business cycles?

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[1] Batten, Hogan, and Jacoby (2005) have emphasized that FI spreads may in principle be defined also as a ratio of (more) risky over less risky (riskless) gross yields but also notice that some type of analysis may be flawed by biases in this last case.

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- Were the interventions by the Federal Reserve (more generally, by US policy-makers including the Treasury department) effective in fighting the disruptive effects of the crisis? In particular, were the Large Scale Asset Purchases (LSAP) programs announced in late 2008 and implemented between early 2009 and mid-2010 effective and when?
- Do any of these questions admit market-specific answers? For instance, are there FI markets that were never affected by the GFC, or for which the crisis seemed to be over well in advance of mid- to late-2009? Similarly, did the European sovereign debt crisis of 2010–2011 revive fears of either a new financial crisis spreading to US markets or of the GFC itself going through a later, deeper phase?

In fact, considerable ambiguity and an intense debate has recently concerned the exact dating of the GFC. The conclusions have often reflect the priors of the different researchers as well as their specific methodological approach. Table 1 offers a synopsis of a few among the papers that have appeared in the literature between 2008 and 2011. Although a simple synopsis cannot claim to be exhaustive, we have systematically searched all papers that have investigated the GFC focusing on the behavior of FI yield spreads. Additionally, many criteria could have been used to sort papers in the table and yet—because most of these manuscripts have been repeatedly revised and updated—we have opted for a simple alphabetical sorting. As a Reader may notice most papers had a distinct policy focus as their objective consisted in drawing a connection between (so-called

Table 1
Synoptic table of recent papers that have relied on (approximate) dating of the great financial crisis.

Paper/author(s)	Implications/statements on dating the GFC	Notes on Dating Policy	Conclusions
Adrian et al. (2010)	Starting date: August 15, 2007.	Event/History-based	CPFF generally effective.
Aït-Sahalia et al. (2009)	End date: second part of 2009. End date: April 2, 2009 (G20 Leaders' Summit on Financial Markets and the World Economy).	Event/History-based (starting date from regime switching mode	International paper, but measures are generally found to have been effective also for the U.S.
Baba (2009)	Starting date: August 9, 2007. No end date stated. Starting date: early August 2007.	Event/History-based	The Fed-activated swap lines met the demand for U.S. dollar funding created by the pull-back in funding from money market mutual funds.
Brave and Genay (2011)	End date: Summer 2009 ("In Summer 2009, the Federal Reserve began to reduce the amount of funds available through the individual programs")	Event/History-based	The announcements of Fed policies during the crisis were associated with significant improvements in broad financial market conditions.
Campbell et al. (2011)	Starting date: Summer of 2007. End date: mid-2009.	Event/History-based	Announcements concerning TALF affected the markets of highly rated ABS and CMBS. TALF may have improved ABS market liquidity, but have not provided substantial subsidies or certification benefits to individual securities. TSLF was extremely effective in raising Treasury repo rate back to levels close to the Fed funds rate.
Cecchetti (2009)	Starting date: August 9, 2007 (the "definitive trigger"). No end date stated.	Event/History-based	TAF helped at first to reduce the LIBOR-Fed funds rate spread. The PDCF played a role in reducing the spreads btw. yields on government agencies and U.S. Treasury securities. TSLF was extremely effective in raising Treasury repo rate back to levels close to the federal funds rate.
Christensen et al. (2009)	First stage of the crisis: August 9, 2007–December 12, 2007 Second stage: December 12, 2007 (Fed commits to unconventional liquidity measures)–? (no end date stated)	Event/History-based	TAF was effective in reducing the LIBOR-Fed funds rate spread.
Dwyer and Tkac (2009)	Starting date: August 9, 2007. End date: March 2009 ("As of this writing in July 2009, it still is a conjecture, even if a more plausible one.")	Event/History-based	The AMLF and the CPFF appear to have been successful in averting a run on money market funds and providing a liquid secondary market to keep the commercial paper funding markets accessible.
Frank and Hesse (2009)	Starting date: July 2007. No end date stated.	Event/History-based	TAF helpful in compressing Libor spreads, but the economic magnitude not very large.
Furceri and Mourougane (2009)	First stage: July 2007–September 15, 2008, a period of financial turmoil and limited spreading. Second stage: 15 September 2008–? (no end date stated).	Event/History-based	International paper, but measures are generally found to have been effective also for the U.S..
Hancock and Passmore (2011)	Starting date: August 2007. End date: May 27, 2009 ("After May 27, 2009, MBS yields largely returned to reflect fundamentals").	Event/History-based	The Fed MBS purchase program removed substantial risk premiums embedded in mortgage rates because of the financial crisis. The Federal Reserve also re-established a robust
Sarkar, 2009	First stage of the crisis: August 2007–September 2008 Second stage: September 2008–? (no end date stated)	Event/History-based	First stage dominated by capital and liquidity shortages; second stage by credit risk. Fed facilities were effective: in particular, the Fed introduced TAF and the bilateral currency swap lines in December 2007 and the TSLF and the PDCF in March 2008, as the type of risks evolved.
Taylor and Williams (2009)	Starting date: August 9, 2007 (called a break-point). No end date stated.	Event/History-based	TAF had no effect on the Libor–Ois spread and did not affect total liquidity, expectations of future overnight rates, or counterparty risk.
Wu (2011)	Starting date: August 9, 2007. No end date stated but market strains "as of 2011, they are mostly gone."	Event/History-based	TAF had strong effects in reducing financial strains in the inter-bank money market. PDCF had less discernible effects than TAF in relieving financial strains in the Libor market.

unconventional) monetary policy liquidity measures and their effects on spreads (the fourth column of this table will be used again in subsequent sections). However, the second column concerning the dating of the GFC immediately reveals that while most papers agreed on early August 2007 as a potential starting date of the GFC, few (or none) of them had traced this claim back to the actual features of spread data.[2] In fact, in a few cases (e.g., Furceri & Mourougane, 2009; Sarkar, 2009) several different stages within the GFC could be isolated and discussed. Only a handful of papers have ventured into establishing an end date for the GFC (see e.g., Aït-Sahalia, Andritzky, Jobst, Nowak, & Tamirisa, 2009; Campbell, Covitz, Nelson, & Pence, 2011), although this was generally possible only for papers written or revised after the beginning of 2009: usually this consists in generic claims to the effect that the crisis would have been re-absorbed around mid-2009. Eventually, most of the papers established this dating in a casual fashion, often with explicit reference to only one spread series of immediate interest, and without resorting to the typical techniques that applied time series econometricians would use to locate a break-point in time.

While the question of precisely dating the GFC remains relevant, it is also important for us to explain the logic of an empirical approach to these issues based on yield spreads, i.e., on bond market-driven estimates of measures of risk premia. There are a number of reasons that can be invoked. A few of them are generally applicable to all research that has focused on FI yield spreads, and others specific to the GFC. First, to filter a financial crisis through the lenses of spread data is implicitly a way to relate financial events to business cycle developments. A feature of U.S. post-WWII business cycle experience that has been widely documented (see e.g., Gilchist, Yankov, & Zakrajsek, 2009; Guha & Hiris, 2002) is the tendency of a number of FI spreads to widen shortly before the onset of recessions and to narrow again before recoveries. One interpretation of these results is that these credit risk spreads measure the default risk on private (relatively risky) debt (see Stock & Watson, 2003). For instance, Philippon (2009) has proposed a model in which the predictive content of corporate bond spreads for economic activity reflects a decline in economic fundamentals stemming from a reduction in the expected present value of corporate cash flows prior to a downturn. Rising credit spreads may also reflect disruptions in the supply of credit resulting from the worsening in the quality of corporate balance sheets or from the deterioration in the health of financial intermediaries—the financial accelerator mechanism emphasized by Bernanke, Gertler, and Gilchrist (1999). Therefore, the information contained in yield spreads is important because it may be indicative of an important channel through which financial prices affect the real side of the economy. More generally, a better understanding of the dynamics of credit and liquidity risk premia incorporated in the prices of FI products (like term deposits and bonds)—specifically, the asymmetric adjustment process that characterizes turbulent crisis periods from more normal states—has a number of practical implications for investors. When market participants perceive an increase in default risk, they will re-allocate to safer assets and the default risk premium will widen.

Our analysis of the changing dynamic properties of a number of commonly reported yield spread series confirms the (possibly obvious) claim that the GFC has been over since the late Spring of 2009 or the Summer of the same, at the latest. Although there is considerable uncertainty as to when exactly the crisis ceased producing its disruptive effects, there is no doubt that after mid-2009 most FI markets have reverted to a normal, pre-crisis state. The financial crisis can be conservatively dated as an August 2007–June 2009 phenomenon, although some yield spread series seem to point to an end of the most serious disruptions as early as December 2008. The LSAP programs implemented by the Fed in the US (agency-supported) residential mortgage market seems to have been considerably effective in the sense that risk premia in this market have been uniquely shielded from the adverse effects of the crisis. Interestingly, this has not occurred in the commercial mortgage market, at least insofar as the private label market for which we have collected data. This in spite of the fact that some of the interventions under the LSAP programs have also specifically targeted the commercial mortgage segment. This may imply that while selective portions of LSAP have produced the desired effects, it may not have been the case across the board. Further bivariate tests reveal that the financial crisis may be characterized as a period in which the yields defining most of the spreads investigated stopped reacting to departures from their (common) "attractor" level in the way they usually did under normal circumstances, always increasing even when the past spread exceeds the long-run attractor yield. On the contrary, in the non-crisis periods and especially in the aftermath of the Great Crisis, we observe that for most spreads, yields tend to adjust in directions—upwards for yields on high (low) default (liquidity) risk bonds, and downward for yields on high (low) default (liquidity) risk bonds—that are compatible with mean-reversion and stationarity of the spreads. Finally, a number of our tests reveal that tensions have recently re-surfaced in a few short-term FI markets between the late Summer of 2011 and the Fall of the same year. For instance, this occurs in the LIBOR and financial commercial paper markets. In the paper we speculate on the causes of these renewed state of turmoil and link it to well-known fears of a possible propagation to the international banking system (therefore also involving US intermediaries) sparked by the re-financing difficulties encountered by a number of European sovereign issuers during the second half of 2011.

Two literatures are related to our goals in this paper. One recent literature—partially summarized in Table 1—has debated whether the liquidity facilities and LSAP program implemented by the Federal Reserve have been as effective as the policy-makers had hoped for. On the one hand, several papers have argued that the short-term liquidity programs implemented by the Fed between 2007 and 2008 have been successful. For instance, Adrian, Kimbrough, and Marchioni (2010) have concluded that the Commercial Paper Funding Facility has been successful and that its declining volumes during 2009 were simply caused by its self-liquidating nature. Christensen, Lopez, and Rudebusch (2009) have assessed the effects of the establishment of the liquidity facilities—in particular, of the Term Auction Facility—on the interbank lending market and, in particular, on term LIBOR spreads over Treasury yields. Their multifactor arbitrage-free model of the term structure of interest rates and bank credit risk reveals that the central bank liquidity facilities established in December 2007 helped lower LIBOR rates. Gagnon, Raskin, Remache, and Sack (2011) have used an event study to argue that the LSAP did reduce U.S. long-term yields. On the other hand, several papers have reached opposite conclusions with reference to the credit facilities and the LSAP. For instance, Taylor and Williams (2009) have reported that the TAF was ineffective in significantly influencing the spread between LIBOR rates and overnight lending rates. Thornton (2009) has stressed that until mid-September 2008, the Fed offset the effect of its TAF lending through the liquidity programs on the total supply of credit through open market operations thus reducing their ability to affect financial markets. However, as already discussed, most of these papers have, also in the light of the fast pace of the events between 2008 and 2009, often escaped the task of documenting the exact starting and—if appropriate—end dates of the GFC, even when their focus is predominantly directed at the behavior of interest rate spreads. Using modern break point econometric tools, our paper takes up such a challenge.

A second literature has proposed increasingly sophisticated models of the dynamics in yield spreads. For instance, Davies (2008) has analyzed the determinants of US credit spreads over an extensive 85 year sample that covers several business cycles. He finds that econometric models are capable of explaining up to one fifth of the movement in

[2] A few exceptions exist: e.g., Aït-Sahalia et al. (2009), Frank and Hesse (2009), Furceri and Mourougane (2009) date the beginning of the subprime crisis back to June-July 2007.

the various spreads considered. This explanatory power derives from autoregressive-type models augmented by relatively small groups of lagged explanatory variables such as changes in riskless interest rates and returns on firms' equities or assets, as in Longstaff and Schwartz (1995).[3] Papageorgiou and Skinner (2006) have studied corporate credit spreads and the Treasury term structure focussing on the evidence of breakpoints in such relationships. Their results suggest that these relations are not constant but change slowly through time. Compared to this literature, our approach is specifically geared towards our opening questions and therefore based on the simplest available set of econometric tools adequate to develop break tests, i.e., *univariate* partial adjustment time series models. These models are useful to simultaneously estimate the persistence of the dynamic spread process (i.e., the implied half-life of a shock) and the long-run spread, thus disregarding the relationship between risk spread curves and the risk-free term structure of interest rates. Moreover, our partial adjustment model can be interpreted as a special, restricted AR(2) process and hence it belongs to the simple class of ARMA models. This has the advantage of allowing us to implement a few well-known breakpoint test methodologies such as Chow's (1960) and Andrews (1993).

The paper has the following structure. Section 2.1 reviews the unfolding of the GFC and proposes a short list of key episodes used in our commentary. Section 2.2 examines how the yield spreads in seven different bond markets have reacted to these key events. Section 3 presents our econometric methodology. Section 4 contains our main empirical findings. It shows that yield spreads can be described as covariance stationary series, that the parameter estimates of a simple partial adjustment model are subject to considerable instability over time, and formally tests for and finds breakpoints that we interpret as data-driven markers of the onset and the end of the GFC. Section 5 concludes.

2. The financial crisis through the yield spread lenses

In this Section we review the main events of the GFC and familiarize with the yield spread data. Our objective is not to exhaustively list all the significant developments or discuss causes and solutions to the crisis. A number of reviews are now available, see e.g., Dwyer and Tkack (2009) and references therein.

2.1. The crisis and the Fed's reaction

The financial crisis began with a downturn in U.S. residential real estate markets as a growing number of banks and hedge funds reported substantial losses on subprime mortgages and mortgage-backed securities (MBS). The crisis had been slowly building up since the early months of 2007. For instance, in late February 2007 the Federal Home Loan Mortgage Corporation (Freddie Mac) had announced that it would no longer buy the most risky subprime mortgages, which meant that a large portion of the process of origination and securitization of subprime MBS would have to be moved to the private sector. In June 2007 Standard and Poor's and Moody's Investor Services had downgraded over 100 bonds backed by second-lien subprime mortgages. However, a major step towards a spiralling crisis was marked by Fitch Ratings' decision in August 2007 to downgrade one of the major firms specialized in mortgage intermediation in the subprime segment, Countrywide Financial Co. As a result, Countrywide was forced to borrow the entire $11.5 billion available in its credit lines with other banks, which was painful evidence that the crisis was destined to spread from the mortgage market to the financial intermediaries backing its operators. Soon the crisis appeared to be able to spread beyond the boundaries of the US mortgage market when it spilled over to the interbank lending market. The London Interbank Offered Rate (LIBOR) and other funding rates spiked after the French bank BNP Paribas announced that it was halting redemptions for three of its investment funds. These two negative developments are labelled as event [1] in our list.

Initially, the Fed's reaction was limited to stressing the availability of the discount window, e.g., by extending the maximum term of primary loans to 30 days, and by lowering the Fed fund rate target by 50 bp. In December 2007, the Fed announced the establishment of reciprocal swap currency agreements with the European Central Bank and the Swiss National Bank to provide a source of dollar funding to European financial markets. Again in December, the Fed announced the creation of the Term Auction Facility (TAF) to lend funds directly to banks for a fixed term. The Fed established the TAF because the volume of discount window borrowing had remained low despite persistent stress in interbank funding markets. This allegedly derived from a perceived stigma associated with borrowing at the window (see Thornton, 2009). These two initial reactions by the Fed in coordination with central banks worldwide are labelled as event [2] in our list. Financial markets remained strained in early 2008. In March, the Federal Reserve established the Term Securities Lending Facility (TSLF) to provide secured loans of Treasury securities to primary dealers for 28-day terms. This is event [3] in our list. Later in March, the Fed established the Primary Dealer Credit Facility (PDCF) to provide secured overnight loans to primary dealers under Section 13(3) of the Federal Reserve Act, which permits the Federal Reserve to lend to any individual, partnership, or corporation "in unusual and exigent circumstances". The PDCF essentially opened the discount window to primary government security dealers.[4] This is event [4] in our list.

The financial crisis intensified during the final four months of 2008. Lehman Brothers, a major investment bank, filed for bankruptcy on September 15. The Lehman bankruptcy immediately produced a victim. On September 16, the Reserve Primary Money Fund announced that the net asset value of its shares had fallen below $1 because of losses incurred on the fund's holdings of Lehman commercial paper and notes. The announcement triggered widespread withdrawals from other money funds, which prompted the U.S. Treasury Department to announce a temporary program to guarantee investments in participating money market mutual funds, the Asset-Backed Commercial Paper Money Market Mutual Fund Liquidity Facility (AMLF), set up to extend non-recourse loans to U.S. depository institutions to finance purchases of asset-backed commercial paper from money market mutual funds. This is event [5]. Financial markets re-plunged in a state of turmoil over the following weeks. To help alleviate financial strains in the commercial paper market, the Fed established the Commercial Paper Funding Facility (CPFF) on October 7, 2008. This facility provided financing for a special-purpose vehicle established to purchase 3-month unsecured and asset-backed commercial paper. On October 21, the Fed created the Money Market Investor Funding Facility (MMIFF). Under the MMIFF, the Fed offered to provide loans to a series of special-purpose vehicles that purchased assets from money market mutual funds. These events are labelled as [6] in our list.

In spite of the beneficial effects produced on the short-end of the FI markets, the situation remained difficult in most other segments.

[3] Batten et al. (2005), Christiansen (2002) and Manzoni (2002) have extended this early literature to incorporate GARCH specifications to accommodate persistence in the conditional variance of yield spread changes.

[4] Again in March, the Federal Reserve Board invoked Section 13(3) when it authorized the Federal Reserve Bank of New York to lend $29 billion to a newly created limited liability corporation (Maiden Lane, LLC) to facilitate the acquisition of the distressed investment bank Bear Stearns by JPMorgan Chase.

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On November 25, the Federal Reserve again invoked Section 13(3) when it announced the creation of the Term Asset-Backed Securities Lending Facility (TALF). Under this facility, the Federal Reserve Bank of New York provided loans on a non-recourse basis to holders of Aaa-rated asset backed securities and recently originated consumer and small business loans. At the same time, the FOMC announced its intention to purchase large amounts of U.S. Treasury and mortgage-backed securities issued by Fannie Mae, Freddie Mac, and Ginnie Mae.[5] This is event [7]. In addition to the Fed's programs to stabilize specific financial markets, the FOMC reduced its target for the federal funds rate in a series of moves that lowered the target rate from 5.25% in August 2007 to a range of 0 to 0.25% in December 2008, event [8] in our list.

Between late 2008 and early 2009 the financial crisis remained at the forefront of policy concerns, as witnessed by the fact that the Federal Reserve Board approved the applications of several large financial firms to become bank holding companies. In February 2009 the Fed announced the extension of all the existing liquidity programs, listed as events [9]–[7]. In the meantime, fears spread that the enormous market for securitized commercial mortgages would be on the brink of collapse. The explicit admission that financial markets remained strained and the consequent extension of the extraordinary measures enacted between December 2007 and December 2008 represents in itself a further significant event, [9], in our list. In fact, in March 2009 the U.S. Treasury and Fed announced the effective launch of the TALF with its first auctions, while in May 2009 the Fed announced that commercial mortgage-backed securities (CMBS) would become eligible collateral under the TALF. These are events [10] and [11], respectively.

The turnaround seems to have occurred after the Spring of 2009. In June 2009 (event [12] in our list) the Fed had still announced a number of modifications to its liquidity programs, even though a novel desire to fine-tune the programs had replaced the tension towards expanding them that had dominated policy-makers until April 2009. The Fed announced that the amounts auctioned at the biweekly auctions of Term Auction Facility (TAF) funds would be reduced from $150 billion to $125 billion, effective with the July 13, 2009 auction. With the situation rapidly improving, in November 2009 the Fed approved a first reduction in the maximum maturity of credit at the discount window. Although the discount window never played a major role in the credit easing policies of the Fed, we take this step as our event [13] because—to the best of our knowledge—it did represent the first official acknowledgement that the financial system was on a healing track. The Federal Reserve completed its purchase of Treasury securities in October 2009. Our final event [14] is dated February 2010, when a number of liquidity programs (CPFF, ABCPMLF, TSLF) expired.[6]

Is it possible to conjecture an end date for the crisis similar to the process that has led us to identify the Summer (say, August) of 2007 as its starting period? The notes above in relation to events [12]-[14] and Fig. 1 lead us to conjecture that the period March-June 2009, and in any event the Spring of 2009 may have marked the end of the crisis. Fig. 1 plots the time series of the total adjusted (in the St. Louis definition) monetary base and the total amount of the outstanding loans under all liquidity/credit facilities between 2008 and early 2010. Clearly, the total size of the credit extended through all liquidity facilities takes off at the end of 2008 (consistently with [7]–[9] above) and peaks after 14–15 months, in March 2009. Then the amount starts declining, and the speed of descent becomes noticeable after June 2009.

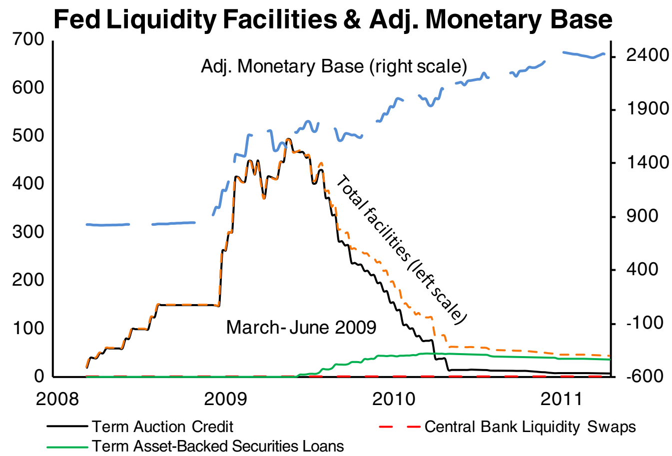

Fig. 1. Quantitative Evolution of Federal Reserve Credit Facilities and Adjusted Monetary Base. The figure plots the total amount (in billions of dollars) of the credit extended to the economy by the Fed through the TAF (Term Auction Facility), the bilateral currency swaps established with a number of central banks between 2007 and 2009, and the TALF (Term Asset-Backed Securities Loan Facility). As a benchmark and because it is directly affected by the securities (Treasuries and mortgage-backed securities) purchases implemented by the Fed in 2008–2010, the chart also plots the total adjusted monetary base in billions of dollars. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

2.2. Yield spread data

In this Section we plot and discuss 7 alternative notions of FI yield spreads—distinct in terms of the FI products they refer to, as well as the maturity of the underlying securities—to describe the unfolding of the crisis and the subsequent healing—if any—of the financial system. We also connect the 7 plots in Fig. 2 with the 14 key events that have listed in Section 2.1. For ease of exposition, such events are summarized at the top of Fig. 2. Finally, Fig. 2 represents a first chance to relate our results and empirical analysis to recent FI market events that have unfolded between late 2010 and 2011. To ease the readability of the graphs, we plot spread data for the 2006–2011 period even when longer time series (in fact, to be used below in our econometric estimates) were available. Each plot in Fig. 2 has a dual-axis structure: the left axis refers to both components of the spread under consideration, plotted in the lower portion of the diagram; the right axis refers to the yield spread itself, plotted in the top portion of the diagram. The sources for all the data series are Haver Analytics and Bloomberg. The frequency of all series is weekly as in many earlier papers, e.g., Christensen et al. (2009), or Longstaff, Mithal, and Neis (2005).

In the following, we emphasize the role played by *credit* and *liquidity spreads*. The credit (default) spread captures the additional compensation required by investors to bear the risk that the issuer of a FI product may default on its obligations. The liquidity spread measures the additional compensation required by investors to bear the risk that the underlying market may not allow a quick and cheap disinvestment, should this be needed. While we would like one simple FI measure for each type of risk premium, there are in fact a multiplicity of such measures used in the literature. For instance, credit risk premia may also be measured using option-adjusted spreads, asset swap spreads, and credit default swap spreads (see e.g., Batten & Hogan, 2002). However, the most popular measures still remain simple yield

[5] The FOMC was later to increase the amount of its purchases in 2009 and again in 2010. The literature has come to refer to this set of programs with the acronym LSAP.

[6] As of the end of 2011, all the liquidity facilities in [2-7] have been closed. At of the end of March 2010, the Federal Reserve has also concluded its LSAPs of $300 billion of Treasury securities, of $1.25 trillion of agency MBS, and of about $175 billion of agency debt. Subsequent large-scale purchase interventions (sometimes dubbed “QE2” by the financial press) were to come in 2010–2011, but these were mostly justified by macroeconomic objectives.

spreads since they attempt to capture the compensation of credit or liquidity qualities by measuring the additional return paid by the riskier security as a spread on some higher quality, lower risk benchmark with identical maturity. One problem with yield spreads is that the benchmark (high quality) security is often chosen to have a maturity close to, but not perfectly coincident with that of the riskier (lower quality) bond. This mismatch means that the measure is biased if the underlying benchmark curve has a non-zero slope. Moreover, the benchmark security may change over time, as the bond rolls down the curve. As a result, the yield spread is often not a consistent measure through time. To overcome the issue of maturity mismatch, it is possible to use a benchmark yield where the correct maturity yield has been interpolated off the appropriate reference curve: The Interpolated Spread is the difference between the yield to maturity of the bond and the linearly interpolated yield to the same maturity. All the yield spreads used in this paper are interpolated spreads.

We now turn to a brief description of the 7 yield spread series in Fig. 2. The Off-On the Run Treasury spread is the difference between the yield of a Treasury with a residual maturity of 10 years but not recently issued and the yield of highly liquid, frequently traded Treasury securities—in this case the most recently issued benchmark with a maturity of exactly 10 years. This spread is commonly interpreted as a measure of the market liquidity risk premium because—given that its definition should try to minimize maturity mis-matches by interpolation—two Treasuries with identical maturity should imply identical credit risk and differ only for the higher "convenience yield" that a highly traded security gives over another security that is traded infrequently. Fig. 2 shows that until late 2007 the off-the-run/on-the-run spread oscillated around its typical, long-run average of 14–18 bp with isolated peaks of 20 bp. However, starting from October 2007, this spread starts exhibiting a modest but noticeable upward trend that leaves it oscillating between 10 and 30 bp for most of 2008, before August. As a result of Lehman's default in September 2008, the liquidity premium spiked, repeatedly exceeding 70 bp and rarely receding below 20 bp throughout the rest of 2008 and until February 2009. During this period, the spread also appears to be exceptionally volatile. Starting in March 2009, the off-the-run/on-the-run spread exhibits a pronounced downward trend that stabilizes it back to 15–30 bp by late 2009. Even though this spread remains more volatile than in the pre-2007 period, from mid-2010 it has settled back to average levels which are consistent with the historical norm.

During the financial crisis, the LIBOR-OIS spread has been a closely watched barometer of distress in money markets.[7] The nature of the LIBOR-OIS spread is not completely clear. At face value, the spread measures a credit risk premium: while the LIBOR, referencing a cash instrument, reflects both credit and liquidity risk, the OIS is a swap rate and as such it has little exposure to default risk because swap contracts do not involve any initial cash flows. However, the typical default risk implicit in LIBOR rates is modest.[8] Fig. 2 shows a pattern for the LIBOR-OIS spread that is qualitatively similar, but considerably more extreme than the off-the-run/on-the-run spread. Until July 2007, the LIBOR-OIS spread moved in narrow corridor, between 1 and 11 bp. At the onset of the crisis, the spread jumped to 90 bp and remained between 50 and 100 bp throughout the Summer of 2008, which is a remarkable 5–10 multiple of the historical pre-crisis norm. However, it is after Lehman's bankruptcy that the LIBOR-OIS spread skyrocketed to an exceptional (but short-lived) 345 bp. In early 2009 the spread still appeared to have remained substantially altered, exceeding 100 bp. After March 2009, the LIBOR-OIS spread started gradually declining, oscillating between 10 and 15 bp in 2010, in line with the pre-crisis experience. However, towards the end of our sample in the Fall of 2011, the plot shows that this spread heads once more north, reaching highs in excess of 40 bp, comparable to those observed at the end of 2009. The connections between such an increase and fears of a funding crisis with European banks spreading over to the US (also through their subsidiaries or parent companies), is probably behind such a late bout of market tensions.

The commercial paper (CP) market is used by commercial banks, non-bank financial institutions, and non-financial institutions to obtain short-term funding. There are two types of CP: unsecured and asset-backed. Unsecured CP consists of promissory notes with a fixed maturity between 1 and 270 days, unless the paper is issued with the option of an extendable maturity. Unsecured CP is not backed by collateral, which makes the credit rating of the originating institution a key variable. A typical spread representative of CP market conditions is the differences between the yields of investment grade (e.g., Aa and higher) 3-month financial unsecured CP and 3-month T-bill yields.[9] Clearly, this spread mostly reflects a compensation for the short-run credit risk of the financial sector. Fig. 2 tells a story that fails to boil down to Lehman's demise. Outstanding CP had peaked with a total market value of $2.2 trillion in August 2007. The market had been growing for years, while spreads had been declining: In Fig. 2 we notice a spread that moves between 10 and 25 bp during the first half of 2007. This was often below the low historical pre-crisis average of 16 bp. At the onset of the crisis, the spread jumps to levels that are between 6 and 12 times larger, oscillating between 70 and 190 bp over the period August 2007–August 2008. Between these dates, the entire CP market experienced a notable decline in terms of volumes issued. As argued by Adrian et al. (2010), the CPFF did substantially reduce the spread, which quickly declined from a new peak of almost 200 bp in early January 2009 to less than 30 bp in late March 2009. Since June 2010, the 3-month financial CP yield spread has tamed to a narrow range of 8–15 bp.

Asset-backed commercial paper (ABCP) is a form of CP that is collateralized by other financial assets (e.g., mortgage-backed securities) and therefore represents secured borrowing. The rise of ABCP has been tightly intertwined with the growth of securitization. In the decade prior to the crisis, ABCP increased from $250 billion in 1997 to over $1 trillion by 2007 (i.e., from roughly 20% to as much as 50% of all outstanding CP), fueled by the considerable distribution of residential mortgage exposure through structured finance products. A typical spread is the difference between the yield of investment grade 3-month ABCP and 3-month T-bill yields, which reflects a compensation for short-run credit and roll-over risks. With reference to ABCP, Fig. 2 shows a dynamics that is similar to unsecured financial CP. Before mid-2007, the average ABCP spread fluctuated between 5 and 40 bp, which is in line with the 19 bp average of the pre-crisis period. However, the ABCP market was one of the first markets hit by the crisis, which is to be expected given its strong connections with the US residential mortgage market.[10] This was immediately

[7] The 3-month LIBOR is the interest rate at which banks borrow unsecured funds from other banks in the London wholesale money market for a period of 3 months. The Overnight Indexed Swap (OIS) rate is the fixed interest rate a bank receives in 3-month swaps between the fixed OIS rate and a (compound) interest payment on the notional amount to be determined with reference to the effective federal funds rate.

[8] A few researchers (e.g., Christensen et al., 2009) have argued that especially during the financial crisis the spikes in the LIBOR rate may have reflected liquidity as well as credit risks.

[9] A justification for focussing on financial CP is offered by Wu and Zhang (2008) who have divided their CP credit rating groups into two broad industry sectors—financial and corporate—and studied whether there are structural differences across the two sectors. They find that credit spreads on financial CP are on average wider and more volatile than the spreads on non-financial CP and that they are more responsive to shocks to economic conditions, which is consistent with our goals.

[10] The ABCP market experienced a sharp decline starting in August 2007. Increasing investor risk aversion to credit exposure, general concerns about the functioning of the ABCP market, and heightened concerns about roll-over risk in the second half of 2007 precipitated a $500 billion reduction in total ABCP outstanding.

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reflected in the ABCP spread, which repeatedly spiked to exceed 150 bp between August 2007 and August 2008, generally oscillating around a new, higher mean of 120–130 bp. Naturally, the collapse of Lehman, one of the major players in the ABCP market, sent spreads to extraordinarily high levels, in excess of 300 bp. However, as in the case of financial CP the creation of the CPFF and of the AMLF in September 2008 greatly helped in bringing the situation under control and lowered the spreads back to "physiological levels" (see Adrian et al., 2010). ABCP spreads have returned below 100 bp around the end of 2008 and after the beginning of the Spring of 2009 they have been oscillating between 10 and 20 bp, in-line with pre-crisis levels.

The market that has been identified as the catalyst of the financial crisis is the US mortgage market. Data on a variety of mortgage rates are available. We have focused our attention on yield spreads derived from two portfolios for which the construction of long time series is possible: a 5-year index of private-label Aaa Fixed Rate CMBS yields computed by Bloomberg/Morgan Stanley, and an index of 30-year fixed rate residential prime mortgage rates computed by Freddie Mac. By construction these latter rates correspond to yields on MBS of Aaa rating and consist of contract interest rates on commitments for fixed-rate 30 years prime mortgages. In the case of private-label Aaa CMBS, we compute a spread with reference to the closest (off-the run) 5-year Treasury. The choice of an off-the run Treasury allows us to attribute the CMBS spread to credit risk in the form of a higher probability of future defaults on the mortgages included in the securitized pools vs. Treasuries. Before mid-2007 the spread oscillated between 50 and 70 bp which is consistently below the 76 bp pre-crisis mean. Interestingly and probably because the epicenter of the crisis was the residential subprime and not the top rated, commercial mortgage market, the spread increased only gradually starting in the late Spring of 2007. It exceeded 450 bp in March 2008, in correspondence to Bear Stearns' collapse. After a brief respite during the Spring of 2008, it spiked again during the Summer of 2008, peaking at a never-seen before level of 1770 bp in the week of Lehman's bankruptcy. Once more, the spread started its gradual decline in March 2009, stabilized around 800 bp in the late Spring of 2009. It subsequently showed a renewed downward trend after the Summer of 2009, down to 300 bp in the late Winter of 2010. Because the Fed announced the expansion of the TALF to include Aaa-rated CMBS in February 2009, the final decline in the spread started only in March 2009. Since then this spread has persisted at a level of approximately 290–310 bp, which however remains elevated relative to historical norms.

In the case of 30-year fixed rate residential prime mortgage rates, we compute a spread with reference to the closest 30-year Treasury bond. The picture offered by Fig. 2 in the case of the 30-year fixed agency mortgage prime spread is different from any of the panels considered before: this spread is barely affected by the crisis. In practice, over the period 2006–May 2008 this spread kept oscillating between 120 and 170 bp, which is close (but slightly more elevated) than the pre-crisis mean of 112 bp.[11] The spread started creeping up during the Spring of 2008 and reached a peak at the pinnacle of the crisis in September 2008 (briefly flirting with the 250 bp threshold). Presumably, the LSAP program announced in November 2008 (and implemented from early 2009) has contributed to drive down the 30-year fixed rate mortgage spread. In fact, this spread not only returned to its normal, pre-crisis levels (around 100 bp) by March 2009, but also subsequently kept declining until stabilizing around 50 bp in early 2010. However, after having touched a record of zero in early 2010, this spread has subsequently climbed to levels between 80 and 110 bp during 2011, which appear to be values similar to the long-run unconditional average for this series. The fact that this spread has not reflected the crisis and it has actually been reduced by the policy-makers' reactions should come as no surprise if LSAP were effective, as argued by Gagnon et al. (2011) among others.

Finally, we have also analyzed the Moody's Baa–Aaa corporate yield spread, the difference between the average yields of two portfolios of corporate bonds maintained and published by Moody's: a portfolio of Baa (i.e., the lowest investment grade rating) corporate bonds with maturities of approximately (at least) 20 years; a portfolio of similar, 20-year maturity bonds with Aaa rating. Given that the spread is based on portfolios that—at least as a first approximation—differ only in their ratings, this is an obvious credit risk premium that compensates for a differential likelihood of default. Fig. 2 shows the familiar GFC-dominated pattern. Until the end of the Summer of 2007, the default spread was oscillating in a narrow range of variation, between 80 and 100 bp. This appears completely typical of pre-crisis experiences, when the mean had been 98 bp. If anything, the spread appeared to gravitate towards the low-end of its typical range of variation, which may indicate some over-pricing of lower credit ratings. The ascent of the default spread started in early October 2007 and was initially measured, bringing it to approximately 150 bp by the end of August 2008. Once more, Lehman's default marked a turning point, as the spread spiked to reach 300 bp during September 2008. The ensuing financial distress took a few months to contaminate the long-term segments of the corporate bond market. The peak was reached in early December 2008, at 347 bp. The aggressive reaction by the Fed lowered the spread below 300 bp in February 2009, although a new local spike in excess of 300 bp occurred in April 2009. From that point on, the default spread stabilized and quickly decreased, reaching a "close-to-normal" level slightly in excess of 100 bp. However, in the second half of 2011 this spread has climbed back up to touch almost 150 bp, which can be explained with the Baa portfolio including bonds issued by banking corporations threatened by the spillovers of the European sovereign debt crisis.

Table 2 performs a comparison between means (medians), volatility (interquartile range) of spreads for three periods, classified according the dominant consensus that has appeared in Table 1: before the crisis (December 2001–July 2007, a sample of 296 weeks), during the crisis (Aug. 2008-June 2009, a sample of 100 weeks), and after the crisis (July 2009–December 2011, a sample of 131 weeks). The before-crisis period is easy to characterize: spreads were on average low, often lower than average spreads over the full-sample periods of data availability (unreported). The medians are also small and not very different from means, which is reflected by the modest and often not statistically significant skewness coefficients. The volatilities of the spreads are tiny, always between 5 and 36 bp per week, with moderate differences when compared to interquartile ranges. In the central crisis-related panel, all mean spreads increase, reaching levels between 2 and 9 times the pre-crisis means. The only exception concerns the 30-year fixed rate mortgage spread, whose mean increases by a timid 44%. In this case, medians are often quite different from the means. This is reflected by many positive and statistically significant skewness coefficients (see e.g., Manzoni, 2002). Moreover, both the standard deviations and the interquartile ranges of the spreads increase enormously during the crisis, ranging from 19 bp per week for the On-/Off-the run Treasury spread to 446 bp for the 5-year CMBS spread. The only exception is the 30-year mortgage rate, where all volatilities increase by a factor of between 2 and 30. Finally, all means and volatilities decline when moving from the crisis to the post-crisis period. In the case of 3 spreads, the post-crisis both means and medians are actually inferior to the pre-crisis ones; even more strikingly, for as many as 5 series out of 7, the post-crisis spread volatilities is inferior to the pre-crisis ones. There is only major exception to this pattern of lower means and volatilities in the post-crisis period: in the case of 5-year CMBS spreads, the post-crisis mean remains an abnormally high 318 bp

[11] However, the mean for this spread over a longer (1985–2007) period exceeds 130 bp, which may be taken to imply that the pre-crisis agency residential mortgage spread should have been considered "normal".

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(to be compared to a pre-crisis mean of 76 bp), while volatility remains high at 88 bp per week vs. 15 bp only before the crisis.

Fig. 3 offers a visual summary. While our comments to Table 1 have stressed means as a measure of location of a series, Fig. 3 presents the same information using two nonparametric statistics of location and dispersion, the median and the interquartile range. The upper panel shows that for 5 spreads out of 7, the crisis marks a clear peak in the spread levels, with the crisis (middle, red) bars ranging between 30% and 200% higher than the pre-crisis (left, green) bars. In most cases, the spreads stabilize back to the pre-crisis level in the post-crisis period (the right, yellow bars). The first exception is the 10-year Off-On-the run Treasury spread, where the visual impression is that there is no effect of the crisis. However, this is only due to the scale of the graph which, to accommodate the enormous variation in the 5-year CMBS spread, largely hides the qualitative variation in the liquidity spread. The second exception is the 30-year fixed rate residential mortgage spread, whose median settles in the post-crisis period to a level lower than before the crisis. The bottom panel of Fig. 3 depicts dynamics in the interquartile range. A pattern emerges: spreads became much more variable during the crisis than they were before. In terms of interquartile range, the increase was often between 5 and 20 times the level of the pre-crisis dispersion measure. Once more, the only exception is the 30-year mortgage spread, which has become *less* variable after the crisis. This should be expected as this outcome is likely to have been caused by the stabilizing effects of the LSAP program that the Fed has implemented during 2008 and up to 2010.

3. The empirical model

We base our empirical tests on a simple univariate time series benchmark for the change in the yield spread index (see e.g., Joutz & Maxwell, 2002; Manzoni, 2002),

$$\Delta s_t = \alpha \Delta s_{t-1} + \beta(s_{t-1} - \gamma) + \epsilon_t \qquad \epsilon_t \sim \text{IID}\left(0, \sigma^2\right), \tag{1}$$

where s_t is a yield spread, $\Delta s_t \equiv s_t - s_{t-1}$ is the change in the spread between week *t-1* and week *t*, α, β and γ are constant parameters to be estimated, and ϵ_t is a white noise shock. (1) has the structure of a classical partial adjustment model, in the sense that it implies that the change in spread between time *t-1* and *t* is *also* explained by the deviation of the spread at time *t-1* from some “benchmark” level, represented by the parameter γ. We have written “also” because the other

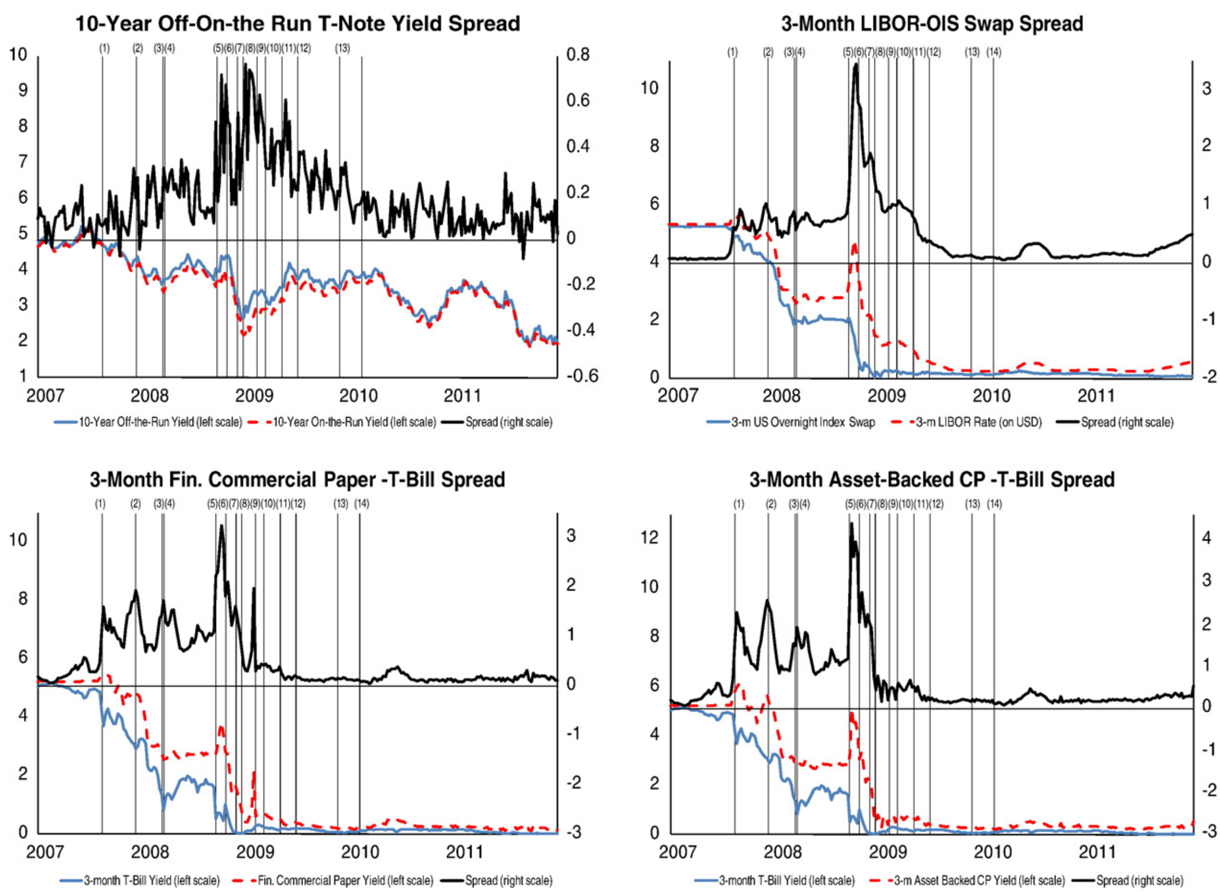

Fig. 2. Plots of yield spreads and of dates of key events in the financial crisis. [1] Aug. 2007 Fitch Ratings downgrades Countrywide Financial Co.; BNP Paribas halts redemptions for 3 investment funds. [2] Dec. 2007 Fed announces creation of Term Auction Facility (TAF); swap lines established with foreign central banks. [3] March 2008 Fed announces the creation of the Term Securities Lending Facility (TSLF); Bear Stearns rescued. [4] March 2008 Fed establishes the Primary Dealer Credit Facility (PDCF). [5] Sept. 2008 Lehman files for bankruptcy; Fed announces the Asset-Backed Commercial Paper Liquidity Facility (AMLF). [6] Oct. 2008 Fed announces the Commercial Paper Funding Facility (CPFF) and the Money Market Investor Funding Facility (MMIFF). [7] Nov. 2008 Fed announces the Term Asset-Backed Securities Lending Facility (TALF); asset purchase program (MBS and Treasuries) announced. [8] Dec. 2008 FOMC votes to establish a target range for the effective federal funds rate of 0% to 0.25%. [9] Feb. 2009 Fed announces extension of the existing liquidity programs. [10] March 2009 U.S. Treasury and Fed announce the launch of the TALF. [11] May 2009 Fed announces that CMBS will be eligible collateral under the TALF. [12] June 2009 Fed announces extensions of and modifications to a number of its liquidity programs. [13] Nov. 2009 Fed approves a reduction in the maximum maturity of credit at the discount window. [14] Feb. 2010 A number of liquidity programs (CPFF, ABCPMLF, TSLF) expire. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

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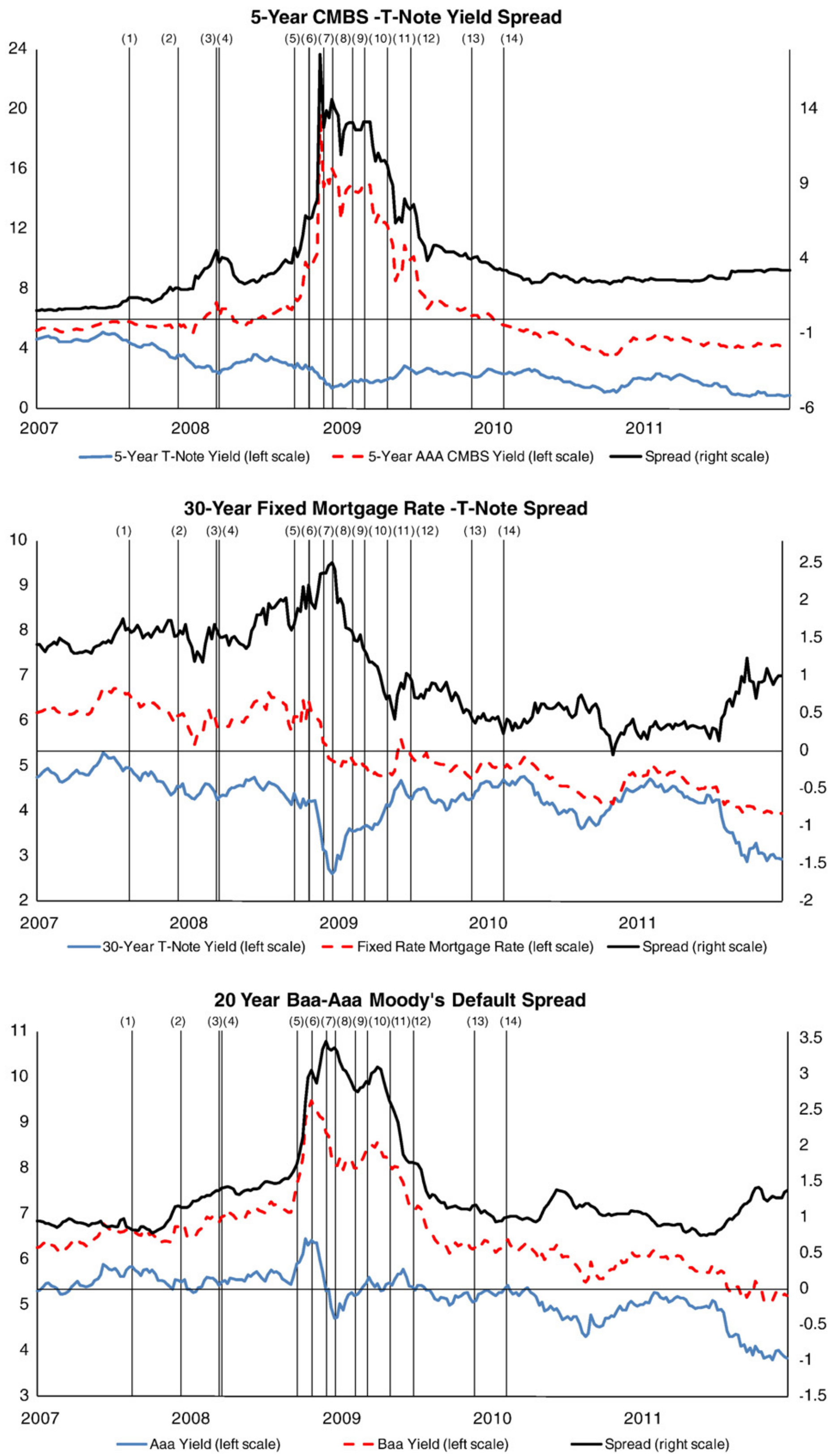

Fig. 2 (*continued*).

component that explains Δs_t is given by $\alpha\Delta s_{t-1}$, which is a traditional autoregressive term. For instance, when $1>\alpha>0$ and $\beta<0$, (1) implies that a portion α of the most recent change in the spread will keep propagating to time t as captured by the term $\alpha\Delta s_{t-1}$. At the same time, if in the previous period the spread has been higher than γ, then the spread will be reduced by $\beta(s_{t-1}-\gamma)<0$; if in the previous period the spread has been lower than γ, then the spread will increase by $\beta(s_{t-1}-\gamma)>0$. This is the sense in which (1) captures *mean reversion*

Table 2
Summary statistics for yield spreads: common pre-crisis, crisis, and post-crisis sample periods.Yield spreads are expressed in percentage, annualized basis points. The *p*-value for the median refers to a Wilcoxon signed rank test.

	Weekly Obs.	Mean	Median	St. Dev.	Interq. Range	Skewness	Excess Kurtosis
	Pre-Crisis Period (December 2001- July 2007)						
10-year Off–On the Run Treasuries	296	0.148 (0.000)	0.129 (0.000)	0.114	0.136	0.892 (0.026)	1.444 (0.185)
3-month LIBOR-OIS	296	0.106 (0.000)	0.098 (0.000)	0.045	0.057	0.983 (0.031)	2.109 (0.121)
3-month Fin. Comm. Paper-Treasury	296	0.157 (0.000)	0.130 (0.000)	0.099	0.120	1.285 (0.002)	1.838 (0.252)
3-month Asset-Backed Comm. Paper-Treasury	296	0.190 (0.000)	0.170 (0.000)	0.098	0.120	1.265 (0.002)	1.728 (0.130)
5-year Comm. MBS Rate-Treasury	296	0.764 (0.000)	0.720 (0.000)	0.153	0.170	1.270 (0.000)	1.432 (0.077)
30-year Fixed Rate Mortgage Rate-	296	1.115 (0.000)	1.125 (0.000)	0.360	0.673	−0.009 (0.932)	−1.442 (0.000)
20-year Aaa–Baa Moody's Default	296	0.977 (0.000)	0.930 (0.000)	0.207	0.273	0.504 (0.001)	0.616 (0.010)

	Weekly Obs.	Mean	Median	St. Dev	Interq. Range	Skewness	Excess Kurtosis
	Crisis Period (August 2007- June 2009)						
10-year Off–On the Run Treasuries	100	0.265 (0.000)	0.230 (0.000)	0.192	0.236	0.791 (0.031)	−0.031 (0.909)
3-month LIBOR-OIS	100	0.944 (0.000)	0.776 (0.000)	0.589	0.310	2.414 (0.038)	6.318 (0.089)
3-month Fin. Comm. Paper-Treasury	100	1.020 (0.000)	0.985 (0.000)	0.625	0.915	0.994 (0.107)	1.397 (0.250)
3-month Asset-Backed Comm. Paper-Treasury	100	1.255 (0.000)	1.075 (0.000)	0.846	1.040	1.344 (0.060)	2.169 (0.192)
5-year Comm. MBS Rate-Treasury	100	5.652 (0.000)	3.715 (0.000)	4.459	6.820	0.879 (0.003)	−0.635 (0.136)
30-year Fixed Rate Mortgage Rate-	100	1.601 (0.000)	1.600 (0.000)	0.401	0.405	−0.328 (0.461)	0.754 (0.250)
20-year Aaa–Baa Moody's Default	100	1.842 (0.000)	1.470 (0.000)	0.854	1.568	0.543 (0.006)	−1.242 (0.000)

	Weekly Obs	Mean	Median	St. Dev.	Interq. Range	Skewnes	Excess Kurtosis
	Post-Crisis Period (July 2009–December 2011)						
10-year Off–On the Run Treasuries	131	0.135 (0.000)	0.123 (0.000)	0.096	0.132	0.650 (0.729)	0.060 (0.039)
3-month LIBOR-OIS	131	0.188 (0.000)	0.149 (0.000)	0.104	0.118	1.113 (0.006)	0.341 (0.826)
3-month Fin. Comm. Paper-Treasury	131	0.148 (0.000)	0.130 (0.000)	0.062	0.030	1.715 (0.306)	2.983 (0.323)
3-month Asset-Backed Comm. Paper-Treasury	131	0.215 (0.000)	0.200 (0.000)	0.074	0.040	1.296 (0.852)	2.292 (0.377)
5-year Comm. MBS Rate-Treasury	131	3.188 (0.000)	2.830 (0.000)	0.876	1.050	2.209 (0.151)	6.393 (0.243)
30-year Fixed Rate Mortgage Rate-	131	0.531 (0.000)	0.490 (0.000)	0.254	0.360	0.445 (0.926)	−1.442 (0.001)
20-year Aaa–Baa Moody's Default	131	1.105 (0.000)	1.070 (0.000)	0.205	0.200	0.079 (0.006)	1.158 (0.000)

towards γ when $\beta<0$ and conversely *mean aversion* away from γ when $\beta>0$.(1) is consistent with Hendry, Pagan, and Sargan's (1984) view of error-correction models as reparameterizations of dynamic linear regression models in terms of differences and levels.[12]

It is easy to devise simulations to show that in the mean-reverting case of $\beta<0$, the spreads tends to converge towards γ and then tends to oscillate around it, while in the mean-averting case of $\beta>0$, any shock will cause the spread to permanently drift away from γ. In particular, when $\beta>0$ and the spread is initialized to be above γ, it diverges to $+\infty$. This is not economically plausible (it means that the price of one of the underlying FI products must vanish). Even worse, if the spread is initialized to be below γ, then it diverges to $-\infty$ and it becomes negative in finite time. Because all the spreads we are examining in this paper have a clear interpretation as risk premia, it is clear that to think of a permanently negative (in fact, diverging) risk premium makes little sense. Therefore (1) is an implausible model unless $\beta<0$. In the knife-edge case of $\beta=0$, (1) simplifies to $\Delta s_t=\alpha\Delta s_{t-1}+\epsilon_t$, which means that Δs_t is a simple AR(1) model. In this case, $s_t=(1+\alpha)s_{t-1}-\alpha s_{t-2}+\epsilon_t$, a (non-stationary) AR(2) model with no intercept and with the two autoregressive coefficients restricted to be linear functions of a single parameter α. In fact, when $\alpha=0$, Δs_t becomes a white noise process with zero mean, $s_t=s_{t-1}+\epsilon_t$, which is a classical random walk process with no drift. This means that in finite time, s_t is bound to become negative. and that its first-moment is not defined. Both are unattractive properties for a yield spread. Because in the case $\alpha=\beta=0$, the spread is a random walk, we also know that it can be written as $s_t=\sum_{j=0}^{t}\epsilon_j$, which shows that any of the shocks ϵ_j will affect the spread forever, i.e., the process has infinite memory. These properties explain why not only $\beta>0$, but also $\beta=0$ has to be thought of as implausible.

Another useful perspective comes from noticing that Eq. (1) can be re-written as

$$s_t=(1+\alpha)s_{t-1}-\alpha s_{t-2}+\beta(s_{t-1}-\gamma)+\epsilon_t=-\beta\gamma+(1+\alpha+\beta)s_{t-1} -\alpha s_{t-2}+\epsilon_t=\phi_0+\phi_1 s_{t-1}+\phi_2 s_{t-2}+\epsilon_t \tag{2}$$

which is an AR(2) model with cross-coefficient restrictions as $\phi_0=-\beta\gamma$, $\phi_1=1+\alpha+\beta$, and $\phi_2=-\alpha$. Interestingly, although the representation in Eq. (1) is the one with the strongest underlying economic intuition, in the applied econometrics literature, the representation in Eq. (2) and its equivalence to (1) is what seems to have drawn the attention to (1) itself.[13] Notice that because γ corresponds to the unconditional mean of the AR(2) representation, assumed to exist, i.e.,

$$E[s_t]=\frac{\phi_0}{1-\phi_1-\phi_2}=\frac{-\beta\gamma}{1-(1+\alpha+\beta)-(-\alpha)}=\gamma, \tag{3}$$

the error correction model may be equivalently interpreted as stating that the change in the spreads is associated with the past movement in the spread plus a portion β of the deviation from the long-run equilibrium level, identical to the unconditional mean $E[s_t]=\gamma$. This is another advantage of the representation in (1): the long-run mean of the process has become an explicit, estimable parameter.

[12] This *univariate* error correction model (ECM) is not the same as (multivariate) ECMs employed in cointegration analysis (e.g., see Joutz & Maxwell, 2002), where a *multivariate* model is internally consistent only if the variables are cointegrated.

[13] For instance, Nickell (1985) has commented that "Since it is almost a stylized fact that aggregate quantity variables in economics follow a second order autoregression with a root close to unity, we may expect to find the error correction mechanism appearing in many different contexts." (p. 124). Nickell also shows that a random walk with a moving average error gives rise to an error correction-type equation that shares many features with Eq. (1).

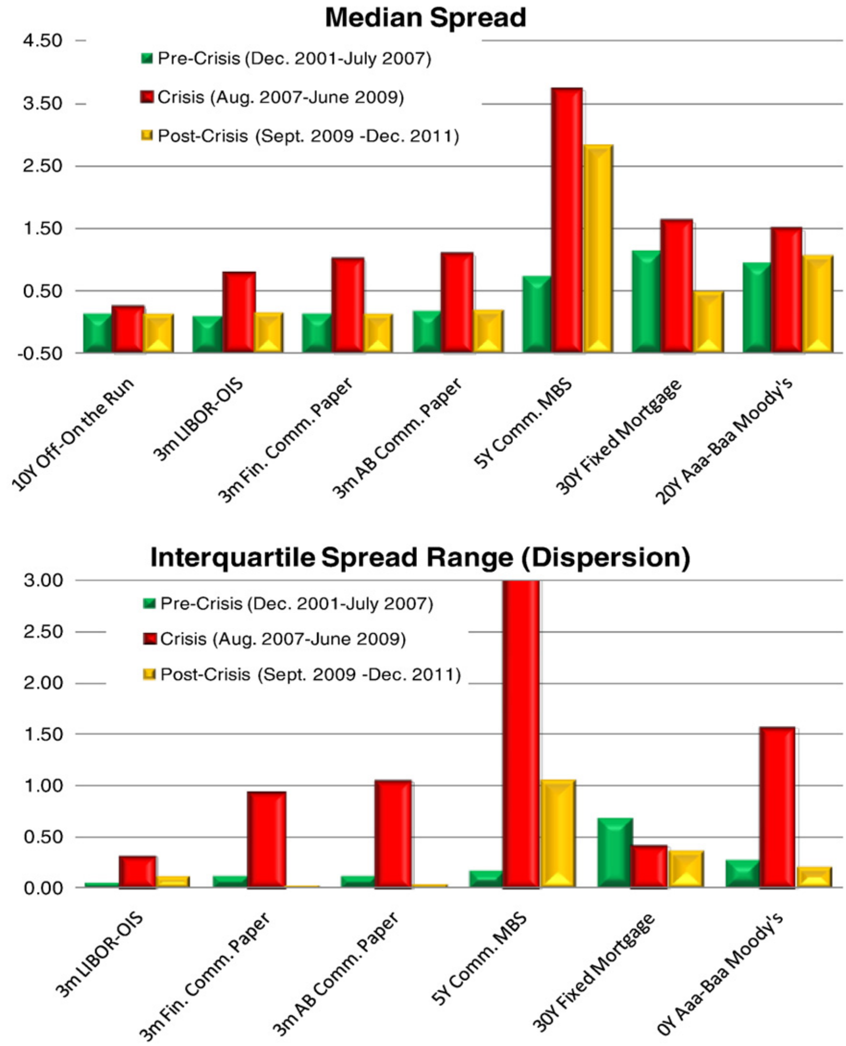

Fig. 3. Nonparametric location and scale statistics for yield spreads. The two bar diagrams compare median spreads (top panel) and the interquartile spread range (i.e., the difference between the 75th and the 25th percentile of their univariate empirical distribution) over three alternative periods: a common pre-crisis period (December 2001–July 2007), the crisis period (August 2007–June 2009), and for the post-crisis period (July 2009–December 2011). (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

3.1. The meaning of $\beta<0$

It is easy to show that $\beta<0$ is a (part of a set of) sufficient condition(s) that guarantees the covariance stationarity of (1). Therefore, $\beta<0$ not only ensures that the process (1) is economically sensible, but also that it is "well behaved". This is easily seen exploiting the (2) representation, $(1-\phi_1 L-\phi_2 L^2)s_t=\phi_0+\epsilon_t$, where L is the lag operator. This stochastic difference equation is stable and the AR(2) process covariance-stationary, provided that the roots of the equation $1-\phi_1 z-\phi_2 z^2=0$ lie outside the unit circle, or

$$|z_{1,2}|=\left|\frac{\phi_1\pm\sqrt{\phi_1^2+4\phi_2}}{2\phi_2}\right|=\left|\frac{(1+\alpha+\beta)\pm\sqrt{(1+\alpha+\beta)^2-4\alpha}}{-2\alpha}\right|>1. \tag{4}$$

If we set $\alpha=0$, then $\phi_2=0$ and the polynomial simplifies to an AR(1) characteristic polynomial, $(1-\phi_1 L)s_t=\phi_0+\epsilon_t$, which is covariance stationary provided that $|1/\phi_1|=|1/(1+\beta)|>1$, and this requires $\beta<0$. In general, when $\alpha\neq 0$, whether or not all the roots from (4) lie outside the unit circle will be a complicated function of both α and β. However, it is easy to compute that a $\alpha_{min}\simeq-0.5$ exists such that if $\alpha_{min}<0<\alpha<1$ and $-1<\beta<0$ (simultaneously) are jointly sufficient (but not necessary) for the roots of the AR(2) characteristic polynomial to fall outside the unit circle. This sufficient condition has an appealing interpretation if applied to the original, partial error correction representation (1): $\alpha_{min}<0<\alpha<1$ is a restriction to the standard stationarity condition within a simple AR(1) model; $-1<\beta<0$ satisfies the same intuition provided above, where $\beta>-1$ is to be considered innocuous as our empirical estimates in Section 4.2 will reveal that β tends to be always negative.

3.2. Testing for break points

Because model (1)–(2) is completely described by its parameters, model stability is equivalent to parameter stability. A large literature has emerged in econometrics that develops tests of model stability. One of the most common tests is Chow's (1960) simple split-sample test. This test is designed to test the null hypothesis of constant parameters against an alternative of a one-time shift in the parameters *at some known time*. The idea of the breakpoint Chow test is to fit a given model separately for each of the two (or $N\geq 2$) subsamples generated by a fixed, given break date and to see whether there are significant differences in the parameters of the estimated

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equations. A significant difference indicates a structural change in the relationship. In the case of (1)–(2), the Chow breakpoint F-statistic is based on the comparison of the restricted and unrestricted sum of squared residuals and in the simplest case involving a single breakpoint, is computed as

$$F = \frac{[\tilde{\epsilon}'\tilde{\epsilon} - (\epsilon'_1\epsilon_1 + \epsilon'_2\epsilon_2)]/3}{(\epsilon'_1\epsilon_1 + \epsilon'_2\epsilon_2)/(T-6),} \tag{5}$$

where ϵ_j is the $T_j \times 1$ vector of residuals when the model is estimated on some sub-samples of T_j observations, $\tilde{\epsilon}'\tilde{\epsilon}$ is the restricted sum of squared residuals when no break is imposed, and $\epsilon'_j\epsilon_j$ is the sum of squared residuals from the subsample $j = 1,2$. Assuming the candidate breakpoint date is exogenous, the F-statistic has an exact finite sample F-distribution if the errors are i.i.d. and normal. The log likelihood ratio (LR) statistic is based on the comparison of the restricted and unrestricted maximum of the (Gaussian) log likelihood function and has an asymptotic distribution with degrees of freedom equal to 3 and *T-6* under the null hypothesis of no structural change.

As an alternative to the classical Chow test, tests for structural change for every breakpoint can be calculated. Although this test was originally proposed by Quandt (1960), a distributional theory has been developed in Andrews (1993) and Hansen (1997). The resulting test is the Quandt–Andrews breakpoint test for one or more *unknown structural breakpoints*. Call $\hat{\theta} \equiv \left[\hat{\alpha}\ \hat{\beta}\ \hat{\gamma}\right]'$ and let $[\pi T] = T_1$ denote the integer part of πT where $0 < \pi_L \le \pi \le \pi_H < 1$. Thus, the proportion π defines sub-period 1, $t = 1,\ldots,T_1$. Under the null hypothesis, (1)-(2) is stable for the entire sample period. Under the alternative hypothesis, the model characterized by the estimator $\hat{\theta}_1(\pi)$ applies to observations 1, …,$[\pi T]$ and model with $\hat{\theta}_2(\pi)$ applies to the remaining $T - [\pi T]$ observations. This describes a nonstandard sort of hypothesis test since under the null hypothesis, the parameter of interest, π, is not part of the model. At this point, the idea is that a single Chow test is performed at every observation between two dates, τ_L and τ_H, where $\tau_L \equiv [\pi_L T]$ and $\tau_H \equiv [\pi_H T]$. The $[(1 - \pi_L - \pi_H)T]$ test statistics from these Chow tests are summarized into one test statistic for a test against the null hypothesis of no breakpoints between τ_L and τ_H, where $\pi_L + \pi_H$ is the percentage of observations set aside and not used to test for breaks. From each individual Chow test, two statistics are usually reported: the Likelihood Ratio F-statistic and the Wald F-statistic. Conditioning on π being fixed, the two test statistics for testing the hypothesis of model constancy against the alternative of structural break at πT are as follows. The Wald statistic is[14]

$$W_T(\pi) = \left[\hat{\theta}_1(\pi) - \hat{\theta}_2(\pi)\right]'\left\{\hat{\mathbf{V}}_1(\pi) + \hat{\mathbf{V}}_2(\pi)\right\}^{-1}\left[\hat{\theta}_1(\pi) - \hat{\theta}_2(\pi)\right], \tag{6}$$

where $\hat{\mathbf{V}}_j(\pi)$ is the (asymptotic) covariance matrix estimator for θ from the 1,…,$[\pi T]$ sample in the case of $j = 1$, and from the $[\pi T]+1$, $[\pi T]+2$, …, T sample in the case of $j = 2$. The likelihood ratio-like statistic is

$$\mathrm{LR}_T(\pi) = \left[L_1\left(\pi\middle|\hat{\theta}_1(\pi)\right) + L_2\left(\pi\middle|\hat{\theta}_2(\pi)\right)\right] - \left[L_1\left(\pi\middle|\hat{\theta}\right) + L_2\left(\pi\middle|\hat{\theta}\right)\right] \tag{7}$$

where $\hat{\theta}$ is based on the full sample. In both cases, the statistic has a limiting chi-squared distribution with K degrees of freedom, where K is the number of parameters in the model, $K = 3$ in the case of Eqs. (1)–(2).

Since π is unknown, the two tests presented above do not solve the problem posed at the outset. Andrews (1993) has derived the behavior of these test statistics by Monte Carlo by simulating it over a range of candidate values for π. This means, for different partitionings of the sample in the interval $[\pi_L,\pi_H]$ and retaining a few functions of the sequences of values obtained, for instance their maximum value for $\pi \in [\pi_L,\pi_H]$. Andrews (1993) and Andrews and Ploberger (1994) have derived the non-standard asymptotic distributions for three statistics that summarize the behavior of $W_T(\pi)$ and $LR_T(\pi)$ as π changes. Among these we find the widely employed maximum (also called Sup) statistics[15]:

$$\mathrm{Max}W_T(\pi) = max_{\pi_L \le \pi \le \pi_H} W_T(\pi) \ \text{ and } \ \mathrm{Max}LR_T(\pi) = max_{\pi_L \le \pi \le \pi_H} LR_T(\pi)\,. \tag{8}$$

Hansen (1997) has provided approximate asymptotic p-values which are used in our empirical work. The distribution of these statistics becomes degenerate as $\pi_L \to 0^+$ or $\pi_H \to 1^-$ i.e., when we approach the beginning or the end of the sample. To compensate for this behavior, it is suggested that the ends of the sample not be included in the testing procedure, by setting $\pi_L = \delta > 0$ and $\pi_H = 1 - \delta < 1$ with the trimming parameter δ typically between 5 and 10% of the sample. We use a 10% trimming throughout.

4. Empirical results

4.1. Are the spreads stationary?

Our first step consists of verifying that it is sensible to model spreads using a covariance stationary model with structure (1)–(2). In particular, because Eq. (2) needs to be covariance stationary, it is important to start by asking whether the FI yield spreads under investigation may contain a unit root.[16] Table 3 reports the results of a standard Augmented Dickey–Fuller (ADF) test, when the number of lags of changes in the spread to be included in the underlying model is selected by minimization of the BIC information criterion with a maximum number of lags equal to 12. The table also reports the results from an alternative, nonparametric Phillips–Perron (PP) test that controls for serial correlation when testing for a unit root induced by violation of the classical Dickey and Fuller's AR(1) framework.[17] In the table, boldfaced p-values indicate that the null of a unit root is rejected with a p-value of 10% or lower, an indication of covariance stationarity for the yield spread series examined.

Table 3 shows that all of the yield spread series under consideration are covariance stationary. In the case of the ADF test, the evidence is overwhelming: in 6 cases out of 7 the ADF p-value is actually lower than 5%, while in only one case the ADF p-value is between 5% and 10%, which still represents evidence against the null of a unit root. The evidence in favor of covariance stationarity of the spreads is confirmed by the PP test.[18] We have also repeated these tests with reference to the

[14] There is a small complication with this result in a time-series context. The two subsamples are generally not independent so using $\hat{\mathbf{V}}_1(\pi) + \hat{\mathbf{V}}_2(\pi)$ as an estimator for the covariance matrix of $\hat{\theta}_1(\pi) - \hat{\theta}_2(\pi)$ may be inappropriate. However, asymptotically the number of observations close to the switch point, if there is one, becomes small, so this is only a finite sample problem.

[15] Two alternatives to the Sup are suggested by Andrews and Ploberger (1994) and Sowell (1996). The average statistics, Avg$W_T(\pi)$ and Avg$LR_T(\pi)$, are computed by taking the sample average of the sequence of values over the $[(1 - \pi_L - \pi_H)T]$ partitions of the sample for $\pi \in [\pi_L,\pi_H]$. The exponential statistics are computed as Exp$W_T(\pi) = ln\{([(1 - \pi_L - \pi_H)T])^{-1} \int_{\pi \in [\pi_L,\pi H]} exp[0.5W_T(\pi)]d\pi\}$ and likewise for the exponential LR statistics. However, Andrews and Ploberger (1994) suggest that the Exp LR and Avg LR versions may often be less than optimal.

[16] In economic terms, we know already the answer: because a spread containing a unit root will eventually become negative and spend an infinite time providing negative compensation to credit and liquidity risks, this hardly makes sense. See also the discussion in Batten et al. (2005).

[17] The PP method estimates the non-augmented DF test equation and modifies the t-ratio of the key coefficient so that serial correlation does not affect the asymptotic distribution of the test statistic. The residual spectrum at frequency zero is estimated using a Bartlett kernel-based sum-of-covariances with a Newey–West bandwidth. In both the ADF and PP tests, the “exogenous regressors” are simply a constant intercept as it is implausible to find time trends in risk premia.

[18] The only series for which the rejection of the null of a unit root occurs with a p-value between 5 and 10% is for the 5-year CMBS-Treasury spread. However, we have to remind ourselves that the vast majority of unit root tests have non-stationarity, i.e., a unit root as their null hypothesis.

Table 3
Unit Root Tests on Yield Spread Series.
The table reports the results from the application of two types of unit root tests to yield spread over the full sample period Jan. 1985–December 2011. The two unit root tests are the standard Augmented Dickey–Fuller (ADF) test, when the number of lags of changes in the spread to be included is selected by minimization of the BIC information criterion with a maximum number of lags equal to 12; and the nonparametric Phillips–Perron (PP) test that controls for serial correlation when testing for a unit root induced by violation of the classical Dickey and Fuller's AR(1) framework. In the case of the PP test, the residual spectrum at frequency zero is estimated using a Bartlett kernel-based sum-of-covariances with a Newey–West bandwidth. In both the ADF and PP tests, the "exogenous regressors" are simply represented by a constant intercept. Boldfaced *p*-values indicate that the null of a unit root may be rejected with a *p*-value of 10% or lower.

	Starting date	Weekly Obs.	Diff.	Augmented Dickey-Fuller Test			Phillips-Perron Test		
				ADF t- Statistic	*p*-Value	BIC-based Lag Length	PP Adj. statistic	*p*-Value	Bandwidth
Full sample analysis (January 1985–December 2011)									
10-year Off-On the Run Treasuries	Jan. 1985	1409	Level	−7.663	0.000	4	−30.545	0.000	25
			First-diff.	−17.027	0.000	11	−213.48	0.000	134
3-month LIBOR-OIS	Dec. 2001	526	Level	−2.908	0.045	4	−2.974	0.038	11
			First-diff.	−13.133	0.000	3	−12.555	0.000	26
3-month Fin. Comm. Paper- Treasury	Jan. 1985	1409	Level	−7.05	0.000	0	−6.542	0.000	13
			First-diff.	−39.604	0.000	0	−49.405	0.000	40
3-month Asset-Backed Comm. Paper-Treasury	Jan. 2001	574	Level	−3.961	0.002	4	−3.914	0.002	11
			First-diff.	−13.651	0.000	3	−23.473	0.000	19
5-year Comm. MBS Rate- Treasury	July 1996	809	Level	−2.282	0.060	9	−2.720	0.071	16
			First-diff.	−7.753	0.000	8	−37.001	0.000	15
30-year Fixed Rate Mortgage Rate-Treasury	Jan. 1985	1409	Level	−3.144	0.024	0	−3.083	0.044	2
			First-diff.	−39.091	0.000	0	−39.101	0.000	3
20-year Aaa–Baa Moody's Default Spread	Jan. 1985	1409	Level	−2.983	0.037	1	−3.211	0.020	20
			First-diff.	−25.162	0.000	0	−25.964	0.000	17

common pre-crisis sample period (December 2001–July 2007) in Table 3, finding identical results. All in all, we conclude that a (1)–(2) representation may be consistent with stationarity of the underlying weekly spread series.[19]

4.2. Model estimates

Estimate model (1) for a few alternative sub-periods.[20] The results are reported in Table 4. A general result emerges: for all sample periods, the estimated model turns out to be covariance stationary, in the sense that the estimated coefficients $\hat{\theta}\equiv\left[\hat{\alpha}\;\hat{\beta}\;\hat{\gamma}\right]'$ map into $\hat{\phi}\equiv\left[\hat{\phi}_0\;\hat{\phi}_1\;\hat{\phi}_2\right]'$ vectors that satisfy Eq. (4). This explains why in Table 4 the estimated half-life of a shock is always a finite value, which is an implication of covariance stationarity. This is a first important finding: even in the midst of the Great 2007–2008 Financial Crisis, FI markets never unravelled to the point of implying non-stationary yield spread dynamics, which would imply an infinite half-life of a shock, i.e., that whatever shock would never be re-absorbed.[21]

The first panel of Table 4 shows full-sample results.[22] $\hat{\beta}$ is negative and statistically significantly negative for all seven spreads. In fact, some yield spreads display a considerable speed of reversion to the mean, in particular the short-term (Off-On the run Treasury, Financial CP-Treasury, and ABCP-Treasury) spreads. These are all characterized by $\hat{\beta}$s below −0.05 and *p*-values of 0.00 that imply half-lives between 2 and 11 weeks, which are relatively short and tell us that in the underlying markets shocks have transient effects on risk premia. The long-term spreads are instead characterized by smaller estimates of $\hat{\beta}<0$, which imply considerably higher half-lives, around 1 year with a maximum of 68 weeks in the case of CMBS spreads. The estimates of the long-run mean γ are all quite plausible—ranging from 17 bp in the case of the Off-On the run spread to 215 bp in the case of the CMBS spread—and statistically significant. Once more the only exception occurs for the CMBS spread, in which case the *p*-value of $\hat{\gamma}$ is 0.07. Most of these values, for instance the approximate 112 bp per year for the Baa–Aaa spread, conform to the priors that are usually reported in the finance literature. A 17 bp per year for the Off–On the run spread confirms the existence of precisely estimated, but also modest, liquidity premium. Finally, the estimates of the autoregressive terms α tend to be "all over the map" (with both positive and negative signs) and in some cases are not statistically significant, even though this parameter plays only an indirect role in the determination of the covariance stationarity of the spread series. Interestingly, even though (1) has a very stylized structure that obviously fails to account for a number of important influences, Table 4 shows that the model generally offers a good fit to the data, with $\bar{R}^2$ peaks in excess of 10% for 3 spreads, consistent with Davies (2008).

The second panel of Table 4 offers similar evidence with reference to a common pre-sample period, December 2001–July 2007. In qualitative terms, there are no major changes from the full-sample period, although here all but one of the estimates of β are lower (more negative) and still highly statistically significant. Together with the values for $\hat{\alpha}$, these estimates imply half-lives of shocks that are systematically lower than before, between 2 and 7 weeks in the case of the short-term spreads (Off-On the run Treasury, LIBOR-OIS, Financial CP, and ABCP), and of 21 weeks for both the CMBS spread and the Baa–Aaa spread. The only exception occurs with reference to the 30-year fixed mortgage spread, whose half-life increases from 68 to 76 weeks, while the corresponding $\hat{\beta}$ increases to only −0.011 and fails to be statistically different from zero (this is the meaning of the coefficient being bold-faced in Table 4). All in all, this is evidence that all yield spreads were strongly mean-reverting before the financial crisis, with only the fixed rate mortgage spread close to the non-stationary borderline, implying substantial persistence of shocks. A further aspect of these estimation results is interesting: the pre-crisis period was characterized by implicit long-run spreads that were

[19] Using daily data, earlier papers (e.g., Joutz & Maxwell, 2002; Manzoni, 2002) have concluded that a range of alternative *daily* yield spreads are I(1) series and have therefore modeled their first-difference. However, these papers often imply that mean spread series are hardly different from zero. In this paper, we model weekly spread series and are able to identify positive, statistically significant and often high risk premia.

[20] The model parameters are estimated by nonlinear least squares (NLS) from Eq. (1). Of course, under the assumption of covariance stationarity, identical parameters can be recovered from MLE estimation of its AR(2) representation. However, we use this mapping in the reverse fashion only to compute the half-life of a shock and to check for covariance stationarity.

[21] Assuming covariance stationarity, one useful measure of persistence of a dynamic process such as Eqs. (1)–(2) is how long does it take for a shock to ϵ_t to be re-absorbed by the dynamic process for the yield spread. The Appendix shows that the half-life $\hat{\tau}$ of a σ_ϵ shock to ϵ_t (i.e., a one-standard deviation shock) can be computed by solving numerically the inequality in (10).

[22] Results across different yield definitions are not directly comparable because the series are available for different sample periods. The second panel of Table 4 shows pre-crisis, common sample evidence that is qualitative similar to the first panel.

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Table 4
Univariate partial correction model estimates: comparing pre- and post-crisis periods.
The table reports nonlinear least squares estimates for the homoskedastic error correction model for yield spread changes:

$$\Delta s_t = \alpha \Delta s_{t-1} + \beta(s_{t-1} - \gamma) + \varepsilon_t,$$

where st is the yield spread and t is a white noise shock with constant variance. *p*-values are obtained from Newey–West HAC standard errors. The "Half-Life" column reports the point estimate of the number of weeks needed for a one-standard deviation shock to t to produce 50% of the long-run effects implied by point parameter estimates reported in the table; the experiment is performed the initial spread equals its long-run expectation (here, the estimated parameter). In the last two panels of the table, besides the implied half-life we also report the change with respect to the previous panel, i.e., the change between the pre-crisis and crisis periods in the third panel, and the change between the post-crisis and the crisis periods in the fourth panel.

	α (persistence coeff.)	*p*-Value	β (mean-reversion coeff.)	*p*-Value	γ (long-run mean)	*p*-Value	Regr. SE	Adj. R^2	Half-life (weeks)
Full sample period (January 1985–December 2011)									
10-year Off–On the Run Treasuries	−0.2877	0.000	−0.3155	0.000	0.1650	0.000	0.110	0.285	2
3-month LIBOR-OIS	0.4699	0.000	−0.0341	0.000	0.2993	0.003	0.079	0.227	11
3-month Fin. Comm. Paper-Treasury	−0.0224	0.401	−0.0657	0.000	0.2832	0.000	0.119	0.033	11
3-month Asset-Backed Comm. Paper-Treas.	0.0900	0.031	−0.0584	0.000	0.3866	0.002	0.175	0.031	11
5-year Comm. MBS Rate-Treasury	−0.2906	0.000	−0.0131	0.045	2.1500	0.066	0.429	0.091	68
30-year Fixed Rate Mortgage Rate-Treas.	−0.0345	0.196	−0.0136	0.002	1.2358	0.000	0.086	0.007	51
20-year Aaa–Baa Moody's Default Spread	0.3828	0.000	−0.0080	0.003	1.1216	0.000	0.041	0.147	54
Common pre-crisis period (December 2001–July 2007)									
10-year Off–On the Run Treasuries	−0.1707	0.003	−0.3687	0.000	0.1450	0.000	0.093	0.241	2
3-month LIBOR-OIS	−0.1147	0.043	−0.1951	0.000	0.1047	0.000	0.027	0.116	4
3-month Fin. Comm. Paper-Treasury	0.1465	0.012	−0.0874	0.000	0.1629	0.000	0.038	0.051	7
3-month Asset-Backed Comm. Paper-Treas.	0.1498	0.010	−0.0961	0.000	0.1950	0.000	0.039	0.056	6
5-year Comm. MBS Rate-Treasury	−0.0592	0.308	−0.0346	0.008	0.7389	0.000	0.034	0.022	21
30-year Fixed Rate Mortgage Rate-Treas.	−0.1510	0.010	**−0.0105**	0.313	1.2013	0.001	0.064	0.021	76
20-year Aaa–Baa Moody's Default Spread	0.2841	0.000	−0.0237	0.025	0.9995	0.000	0.037	0.083	21
Crisis Period (July 2007–June 2009)									
10-year Off-On the Run Treasuries	−0.4257	0.000	−0.2269	0.004	0.2662	0.000	0.138	0.331	4 (+2)
3-month LIBOR-OIS	0.5108	0.000	−0.0613	0.003	0.9190	0.000	0.163	0.294	6 (+2)
3-month Fin. Comm. Paper-Treasury	0.0058	0.954	−0.1041	0.012	0.9652	0.000	0.322	0.046	7 (==)
3-month Asset-Backed Comm. Paper-Treas.	0.1453	0.143	−0.1055	0.009	1.1879	0.000	0.392	0.054	6 (==)
5-year Comm. MBS Rate-Treasury	−0.3097	0.001	**−0.0316**	0.221	8.0403	0.057	1.160	0.101	29 (+8)
30-year Fixed Rate Mortgage Rate-Treas.	−0.0878	0.388	**−0.0476**	0.185	1.5055	0.000	0.139	0.012	16 (−60)
20-year Aaa–Baa Moody's Default Spread	0.6105	0.000	**−0.0104**	0.131	1.9802	0.001	0.071	0.463	27 (+6)
Post-crisis period (July 2009–December 2011)									
10-year Off-On the Run Treasuries	−0.3741	0.000	−0.2886	0.000	0.1314	0.000	0.075	0.326	3 (−1)
3-month LIBOR-OIS	0.4901	0.000	**−0.0099**	0.480	0.2630	0.158	0.016	0.238	36 (+30)
3-month Fin. Comm. Paper-Treasury	−0.2052	0.017	−0.1687	0.003	0.1457	0.000	0.038	0.132	5 (−2)
3-month Asset-Backed Comm. Paper-Treas.	−0.4568	0.000	**−0.0640**	0.259	0.2509	0.000	0.042	0.193	16 (+10)
5-year Comm. MBS Rate-Treasury	0.1318	0.105	−0.0814	0.000	2.9120	0.000	0.186	0.197	8 (−21)
30-year Fixed Rate Mortgage Rate-Treas.	−0.1854	0.03	−0.0588	0.100	0.5594	0.000	0.101	−0.052	14 (−2)
20-year Aaa–Baa Moody's Default Spread	0.5228	0.000	−0.0405	0.005	1.0846	0.000	0.035	0.336	9 (−18)

very small, possibly surprisingly so. One is tempted to argue that they may have been "excessively" small, although the absence of a benchmark theoretical model is an obstacle to such a conclusion. All the estimates of γ are highly statistically significant.

The third panel concerns the 2007–2009 crisis period and contains some of our key results. Here, once again, the fundamental contrast is between the fixed rate mortgage spreads and all the remaining spreads. In general, all the $\hat{\beta}$ estimates uniformly increase (towards 0) when going from the pre-crisis to the crisis period. This implies a diminished speed of reversion towards the long run mean. Interestingly, most $\hat{\alpha}$ estimates increase in absolute value and 3 of them stop being statistically significant (i.e., during the crisis there is more autoregressive-type persistence in spread changes). Both effects contribute to a discrete jump in the half-life of shocks of most spreads, from +2 weeks in the case of Off-On the run and LIBOR-OIS spreads to +6 and 8 weeks for CMBS and corporate default spreads. In fact, in these two latter cases, $\hat{\beta}$ estimates remain negative but fail to be significant. The implication is that for 6 spreads, the financial crisis has implied a higher persistence of changes in the spread and a lower speed of reversion to its long-run mean for the level of the spread itself. In the perspective of a partial adjustment model such as (1), this is what a financial crisis is all about in FI markets: the risk premia (for both credit and liquidity risks) become highly persistent in the sense that any shocks—and during a crisis we can presume that many of these shocks will carry a positive sign—take a longer time to be re-absorbed. Needless to say, higher risk premia mean higher risk-adjusted discount rates when evaluating bonds, and lower (depressed) market valuations for riskier bonds.

Another—possibly obvious—way in which a financial crisis manifests itself is through the implied estimates of the long-run mean of the spreads, the $\hat{\gamma}$s. These all increase by a factor between 1.8 and 9 when we compare the estimates for the pre-crisis with the crisis sample; the smallest increase is a stunning 84% in the case of the Off-On the run spread (from 15 to 27 bp), while the largest increase—+878% (from 11 to 92 bp) for the LIBOR-OIS spread—hardly deserves any comment and has been the focus of considerable debate (see e.g., Christensen et al., 2009). The very levels of the $\hat{\gamma}$s are symptomatic of the crisis, with 3 short-term spreads close to 100 bp per year, two long-term spreads in excess of 150 bp, and the CMBS spread jumping to an unprecedented 804 bp. Yet, it is remarkable that (1) fits the data rather well during the financial crisis, with 4 $\bar{R}^2$ exceeding 10% and an impressive 46% for the corporate default spread.

The exception to the broad picture commented here deserves attention because it may have important implications for the effectiveness of the LSAP programs. The only yield spread series for which we record a substantial decline in both the implied half-life of a shock (persistence) and a negligible (+25%) increase in its long-run mean is the 30-year fixed mortgage rate spread, which seems to have been left relatively unscathed by the Great Financial Crisis. In fact, for this spread $\hat{\beta}$ even declines when going from the pre- to the crisis period (from −0.011 to −0.048, even though both estimates are not significant). This explains the dramatic decline in the half-life

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estimate from 76 to 16 weeks.[23] We attribute this singularity in the shifts undergone by the dynamic process characterizing the prime mortgage spread to the effectiveness of the LSAP programs implemented by the Fed. We return to this point in Section 5.3.

Obviously, it is difficult to miss the fact that a simple inspection of the second and third panels of Table 3 reveals an enormous amount of instability in most estimated coefficients as well as in the implied summary statistics. Some dramatic event—we now know it as the Great Financial Crisis—has enveloped the FI markets and structurally changed their dynamic properties in ways that would have been difficult to anticipate. This interpretation is further validated by a comparison of the third and fourth panels of the table: after the crisis was over, the model parameters shifted once more, in this case towards the pre-crisis levels (see below for specific comments). We formally test these hypotheses in Section 5.3.

Finally, the last panel of Table 4 reports estimation results for the post-crisis period, July 2009–December 2011. At least in a qualitative sense, all the relevant parameters revert back to values typical of the pre-crisis period. Most $\hat{\beta}$ estimates decline thus marking a renewed strength in the mean reversion speed of spreads. In fact, the $\hat{\beta}$ estimates for 5 out of 7 stabilize to levels that are below the ones estimated over the pre-crisis period.[24] For these 5 spread series, the implied half-life of a shock is indeed below the pre-crisis estimates with values between 3 and 14 weeks. In fact, also the half-life of shocks to mortgage rate spreads has substantially declined from 76 to 14 weeks. This means that these declining evolution of the $\hat{\beta}$ estimates has not been reversed by parallel breaks in the $\hat{\alpha}$ estimates reported in Table 4, fourth panel. We can summarize these developments by saying that by the second half of 2009, the financial crisis had stopped exercising its effects on the ability of (US) FI markets to self-correct towards their long-run equilibria. This is also visible in the estimates of the long-run yield spreads implied by (1): they all decline towards their pre-crisis levels, although in many cases they have remained above their pre-crisis estimates. However, four cases can be found (Off–On the run, 3-month financial CP, and 30-year fixed rate mortgages) in which the implied long-run spreads have actually declined below their 2002–2007 implied mean levels. These reversions of the estimated long-run spreads towards pre-crisis levels represents a further—in a sense, more obvious—way in which the financial crisis seems to have been over by June 2009. It is more ambiguous whether policy makers may rationally be concerned for the fact that a few of the $\hat{\gamma}$ estimates appear to have traced back to long-run levels that are *inferior* to their already modest pre-crisis levels.[25] It should not be considered surprising that $\hat{\gamma}$ for 30-year mortgage spreads has declined between 2009 and 2011 to an exceptionally low level of 56 bp. This low level is a likely result of the LSAP programs. However, Table 4 also stresses that any effects of the Fed policies did not really (or not only) affect the average spreads, but also and especially their "deep" dynamic properties as revealed by structural changes in the half-life of shocks to fixed mortgage rate spreads.[26]

Finally, unreported tests emphasize that the increase in the implied half-life of shocks for LIBOR-OIS and 3-month ABCP spreads in the post-crisis period vs. the pre-crisis one, entirely derives from the inclusion in our sample of 2011 data. Even though for these two spreads, the estimated $\hat{\gamma}$ (26 and 25 bp, respectively) remain considerably inferior to those recorded for the pre-crisis period (91 and 119 bp, respectively, i.e., between one-third and one-fourth), there is very recent evidence that limited segments of the US FI market may have been engulfed back into a state of turmoil in which shocks tend to be slowly re-absorbed, with half-lives between 4 and 8 months. Yet, the fact that only market segments more directly involved with banking funding operations may point to the fact that such properties of the dynamic spread process may not be the result of lingering GFC effects, but of a new financial storm at the horizon, in this case emanating from the European sovereign debt-induced tensions and no longer from the residential US housing markets.

4.3. Breakpoint tests

Table 5 formally tests for the presence of breaks in (1) and contains the other key result of the paper. The left portion of the table presents Andrews–Quandt break test results, when the date of the break is not assumed to be known and its assessment ("estimation") must be based on the data.[27] This is truly an "ignorance prior test" because it does not impose any structure on our search for evidence of potential breaks. The right portion of the table resorts instead to the more traditional Chow break test, in which the researcher needs to contribute her knowledge of the potential date of the breakpoints, with all the perils of the assumption. Clearly, Chow tests are much more efficient in a statistical sense when the researcher is able to feed sensible candidate break dates to the testing procedure.

The Andrews–Quandt break tests reveal evidence of only one endogenously determined break in the case of one series, the 3-month ABCP spread. There is evidence of 3 breaks in the remaining six series: the second break is obtained from an Andrews–Quandt test that conditions on a first break; the third breakpoint is estimated conditioning concerns on the first two breaks. In general, the three breakpoint dates obtained confirm the boom–bust–boom cycle that we would expect when a serious financial crisis impacts markets and resolves later on. However, for a few of the spreads, we have also evidence of a more recent, 2011 break that we may relate to the European sovereign and bank debt woes. The first break occurs between April 2007 (in the case of the 5-year CMBS spread) and early 2009 (for the 3-month financial CP spread).[28] All these breaks are detected at a very high level of statistical significance. Casual evidence matched with the empirical estimates in Table 4 leads us to label this break as a "crisis onset" shift. It is also sensible to find that this first break occurs first in 5-year private label CMBS spreads than elsewhere, as the crisis originated in the mortgage markets in which private originators were particularly active. A second break affects at least 5 of the spread series with a dating that ranges between March 2009 (for the 3-month LIBOR-OIS spread) and December 2009 (for the off-on the run Treasury spread). Based on the evidence in Table 4, it is sensible to interpret the second set of breakpoint dates in the light of FI

[23] However, the already low $\bar{R}^2$ (2.1%) of the pre-crisis sample further declines (to 1.2%) in the crisis sample.

[24] The two exceptions are the LIBOR-OIS and the ABCP spreads for which the post-crisis $\hat{\beta}$s are −0.01 and −0.06, respectively, vs. a pre-crisis estimates of −0.20 and −0.10. Oddly enough, the LIBOR-OIS spread half-life has increased from 4 to 36 weeks. This may be related to the growing pressure on the European fixed income markets later surfaced in the Spring of 2010 with reference to Greek bail-out and the refinancing difficulties experienced by a few other EU countries, such as Spain and Portugal. Because markets have feared a contagion to the global banking system, the LIBOR-OIS and ABCP spreads may have captured the price of escalating funding risks.

[25] Some commentators (see e.g., Courtois, Gaines, & Hatchando, 2010) have written about the hazards of re-inflating asset price bubbles by pushing bond prices (risk premia) too high (low).

[26] However, because LSAP programs have also concerned commercial MBS through extensions of the TALF program during 2009, it is unclear why they have so far failed to produce a repairing influence on the CMBS market of a comparable extent to the impact caused on the fixed rate residential mortgage market.

[27] Both types of break tests are applied sequentially, in the sense that when the occurrence of a break is isolated (i.e., the null of no break is rejected), tests for additional breaks are applied conditioning on the date of the first break. The Andrews-Quandt test is applied to sample observations after cutting the first 5% and the last 5% of the available observations. The last column of the table shows the possible ranges for break dates in the conditional mean function isolated by both sets of break tests.

[28] Rather oddly, the 3-month ABCP spread seems to have been subject to a break in early October 2008 and not to have emerged from such a new "regime" so far. This is inconsistent with both the empirical estimates in Table 4 and with the results from the Chow test to be reported shortly. In the case of the LIBOR-OIS spread, the tests also locate a break between April 2003 and June 2004 for which there is little intuition, even though it is clear from the data that this spread tends to be characterized by a lower mean and persistence after 2004 vs. 2001–2003.

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Table 5
Break Tests Applied to Univariate Partial Correction Models.
The table the outcomes of two break tests applied to nonlinear least squares estimates of a homoskedastic error correction model for yield spread changes:

$$\Delta s_t = \alpha \Delta s_{t-1} + \beta(s_{t-1} - \gamma) + \varepsilon_t,$$

where st is the yield spread and t is a white noise shock with constant variance. Both types of break tests are applied sequentially, in the sense that when the occurrence of a break is isolated (or fails to be rejected), tests for additional breaks are applied conditioning on the date (assumed or endogenously determined) of the first break. The left block of the table reports the outcomes of an Andrews–Quandt test in which break dates are left unspecified. The test is applied to sample observations after cutting the first 5% and the last 5% of the available observations. In the table, we report the Maximum LR statistic. The right block of the table reports instead the outcomes of a standard Chow break test in which the break dates are exogenously specified to correspond to the first week of August 2007 and the last week of June 2009. In the case of the Chow test, both the log-likelihood and the F test statics are reported. The last column of the table shows instead the possible ranges for break dates in the conditional mean function isolated by both sets of break tests.

	Andrews–Quandt Test		Andrews–Quandt Test Conditioning on First Break		Andrews–Quandt Test Conditioning on Second Break		Chow Break Test						Break Ranges
	Maximum LR *F*-statistic	Maximum Wald *F*-statistic	Maximum LR *F*-statistic	Maximum Wald *F*-statistic	Maximum LR FStatistic	Maximum Wald *F*-statistic	Dates	F Statistic	Log-LR Stat	Dates	F Statistic	Log-LR Stat	
10-year Off–On the	6.900 (0.005)	34.974 (0.000)	10.681 (0.000)	30.808 (0.000)	6.470 (0.009)	39.860 (0.000)	8/3/2007	**4.026**	**12.079**	8/3/2007	**2.977**	**17.86**	April 2003–July 2008
Run Treasuries	6/11/2004	4/11/2003	3/7/2008	3/7/2008	12/25/2009	1/1/2010		(0.007)	(0.008)	6/26/2009	(0.007)	(0.007)	June 2009–Dec. 2009
3-month LIBOR-OIS	29.802 (0.000)	89.406 (0.001)	7.781 (0.001)	23.343 (0.001)	8.764 (0.000)	26.292 (0.000)	8/3/2007	**9.282**	**27.438**	8/3/2007	**8.705**	**50.608**	Aug. 2007–Oct. 2008
	10/17/2008	10/17/2008	3/20/2009	3/20/2009	8/5/2011	8/5/2011		(0.000)	(0.000)	6/26/2009	(0.000)	(0.000)	March 2009–June 2009
													Third break: August 2011
3-month Fin. Comm.	50.902 (0.000)	152.702 (0.000)	49.865 (0.000)	149.595 (0.000)	9.244 (0.000)	27.731 (0.000)	8/3/2007	2.067	6.215	8/3/2007	**6.308**	**37.584**	Aug. 2007–Feb. 2009
Paper-Treasury	2/6/2009	2/6/2009	5/29/2009	5/29/2009	9/23/2011	9/23/2011		(0.103)	(0.102)	6/26/2009	(0.000)	(0.000)	May 2009–June 2009
													Third break: Sept. 2011
3-month Asset-Backed	59.590 (0.000)	178.771 (0.000)	0.522 (0.999)	1.567 (0.999)	___	___	8/3/2007	2.066	6.229	8/3/2007	**4.843**	**28.786**	Aug. 2007–Aug. 2008
Comm. Paper-Treasury	10/3/2008	10/3/2008	___	___	___	___		(0.104)	(0.101)	6/26/2009	(0.000)	(0.000)	June 2009
3-month LIBOR-OIS													
5-year Comm. MBS	9.873 (0.000)	29.618 (0.000)	6.388 (0.009)	19.164 (0.009)	5.195 (0.032)	15.585 (0.032)	8/3/2007	1.980	5.961	8/3/2007	**4.883**	**29.096**	Aug. 2007–April 2008
Rate-Treasury	4/27/2007	4/27/2007	6/12/2009	6/12/2009	9/18/2009	9/18/2009		(0.116)	(0.114)	6/26/2009	(0.000)	(0.000)	March 2009–June 2009
													Third break: Sept. 2009
30-year Fixed Rate Mortgage	7.574 (0.02)	21.720 (0.02)	10.693 (0.000)	32.079 (0.000)	8.399 (0.001)	25.196 (0.001)	8/3/2007	**3.666**	**11.003**	8/3/2007	**2.527**	**15.175**	May 2007–Dec. 2008
Rate-Treasury	5/18/2007	5/18/2007	1/9/2009	1/9/2009	8/5/2011	8/5/2011		(0.012)	(0.012)	6/26/2009	(0.020)	(0.019)	June 2009–Aug. 2011
20-year Aaa–Baa Moody's	41.281 (0.000)	124.485 (0.000)	14.188 (0.000)	42.563 (0.000)	11.276 (0.000)	33.827 (0.000)	8/3/2007	**38.771**	**112.216**	8/3/2007	**21.168**	**122.346**	July 2007–Oct. 2008
Default Spread	10/26/2007	10/26/2007	10/24/2008	10/24/2008	4/17/2009	4/17/2009		(0.000)	(0.000)	6/26/2009	(0.000)	(0.000)	April 2009–June 2009

markets having re-emerged from the crisis. More puzzling is the fact that three markets seemed to have been subject to one additional break: in September 2009 in the case of the 5-year CMBS series, in August 2011 for the LIBOR-OIS spread, and in September 2011 in the case of the 3-month financial commercial paper spread. The explanations for these additional breakpoints are heterogeneous. In the case of the late 2009 breakpoint, this may be due to the effects of the Fed expanding the TALF to include Aaa-rated CMBS during 2009, which has slowly improved conditions in many securitized mortgage product markets, including the private label CMBS one. As already commented above, both LIBOR-OIS and the 3-month financial commercial paper spreads are tightly linked to the access by US banks to coverage of their short-term funding needs. In this case, the Summer of 2011 has been marked by increasing propagation of market tensions caused by the European sovereign debt jitters, and the additional breakpoints isolated in Table 5 may indicate—in accordance with Table 4—that a few spread have returned in the spotlight of market tensions towards the end of our sample. However, many Readers may still find confusing the fact that the same test may generate rather different numbers of breakpoints as well as "estimated" dates across different spread series. Of course, this may be due to the low power that the Andrews–Quandt test tends to have because the breakpoint date is left unspecified.

To remedy to this drawback, the right portion of Table 5 reports the outcomes of a standard Chow break test in which the break dates are exogenously specified to correspond to the first week of August 2007 and the last week of June 2009. The first date is taken to represent the onset of the crisis; the second date is a candidate date for the end of the crisis. These exogenously picked breaks are consistent with the literature summarized in Table 1 as well as with the preliminary results in Table 5 from Andrews-style tests. Here the results conform with the evidence in Table 4: for 4 out of 7 yield spreads series, there is highly significant (i.e., with a p-value below 5%, the boldfaced statistics in the table) evidence of a break in early August 2007; weaker evidence (with a p-value around 10%) can also be found for 2 additional series. There are no economically important differences between the F and Log-LR versions of the Chow test. Additionally, for all the spreads under consideration, conditioning on a first break occurring in the first week of August 2007, we proceed to test for another breakpoint at the end of June 2009. We always find evidence of a second break, which we interpret as evidence of the end of the GFC. Interestingly, when both break dates are jointly specified and tested, both the F and Log-LR tests always yield p-values that are very small.

The last column of Table 5 provides a summary of the breakpoint test results across different methodologies. Clearly, this summary may provide economic intuition, but has no statistical foundation, as it is impossible to take "averages" of breakpoint dates across methodologies. For all spreads, there is ample evidence of a range of break dates. For six out of seven series there is compelling evidence of one "crisis-triggering" break occurring between May 2007 and August 2008. While real estate spreads seem to break first, for a few other spreads related to the funding of financial institutions (the 3-month CP and ABCP spreads) and of corporate entities in general (the Baa–Aaa spread) the evidence points towards a later break, as late as the Fall of 2008, in connection to Lehman's collapse. This dating of course fits our summary of the main events in Section 2.1. Furthermore, all spread series seem to have left a crisis state between March and May 2009 (for the 5-year CMBS, the 3-month LIBOR-OIS, and the financial CP spreads) and June 2009, which is what we would have conjectured on the basis of the commentary of Section 2.1. Yet, three series seemed to be affected by one additional break occurring between late 2009 and the Summer of 2011. In the latter case (affecting the LIBOR-OIS and the 3-month financial commercial paper spreads), it is possible to link this additional evidence of instability to market turmoil caused by European sovereign debt concerns.

4.4. Exploring the mechanism: how did the crisis affect bond markets?

A finding that any of the coefficients in $\theta \equiv [\alpha\,\beta\,\gamma]'$ in Eq. (1) is subject to one or more structural breaks over the sample period, is not completely informative because it fails to give adequate information on whether the breakpoint in the conditional mean process for $\Delta s_t \equiv \Delta y_t^{\text{high}} - \Delta y_t^{\text{low}}$—where Δy_t^{high} is the series of changes in yields for the bond with the highest (lowest) credit risk (liquidity) and Δy_t^{low} is the series changes in yields for the bond with the lowest (highest) credit risk (liquidity)—derives from the presence of breakpoints in the dynamic process for either Δy_t^{high} or Δy_t^{low}. We therefore proceed to a further decomposition of our results through the estimation of the bivariate seemingly unrelated models

$$\begin{cases} \Delta y_t^{\text{high}} = \alpha \Delta ffr_{t-1} + \beta^{\text{h}}(s_{t-1} - \gamma) + \epsilon_t^{\text{h}} \\ \Delta y_t^{\text{low}} = \alpha \Delta ffr_{t-1} + \beta^{\text{l}}(s_{t-1} - \gamma) + \epsilon_t^{\text{l}}, \end{cases} \tag{9}$$

where ffr_t is the effective federal funds rate and $\epsilon_t \equiv [\epsilon_t^h \epsilon_t^l]' \sim$ IID $N(0,\Sigma)$. Notice that the off-diagonal element of Σ, $Cov[\epsilon_t^h, \epsilon_t^l]$, represents the covariance of shocks affecting the two yield series. Eq. (9) represents a restricted SUR bivariate regression because the coefficient α loading on past changes in *ffr* and the coefficient γ to which the reversion yields approaches are common across the two equations in the model. Eq. (9) is similar in spirit to the partial error correction model (1) used early on, although there are important differences. First, current changes in yields are modeled as depending in time *t-1* changes in *ffr*, where *ffr* is used to capture expected movements in interest rates that propagate from monetary policy actions to the entire yield curve.[29] Second, in this case the correction term has structure $\beta^n(s_{t-1} - \gamma)$ $(n = h,l)$ indicating that when $s_{t-1} \gtrless \gamma$ then the yield would decrease (increase) when $\beta^n < 0$, and it would increase (decrease) when $\beta^n > 0$. However, because it models yield changes on the left-hand side as a function of deviations of spread from some benchmark level γ, Eq. (9) does not represent a formal error correction model, it does not have a (vector) autoregressive equivalent representation, and γ does not represent the long-run conditional mean of any of the two yield series in EQ. (9). However, Eq. (9) does capture the logic that (risky) yields adjust when the short-end of the risk-free yield curve moves (e.g., by an expectations hypothesis effect) and when past credit and liquidity risk premia appear to have deviated from some long-run "norm".

Although the structure of (9) does not allow us to formally connect the sign or magnitude of the coefficients β^{h} and β^{l} to the covariance stationarity of the process, we have two testable formal hypotheses concerning these adjustment coefficients:

1. If the yield spread is stationary and mean-reverting, we would expect that $\beta^{\text{h}} \leq 0$ and $\beta^{\text{l}} \geq 0$ (with at least one strict inequality), i.e. when the spread exceeds some norm γ, both yields should contribute to the adjustment, the yield on riskier (less liquid) bond by adjusting downwards, and the yield on the less risky (more liquid) bond by soaring.
2. Unless $\beta^{\text{h}} \neq \beta^{\text{l}}$ no adjustment in the spread is possible, although this is a weak necessary condition (one may formulate a sharper condition that $\beta^{\text{h}} < \beta^{\text{l}}$, although signs matter as much as magnitudes).

It may be also of interest to test $\alpha > 0$ (or $\alpha \neq 0$), which is equivalent to an expectation hypothesis effect on yield changes, if we take *ffr* as the rate representative of the short-end of the yield curve.

[29] We have also estimated a variety of models like (9) in which current changes in yields dependend on their most recent change (e.g., Δy_t^{high} on $\Delta y_{t-1}^{\text{high}}$, etc.) or on the most recent change of the spread, but did not find any substantial differences. When α has been allowed to differ across equations (i.e., they become $\Delta y_t^n = \alpha^n \Delta ffr_{t-1} + \beta^n(s_{t-1} - \gamma) + \epsilon_t^n$, $n = h,l$) Wald tests of the null that the two αs are equal lead to no rejections.

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Table 6
Bivariate partial correction model estimates: comparing pre- and post-crisis periods.
The table reports restricted SUR least squares estimates for the bivariate error correction model for yield changes:

$$\begin{cases} \Delta y_t^{high} = \alpha \Delta ffr_{t-1} + \beta^{high}(s_{t-1} - \gamma) + \varepsilon_t^{high} \\ \Delta y_t^{low} = \alpha \Delta ffr_{t-1} + \beta^{low}(s_{t-1} - \gamma) + \varepsilon_t^{low} \end{cases}$$

where st yr^{high}–yr^{low} is the yield spread between the yield yr^{high} and the yield yr^{low}, where yr^{high} is the yield of the bond that is either riskier or less liquid (or both), yt *low* is the yield of the bond that is either less risky or more liquid (or both), t^{high} and t^{low} are white noise shocks with constant variances and constant correlation. *p* Values are obtained from Newey–West HAC standard errors. *ffrt* is the effective Federal funds rate and it proxies expectations for anticipated changes in monetary policy and fixed income market conditions. The restriction consists of imposing that the coefficients and are common across equations.

Spread	Components	α (FFR loading coeff.)	β (Mean-reversion coeff.)	*p*-Value	γ (Long-run mean)	Test $\circleddash_H = \circleddash_L$ (*p*-value)	Regr. SE	Correlation of residuals	Adj. R^2
Common Pre-Crisis Period (December 2001–July 2007)									
10-year Off-On the Run Treas.	Off-the-run	0.0811	−0.1549	0.003	0.1500	0.000	0.101	0.440	0.021
	On-the-run	(0.165)	0.5984	0.000	(0.000)		0.071		0.472
3-month LIBOR-OIS	LIBOR	0.0070	−0.5924	0.000	0.1244	0.000	0.021	0.553	0.613
	OIS	(0.701)	0.3890	0.000	(0.000)		0.034		0.161
3-month Fin. CP-Treas.	Fin. CP	0.0424	**0.0648**	0.001	0.0261	**0.104**	0.037	0.593	0.005
	3-m T-Bill	(0.166)	0.0891	0.001	(0.057)		0.047		0.050
3-month Asset Bckd. CP -	ABCP	0.0367	**0.0900**	0.000	0.1103	0.014	0.040	0.573	0.015
	3-m T-Bill	(0.253)	0.1373	0.000	(0.000)		0.046		0.088
5-year CMBS-Treas.	CMBS	−0.0779	−0.0953	0.028	0.7554	0.006	0.114	0.954	0.009
	5y Treasury	(0.388)	−0.0599	0.140	(0.000)		0.106		0.001
30-year mortg.-Treas.	Fixed mtg.	−0.0537	−0.0057	0.683	1.1780	**0.206**	0.088	0.710	−0.009
	30y Treasury	(0.402)	0.0074	0.560	(0.000)		0.079		−0.001
10-year Baa–Aaa Corp.	Baa	−0.0567	−0.0267	0.235	0.9506	0.094	0.079	0.886	−0.006
	Aaa	(0.397)	**−0.0085**	0.718	(0.000)		0.083		−0.004
Post-Crisis Period (July 2009–December 2011)									
10-year Off-On the Run Treas.	Off-the-run	0.0246	0.0779	0.311	0.2856	0.000	0.149	0.330	−0.017
	On-the-run	(0.642)	0.4885	0.000	(0.000)		0.115		0.397
3-month LIBOR-OIS	LIBOR	0.0470	−0.0606	0.010	0.8091	**0.403**	0.186	0.120	0.058
	OIS	(0.325)	**−0.0455**	0.003	(0.982)		0.098		0.026
3-month Fin. CP-Treas.	Fin. CP	−0.0510	−0.1007	0.019	0.4321	0.084	0.270		0.042
	3-m T-Bill	(0.591)	−0.0242	0.383	(0.023)		0.226		−0.060
3-month Asset Bckd. CP -	ABCP	0.0911	−0.0850	0.030	0.6238	0.080	0.333	−0.025	0.059
	3-m T-Bill	(0.358	−0.0095	0.674	(0.019)		0.228		−0.079
5-year CMBS-Treas.	CMBS	0.1023	**−0.0370**	0.161	6.3525	**0.135**	1.083	−0.102	0.005
	5y Treasury	(0.157)	**0.0036**	0.277	(0.006)		0.148		−0.006
30-year mortg.-Treas.	Fixed mtg.	0.0546	−0.1471	0.000	1.5414	**0.117**	0.151	0.485	0.129
	30y Treasury	(0.366)	**−0.0926**	0.003	(0.000)		0.127		0.067
10-year Baa–Aaa Corp.	Baa	0.0025	**−0.0248**	0.170	2.2011	**0.229**	0.157	0.773	−0.001
	Aaa	(0.970)	**−0.0112**	0.442	(0.001)		0.132		−0.015
Post-Crisis Period (July 2009–December 2011)									
10-year Off-On the Run Treas.	Off-the-run	**−1.166**	(0.901)	0.507	0.1707	0.005	0.128	0.763	0.007
	On-the-run	(0.012)	0.467	0.000	(0.000)		0.098		0.176
3-month LIBOR-OIS	LIBOR	0.007	−0.018	0.117	0.139	**0.687**	0.018	0.246	0.009
	OIS	(0.918)	−0.012	0.391	(0.044)		0.013		0.002
3-month Fin. CP-Treas.	Fin. CP	−0.034	−0.200	0.000	0.143	0.000	0.038	0.146	0.088
	3-m T-Bill	(0.708)	0.016	0.493	(0.000)		0.017		−0.020
3-month Asset Bckd. CP-	ABCP	0.056	−0.123	0.028	0.233	0.011	0.045	0.115	0.027
	3-m T-Bill	(0.550)	0.024	0.248	(0.000)		0.017		−0.013
5-year CMBS-Treas.	CMBS	0.033	−0.019	0.000	1.601	0.002	0.4460	0.115	0.008
	5y Treasury	(0.578	−0.001	0.826	(0.008)		0.115		−0.006
30-year mortg.-Treas.	Fixed mtg.	−0.530	−0.077	0.022	0.407	0.038	0.072	0.371	0.055
	30y Treasury	(0.171)	−0.010	0.750	(0.000)		0.103		−0.017
10-year Baa–Aaa Corp.	Baa	−0.070	−0.014	0.714	1.022	0.005	0.1	0.913	−0.031
	Aaa	(0.901)	−0.064	0.109	(0.000)		0.103		−0.017

We have estimated Eq. (9) for each of the 7 yield spread series (i.e., this is total of 14 underlying yield series) analyzed in this paper. Table 6 presents the results for the three sub-samples already used in Tables 2 and 4, distinguishing between the pre-crisis, crisis, and post-crisis periods. In the Table we have boldfaced "rejections" of the two hypothesis (mean reversion: either $\beta^h = 0$ and $\beta^l > 0$ or $\beta^h < 0$ and $\beta^l = 0$; the weaker condition $\beta^h \neq \beta^l$).[30] A glance at the table reveals that the financial crisis may be characterized as a period in which β^h and/or β^l often have an incorrect sign, and in which the hypothesis that $\beta^h = \beta^l$ is often not rejected. This means that yields stop reacting to departures from the attractor γ in the way they should, essentially increasing even when the past spread largely exceeds γ. On the contrary, in the non-crisis periods and especially in the aftermath of the Great Crisis, we observe that for most spreads, the conditions $\beta^h \leq 0$ and $\beta^l \geq 0$ hold, while $\beta^h \neq \beta^l$ which is compatible with mean-reversion and stationarity of the spreads. In particular, the pre-crisis period represents a sample in which most bond markets displayed orderly conditions. There is only some evidence that yields on financial CP and ABCP may have failed to move in directions opposite to $(s_{t-1} - \gamma)$. However, only in the case of the financial CP spread, the null of $\beta^h = \beta^l$ cannot be rejected, and even this occurs at a marginal *p*-value of 0.104.[31] The estimates of α are never found

[30] We are boldfacing the failure to reject the null hypothesis of $\beta^h = \beta^l$.

[31] Interestingly, $\beta^h = \beta^l$ seems also to hold for the 30-year fixed mortgage rate spread, even though in this case the signs of $\beta^{30Y\ Treas} > 0$ and of $\beta^{30Y\ mort} < 0$ are correct (yet, these are not statistically significant also on an individual basis).

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Table 7
Break Tests Applied to Bivariate Partial Correction Models.
The table the outcomes of two break tests applied to restricted SUR least squares estimates of the bivariate error correction model for yield changes:

$$\begin{cases} \Delta y_t^{\text{high}} = \alpha \Delta ffr_{t-1} + \beta^{\text{high}}(s_{t-1} - \gamma) + \varepsilon_t^{\text{high}} \\ \Delta y_t^{\text{low}} = \alpha \Delta ffr_{t-1} + \beta^{\text{low}}(s_{t-1} - \gamma) + \varepsilon_t^{\text{low}} \end{cases}$$

Both types of break tests are applied sequentially to each univariate series of residuals, in the sense that when the occurrence of a break is isolated (or fails to be rejected), tests for additional breaks are applied conditioning on the date (assumed or endogenously determined) of the first break. The left block of the table reports the outcomes of a Andrews–Quandt test in which break dates are left unspecified. The test is applied to sample observations after cutting the first 5% and the last 5% of the available observations. In the table, we report the Maximum and Average LR statistic. The right block of the table reports instead the outcomes of a standard Chow break test in which the break dates are exogenously specified to correspond to the first week of August 2007 and the last week of June 2009. In the case of the Chow test, both the log-likelihood and the F test statics are reported. The last column of the table shows instead the possible ranges for break dates in the conditional mean function isolated by both sets of break tests.

		Andrews–Quandt Test		Andrews–Quandt Test Conditioning on First Break		Chow Break Test						Break Ranges
		Maximum LR *F*-statistic	Average LR FStatistic	Maximum LR *F*-statistic	Average LR *F*-statistic	Dates	*F*-Statistic	Log-LR Stat	Dates	*F*-Statistic	Log-LR Stat	
10-year Off-On the Run Treas.	Off-the-run	65.095 (0.000)	12.685 (0.000)	34.331 (0.000)	11.103 (0.000)	8/3/2007	1.642	4.937	8/03/2007	1.697	10,211	Aug. 2007
8/3/2007 Off-the-run		8/3/2007	8/3/2007	10/3/2008	10/3/2008		(0.178)	(0.177)	26/06/2009	(0.118)	(0.116)	Oct. 2008
	On-the-run	22.417 (0.000)	112.18 (0.000)	36.610 (0.000)	103.147 (0.000)		**15.725**	**46598**		**11.747**	**69.205**	Aug. 2007–June 2008
		5/30/2003	4/4/2003	6/20/2008	2/22/2008		(0.000)	(0.000)		(0.000)	(0.000)	June 2009
3-month LIBOR-OIS	LIBOR	7.168 (0.120)	1.578 (0.180)	–	–	8/3/2007	**4.712**	**14.108**	8/03/2007	**4.251**	**25.33**	Aug. 2007
		–	–	–	–		(0.003)	(0.003)	26/06/2009	(0.000)	(0.000)	June 2009
	OIS	13.933 (0.005)	3.965 (0.016)	8.170 (0.074)	2.616 (0.057)		2.399	7.23		1.296	7,855	Aug. 2007–March 2008
		8/3/2007	8/3/2007	3/28/2008	3/28/2008		(0.067)	(0.065)	8/03/2007	(0.257)	(0.249)	
3-month Fin. Comm.	3m Fin. CP	7.859 (0.091)	1.673 (0.144)	4.454 (0.384)	0.864 (0.484)	8/3/2007	0.898	2.704	26/06/2009	1.399	8.425	
Paper-Treas.		10/17/2008	___	___	___		(0.441)	(0.439)		(0.211)	(0.209)	
	3m T-Bill	13.979 (0.004)	4.121 (0.013)	8.007 (0.088)	2.805 (0.064)		**2.444**	7.343		**1.875**	**17.251**	Aug. 2007–March 2008
		8/3/2007	8/3/2007	3/28/2008	3/28/2008		(0.006)	(0.061)	8/03/2007	(0.009)	(0.008)	June 2009
3-month Asset Bcked	3m ABCP	7.332 (0.099)	1.483 (0.184)	7.314 (0.108)	1.136 (0.314)	8/3/2007	**4.172**	**12,510**	26/06/2009	**2.653**	**15.951**	Aug. 2007
Comm. Paper-Treas.		10/3/2008	–	–	–		(0.006)	(0.006)		(0.015)	(0.014)	June 2009
	3m T-Bill	13.979 (0.004)	4.121 (0.013)	8.007 (0.088)	2.805 (0.064)		**2.444**	7.343		**1.875**	**17.251**	Aug. 2007
		8/3/2007	8/3/2007	3/28/2008	3/28/2008		(0.006)	(0.061)		(0.009)	(0.008)	June 2009
5-year Commercial	5Y CMBS	3.301 (0.574)	1.308 (0.240)	–	–	8/3/2007	1.849	5.568	8/03/2007	**2.996**	**17.976**	Aug. 2007
MBS-Treas.		–	–	–	–		(0.137)	(0.135)	26/06/2009	(0.007)	(0.006)	June 2009
	5y Treasury	7.764 (0.087)	2.208 (0.093 7	7.489 (0.104)	1.399 (0.208)		**2.757**	**8.291**		0.758	4.602	Aug. 2007
		8/3/2007	8/3/2007	3/28/2008	3/28/2008		(0.041)	(0.040)		(0.603)	(0.596)	
30-year fixed rate	30Y Mortg.	5.164 (0.260)	1.905 (0.127)	–	–	8/3/2007	2.003	6.021	8/03/2007	**5.541**	**33.069**	Aug. 2007
mortgage-Treas.									26/06/2009			June 2009
		–	–	–	–		(0.111)	(0.111)		(0.000)	(0.000)	
	30y Treas.	12.606 (0.009)	0.849 (0.421)	13.794 (0.004)	4.849 (0.005)		**2.757**	**8.291**		**3.237**	**19.411**	Aug. 2007–Jan. 2009
		1/2/2009	–	2/27/2009	2/13/2009		(0.041)	(0.040)		(0.003)	(0.004)	Feb. 2009–June 2009
10-year Baa–Aaa Corporate	10Y Baa	21.974 (0.000)	1.798 (0.133)	22.740 (0.000)	10.762 (0.000)	8/3/2007	0.361	7.073	8/03/2007	**2.318**	**13.929**	Aug. 2007
		4/10/2009	–	8/21/2009	8/21/2009		(0.071)	(0.070)	26/06/2009	(0.031)	(0.031)	April 2009–Aug. 2009
	10Y Aaa	17.680 (0.000)	1.477 (0.184)	9.840 (0.031)	2.906 (0.043)		1.361	7.007		0.591	3.572	Jan. 2009
		1/2/2009	–	6/19/2009	6/19/2009		(0.071)	(0.070)		(0.737)	(0.734)	June 2009

to be statistically significant and $\hat{\alpha}$ is generally rather small, if not negative. This implies that an expectations hypothesis-like effect on the yields investigated is small at best. Finally, the estimates of γ in the pre-crisis panel of the table are generally moderate and consistent with the estimates in Table 4, which reinforces our interpretation of γ as a long-run attractor value for the spread.

During the crisis, *all* the yield spread series are affected by a rejection of either $\beta^{h}<0$ or of $\beta^{l}>0$. This is consistent with Table 4: yield spreads simply stopped being reverting and this is also shown by the fact that many yield series have stopped reacting to $(s_{t-1}-\gamma)$ at all or, worse, with a sign that is incompatible with stationarity of the spread. Interestingly, with one exception only (the off-on the run spread), this failure may actually be imputed to the fact that $\beta^{l}<0$, i.e., it is the yield on the less risky (or more liquid) bond that stops reacting to abnormally high spreads. Additionally, for 4 out of 7 spreads, $\beta^{h}=\beta^{l}$ cannot be rejected. Table 6 also shows that $\hat{\gamma}$ during the crisis period increased by an order of magnitude (between 31% and 1656%) vs. the pre-crisis period, which is consistent with our findings in Table 4. Interestingly, both the correlations of residuals in (9) and the adjusted R^2 substantially decline during the crisis, which is to be expected in a period of highly turbulent yield spreads. The lower panel of Table 6 concludes by showing a substantial return to normal market conditions in the post-crisis period, with either $\beta^{h}\leq 0$ or $\beta^{l}\geq 0$ satisfied for all spreads and $\beta^{h}\neq\beta^{l}$ failing in only one case (the 3 month LIBOR-OIS spread).

Table 7 repeats the break-point test analysis in Table 5 for the restricted bivariate SUR model in (9). Because break-point tests for multivariate models tend to be tricky, we have resorted to testing for breaks in each of the equations appearing in (9) separately, applying the same tests—Andrews–Quandt with no exogenously fixed date and Chow tests with dates suggested by the anecdotal evidence in the literature as well as by the results in Table 5—as in Section 5.3. Using Andrews–Quandt tests and *p*-values of 0.10 or less, we find evidence of two breaks in at least one yield that is part of the definition of *all* the yield spreads under investigation but one, once again the 3-month asset-backed commercial paper spread; in the case of the Off-On the run Treasury spread and the Baa–Aaa corporate default spread we actually find evidence of two breaks both the (generally) high and low yields that define the spread; in two other cases (3-month financial CP and 3-month ABCP yield spreads), both components of the spread definition show at least some evidence of a breakpoint. Strikingly, in the case of all the spreads at least one of the components is affected by a first break in correspondence to August 2007 or earlier, which confirms our previous analysis. There is considerable more uncertainty on the dating the second break, which—according to a bust-boom logic and Table 6—ought to be interpreted as the exit date from the crisis. In a few cases, markets seem to leave the crisis as early as the Spring of 2008 (e.g., this happens for the OIS, the 3-month T-bill, and to some extent the 5-year Treasury rates; the 3-month T-bill rate actually appears in the definition of two spreads).

Results are qualitatively similar in the right panel of Table 7. For almost spread series (the only stark exception is the 30-year fixed mortgage rate spread) a Chow test rejects the null of no break in correspondence to early August 2007. The breaks affect both riskier and less liquid FI instruments (e.g., the LIBOR and 3-month ABCP) as well as the yields on Treasury bills and notes. This is not as counter-intuitive as this may appear because these breaks in the rate process for less risky Treasuries (and/or more liquid, like in the case of the OIS rate) are consistent with a liquidity crisis in which there is a massive flight to the safety of Treasuries. In the case of longer-term bonds, it is also possible that the LSAP interventions may have weighted on the breaks we have isolated. There is instead no evidence of a mid-2007 break in fixed rate mortgage, Baa, and Aaa rates. In the case of the 10-year Off-On the run spread, there is also evidence of an additional break in the Spring of 2008, which is consistent with Table 6.

5. Conclusions

This paper has employed simple breakpoint tests applied to univariate and bivariate partial correction models of individual, weekly yield spread series to ask how does a financial crisis affect bond risk (both liquidity and credit) premia and whether it is possible to "date" a financial crisis. Two insights are important and would probably deserve further investigation. First, although most commentaries during the crisis have insisted upon drawing our attention to *level* of yield spreads as indicators of market disruption, our empirical results show that the crisis has had the power to affect the *persistence structure*—more precisely, the typical average duration of shocks—of the dynamic process followed by the spreads. In a policy perspective, this means that not only do (bond) risk premia increase during a financial crisis—as everybody would expect—but also that any shock that may cause these premia to depart from their normal levels, is destined to produce long-lived effects. Second, we have uncovered evidence that while one market—the prime (agency-sponsored) fixed-rate residential mortgage market—seems to have escaped the crisis altogether, in a few other FI segments the crisis is not only over, but the dynamics of spreads seems to have already completed reverted to the patterns that have characterized the pre-crisis periods. The finding that some markets may have "under"-shot to the bubble-like conditions of the pre-crisis period may instead provide reason for concern. Yet, other FI markets were found (e.g., the LIBOR, OIS or the 3-month financial commercial paper markets) to affected by renewed tensions starting from the late Summer of 2011. Although our analysis has lacked of sufficient to draw firm conclusions on whether these signals may mark a new (potentially, third) stage of the GFC and not a new, possibly European sovereign-driven crisis, our econometric methodology has thus proven flexible enough to provide a real-time tool to monitor the existence of tensions in US bond markets.

Of course, this paper could be extended also in ways that do not involve its methods but instead require additional data. First, a number of papers (e.g., Gilchrist et al., 2009) have not used standard index data to build yield spread series but instead carefully constructed credit and liquidity spread indices for different sectors of economic activity, rating categories, and alternative maturities directly from raw data sets that include individual (corporate) bond prices. It would be important to pursue our question of whether and how a crisis affects the persistence structure of spread dynamics across different risk (rating) classes and over the entire term structure of spreads (see e.g., Ahn, Dieckmann, & Perez, 2009). Second, even if one limits herself to spreads commonly reported in the literature, a number of additional spreads could have been examined in addition to the seven series used in this paper, such as swap rate spreads (vs. Treasuries) or medium-term REFCO liquidity (vs. Treasuries) spreads as in Longstaff et al. (2005), short-term spreads (vs. Treasuries) for adjustable mortgage rates which are popular in the real estate literature (see e.g., Lehnert, Passmore, & Sherlund, 2008), and corporate default spreads that also involve non-quality grade bonds (e.g., a Aa-Bbb junk spread), as in Joutz and Maxwell (2002).

Appendix A. Half-Life of a Shock in Partial Error-Correction Model

Exploiting the invertibility of a covariance-stationary process, the spread process (1)-(2) can be written as

$$s_t=\left(1-\phi_1 L-\phi_2 L^2\right)^{-1}(\phi_0+\epsilon_t)=\sum_{j=0}^{\infty}\psi_j L^j\phi_0+\sum_{j=0}^{\infty}\psi_j L^j\epsilon_t=\frac{\phi_0}{1-\phi_1-\phi_2}+\sum_{j=0}^{\infty}\psi_j L^j\epsilon_t,$$

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where

$$\psi_j = \frac{\lambda_1}{(\lambda_1 - \lambda_2)}\lambda_1^j + \frac{\lambda_2}{(\lambda_2 - \lambda_1)}\lambda_2^j,$$

and the λ_is ($i=1,2$) are the eigenvalues of the characteristic matrix

$$\mathbf{F} \equiv \begin{bmatrix} 1+\alpha+\beta & -\alpha \\ 1 & 0 \end{bmatrix}.$$

This can be seen by defining $\xi_t \equiv [s_t\, s_{t-1}]'$ and re-writing (2) as

$$\xi_t = \begin{bmatrix} \phi_0 \\ 0 \end{bmatrix} + \begin{bmatrix} \phi_1 & \phi_2 \\ 1 & 0 \end{bmatrix}\xi_{t-1} + [\epsilon_t 0] = \mu^* + \mathbf{F}\xi_{t-1} + \mathbf{v}_t,$$

where $\mu^* \equiv [\phi_0 0]'$ and $\mathbf{v}_t \equiv \left[\epsilon_t 0\right]'$. At this point, standard results (see Hamilton, 1994) give

$$\begin{aligned}\xi_{t+\tau} &= \sum_{j=0}^{\tau-1}\mathbf{F}^{\tau-j}\mu^* + \mathbf{F}^{\tau}\xi_t + \sum_{j=0}^{\tau-1}\mathbf{F}^{\tau-j}\mathbf{v}_{t+j} \\ &= \sum_{j=0}^{\tau-1}\mathbf{F}^j L^j\mu^* + \mathbf{F}^{\tau}\xi_t + \sum_{j=0}^{\tau-1}\mathbf{F}^j L^j\mathbf{v}_{t+\tau},\end{aligned}$$

where $\mathbf{F}^0 = \mathbf{I}_2$ and in which the first equation is:

$$\begin{aligned}s_{t+\tau} = \mathbf{e'}_1\xi_{t+\tau} &= \sum_{j=0}^{\tau-1}\mathbf{e'}_1\mathbf{F}^j L^j\mu^* + \sum_{j=0}^{\tau-1}\mathbf{e'}_1\mathbf{F}^j L^j\mathbf{v}_{t+\tau} \\ &= \sum_{j=0}^{\tau-1}\psi_j L^j\phi_0 + \sum_{j=0}^{\tau-1}\psi_j L^j\epsilon_{t+\tau}.\end{aligned}$$

This leads to the identification:

$$\psi_j = \mathbf{e}'_1\mathbf{F}^j\mathbf{e}_1 = \mathbf{e}'\left\{\prod_{k=1}^j\begin{bmatrix} \phi_1 & \phi_2 \\ 1 & 0 \end{bmatrix}\right\}\mathbf{e}_1,$$

i.e., the [1,1] element of the j-th power of the characteristic matrix **F**. Such a value is easy to obtain in terms of the eigenvalues of **F**. Recall that the eigenvalues of a matrix **F** are those numbers λ for which:

$$det(\mathbf{F} - \lambda\mathbf{I}_2) = 0.$$

Clearly, $det(\mathbf{F} - \lambda\mathbf{I}_2) = -\lambda(\phi_1 - \lambda) - \phi_2 = \lambda^2 - \lambda\phi_1 - \phi_2 = 0$, which has solutions:

$$\lambda_{1,2} = \frac{\phi_1 \pm \sqrt{\phi_1^2 + 4\phi_2}}{2} = \begin{cases} 0.5\phi_1 + 0.5\sqrt{\phi_1^2 + 4\phi_2} \\ 0.5\phi_1 - 0.5\sqrt{\phi_1^2 + 4\phi_2} \end{cases}.$$

When ϕ_1 and ϕ_2 are distinct, there exists a nonsingular 2×2 matrix **T** such that $\mathbf{F} = \mathbf{T}\Lambda\mathbf{T}^{-1}$ where $\Lambda \equiv diag\{\lambda_1, \lambda_2\}$, so that

$$\mathbf{F}^2 = \left(\mathbf{T}\Lambda\mathbf{T}^{-1}\right)\left(\mathbf{T}\Lambda\mathbf{T}^{-1}\right) = \mathbf{T}\Lambda\Lambda\mathbf{T}^{-1} = \mathbf{T}\Lambda^2\mathbf{T}^{-1}.$$

The diagonal structure of Λ implies that Λ^2 is also a diagonal matrix whose elements are the squares of the eigenvalues of **F**. By induction, assuming that $\mathbf{F}^{j-1} = \mathbf{T}\Lambda^{j-1}\mathbf{T}^{-1}$, it is easy to show that:

$$\mathbf{F}^j = \mathbf{F}\mathbf{F}^{j-1} = \left(\mathbf{T}\Lambda\mathbf{T}^{-1}\right)\left(\mathbf{T}\Lambda^{j-1}\mathbf{T}^{-1}\right) = \mathbf{T}\Lambda\Lambda^{j-1}\mathbf{T}^{-1} = \mathbf{T}\Lambda^j\mathbf{T}^{-1}.$$

Let t_{ij} denote the row i, column j element of **T** and let t^{ij} denote the row i, column j element of $\mathbf{T}^{-1}$. Then $\mathbf{T}\Lambda^j\mathbf{T}^{-1}$ can be written out to give a [1,1] element with structure:

$$\psi_j = t_{11}\lambda_1^j t^{11} + t_{12}\lambda_2^j t^{21} = c_1\lambda_1^j + c_2\lambda_2^j,$$

where $c_q \equiv t_{1q}t^{q1}$, $q=1,2$. However, it is evident that $c_1+c_2=1$ because $\mathbf{T}\mathbf{T}^{-1} = \mathbf{I}_2$ and c_1+c_2 represents the [1,1] element of $\mathbf{T}\mathbf{T}^{-1}$. This means that ψ_j can be characterized as a weighted average of the eigenvalues λ_1^j and λ_2^j for the characteristic matrix $\mathbf{F}^j$, with weights $c_q \equiv t_{1q}t^{q1}$, $q=1,2$. Moreover, exploiting the fact that the vector $\mathbf{t}_i$ (with generic element t_{iq}, $q=1,2$) is an eigenvector of **F** associated with the eigenvalue λ_i ($i=1,2$), Hamilton (1994, pp. 22–23) shows that the coefficients c_qs can be alternatively written as:

$$c_1 = \frac{\lambda_1}{(\lambda_1 - \lambda_2)} \quad c_2 = \frac{\lambda_2}{(\lambda_2 - \lambda_1)}.$$

The effect on the present value of s_t (when the discount factor is $\rho \in (0,1)$) of a change in ϵ_t is given by:

$$\frac{\partial\sum_{\tau=0}^{\infty}\rho^{\tau}s_{t+\tau}}{\partial\epsilon_t} = \mathbf{e}'_1\frac{\partial\sum_{\tau=0}^{\infty}\rho^{\tau}\xi_{t+\tau}}{\partial\mathbf{v}'_t}\mathbf{e}_1 = \frac{1}{1-\rho\phi_1-\rho^2\phi_2}.$$

provided all the eigenvalues of **F** are less than $1/\rho$ in modulus. The cumulative effect of a one-time change in ϵ_t on $s_t, s_{t+1}, \ldots$ can be considered a special case of this result with no discounting. Setting $\rho=1$, we have that, provided the eigenvalues of **F** are all less than 1 in modulus, the cumulative effect of a one-time change in ϵ_t on the spread is given by $1/(1-\phi_1-\phi_2)$, which can alternatively be interpreted as giving the eventual long-run effect on the spread of a permanent change in ϵ_t. Therefore the half-life $\hat{\tau}$ of a σ_ϵ shock to ϵ_t (i.e., a one-standard deviation shock) is defined as

$$\hat{\tau}\partial\sum_{\mathbf{j}=0}^{\hat{\tau}}\frac{\partial\mathbf{s_{t+j}}}{\partial\epsilon_\mathbf{t}} = \sum_{\mathbf{j}=0}^{\hat{\tau}}\psi_\mathbf{j} = \sum_{\mathbf{j}=0}^{\hat{\tau}}\left[\frac{\lambda_1}{(\lambda_1-\lambda_2)}\lambda_1^\mathbf{j} + \frac{\lambda_2}{(\lambda_2-\lambda_1)}\lambda_2^\mathbf{j}\right] \geq \frac{1}{2}\frac{1}{1-\phi_1-\phi_2},$$

where $\lambda_1 \equiv 0.5(1+\alpha+\beta) + 0.5\sqrt{(1+\alpha+\beta)^2 - 4\alpha}\lambda_2 \equiv 0.5(1+\alpha+\beta) - 0.5\sqrt{(1+\alpha+\beta)^2 - 4\alpha}$, which can be shown to be equivalent to:

$$\frac{\lambda_1}{(\lambda_1-\lambda_2)}\frac{1-\lambda_1^{\hat{\tau}+1}}{1-\lambda_1} + \frac{\lambda_2}{(\lambda_2-\lambda_1)}\frac{1-\lambda_2^{\hat{\tau}+1}}{1-\lambda_2} \geq \frac{1}{2}\frac{1}{1-\phi_1-\phi_2}. \tag{10}$$

This equation can be solved numerically to find the minimum $\hat{\tau}$ such that the right-hand side exceeds half of the long-run effect of a shock.

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