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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 the crisis on the dynamic properties of a number of yield spreads. We find that the crisis has caused 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 the financial crisis has been over approximately since the Spring of 2009. The financial crisis can be conservatively dated as a August 2007 - June 2009 phenomenon, although some yield spread series seem to point out to an end of the most serious disruptions as early as in December 2008. We uncover evidence that the LSAP program implemented by the Fed in the US 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.

JEL Classification: E40, E52, C23.
Key Words: yield spreads, credit risk, liquidity risk, break-point tests, partial adjustment models.

1. Introduction

The financial crisis of 2007-2009 is viewed as the worst financial disruption since the Great Depression of 1929-33. The banking crises of the Great Depression involved runs on banks by depositors, whereas the crisis of 2007-2009 reflected panic in wholesale funding markets that left banks unable to roll over short-term debt. That has deteriorated to engulf most fixed income markets, both in the US and internationally. 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. Because a number of such interventions have directly involved the segments of the fixed

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income (FI) markets more severely affected by the crisis, 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. The dimensions of risk that are measured by yield spreads may be many, but they can be grouped as being either default or liquidity risks. We ask five questions:

- Is the financial crisis over? Here the answer ("yes") may seem obvious in hindsight, but one question remains: what is a financial crisis, at least in the perspective provided by bond prices and yields?
- Can we date the financial crisis? 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 the dating as it is usually required of business cycles?
- How can we date a financial crisis, at least on the basis of the yield spread data perspective adopted in this paper? This relates to the general question of what features and properties of yield spreads may be affected by a financial crisis.
- 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 how?
- Finally, do any of these questions admit answers that may be market-specific? For instance, are there FI markets never affected by the crisis, or for which the crisis does not seem to be over yet?

While the questions listed above are of paramount importance, it remains interesting to ask: Why a perspective on 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 focussed on FI yield spreads, and others specific to the recent financial crisis. 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., Friedman and Kuttner, 1993; Guha and Hiris, 2002; Gilchrist, Yankov and Zakrajsek, 2009) is the tendency of a number of yield spreads (e.g., between the interest rate on commercial paper and Treasury bills) to widen shortly before the onset of recessions and to narrow again before recoveries. One interpretation of these results is that these credit risk spread measure the default risk on private (relatively risky) debt. If private lenders can accurately assess increased default risks for individual firms or industries, these changes will, after aggregation, be reflected by increases in the spread. 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. In fact, the evidence of predictive power from yield spreads to real economic activity also holds out-of-sample. For instance, King, Levin, and Perli (2007) have recently shown this fact by considering models of recession risk based on 54 variables that reflect financial markets' perceptions, including spreads on 5- and 10-year corporate bonds with various credit ratings. These models also tend to perform well out of sample.[1]

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. Hence, investors and portfolio managers that employ such yield spread strategies where they swap one bond for another when the yield spread is out of line with their historical yields would benefit from an understanding the dynamic behavior of the FI risk premia.

Our key results are easy to summarize. The econometric analysis of the changing dynamic properties of a number of commonly reported yield spread series confirms the (possibly obvious) claim that the financial crisis is over. Although there is considerable uncertainty as to when exactly the crisis ceased producing its disruptive effects, there is no doubt that after the Spring of 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

[1] See also the evidence in Mody and Taylor (2003). A number of papers have stressed that results vary across different financial instruments underlying the credit spreads as well as across different time periods. See e.g., Stock and Watson (2003).

bonds—that are compatible with mean-reversion and stationarity of the spreads.

Two literatures are related to our goals in this paper. One recent literature 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 (2010) 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 when the Fed makes a sterilized TAF loan to a depository institution, it directly allocates credit to that institution. Until mid-September 2008, the Fed offset the effect of its lending through the liquidity programs on the total supply of credit through open market operations thus reducing their ability to affect financial markets.

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. His analysis demonstrates that econometric models are capable of explaining up to one fifth of the movement in 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).[2] Morris, Neal, and Rolph (1998) have used a standard, linear cointegration approach to investigate how monthly corporate credit spreads respond to movements in short-term riskless interest rates. Papageorgiou and Skinner (2005) 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 (in terms of the implied half-life of a shock) and the long-run spread, thus disregarding the connections between different segments of the FI market as well as the

[2] 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.

relationship between credit risk spread curves and the risk-free terms 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 *linear* 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 2007-2009 financial crisis and proposes a short list of key episodes. 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 in correspondence to the onset and the end of the financial crisis. In particular, Section 5.4 asks whether the failure of yield spreads to be mean-reverting may be decomposed across the yields that enter the definition of the spread. Section 6 concludes.

2. **The Financial Crisis Through the Yield Spread Lenses**

In this Section we review the main events of the 2007-2009 financial crisis and proceed to familiarize with the yield spread series that we investigate. Our objective is not to exhaustively list all the significant developments or discuss causes and solutions to the crisis. A number of excellent analysis are available, see e.g., Gorton (2009) and Wheelock (2010).

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 and—by wide consensus among researchers (see e.g., Wheelock, 2010)—they mark an arbitrary but useful onset date for the crisis.

Initially, the Fed's reaction was limited to stressing the availability of the discount window. This was done by extending the maximum term of primary loans to 30 days and lowering the Fed fund rate target, initially by 50 basis points. Financial strains eased in September and October 2007 but reappeared in November. 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 in part 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 discount window (see Thornton, 2009). These two initial reaction 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. This is event [4] in our list.[3]

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] in our list. 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.

[3] 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.

In spite of the beneficial effects produced on the short-end of the FI markets, the situation remained difficult in most other segments. 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.[4] 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 percent in August 2007 to a range of 0 to 0.25 percent 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 (e.g., Goldman Sachs, Morgan Stanley, and GMAC). In February 2009 the Fed announced the extension of all the existing liquidity programs, listed as events [2]-[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 healing. 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.[5]

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

[5]As of the end of the Spring 2010, the liquidity facilities in [2]-[7] have been closed. The minor exception was the TALF that has been closed on March 31, 2010 according to schedule, but that has remained open for newly issued CMBS until June 30, 2010. As of the end of March 2010, the Federal Reserve has also concluded its LSAPs of $300 billion of Treasury securities, of

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 Figure 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. Figure 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.

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 Figure 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 Figure 2. Each plot in Figure 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 other papers, e.g., Christensen, Lopez and Rudebusch (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 and Hogan, 2002). However, the most popular measures are the yield 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 (e.g., defaultable) bond. This mismatch means that the measure is biased if the underlying benchmark curve is sloped. Moreover, the benchmark security can 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 the 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

$1.25 trillion of agency MBS, and of about $175 billion of agency debt.

Spread or I-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 data used in this paper are interpolated spreads.

We now turn to a brief description of the 7 yield spread series in Figure 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 security with a 10-year maturity. 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. Figure 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 goes through the roof, repeatedly peaking at levels in excess of 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.

During the financial crisis, the LIBOR-OIS spread has been a closely watched barometer of distress in money markets. 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. 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.[6] Figure 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 alterated, 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.

[6] 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 risks as well as credit risks.

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 (bank) unsecured CP and 3-month T-bill yields.[7] Clearly, this spread mostly reflects a compensation for the short-run credit risk of the financial sector. Figure 2 tells a story that fails to boil down to the 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 Figure 2 we notice a spread that moves between 10 and 25 bp in the period 2006-July 2007. This is 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 August 2007 and September 2008, 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 and therefore represents secured borrowing. Historically, senior tranches of asset-backed securities (e.g., MBS) have frequently served as collateral to ABCP. 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 percent to as much as 50 percent 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, Figure 2 shows a dynamics that is similar to unsecured financial CP. Between 2006 and 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: 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 rollover risk in the second half of 2007 precipitated a $500 billion reduction in total ABCP outstanding. This was immediately reflected in the ABCP spread, which repeatedly spiked to exceed 150 bp between August 2007 and August 2008, generally

[7] A justification for focussing on financial CP is offered by Wu and Zhang (2005) 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.

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., 2009). 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.[8] 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. Between 2006 and 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 1,770 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.

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 Figure 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.[9] 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) have contributed to drive down the 30-year fixed rate mortgage spread.

[8] Portfolio index series also exist for lower-rated private-label MBS and CMBS, but these time series proved too short for the application of the econometric methods in this paper.

[9] 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".

In fact, this spread not only returned to its normal, pre-crisis levels (around 100 bp) by March 2009, but subsequently it kept declining until stabilizing around 50 bp in early 2010, which are incredibly low levels from a historical perspective.[10] 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.

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 issued by corporations. 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 a differential likelihood of default. Figure 2 shows the familiar 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. It is interesting to notice that financial distress took a few months to contaminate the long-term segments of the corporate bond market. The peak was in fact reached in early December 2008, at 347 bp. The aggressive reaction by the Fed lowered the spread below 300 bp during 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.

Table 1 performs a comparison between means (medians), volatility (interquartile range) of spreads for three periods: before the crisis (Dec. 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-Febr. 2010, a sample of 33 weeks). The before-crisis period is easy to characterize: spreads were on average low, often lower than average spreads over the full-sample periods (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

[10] In a long, 1985-2010, weekly time series for this spread, we have that the minimum historical observation has occurred in mid-December 2009 (at 37 bp). The other two periods of low agency residential mortgage spreads have been 1992 and 2003-2004, when the 30-year spread persistently declined below 100 bp, with troughs of 40-50 bp.

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. Although the sample becomes short, all means and volatilities decline when moving from the crisis to the post-crisis period.

Figure 3 offers a visual summary. While our comments to Tables 1-2 have stressed means as a measure of location of a series, Figure 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 percent 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. The bottom panel of Figure 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 during 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 2009.

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 and Maxwell, 2002; Manzoni, 2002),

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

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 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* 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 reparameterization of dynamic linear regression models in terms of differences and levels.[11]

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 the underlying FI product 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 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 (1) can be re-written as

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

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 (1) is the one with the strongest underlying economic intuition, in the applied econometrics literature, the representation in (2) and its equivalence to (1) is what seems to have drawn the attention to (1) itself.[12] Notice that because γ corresponds to the unconditional mean of the AR(2) representation, assumed to exist, i.e.,

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

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

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

[12]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 also gives rise to an error correction-type equation that shares many features with (1).

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.

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 the process defined by (1) 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 $\alpha = 0$ (as we have done in Figure 4), 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 always be negative.

3.2. *Testing for Instability*

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 $N \geq 2$) sub-samples generated by a fixed break data and to see whether there are significant differences in the parameters of the estimated 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

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

where $\boldsymbol{\epsilon}_j$ is the $T_j \times 1$ vector of residuals when the model is estimated on some sub-samples of T_j observations, $\tilde{\boldsymbol{\epsilon}}'\tilde{\boldsymbol{\epsilon}}$ is the restricted sum of squared residuals when no break is imposed, and $\boldsymbol{\epsilon}_j'\boldsymbol{\epsilon}_j$ 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 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 is the 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 $\hat{\boldsymbol{\theta}} \equiv [\hat{\alpha}\ \hat{\beta}\ \hat{\gamma}]'$ and let $[\pi T] = T_1$ denote the integer part of πT where $0 < \pi_L \leq \pi \leq \pi_H < 1$. Thus, the proportion π defines sub-period 1, $t = 1, ..., T_1$. Under the null hypothesis, (1)-(2) is stable for the entire sample period. Under the alternative hypothesis, the model characterized by the estimator $\hat{\boldsymbol{\theta}}_1(\pi)$ applies to observations $1, ..., [\pi T]$ and model with $\hat{\boldsymbol{\theta}}_2(\pi)$ 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[13]

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

where $\hat{\mathbf{V}}_j(\pi)$ is the (asymptotic) covariance matrix estimator for $\boldsymbol{\theta}$ 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

$$LR_T(\pi) = [L_1(\pi|\hat{\boldsymbol{\theta}}_1(\pi)) + L_2(\pi|\hat{\boldsymbol{\theta}}_2(\pi))] - [L_1(\pi|\hat{\boldsymbol{\theta}}) + L_2(\pi|\hat{\boldsymbol{\theta}})] \tag{7}$$

where $\hat{\boldsymbol{\theta}}$ 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 (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

[13] 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 $\hat{\boldsymbol{\theta}}_1(\pi) - \hat{\boldsymbol{\theta}}_2(\pi)$ 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.

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 there are the widely employed maximum (also called Sup) statistics:[14]

$$MaxW_T(\pi) = \max_{\pi_L \leq \pi \leq \pi_H} W_T(\pi) \quad \text{and} \quad MaxLR_T(\pi) = \max_{\pi_L \leq \pi \leq \pi_H} LR_T(\pi). \tag{8}$$

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 equation 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, since (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.[15] Table 2 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.[16] 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 2 shows that most (all) of the yield spread series under consideration are covariance stationary. In the case of the ADF test, the evidence is overwhelming: in 4 cases out of 7 the ADF p-value is actually lower than 5%, while in other 3 cases the ADF p-value is between 5 and 10%, which still represents evidence

[14] Two alternatives to the Sup are suggested by Andrews and Ploberger (1994) and Sowell (1996). The average statistics, $AvgW_T(\pi)$ and $AvgLR_T(\pi)$, 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 $ExpW_T(\pi) = \ln\{([(1 - \pi_L - \pi_H)T])^{-1} \int_{\pi \in [\pi_L, \pi_H]} \exp[0.5 W_T(\pi)] d\pi\}$ 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.

[15] 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.

[16] 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.

against the null of a unit root. The evidence in favor of covariance stationarity of the spreads is even stronger when the PP test is applied. Five yield spread series out of 7 lead to p-values below 5%. However, in this case one of the two cases left (the 3-month LIBOR-OIS spread) produces a p-value of 0.078, while the other (the 5-year CMBS-Treasury spread) actually shows evidence of containing a unit root (the p-value is 0.108 and does not allow to reject the null). However, even in this case of conflicting evidence from ADF vs. PP tests, 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. Because the traditional classical methodology accepts the null unless there is strong evidence against it, unit root tests usually tend to conclude that there is a unit root. The problem is exacerbated by the fact that unit root tests generally have low power. In this sense, one may be favorable to resolve the tension between the 0.080 ADF p-value and the 0.108 PP p-value for the 5-year CMBS-Treasury spread in favor of stationarity. We have also repeated these tests with reference to the common pre-crisis sample period (December 2001 - July 2007) in Table 2, finding identical results. All in all, we conclude that a (1)-(2) representation may be consistent with stationarity of the underlying monthly spread series.[17]

4.2. *Model Estimates*

We estimate model (1) for a few alternative sub-periods.[18] The results are reported in Table 3. A general result emerges: for all sample periods, the estimated model turns out to be covariance stationary, in the sense that the estimated coefficients $\hat{\boldsymbol{\theta}} \equiv [\hat{\boldsymbol{\alpha}}\ \hat{\beta}\ \hat{\gamma}]'$ map into $\hat{\boldsymbol{\phi}} \equiv [\hat{\phi}_0\ \hat{\phi}_1\ \hat{\phi}_2]'$ vectors that satisfy (4). This explains why in Table 3 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.[19]

The first panel of Table 3 shows full-sample results.[20] $\hat{\beta}$ is negative for all seven spreads and only in one case (for the 5-year CMBS spread, which is in some sense consistent with Table 2) $\hat{\beta} < 0$ fails to be statistically significant (but the p-value is 0.06). In fact, some yield spreads display a considerable speed of reversion to the mean, in particular the short-term (Off-On the run Treasury, LIBOR-OIS, Financial CP-Treasury, and ABCP-Treasury) spreads. These are all characterized by $\hat{\beta}$s below -0.06 and p-values of 0.00

[17] Using daily data, earlier papers (e.g., In et al. , 2003; Joutz and Maxwell, 2002; Manzoni, 2002) have concluded that a range of alternative *daily* yield spreads are I(1) series and have therefore modelled 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 FI risk premia.

[18] The model parameters are estimated by nonlinear least squares (NLS) from (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.

[19] Assuming covariance stationarity, one useful measure of persistence of a dynamic process such as (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)..

[20] 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.

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 206 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 0.12. Most of these values, for instance the roughly 100 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 contributing to 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 3 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 3 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 3). 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 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 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, the $\hat{\beta}$ estimates remain negative but fail to be statistically significant. The implication is that for 6 spreads out of 7, 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 the 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 negative 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 estimate from 76 to 16 weeks.[21] We attribute this singularity in the dynamics 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 hypothesis in Section 5.3.

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

Finally, the last panel of Table 3 reports estimation results for the post-crisis period, July 2009 - February 2010. At least in a qualitative sense, all the relevant parameters revert back to values typical of the pre-crisis period. All the $\hat{\beta}$ estimates decline and—once more, with one exception, 30-year mortgage rate spreads for which $\hat{\beta}$ declines, but fails to be statistically significant—mark a renewed strength in the mean reversion speed of spreads. In fact the $\hat{\beta}$ estimates for 6 out of 7 stabilize to levels that are below the ones estimated over the pre-crisis period.[22] For these 6 spread series, the implied half-life of a shock is indeed below the pre-crisis estimates with values between 1 (i.e., no persistence whatsoever) and 9 weeks. In fact, also the half-life of shocks to mortgage rate spreads has substantially declined from 16 to 9 weeks. This means that these declining evolution of the $\hat{\beta}$ estimates have not been reversed by parallel breaks in the $\hat{\alpha}$ estimates reported in Table 3, fourth panel. In fact, most estimates of the α coefficients fail to be statistically significant in the post-crisis sample. 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 2 cases (Off-On the run and CMBS spreads) they have remained above the pre-crisis levels. In another case (the corporate default spread), the implied long-run spread has simply stabilized back to the 2002-2007 levels (approximately between 95 and 100 bp). 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 should develop any concerns for the fact that the $\hat{\gamma}$ estimates for the LIBOR-OIS, the Financial CP, and the ABCP spreads appear to have traced back to long-run levels that are *inferior* to their already modest pre-crisis levels.[23] It should not be considered surprising that $\hat{\gamma}$ for 30-year mortgage spreads has declined between 2009 and 2010 to an exceptionally low level of 54 bp. This low level is a likely result of the LSAP programs. However, Table 3 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.[24]

4.3. *Breakpoint Tests*

Table 4 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

[22]The only exception is the LIBOR-OIS spread for which the post-crisis $\hat{\beta}$ is -0.06 vs. a pre-crisis estimate of -0.20. Oddly enough, the LIBOR-OIS spread half-life has increased from 6 to 10 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.

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

[24]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 reparing influence on the CMBS market of a comparable extent to the impact caused on the fixed rate residential mortgage market.

to be known and its assessment ("estimation") must be based on the data.[25] This a truly "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 break in the case of two series, the 3-month ABCP and the 20-year corporate Baa-Aaa spreads; they give instead evidence of two breaks in the case of other two series, the 3-month LIBOR-OIS and the 5-year CMBS spreads. There is no evidence of breaks in the remaining 3 spread series, i.e., 10-year Off-On the run, 3-month Financial CP, and 30-year fixed mortgage spreads. The two series subject to two breaks confirm the boom-bust-boom cycle that we would expect when a serious financial crisis impacts markets and resolves later on: the first break occurs in one case in October 2008 (the LIBOR-OIS spread), and in the other case (Baa-Aaa spread) in November 2008. Both breaks are detected at a very high level of statistical significance. In the former case, a second break (but only using a Maximum Wald Statistic) is detected in correspondence to late 2008 after the full deployment of the short-term liquidity facilities (e.g., CPFF and MMIFF); in the latter case, a second break is detected in correspondence of the Spring of 2009. Based on the evidence in Table 3, it is sensible in both cases to interpret the first breakpoint date as the date in which the corresponding FI markets (LIBOR and CMBS) have entered the crisis, and the second date as the date in which they have emerged from the crisis. More puzzling is the fact that two markets seemed to have entered the crisis—in October 2008 in the case of Financial CP and in July 2007 in the case of the corporate bond market, with both breakpoint dates detected with very low and reliable p-values—but not to have left it. This is shown by the fact that a second Andrews-Quandt break test that conditions on a first break returns no evidence of further breakpoints for these two series. Moreover, the fact that in Table 4 three yield series seem to have not been subject to any breaks does not square well with the evidence in Table 3. Of course, this may 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 4 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. Here the results fully conform with the evidence in Table 3: for 6 out of 7 yield spreads series, there is evidence of a break in early August 2007. The corresponding p-values are below 1% for 4 series, while for other 2 series they are between 1 and 5%. Not surprisingly, the only yield spread not affected by a break in the Summer of 2007 is the 30-year mortgage spread. There are

[25] 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.

no economically important differences between the F and Log-LR versions of the Chow test. For these 6 spreads, conditioning on a first break occurring in the first week of August 2007, we further proceed to test for another breakpoint at the end of June 2009.[26] For all of the 6 spread series we find evidence of a second break, which we interpret as evidence of the end of the financial crisis.

The last column of Table 4 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 only one spread, there is no evidence of breaks in (1). This series is the 30-year fixed mortgage rate spread. Assuming—in the light of the evidence for the 5-year CMBS spread—that in the absence of the LSAP programs all of the US residential mortgage market would have been significantly affected by the 2007-2008 financial crisis, the results in Tables 3-4 show the substantial success of these policy measures. The remaining 6 series present the typical boom-bust-boom that we would expect of financial markets when the available data span a complete financial crisis. They all enter a crisis period—characterized by high persistence of shocks and high long-run mean spreads (see Table 3)—between August 2007 and November 2008, which fits our summary of the main events in Section 2.1. They all leave the crisis between December 2008 (for the 3-month LIBOR-OIS spread) and June 2009, which is what we would have conjectured on the basis of the commentary of Section 2.1.

4.4. *How Did the Crisis Affect Bond Markets?*

A finding that any of the coefficients in $\boldsymbol{\theta} \equiv [\boldsymbol{\alpha}\ \beta\ \gamma]'$ in (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 $\Delta s_t \equiv \Delta y_t^{high} - \Delta y_t^{low}$—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

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

where ffr_t is the effective federal funds rate and $\boldsymbol{\epsilon}_t \equiv [\epsilon_t^h\ \epsilon_t^l]' \sim$ IID $N(\mathbf{0}, \boldsymbol{\Sigma})$. Notice that the off-diagonal element of $\boldsymbol{\Sigma}$, $Cov[\epsilon_t^h, \epsilon_t^l]$, represents the covariance of shocks affecting the two yield series. (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. (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

[26] In the case of the fixed rate mortgage spread we anyway proceed to test for the presence of a first break in 2008 or 2009 and find no evidence of a breakpoint.

ffr is used to capture expected movements in interest rates that propagate from monetary policy actions to the entire yield curve.[27] Second, the correction term has in this case 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, (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 (9).[28] However, (9) does capture the logic that (risky) yields adjust when the short-end of the riskfree 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 $\beta^h \leq 0$ and $\beta^l \geq 0$, i.e. when the spread exceeds some historical 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 inching up. Of course, $\beta^h < 0$ and $\beta^l = 0$ represents a realistic possibility.

2. Unless $\beta^h \neq \beta^l$ no adjustment in the yield spread is possible, although this is a rather weak necessary condition (one may formulate a sharper condition that $\beta^h < \beta^l$, although signs matter as much as magnitudes in this case).

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.

We have estimated (9) for each of the 7 yield spread series (i.e., this is total of 14 underlying yield series) analyzed in this paper. Table 5 presents the results for the three sub-samples already used in Tables 1 and 3, 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 $\beta^h \neq \beta^l$).[29] 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

[27]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 Δy_{t-1}^{high}, etc.) or on the most recent change of the spread, but did not find any substantial differences. In particular, we have noticed that in the majority of the cases, 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 were not statistically different would lead to no rejections.

[28]In a bivariate model one would expect a vector to possibly represent the unconditional means of the series, while γ is a scalar. More importantly, notice that γ is compared to the past yield *spread* and not to the yield in the error correction term.

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

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.[30] The estimates of α are never found 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 3, 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 $\beta^h \leq 0$ or of $\beta^l \geq 0$. This is consistent with Table 3: 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, 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 5 also shows that $\hat{\gamma}$ during the crisis period tended to increase to an order of magnitude (between 31 and 1656 percent) higher vs. the pre-crisis period, which is consistent with our findings in Table 3. 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 5 concludes by showing a complete 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 30-year fixed mortgage spread).

Table 6 repeats the break-point test analysis in Table 4 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 4—as in Section 5.3.[31] Using Andrews-Quandt tests, we find evidence of two breaks in at least one yield that is part of the definition of *all* the yield spreads under investigation; in the case of the Baa-Aaa corporate default spread we actually find evidence of two breaks both in Baa and in Aaa yields; 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 5 spreads at least one of the components is affected by a first break in correspondence to August 2007, which confirms our previous analysis. However, in the case of the 30-year mortgage spread and for the Baa-Aaa spread the evidence is in favor of a break in early 2009. This latter result may be related to the tendency of corporate

[30]Interestingly, $\beta^h = \beta^l$ seems also to hold for the CMBS spread, even though the hypothesis of $\beta^{30Y\ Treas} > 0$ cannot be rejected while β^{CMBS} is not statistically different from zero.

[31]In this case we report both the Maximum LR and the Average LR statistics, as a robustness check.

bond markets during the crisis to reflect with a long lag the mounting stress in shorter-term FI markets. There is considerable more uncertainty on the dating the second break, which—according to a bust-boom logic and Table 5—ought to be interpreted as the exit date from the crisis. In four cases, markets seem to leave the crisis as early as the Spring of 2008 (this happens for the OIS, the 3-month T-bill, and the 5-year Treasury rates; the 3-month T-bill rate actually appears in the definition of two spreads). In the case of the LIBOR rate, the second break is estimated to have occurred in October 2008, for the 30-year Treasury rate in February 2009, and for Baa and Aaa corporate yields between June and August 2009.

Results are qualitatively similar in the right panel of Table 6. In four cases (Off-On the run, LIBOR-OIS, financial CP, and ABCP spreads) a Chow test rejects the null of no break in correspondence to early August 2007; in a fifth case (CMBS spread) there is also evidence of a break, although the corresponding p-value is between 0.05 and 0.10. The breaks affect the yields on Treasury bills and notes and the OIS rate. 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 2009, which is consistent with Table 5.

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. We have argued that our finding of no breakpoints in the process of residential mortgage spreads may be tentatively ascribed to the success of the portions of the LSAP programs that led the Fed and the Treasury to intervene in the prime, agency-backed MBS markets and to provide unlimited financial backstop to losses incurred by the government-sponsored agencies involved in the housing market. 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.

Of course, this paper could be extended also in ways that do not involve its methods but instead require extensions to additional data. First, a number of papers (e.g., Gilchrist et al., 2009) have not used standard index (portfolio) data to build yield spread series but instead carefully constructed credit and liquidity yield 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 the spreads (see e.g., Ahn, Dieckmann, and 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 In, Brown, and Fang (2004) and Longstaff, Mithal, and Neis (2005), short-term spreads (vs. Treasuries) for adjustable mortgage rates which are popular in the real estate finance literature (see e.g., Lehnert, Passmore, and 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) and Mody and Taylor (2003).

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

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

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 $\boldsymbol{\xi}_t \equiv [s_t \; s_{t-1}]'$ and re-writing (2) as

$$\boldsymbol{\xi}_t = \begin{bmatrix} \phi_0 \\ 0 \end{bmatrix} + \begin{bmatrix} \phi_1 & \phi_2 \\ 1 & 0 \end{bmatrix}\boldsymbol{\xi}_{t-1} + \begin{bmatrix} \epsilon_t \\ 0 \end{bmatrix} = \boldsymbol{\mu}^* + \mathbf{F}\boldsymbol{\xi}_{t-1} + \mathbf{v}_t,$$

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

$$\boldsymbol{\xi}_{t+\tau} = \sum_{j=0}^{\tau-1}\mathbf{F}^{\tau-j}\boldsymbol{\mu}^* + \mathbf{F}^{\tau}\boldsymbol{\xi}_t + \sum_{j=0}^{\tau-1}\mathbf{F}^{\tau-j}\mathbf{v}_{t+j} = \sum_{j=0}^{\tau-1}\mathbf{F}^{j}L^j\boldsymbol{\mu}^* + \mathbf{F}^{\tau}\boldsymbol{\xi}_t + \sum_{j=0}^{\tau-1}\mathbf{F}^{j}L^j\mathbf{v}_{t+\tau},$$

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

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

This leads to the identification:

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

i.e., the [1,1] element of the j-th power of the characteristic matrix $\mathbf{F}$. Such a value is easy to obtain in terms of the eigenvalues of $\mathbf{F}$. Recall that the eigenvalues of a matrix $\mathbf{F}$ 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 $\mathbf{T}$ such that $\mathbf{F} = \mathbf{T\Lambda T}^{-1}$ where $\mathbf{\Lambda} \equiv diag\{\lambda_1, \lambda_2\}$, so that

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

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

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

Let t_{ij} denote the row i, column j element of $\mathbf{T}$ 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 $\mathbf{TT}^{-1} = \mathbf{I}_2$ and $c_1 + c_2$ represents the [1,1] element of $\mathbf{TT}^{-1}$. 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 $\mathbf{F}$ 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:

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

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:

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

provided all the eigenvalues of $\mathbf{F}$ are less than $1/\rho$ in modulus. The cumulative effect of a one-time change in ϵ_t on s_t, s_{t+1}, ... can be considered a special case of this result with no discounting. Setting $\rho = 1$, we have that, provided the eigenvalues of $\mathbf{F}$ 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

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

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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Table 1

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	0.129	0.114	0.136	0.892	1.444
		(0.000)	(0.000)			(0.026)	(0.185)
3-month LIBOR-OIS	296	0.106	0.098	0.045	0.057	0.983	2.109
		(0.000)	(0.000)			(0.031)	(0.121)
3-month Fin. Comm. Paper-Treasury	296	0.157	0.130	0.099	0.120	1.285	1.838
		(0.000)	(0.000)			(0.002)	(0.252)
3-month Asset-Backed Comm. Paper-Treasury	296	0.190	0.170	0.098	0.120	1.265	1.728
		(0.000)	(0.000)			(0.002)	(0.130)
5-year Comm. MBS Rate-Treasury	296	0.764	0.720	0.153	0.170	1.270	1.432
		(0.000)	(0.000)			(0.000)	(0.077)
30-year Fixed Rate Mortgage Rate-	296	1.115	1.125	0.360	0.673	-0.009	-1.442
		(0.000)	(0.000)			(0.932)	(0.000)
20-year Aaa-Baa Moody's Default	296	0.977	0.930	0.207	0.273	0.504	-0.616
		(0.000)	(0.000)			(0.001)	(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	0.230	0.192	0.236	0.791	-0.031
		(0.000)	(0.000)			(0.031)	(0.909)
3-month LIBOR-OIS	100	0.944	0.776	0.589	0.310	2.414	6.318
		(0.000)	(0.000)			(0.038)	(0.089)
3-month Fin. Comm. Paper-Treasury	100	1.020	0.985	0.625	0.915	0.994	1.397
		(0.000)	(0.000)			(0.107)	(0.250)
3-month Asset-Backed Comm. Paper-Treasury	100	1.255	1.075	0.846	1.040	1.344	2.169
		(0.000)	(0.000)			(0.060)	(0.192)
5-year Comm. MBS Rate-Treasury	100	5.652	3.715	4.459	6.820	0.879	-0.635
		(0.000)	(0.000)			(0.003)	(0.136)
30-year Fixed Rate Mortgage Rate-	100	1.601	1.600	0.401	0.405	-0.328	0.754
		(0.000)	(0.000)			(0.461)	(0.250)
20-year Aaa-Baa Moody's Default	100	1.842	1.470	0.854	1.568	0.543	-1.242
		(0.000)	(0.000)			(0.006)	(0.000)

	Weekly Obs.	Mean	Median	St. Dev.	Interq. Range	Skewness	Excess Kurtosis
	Post-Crisis Period (July 2009 - February 2010)						
10-year Off-On the Run Treasuries	33	0.241	0.221	0.087	0.137	0.144	-0.901
		(0.000)	(0.000)			(0.729)	(0.039)
3-month LIBOR-OIS	33	0.171	0.131	0.092	0.118	1.156	-0.106
		(0.000)	(0.000)			(0.0063)	(0.826)
3-month Fin. Comm. Paper-Treasury	33	0.132	0.130	0.026	0.030	1.288	2.118
		(0.000)	(0.000)			(0.306)	(0.323)
3-month Asset-Backed Comm. Paper-Treasury	33	0.181	0.180	0.029	0.040	0.110	-0.283
		(0.000)	(0.000)			(0.852)	(0.377)
5-year Comm. MBS Rate-Treasury	33	4.507	4.370	1.057	1.050	1.604	2.363
		(0.000)	(0.000)			(0.151)	(0.243)
30-year Fixed Rate Mortgage Rate-	33	0.674	0.710	0.199	0.360	0.032	-1.287
		(0.000)	(0.000)			(0.926)	(0.001)
20-year Aaa-Baa Moody's Default	33	1.243	1.160	0.246	0.200	1.160	0.221
		(0.000)	(0.000)			(0.006)	(0.000)

Table 2

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 – Feb. 2010. 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. t-statistic	P-value	Band-width
10-year Off-On the Run Treasuries	Jan. 1985	1310	Level	-7.379	**0.000**	4	-29.389	**0.000**	25
			First-diff.	-20.539	**0.000**	6	-199.15	**0.000**	121
3-month LIBOR-OIS	Dec. 2001	427	Level	-2.619	**0.090**	4	-2.680	**0.078**	11
			First-diff.	-11.862	**0.000**	3	-11.311	**0.000**	25
3-month Fin. Comm. Paper-Treasury	Jan. 1985	1310	Level	-6.812	**0.000**	0	-6.328	**0.000**	13
			First-diff.	-38.087	**0.000**	0	-47.123	**0.000**	39
3-month Asset-Backed Comm. Paper-Treasury	Jan. 2001	475	Level	-3.596	**0.006**	4	-3.619	**0.006**	10
			First-diff.	-12.478	**0.000**	3	-21.163	**0.000**	18
5-year Comm. MBS Rate-Treasury	July 1996	710	Level	-2.670	**0.080**	9	-2.533	0.108	15
			First-diff.	-7.224	**0.000**	8	-34.775	**0.000**	14
30-year Fixed Rate Mortgage Rate-Treasury	Jan. 1985	1310	Level	-2.936	**0.042**	0	-2.917	**0.044**	3
			First-diff.	-36.739	**0.000**	0	-36.743	**0.000**	3
20-year Aaa-Baa Moody's Default Spread	Jan. 1985	1310	Level	-2.854	**0.051**	1	-3.097	**0.027**	20
			First-diff.	-24.566	**0.000**	0	-25.439	**0.000**	17

Table 3

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 s_t is the yield spread and ε_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 ε_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 γ). 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 - February 2010)								
10-year Off-On the Run Treasuries	-0.2827	0.000	-0.3176	0.000	0.1700	0.000	0.113	0.283	2
3-month LIBOR-OIS	0.4696	0.000	-0.0347	0.000	0.3094	0.012	0.087	0.227	11
3-month Fin. Comm. Paper-Treasury	-0.0197	0.477	-0.0662	0.000	0.2929	0.000	0.123	0.033	11
3-month Asset-Backed Comm. Paper-Treas.	0.0956	0.038	-0.0591	0.000	0.4066	0.007	0.191	0.032	11
5-year Comm. MBS Rate-Treasury	-0.2929	0.000	-0.0132	0.061	2.0594	0.118	0.456	0.092	68
30-year Fixed Rate Mortgage Rate-Treas.	-0.0081	0.770	-0.0141	0.004	1.2646	0.000	0.085	0.005	50
20-year Aaa-Baa Moody's Default Spread	0.3718	0.000	-0.0078	0.004	0.9827	0.000	0.041	0.139	56
	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 - February 2010)								
10-year Off-On the Run Treasuries	-0.1318	0.474	-0.5027	0.018	0.2361	0.000	0.081	0.249	1 (-3)
3-month LIBOR-OIS	0.1129	0.485	-0.0620	0.052	0.0642	0.360	0.016	0.123	10 (+4)
3-month Fin. Comm. Paper-Treasury	-0.3295	0.013	-0.3269	0.006	0.1223	0.000	0.018	0.311	3 (-4)
3-month Asset-Backed Comm. Paper-Treas.	-0.4162	0.002	-0.5967	0.000	0.1749	0.000	0.025	0.519	1 (-5)
5-year Comm. MBS Rate-Treasury	0.1600	0.336	-0.1291	0.019	3.7904	0.000	0.339	0.15	5 (-24)
30-year Fixed Rate Mortgage Rate-Treas.	-0.0531	0.774	**-0.0822**	0.322	0.5391	0.035	0.088	-0.022	9 (-7)
20-year Aaa-Baa Moody's Default Spread	0.3248	0.054	-0.0542	0.083	0.9641	0.000	0.056	0.186	9 (-18)

Table 4

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 s_t is the yield spread and ε_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 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 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	Maximum Wald F-Statistic	Maximum LR F-Statistic	Maximum Wald F-Statistic	Dates	F Statistic	Log-LR Stat	Dates	F Statistic	Log-LR Stat	
10-year Off-On the Run	8.256 (0.428)	7.732 (0.494)	___	___	8/3/2007	5.140	15.399	8/3/2007	2.801	16.811	Aug. 2007
Treasuries	___	___	___	___		(0.002)	(0.002)	6/26/2009	(0.010)	(0.010)	June 2009
3-month LIBOR-OIS	23.006 (0.001)	25.035 (0.001)	12.263 (0.116)	19.265 (0.007)	8/3/2007	9.628	28.331	8/3/2007	7.125	41.571	Aug. 2007 - Oct. 2008
	10/24/2008	10/24/2008	___	12/19/2008		(0.000)	(0.000)	6/26/2009	(0.000)	(0.000)	Dec. 2008 - June 2009
3-month Fin. Comm.	4.874 (0.879)	7.105 (0.578)	___	___	8/3/2007	4.874	14.607	8/3/2007	5.660	33.754	Aug. 2007
Paper-Treasury	___	___	___	___		(0.002)	(0.002)	6/26/2009	(0.000)	(0.000)	June 2009
3-month Asset-Backed	46.299 (0.000)	46.269 (0.000)	7.122 (0.576)	9.944 (0.257)	8/3/2007	3.512	10.553	8/3/2007	3.846	22.959	Aug. 2007 - Oct. 2008
Comm. Paper-Treasury	10/3/2008	10/3/2008	___	___		(0.015)	(0.014)	6/26/2009	(0.001)	(0.001)	June 2009
5-year Comm. MBS Rate-	30.462 (0.000)	29.493 (0.000)	7.576 (0.051)	21.003 (0.003)	8/3/2007	2.643	7.952	8/3/2007	4.252	25.377	Aug. 2007 - Nov. 2008
Treasury	11/28/2008	11/28/2008	3/27/2009	3/27/2009		(0.048)	(0.047)	6/26/2009	(0.000)	(0.000)	March 2009 - June 2009
30-year Fixed Rate	11.223 (0.117)	8.165 (0.439)	___	___	8/3/2007	1.259	3.788	8/3/2007	1.293	7.786	___
Mortgage Rate-Treasury	___	___	___	___		(0.287)	(0.285)	6/26/2009	(0.257)	(0.254)	___
20-year Aaa-Baa Moody's	37.361 (0.000)	37.944 (0.000)	5.347 (0.822)	4.195 (0.945)	8/3/2007	37.029	107.09	8/3/2007	20.293	117.19	July 2007 - Aug. 2007
Default Spread	7/27/2007	7/27/2007	___	___		(0.000)	(0.000)	6/26/2009	(0.000)	(0.000)	June 2009

Table 5

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 $s_t \equiv y_r^{high} - y_r^{low}$ is the yield spread between the yield y_r^{high} and the yield y_r^{low}, where y_r^{high} is the yield of the bond that is either riskier or less liquid (or both), y_t^{low} is the yield of the bond that is either less risky or more liquid (or both), ε_t^{high} and ε_t^{low} are white noise shocks with constant variances and constant correlation. p-values are obtained from Newey-West HAC standard errors. ffr_t 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 α and γ are common across equations.

Spread	Components	α (FFR loading coeff.)	β (Mean-reversion coeff.)	p-value	γ (Long-run mean)	Test $\beta_H = \beta_L$ (p-value)	Regr. SE	Correlation of residuals	Adj. R^2
		Common Pre-Crisis Period (December 2001 - July 2007)							
10-year Off-On	Off-the-run	0.0811	-0.1549	0.003	0.1500	0.000	0.101	0.440	0.021
the Run Treas.	On-the-run	(0.165)	0.5984	0.000	(0.000)		0.071		0.472
3-month	LIBOR	0.0070	-0.5924	0.000	0.1244	0.000	0.021	0.553	0.613
LIBOR-OIS	OIS	(0.701)	0.3890	0.000	(0.000)		0.034		0.161
3-month Fin.	Fin. CP	0.0424	**0.0648**	0.001	0.0261	**0.104**	0.037	0.593	0.005
CP - Treas.	3-m T-Bill	(0.166)	0.0891	0.001	(0.057)		0.047		0.050
3-month Asset	ABCP	0.0367	**0.0900**	0.000	0.1103	0.014	0.040	0.573	0.015
Bckd. CP - Treas.	3-m T-Bill	(0.253)	0.1373	0.000	(0.000)		0.046		0.088
5-year CMBS -	CMBS	-0.0779	-0.0953	0.028	0.7554	0.006	0.114	0.954	0.009
Treas.	5y Treasury	(0.388)	-0.0599	0.140	(0.000)		0.106		0.001
30-year mortg. -	Fixed mtg.	-0.0537	**-0.0057**	0.683	1.1780	**0.206**	0.088	0.710	-0.009
Treas.	30y Treasury	(0.402)	0.0074	0.560	(0.000)		0.079		-0.001
10-year Baa -	Baa	-0.0567	-0.0267	0.235	0.9506	0.094	0.079	0.886	-0.006
Aaa Corp.	Aaa	(0.397)	-0.0085	0.718	(0.000)		0.083		-0.004
		Crisis Period (July 2007 - June 2009)							
10-year Off-On	Off-the-run	0.0246	**0.0779**	0.311	0.2856	0.000	0.149	0.330	-0.017
the Run Treas.	On-the-run	(0.642)	0.4885	0.000	(0.000)		0.115		0.397
3-month	LIBOR	0.0470	-0.0606	0.010	0.8091	**0.403**	0.186	0.120	0.058
LIBOR-OIS	OIS	(0.325)	**-0.0455**	0.003	(0.982)		0.098		0.026
3-month Fin.	Fin. CP	-0.0510	-0.1007	0.019	0.4321	0.084	0.270	0.101	0.042
CP - Treas.	3-m T-Bill	(0.591)	**-0.0242**	0.383	(0.023)		0.226		-0.060
3-month Asset	ABCP	0.0911	-0.0850	0.030	0.6238	0.080	0.333	-0.025	0.059
Bckd. CP - Treas.	3-m T-Bill	(0.358)	**-0.0095**	0.674	(0.019)		0.228		-0.079
5-year CMBS -	CMBS	0.1023	-0.0370	0.161	6.3525	**0.135**	1.083	-0.102	0.005
Treas.	5y Treasury	(0.157)	**0.0036**	0.277	(0.006)		0.148		-0.006
30-year mortg. -	Fixed mtg.	0.0546	-0.1471	0.000	1.5414	**0.117**	0.151	0.485	0.129
Treas.	30y Treasury	(0.366)	**-0.0926**	0.003	(0.000)		0.127		0.067
10-year Baa -	Baa	0.0025	-0.0248	0.170	2.2011	**0.229**	0.157	0.773	-0.001
Aaa Corp.	Aaa	(0.970)	**-0.0112**	0.442	(0.001)		0.132		-0.015
		Post-Crisis Period (July 2009 - February 2010)							
10-year Off-On	Off-the-run	-0.6110	0.2191	0.213	0.2471	0.000	0.096	0.547	0.058
the Run Treas.	On-the-run	(0.313)	0.7996	0.000	(0.000)		0.069		0.474
3-month	LIBOR	0.0077	-0.1115	0.000	0.0644	0.000	0.012	0.016	0.350
LIBOR-OIS	OIS	(0.914)	-0.0178	0.178	(0.044)		0.011		-0.008
3-month Fin.	Fin. CP	-0.0012	-0.3666	0.002	0.1270	0.267	0.021	0.501	-0.034
CP - Treas.	3-m T-Bill	(0.993)	0.0922	0.401	(0.000)		0.017		-0.040
3-month Asset	ABCP	0.2026	-0.5904	0.000	0.1790	0.000	0.028	0.254	0.279
Bckd. CP - Treas.	3-m T-Bill	(0.164)	0.1377	0.161	(0.000)		0.017		-0.014
5-year CMBS -	CMBS	1.2270	-0.1469	0.001	3.6916	0.002	0.2960	-0.300	0.251
Treas.	5y Treasury	(0.156)	-0.0117	0.402	(0.000)		0.1082		-0.087
30-year mortg. -	Fixed mtg.	-0.0133	-0.1563	0.012	0.6101	**0.182**	0.069	0.413	0.122
Treas.	30y Treasury	(0.982)	-0.0604	0.412	(0.000)		0.091		-0.044
10-year Baa -	Baa	0.1182	-0.1145	0.025	0.9307	0.012	0.089	0.856	0.074
Aaa Corp.	Aaa	(0.864)	-0.0413	0.221	(0.000)		0.075		-0.023

Table 6

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:

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

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 F-Statistic	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) 8/3/2007	12.685 (0.000) 8/3/2007	34.331 (0.000) 10/3/2008	1/11/1900 10/3/2008	8/3/2007	65.095 (0.000)	63.622 (0.000)	6/26/2009	7.509 (0.006)	7.499 (0.006)	Aug. 2007 Oct. 2008 - June 2009
	On-the-run	1.805 (0.893) ___	0.421 (0.750) ___	___ ___	___ ___		0.010 (0.922)	0.010 (0.922)		0.001 (0.970)	0.001 (0.970)	
3-month LIBOR - OIS	LIBOR	7.019 (0.124) ___	1.555 (0.176) ___	___ ___	___ ___	8/3/2007	0.392 (0.532)	0.393 (0.531)	6/26/2009	1.164 (0.281)	1.168 (0.280)	
	OIS	13.933 (0.005) 8/3/2007	3.965 (0.016) 8/3/2007	8.170 (0.074) 3/28/2008	2.616 (0.057) 3/28/2008		13.933 (0.000)	13.775 (0.000)		1.463 (0.229)	1.477 (0.224)	Aug. 2007 March 2008
3-month Fin. Comm. Paper - Treas.	3m Fin. CP	7.364 (0.096) 10/17/2008	1.593 (0.168) ___	4.614 (0.342) ___	0.890 (0.401) ___	8/3/2007	0.161 (0.688)	0.162 (0.687)	6/26/2009	0.539 (0.463)	0.541 (0.462)	Oct. 2008?
	3m T-Bill	13.933 (0.005) 8/3/2007	3.965 (0.016) 8/3/2007	8.170 (0.082) 3/28/2008	2.616 (0.057) 3/28/2008		13.933 (0.000)	13.775 (0.000)		0.064 (0.800)	0.064 (0.800)	Aug. 2007 March 2008
3-month Asset Bcked Comm. Paper - Treas.	3m ABCP	7.373 (0.096) 10/3/2008	1.425 (0.205) ___	7.228 (0.113) ___	1.198 (0.270) ___	8/3/2007	0.312 (0.577)	0.313 (0.580)	6/26/2009	0.505 (0.478)	0.507 (0.476)	Oct. 2008?
	3m T-Bill	7.428 (0.093) 8/3/2007	2.069 (0.099) 8/3/2007	7.524 (0.099) 3/28/2008	1.363 (0.221) 3/28/2008		6.889 (0.009)	6.866 (0.009)		0.460 (0.499)	0.467 (0.500)	Aug. 2007 March 2008
5-year Commercial MBS - Treas.	5Y CMBS	3.328 (0.558) ___	1.360 (0.222) ___	___ ___	___ ___	8/3/2007	2.520 (0.091)	2.624 (0.100)	6/26/2009	0.976 (0.325)	0.987 (0.320)	Aug. 2007?
	5y Treasury	7.428 (0.093) 8/3/2007	2.069 (0.099) 8/3/2007	7.524 (0.099) 3/28/2008	1.363 (0.221) 3/28/2008		1.341 (0.247)	1.346 (0.246)		0.100 (0.753)	0.100 (0.752)	Aug. 2007 March 2008
30-year fixed rate mortgage - Treas.	30Y Mortg.	4.883 (0.307) ___	1.745 (0.142) ___	___ ___	___ ___	8/3/2007	0.084 (0.772)	0.084 (0.772)	6/26/2009	2.377 (0.124)	2.382 (0.123)	
	30y Treas.	12.245 (0.011) 1/2/2009	0.762 (0.475) ___	13.548 (0.006) 2/13/2009	4.556 (0.009) 2/13/2009		0.298 (0.586)	0.300 (0.585)		0.001 (0.973)	0.001 (0.973)	Jan. 2009? Feb. 2009
10-year Baa-Aaa Corporate	10Y Baa	21.913 (0.000) 4/10/2009	1.769 (0.138) ___	22.512 (0.000) 8/21/2009	10.483 (0.000) 8/21/2009	8/3/2007	0.681 (0.410)	0.683 (0.409)	6/26/2009	6.292 (0.013)	6.276 (0.012)	Apr. - June 2009 Aug. 2009
	10Y Aaa	17.161 (0.001) 1/2/2009	1.433 (0.203) ___	9.547 (0.040) 6/19/2009	2.839 (0.046) 6/19/2009		0.122 (0.727)	0.123 (0.726)		1.112 (0.292)	1.116 (0.291)	Jan. 2009? June 2009

Figure 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.

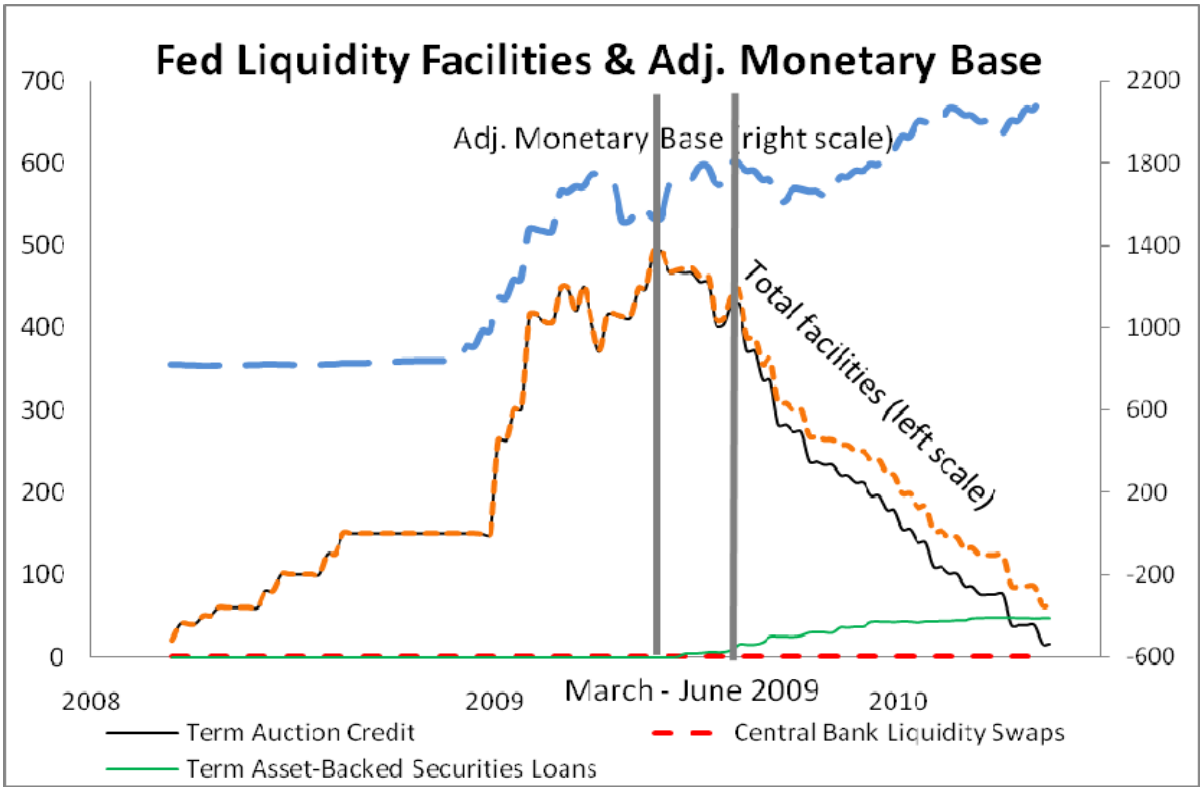
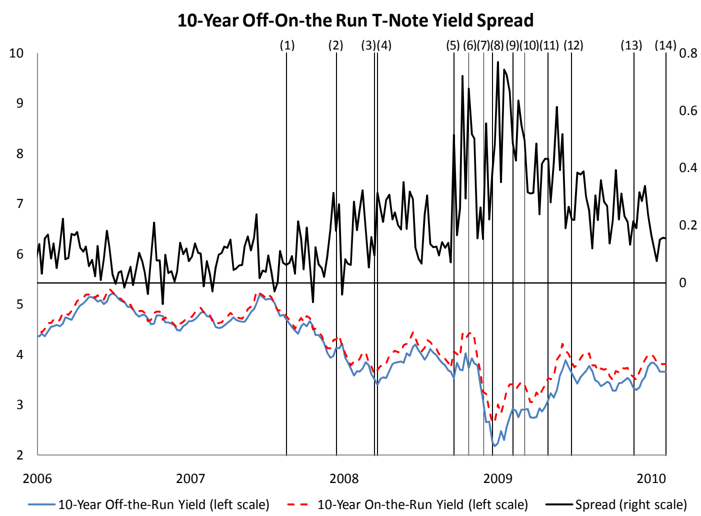
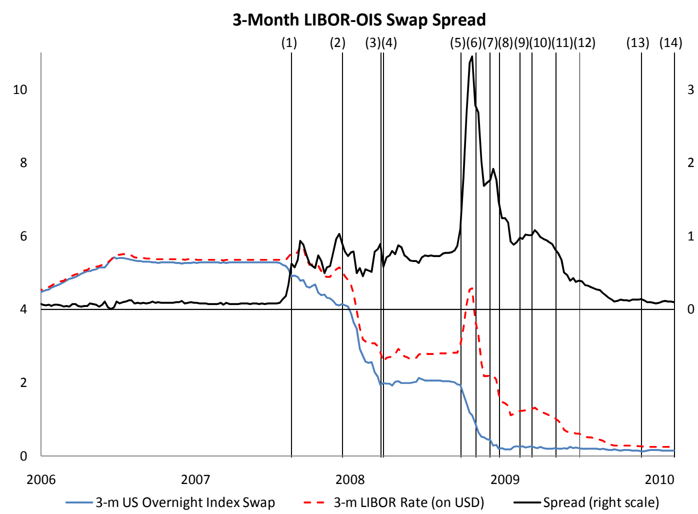
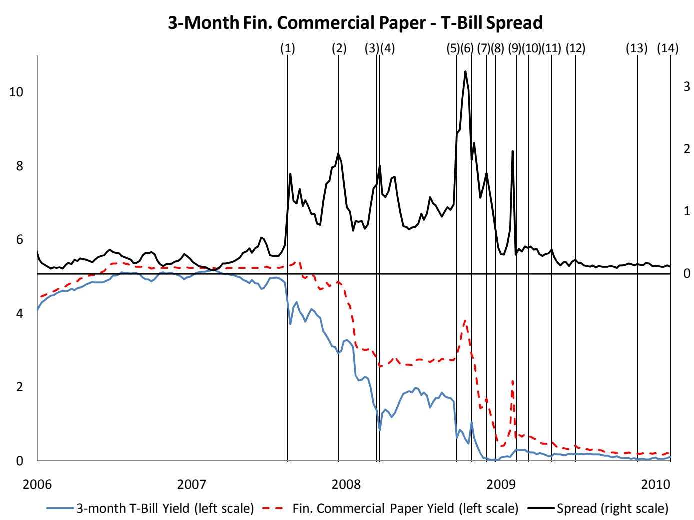


Figure 2
Plots of Yield Spreads and of Dates of Key Events in the Financial Crisis

[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 percent
[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







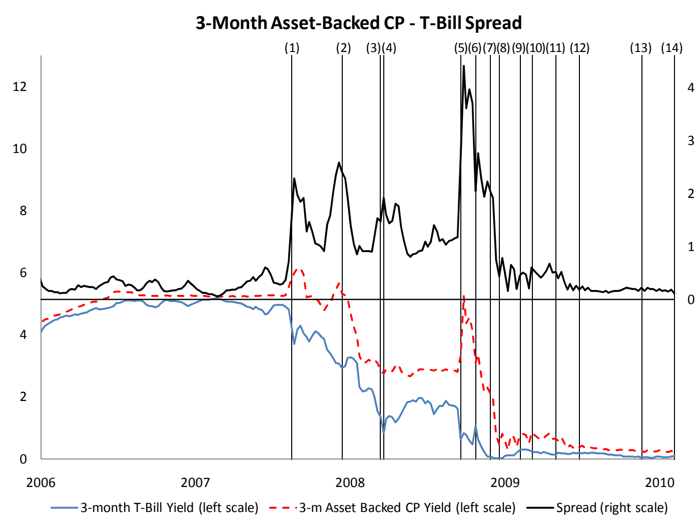
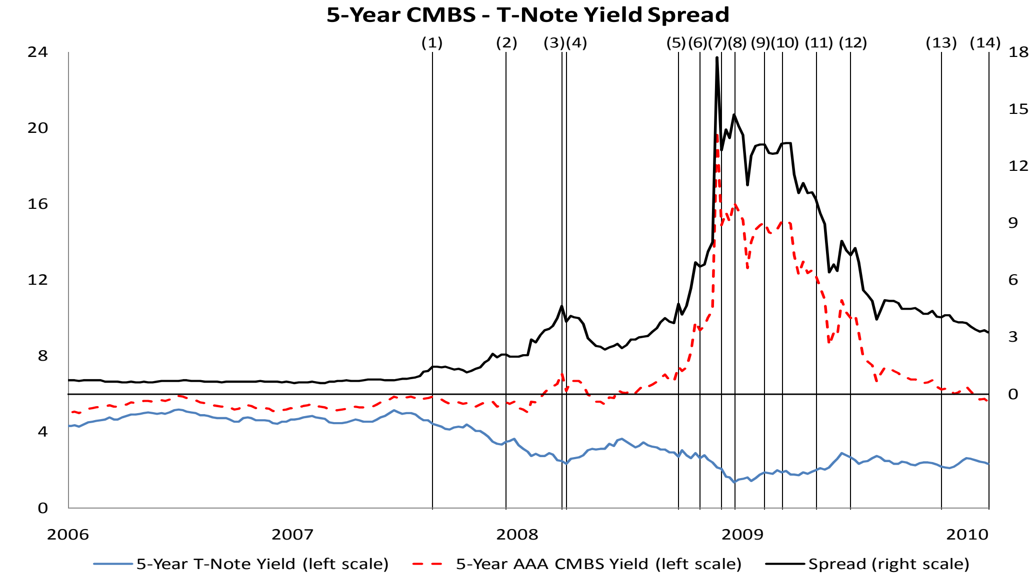
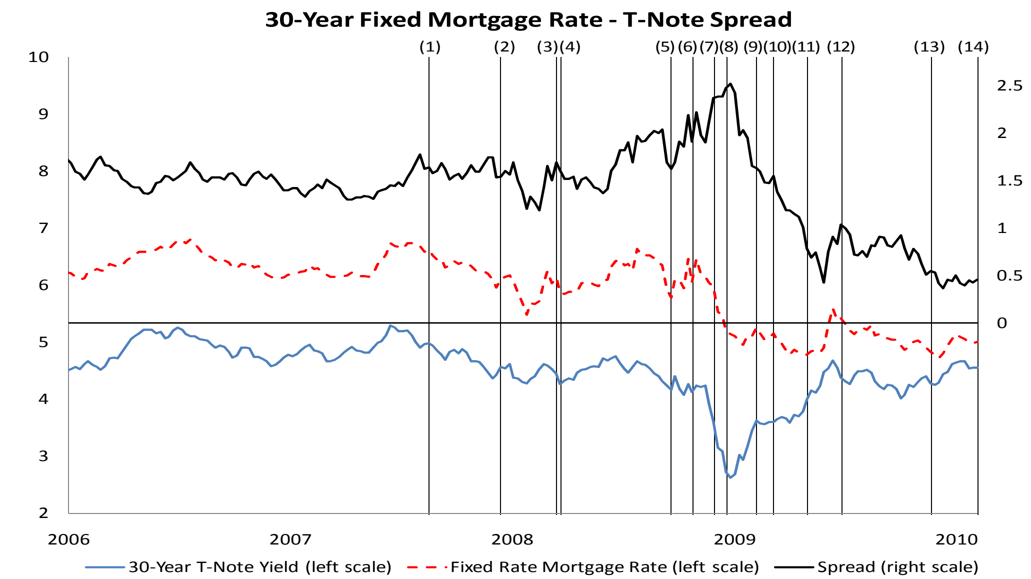


Figure 2 [continued]

Plots of Yield Spreads and of Date of Key Events in the Financial Crisis





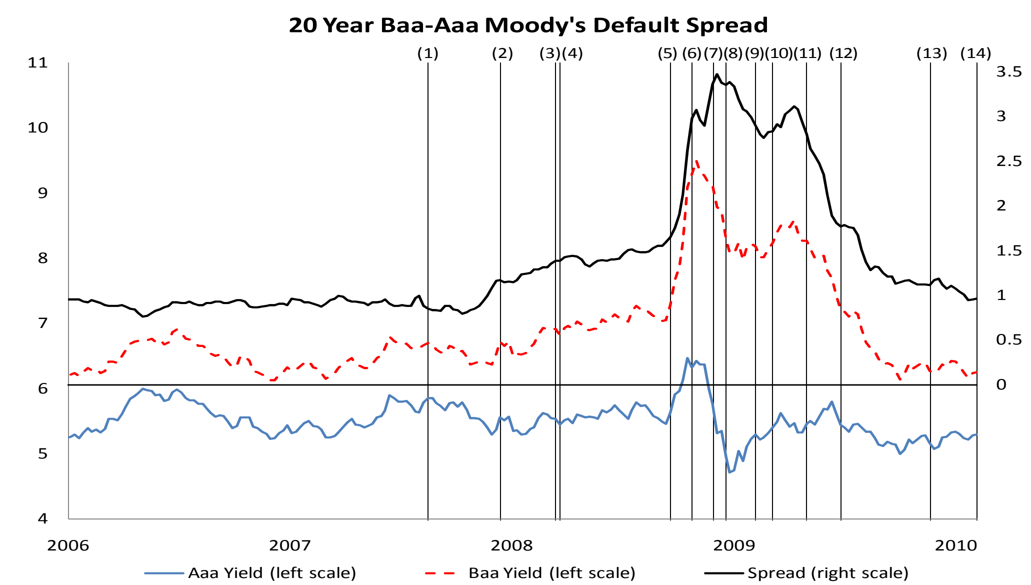


Figure 3

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 – September 2009), and for the post-crisis period (September 2009 – February 2010).