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Abstract

Estimating the effect of Federal Reserve's announcements of Large-Scale Asset Purchase (LSAP) programs on corporate credit risk is complicated by the simultaneity of policy decisions and movements in prices of risky financial assets, as well as by the fact that both interest rates of assets targeted by the programs and indicators of credit risk reacted to other common shocks during the recent financial crisis. This paper employs a heteroskedasticity-based approach to estimate the structural coefficient measuring the sensitivity of market-based indicators of corporate credit risk to declines in the benchmark market interest rates prompted by the LSAP announcements. The results indicate that the LSAP announcements led to a significant reduction in the cost of insuring against default risk—as measured by the CDX indexes—for both investment- and speculative-grade corporate credits. While the unconventional policy measures employed by the Federal Reserve to stimulate the economy have substantially lowered the overall level of credit risk in the economy, the LSAP announcements appear to have had no measurable effect on credit risk in the financial intermediary sector.

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Keywords: credit default swap (CDS), default risk channel, LSAPs, quantitative easing, identification through heteroskedasticity

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1 Introduction

The extraordinary turmoil that roiled global financial markets during the 2007–09 crisis and the subsequently sluggish pace of economic recovery led the Federal Reserve to take a number of unprecedented steps to improve market functioning and support economic activity. In terms of both sheer scale and prominence, the attempts to stimulate the economy by purchasing large quantities of government-backed securities in the secondary market—after the target federal funds rate was lowered to its effective zero lower bound at the end of 2008—have arguably been the most important unconventional policy measures employed by the Federal Open Market Committee (FOMC) in recent years; see D'Amico, English, López Salido, and Nelson [2012] for a thorough discussion of the role of asset purchases in the broader context of monetary policy strategy.

Formally referred to as the Large-Scale Asset Purchase (LSAP) programs, or "quantitative easing" in popular parlance, the programs were designed to lower longer-term market interest rates by purchasing debt obligations of the government-sponsored housing agencies (GSEs), mortgage-backed securities (MBS) issued by those agencies, and coupon securities issued by the United States Treasury. In addition to conducting two LSAP programs during the 2008–10 period, the FOMC in the autumn of 2011 also initiated a Maturity Extension Program (MEP), in an effort to put further downward pressure on longer-term interest rates and thereby provide additional stimulus to economic growth.[1]

The rationale underlying LSAPs hinges on a presumption that the relative prices of financial assets are to an important extent influenced by the quantity of assets available to investors. Implicit in this view is a departure from the expectation hypothesis of the term structure of interest rates and an appeal to theories of "imperfect asset substitution," "portfolio choice," or "preferred habitat," theories that recently have received renewed attention and rigorous micro foundations in the work of Andrés, López Salido, and Nelson [2004] and Vayanos and Vila [2009]. Indeed, in their communication of the likely effects of LSAPs on longer-term interest rates, policymakers have repeatedly invoked the preferred-habitat models of interest rate determination, as the canonical arbitrage-free term structure framework leaves essentially no scope for the relative supply of deeply liquid financial assets—such as nominal Treasuries—to influence their prices (Kohn [2009] and Yellen [2011]).[2]

[1] Under the MEP, which came to an end at the end of 2012, the Federal Reserve was selling shorter-term Treasury securities and using the proceeds to buy longer-term Treasury securities, thereby extending the average maturity of the securities in its portfolio.

[2] As formalized by Vayanos and Vila [2009], the preferred habitat models rely on two important assumptions: (1) the existence of investors that exhibit preferences for assets of only certain maturities; and (2) prices of assets of a given maturity are determined by the confluence of demand and supply shocks specific to that maturity. Within this framework, supply effects matter because a negative supply shocks to the stock of assets of a particular maturity—a large purchase of long-term Treasuries by the Federal Reserve, for example—creates a shortage of those assets, which at prevailing prices cannot be entirely offset by substitution of other securities. Examples of such preferred habitat investors include pension funds and insurance companies, a class of financial intermediaries that has an explicit preference for longer-term assets in order to match their long-duration liabilities. Short-term investors such as money

Given the unprecedented nature of the Federal Reserve's unconventional policy measures, a rapidly growing literature has emerged that tries to evaluate empirically the effects of the various asset purchase programs on financial asset prices. Perhaps not too surprisingly, the initial phase of this research has focused on the question of whether purchases of large quantities of Treasury coupon securities have altered the level of longer-term Treasury yields. Employing a high-frequency, event-style methodology Gagnon, Raskin, Remache, and Sack [2011], Swanson [2011], Krishnamurthy and Vissing-Jorgensen [2011], and Wright [2012] present compelling evidence that the Federal Reserve's LSAP announcements had economically and statistically significant effects on Treasury yields. Consistent with this evidence, Greenwood and Vayanos [2010a], Gagnon, Raskin, Remache, and Sack [2011], Krishnamurthy and Vissing-Jorgensen [2011], and Hamilton and Wu [2012] also show that Treasury supply factors have important effects on Treasury yields and the associated term premiums at lower frequencies and over longer sample periods.[3]

By cleverly exploiting the variation in prices across individual securities (CUSIPs) induced by the Federal Reserve's purchases of Treasury coupon securities, D'Amico and King [2013], find strong evidence of localized supply effects in the Treasury market—that is, purchases of specific CUSIPs in the secondary market had economically and statistically significant effect on yields of both purchased securities and those at nearby maturities.[4] Using a similar micro-level approach, D'Amico, English, López Salido, and Nelson [2012] attempt to disentangle the transmission channels involved in LSAPs and find that a significant portion of the variation in local supply and aggregate duration of Treasury securities was transmitted to longer-term Treasury yields through the term-premium component.[5]

Taking a different tack, Li and Wei [2013] develop and estimate an arbitrage-free term structure model of interest rates that, in addition to observable yield curve factors, incorporates variables involving the relative supply of Treasury and agency mortgage-backed securities. The inclusion of the Treasury-supply factor is motivated by the work of Vayanos and Vila [2009], while the inclusion of the MBS-supply factor reflects the market participants' perception of agency MBS as "safe"

market mutual funds, on the other hand, tend to hold Treasury bills and other short-dated claims to maintain a high degree of liquidity in their portfolios. Early treatment of these ideas can be found in Tobin [1961, 1963] and Modigliani and Sutch [1966, 1967].

[3]These findings are consistent with the work of Laubach [2009] and Krishnamurthy and Vissing-Jorgensen [2012], who find that fluctuations in the total supply of Treasury debt—conditional on the standard yield curve factors—have appreciable explanatory power for movements in Treasury yields. Relatedly, Greenwood and Vayanos [2010a,b] and Hamilton and Wu [2012] show that changes in the maturity structure of Treasury debt outstanding have a similar effect.

[4]Using the micro-level approach of D'Amico and King [2013], Meaning and Zhu [2011] find similar effects for the Bank of England's Asset Purchase Facility. Using an event-style analysis, Glick and Leduc [2012] present evidence that asset purchases by the Federal Reserve and the Bank of England—in addition to lowering longer-term government bond yields—also lowered the exchange value of the dollar and the pound as well as exerted downward pressure on commodity prices.

[5]According to their results, most of the decline in the overall term premium can be attributed to a reduction in the real term-premium component; the effect of LSAPs on the inflation risk premium, by contrast, was economically small and imprecisely estimated.

assets—and therefore close substitutes for Treasuries—owing to their earlier implicit and later explicit government guarantee. Both of these supply factors affect the term structure of interest rates primarily through the term-premium component, and according to Li and Wei [2013] estimates, the combined effect of the three LSAPs resulted in a significant reduction in longer-term Treasury yields.

While economists have devoted the lion's share of attention to evaluating the effects of LSAPs on Treasury yields, considerably less attention has been paid to the question of whether LSAPs had an effect on yields of riskier assets.[6] As emphasized by Krishnamurthy and Vissing-Jorgensen [2011], LSAPs can affect private yields through different channels. In this paper, we focus on one particular channel—the "default risk" channel. Specifically, we quantify the effect that the announcements of the three asset purchase programs—through their impact on the risk-free rates—had on market-based measures of corporate credit risk, both in its broad economy-wide sense and on credit risk specific to the financial sector.[7]

The focus on the former is motivated by the fact that if LSAPs were successful in stimulating the economy by lowering the general level of interest rates, we should observe a reduction in expected defaults and, as a result, a decline in corporate borrowing costs. Moreover, as the economic recovery gains traction, some standard asset pricing models imply an associated reduction not only in the compensation demanded by investors for expected default risk, but also in the average price of bearing exposure to corporate credit risk, above and beyond the compensation for expected defaults—that is, a reduction in the default risk premium. This increase in investor risk appetite—by lowering the price of default risk—should put additional downward pressure on corporate borrowing rates and thereby further stimulate business fixed investment, an especially cyclically-sensitive component of aggregate demand.

[6] An important early exception is Hancock and Passmore [2011], who analyze the effect of purchases of agency MBS on mortgage interest rates.

[7] In addition to the default risk channel, Krishnamurthy and Vissing-Jorgensen [2011] identify the following six channels through which LSAPs might affect market interest rates: (1) the "signaling" channel, according to which the unconventional monetary policy actions such as LSAPs exert downward pressure on longer-term interest rates only if such actions credibly signal the commitment that the monetary authority will maintain the short-term policy rate below its equilibrium level even after the output returns to its potential (Clouse, Henderson, Orphanides, and Tinsley [2003] and Eggertsson and Woodford [2003]); (2) the "duration risk" channel, in which purchases of Treasuries, agency debt, and agency MBS lower longer-term interest rates by reducing the amount of duration risk in the hands of preferred-habitat investors (Vayanos and Vila [2009]); (3) the "liquidity" channel, which posits that because the monetary authority finances its purchases of longer-term assets by increasing reserve balances of depository institutions, the resulting increase in the liquidity should lower liquidity premium on the most liquid assets; (4) the "safety" channel, a variation on the preferred habitat model, in which the preferred habitat applies to near-zero default-risk assets (Krishnamurthy and Vissing-Jorgensen [2012]); (5) the "prepayment risk premium" channel, another variant on the preferred habitat theme that requires segmentation of the MBS market, and in which purchases of MBS reduce the amount of prepayment risk, thereby lowering MBS yields (Gabaix, Krishnamurthy, and Vigneron [2007]); and (6) the "inflation" channel, according to which LSAPs influence nominal interest rates though changes in inflation expectations and uncertainty over inflation outcomes. It is worth emphasizing that because our paper is concerned with the response of market-based indicators of corporate credit risk to changes in the benchmark market interest rates induced by LSAPs, we do not have to take the stand on which of these channels is primarily responsible for the decline in interest rates.

The focus on credit risk specific to the financial sector, on the other hand, is motivated by an influential recent theoretical literature that stresses the implications of the capital position of financial intermediaries for asset prices and macroeconomic stability; see, for example, Brunnermeier and Sannikov [2011] and He and Krishnamurthy [2012, 2013]. The common thread running through these theories is that a deterioration in macroeconomic conditions, by depressing the capital base of financial intermediaries, induces a reduction in the risk-bearing capacity of the financial sector. To the extent that financial intermediaries are the marginal investors in asset markets, this effective increase in risk aversion causes a jump in the conditional volatility and correlation of asset prices and a sharp widening of credit spreads, a worsening of financial conditions that amplifies the effect of the initial shock on the macroeconomy.[8]

Any empirical investigation of the effect of LSAPs on corporate credit risk confronts a serious econometric challenge. First, yields of assets targeted by the central bank purchases—typically safe assets—may be simultaneously influenced by the movements in prices of risky financial assets, resulting in a difficult endogeneity problem. Second, the identification of the responsiveness of credit risk indicators to such policy interventions is complicated by the fact that a number of other factors, including news about the economic outlook and "flight-to-quality" consideration, likely had a significant effect on both the benchmark interest rates and market-based indicators of corporate credit risk during the period in which LSAPs were implemented.

In such circumstances, as we show below, the standard high-frequency, event-style regression analysis—which effectively assumes that the LSAP announcements are the sole source of volatility in the benchmark market interest rates on those days—will yield a downward-biased estimates of the coefficients measuring the effect of LSAPs on corporate credit risk indicators. Consistent with that observations, our event-style regression results indeed imply that the LSAP announcements had no impact on the cost of insuring against a broad-based incidence of defaults, as measured by the response of tradable credit derivative (CDX) indexes that are used widely by investors for hedging of and investing in corporate credit risk.[9]

To address these identification issues, we employ an alternative econometric approach developed by Rigobon [2003] and Rigobon and Sack [2003, 2004], which allows us to identify the parameter of interest—the structural response coefficient measuring the reaction of CDX indexes to declines in the benchmark interest rates induced by the LSAP announcements—under a weaker set of assumptions than those employed in our event-style analysis. In this so-called identification-through-heteroskedasticity approach, the response of credit risk indicators to policy interventions is identified vis-à-vis the shift in the variance of monetary policy shocks associated with policy announcements. As in Wright [2012], our identification strategy involves a natural assumption that the volatility

[8] Recent work by Adrian, Moench, and Shin [2010a,b] and Gilchrist and Zakrajšek [2012] provides empirical evidence supporting these types of intermediary asset pricing theories.

[9] This result stands in contrast to that of Krishnamurthy and Vissing-Jorgensen [2011], who using a similar methodology report a significant decline in the single-name credit default swap premiums—especially for riskier borrowers—in response to the LSAP announcements.

of monetary policy shocks increased on the days of the LSAP announcements, precisely because a larger portion of the news hitting financial markets was about monetary policy.[10]

In contrast to our event-style regression results, the heteroskedasticity-based approach implies that the declines in benchmark market interest rates induced by the LSAP announcements led to economically large and statistically significant reductions in the CDX indexes, both for the investment and speculative-grade segments of the U.S. corporate sector. The stark difference in the results from the two econometric approaches underscores the difficult identification issue of the so-called default risk channel of monetary policy transmission during the recent financial crisis, a period in which both policy rates—in this case yields on Treasuries, MBS, and agency debt—and credit risk indicators were likely reacting simultaneously to common shocks during days surrounding policy announcements. At the same time, our heteroskedasticity-based identification strategy implies that while the unconventional policy measures employed by the Federal Reserve to stimulate the economy in recent years have substantially lowered the overall level of credit risk in the economy, the LSAP announcements—somewhat to our surprise—had no measurable effect on credit risk in the financial intermediary sector.

The remainder of the paper is organized as follows. In Section 2, we discuss the LSAP announcement dates used in the analysis and present the necessary background evidence regarding the effect of those announcements on the key benchmark interest rates. Section 3 describes the construction of our credit risk indicators, both for the broad U.S. corporate sector and those pertaining to the financial sector. Section 4 contains our main results. It begins with an event-style analysis of the impact of the LSAP announcements on corporate credit risk and then shows why such an analysis may lead to a downward bias in the OLS estimator of the coefficient measuring the response of credit risk indicators to the changes in the benchmark market interest rates prompted by the LSAP announcements. To address this issue, it proposes a heteroskedasticity-based estimator of this effect and presents our key findings; to examine the robustness of our result, subsection 4.3 zeroes in on the five largest U.S. financial institutions, which play a key role in both the traditional bank-like credit intermediation process and in the arm's length finance that takes place in securities markets. Section 5 offers a brief conclusion.

2 Asset Purchase Programs and Benchmark Interest Rates

In this section, we present evidence from an event study of major announcements associated with the Federal Reserve's three asset purchase programs (LSAP-I, LSAP-II, and MEP). Gauging the effect of LSAP announcements on yields of assets purchased by the Federal Reserve—that is, Treasury

[10] Wright [2012] employs a structural VAR approach—with the identifying assumption that monetary policy shocks are heteroskedastic—to measure the effect of monetary policy shocks in the zero lower bound environment. In contrast to our paper, which seeks to quantify the impact of the LSAP announcements on corporate credit risk, Wright [2012] considers the total effects of FOMC news on a broader array of asset prices.

coupon securities, agency MBS, and agency debt—provides the necessary backdrop against which to evaluate the effects of the three asset purchase programs on corporate credit risk.

To maintain comparability with previous studies, we focus on the event dates identified by Krishnamurthy and Vissing-Jorgensen [2011]. Within the standard taxonomy of the Federal Reserve's asset purchase programs, these event dates are as follows:

- The first asset purchase program (LSAP-I):[11]

    1. Nov-25-2008: The initial announcement that the Federal Reserve would purchase up to \$100 billion of agency debt and up to \$500 billion of agency MBS.
    2. Dec-1-2008: Chairman Bernanke's speech on the "Federal Reserve Policies in the Financial Crisis," which suggested that the Federal Reserve could purchase longer-term Treasury securities in substantial quantities in order to stimulate the economy.
    3. Dec-16-2008: The FOMC statement that indicated "The Federal Reserve will continue to consider ways of using its balance sheet to further support credit markets and economic activity."
    4. Jan-28-2009: The FOMC statement that was interpreted by some market participants as disappointing because of its lack of concrete language regarding the possibility and timing of purchases of longer-term Treasuries in the secondary market.
    5. Mar-18-2009: The FOMC statement, which announced purchases of Treasury securities of up to \$300 billion and increased the size of purchases of agency MBS and agency debt to up to \$1.2 trillion and \$200 billion, respectively.

- The second asset purchase programs (LSAP-II):

    1. Aug-10-2010: The FOMC statement that was interpreted as suggesting a more downbeat assessment of the economic outlook than expected because it stated "To help support economic recovery in the context of price stability, the Committee will keep the Federal Reserve's holdings of securities at their current level by reinvesting principal payments from agency debt and agency mortgage-backed securities in longer-term Treasury securities. The Committee will continue to roll over the Federal Reserve's holdings of Treasury securities as they mature."
    2. Sep-21-2010: The FOMC statement that indicated the Committee will maintain its existing policy of reinvesting principal payments from its securities holdings.

[11] In their study of the first asset purchase program, Gagnon, Raskin, Remache, and Sack [2011] identified eight event dates, beginning with the Nov-25-2008 announcement and continuing through the autumn of 2009. Using intraday price and volume data, Krishnamurthy and Vissing-Jorgensen [2011] showed that significant changes in Treasury yields occurred only on the first five of the eight event dates and, therefore, leave out the last three event dates associated with the LSAP-I.

3. Nov-3-2010: The FOMC statement that indicated that in addition to maintaining the existing reinvestment policy, "[T]he Committee intends to purchase a further $600 billion of longer-term Treasury securities by the end of the second quarter of 2011, a pace of about $75 billion per month."

- The maturity extension program (MEP):

1. Sep-21-2011: The FOMC statement that indicated "To support a stronger economic recovery and to help ensure that inflation, over time, is at levels consistent with the dual mandate, the Committee decided today to extend the average maturity of its holdings of securities. The Committee intends to purchase, by the end of June 2012, $400 billion of Treasury securities with remaining maturities of 6 years to 30 years and to sell an equal amount of Treasury securities with remaining maturities of 3 years or less."

To obtain an estimate of the effect of various LSAP announcements on benchmark interest rates, we run the following event-style regression:

$$\Delta i_t = \theta_0 + \theta_1 \text{LSAP-I}_t + \theta_2 \text{LSAP-II}_t + \theta_3 \text{MEP}_t + \epsilon_t, \tag{1}$$

where Δi_t denotes the daily change in the specified interest rate; LSAP-I_t is a 0/1-variable that equals one on the five LSAP-I announcement days; LSAP-I_t is a 0/1-variable that equals one on the three LSAP-II announcement days; and MEP_t is a 0/1-variable that equals one on the MEP announcement day.[12] This specification implies that the coefficients θ_1, θ_2, and θ_3 measure the average effect of each LSAP program on the specified interest rate. We estimate equation (1) by OLS over the period from Jan-02-2008 to Dec-30-2011.

Table 1 summarizes the average effects of the program-specific LSAP announcements on yields of assets targeted by the three purchase programs.[13] According to the entries in the table, the effect of the first purchase program (LSAP-I) on market interest rates was substantial in economic terms: The average decline in longer-term Treasury yields induced by the five announcements was about 20 basis points, while yields on agency MBS and longer-term agency debt fell almost 25 basis

[12] In this type of event-style analysis, it is entirely possible that there are other "true"events that have been omitted. As discussed by Krishnamurthy and Vissing-Jorgensen [2011], failing to include potentially relevant dates into the analysis could result in an upward or a downward bias to the estimate of the overall effect of LSAPs on interest rates; importantly, the direction of the bias depends on how the events on the omitted dates affected the market participants' perception of the likelihood and magnitude of the purchase program. As a robustness check, we also considered a number of other potential event dates—especially those pertaining to the LSAP-II—but the inclusion of those additional event dates had no discernible effect on any of the results reported in the paper.

[13] All Treasury yields are derived from the smoothed Treasury yield curve estimated by Gürkaynak, Sack, and Wright [2007]; the agency MBS yield corresponds to the Fannie Mae 30-year current coupon MBS rate obtained from Barclays; and agency debt yields correspond to yield indexes based on individual senior unsecured bonds issued by Fannie Mae calculated internally at the Federal Reserve Board; we are grateful to Diana Hancock for providing us with these data.

Table 1: The Effect of LSAP Announcements on Selected Interest Rates
(*Event-Style Regression Analysis*)

Interest Rate	LSAP-I	LSAP-II	MEP	Pr > W^{a}	R^2
Treasury (1y)	-0.080^{***}	-0.002	0.037^{***}	0.000	0.012
	(0.024)	(0.002)	(0.002)		
Treasury (5y)	-0.190^{**}	-0.066^{***}	0.020^{***}	0.000	0.032
	(0.080)	(0.013)	(0.002)		
Treasury (10y)	-0.191^{**}	-0.044	-0.074^{***}	0.000	0.033
	(0.087)	(0.033)	(0.002)		
Agency MBS[b]	-0.239^{**}	-0.066	-0.136^{***}	0.000	0.040
	(0.098)	(0.040)	(0.003)		
S-T Agency[c]	-0.168^{***}	0.001	0.041^{***}	0.000	0.031
	(0.057)	(0.015)	(0.002)		
M-T Agency[d]	-0.197^{**}	-0.040^{**}	0.029^{***}	0.000	0.032
	(0.083)	(0.016)	(0.002)		
L-T Agency[e]	-0.267^{**}	-0.042	-0.037^{***}	0.000	0.057
	(0.109)	(0.029)	(0.002)		

NOTE: Sample period: daily data from Jan-02-2008 to Dec-30-2011. The dependent variable in each regression is the 1-day change in the specified interest rate. Entries in the table denote the OLS estimates of the average effect (in percentage points) of LSAP-I, LSAP-II, and MEP announcements. LSAP-I = a 0/1-variable that equals one on the following five announcement days: Nov-25-2008, Dec-01-2008, Dec-16-2008, Jan-28-2009, and Mar-18-2009; LSAP-II = a 0/1-variable that equals one on the following three announcement days: Aug-10-2010, Sep-21-2010, and Nov-03-2010; and MEP = a 0/1-variable that equals one on the announcement date of Sep-21-2011. All specifications include a constant (not reported). Heteroskedasticity-consistent asymptotic standard errors are reported in parentheses: * $p < 0.10$; ** $p < 0.05$; and *** $p < 0.01$.

[a] p-value of the test of the null hypothesis that the LSAP-I, LSAP-II, and MEP announcement effects are jointly equal to zero.

[b] Fannie Mae 30-year current coupon MBS rate.

[c] Yield index on Fannie Mae's shorter-term (maturity 1–3 years) senior unsecured bonds.

[d] Yield index on Fannie Mae's medium-term (maturity 3–5 years) senior unsecured bonds.

[e] Yield index on Fannie Mae's longer-term (maturity greater than 5 years) senior unsecured bonds.

points, on average; as indicated by the associated standard errors, these declines are all statistically significant at conventional levels.

The effects of the second purchase program (LSAP-II) on interest rates were also economically important, though not as widespread as those of the LSAP-I. According to our estimates, the average decline in the 5-year Treasury yield in response to the three LSAP-II announcements was about 7 basis points, while yields on medium-term (3–5 years) agency debt fell 4 basis points.

The last program undertaken by the Federal Reserve during our sample period (MEP) also had predictable effects, as the declines in interest rates were concentrated at the longer-end of the maturity spectrum. Indeed, as envisioned by the FOMC, the MEP significantly flattened the Treasury yield curve, both by depressing the long-end and by inducing a small rise at the short- and intermediate-end of the yield curve. All told, these results are consistent with the recent work of Gagnon, Raskin, Remache, and Sack [2011], Krishnamurthy and Vissing-Jorgensen [2011], D'Amico, English, López Salido, and Nelson [2012], Wright [2012], and D'Amico and King [2013], who document that asset purchase programs had significantly altered the level of longer-term government bond yields. They are also consistent with the event-style evidence presented by Swanson [2011], who shows that the Federal Reserve's large purchases of longer-term Treasury securities during the 1961 Operation Twist had a major effect on financial markets.

3 Measuring Corporate Credit Risk

This section describes the construction of credit risk indicators used in our analysis. In all instances, the indicators are based on financial derivatives on credit risk—that is, (single-name) credit default swaps (CDS)—instruments used extensively by investors for hedging of and investing in credit risk. A CDS is simply an insurance contract between two parties: a protection buyer, who makes fixed, periodic payments; and a protection seller, who collects these premiums in exchange for making the protection buyer whole in case of default.[14] Although akin to insurance, CDS are not regulated by insurance regulators—they are over-the-counter (OTC) transactions—and unlike standard insurance contracts, it is not necessary to own the underlying debt in order to buy protection using CDS. In general, CDS trades take place between institutional investors and dealers, with the legal framework for trading governed by the International Swaps and Derivatives Association (ISDA).

Using CDS contracts to measure credit risk has a number of advantages over other commonly-used indicators of credit risk such corporate bond credit spreads. First, it is far easier to buy credit protection than to short corporate bonds. As a result, CDS are a natural vehicle for shorting default risk, which allows investors to take a specific view on the outlook for credit quality of a specific company. Second, the use of CDS contracts allows users to avoid triggering tax or accounting implications that arise from sales of actual bonds. Third, CDS contracts offer easy access to hard to find credits, reflecting a limited supply of bonds in many instances. And lastly, investors can more easily tailor their credit exposure to maturity requirements and desired seniority in the firm's capital structure. It should be noted that although CDS are available at various maturities, the 5-year contract is by far the most commonly traded and liquid segment of the market.

3.1 Overall Credit Risk

We rely on *credit derivative indexes* owned and managed by Markit as a comprehensive measure of credit risk at the broad, economy-wide level. Compared with other commonly-referenced financial indexes, such as indexes of corporate bond yields and spreads or equity indexes, credit derivative

[14] CDS contracts generally trade based on a spread, which represents the cost a protection buyer has to pay the protection seller (i.e., the premium paid for protection). For the protection buyer, the value of the CDS contract rises if the spread increases; for example, a protection buyer paying a spread of 100 basis points, when the current spread is 150 basis points, would be able to unwind the position at a higher spread level.

indexes are *tradable* products. Buying and selling of the credit derivative index is comparable to buying and selling portfolios of corporate cash instruments: By buying the index the investor takes on the credit exposure—is exposed to defaults—a position similar to that of buying a portfolio of bonds; by selling the index, the credit exposure is passed on to another party. As a result, investors can take positions directly in the indexes without having to trade a large number of underlying components—in fact, index trading is often thought to lead single-name CDS trading.

To capture the full spectrum of credit quality in the U.S. corporate market, we consider two indexes: the (North American) 5-year CDX investment-grade index (CDX-IG); and the (North American) 5-year CDX speculative-grade index (CDX-SG). The investment-grade CDX index references 125 CDS on firms that have an investment-grade rating from both Moody's and Standard & Poor's at the time the index starts trading, while the speculative-grade CDX index references 100 CDS on firms that have a "junk" rating from at least one rating agency. Importantly, the component firms must have highly liquid single-name CDS trading in their name, and the composition of both indexes—which is determined by a dealer poll—reflects the composition of the U.S. corporate sector.

All firms in the indexes are equally weighted, and the composition of both indexes is fixed once the indexes start trading.[15] However, new vintages of the indexes are introduced every six months, and the new vintages may have different components than the old vintages. When a new vintage is introduced, it becomes the "on-the-run" vintage; previous versions of the indexes continue trading as "off-the-run" vintages.[16] To ensure the maximum liquidity for our broad indicators of corporate credit risk, we spliced together the on-the-run vintages for both the investment- and speculative-grade CDX indexes.[17]

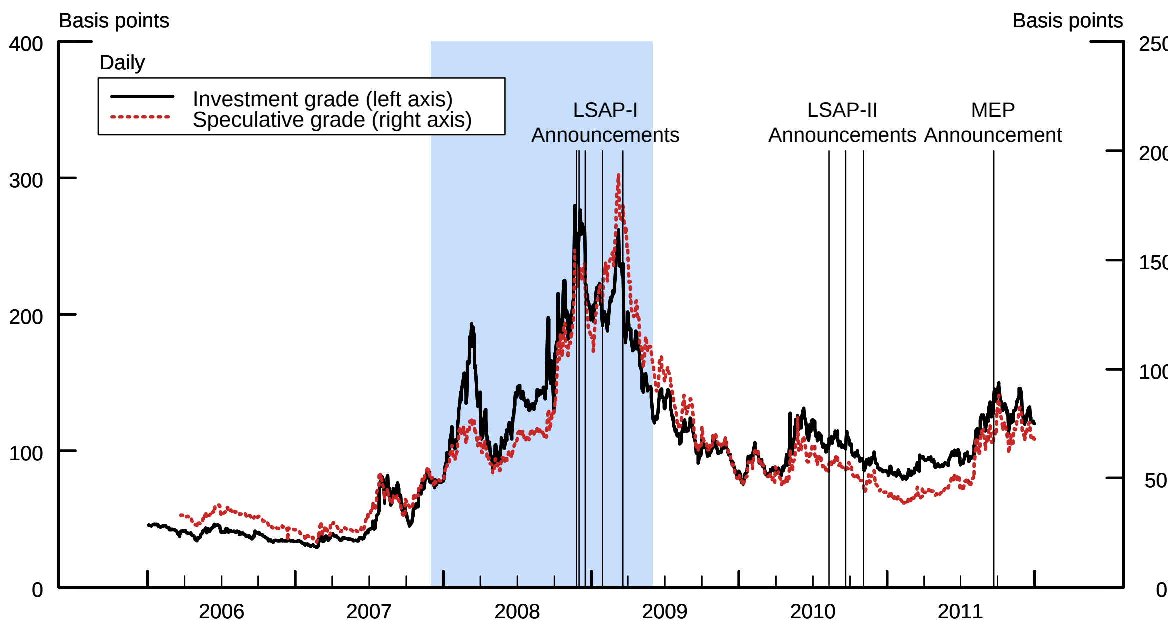
Figure 1 shows our two main credit risk indicators over the 2007–11 period. According to these two measures, credit risk in the U.S. corporate sector increased noticeably with the onset of the financial crisis in the summer of 2007, likely reflecting the rapidly deteriorating outlook for the housing sector and associated concerns about the possible spill-over effects on financial institutions and the broader economy. Both indicators were exceptionally volatile during the subsequent recession and spiked sharply at critical events of the crisis: the collapse of Bear Stearns investment bank in March 2008; the bankruptcy of Lehman Brothers in mid-September 2008; and in early 2009, when continued pressures on already-strained balance sheets of financial intermediaries threatened

[15] Because the indexes are equally weighted, in the event of a default of one index component, the notional value of the contract is reduced by $1/n$, where n is the number of firms in the index; the index will then continue trading with $n-1$ remaining firms—that is, there is no replacement for the defaulted firm.

[16] As noted above, the composition of each index vintage is determined by a dealer poll, starting from the composition of the previous index. Typically, firms that became illiquid in the single-name CDS market or that have been downgraded/upgraded by one or more rating agencies are removed from the index and replaced with new firms.

[17] While this approach minimizes, to the extent possible, the role of liquidity effects in our indicators of default risk, it does introduce the compositional changes in the CDX index that occur every six months when a new vintage is launched. These compositional effects, however, appear to be fairly small, reflecting the relatively small number of component changes from one vintage to another. As a robustness check, we re-did the analysis using the off-the-run CDX indexes, and the results were essentially the same.

Figure 1: Market-Based Indicators of Credit Risk in the U.S. Corporate Sector



NOTE: The solid line depicts the 5-year (on-the-run) investment-grade CDX index (CDX-IG), while the dotted line depicts the 5-year (on-the-run) speculative-grade CDX index (CDX-SG). The shaded vertical bar represents the 2007–09 NBER-dated recession. LSAP-I announcement days are Nov-25-2008, Dec-01-2008, Dec-16-2008, Jan-28-2009, and Mar-18-2009; LSAP-II announcement days are Aug-10-2010, Sep-21-2010, and Nov-03-2010; and the MEP announcement date is Sep-21-2011.

the viability of some of the institutions, a situation that was greatly exacerbated by the lingering effects of the deep economic contraction that materialized in the second half of 2008.

Indeed, as shown by the thin vertical lines, the first round of asset purchases (LSAP-I) was launched in response to these adverse economic developments and to help stimulate economic growth. The flare-up in CDX spreads in the spring of 2010, which was part of a deterioration in broad financial conditions, largely reflected investors' concerns about European sovereign debt and banking issues as well as concerns about the durability of the global recovery. Although financial market conditions improved somewhat early in the second half of 2010—partly as investors increasingly priced in further monetary policy accommodation—the Federal Reserve initiated LSAP-II in the second half of the year, in response to indications of a slowing pace of economic recovery and persistent disinflationary pressures.

Financial markets were buffeted again over the second half of 2011 by changes in investors' assessments of the ongoing European crisis as well as in their evaluation of the U.S. economic outlook. As a result, the credit outlook for the corporate sector deteriorated markedly. With economic activity expanding only at a slow pace and with labor market conditions remaining weak, the FOMC launched the MEP with the intent to put further downward pressure on longer-term interest rates and to help make broader financial conditions more accommodative.

3.2 Financial Sector Credit Risk

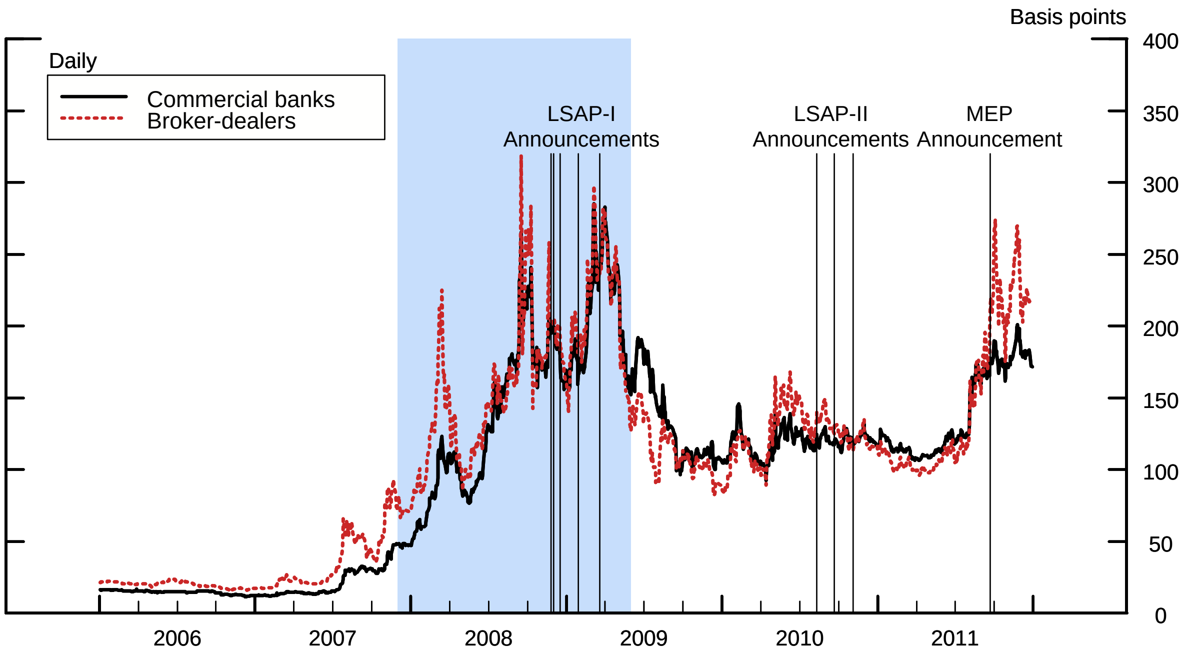
We now turn to the construction of credit risk indicators specific to the financial sector. In light of the above discussion, the most natural such indicator would be a credit derivative index, with its components referencing CDS contracts of a broad array of U.S. financial institutions. Such an index, unfortunately, does not exist. As an alternative, we use the single-name (North American) 5-year CDS contracts to construct simple indexes of credit risk for two types of financial institutions: commercial banks (CDS-BK) and securities broker-dealers (CDS-BD).

Our focus on these two types of financial intermediaries is motivated by the fact that a significant portion of the credit extended to businesses and households—both on- and off-balance-sheet—is through the commercial banking sector (Bassett, Gilchrist, Weinbach, and Zakrajšek [2011]). Thus, our commercial bank CDS index captures the market-based assessment of the credit risk for the traditional class of financial intermediaries. In contrast, securities broker-dealers, a class of highly leveraged financial intermediaries, buy and sell a large array of securities for a fee, hold an inventory of securities for resale, and differ from other types of institutional investors by their active pro-cyclical management of leverage. As documented by Adrian and Shin [2010], expansions in broker-dealer assets are associated with increases in leverage as broker-dealers take advantage of greater balance sheet capacity; conversely, contractions in their assets are associated with de-leveraging of their balance sheets. Reflecting their role as a "marginal investor," broker-dealers play an important role in most financial markets, and, as shown by Gilchrist and Zakrajšek [2012], changes in their creditworthiness—as measured by the movements in their CDS spreads—are closely related to fluctuations in the effective risk-bearing of the broader financial sector.

To construct these credit risk indicators, we selected from the Markit's single-name database the daily 5-year CDS quotes for a sample of 26 U.S. commercial banks and nine U.S. broker-dealers. In terms of the triggering events, we focus on the contracts with the Modified Restructuring (MR) clause, which, in addition to an outright default, considers any change in the nature of a company's financial liabilities in the absence of default as a credit event.[18] Using these micro-level data, we

[18]Restructuring without an outright default could include a reduction in interest rate or principal; postponement or deferral of payment; a change in the priority of an obligation; or a change in the composition of a principal or interest payment. On April 8, 2008, ISDA implemented the so-called Big Bang Protocol in order to increase standardization to single-name CDS contracts and trading conventions. For our purposes, the most important change was a move toward "XR" contracts, which do not include restructuring as a credit event. Historically, the market convention has been for (North American) investment-grade names—the segment of the credit quality spectrum that is relevant for the financial sector—to trade with an MR clause and (North American) speculative-grade names to trade with an XR clause. The Big Bang made XR contracts the standard convention for all names. All else equal, CDS contracts with an MR clause should trade at a premium to XR contracts because they provide the protection buyer with coverage for an additional type of credit event; moreover, there is some evidence suggesting that since the Big Bang, XR contracts have become the most liquid segment of the market. As a robustness check, we re-did the analysis using CDS spreads based on the XR contracts, but this change had a minimal effect on our results.

Figure 2: Market-Based Indicators of Credit Risk in the U.S. Financial Sector



NOTE: The solid line depicts the average (5-year) CDS index for a sample of 26 U.S. commercial banks (CDS-BK), while the dotted line depicts the average (5-year) CDS index for a sample of nine U.S. broker-dealers (CDS-BD). The shaded vertical bar represents the 2007–09 NBER-dated recession. LSAP-I announcement days are Nov-25-2008, Dec-01-2008, Dec-16-2008, Jan-28-2009, and Mar-18-2009; LSAP-II announcement days are Aug-10-2010, Sep-21-2010, and Nov-03-2010; and the MEP announcement date is Sep-21-2011.

calculate CDS indexes for the portfolios of both commercial banks and broker-dealers as simple (unweighted) cross-sectional averages of the available spreads at each day.

As shown in Figure 2, the behavior of these two credit risk indicators over the recent crisis closely mimics that of the broader investment-grade U.S. corporate sector—financial firms are rated almost exclusively as investment grade by the major rating agencies. One problem with the construction of these indexes is that the underlying micro-level panels are of unbalanced nature, as smaller and less prominent institutions have occasional gaps in their CDS series. The lack of reliable CDS quotes for certain institutions on certain days is most likely due to the sporadic impairment in the functioning of the credit derivatives market, especially during the most acute phases of the financial crisis. While the cross-sectional average of the component quotes provides an unbiased estimate of the average *level* of CDS spreads at any given point in time, our formal analysis relies on *changes* in the credit risk indicators. Because of potential changes in the composition of the indexes between two periods, taking first difference of the indexes shown in Figure 2 would, consequently, introduce a significant amount of noise in the differenced series, thereby complicating the interpretation of the results.

Table 2: The Effect of LSAP Announcements on Corporate Credit Risk
(*Event-Style Regression Analysis*)

Announcement	CDX-IG	CDX-SG	CDS-BK	CDS-BD
LSAP-I	−0.049	−0.237	−0.020	−0.015
	(0.068)	(0.257)	(0.205)	(0.027)
LSAP-II	0.011	0.030	0.002	−0.012
	(0.009)	(0.054)	(0.020)	(0.028)
MEP	0.063***	0.331***	0.051***	0.175***
	(0.002)	(0.009)	(0.001)	(0.003)
Pr $> W^{a}$	0.000	0.000	0.000	0.000
R^2	0.005	0.001	0.003	0.005

NOTE: Sample period: daily data from Jan-02-2008 to Dec-30-2011. The dependent variable in each regression is the 1-day change in the specified credit risk indicator: CDX-IG = 5-year (on-the-run) investment-grade CDX index; CDX-SG = 5-year (on-the-run) speculative-grade CDX index; CDS-BK = 5-year CDS index for 26 commercial banks; and CDS-BD = 5-year CDS index for 9 broker-dealers. Entries in the table denote the OLS estimates of the average effect (in percentage points) of LSAP-I, LSAP-II, and MEP announcements. LSAP-I = a 0/1-variable that equals one on the following five announcement days: Nov-25-2008, Dec-01-2008, Dec-16-2008, Jan-28-2009, and Mar-18-2009; LSAP-II = a 0/1-variable that equals one on the following three announcement days: Aug-10-2010, Sep-21-2010, and Nov-03-2010; and MEP = a 0/1-variable that equals one on the announcement date of Sep-21-2011. All specifications include a constant (not reported). Heteroskedasticity-consistent asymptotic standard errors are reported in parentheses: * $p < 0.10$; ** $p < 0.05$; and *** $p < 0.01$.
[a] p-value of the test of the null hypothesis that the LSAP-I, LSAP-II, and MEP announcement effects are jointly equal to zero.

To deal with this problem, we first compute the difference of the daily CDS spreads for each component of the two indexes and then compute the cross-sectional (unweighted) averages for the two portfolios of financial firms. When calculating the cross-sectional averages of the 1-day changes in CDS spreads, we use trimmed means, which delete the smallest and largest change in CDS spreads from the sample at each point in time. By using such a robust estimator of the population mean, we mitigate the effects of extreme values that might arise from abrupt changes in the liquidity of the single-name CDS market.[19]

4 Asset Purchase Programs and Corporate Credit Risk

To examine the effect of the LSAP announcements on corporate credit risk, we first re-estimate the event-style regression specification (1), using changes in our four credit risk indicators as dependent variables. The results of this exercise are reported in Table 2. According to these results, the announcement effects of the first two asset purchase programs (LSAP-I and LSAP-II) had no

[19]As an additional sensitivity check, we also computed winsorized means of the underlying CDS changes and obtained essentially the same results; see Maronna, Martin, and Yohai [2006] for a useful exposition of robust estimators.

Table 3: Changes in Interest Rates and Corporate Credit Risk

Interest Rate	CDX-IG	CDX-SG	CDS-BK	CDS-BD
Treasury (5y)	-0.281^{***}	-1.106^{***}	-0.157^{***}	-0.392^{***}
	(0.042)	(0.136)	(0.034)	(0.065)
R^2	0.153	0.116	0.102	0.140
Agency MBS[a]	-0.099^{***}	-0.435^{***}	-0.069^{***}	-0.176^{***}
	(0.036)	(0.148)	(0.023)	(0.043)
R^2	0.024	0.023	0.025	0.036
Agency Debt[b]	-0.170^{***}	-0.714^{***}	-0.116^{***}	-0.282^{***}
	(0.041)	(0.170)	(0.030)	(0.059)
R^2	0.058	0.051	0.058	0.076

NOTE: Sample period: daily data from Jan-02-2008 to Dec-30-2011. The dependent variable in each regression is the 1-day change in the specified credit risk indicator: CDX-IG = 5-year (on-the-run) investment-grade CDX index; CDX-SG = 5-year (on-the-run) speculative-grade CDX index; CDS-BK = 5-year CDS index for 26 commercial banks; and CDS-BD = 5-year CDS index for 9 broker-dealers. Entries in the table denote the OLS estimates of the coefficients associated with the 1-day change in the specified interest rate. All specifications include a constant (not reported). Heteroskedasticity-consistent asymptotic standard errors are reported in parentheses: * $p < 0.10$; ** $p < 0.05$; and *** $p < 0.01$.

[a] Fannie Mae 30-year current coupon MBS rate.

[b] Yield index on Fannie Mae's longer-term (maturity greater than 5 years) senior unsecured bonds.

statistically discernible effect on the broad indicators of corporate credit risk or on the average CDS premiums of commercial banks and broker-dealers, despite the fact that the news of these two programs induced a significant decline in longer-term market interest rates.

In contrast, the MEP announcement is associated with a statistically significant increase in all four credit risk indicators, with the effect being most pronounced for the speculative-grade CDX index (33 basis points) and for the CDS spreads of broker-dealers (17 basis points). These sizable increases in the cost of insuring against corporate defaults induced by the MEP announcement are likely due to the fact that the announced size of the program was somewhat less than than markets had originally anticipated. At first glance, the results in Tables 1 and 2 suggest that while the LSAP announcements led to significant declines in the benchmark market interest rates, they had no obvious effect on the cost of insuring against a specter of corporate defaults, both in an economy-wide sense or against defaults specific to the financial sector.

To understand the lack of response of CDS spreads to the LSAP announcements, it is instructive to examine the relationship between changes in interest rates and changes in CDS spreads during the entire 2008–11 period. In Table 3, we report the coefficients from the regression of the daily change in each of our four credit risk indicators on the daily change in the three benchmark interest rates: the 5-year Treasury yield, the 30-year MBS rate, and the rate on longer-term agency bonds.

Several points about these results are worth noting. First, all the coefficients on interest rate changes are negative and highly significant, both in economic and statistical terms. This strong negative relationship implies that when longer-term risk-free rates were falling during the crisis, the cost of insurance against corporate defaults was rising sharply. Second, in terms of the type of interest rate, changes in Treasury yields appear to have had the largest economic impact on the CDS spreads, followed by changes in yields on longer-term agency bonds. And lastly, the largest (absolute) coefficients are associated with the speculative-grade firms (CDX-SG) and with the broker-dealers (CDS-BD), two relatively highly leveraged segments of the U.S. corporate sector.

Two natural and related interpretations of these results spring to mind. The first is that the negative relationship between changes in the benchmark market interest rates and corporate credit risk is driven by adverse news to economic fundamentals, which signals a deterioration in the outlook for credit quality, reflecting a downward revision to future growth prospects. As a result, the cost of default insurance increases, while longer-term risk-free rates decline. This interpretation is consistent with the result that the negative relationship between changes in risk-free interest rates and CDS spreads is most pronounced for lower-rated corporate credits, a segment of the market that was especially vulnerable to adverse macroeconomic shocks during this period.

The second interpretation is that there are shocks to risk premiums that trigger a "flight-to-quality," a phenomenon that causes investors to dump risky assets to purchase safer investments. In that case, the expected returns on risky assets increase, while those on riskless assets fall. This interpretation is consistent with the result that, in absolute terms, the largest coefficients on interest rate changes are associated with Treasuries and longer-term agency bonds.[20] As we show in the next section, both of these mechanisms will lead to a downward bias in the OLS estimates of the coefficient measuring the response of CDS spreads to changes in market interest rates induced by the LSAP announcements.

4.1 Identification Through Heteroskedasticity

As emphasized by Rigobon and Sack [2003, 2004], causal inference regarding the impact of policy announcements on asset prices may be difficult in an environment where both policy rates—in our case yields on Treasuries, MBS, and agency debt—and asset prices endogenously respond to common shocks in periods surrounding policy announcements. To illustrate the argument more formally, let Δi_t denote the change in yields on safe assets that are directly influenced by the LSAP announcements, and let Δs_t denote the change in yields on risky corporate assets, as measured by the changes in the relevant CDS spreads. Furthermore, let x_t denote a common shock that simultaneously affects both CDS spreads and risk-free interest rates, and let ϵ_t represent policy shock—the LSAP announcement—while η_t is a shock to CDS spreads that is independent of the common shock x_t. It is assumed that disturbances x_t, ϵ_t, and η_t are homoskedastic with variances σ_x^2, σ_ϵ^2, and σ_η^2, respectively.

[20]While the government guarantee behind agency MBS was certainly prized by investors, these securities carry a significant prepayment risk. They were therefore unlikely, during flight-to-quality episodes, to be viewed by investors as safe as Treasuries or agency bonds.

The response of interest rates and CDS spreads to various shocks is captured by a simultaneous system of equations of the form:

$$\Delta i_t = \beta \Delta s_t + \gamma x_t + \epsilon_t; \tag{2}$$

$$\Delta s_t = \alpha \Delta i_t + x_t + \eta_t. \tag{3}$$

The aim of the exercise is to estimate the structural response coefficient α, which measures the direct effect of changes in risk-free interest rates on changes in CDS spreads vis-à-vis the announcement effect ϵ_t. In this context, a positive realization of x_t can be interpreted as an adverse shock to economic fundamentals (i.e., growth prospects), which increases the likelihood of future defaults. If risk-free rates decline in response to a deterioration in economic outlook, it follows that $\gamma < 0$. The shock η_t, by contrast, can be interpreted as an innovation to risk premiums, which captures movements in credit spreads unrelated to economic fundamentals x_t.[21] Assuming that risk-free rates fall when credit spreads widen, we have $\beta < 0$. This may occur either because market interest rates respond to deteriorating credit conditions or because monetary policy responds directly to an increase in credit spreads.

The reduced-form of the system (2)–(3) is given by

$$\begin{aligned} \Delta i_t &= \frac{1}{1-\alpha\beta}\big[(\beta+\gamma)x_t + \beta\eta_t + \epsilon_t\big]; \\ \Delta s_t &= \frac{1}{1-\alpha\beta}\big[(1+\alpha\gamma)x_t + \eta_t + \alpha\epsilon_t\big], \end{aligned}$$

and the bias of the response coefficient α obtained from an OLS regression of Δs_t on Δi_t can be understood by computing the probability limit of the coefficient estimate:

$$\operatorname{plim} \alpha_{OLS} = \alpha + (1-\alpha\beta)\left[\frac{\beta\sigma_\eta^2 + (\beta+\gamma)\sigma_x^2}{\sigma_\epsilon^2 + \beta^2\sigma_\eta^2 + (\beta+\gamma)^2\sigma_x^2}\right].$$

A natural assumption in our context is that $\beta < 0$ and $\gamma < 0$. Hence, assuming that the true response coefficient $\alpha > 0$—so that policy actions aimed at lowering market interest rates (i.e. a negative shock to ϵ_t) also reduce credit risk—the presence of simultaneity bias (owing to $\sigma_x^2 > 0$) and omitted variable bias (owing to $\sigma_\eta^2 > 0$) implies a downward bias in the OLS estimate of α. Alternatively, assume that $\alpha < 0$. In that case, as long as policy actions effectively reduce risk-free rates so that $(1 - \alpha\beta) > 0$, the OLS estimate of the response coefficient α will also be biased downward.

[21] Alternatively, one may view η_t as a "liquidity" shock that affects the CDS market but does not influence risk-free rates and interpret x_t as a "flight-to-quality shock" that raises risk premiums on risky assets and lowers yields on safe assets. As we show below, the bias of the OLS coefficient obtained by regressing changes in the CDS spreads on changes in risk-free rates depends on the sign of the coefficients β and γ and not on the specific interpretation of the two shocks.

Note that the bias of the OLS estimate of the coefficient α is a decreasing function of σ_ϵ^2, the variance of the policy shock. As σ_ϵ^2 becomes large relative to σ_η^2 and σ_x^2, the bias disappears. As emphasized by Rigobon and Sack [2004], high-frequency event studies that use OLS to estimate the effect of changes in policy rates on financial asset prices effectively assume that ϵ_t is the only source of volatility on policy announcement days. Given the heightened volatility and strains that characterized financial markets during this period, this seems a questionable assumption.

Building on the work of Rigobon [2003], Rigobon and Sack [2004] propose an estimator for the response coefficient α that is identified through the fact that the volatility of policy shocks increases on policy announcement days.[22] The essential idea is that by knowing *a priori* on which dates the variance of policy shocks shifts, the researcher is able to identify the response of asset prices to changes in policy rates under a fairly weak set of assumptions by looking at changes in the comovement between policy rates and financial asset prices.

More concretely, let P denote a subset of T_P policy announcement days and $\tilde{P}$ denote a subset of $T_{\tilde{P}}$ non-announcement days. Furthermore, let

$$\sigma^2_{\epsilon|P} = E[\epsilon_t^2 \,|\, t \in P] \quad \text{and} \quad \sigma^2_{\epsilon|\tilde{P}} = E[\epsilon_t^2 \,|\, t \in \tilde{P}]$$

denote the conditional variances of policy shocks in the P and $\tilde{P}$ subsamples, respectively, while

$$\mathbf{V}_P = E\left\{ \begin{bmatrix} \Delta i_t \\ \Delta s_t \end{bmatrix} \begin{bmatrix} \Delta i_t & \Delta s_t \end{bmatrix} \,\middle|\, t \in P \right\} \quad \text{and} \quad \mathbf{V}_{\tilde{P}} = E\left\{ \begin{bmatrix} \Delta i_t \\ \Delta s_t \end{bmatrix} \begin{bmatrix} \Delta i_t & \Delta s_t \end{bmatrix} \,\middle|\, t \in \tilde{P} \right\}$$

are the covariance matrices of the vector $[\Delta i_t \;\, \Delta s_t]'$ across the two subsamples. Then under the identifying assumption that $\sigma^2_{\epsilon|P} \neq \sigma^2_{\epsilon|\tilde{P}}$,

$$\Delta\mathbf{V} = \mathbf{V}_P - \mathbf{V}_{\tilde{P}} = \left(\frac{\sigma^2_{\epsilon|P} - \sigma^2_{\epsilon|\tilde{P}}}{(1-\alpha\beta)^2} \right) \begin{bmatrix} 1 & \alpha \\ \alpha & \alpha^2 \end{bmatrix},$$

and

$$\hat{\alpha}_i = \frac{\Delta\widehat{\mathbf{V}}_{12}}{\Delta\widehat{\mathbf{V}}_{11}} \quad \text{or} \quad \hat{\alpha}_s = \frac{\Delta\widehat{\mathbf{V}}_{22}}{\Delta\widehat{\mathbf{V}}_{12}},$$

where $\Delta\widehat{\mathbf{V}}_{ij}$ denotes the (i,j)-th element of the sample analogue of the matrix $\Delta\widehat{\mathbf{V}} = \widehat{\mathbf{V}}_P - \widehat{\mathbf{V}}_{\tilde{P}}$, are both consistent estimators of the structural coefficient α. Note that a necessary condition for identifications is that the coefficients α, β, and γ and the variances of underlying shocks σ_x^2 and σ_η^2 are stable across the two sets of dates.

As shown by Rigobon and Sack [2004], these two estimators can be obtained from a simple instrumental variables (IV) procedure. To see this, define $\mathbf{\Delta i}$ and $\mathbf{\Delta s}$ as the $T \times 1$ vectors of stacked data corresponding to the two subsamples (i.e., $T = T_P + T_{\tilde{P}}$); let $\mathbf{d}_P$ denote the $T \times 1$

[22] For extensions of this identification approach see Sentana and Fiorentini [2001] and Klein and Vella [2010].

vector such that $d_{t,P} = 1$ if $t \in P$ and zero otherwise; and let $\mathbf{d}_{\tilde{P}} = \boldsymbol{\iota} - \mathbf{d}_P$, where $\boldsymbol{\iota}$ denotes the $T \times 1$ vector of ones. Define

$$\mathbf{z}_i = [\mathbf{d}_P \odot \Delta\mathbf{i} - \mathbf{d}_{\tilde{P}} \odot \Delta\mathbf{i}] \quad \text{and} \quad \mathbf{z}_s = [\mathbf{d}_P \odot \Delta\mathbf{s} - \mathbf{d}_{\tilde{P}} \odot \Delta\mathbf{s}],$$

where $\odot$ denotes the element-wise multiplication (i.e., the Hadamard product) of two vectors. Then the estimate of α_i can be obtained directly from an IV regression of $\Delta\mathbf{s}$ on $\Delta\mathbf{i}$ using $\mathbf{z}_i$ as an instrument, while an IV regression of $\Delta\mathbf{s}$ on $\Delta\mathbf{i}$ using $\mathbf{z}_s$ as an instrument yields an estimate of α_s:

$$\hat{\alpha}_i = (\mathbf{z}_i'\Delta\mathbf{i})^{-1}(\mathbf{z}_i'\Delta\mathbf{s}) \quad \text{and} \quad \hat{\alpha}_s = (\mathbf{z}_s'\Delta\mathbf{i})^{-1}(\mathbf{z}_s'\Delta\mathbf{s}).$$

In this so-called identification-through-heteroskedasticity approach, the response coefficients α_i and α_s can be estimated separately by 2SLS, but one can also impose a restriction that $\alpha_i = \alpha_s = \alpha$ and estimate α using GMM:

$$\hat{\alpha}_{GMM} = \arg\min_{\alpha} \Big[(\Delta\mathbf{s} - \alpha\Delta\mathbf{i})'\mathbf{Z}\Big] W_T \Big[\mathbf{Z}'(\Delta\mathbf{s} - \alpha\Delta\mathbf{i})\Big],$$

where $\mathbf{Z} = [\mathbf{z}_i \; \mathbf{z}_s]$ is the $T \times 2$ matrix of valid instruments, and W_T is an appropriately defined (data-dependent) weighting matrix. Because the coefficient α is over-identified, one can test the restriction that $\alpha_i = \alpha_s$. And lastly, in the case where $\Delta\mathbf{s}$ is a $T \times k$ matrix of asset prices—so that $\boldsymbol{\alpha} = [\alpha_1 \; \alpha_2 \; \ldots \; \alpha_k]'$ is a $k \times 1$ vector of policy response coefficients—it is natural to implement this approach within a system of equations estimated by GMM, where $\mathbf{Z} = [\mathbf{z}_i \; \mathbf{z}_{1,s} \; \mathbf{z}_{2,s} \; \ldots \; \mathbf{z}_{k,s}]$ is $T \times (1 + k)$ matrix of valid instruments.

To operationalize this approach, we must specify the P and $\tilde{P}$ subsamples. The subsample P naturally contains the nine LSAP announcement dates used in the event-style analysis. The subsample $\tilde{P}$ comprises the remainder of the Jan-02-2008–Dec-30-2011 sample period with the following exceptions: First, we eliminated all days associated with non-LSAP policy announcements that could have left an imprint in financial markets. These non-LSAP announcements include communication associated with the FOMC meetings, release of the FOMC minutes, and major speeches and Congressional testimonies by the FOMC participants. The exclusion of these days, most of which contain some indirect news about unconventional policy measures, serves to sharpen the distinction between the two covariance matrices that is crucial for identification.

And lastly, given that our sample period is characterized by an exceptional turmoil in financial markets, we also eliminated from the subsample $\tilde{P}$ a small number of dates associated with extreme changes in the two CDX indexes, a move designed to mitigate the effect of outliers on our estimates. Specifically, we dropped from the subsample $\tilde{P}$ all days where the change in either investment- or speculative-grade CDX index was below the 1st or above the 99th percentile of its respective distribution. All told, this procedure yielded 673 observations for the final subsample $\tilde{P}$.[23]

4.2 Results

In presenting our main results, we consider three estimators of α, the coefficient measuring the response of credit risk indicators to changes in the risk-free rates induced by the LSAP announcements: (1) **HET-1**: a 2SLS estimator of the coefficient α obtained from an IV regression of Δs_t on Δi_t, using $\mathbf{z}_i$ as an instrument (this corresponds to the estimator α_i above); (2) **HET-2**: a single-equation GMM estimator of α that uses both $\mathbf{z}_i$ and $\mathbf{z}_s$ as instruments (this corresponds to the estimator α_{GMM} above); and (3) **HET-3**: a system GMM estimator of the vector of coefficient $\boldsymbol{\alpha}$ that uses all valid instruments when analyzing the response of multiple credit risk indicators to the changes in the benchmark market interest rates. In both instances involving GMM estimation, the optimal weighting matrix W_T is obtained from a two-stage procedure that sets the weighting matrix equal to the identity matrix in the first stage and then computes the optimal weighting matrix based on the first-stage estimates.

The top panel of Table 4 reports the results for the case where Δi_t corresponds to change in the 5-year Treasury yield. Using the 2SLS estimator (HET-1), the estimates of α for both investment- and speculative-grade CDX indexes are positive but imprecisely estimated and hence are not statistically different from zero at conventional significance level. Although the standard errors are of the same order of magnitude across the three estimators, the point estimates of the response coefficient α increase notably as we move from a single-equation 2SLS estimation to either the single-equation or system GMM estimation methods. In particular, the system GMM estimator (HET-3) yields an estimate of the response coefficient for the investment-grade CDX index (CDX-IG) of 0.382 and an estimate of 1.285 for its speculative-grade counterpart (CDX-SG). In addition to being statistically significant, one cannot reject—using the Hansen [1982] J-test—the over-identifying restrictions for either the single-equation or system GMM estimators.

As shown in the bottom two panels, using the agency MBS rate or the agency bond yield in place of the 5-year Treasury yield to measure changes in risk-free interest rates produces very similar results. In all instances, the estimates of the coefficient α are positive, economically meaningful, and they become statistically highly significant with the use of more-efficient GMM methods. Again, as indicated by the p-values of the J-test, we do not reject the over-identifying restrictions in all the cases.

The results in Table 4 imply that the cost of insuring against broad-based incidence of defaults in the U.S. corporate sector declined significantly in response to a drop in yields on safe assets induced

[23]Note that in our notation, $T_P = 9$ and $T_{\tilde{P}} = 673$. Because the two samples are of unequal sizes, all endogenous variables and instruments in the P subsample are normalized by $\sqrt{T_P}$, while those in the $\tilde{P}$ subsample are normalized by $\sqrt{T_{\tilde{P}}}$. As a robustness check, we also considered subsamples $\tilde{P}$ comprising of a smaller number of days in periods immediately surrounding the nine LSAP announcements and found very similar results.

Table 4: LSAP Announcements, Interest Rates, and Corporate Credit Risk
(*Identification Through Heteroskedasticity*)

Interest Rate: Treasury (5y)

Credit Risk Indicator	HET-1	HET-2	HET-3
CDX-IG	0.157	0.351*	0.382**
	(0.215)	(0.185)	(0.153)
CDX-SG	0.361	1.023	1.285*
	(0.822)	(0.760)	(0.663)
$\Pr > W$[a]	.	.	0.004
$\Pr > J_T$[b]	.	0.185	0.725
	.	0.127	.

Interest Rate: Agency MBS[c]

Credit Risk Indicator	HET-1	HET-2	HET-3
CDX-IG	0.172	0.286***	0.388***
	(0.140)	(0.097)	(0.084)
CDX-SG	0.689	1.181***	1.530***
	(0.618)	(0.182)	(0.340)
$\Pr > W$[a]	.	.	0.000
$\Pr > J_T$[b]	.	0.203	0.683
	.	0.181	.

Interest Rate: L-T Agency[d]

Credit Risk Indicator	HET-1	HET-2	HET-3
CDX-IG	0.100	0.246***	0.265***
	(0.162)	(0.095)	(0.049)
CDX-SG	0.365	0.934**	1.033***
	(0.707)	(0.388)	(0.226)
$\Pr > W$[a]	.	.	0.000
$\Pr > J_T$[b]	.	0.173	0.635
	.	0.140	.

NOTE: Obs = 682. The dependent variable in each regression is the 1-day change in the specified credit risk indicator: CDX-IG = 5-year (on-the-run) investment-grade CDX index; and CDX-SG = 5-year (on-the-run) speculative-grade CDX index. Entries in the table denote the IV estimates of the coefficients associated with the 1-day change in the specified interest rate: HET-1 = single-equation 2SLS; HET-2 = single-equation GMM; and HET-3 = two-equation GMM system (see text for details). All specifications include a constant (not reported). Heteroskedasticity-consistent asymptotic standard errors are reported in parentheses: * $p < 0.10$; ** $p < 0.05$; and *** $p < 0.01$.

[a] p-value for the joint significance test of coefficients associated with interest rate changes in the two-equation GMM system.

[b] p-value for the Hansen [1982] J-test of the over-identifying restrictions. In the single-equation GMM estimation (HET-2), the top/bottom p-value corresponds to IG/SG CDX index.

[c] Fannie Mae 30-year current coupon MBS rate.

[d] Yield index on Fannie Mae's longer-term (maturity greater than 5 years) senior unsecured bonds.

Table 5: LSAP Announcements, Interest Rates, and Financial Sector Credit Risk
(*Identification Through Heteroskedasticity*)

Interest Rate: Treasury (5y)			
Credit Risk Indicator	HET-1	HET-2	HET-3
CDX-BK	−0.042	−0.009	−0.011
	(0.040)	(0.028)	(0.042)
CDX-BD	−0.080	−0.077	−0.051
	(0.057)	(0.058)	(0.108)
$\Pr > W$[a]	.	.	0.451
$\Pr > J_T$[b]	.	0.293	0.864
	.	0.549	.
Interest Rate: Agency MBS[c]			
Credit Risk Indicator	HET-1	HET-2	HET-3
CDX-BK	−0.021	0.002	−0.018
	(0.041)	(0.023)	(0.101)
CDX-BD	−0.076	−0.048	−0.084
	(0.081)	(0.053)	(0.214)
$\Pr > W$[a]	.	.	0.005
$\Pr > J_T$[b]	.	0.293	0.750
	.	0.527	.
Interest Rate: L-T Agency[d]			
Credit Risk Indicator	HET-1	HET-2	HET-3
CDX-BK	−0.028	−0.004	−0.012
	(0.036)	(0.021)	(0.044)
CDX-BD	−0.065	−0.054	−0.052
	(0.053)	(0.047)	(0.108)
$\Pr > W$[a]	.	.	0.522
$\Pr > J_T$[b]	.	0.291	0.829
	.	0.564	.

NOTE: Obs = 682. The dependent variable in each regression is the 1-day change in the specified credit risk indicator: CDS-BK = 5-year CDS index for 26 commercial banks; and CDS-BD = 5-year CDS index for 9 broker-dealers. Entries in the table denote the IV estimates of the coefficients associated with the 1-day change in the specified interest rate: HET-1 = single-equation 2SLS; HET-2 = single-equation GMM; and HET-3 = two-equation GMM system (see text for details). All specifications include a constant (not reported). Heteroskedasticity-consistent asymptotic standard errors are reported in parentheses: * $p < 0.10$; ** $p < 0.05$; and *** $p < 0.01$.

[a] p-value for the joint significance test of coefficients associated with interest rate changes in the two-equation GMM system.

[b] p-value for the Hansen [1982] J-test of the over-identifying restrictions. In the single-equation GMM estimation (HET-2), the top/bottom p-value corresponds to BK/BD CDS index.

[c] Fannie Mae 30-year current coupon MBS rate.

[d] Yield index on Fannie Mae's longer-term (maturity greater than 5 years) senior unsecured bonds.

by the LSAP announcements. Moreover, the estimates of the response coefficients that rely on the more-efficient GMM methods indicate that the decline in the speculative-grade CDX index is three to four times as large as in the investment-grade segment of the corporate sector. In principle, this sizable reduction in the cost of default insurance should be reflected in a substantially lower business borrowing costs, especially for riskier credits. According to our estimates, a reduction of 25 basis points in longer-term market interest rates induced by an LSAP announcement lowers the 5-year CDS premium for a typical junk-rated firm by about the same amount, translating into a drop of 50 basis points or more in the level of interest rates faced by such a firm.

These results are consistent with those of Krishnamurthy and Vissing-Jorgensen [2011], who find that the first asset purchase program (LSAP-I) significantly lowered CDS spreads on lower-rated corporate bonds. They are also consistent with the recent work of Wright [2012], who shows that monetary policy shocks had a significant—though fairly short lived—effect on corporate bond yields during the period when short-term rates were stuck at the zero lower bound. In general, the results in Table 4 comport with our earlier discussion, which argued that an OLS event-type estimator of changes in CDS spreads on the LSAP announcement indicators is likely to be biased downward, relative to an estimator that controls for the simultaneity between changes in the benchmark market interest rates and CDS spreads during the crisis period.

We now turn to the impact of the LSAP announcements on the market perception of credit risk in the financial sector. Table 5 summarizes the results from IV regressions, in which changes in the CDS spreads for our two types of financial intermediaries are regressed on changes in the benchmark interest rates. The striking feature of these results is that, regardless of the estimation procedure or the choice of the benchmark interest rate, all estimates of the structural response coefficient α are statistically indistinguishable from zero; moreover, the estimates are essentially zero in economic terms. Thus, in contrast to the response of broad, economy-wide indicators of corporate credit risk, the results in Table 5 indicate that the declines in risk-free rates induced by the LSAP announcements had no discernible effect on the CDS spreads of U.S. commercial banks or broker-dealers. In fact, our results imply that in a relative sense, the market views the financial intermediary sector as riskier as a result of LSAPs.

To the extent that loans extended by financial intermediaries, along with their other financial investments, should be less likely to default or deliver subpar returns when the broad-based indicators of corporate credit risk fall, these results appear puzzling. One possible explanation for these findings is that the profitability of the financial sector—the primary purpose of which is to perform maturity transformation—declines when longer-term interest rates fall relative to short-term interest rates. This interpretation is consistent with the recent work of English, Van den Heuvel, and Zakrajšek [2012], who document that the return on assets in the U.S. commercial banking sector drops sharply in response to the flattening of the Treasury yield curve, reflecting the ensuing compression of banks' net interest margins and deposit disintermediation that shrinks banks' balance sheets.

At the time when the financial sector is already facing significant capital and liquidity pressures, an LSAP-induced reduction in longer-term interest rates would put a further downward pressure on the sector's profitability, which may cause the CDS spreads of financial institutions to remain unchanged because the deterioration in their near-term creditworthiness outweighs the improvement in the economic outlook. An alternative possibility is that the various LSAP announcements reinforced the investors' perception that the government may impose significant losses on the holders of unsecured debt claims issued by the financial sector because LSAPs eliminated the extreme tail risk associated with the systemic financial crises. If market participants believed that the wholesale government bailout of the financial sector—which would have been more likely in the case of an extreme deterioration in economic conditions and in which all creditors would also likely be made whole—was less probable as a consequence of LSAPs, the financial sector CDS spreads might not fall, even as the LSAP announcements induce a decline in the broad market-based indicators of corporate credit risks

4.3 A Case Study of the Top 5 Financial Holding Companies

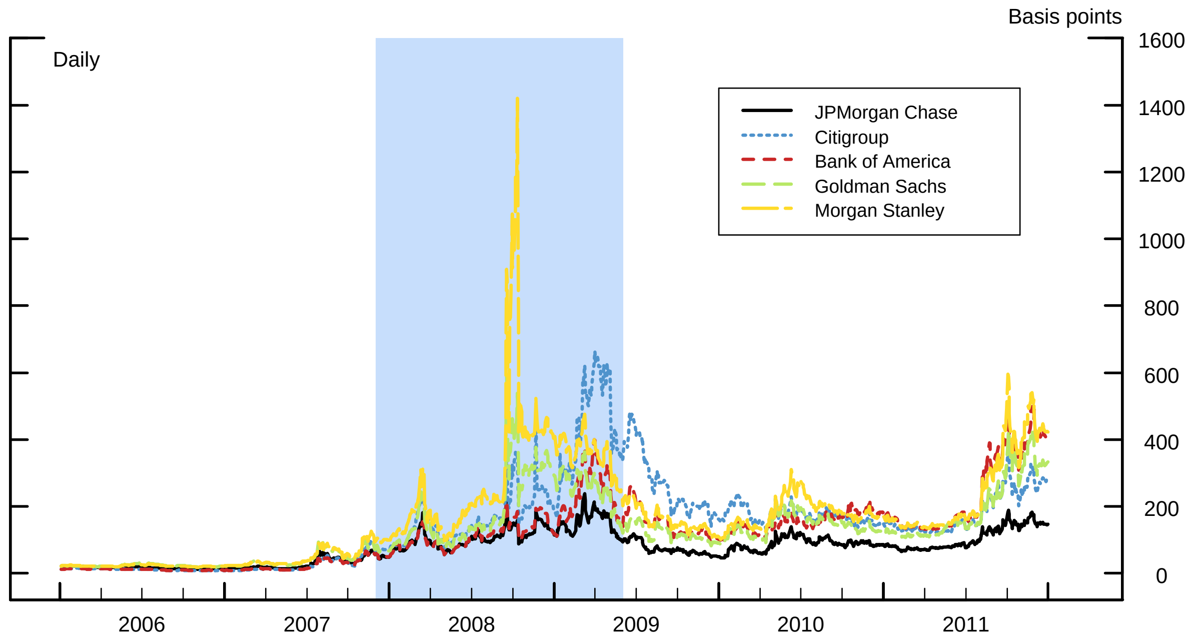
One potential critique of the above analysis is that the single-name CDS spreads of banks and broker-dealers in the two sectoral indexes are not as liquid as the components of the tradable CDX index. As a result, our credit risk indicators for the financial sector may not respond to the LSAP announcements in a sufficiently timely manner.[24] In addition, to the extent that credit risk in the U.S. financial sector during the crisis was concentrated at the largest institutions, the focus on the average change in bank or broker-dealer CDS spreads may not be very indicative of how LSAPs may have altered the market's perception of credit risk in the financial sector.

As a final exercise, therefore, we focus on the CDS spreads of the five largest and most prominent U.S. Financial Holding Companies (FHCs): JPMorgan Chase & Co., Citigroup Inc., Bank of America Corp., Goldman Sachs Group Inc., and Morgan Stanley. Reflecting their systemic importance, the financial health of these FHCs was of direct concern to both policymakers and market participants during the crisis. As a result, the CDS contracts written on these companies are highly liquid.[25] Through their commercial bank subsidiaries, these FHCs also engage in the traditional provision of credit to businesses and households, while at the same time pursuing nonbanking activities that offer customers a wide range of financial services, including the opportunity to invest in securities and, in some instances, to purchase insurance products; they also operate highly leveraged broker-dealer subsidiaries, which as argued above, play an important role in financial markets.

[24] As a simple check of this hypothesis, we re-did our event-style analysis using the properly constructed 2-day changes of the sectoral CDS spread indexes, but the results were essentially the same as those reported in Table 2.

[25] As before, our analysis is based on the CDS contracts with the MR clause. As a robustness check, we re-did the exercise using the CDS spreads based on the contracts with the XR clause, but that change had immaterial effects on the results.

Figure 3: CDS Spreads for the Top 5 Financial Holding Companies



NOTE: The figure depicts the 5-year CDS spreads for the five largest U.S. Financial Holding Companies. The shaded vertical bar represents the 2007–09 NBER-dated recession.

Figure 3 shows the CDS spreads for these systemically important global financial institutions. The investors' perception of credit risk associated with these five institutions clearly changed significantly with the onset of the financial crisis in the summer of 2007. Among them, the two former investment banks, Goldman Sachs and Morgan Stanley, had the most volatile CDS spreads, reflecting, in part, their business models that involved the use of high leverage and heavy reliance on short-term funding to engage in maturity transformation. Not surprisingly, there is a high degree of comovement in the CDS spreads of these five institutions; indeed, reflecting their common exposure to the macroeconomic risk factors, the first principal component explains about 75 percent of the variability in CDS spread changes over the 2008–11 sample period.

To examine formally the impact of the LSAP announcements of the CDS spreads of these institutions, we employ a system-GMM estimator, which allows us to estimate simultaneously the response of the institutions-specific CDS spreads to the LSAP-induced changes in market interest rates. In the estimation, we allow the response coefficient α to differ across the five FHCs. According to the entries in Table 6, the estimated response of the CDS spreads to a decline in the benchmark market interest rates prompted by the LSAP announcements is negative, economically meaningful and consistently statistically significant for only Citigroup and Bank of America; for the remaining three institutions in our sample, the estimates of the response coefficients are generally much smaller

Table 6: LSAP Announcements, Interest Rates, and Credit Risk at the Top 5 FHCs
(*Identification Through Heteroskedasticity*)

Financial Holding Company	Treasury	Agency MBS	L-T Agency
JPMorgan Chase	−0.131*	−0.054	−0.057
	(0.072)	(0.041)	(0.064)
Citigroup	−0.184***	−0.201**	−0.150***
	(0.035)	(0.043)	(0.044)
Bank of America	−0.102*	−0.198***	−0.106***
	(0.062)	(0.072)	(0.035)
Goldman Sachs	−0.140	−0.121*	−0.089
	(0.096)	(0.062)	(0.060)
Morgan Stanley	0.008	0.051	0.054
	(0.130)	(0.105)	(0.128)
$\Pr > W^{a}$	0.000	0.000	0.000
$\Pr > J_T{}^{b}$	0.009	0.002	0.003

NOTE: Obs = 682. The dependent variable in each regression is the 1-day change in the 5-year CDS spread of the specified financial holding company. Entries in the table denote the GMM estimates—from a five-equation system—of the coefficients associated with the 1-day change in the specified interest rate: Treasury = 5-year Treasury yield; Agency MBS = Fannie Mae 30-year current coupon MBS rate; and L-T Agency = yield index on Fannie Mae's longer-term (maturity greater than 5 years) senior unsecured bonds. All specifications include a constant (not reported). Heteroskedasticity-consistent asymptotic standard errors are reported in parentheses: * $p < 0.10$; ** $p < 0.05$; and *** $p < 0.01$.

[a] p-value for the joint significance test of coefficients associated with interest rate changes in the five-equation GMM system.

[b] p-value for the Hansen [1982] J-test of the over-identifying restrictions.

in economic terms and in most cases statistically indistinguishable from zero.

These results are broadly consistent with those in Table 5, which showed that the average CDS spreads in the financial intermediary sector did not react to the LSAP-induced reductions in market interest rates. The fact that the cost of insuring against the default of Citigroup and Bank of America rose in response to the LSAP announcements is likely due to the widespread market perception that these two institutions were particularly battered by the recent financial crisis, a perception buttressed by their inability to pass the Federal Reserve's stress tests during that period. To the extent that the financial sector's return to high and sustained profitability was likely to take a considerable amount of time in an environment characterized by a flat term structure and weak economic growth, it is plausible that the LSAP-induced declines in longer-term interest rates caused market participants to reassess the credit risk of these two specific institutions, especially given their relatively weak capital positions.

5 Conclusion

In this paper, we analyzed the impact of changes in the benchmark market interest rates prompted by the LSAP announcements on the market-based indicators of corporate credit risk. Importantly, we used the identification-through-heterogeneity approach advocated by Rigobon [2003] and Rigobon and Sack [2003, 2004] to correct for the simultaneity bias that plagues the standard event-style analysis. This approach, which allows us to identify more cleanly the structural response of CDS spreads to the LSAP-induced declines in market interest rates, indicates that the policy announcements led to a significant reduction in the cost of insuring against defaults for both investment- and speculative-grade corporate credits. In conjunction with the results of Hancock and Passmore [2011], who find that the Federal Reserve's purchases of agency MBS led to a significant reduction in residential mortgage rates, our results thus support the view that LSAPs induced a significant easing of financial conditions in both the household and business sectors.

While the unconventional policy measures employed by the Federal Reserve to stimulate the economy appeared to have lowered the overall level of credit risk in the economy, they had no measurable impact on credit risk specific to the financial sector. This apparently puzzling result could reflect the fact that the flattening of the yield curve engineered by LSAPs reduced the future profitability of financial institutions that intermediate funds across maturities, which outweighed the improvement in the economic outlook, leaving the CDS spreads of financial firms unchanged on balance.

Alternatively, the CDS spreads of financial institutions may have failed to decline (or even increased) because the announcements of the purchase programs led market participants to lower their perceived likelihood of wholesale bailouts of the financial sector, situations in which the bondholders would likely suffer only minimal losses. To the extent that LSAPs eliminated the extreme tail risk associated with the systemic financial crises, investors may have realized that the government is more likely to impose greater losses on the holders of unsecured debt claims issued by financial firms, a reassessment of risk that would have boosted the financial sector CDS spreads relative to broad market-based measures of corporate credit risk.

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