The Fed Takes on Corporate Credit Risk: An Analysis of the Efficacy of the SMCCF

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Abstract

We evaluate the efficacy of the Secondary Market Corporate Credit Facility (SMCCF), a program designed to stabilize the U.S. corporate bond market during the Covid-19 pandemic. We show that the program announcements on March 23 and April 9, 2020, significantly lowered credit and bid-ask spreads across the maturity spectrum and ultimately restored the upward-sloping term structure of credit spreads. Using intraday event study methodology, we also document that actual program purchases reduced credit spreads of eligible bonds by about two basis points more than those of ineligible bonds. To shed light on the underlying mechanism, we calibrate a variant of the preferred-habit model and show that a "dash for cash," a sell-off of shorter-term lowest-risk investment-grade bonds, combined with a spike in the arbitrageurs' risk aversion, can account for the inversion of the credit curve during the height of the pandemic-induced turmoil in the market. Consistent with the empirical findings, the Fed's announcements, by reducing risk aversion and alleviating market segmentation, helped restore the upward-sloping credit curve in the investment-grade segment of the market.

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1 Introduction

The Covid-19 shock in early 2020 severely strained the U.S. corporate bond market. The pandemic-induced "dash for cash" triggered a selloff in U.S. fixed income (and other) markets, with fixed-income mutual funds registering large outflows (see Ma et al., 2022). A number of structural factors exacerbated strains in the corporate bond market. Importantly, the structure of ownership in the market has changed significantly since the Global Financial Crisis (GFC), as holdings of corporate bond mutual funds have risen substantially over the past two decades (see Liang, 2020). Because investors in corporate bond mutual funds are offered daily liquidity—despite the fact that the underlying assets are significantly less liquid—the resulting "liquidity mismatch" made these funds especially vulnerable to runs (see Falato et al., 2021).

Equally important is the fact that since the GFC, intermediation in the market has remained concentrated in about a dozen or so primary dealers, most of whom are affiliated with major banks. The rapid growth of the U.S. Treasury market in the years before the pandemic has outstripped the intermediation capacity of these bank-affiliated dealers, which was already constrained by the post-GFC regulations (see Duffie, 2020). Unsurprisingly, as the risk-off sentiment swept through financial markets in March 2020, prices of corporate bonds fell and credit spreads increased sharply.[1]

The Fed reacted swiftly to the turmoil roiling financial markets, unveiling a broad array of measures to limit the economic damage from the pandemic (see Clarida et al., 2021). Although these actions averted a wider market meltdown, liquidity in the corporate bond market, which is limited in the best of circumstances, continued to deteriorate and credit spreads widened further. In response to these escalating strains, the Fed announced on March 23, 2020, what is arguably its most sweeping intervention in the economy to date: the creation of the Primary Market Corporate Credit Facility (PMCCF) and the Secondary Market Corporate Credit Facility (SMCCF).[2]

The announcement, characterized by market participants as "whatever it takes" and "throwing the kitchen sink" at the markets, had an immediate effect, significantly boosting stock prices, raising intermediate- and longer-dated Treasury yields, and compressing credit spreads. Nonetheless, conditions in the corporate bond market remained strained. In response, the Fed moved further into uncharted territory and on April 9 announced updated terms for the two corporate bond-buying facilities. In this additional move aimed at unfreezing corporate credit markets, the Fed indicated that P/SMCCF-eligible issuers now include companies recently downgraded from investment grade to "junk," the so-called fallen angels.

In this paper, we evaluate the efficacy of the SMCCF and analyze the mechanism through which it affected the corporate bond market. We focus on the SMCCF because of its historic importance—the first time the Fed directly supported corporate credit markets by signaling a willingness to purchase outstanding corporate debt and potentially take a material amount of credit risk on its

[1]As discussed by Schrimpf et al. (2020), large sales of U.S. Treasuries by some leveraged non-bank investors and foreign holders in early March further strained the balance sheet capacity of bank-affiliated dealers.

[2]The objective of the PMCCF was to support credit to businesses through the issuance of bonds and loans in the primary market. The SMCCF, by contrast, was established to provide liquidity to the market for outstanding corporate bonds. Both facilities were initially opened to the U.S. investment-grade companies only.

balance sheet. Understanding the efficacy of such programs and channels through which they affect broad financial conditions is thus critical for policy going forward.[3]

Formally evaluating the impact of the SMCCF on the corporate bond market is complicated by the fact that the Fed announced, expanded, and operated the SMCCF in conjunction with a number of other emergency measures.[4] Further complicating the matter is the fact that the relationship between investment-grade credit spreads and the bonds' remaining maturity—the so-called credit curve, which, in general, is upward sloping—inverted abruptly in early March 2020, with the long-short credit spread differential dropping deep into the negative territory. The inversion was especially pronounced in the safest segment of the investment-grade portion of the market, as investors amidst the panic first tried to liquidate their holdings of most liquid securities (i.e., shorter-maturity high-quality investment-grade bonds).

The pandemic-induced inversion of the investment-grade credit curve presents an important confounding factor when using a difference-in-differences (DiD) methodology that relies on the program's two key eligibility requirements—eligible bonds had to have an investment-grade credit rating and a remaining maturity of less than or equal to five years when purchased—to estimate the direct effects of the program on corporate bond prices. At the same time, it hints at a powerful channel through which announcements of such policies may affect credit markets in times of widespread financial distress. To control for this confounding effect, we augment the baseline DiD specification that relies on the program's two eligibility criteria with interaction terms, which allow the slope of the credit curve to rotate in the post-announcement window and across the program eligible and ineligible maturity segments of the curve.

Our empirical analysis of the SMCCF announcement effects offers three main takeaways. First, the March 23 and April 9 announcements significantly narrowed investment-grade credit and bid-ask spreads across the maturity spectrum—the "level" effect. Second, and more importantly, the two announcements significantly rotated the investment-grade credit curve—the "slope" effect—restoring the normal upward-sloping term structure of credit spreads. The March 23 announcement, in particular, reduced credit spreads on shorter-term program-eligible bonds relative to their longer-term ineligible counterparts. The April 9 announcement, in contrast, induced a steepening of the

[3]It is worth noting that the Bank of Japan, the Bank of England, and the European Central Bank have in the past launched similar corporate bond-buying programs in an effort to ease broad financial conditions and stimulate their economies. In fact, "credit-easing" programs are now a standard part of the toolkit used by central banks to deliver monetary stimulus when constrained by the effective lower bound on nominal interest rates. Given their ubiquitous use, Brunnermeier and Krishnamurthy (2020) develop a formal corporate finance framework to guide central banks' interventions in credit markets in response to shocks such as the Covid-19 pandemic.

[4]In announcing the establishment of its corporate bond purchase programs on March 23, 2020, the Fed also revived the Term Asset-Backed Securities Loan Facility and expanded its QE program—launched on March 15—to include purchases of commercial mortgage-backed securities in its mortgage-backed security purchases; at the same time, the Fed noted that it expects to announce shortly another emergency lending program—to be called the Main Street Business Lending Program—designed to support credit to small and medium-sized businesses. Further complicating matters is the fact that in the days leading to the March 23 announcement, the Fed revived the Commercial Paper Funding Facility (March 17), the Primary Dealer Credit Facility (March 17), and the Money Market Mutual Fund Liquidity Facility (March 18). Similarly, the expansion of the SMCCF to "fallen angels" announced on April 9 was accompanied by the establishment of the Municipal Liquidity Facility and the Paycheck Protection Program Liquidity Facility.

entire investment-grade credit curve, irrespective of the SMCCF's maturity-eligibility criterion. And third, our estimates indicate that the announcement-induced narrowing of credit spreads was due almost entirely to a reduction in credit risk premia as opposed to a reduction in default risk.

In combination, these findings have important implications for gauging the efficacy of the SMCFF and similar corporate bond-buying programs. The result that the March 23 announcement had a significant differential effect on the credit spreads of program-eligible bonds—even after controlling for the announcement-induced shifts in the credit curve—indicates that in designing such programs, eligibility criteria matter. They matter in the sense that they can be an important part of the causal mechanism through which announcements of such programs affect prices of eligible assets.

At the same time, the result that the April 9 announcement—through its impact on the slope of the entire investment-grade credit curve—led to a narrowing of credit spreads of both eligible and ineligible securities suggest a broader mechanism at work. As emphasized by Hanson et al. (2020), the establishment of such credit-support facilities provides investors with a valuable asymmetric put option that mitigates severe downside or tail risks, thereby reducing both the uncertainty and the associated risk premia in credit markets. Our results that the announcement-induced declines in credit spreads appear to mainly reflect reductions in credit risk premia and that the differential effect on program-eligible bonds in response to the April 9 announcement is fully subsumed by the slope effect are consistent with this interpretation.

Our last empirical exercise quantifies the effect of facility's actual purchases of individual corporate bonds on credit and bid-ask spreads. Using intraday transactions data that precisely identify the Fed's purchases of individual corporate bonds, we show that credit spreads of program-eligible bonds narrowed, on average, five basis points upon purchase, though about two basis points of this decline was reversed within five hours of the purchase. Over the same window, credit spreads of program-ineligible bonds narrowed about one basis point. Concentrated at the lower-end of the investment-grade quality spectrum (i.e., A/A and Baa/BBB rating categories), the differential of two basis points represents a sizable effect, given the modest size—by QE standards—of the Fed's actual purchases. All told, our empirical results speak to the extraordinary power of modern central banks: when markets have trust in the central bank's ability to deliver on its promise, as exemplified by the game-changing "whatever it takes" remark by Mario Draghi, the central bank needs to do less (if anything) to deliver on its promise.

In the last part of the paper, we synthesize our empirical findings through the lens of the preferred-habitat framework, as formalized recently by the influential work of Vayanos and Vila (2021). In this framework, risk-averse arbitrageurs integrate an otherwise segmented market owing to investor clienteles—the so-called preferred-habitat investors—who only demand bonds in a specific maturity sector. The extent to which the market is segmented in the maturity dimension depends on the arbitrageurs' risk aversion, in effect their effective risk-bearing capacity.

As shown by Vayanos and Vila (2021), when these arbitrageurs have ample risk-bearing capacity—that is, they are not very risk averse—a demand shock in a given maturity sector affects the entire term structure of interest rates through its impact on the duration of arbitrageurs' portfolios. However, when the arbitrageurs become more risk averse, they pull back from risk-taking, and the market becomes segmented in the maturity dimension due to limits to arbitrage. This segmentation gives rise to "localized" demand effects, whereby prices of securities in the maturity sector where the shock originated experience the largest changes.

We extend their version of the model that focuses on the Treasury bond market to include a parallel market for high-quality investment-grade corporate bonds. Using a calibrated version of this extended model, we look for a configuration of fundamental shocks that can reproduce movements in the model-implied credit curve around the two SMCCF announcements. We show quantitatively that the inversion of the credit curve at the onset of the pandemic can be explained by the combination of two empirically documented shocks: a negative demand shock for short-term high-quality investment-grade paper—the aforementioned the dash for cash—and a simultaneous jump in the arbitrageurs' risk aversion. In combination, these two shocks generate a localized effect concentrated at the short-end of the market that is sufficient to invert the credit curve as seen in the data.

We model the Fed's subsequent announcements as interventions that directly reduce the degree of arbitrageurs' risk aversion. To account for the estimated rotation of the credit curve in response to the March 23 announcement, our calculations imply that this announcement offset nearly three-quarters of the pandemic-induced jump in the arbitrageurs' risk aversion. The April 9 announcement further reduced risk aversion, ultimately restoring the credit curve to its pre-pandemic shape and level.

Relation to literature: Our paper contributes to literature on the pandemic-induced dislocations in the U.S. corporate bond market and the Fed's response to the crisis.[5] In this regard, Boyarchenko et al. (2022) analyze how the pandemic disrupted the ability of U.S. nonfinancial corporations to access primary credit markets. They document a marked, though gradual, improvement in primary bond market conditions following the Fed's March 23 announcement of the two corporate bond-buying programs and attribute the improvement to better benchmarking of primary market prices to secondary market prices of similar bonds and to increased willingness of broker-dealers to underwrite bond issuance. D'Amico et al. (2020) focus on the effects of the March 23 and April 9 announcements on the corporate bond exchange-traded funds (ETFs) and CDX indexes and document that the two announcements had a significant positive effect on the directly eligible ETFs, as well as on the ETFs

[5]The pandemic, of course, greatly affected functioning of other asset markets as well. Its impact on the U.S. Treasury market is analyzed in detail by Duffie (2020), Fleming and Ruela (2020), Schrimpf et al. (2020), He et al. (2021), Kruttli et al. (2021), and Vissing-Jorgensen (2021); Augustin et al. (2022) focus on non-U.S. government bonds markets, while Bahaj and Reis (2020) analyze the pandemic-induced strains in dollar funding markets; Gormsen and Koijen (2020) and Cox et al. (2020) study the impact of the Covid-19 shock on U.S. equity markets; and Bi and Marsh (2020), Li and Lu (2020), and Wei and Yue (2020) examine disruptions in the U.S. municipal bond market. For related research on the effects of credit easing programs launched by the Bank of England and the European Central Bank in 2016, see D'Amico and Kaminska (2019), Grosse-Rueschkamp et al. (2019), Adelino et al. (2020), and Todorov (2020). The impact of the Bank of Japan's corporate bond-purchase program launched in 2010 is analyzed by Suganuma and Ueno (2018).

holding eligible bonds and their close substitutes.

Nozawa and Qiu (2021) also analyze the reaction of corporate bond credit spreads to the Fed's March 23 and April 9 announcements. Their main findings stress the heterogeneous response of spreads across different subsamples of corporate bonds to both announcements, which they attribute to market segmentation, especially across credit ratings. Using a VAR-based variance decomposition approach of Nozawa (2017) to estimate the expected default loss and risk premium components of credit spreads, they find that about one-half of the decline in credit spreads at the aggregate level in response to the two announcements is attributable to a decline in the expected default loss component. The remaining half is accounted for by a decline in the aggregate credit risk premium, which they argue indicates that the default-risk channel of monetary policy played an important role during the Covid-19 pandemic.

While complementary in certain ways, our analysis yields important new and different insights into the mechanics of the SMCCF. On the empirical front, we show that the widely used DiD identification strategy underpinning the existing research on the causal effects of the SMCCF—an approach that does not control for the announcement-induced rotations of the credit curve—does not fully uncover the causal mechanism of how the Fed's announcements affected the pricing of securities in the secondary market. Compared with Nozawa and Qiu (2021), our alternative approach of decomposing credit spreads into the expected default and risk premium components assigns a much greater role to the risk-premium channel than to the default-risk channel. On this point, our use of a calibrated preferred-habitat model to quantify the relative importance of local demand shocks versus changes in the arbitrageurs' risk aversion offers an especially useful perspective on the two sets of results.

The way in which we introduce the pandemic-induced disruptions in the corporate bond market and the ensuing announcement effects into the preferred-habitat framework is informed importantly by the work of Haddad et al. (2021), O'Hara and Zhou (2021), and Kargar et al. (2021). Haddad et al. (2021) document that during the most acute period of market turmoil in early and mid-March, corporate bonds traded at a large discount to their corresponding credit default swap (CDS) contracts; moreover, this so-called bond-CDS basis widened most for bonds at the safer end of the credit quality spectrum. This result is consistent with a negative local demand shock at the short-end of the investment-grade credit curve, whereby investors in an effort to obtain cash first sold safe and more liquid corporate bonds, as opposed to more illiquid synthetic credit derivatives. In our model, the increase in the arbitrageurs' risk aversion—in effect, a reduction in their risk-bearing capacity—is consistent with O'Hara and Zhou (2021) and Kargar et al. (2021), who document the unwillingness of broker-dealers to absorb the selling pressure at the height of the crisis.

Beyond the pandemic-related disruptions in the corporate bond market, our paper also provides evidence in support of the preferred-habitat theories of the determination of the term structure of interest rates. As shown formally by Vayanos and Vila (2021), the preferred-habitat models feature localized demand effects that are key to understanding the impact of the central banks' QE programs on longer-term government bond yields, a point demonstrated empirically by D'Amico and King (2013).[6] Our contribution to this broader literature is to show that the pandemic-induced inversion of the investment-grade credit curve can be explained through the combination of a large negative shock to the preferred-habitat investors' demand for short-term high-quality investment-grade bonds, coupled with a sharp increase in the arbitrageurs' risk aversion. Consistent with Hanson et al. (2020), the Fed's announcements of the corporate bond-buying programs effectively reduced both the uncertainty and the associated risk premia in credit markets, in effect reducing the arbitrageurs' risk aversion and thereby alleviating market segmentation.

The road map for the remainder of the paper is as follows. Section 3 outlines our DiD methodology and presents the results, which quantify the impact of the March 23 and April 9 announcements on credit and bid-ask spreads; the section also contains the intraday-day event study, which quantifies the SMCCF's purchase effects. In Section 4, we augment the preferred-habitat model of the Treasury bond market due to Vayanos and Vila (2021) with a market for high-quality investment-grade corporate bonds. We use a calibrated version of the model to reproduce the observed movements in the investment-grade credit curve around the two announcements, providing a theoretical illustration of the mechanism of how such corporate bond-buying programs affect market prices. Section 5 concludes.

2 Overview of the Fed's Corporate Bond-Buying Programs

On March 23, 2020, the Fed announced an unprecedented corporate bond-buying program in response to severe strains in the U.S. corporate bond market. By establishing two emergency lending facilities pursuant to Section 13(3) of the Federal Reserve Act—the Primary Market Corporate Credit Facility and the Secondary Market Corporate Credit Facility—the Fed committed to buying a substantial amount of corporate debt in both the primary and secondary markets.[7]

Eligible bonds were required to have been issued by U.S. companies and had to have a remaining maturity of five years or less. The maximum amount of bonds that the SMCCF was allowed to purchase in the secondary market of any eligible issuer was capped at ten percent of the issuer's maximum dollar amount of bonds outstanding on any day between March 22, 2019, and March 22, 2020. The March 23 announcement stipulated that the two corporate bond-buying facilities were open to only investment-grade U.S. companies.

[6] See Culbertson (1957) and Modigliani and Sutch (1966, 1967) for early more informal treatments of the preferred-habitat view of the Treasury term structure.

[7] As discussed by Sastry (2018), Section 13(3) of the Federal Reserve Act, which was added to the act at the height of the Great Depression in 1932, granted the Fed enormous emergency lending powers. Notably, it granted the 12 Federal Reserve Banks the authority to "discount" for any "individual, partnership, or corporation" notes "endorsed or otherwise secured to the satisfaction of the Federal Reserve Bank[s]," subject to a determination by the Board of Governors of the Federal Reserve System of "unusual and exigent circumstances." While the Fed's aggressive use of Section 13(3) during the 2008–09 financial crisis successfully stabilized the financial system, the Congress responded to the Fed's use of Section 13(3) by narrowing that authority in the Dodd-Frank Act of 2010. Most importantly, any emergency lending must now be made through a "program or facility with broad-based eligibility," it cannot "aid a failing financial company" or "borrowers that are insolvent," and it cannot have "a purpose of assisting a single and specific company avoid bankruptcy." In addition, the Fed is prohibited from establishing a Section 13(3) program without the prior approval of the secretary of the Treasury.

Table 1: The Composition of the Initial Broad Market Listing

Sector	No. of issuers	Weight (%)	Issuer with the largest weight
Basic Industries	41	3.6	DuPont De Nemours
Capital Goods	70	7.4	General Electric
Communications	33	7.8	AT&T
Consumer Cyclical	73	16.2	Toyota Motor Credit
Consumer Non-Cyclical	101	20.4	AbbVie
Energy	78	9.5	BP Capital Markets America
Insurance	72	8.0	Met Life Global Funding
Nonbank Financials	41	2.1	Int. Lease Finance Corp.
REITs	56	3.2	WEA Finance
Technology	55	9.2	Apple
Transportation	18	2.6	Burlington North Santa Fe
Utilities	156	10.4	NextEra Energy Capital

Note: This table reports the sectoral composition of the initial Broad Market Listing, announced on June 28, 2020, and effective as of June 5, 2020. See the text for details.
Source: Authors' calculations using data from the Federal Reserve Bank of New York.

On April 9, 2020, the Fed announced that the PMCCF and SMCCF would support $500 billion of primary market purchases and $250 billion of secondary market purchases, respectively, backed by $75 billion provided by the Treasury Department using funding from the Coronavirus Aid, Relief, and Economic Security Act (CARES Act). In addition, the Fed expanded the two facilities to include certain fallen angels—companies that were rated at least Baa3/BBB- as of March 22, 2020, and were rated at least Ba3/BB- as of the date on which the Fed purchased their bonds. The SMCCF started buying corporate bond ETFs on May 12 and individual corporate bonds on June 16. On July 28, the Fed announced an extension of the two corporate bond-buying facilities—which were initially scheduled to expire on or around September 30, 2020—through December 31, 2020.[8]

The term sheet of the SMCCF stipulated that the facility's direct purchases of individual securities in the secondary market will attempt to track "a broad, diversified market index of U.S. corporate bonds." To operationalize this notion, the Federal Reserve Bank of New York published on June 28, 2020, the initial Broad Market Listing (BML), a set of corporate bonds eligible for purchase by the SMCCF.[9] To get a sense of what credits the facility was targeting, we report in Table 1 the composition of the initial BML. This first listing of eligible bonds, which went into effect on June 5, 2020, included securities issued by 794 U.S. companies in 12 broad sectors. The "Consumer Cyclical" and "Consumer Non-Cyclical" sectors had the largest weights of 16 percent and 20 percent, respectively. In the Consumer Cyclical sector, Toyota Motor Credit Corp. was the largest issuer, while AbbVie Inc., a biopharmaceutical company originated as a spinoff of Abbott Laboratories,

[8]The PMCCF was slated to commence purchases in the primary market on June 29, 2020, but during its operational phase did not execute a single transaction.

[9]The Federal Reserve Bank of New York published an updated Broad Market Listing roughly once a month through the remainder of the year.

was the largest issuer in the Consumer Non-Cyclical sector.[10]

3 The SMCCF Announcement and Purchase Effects

In this section, we use daily transaction-level bond data to quantify the how the March 23 and April 9 announcements affected the *level* of credit and bid-ask spreads across the program-eligible and ineligible bonds. To isolate and estimate the direct effects of the program on corporate bond prices and market liquidity measures, we begin the analysis by estimating the differential effects of the March 23 and April 9 announcements on program-eligible and ineligible bonds. Specifically, we use bond-level transactions data provided by the Trade Reporting and Compliance Engine (TRACE) to construct pairs of eligible and ineligible securities trading in the secondary market, where both types of securities were issued by the *same* company.[11] Using this DiD methodology—which allows us to control for industry characteristics, as well as firm-specific characteristics such as size, age, and the overall degree of credit risk exposure faced by the firm—we obtain a set of baseline results of how the two announcements affected trading conditions in the secondary market.

To better understand these announcement effects, we use the empirical framework of Gilchrist and Zakrajšek (2012) to decompose credit spreads into two components: a component capturing issuer-specific default risk and a residual component capturing credit risk premia or investor sentiment. Next, we delineate the key way through which the Fed's announcements affected pricing of investment-grade corporate securities in the secondary market. We do so by extending the baseline DiD analysis to also consider how the two announcements affected the *slope* of credit curve within the program-eligible and ineligible segments of the market. Lastly, we quantify the facility's actual purchases on credit and bid-ask spreads using an intraday event-study methodology.

3.1 Data Sources

The pricing data used in our analysis come from TRACE, a database containing information about individual corporate bond transactions in the secondary market. Most importantly, the TRACE database records the date and time of individual transactions, transaction prices and volumes, the direction of a transaction (buy or sell), as well as information about whether a transaction is "dealer-to-customer" or "dealer-to-dealer." After running the TRACE data through filters developed by Dick-Nielsen and Poulsen (2019), we combine the resulting security-level transactions data with the information from the Mergent's Fixed Income Securities Database to obtain bond characteristics, such as bond type, coupon frequency and payout dates, seniority, date and amount of issuance, maturity date, and credit ratings.

We restrict our TRACE sample to transactions involving senior unsecured bonds with fixed

[10]The subsequent Broad Market Listings had essentially the same sectoral composition.

[11]TRACE is the vehicle developed by the Financial Industry Regulatory Authority (FINRA) that facilitates the mandatory reporting of over-the-counter transactions in eligible fixed income securities. According to an SEC-approved set of rules, all broker-dealers who are FINRA member firms have an obligation to report transactions in TRACE-eligible securities.

coupon schedules that were issued by U.S. companies. From this sample, we drop all transactions involving bonds with a remaining maturity of less that one year or more than 12 years. These filters ensure that prices in our sample are not unduly influenced by the potential liquidity anomalies arising from the bond's special features, such as an impending redemption, unusually long maturity by the standards of fixed income markets, or changes in its promised cash flows.[12]

The daily price for each bond in our sample is defined as the last transaction price recorded between 9 a.m. and 4:00 p.m. on a given business day. We refer to the corresponding dollar amount traded as the transaction amount or transaction volume.[13] Following Gilchrist and Zakrajšek (2012), we construct a credit spread for each bond on each trading day as the difference between the bond's yield-to-maturity implied by its daily price and the yield-to-maturity of a synthetic risk-free security that mimics exactly the cash flows of the corresponding corporate bond. The yield of this synthetic risk-free security is calculated from its hypothetical price, which is equal to the present value of the promised cash flows, discounted by the term structure of zero-coupon U.S. Treasury yields, as estimated on that day by Gürkaynak et al. (2007).

To measure liquidity at the security level, we utilize information about the type of counterparties involved in each recorded transaction. Specifically, on each business day, we define the bond's "bid" price as an arithmetic average of all prices generated by transactions involving dealers buying that bond from a non-dealer customer. The bond's corresponding "ask" price, by contrast, is defined as an arithmetic average of all prices generated by transactions involving non-dealer customers buying that same bond from a dealer. Lastly, we define the bond's "mid" price as an arithmetic average of all prices involving dealer-to-dealer transactions. Our proxy for the bond-specific bid-ask spread is then calculated as the difference between the bond's ask and bid prices, divided by the mid price.

In Panel A of Table 2, we report summary statistics for selected bond characteristics using a subsample of investment-grade bonds for which we constructed credit spreads; the corresponding statistics for the subsample of investment-grade bonds for which we were able to construct bid-ask spreads are reported in Panel B.[14] In each case, we focus on two sample periods: a pandemic period running from January through the end of July of 2020 and a comparable pre-pandemic period in 2019. According to Panel A, the average credit spread in our sample of bonds was about 100 basis points before the pandemic but shot up to almost 160 basis points over the first seven months of 2020. In general, the Covid-19 shock shifted the entire distribution of credit spreads notably to the right and significantly increased the dispersion of credit spreads in our sample.

[12]In fact, a vast majority of bonds purchased by the SMCCF were senior unsecured bonds with fixed coupon schedules; restricting our sample to fixed-coupon bonds thus facilitates comparisons with the sample of bonds purchased by the facility.

[13]As a robustness check, we also defined the daily price for each bond as a weighted average of all of its transaction prices between 9:00 a.m. and 4:00 p.m. on a given day, with weights equal to the corresponding transaction amounts. Using this alternative definition had a negligible effect on all the results reported in the paper.

[14]For the purposes of analyzing the efficacy of the SMCCF, investment-grade bonds are the relevant segment of the U.S. corporate bond market. To ensure that our results are not unduly influenced by a small number of extreme observations, we drop from the credit spread sample all observations with credit spreads of less than one basis point or with credit spreads exceeding 2,000 basis points. From the bid-ask spread sample, we drop all observations with bid-ask spreads of less than one basis point or with bid-ask spreads exceeding 500 basis points.

Table 2: Summary Statistics of Selected Bond Characteristics

	Pre-Pandemic Period					Pandemic Period				
Variable	Mean	SD	P25	P50	P75	Mean	SD	P25	P50	P75
A. Sample of credit spreads[a]										
Credit spread (bps.)	99.4	63.2	56.1	84.7	128.0	158.8	152.9	61.4	109.3	204.9
Time-to-maturity (years)	4.8	2.5	2.7	4.4	6.8	4.5	2.6	2.3	4.1	6.5
Age (years)	4.3	4.2	1.7	3.4	5.9	4.6	4.4	1.8	3.7	6.3
Coupon rate (pct.)	3.7	1.1	3.0	3.5	4.1	3.6	1.1	2.9	3.5	4.0
Par amount ($ millions)	846.4	776.1	400.0	650.0	1000.0	845.2	766.3	420.0	650.0	1000.0
B. Sample of bid-ask spreads[b]										
Bid-ask spread (bps.)	45.4	48.1	16.1	29.4	55.6	61.1	75.6	15.5	32.2	73.9
Time-to-maturity (years)	4.6	2.5	2.6	4.2	6.5	4.3	2.6	2.1	3.9	6.2
Age (years)	4.0	3.5	1.8	3.3	5.6	4.4	3.7	1.9	3.8	6.0
Coupon rate (pct.)	3.5	1.0	2.9	3.4	4.0	3.5	1.0	2.8	3.4	4.0
Par amount ($ millions)	1129.1	936.7	500.0	990.5	1500.0	1106.8	926.5	500.0	850.0	1348.4

Note: The table reports summary statistics of selected bond characteristics for our sample of outstanding investment-grade corporate bonds issued by U.S. companies. Summary statistics are reported for two non-overlapping sample periods of equal length: the pre-pandemic period (Jan–Jul 2019) and the pandemic period (Jan–Jul 2020).

[a] The sample of credit spreads corresponds to a set of investment-grade bonds for which we could compute daily credit spreads (see the text for details). Panel dimensions (Jan–Jul 2019): No. of bonds = 4,996; No. of firms = 957; and Observations =429,547. Panel dimensions (Jan–Jul 2020): No. of bonds = 5,453; No. of firms = 975; and Observations = 451,917.

[b] The sample of bid-ask spreads corresponds to a set of bonds investment-grade for which we could compute daily bid-ask spreads (see the text for details). Panel dimensions (Jan–Jul 2019): No. of bonds = 4,425; No. of firms = 912; and Observations = 209,829. Panel dimensions (Jan–Jul 2020): No. of bonds = 4,908; No. of firms = 922; and Observations = 218,917.

As shown in Panel B, a similar, though less pronounced, shift also occurred in the distribution of bid-ask spreads. The more muted response of bid-ask spreads owes importantly to the fact that the sample of bonds for which we are able to calculate bid-ask spreads is by construction smaller than the sample of bonds for which we can compute credit spreads.[15] Note that the par values of bonds in the former sample are systematically larger than the par values of bonds in the latter sample, as this sample of bonds by construction includes securities that trade more frequently and thus are more liquid. Despite these differences, the remaining bond characteristics are very similar across the two samples.

3.2 Empirical Methodology and Baseline Results

Our first pass at quantifying the effects of the March 23 and April 9 announcements on the corporate bond market involves estimating

$$Y_{i,j,t} = \beta_1 \mathbb{1}[t \geq t^*] + \beta_2 \big(\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]\big) + \theta' \mathbf{X}_{i,j,t} + \eta_i + \epsilon_{i,j,t}, \tag{1}$$

where i indexes issuers, j indexes their outstanding bonds, and t indexes business days. In this specification, $Y_{i,j,t}$ denotes the outcome variable of interest, either a credit spread ($\mathrm{CS}_{i,j,t}$) or the log of a bid-ask spread ($\ln \mathrm{BAS}_{i,j,t}$) on bond j (a liability of issuer i) on business day t.[16] Turning to the key explanatory variables, the 0/1-indicator $\mathbb{1}[t \geq t^*]$ equals one if the date t is greater than or equal to the specified announcement date t^*, either March 23 or April 9. The 0/1-indicator $\mathbb{1}[j \in E]$ equals one if bond j was eligible for purchase by the SMCCF.

In this canonical DiD specification, the effects of the March 23 and April 9 announcements on conditions in the corporate bond market are gauged through the program's key eligibility criterion—the five-year maturity cutoff captured by the 0/1-indicator $\mathbb{1}[j = E]$. By exploiting the fact many U.S. investment-grade companies had outstanding bonds with maturity of less than or equal to five years—which were eligible for purchase by the SMCCF—as well as outstanding bonds with maturity greater than five years that were ineligible, we can, in principle, identify the causal impact of the two announcements on the outcome variable of interest. In that case, the coefficient β_2 on the interaction term $\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$ quantifies the difference in the specified outcome variable between the program-eligible and ineligible bonds of the same issuer in response to the specified announcement, a natural metric by which to judge the efficacy of the program. To gauge the persistence of these announcement effects, we estimate specification (1) using symmetric two-, five-, and ten-day windows bracketing the March 23 and April 9 announcements.

Specification (1) also includes a vector of covariates, denoted by $\mathbf{X}_{i,j,t}$, comprising of pre-

[15] Recall that to construct bid-ask spreads, we require a minimum of three distinct transactions on each day: (*i*) a sale of the bond by a dealer to a non-dealer customer; (*ii*) a sale of the same bond by a non-dealer customer to a dealer; and (*iii*) a sale of the same bond between two dealers. As a result, the sample of bid-ask spreads will be smaller than the corresponding sample of credit spreads, as the construction of the latter requires only a single daily transaction.

[16] Taking logs of bid-ask spreads provides a useful transformation to control for heteroskedasticity, given that the distribution of bid-ask spreads is highly skewed.

determined bond characteristics that can affect the level of credit or bid-ask spreads. These include the bond's (fixed) coupon rate, its remaining maturity, age, and the log of par value, as well as 0/1-indicator variables for whether the bond is callable, has credit enhancements, or is subject to covenants. The vector $\mathbf{X}_{i,j,t}$ also includes the indicator variable $\mathbb{1}[j = E]$, which controls for common factors affecting SMCCF-eligible bonds across the pre- and post-announcement segments of each estimation window. Issuer fixed effect η_i captures all (time-invariant) unobservable issuer characteristics, in effect differencing out all issuer-specific effects of policy announcements within each estimation window.

3.2.1 "Treatment" and "Control" Groups

To implement the above approach, we use our TRACE data set to construct an issuer-matched sample of program-eligible (i.e., treated) and ineligible (i.e., control) bonds. Specifically, for the treated group, we select all outstanding bonds that satisfy the following two conditions: (*i*) their issuers had an investment-grade rating as of March 22, and they maintained that rating during the specified post-announcement window; and (*ii*) the bonds' remaining maturity as of the March 23 announcement was less than or equal to five years. This sample of bonds was eligible for the purchase by the SMCCF as of the March 23 announcement. For each bond in this sample, we then identify all outstanding bonds issued by the same company, but whose remaining maturity is greater than five years; this second sample of bonds was not eligible to be purchased by the SMCCF and constitutes the control group.

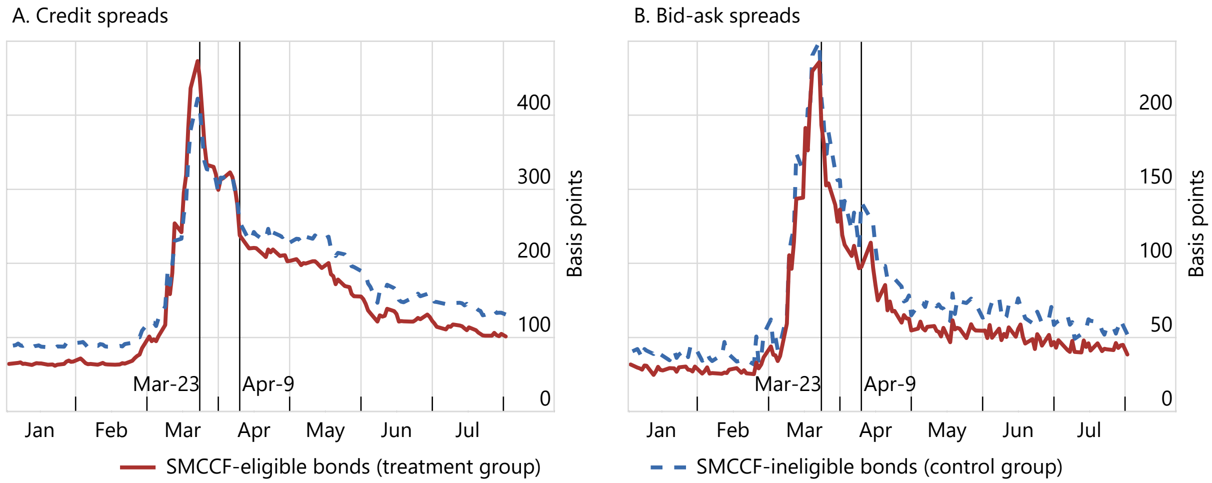
Using these two samples of bonds, we construct the *narrow* treatment and control groups as follows. For the narrow treatment group, we select from issuers with multiple bonds in the eligible sample, a bond with the remaining maturity closest to five years. Analogously, if there are multiple bonds in the ineligible sample that can be paired with the bond in the narrow treatment group, we keep only the bond with the remaining maturity closest to five years—these bonds make up the narrow control group. This selection procedure yielded 3,225 pairs of bonds, which between January and August 2022 were outstanding liabilities of 545 U.S. investment-grade companies. For the rest of the paper, we refer to this as the "narrow" sample.

One advantage of the narrow sample is that pairs of the program-eligible and ineligible bonds are as close as possible in terms of their remaining maturities and thus are to large extent subject to the same underlying default risk. That is, if issuer-specific default risk is roughly constant around the five-year horizon, it would be absorbed by issuer fixed effects in specification (1), sharpening the identification of the SMCFF-specific announcement effects on conditions in the corporate bond market.[17]

The red line in Panel A of Figure 1 shows the daily average credit spread of bonds in the narrow treatment group, while the blue line shows the corresponding average credit spread in the narrow control group. The red and blue lines in Panel B show the evolution of the respective average bid-

[17] The average (median) difference in the remaining maturities across pairs of bonds is 2.7 (2.3) years, while the 5th and 95th percentiles are 0.5 6 years, respectively.

Figure 1: SMCCF-Eligible vs. SMCCF-Ineligible Corporate Bonds



Note: The red solid line in Panel A shows the daily average credit spread of SMCCF-eligible corporate bonds (i.e., the narrow treatment group), while the dashed blue line shows the daily average credit spread of SMCCF-ineligible corporate bonds (i.e., the narrow control group). The corresponding lines in Panel B show the daily average bid-ask spreads for the same two groups of bonds. See the text for details regarding the construction of the narrow treatment and control groups. Vertical lines at specified dates: Mar-23 = Fed announces the establishment of the P/SMCCF; and Apr-9 = Fed expands the facilities to include corporate bonds of issuers that were rated investment grade as of March 22 but were subsequently downgraded to junk.
Source: Authors' calculations using TRACE data.

ask spreads. Before the realization of the potential economic impact of the Covid-19 shock rattled investor confidence in late February, the average credit spread in the control group was consistently above that in the treatment group. The gap between the spreads in the two samples was very stable around the average of about 25 basis points. This pattern is consistent with the fact that the average bid-ask spread in the treated group was systematically below the average bid-ask spread in the control group during this period. A likely interpretation is that bonds in the treated group were, on average, more liquid than their counterparts in the control group due to their shorter maturity and possibly other bond or issuer characteristics.

In early March, when fears over the impact of the Covid-19 outbreak sparked a broad sell-off in risky assets, the gap between the two credit spread series started to close, disappearing completely during the bout of turmoil that swept through financial markets in mid-March. This acute risk-off period also saw a widespread deterioration in market liquidity, as the average bid-ask spreads in both samples shot up and converged at elevated levels. Following the Fed's March 23 announcement, credit spreads in both the treatment and control groups declined significantly. Interestingly, the size of the drop in the average credit spreads in the immediate aftermath of the announcement was virtually the same across the two groups.

The commensurate drop in credit spreads across the two groups in the wake of the March 23 announcement would suggest that what caused the spreads to narrow was not the announcement of the corporate bond-buying program per se. Rather, it was the Fed's "whatever it takes" pledge to keep the economy from collapsing under the weight of the Covid-19 pandemic, reflected in the opening sentence of the announcement, which stated that the Fed is “committed to using its full range of tools to support households, businesses and the U.S. economy overall.” This interpretation is consistent with the decline in the average bid-ask spread in both samples in the days following the March 23 announcement, an indication that this extraordinary announcement significantly improved the overall functioning of the corporate bond market.

The Fed’s April 9 announcement, by contrast, appears to have had a more differential effect on credit spreads in the treatment and control groups. In particular, the average credit spread in the treated group fell more that the average credit spread in the control group. At first glance, this would suggest that the April 9 follow-up announcement had a distinct impact on the corporate bond market. At the same time, the April 9 announcement appears to have had no differential effect on the average bid-ask spreads in the two groups of corporate bonds.

3.2.2 Baseline Estimates of the Announcement Effects

The formal results of how the two announcements affected conditions in the U.S. corporate bond market are reported in Tables 3 and 4. Consider the impact on credit spreads as reported in Panel A of Table 3. It is instructive to first examine the estimates of β_1, the coefficient on the indicator $\mathbb{1}[t \geq t^*]$, which measures the change in the average credit spread between the pre- and post-announcement segments of the windows bracketing each announcement—the *level* effect. In response to the March 23 announcement (columns 1–3), credit spreads narrowed, on average, about 30 basis points within the first two days following the announcement. Within five days, however, this effect has fully dissipated, and within ten days of the announcement, spreads were, on average, 55 basis points higher than they were over the ten days before the announcement. This reversal in credit conditions is indicative of the turmoil that roiled the corporate bond market in late March as the news of the pandemic and associated policy responses unfolded.

The April 9 announcement (columns 4–6), by contrast, led to a clear improvement in credit conditions, as evidenced by the steady narrowing of the average credit spread in the post-announcement windows relative to the pre-announcement windows. The estimates of coefficient β_1 imply that the average credit spread fell more than 50 basis points in the two-day window and more than 70 basis points in the ten days following the April 9 announcement. These estimates serve as useful benchmarks when assessing the additional impact of the SMCCF announcements through the maturity-eligibility criterion.

Turning to these estimates, the entries in columns (1)–(3) indicate that the March 23 announcement induced a significant—in both statistical and economic terms—narrowing of credit spreads on SMCCF-eligible bonds compared with their ineligible counterparts. Within the two- and five-day windows bracketing the announcement, the estimates of β_2—the coefficient on the interaction term $\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$—imply an additional narrowing of credit spreads of 26 and 23 basis points, respectively, on SMCCF-eligible bonds relative to ineligible bonds issued by the same set of companies. Consistent with the finding that the March 23 announcement led to only a short-lived

Table 3: Effect of the SMCCF Announcements on Credit Spreads
(Narrow Treatment and Control Groups)

	Mar-23 Announcement			Apr-9 Announcement		
	2-day	5-day	10-day	2-day	5-day	10-day
Explanatory Variable	(1)	(2)	(3)	(4)	(5)	(6)
A. Without time fixed effects						
$\mathbb{1}[t \geq t^*]$	-0.31^{***}	0.05	0.55^{***}	-0.52^{***}	-0.69^{***}	-0.72^{***}
	(0.04)	(0.03)	(0.04)	(0.02)	(0.02)	(0.02)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	-0.26^{***}	-0.23^{***}	-0.08^{**}	-0.11^{***}	-0.17^{***}	-0.24^{***}
	(0.07)	(0.04)	(0.03)	(0.03)	(0.03)	(0.03)
R^2	0.76	0.70	0.65	0.91	0.90	0.89
No. of firms	487	523	544	513	537	552
No. of bonds	1,395	1,812	2,181	1,477	1,813	2,146
Observations	3,934	8,656	16,466	4,174	9,106	17,316
B. With time fixed effects						
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	-0.25^{***}	-0.21^{***}	-0.09^{***}	-0.06^{**}	-0.13^{***}	-0.20^{***}
	(0.06)	(0.04)	(0.03)	(0.03)	(0.03)	(0.03)
R^2	0.80	0.79	0.79	0.92	0.91	0.90
No. of firms	479	513	533	498	527	546
No. of bonds	1,271	1,606	1,892	1,331	1,604	1,854
Observations	3,803	8,350	15,873	4,004	8,760	16,670

Note: The dependent variable in all specifications is $\mathrm{CS}_{i,j,t}$, the credit spread of bond j (issued by firm i) on business day t. The entries in the table denote the OLS estimates of coefficients associated with the specified explanatory variable: $\mathbb{1}[t \geq t^*]$ = 0/1-indicator that equals one if date t is greater than or equal to the specified announcement date t^* and zero otherwise; and $\mathbb{1}[j = E]$ = 0/1-indicator that equals one if bond j was eligible for purchase by the SMCCF as of March 22 and zero otherwise. All specifications include a vector of bond-specific controls (not reported) and issuer fixed effects. Asymptotic standard errors reported in parentheses are clustered at the issuer level: * $p < .10$; ** $p < .05$; and *** $p < .01$.

Table 4: Effect of the SMCCF Announcements on Bid-Ask Spreads
(Narrow Treatment and Control Groups)

	Mar-23 Announcement			Apr-9 Announcement		
	2-day	5-day	10-day	2-day	5-day	10-day
Explanatory Variable	(1)	(2)	(3)	(4)	(5)	(6)
A. Without time fixed effects						
$\mathbb{1}[t \geq t^*]$	−0.29***	−0.25***	−0.12***	0.02	−0.20***	−0.46***
	(0.06)	(0.04)	(0.03)	(0.05)	(0.03)	(0.03)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	−0.11	−0.14**	−0.10**	0.00	−0.00	−0.05
	(0.09)	(0.06)	(0.04)	(0.07)	(0.05)	(0.04)
R^2	0.11	0.12	0.12	0.31	0.31	0.33
No. of firms	310	359	395	337	381	423
No. of bonds	985	1,426	1,855	1,089	1,527	1,950
Observations	1,888	4,134	8,052	2,150	4,904	9,468
B. With time fixed effects						
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	−0.09	−0.15***	−0.11***	0.03	0.02	−0.02
	(0.09)	(0.06)	(0.04)	(0.07)	(0.05)	(0.04)
R^2	0.14	0.16	0.22	0.33	0.34	0.38
No. of firms	288	338	383	316	365	410
No. of bonds	870	1,228	1,596	962	1,328	1,665
Observations	1,741	3,815	7,498	1,988	4,559	8,810

Note: The dependent variable in all specifications is $\ln \text{BAS}_{i,j,t}$, log of the bid-ask spread of bond j (issued by firm i) on business day t. The entries in the table denote the OLS estimates of coefficients associated with the specified explanatory variable: $\mathbb{1}[t \geq t^*] = 0/1$-indicator that equals one if date t is greater than or equal to the specified announcement date t^* and zero otherwise; and $\mathbb{1}[j = E] = 0/1$-indicator that equals one if bond j was eligible for purchase by the SMCCF as of March 22 and zero otherwise. All specifications include a vector of bond-specific controls (not reported) and issuer fixed effects. Asymptotic standard errors reported in parentheses are clustered at the issuer level: * $p < .10$; ** $p < .05$; and *** $p < .01$.

improvement in overall credit conditions, the estimated announcement effect for the eligible bonds shrinks to eight basis points in the ten-day window, though it remains statistically significant at conventional levels.

Whereas the March 23 announcement effects on credit spreads are estimated to dissipate over time, we see the opposite pattern in response to the April 9 announcement. As shown in columns (4)–(6), the estimated announcement effects for SMCCF-eligible bonds increase (in absolute value) with the window length. In the two-day window, the April 9 announcement induced an additional decline in credit spreads on the eligible bonds of 11 basis points, which increased to 17 basis points in the five-day window and to 24 basis points in the ten-day window. These results indicate that the April 9 announcement had a much more lasting effect on the level of credit spreads than the March 23 announcement, a finding consistent with the persistent gap between the red and blue lines in the left panel of Figure 1 that emerged after April 9.

The April 9 differential announcement effects are estimated quite precisely—the standard error is a mere three basis points in all three estimation windows. Moreover, these estimates are quite large in economic terms compared with the overall level effect of the announcement. In the ten-day window, for instance, the April 9 announcement induced a decline of 72 basis points across all credit spreads and an additional decline of 24 basis points in spreads on SMCCF-eligible bonds.

As documented by Baker et al. (2020), financial markets in March and early April were buffeted by a cascade of news about mandatory business closures and other restrictions on commercial activity aimed to slow or contain the pandemic. To control for the myriad of common shocks during the estimation windows, including news about actual or prospective fiscal and monetary policy actions, Panel B reports the estimation results of the baseline specification (1) augmented with a full set of time fixed effects.[18] Note that the estimates of β_2—the coefficient on the interaction term $\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$—are very similar across the two specifications. The robustness of these estimates indicates that reactions in financial markets to news about the course of the pandemic and associated policy responses are unlikely to be confounding the identification of the SMCCF-specific announcement effects.

In Table 4, we report the results from analogous exercise using the log of bid-ask spreads as the dependent variable, under both the baseline specification (Panel A) and the alternative specification that includes time fixed effects (Panel B). The log transformation of bid-ask spreads implies that the estimated announcement effects reported in the table are expressed in *percentage* changes in bid-ask spreads. To convert them back to original units (basis points), we multiply the relevant coefficients by the sample mean of bid-ask spreads in the specified window.

According to the estimates of β_1, the coefficient on the announcement indicator $\mathbb{1}[t \geq t^*]$, the March 23 announcement (columns 1–3 in Panel A) significantly improved liquidity conditions in the corporate bond market. On average, bid-ask spreads fell almost 60 basis points within two days of the announcement, and while some of that decline was subsequently reversed, we still observe a decline of nearly 25 basis points in the average bid-ask spread within the ten-day window. The

[18]Note that in this case, the coefficient β_1 on the indicator $\mathbb{1}[t \geq t^*]$ is not separately identified.

improvement in liquidity is even more pronounced after the April 9 announcement (columns 4–6), as the average bid-ask spread fell about 50 basis points within the ten-day window. Note that these announcement effects are all highly statistically significant.

While the two announcements significantly improved overall liquidity conditions in the market, the effect on SMCCF-eligible bonds is limited to the March 23 announcement. According to our estimates, the March 23 announcement compressed bid-ask spreads of eligible bonds by an additional 25 basis points or so within the five-day window, with the effect diminishing to about 20 basis points in the ten-day window. Consistent with the market commentary that the April 9 announcement was focused more on credit risk as opposed to broad liquidity concerns in financial markets, the estimated effect on bid-ask spreads for SMCCF-eligible bonds is essentially zero in all windows following the announcement. Lastly, it is worth pointing out that standard errors associated with the estimated announcement effects for bid-ask spreads tend to be somewhat larger than their counterparts in the credit-spread regressions. As discussed above, this partly reflects smaller sample sizes used in the estimation of the bid-ask-spread regressions.

All told, our baseline results indicate that the March 23 and April 9 announcements significantly lowered the average level of credit and bid-ask spreads in the investment-grade segment of the U.S. corporate bond market, a finding in line with the literature (see Nozawa and Qiu, 2021; Boyarchenko et al., 2022). The March 23 announcement also led to an economically significant narrowing of both credit and bid-ask spreads on program-eligible bonds, while the effect of the April 9 announcement was for the most part concentrated on credit spreads of such bonds. Across the two announcements, our ten-day window estimates imply a total decline of 45 basis points in credit spreads and roughly a 25 basis point reduction in bid-ask spreads for SMCCF-eligible bonds relative to their ineligible counterparts. By any stretch of the imagination, these are sizable program-specific effects, especially since the Fed has yet to purchase a single corporate bond in that time frame.[19]

3.2.3 Announcement Effects on the "Fallen Angels"

Unlike the "whatever it takes" message implied by the March 23 announcement, the April 9 announcement primarily clarified a number of key aspects of the Fed's corporate bond-buying program. Most importantly, the Fed extended the facility to certain "fallen angels," in effect signaling to investors that it may be willing to support the speculative-grade segment of the corporate bond market and thus take on a potentially significant amount of credit risk on its balance sheet. In this section, we zero in on this aspect of the April 9 announcement and examine its impact on the fallen angels' credit and bid-ask spreads.

According to the SMCCF's term sheet, an eligible fallen angel is a U.S. company that had an

[19] As a robustness check, Appendix A contains estimation results of specification (1) based on the *full-sample* definition of treatment and control groups. The full-sample treatment and control groups are constructed by assigning—based on the five-year maturity cutoff—all outstanding investment-grade bonds with remaining maturity between one and 12 years into the relevant group; that is, they relax the assumption that an SMCCF-eligible company must have a pair of outstanding bonds on either side of the five-year maturity cutoff. As shown in Tables A-1 and A-2, these full-sample estimates of the announcement level effects are quite comparable to their corresponding estimates based on the narrow definitions of the two groups reported in Tables 3 and 4.

investment-grade credit rating as of March 22 but was subsequently downgraded to junk.[20] To gauge the effects of the two announcements on eligible fallen angels, we focus on companies rated investment grade as of March 22, but which were downgraded to the eligible fallen-angel category between March 23 and April 9. We identified 14 such companies, all of which were downgraded to junk within a couple of days of the March 23 announcement.

For these 14 issuers, we add pairs of their outstanding bonds that are closest to the five-year maturity cutoff to our narrow treatment and control groups and estimate

$$\begin{aligned} Y_{i,j,t} &= \beta_1 \mathbb{1}[t \geq t^*] + \beta_2\big(\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]\big) \\ &\quad + \beta_3\big(\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E] \times \mathbb{1}[i = \mathrm{FA}]\big) + \theta' \mathbf{X}_{i,j,t} + \eta_i + \epsilon_{i,j,t}, \end{aligned} \tag{2}$$

where $\mathbb{1}[i = \mathrm{FA}]$ denotes a 0/1-indicator that equals one if issuer i is an eligible fallen angel and zero otherwise. To facilitate the comparison between March 23 and April 9 announcement effects, we estimate specification (2) in symmetric five-day windows bracketing each announcement.[21]

The results of this exercise are reported in Table 5. According to the entries in column (1), the estimate of β_3, the coefficient on the interaction term $\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E] \times \mathbb{1}[i = \mathrm{FA}]$, is positive, economically large, and statistically significant. The point estimate of 3.35 implies that credit spreads on SMCCF-eligible bonds issued by fallen angels *increased* 335 basis points relative to their non-eligible counterparts in response to the March 23 announcement.

The large estimated differential increase in credit spreads in response to the March 23 announcement likely reflects the confluence of two factors. First, the actual downgrade to junk status would, all else equal, lead to an increase in credit spreads in both the treatment and control groups. The much larger estimated increase in credit spreads on the fallen angels' treated bonds is likely due to investors' perception that the increase in default risk that led to the downgrade was heavily concentrated in the near term. Second, following the downgrade, the fallen angels' SMCCF-eligible bonds were no longer eligible for purchase by the facility. The loss of program eligibility for bonds in the treatment group would additionally drive up their credit spreads relative to their counterparts in the control group. Both factors—the increase in the near-term risk of default and the loss of bonds' eligibility status—could thus induce a differential effect between the fallen angels' credit spreads in the treatment and control groups.

The estimated effects of the April 9 announcement reported in column (2), by contrast, are as expected. In the five-day window bracketing the announcement, credit spreads on the fallen angels' SMCCF-eligible bonds are estimated to have narrowed more than 160 basis points relative to their ineligible counterparts. All told, the April 9 announcement is estimated to have reversed about one-half of the relative increase in credit spreads for fallen angels that occurred in the aftermath of the March 23 announcement.

Columns (3) and (4) contain the corresponding announcement effects for the bid-ask spreads.

[20] The eligible fallen angel's credit rating would still had to be at least Ba3/BB- as of the date on which the facility purchased their eligible bonds.

[21] Because specification (2) includes issuer fixed effects, the 0/1-indicator $\mathbb{1}[i = \mathrm{FA}]$ is not separately identified.

Table 5: Effect of the SMCCF Announcements on Fallen Angels

	Credit Spreads		Bid-Ask Spreads	
	Mar-23	Apr-9	Mar-23	Apr-9
Explanatory Variable	(1)	(2)	(3)	(4)
$\mathbb{1}[t \geq t^*]$	0.05	-0.72^{***}	-0.26^{***}	-0.21^{***}
	(0.04)	(0.03)	(0.05)	(0.04)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	-0.23^{***}	-0.13^{***}	-0.15^{**}	0.02
	(0.04)	(0.03)	(0.06)	(0.05)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E] \times \mathbb{1}[i = \text{FA}]$	3.35^{***}	-1.63^{**}	0.61^{***}	-0.12
	(0.99)	(0.51)	(0.09)	(0.11)
R^2	0.72	0.90	0.13	0.32
No. of firms	507	521	338	365
No. of bonds	1,592	1,589	1,228	1,328
Observations	8,134	8,534	3,815	4,559

Note: The dependent variable in specifications (1) and (2) is $\text{CS}_{i,j,t}$, the credit spread of bond j (issued by firm i) on business day t, while in specifications (3) and (4), the dependent variables is $\ln \text{BAS}_{i,j,t}$, the log of the corresponding bid-ask spread. The entries in the table denote the OLS estimates of coefficients on the specified explanatory variable: $\mathbb{1}[t \geq t^*]$ = 0/1-indicator that equals one if date t is greater than or equal to the specified announcement date t^* and zero otherwise; $\mathbb{1}[j = E]$ = 0/1-indicator that equals one if bond j was eligible for purchase by the SMCCF as of March 22 and zero otherwise; and $\mathbb{1}[i = \text{FA}]$ = 0/1-indicator that equals one if issuer i became fallen angel within five business days of the March 23 announcement. All specifications include a vector of bond-specific controls (not reported) and issuer fixed effects. Asymptotic standard errors reported in parentheses are clustered at the issuer level: * $p < .10$; ** $p < .05$; and *** $p < .01$.

Consistent with the credit spread results reported in columns (1) and (2), the March 23 announcement is estimated to have boosted bid-ask spreads on the fallen angels' SMCCF-eligible bonds about 60 percent—relative to their ineligible counterparts—in the five days following the announcement. Unlike in the case of credit spreads, this deterioration in relative liquidity was not ameliorated by the April 9 announcement.[22]

3.3 Default Risk vs. Credit Risk Premia

With these baseline results in hand, we now ask a question to what extent are the announcement-induced declines in credit spreads due to a reduction in default risk as opposed to a decline in credit risk premiums? To get at this question, we follow Gilchrist and Zakrajšek (2012) and decompose investment-grade credit spreads into a component that captures issuer-specific time-varying

[22] In Appendix B, we report results from estimating specification (1) using the narrow treatment and control groups constructed from outstanding bonds of issuers with a speculative-grade credit rating as of March 22, 2020. As shown in Table B-1, the March 23 announcement led to an increase of more than 50 basis points in the average high-yield credit spread, a result consistent with the fact that the March 23 announcement explicitly limited future purchases to investment-grade bonds. In response to the April 9 announcement, in contrast, the average high-yield credit spread fell more than 140 basis points; moreover, spreads on high-yield bonds with remaining maturity of less than five years registered an additional decline of 33 basis points compared with their longer maturity counterparts. These results suggest that investors read the April 9 announcement as indicating that the Fed would be ready to support the entire speculative-grade segment of the market if conditions continued to worsen.

default risk and a residual component that can be thought of as capturing investor attitudes toward corporate credit risk in that segment of the market. Specifically, we estimate the following regression:

$$\ln \mathrm{CS}_{i,j,t}^{(\tau)} = \alpha_0 + \alpha_1 \mathrm{DD}_{i,t}^{(\tau)} + \lambda' \mathbf{Z}_{i,j,t} + \nu_{i,j,t}^{(\tau)}, \tag{3}$$

where $\mathrm{CS}_{i,j,t}^{(\tau)}$ is the credit spread of bond j (issued by firm i) with the remaining maturity of τ years, and $\mathrm{DD}_{i,t}^{(\tau)}$ denotes the *distance-to-default* for issuer i over horizon τ, an option-theoretic default-risk indicator based on the firm's equity valuations and its volatility, as well as the firm's leverage (see Merton, 1974). The vector $\mathbf{Z}_{i,j,t}$ includes standard pre-determined bond characteristics that can influence credit spreads though liquidity premia (see Appendix C for details).

We estimate equation (3) by OLS using daily TRACE data on investment-grade corporate bonds issued by publicly listed U.S. companies over the sample period June 2002 to December 2019 and then use the resulting parameter estimates to predict credit spreads over the January–July 2020 period.[23] As discussed by Gilchrist and Zakrajšek (2012), this flexible empirical approach removes from credit spreads equity investors' assessment of the underlying default risk. The estimated residual $\hat{\nu}_{i,j,t}^{(\tau)}$, the (log) credit spread "pricing error," reflects a portion of the credit spread that is not attributable to issuer's default risk and which we interpret as an estimate of the credit risk premium. When averaged across issuers, the resulting average residual credit spread—the so-called excess bond premium (EBP)—captures fluctuations in the average price of bearing corporate credit risk, above and beyond the compensation that investors in the corporate bond market require for expected defaults.

Figure 2 plots the daily estimate of the credit risk premium in the investment-grade segment of the market during the Covid-19 pandemic. Consistent with the benign—in fact, some may say frothy—conditions that characterized credit markets on the eve of the pandemic, the average credit risk premium in the investment-grade segment of the corporate bond market was essentially zero in early 2020. The credit risk premium started to rise in late February and took off amidst the bout of turmoil that swept through global financial markets in early March. In fact, the March run-up in our estimate of the average investment-grade credit risk premium is comparable in magnitude to the increase in the Gilchrist and Zakrajšek (2012) original estimate of the EBP in the aftermath of the collapse of Lehman Brothers in September 2008.

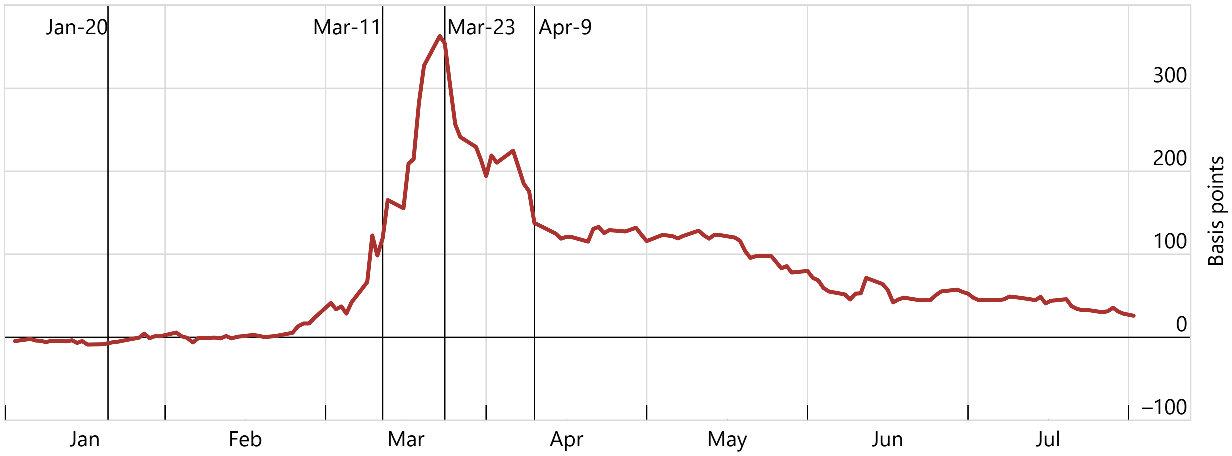
The increase in the average credit risk premium during the pandemic accounts for roughly three-fourth of the total rise in the average credit spread for our sample of bonds. This suggests that much of the rise in investment-grade credit spreads in response to the Covid-19 shock can be attributed to increases in credit risk premia, or a deterioration in credit market sentiment, as opposed to increases in the likelihood of default.

In Table 6, we report the results from estimating specification (1), using the residual credit spreads as the dependent variable.[24] It is important to note that because the distance-to-default

[23] We start the estimation in June 2002, when TRACE data first became available and stop in December 2019 to avoid any "look-ahead" bias when predicting credit spreads during the pandemic.

[24] Because measuring distance-to-default requires equity prices, the sample of bonds used in this exercise corresponds

FIGURE 2: Investment-Grade Credit Risk Premium During the Covid-19 Pandemic



NOTE: The solid red line shows the time-series of the cross-sectional average of the residual credit spreads, a proxy for the credit risk premium (see the text for details). Vertical lines at specified dates: Jan-20 = Chinese officials acknowledge that Covid-19 might be transmissible between humans; Mar-11 = WHO declares Covid-19 a pandemic; Mar-23 = Fed announces the establishment of the P/SMCCF; and Apr-9 = Fed expands the facilities to include corporate bonds of issuers that were rated investment grade as of March 22 but were subsequently downgraded to junk.
SOURCE: Authors' calculations using data from TRACE, CRSP, and S&P's Compustat.

increases with the horizon (i.e., the bond's remaining maturity τ), such variation in default risk is not automatically picked up by the inclusion of issuer fixed effects in the regression specification. Nonetheless, the estimates of coefficients β_1 and β_2 using the residual credit spreads as the dependent variable are almost identical—in terms of both their magnitudes and temporal patterns—as those that use the actual credit spreads as the dependent variable (see Panel A of Table A-1).

These findings imply that the announcement-induced declines in the average credit spread, as well as the additional declines in credit spreads of SMCCF-eligible bonds, are due primarily to a reduction in credit risk premia, or an improvement in market sentiment, rather than to a reduction in default risk, at least as perceived by equity markets.[25] As such, they are consistent with the theoretical framework of Hanson et al. (2020), in which the main mechanism through which the creation of a corporate bond-buying facility affects the market is through a reduction in uncertainty and the associated credit risk premia. The announcement-induced declines in credit risk premia could also reflect a reduction in the risk-aversion of broker-dealers—the marginal investors in the corporate bond market—whose pullback from market-making during the height of the pandemic has

to a subset of U.S. issuers in the TRACE database that are publicly listed. Given this more restrictive sample, we use the full-sample definitions of the treatment and control groups to estimate the SMCCF announcement effects (see Appendix A).

[25]Figure C-1 in Appendix C provides further evidence in support of this interpretation. In particular, it shows the distribution of the distance-to-default in the high- and low-quality segments on the investment-grade corporate bond market in the five-day windows bracketing the two announcements. The relatively small shifts in the distribution of default risk in response to both announcements are consistent with our result that the announcement-induced declines in credit spreads reflect primarily a reduction in credit risk premia.

Table 6: Effect of the SMCCF Announcements on Credit Risk Premia

	2-day	5-day	10-day
Explanatory Variable	(1)	(2)	(3)
A. Mar-23 announcement			
$\mathbb{1}[t \geq t^*]$	-0.30^{***}	0.02	0.46^{***}
	(0.03)	(0.04)	(0.05)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	-0.47^{***}	-0.42^{***}	-0.28^{***}
	(0.04)	(0.04)	(0.05)
R^2	0.66	0.59	0.55
No. of firms	496	543	565
No. of bonds	2,555	2,785	2,926
Observations	9,889	21,473	40,452
B. Apr-9 announcement			
$\mathbb{1}[t \geq t^*]$	-0.48^{***}	-0.63^{***}	-0.67^{***}
	(0.02)	(0.02)	(0.03)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	-0.05^{*}	-0.15^{***}	-0.24^{***}
	(0.02)	(0.03)	(0.03)
R^2	0.88	0.87	0.85
No. of firms	516	555	569
No. of bonds	2,596	2,781	2,942
Observations	10,037	22,181	42,469

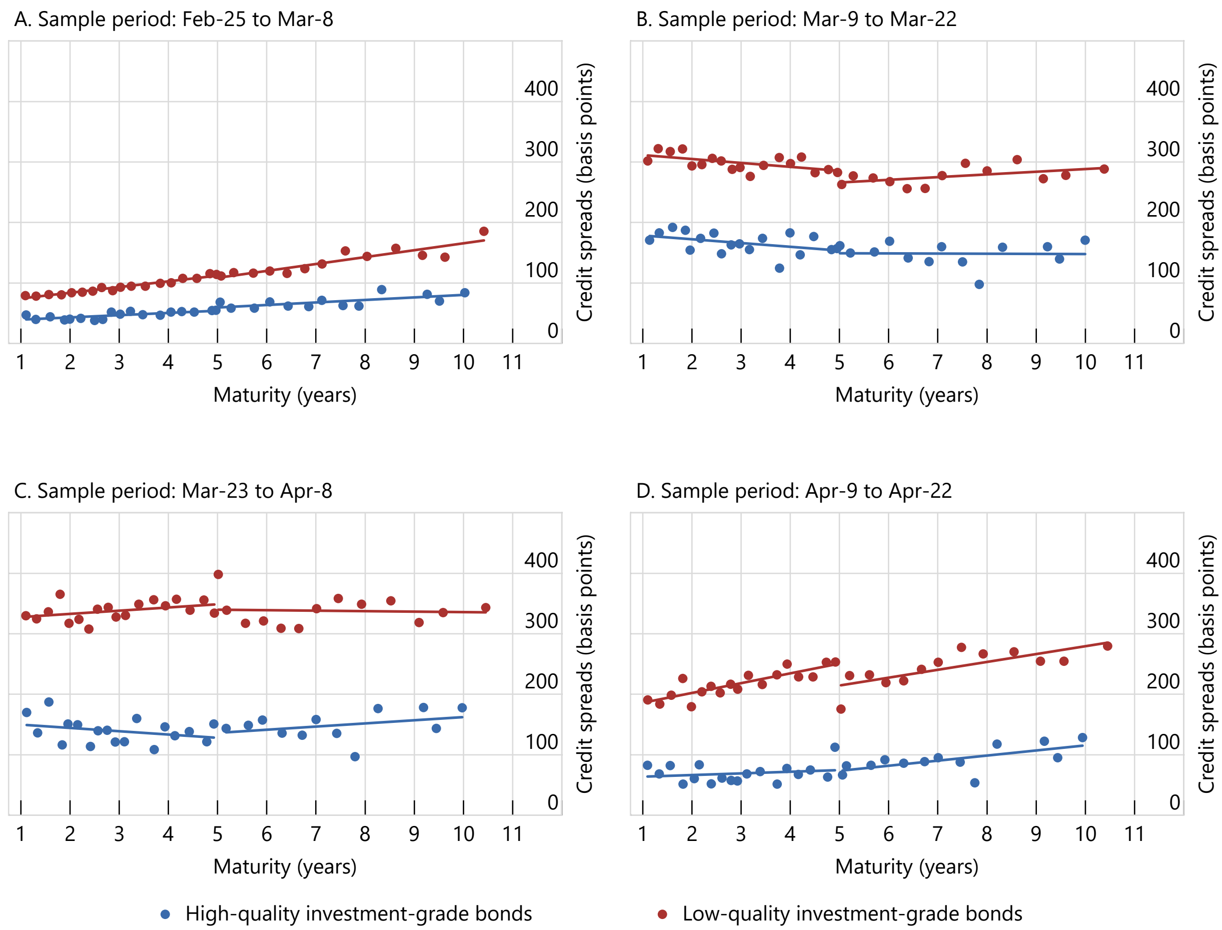
Note: The dependent variable in all specifications is $\text{RCS}_{i,j,t}$, the residual credit spread of bond j (issued by firm i) on business day t (see the text for details). The entries in the table denote the OLS estimates of coefficients associated with the specified explanatory variable: $\mathbb{1}[t \geq t^*] = 0/1$-indicator that equals one if date t is greater than or equal to the specified announcement date t^* and zero otherwise; and $\mathbb{1}[j = E] = 0/1$-indicator that equals one if bond j was eligible for purchase by the SMCCF as of March 22 and zero otherwise. All specifications include a vector of bond-specific controls (not reported) and issuer fixed effects. Asymptotic standard errors reported in parentheses are clustered at the issuer level: * $p < .10$; ** $p < .05$; and *** $p < .01$.

been documented by O'Hara and Zhou (2021) and Kargar et al. (2021); we explore this mechanism quantitatively within a preferred-habitat framework in Section 4.

3.4 Announcements Effects and the Slope of the Credit Curve

The results reported above emphasized the differential impact of the two SMCCF announcements on the average *level* of credit spreads between program-eligible and ineligible bonds. In this section, we extend the analysis by considering how the two announcements affected the *slope* of the term structure of investment-grade credit spreads. That is, we ask a question whether the estimated announcement effects on credit spreads reported in Table 3 are uniform across maturities within the program-eligible and ineligible groups of bonds, an assumption implicit in the DiD methodology that relies on the maturity-eligibility cutoff as a treatment effect.

Figure 3: Investment-Grade Credit Curve During the Covid-19 Pandemic



Note: Each panel shows the binscatter plot of the relationship between credit spreads and maturity in the investment-grade segment of the U.S. corporate bond market during the specified period of the Covid-19 pandemic. High-quality investment-grade bonds are those in Aaa/AAA or Aa/AA rating categories, while low-quality investment-grade bonds are those in A/A or Baa/BBB categories.
Source: Authors' calculations using TRACE data.

To do so, we consider the full-sample treatment and control groups, which assign—based on the five-year maturity cutoff—all outstanding investment-grade bonds with remaining maturity between one and 12 years into the relevant group (see Appendix A). Motivation for this analysis is provided by Figure 3, which shows the fitted credit curve—the cross-sectional relationship between credit spreads and maturity—estimated on a pooled sample of bonds in the full-sample treatment and control groups during the various phases of the pandemic. We distinguish between "high" and "low" quality investment-grade bonds, with the former plotted in blue and the latter in red.[26]

[26]High-quality investment-grade bonds are those whose average credit rating across Moody's, S&P, and Fitch ratings categories falls in the Aaa/AAA or Aa/AA categories, while low quality are those whose average credit rating falls in the A/A or Baa/BBB categories. In case the average of issuer ratings across the three rating agencies was not an integer, we applied the "floor" function to the resulting average.

Panel A focuses on the early phase of the pandemic-induced turmoil in the corporate bond market. While credit spreads had widened some during this period, the slope of credit curve in both segments of the investment-grade market remained stable and upward sloping. In mid-March, as the crisis gathered momentum, credit spreads spiked, and the credit curve inverted, as spreads of shorter-maturity bonds increased considerably more than spreads of their longer-maturity counterparts (Panel B). The inversion was, on balance, somewhat more pronounced in the high-quality investment-grade segment of the market, where the increase in spreads on shorter-maturity bonds was especially large in relative terms. These patterns are consistent with the well-documented dash for cash, whereby investors at the nadir of the pandemic-induced panic first tried to liquidate their holdings of most liquid securities, namely shorter-maturity high-quality investment-grade bonds (see Haddad et al., 2021).[27]

Panel C captures the period following the March 23 announcement but before the April 9 announcement. During this period, credit spreads widened further, on balance, while the inversion of the credit curve lessened somewhat. Lastly, Panel D focuses on the couple of weeks following the April 9 announcement. Although credit spreads remained elevated, credit curves for both high- and low-quality investment-grade bonds are again upward sloping, with slopes of comparable magnitude to those seen during the initial phase of the crisis shown in Panel A. These rotations suggest that one of the key aspects of the two announcements was to restore the normal upward slope of the investment-grade credit curve. They also imply that one must control for such shifts in the credit curve when assessing the impact of the announcement effects through the five-year eligibility cutoff.

To formally do so, we augment the baseline specification (1) with an interaction term that allows the slope of the credit curve to shift in the post-announcement window. Specifically, we estimate:

$$\begin{aligned}\mathrm{CS}_{i,j,t} = {} & \beta_1 \mathbb{1}[t \geq t^*] + \beta_2\big(\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]\big) + \beta_3\big(\mathbb{1}[t \geq t^*] \times \tau_{i,j,t}\big) \\ & + \beta_4\big(\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E] \times \tau_{i,j,t}\big) + \theta'\mathbf{X}_{i,j,t} + \eta_i + \epsilon_{i,j,t},\end{aligned} \tag{4}$$

where $\tau_{i,j,t}$ denotes the remaining maturity of bond j on business day t. Table 7 summarizes results of this exercise.

According to columns (1)–(3), the coefficients β_2 and β_4 on the interaction terms $\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$ and $\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E] \times \tau_{i,j,t}$, respectively, are both statistically highly significant in all three estimation windows. The fact that $\hat{\beta}_2 < 0$ and $\hat{\beta}_4 > 0$ implies that the March 23 announcement steepened the program-eligible segment of the investment-grade credit by compressing credit spreads at the short-end of the curve. The estimates of β_2 and β_4 in columns (4)–(6), by contrast, are all statistically indistinguishable from zero, implying no such slope effect in response to the April 9 announcement. As evidenced by the statistically significant coefficients β_1 and β_3 on the the interaction terms $\mathbb{1}[t \geq t^*]$ and $\mathbb{1}[t \geq t^*] \times \tau_{i,j,t}$, respectively, the entire investment-grade credit curve steepened in response to the April 9 announcement, with no differential effect between the program-eligible and ineligible segments of the curve.

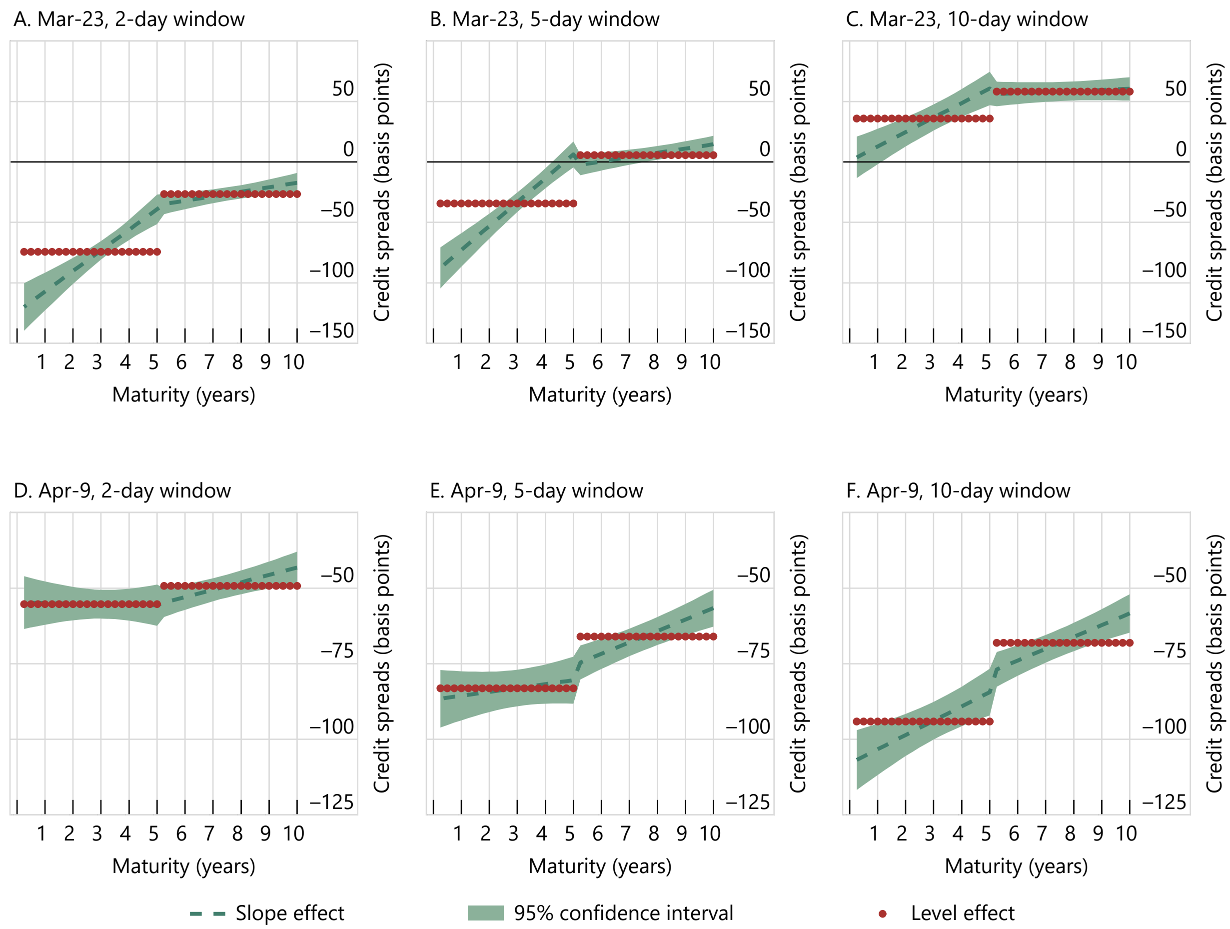
[27] Cesa-Bianchi and Eguren-Martin (2021) document a related "dash for dollars," by comparing the performance of U.S. dollar-denominated bonds with that of non-dollar bonds during the pandemic.

TABLE 7: Effect of the SMCCF Announcements on the Slope of the Credit Curve

	Mar-23 Announcement			Apr-9 Announcement		
	2-day	5-day	10-day	2-day	5-day	10-day
Explanatory Variable	(1)	(2)	(3)	(4)	(5)	(6)
$\mathbb{1}[t \geq t^*]$	−0.54***	−0.22*	0.50***	−0.67***	−0.94***	−0.96***
	(0.11)	(0.10)	(0.11)	(0.06)	(0.06)	(0.07)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	−0.70***	−0.71***	−0.49***	0.13	0.07	−0.12
	(0.14)	(0.12)	(0.12)	(0.07)	(0.07)	(0.08)
$\mathbb{1}[t \geq t^*] \times \tau_{i,j,t}$	0.04**	0.04**	0.01	0.02***	0.04***	0.04***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E] \times \tau_{i,j,t}$	0.13***	0.16***	0.11***	−0.03	−0.02	0.01
	(0.03)	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)
R^2	0.72	0.66	0.61	0.91	0.90	0.88
No. of firms	747	840	887	785	872	907
No. of bonds	3,717	4,186	4,509	3,819	4,203	4,529
Observations	13,370	29,034	55,319	13,861	30,504	58,134

NOTE: The dependent variable in all specifications is $\text{CS}_{i,j,t}$, the credit spread of bond j (issued by firm i) on business day t. The entries in the table denote the OLS estimates of coefficients associated with the specified explanatory variable: $\mathbb{1}[t \geq t^*] = 0/1$-indicator that equals one if date t is greater than or equal to the specified announcement date t^* and zero otherwise; $\mathbb{1}[j = E] = 0/1$-indicator that equals one if bond j was eligible for purchase by the SMCCF as of March 22 and zero otherwise; and $\tau_{i,j,t}$ = bond j's remaining maturity (in years). All specifications include a vector of bond-specific controls (not reported) and issuer fixed effects. Asymptotic standard errors reported in parentheses are clustered at the issuer level: * $p < .10$; ** $p < .05$; and *** $p < .01$.

Figure 4: Announcement-Induced Shifts in the Slope of the Investment-Grade Credit Curve



Note: The green dashed lines in Panel A–C show the estimated effect of the March 23 announcement on the slope of the investment-grade credit curve, whereas those in Panels D–F show the corresponding effects of the April 9 announcement; the green shaded bands represent the corresponding 95 percent confidence intervals (see specification (4) and the corresponding estimates reported in Table 7 for details). The red dotted lines show the baseline estimates of the level effect implied by the two announcements (see specification (1) and estimates reported in Panel A of Table A-1 in Appendix A for details).
Source: Authors' calculations using TRACE data.

Figure 4 displays the estimated announcement effects across the full range of maturities considered in the estimation. Specifically, the green dashed line in each panel uses the relevant estimates of $\beta_1, \ldots, \beta_4$ reported in Table 7 to trace out the announcement-induced shift in the investment-grade credit curve in the specified estimation window; the green shaded bands represent the associated 95-percent confidence intervals. For comparison purposes, the red dotted lines show the announcement-induced changes in the average level of credit spreads for program-eligible and ineligible bonds based on the baseline DiD specification (1), which does not control for the associated changes in the slope of the credit curve.

The figure clearly illustrates the main mechanism by which the Fed's announcements affected pricing of investment-grade corporate securities in the secondary market. Turning first to the March 23 announcement (Panels A–C), the estimated rotations of the credit curve imply economically large differences in the announcement effects across maturities. In the two-day window following the March 23 announcement, our estimates imply a reduction of more than 100 basis points in credit spreads for bonds with a remaining maturity of one year, compared with a reduction of about 15 basis points in spreads of bonds with a remaining maturity of ten years. And while the entire curve shifts noticeably higher and flattens somewhat as the estimation window lengthens, the slope effect induced by the March 23 announcement on program-eligible bonds remains considerably stronger and statistically different from the slope effect on program-ineligible bonds.

In response to the April 9 announcement (Panels D–E), by contrast, the entire term structure of credit spreads steepens and moves noticeably lower as the estimation window lengthens. And while the March 23 announcement had a differential effect on credit spreads of shorter-term program-eligible bonds, the April-9 announcement had a uniform effect on the term structure of investment-grade credit spreads. When summed up across the two announcements, our estimates based on the ten-day window imply an announcement-induced differential of about 100 basis points between bonds with remaining maturities of one and ten years.

To sum up, the results in Table 7 and Figure 4 indicate that the March 23 announcement had a significantly stronger slope effect on the program-eligible segment of the credit curve, which helped to restore the upward-sloping term structure of credit spreads, especially at the short-end of the curve. The April 9 announcement, in contrast, affected the entire term structure of credit spreads, irrespective of whether the underlying securities were eligible or ineligible for purchase by the SMCCF.

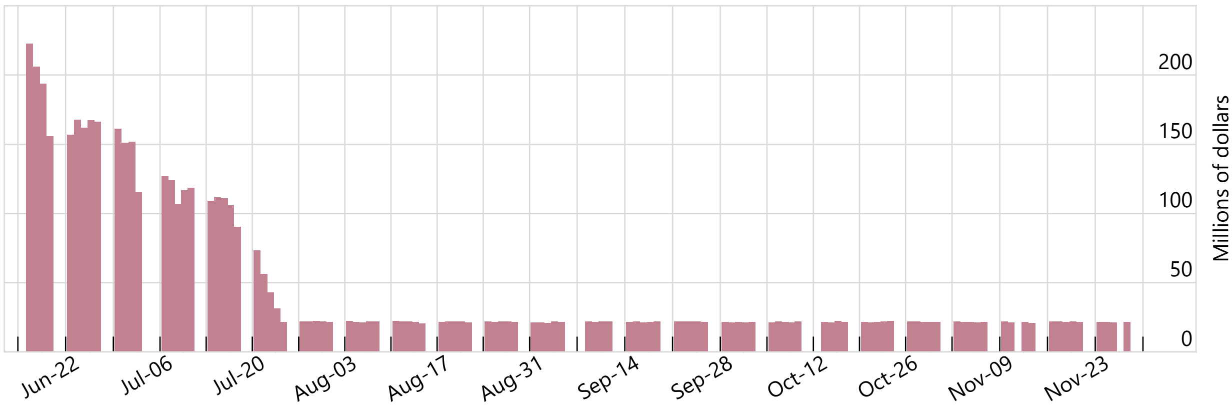
3.5 The SMCCF's Purchase Effects

The previous section focused on the impact of the two bond-buying program announcements on conditions in the corporate bond market. We now turn to quantifying the impact of the Fed's actual purchases on credit and bid-ask spreads.

As noted above, the Fed started to purchase individual corporate bonds on June 16, 2020, and ended all purchases on November 27, 2020. Figure 5 shows the dollar amount of corporate bonds purchased by the facility during that period. In the latter half of June, the SMCCF purchased about \$170 million of corporate bonds during an average day. The average pace of purchases tapered off to about \$120 million per day during the first half of July before dropping to about \$20 million by the end of July, a pace that was maintained through the end of the facility's purchase operations. In total, the SMCCF purchased about \$5.4 billion corporate bonds in a span of five and a half months, with almost 70 percent of those purchases (about \$3.6 billion) taking place in the second half of June and July.

To identify the effect of these purchases on credit and bid-ask spreads, we utilize the *intraday* TRACE transactions data between June 16 and July 31 to identify the Fed's purchases of individual bonds, a period during which the facility made most of its purchases. By matching the bond's

Figure 5: SMCCF's Purchases of Corporate Bonds



Note: The vertical bars show the total daily dollar amount of individual corporate bonds purchased by the SMCCF between June 16, 2020 and November 27, 2020.
Source: Authors' calculation using data from the Federal Reserve Bank of New York.

CUSIP, purchase date and time, transaction price and quantity in dealer-to-customer transactions, we are able to identify almost all of the Fed's purchases of individual corporate bonds during this period.[28] Using the exact time of each purchase, we perform a simple event study, whose major advantage is that we are able to estimate a precise average purchase effect.

The results of this exercise are summarized in Figure 6. The red line in Panel A of shows the average credit spread on bonds purchased by the SMCCF within the event window that spans 20 hours before and 20 hours after the purchase time, which is normalized to be equal to zero. According to this figure, the credit spread on an average purchased bond declined about five basis points upon the actual purchase. Over the subsequent five hours, the spread edged up about two basis points before stabilizing over the remainder of the event window for a net decline of about three basis points. The blue line shows the corresponding average spread in the control group—that is, bonds issued by the same set of issuers but whose remaining maturity is greater than five years.[29] Interestingly, the actual purchases appear to have also had a delayed effects on the credit spreads of ineligible bonds, though this effect is very small, a mere basis point or so.[30]

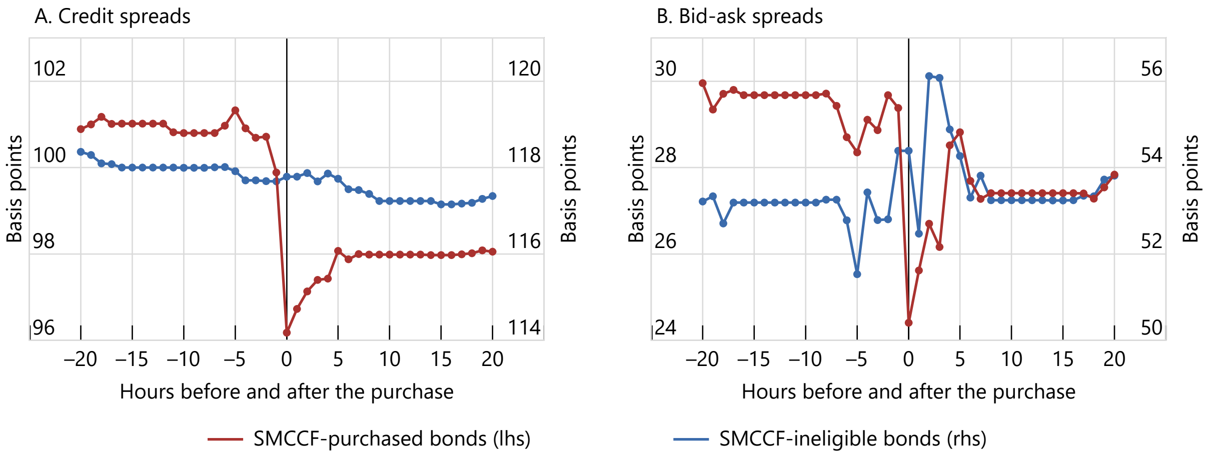
Panel B shows the same event study for the bid-ask spreads. Though considerably more noisy,

[28] We identify all of the facility's 1,351 purchases, except for a single purchase on June 29; this transaction involved the bond with CUSIP 126650CT5, issued by the CVS Health Corporation, which had two matches at slightly different times: 11:33:39 a.m. and 11:59:12 a.m. We dropped this transaction from the analysis.

[29] As before, we construct the control group by pairing each bond purchased by the SMCCF—the treatment group—with a bond issued by the same company but whose remaining maturity is greater than five years. There are 482 unique issuers in our treatment group. If an issuer has multiple bonds purchased by the SMCCF, we choose the bond with remaining maturity as close to five years as possible. Similarly, if there are multiple bonds that can be paired up with a given bond purchased by the SMCCF, we choose the bond with a remaining maturity as close to five years as possible.

[30] In Appendix D, we show that the purchase effects on credit spreads are attributable entirely to declines in corporate bond yields and were concentrated in the low quality segment of the investment-grade corporate bond market.

FIGURE 6: The Impact of the SMCCF's Corporate Bond Purchases



NOTE: The red line in Panel A shows the average credit spread on bonds purchased by the SMCCF in a symmetric 20-hour window bracketing their purchases, while the blue line shows the corresponding average credit spread on the SMCCF-ineligible bonds issued by the same set of issuers. The corresponding lines in Panel B show the average bid-ask spreads for the same two groups of bonds.
SOURCE: Authors' calculations using TRACE data and data from the Federal Reserve Bank of New York.

the average bid-ask spread on bonds actually purchased by the SMCCF (the red line) is estimated to have declined about five basis points upon purchase before bouncing back over the subsequent several hours. The average bid-ask spread on bonds in the control group (the blue line), by contrast, shows no discernible pattern around the purchase time.

All told, our empirical analysis indicates that the vast majority of the SMCCF's impact on the corporate bond market—as reflected in a substantially narrower credit and bid-ask spreads—is due to the announcement effects, which occurred well before the Fed directly intervened in the market. Nonetheless, we find evidence that the actual purchases of bonds in the secondary market also had an effect on credit spreads. In particular, the Fed's actual purchases lowered, on average, credit spreads on purchased bonds two basis points relative to program-ineligible bonds of the same set of issuers. Given the modest—by QE standards—scale of these purchases this is a sizable effect, corroborating the overall efficacy of the program.

4 Inspecting the Mechanism

In this section, we use the preferred-habitat framework of Vayanos and Vila (2021) to shed light on the mechanism of the SMCCF. We first extend their model of the Treasury bond market to also include market for high-grade corporate debt. We then calibrate the extended model to match key moments in the data and use the model to match the observed shifts in the investment-grade credit curve around March 23 and April 9 announcements.

4.1 The Setup

We consider a version of the preferred-habitat model with two markets: a market for Treasury bonds and a parallel market for high-quality (i.e., Aaa/AAA or Aa/AA rated) investment-grade corporate bonds. By focusing on the safest segment of the corporate bond market, we abstract from default risk. In our framework, credit spreads are driven by the differential exposure of corporate and comparable-maturity Treasury bonds to exogenous fluctuations in the short rate and by the market-specific demand shocks. Differences in demand across Treasury and corporate bond markets are motivated by differences in investors' liquidity preferences.

In terms of notation, we use the tilde symbol on top of a variable to refer to a variable or a process in the corporate bond market, while the same variable without the tilde symbol denotes the corresponding variable or process in the Treasury market. Specifically, let $P_t^{(\tau)}$ and $y_t^{(\tau)} = -\ln P_t/\tau$ denote the time-t price and yield of a (zero-coupon) Treasury bond with remaining maturity τ, respectively, and $\widetilde{P}_t^{(\tau)}$ and $\tilde{y}_t^{(\tau)} = -\ln \widetilde{P}_t^{(\tau)}/\tau$ denote the respective price and yield of the same maturity (zero-coupon) corporate bond. As in Vayanos and Vila (2021), we assume there are two types of agents in each market: arbitrageurs and preferred-habitat investors. To minimize the departure from their model, we assume that the two markets have separate groups of arbitrageurs and preferred-habitat investors.

Consider first the corporate bond market with a continuum of corporate bonds in zero net supply and with maturities ranging between 0 and T. Arbitrageurs in the corporate bond market can invest in the short rate and corporate bonds and do so to maximize a mean-variance objective over instantaneous changes in their wealth:

$$\max_{\{\widetilde{X}_t^{(\tau)}\}_{\tau=0}^{T}} \left\{ E_t\left[d\widetilde{W}_t\right] - \frac{\tilde{a}}{2}\operatorname{Var}_t\left[d\widetilde{W}_t\right] \right\}$$
$$\text{subject to} \quad d\widetilde{W}_t = \widetilde{W}_t r_t dt + \int_0^T \widetilde{X}_t^{(\tau)} \left(\frac{d\widetilde{P}_t^{(\tau)}}{\widetilde{P}_t^{(\tau)}} - r_t dt \right) d\tau,$$

where $\widetilde{W}_t$ and $\widetilde{X}_t^{(\tau)}$ denote the arbitrageurs' wealth and position in maturity-τ corporate bonds, respectively, and $\tilde{a} \geq 0$ is the arbitrageurs' risk-aversion coefficient, a proxy for their risk-bearing capacity. The short rate r_t follows an exogenous Ornstein-Uhlenbeck process of the form

$$dr_t = \kappa_r(\bar{r} - r_t)dt + \sigma_r dB_{r,t},$$

where $\bar{r}$ denotes its long-run average and κ_r and σ_r denote the mean-reversion and diffusion parameters, respectively.

The preferred-habitat investors in the corporate bond market constitute a maturity clientele, in the sense that the maturity-τ investors demand only corporate bonds with maturity τ, according to

$$\widetilde{Z}_t^{(\tau)} = \tilde{\alpha}(\tau)\tau\tilde{y}_t^{(\tau)} - \tilde{\beta}_t^{(\tau)},$$

where $\tilde{\alpha}(\tau)$ and $\tilde{\beta}_t^{(\tau)}$ denote the *slope* and *intercept* of their demand function at time t, respectively. The stochastic demand intercept $\tilde{\beta}_t^{(\tau)}$ is given by

$$\tilde{\beta}_t^{(\tau)} = \tilde{\theta}_0(\tau) + \tilde{\theta}(\tau)\tilde{\beta}_t,$$

where the demand risk factor $\tilde{\beta}_t$ follows an Ornstein-Uhlenbeck process of the form

$$d\tilde{\beta}_t = -\kappa_{\tilde{\beta}}\tilde{\beta}_t dt + \sigma_{\tilde{\beta}} dB_{\tilde{\beta},t}.$$

As shown by Vayanos and Vila (2021), bond prices in this framework are an exponential-linear function of the short rate r_t and the demand risk factor $\tilde{\beta}_t$. In particular, the price of the maturity-τ corporate bond is given by

$$\widetilde{P}_t^{(\tau)} = e^{-\left(\widetilde{A}_r(\tau)r_t + \widetilde{A}_{\tilde{\beta}}(\tau)\tilde{\beta}_t + \widetilde{C}(\tau)\right)},$$

where the coefficients $\widetilde{A}_r(\tau)$, $\widetilde{A}_{\tilde{\beta}}(\tau)$, and $\widetilde{C}(\tau)$ are determined endogenously in equilibrium. As a result, the corresponding yield is affine in the state variables and is given by

$$\tilde{y}_t^{(\tau)} = \frac{1}{\tau}\left[\widetilde{A}_r(\tau)r_t + \widetilde{A}_{\tilde{\beta}}(\tau)\tilde{\beta}_t + \widetilde{C}(\tau)\right].$$

Because the Treasury market has the same structure, Treasury bond prices and yields have analogous expressions: $P_t^{(\tau)} = \exp\left(-[A_r(\tau)r_t + A_\beta(\tau)\beta_t + C(\tau)]\right)$ and $y_t^{(\tau)} = \frac{1}{\tau}[A_r(\tau)r_t + A_\beta(\tau)\beta_t + C(\tau)]$, where β_t represents the exogenous demand risk factor in the Treasury bond market, which evolves according to $d\beta_t = -\kappa_\beta\beta_t dt + \sigma_\beta dB_{\beta,t}$. Therefore, credit spread of the maturity-τ corporate bond is given by

$$\begin{aligned} s_t^{(\tau)} &\equiv \tilde{y}_t^{(\tau)} - y_t^{(\tau)} \\ &= \frac{1}{\tau}\left[\left(\widetilde{A}_r(\tau) - A_r(\tau)\right)r_t + \left(\widetilde{A}_{\tilde{\beta}}(\tau)\tilde{\beta}_t - A_\beta(\tau)\beta_t\right) + \left(\widetilde{C}(\tau) - C(\tau)\right)\right]. \end{aligned}$$

In other words, credit spreads are determined by the differential exposure of corporate and comparable-maturity Treasury bonds to exogenous fluctuations in the short rate and by the idiosyncratic demand shocks in the two markets. As we show below, a reasonable calibration of this simple framework is capable of matching the dynamics of investment-grade credit spreads during the height of the pandemic.[31]

Following Vayanos and Vila (2021), we parameterize the demand-side parameters in the corpo-

[31] This framework could be extended in a number of dimensions to make it more realistic. For example, we can introduce another group of arbitrageurs who arbitrage between the corporate and Treasury bond markets, thereby intrinsically linking prices of the two types of securities through the risk-bearing capacity of these cross-market arbitrageurs. Moreover, one can introduce exogenous default risk factor(s), so that credit spreads partly reflect expected defaults. In this paper, however, we abstract from corporate default risk because our aim is to characterize movements in the credit curve of high-quality, investment-grade bonds, a segment of the corporate bond market that experienced most severe dislocations at the onset of the pandemic. Default risk of high-quality investment-grade bonds is, in general, quite low and as shown in Panels A and C of Figure C-1 in Appendix C did not change materially in the narrow windows surrounding the March 23 and April 9 announcements.

rate bond market as

$$\tilde{\alpha}(\tau) = \tilde{\alpha} e^{-\tilde{\delta}_\alpha \tau}; \quad \tilde{\theta}_0(\tau) = \tilde{\theta}_0 \left(e^{-\tilde{\delta}_\alpha \tau} - e^{-\tilde{\delta}_\theta \tau} \right); \quad \tilde{\theta}(\tau) = \tilde{\theta} \left(e^{-\tilde{\delta}_\alpha \tau} - e^{-\tilde{\delta}_\theta \tau} \right), \text{ for } 0 \leq \tau \leq T.$$

We use the same functional forms for $\alpha(\tau)$, $\theta_0(\tau)$, and $\theta(\tau)$ in the Treasury market. Because a vast majority of outstanding corporate bonds has a remaining maturity less than 20 years, we set $T = 20$.

In calibrating the parameters of these functional forms, we use TRACE data from July 2002 to January 2020 on corporate bond yields and the corresponding yields of the duration-matched synthetic Treasuries constructed following Gilchrist and Zakrajšek (2012), which were used to calculate the bond-level credit spreads. Because we have abstracted from default risk, we restrict the TRACE data to high-quality investment-grade corporate bonds issued by U.S. companies. For the resulting sample of corporate bonds and the associated sample of duration-matched synthetic Treasuries, we use the Nelson-Siegel-Svensson framework to estimate a monthly yield curve, from which we then calculate the implied yields at maturities of $1, 2, \ldots, 20$ years in each market.[32]

As Vayanos and Vila (2021), we calibrate parameters in the Treasury bond market (i.e., κ_r, σ_r, κ_β, a, α, θ_0, θ, δ_α, and δ_θ) to match the selected moments of yields and trading volume in the Treasury bond market or we set them equal to the values from the literature. We use an analogous set of moments from the corporate bond market to calibrate the corresponding parameters in that market (i.e., $\kappa_{\tilde{\beta}}$, $\tilde{a}$, $\tilde{\alpha}$, $\tilde{\theta}_0$, $\tilde{\theta}$, $\tilde{\delta}_\alpha$, and $\tilde{\delta}_\theta$). Table E-1 in Appendix E conveniently summarizes the calibrated parameters and the empirical moments used to determine them.

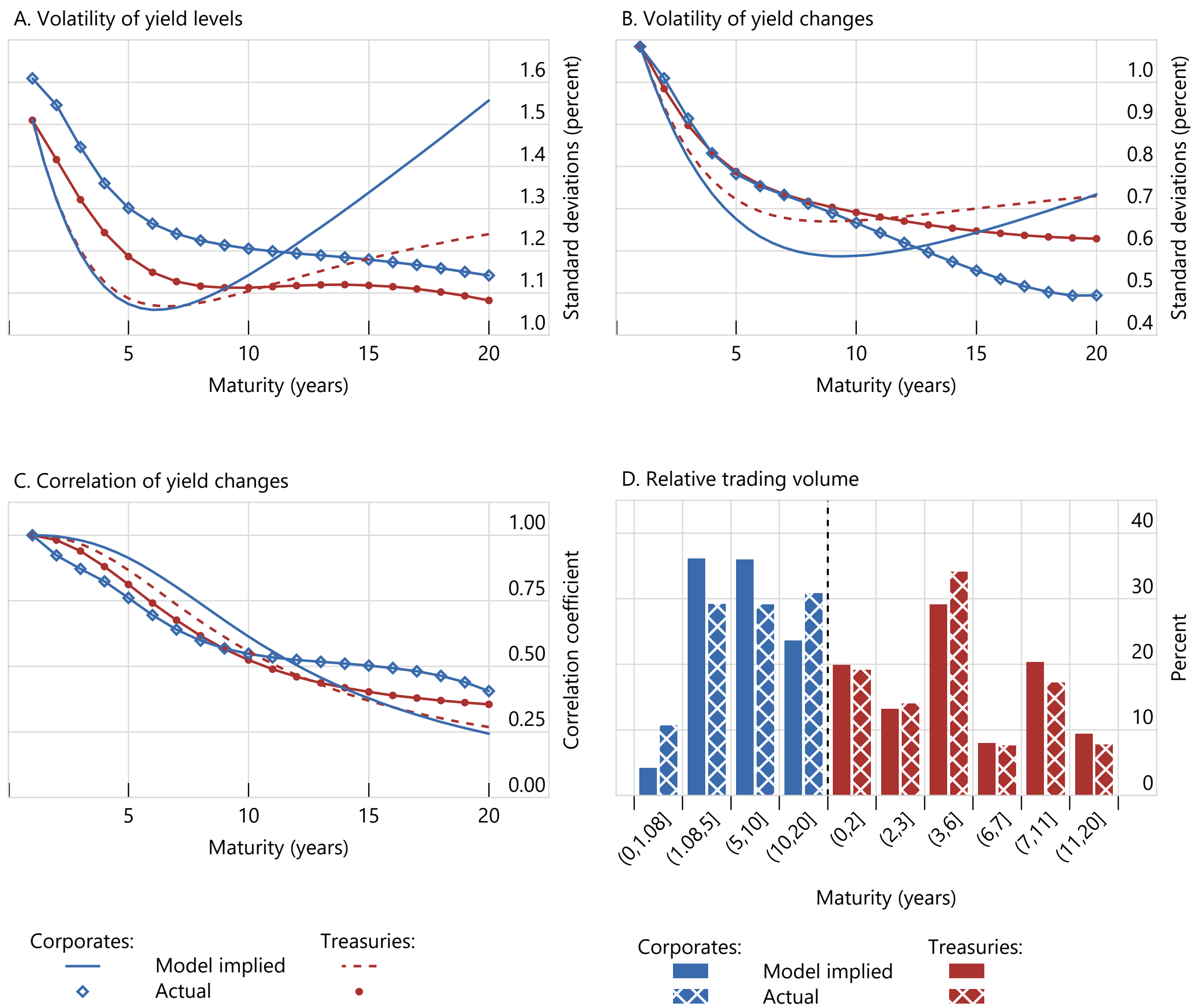
Figure 7 shows the selected model-implied moments and their empirical analogues across the range of maturities considered in our calibration procedure. As shown in Panels A and B, our calibrated model is able to generate the downward-sloping volatility profile of yields and yield changes at short and intermediate maturities, a segment of the market where the Fed's announcement induced the largest movements in the investment-grade credit curve. According to Panel C, the model is also able to produce a correlation structure of yield changes across maturities that closely matches that observed in the data. And as shown in Panel D, the model-implied relative trading volumes across maturity buckets accord, on average, quite well with those observed in the two markets before the pandemic.

4.2 Understanding the Mechanism of the SMCCF

We now use the calibrated model to shed light on the potential mechanisms underlying the dynamics of the investment-grade credit curve during the latter part of March and the first half of April, 2020. Figure 8 highlights the key movements in the credit curve that are focus of this exercise. Specifically, the figure shows the fitted credit curve at specified dates, obtained from a cross-sectional regression of credit spreads on high-quality investment-grade corporate bonds on a quadratic polynomial in

[32] As a robustness check, we compared the implied yields on the synthetic Treasuries at maturities $1, 2, \ldots, 20$ years to the corresponding yields estimated by Gürkaynak et al. (2007) using prices of the actual U.S. Treasuries. At all maturities and across time, differences in the implied yields were negligible.

Figure 7: Selected Model-Implied and Empirical Moments

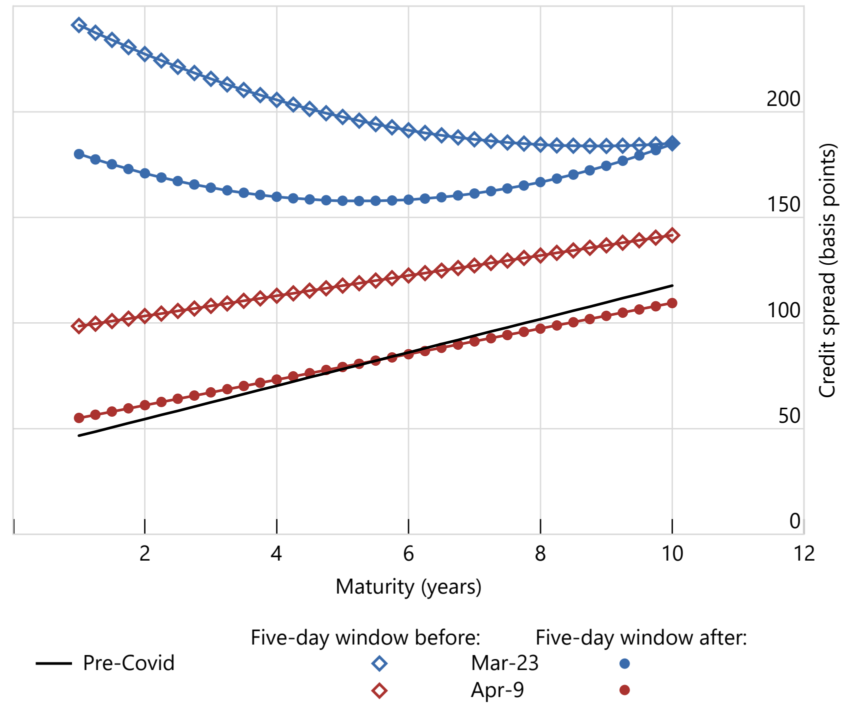


Note: The panels of figure compare the selected model-implied moments with their empirical counterparts. See the main text and Table E-1 in Appendix E for details.
Source: Authors' calculations using TRACE data and data from the Federal Reserve Bank of New York.

the bonds' remaining maturity.

As shown by the solid black line, the average pre-pandemic credit curve in the high-quality segment of the investment-grade corporate bond market shows a linear upward-sloping relationship between credit spreads and maturity. The estimate of the same curve based in the five-day window before the March 23 announcement (the blue line with $\diamond$'s) shows how the pandemic roiled the corporate bond market: not only has the whole curve shifted markedly higher, but it also inverted, with credit spreads on shorter-maturity corporate bonds significantly exceeding those on their longer-maturity counterparts. Within the five days following the March 23 announcement (the blue line with $\bullet$'s), the short and intermediate end of the curve moved significantly lower, noticeably lessening

Figure 8: Investment-Grade Credit Curve During the Covid-19 Pandemic



Note: The red and blue lines show the fitted high-quality investment-grade credit curve, implied by a regression of credit spreads on the quadratic polynomial in the bonds' remaining maturity in five-day windows bracketing the March 23 and April 9 announcements. The five (business) day window before March 23 runs from March 16 to March 22, while the five (business) day window after March 23 runs from March 23 to March 30. The corresponding five (business) day windows bracketing the April 9 announcements are April 2 to April 8 and April 9 to April 17. The solid black line shows the fitted linear credit curve based on daily data from July 2002 through January 2020.
Source: Authors' calculations using TRACE data.

the inversion. By early April, the credit curve, though still elevated by pre-pandemic standards, has resumed its normal upward-sloping shape (the red line with $\diamond$'s). And within five days of the April 9 announcement, the fitted credit curve (the red line with $\bullet$'s) is virtually indistinguishable from its pre-pandemic estimate.

Using the calibrated model, we explain these dynamics through the interaction of two factors. First, the surge in the demand for cash in mid-March had a disproportionately large effect on the short-end of the credit curve, as investors en masse liquidated their holdings of most liquid securities (i.e., shorter-maturity high-quality, investment-grade bonds.) And second, faced with the massive sell-off across fixed income markets, broker-dealers became increasingly unwilling to take on inventory, a reduction in their risk-bearing capacity that is consistent with the spike in the credit risk premium shown in Figure 2.

Formally, we model the pandemic-induced surge in the demand for cash as an unanticipated drop in the preferred-habitat investors' demand for short-term corporate debt. Specifically, we posit a one-off increase, denoted by $\Delta\tilde{\theta}_0(\tau)$, in the intercept of the preferred-habitat investors' demand curve for corporate bonds, where $\Delta\tilde{\theta}_0(\tau)$ is a linear combination of Dirac functions with mass Δ_1 and Δ_2 at maturities τ_1 and τ_2, respectively:

$$\Delta\tilde{\theta}_0(\tau) = \Delta_1 \times \mathbb{1}_{\{\tau=\tau_1\}} + \Delta_2 \times \mathbb{1}_{\{\tau=\tau_2\}}.$$

We set $\tau_1 = 1/4$ and $\tau_2 = 5$ and calibrate Δ_1 and Δ_2 to match the increases in the one- and five-year corporate bond yields from their average pre-pandemic levels (the solid black line in Figure 8) to levels registered during the five-day window immediately preceding the March 23 announcement (the blue line with $\diamond$'s).

We assume that this negative demand shock for shorter-term corporate debt subsequently dissipates at a deterministic rate $\kappa_{\tilde{\theta}}$, the value of which is chosen to match the increase in the ten-year corporate bond yield upon the impact of shock. To capture the concomitant reduction in the broker-dealers' risk-bearing capacity—that is, their inability or unwillingness to absorb the additional inventory on their balance sheets—we assume that the risk-aversion coefficient of arbitrageurs in the corporate bond market $\tilde{a}$ jumps from its baseline value of 3.3 to 1,000.

The solid blue line in Panel A of Figure 9 shows the resulting model-implied credit curve. Its congruence with the actual fitted credit curve (the red line with $\bullet$'s) indicates that the combination of these two shocks can fully account for both the upward shift and the inversion of the credit curve during the height of the pandemic in mid-March. To match these movements in the credit curve, we must set the demand shock parameters $\Delta_1 = 0.41$ and $\Delta_2 = 1.11$, values that are about 14 and 38 times the unconditional volatility of the demand risk factor $\tilde{\beta}_t$ (i.e., $\sqrt{\sigma^2_{\tilde{\beta}}/(2\kappa_{\tilde{\beta}})} = 0.029$).[33]

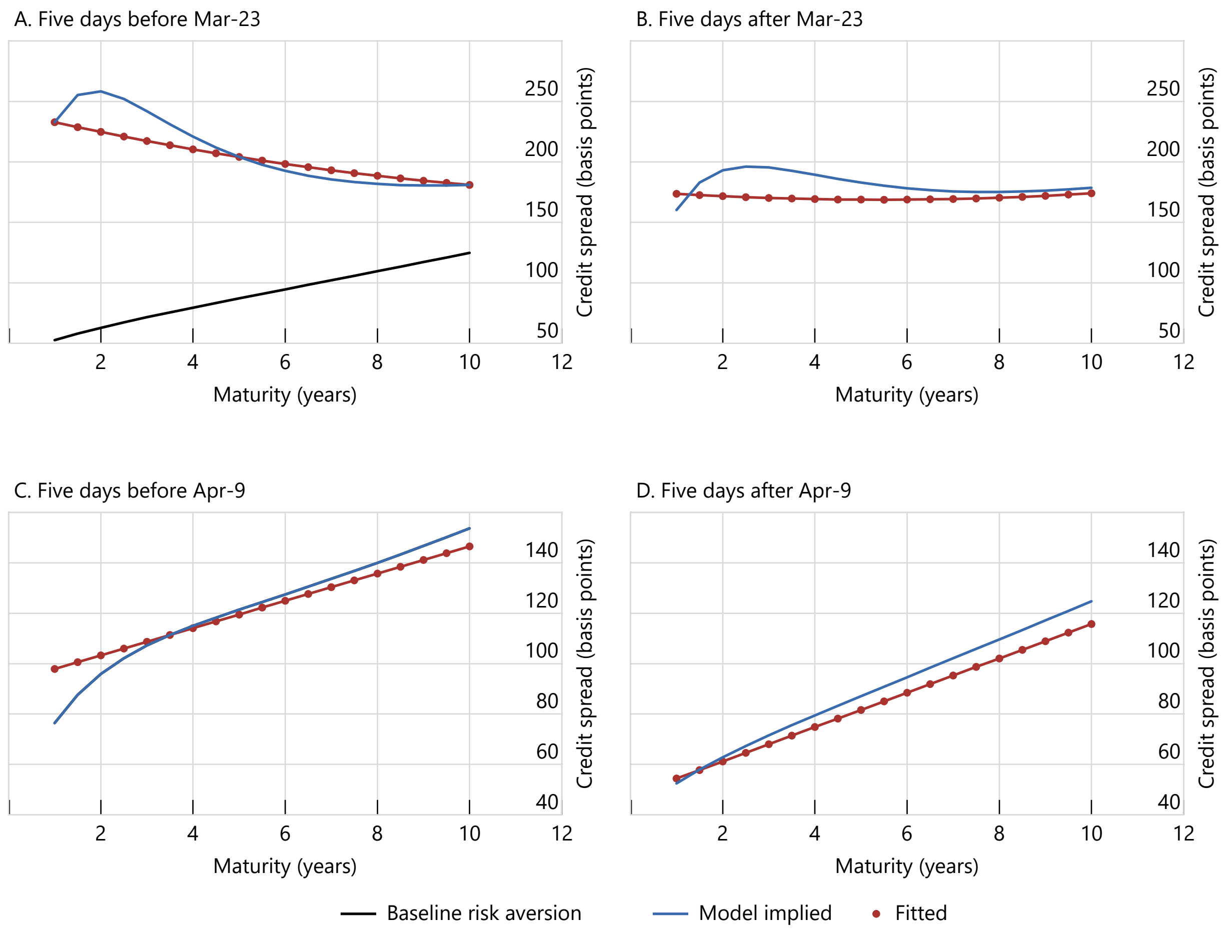
To gauge the importance of the drop in the preferred-habitat investors' demand for short-term corporate debt vis-à-vis the increase in the arbitrageurs' risk aversion, the solid black line in Panel A shows the credit curve from an experiment featuring a local demand shock of the same magnitude as before (i.e., $(\Delta_1, \Delta_2) = (0.41, 1.11)$), but where the arbitrageurs' risk aversion stays at its baseline (pre-pandemic) level of $\tilde{a} = 3.3$. Note that the location and slope of the curve from this experiment both correspond closely to the pre-pandemic fitted curve shown by the black solid line in Figure 8. This implies that the spike in the arbitrageurs' risk aversion plays a critical role in both the upward shift and the inversion of the credit curve, according to our calculations. Intuitively, with the risk-aversion coefficient at its pre-pandemic level, the arbitrageurs are willing to take exceptionally large positions in the corporate bond market and can thus greatly dampen the transmission of a negative demand shock concentrated at the short-end of the credit curve to longer-term spreads.[34]

As noted above, we model the Fed's announcements as "calming" the market by reducing the arbitrageurs' risk aversion. Viewed through this lens, we look in Panel B for the value of $\tilde{a}$ that minimizes deviations of the model-implied one-, five- and ten-year credit spreads from their empirical counterparts in the five-day window following the March 23 announcement. The result of this

[33]To match the increase in the ten-year corporate bond yield, this configuration of the shock to the preferred-habitat investors' demand for short-term corporate bonds implies the deterministic decay parameter of the shock $\kappa_{\tilde{\theta}} = 0.93$, yielding a half-life 0.74 years.

[34]In fact, to match the observed increases in the one- and five-year corporate bond yields with $\tilde{a} = 3.3$, we must set $(\Delta_1, \Delta_2) = (468.6, 129.9)$, a configuration implying a local demand shock of implausible magnitude.

Figure 9: Model-Implied vs. Actual Movements of the Credit Curve



Note: The solid blue lines show the response of the model-implied credit curve to an unanticipated drop in the preferred-habitat investors' demand for shorter-term corporate bonds and a concomitant jump in the risk-aversion coefficient of arbitrageurs in the corporate bond market. The solid black line shows the response of the model-implied credit curve assuming that the negative demand shock was not accompanied by an increase in the arbitrageurs' risk aversion. The red dotted lines show the actual fitted credit curves (see Figure 8).
Source: Authors' calculations.

experiment implies a reduction in $\tilde{a}$ from 1,000 to 285.7. The drop of such magnitude would be consistent with extensive market commentary, which at the time noted that investors read the March 23 announcement as a pledge that the Fed will do whatever it takes to keep the economy from collapsing under the weight of the pandemic.

We model movements in the credit curve around the April 9 announcement in the same vein. According to Panel C, credit spreads on high-quality investment-grade bonds narrowed further in the days leading to the April 9 announcement. This narrowing was especially pronounced at the short-end of the maturity spectrum (e.g., the one-year credit spread fell 75 basis points), resulting in a noticeable steepening of the credit curve. To generate such a downward shift and rotation of the credit curve, our calculations imply a further decline in the arbitrageurs' risk aversion, from 286.7

to 39.7, in the five days before the April 9 announcement.

As noted above, the April 9 announcement effectively restored the investment-grade credit curve to its pre-pandemic shape and level. In the context of our model, the restoration owes primarily to the continued improvement in the arbitrageurs' risk-bearing capacity, as reflected in the return of their risk-aversion coefficient $\tilde{a}$ to its pre-pandemic level of 3.3. Indeed, as shown in Panel D, this is sufficient to bring the model-implied credit curve very close to its empirical counterpart.

5 Conclusion

The aim of this paper is to quantify the announcement and purchase effects of the SMCCF on prices and market liquidity measures in the U.S. corporate bond market. Using a matched sample of program-eligible and program-ineligible securities trading in the secondary market—with both types of securities issued by the same company—and a DiD methodology that explicitly controls for the announcement-induced shifts in the credit curve, we isolate and estimate the direct effects of the March 23 and April 9 announcements on credit and bid-ask spreads.

The results from this empirical analysis indicate that the two announcements significantly reduced investment-grade credit and bid-ask spreads across the maturity spectrum—the so-called level effect. More importantly, through the so-called slope effect, the two announcements rotated the credit curve and restored the normal upward-sloping term structure of credit spreads in the investment-grade segment of the market. The March 23 announcement, in particular, lowered credit spreads on shorter-term program-eligible bonds relative to their longer-term ineligible counterparts; the April 9 announcement, by contrast, induced a steepening of the entire investment-grade credit curve, irrespective of the SMCCF's maturity-eligibility criterion.

Using a flexible empirical approach to decompose credit spreads into a component capturing issuer-specific default risk and a residual component capturing credit risk premia, we find that the announcement-induced narrowing of credit spreads is due almost entirely to a decline in credit risk premia, as opposed to a reduction in the likelihood of default. Utilizing an event-style methodology that precisely identifies the Fed's purchases of individual corporate bonds, we document that the purchases had sizable effects on credit spreads, particularly for corporate bonds at the lower-end of the investment-grade quality spectrum.

Lastly, we show that our empirical findings can be rationalized within the preferred-habitat framework of Vayanos and Vila (2021). In particular, the pandemic-induced inversion of the credit curve can be explained by a negative demand shock to the preferred-habitat investors' demand for short-term high-quality corporate bonds, coupled with a sharp increase in the arbitrageurs' risk aversion. In this framework, the Fed's subsequent announcements reduce the arbitrageurs' risk aversion and alleviate the extreme form of market segmentation. All told, our results imply that the primary effect of the Fed's announcements was to restore investor confidence and improve market sentiment, in the process making it substantially easier for companies to borrow in the corporate bond and other debt markets.

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Supplementary Material – For Online Publication Only

This section contains five appendixes (A–E). In Appendix A, we document that our baseline estimates of the SMCCF announcement effects based on the narrow treatment and control groups are robust to defining treatment and control groups so as to include issuers' all eligible and ineligible bonds, respectively. In Appendix B, we present estimates of the impact of the two announcements on pricing and liquidity of speculative-grade bonds, a category of bonds that was ineligible for purchase by the SMCCF. Appendix C provides details regarding the construction of the residual credit spreads, our proxy for the credit risk premia, while Appendix D presents additional results regarding the SMCCF's purchase effects. Lastly, Appendix E provides a detailed explanation of how we calibrated the extended version of the Vayanos and Vila (2021) preferred-habitat model.

A SMCCF Announcement Effects Based on the Full Sample

In this section, we examine the robustness of our baseline estimates of the announcement effects reported in Tables 3 and 4 to an alternative definitions of the treatment and control groups. Specifically, we re-define the narrow treatment and control groups by sorting all available bonds with remaining maturity between one and 12 years into one of the two groups. Using these full-sample definitions of the treatment and control groups, we re-estimate specification (1) and report the results in Tables A-1 and A-2. Broadly speaking, the results for both credit spreads (Table A-1) and bid-ask spreads (Table A-2) based on the full sample are fully consistent with—if anything, they strengthen—the findings reported in Tables 3 and 4 of the main text.

Turning first to credit spreads in Panel A of Table A-1, note that the estimates of β_1—the coefficient on the indicator $\mathbb{1}[t \geq t^*]$ measuring the average response of credit spreads to the two announcements—exhibit the same pattern and are of similar magnitudes as the estimates based on the narrow definitions of the treatment and control groups. In response to the March 23 announcement (columns 1–3), credit spreads narrowed, on average, 26 basis points within the two-day window before widening 56 basis points, on average, within the ten-day window. In response to the April 9 announcement (columns 4–6), by contrast, the average credit spread fell steadily: 50 basis points with the two-day window and more than 70 basis points within the ten-day window.

We see a similar pattern in the estimates of coefficient β_2 on the interaction term $\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$, which measures the additional response of credit spreads of SMCCF-eligible bonds. In response to the March 23 announcement (columns 1–3), credit spreads on such bonds narrowed an additional 48 basis points within the two-day window, with the announcement effect waning to 18 basis points within the ten-day window—note that these estimates are about twice as large as those based on the narrow sample. In contrast, the estimated effects for SMCCF-eligible bonds following the April 9 announcement based on the full sample (columns 4–6) are roughly of the same magnitude as those based on the narrow sample. The full-sample estimates imply an additional statistically significant narrowing of credit spreads for SMCCF-eligible bonds of six basis points within two days of the April 9 announcement, 18 basis points within five days, and 26 basis points within ten days of the announcement. As shown in Panel B, the inclusion of time fixed effects has a negligible effect on the estimates of β_2.

Panel A of Table A-2 reports the estimation results for the log of bid-ask spreads based on the specification without time fixed effects. As in the case of credit spreads, the estimated overall effects of the two announcements—as captured by the coefficient β_1 on the announcement indicator $\mathbb{1}[t \geq t^*]$—based on the full sample are quite similar to those based on the narrow sample. In the ten-day window, the average bid-ask spread is estimated to decline more than 15 basis points in response to the March 23 announcement (column 3), while the April 9 announcement is estimated to reduce the average bid-ask spread about 45 basis points over the same horizon (column 6).

The full-sample estimates of coefficient β_2 on the interaction term $\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$ are also comparable to their corresponding estimates based on the narrow sample. At the ten-day horizon, the full-sample estimates of β_2 imply an additional statistically significant narrowing of bid-ask spreads of nearly 15 basis points for SMCCF-eligible bonds in response to the March 23 announcement (column 3) and six basis points in response to the April 9 announcement (column 6). These results confirm that liquidity conditions for program-eligible bonds improved significantly more in response to the March 23 announcement compared with the April 9 announcement. As before, the inclusion of time fixed effects (Panel B) yields nearly identical estimates of β_2.

Table A-1: Effect of the SMCCF Announcements on Credit Spreads
(Full-Sample Treatment and Control Groups)

	Mar-23 Announcement			Apr-9 Announcement		
	2-day	5-day	10-day	2-day	5-day	10-day
Explanatory Variable	(1)	(2)	(3)	(4)	(5)	(6)
A. Without time fixed effects						
$\mathbb{1}[t \geq t^*]$	-0.26^{***}	0.06^{*}	0.59^{***}	-0.50^{***}	-0.67^{***}	-0.68^{***}
	(0.03)	(0.03)	(0.04)	(0.02)	(0.02)	(0.02)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	-0.48^{***}	-0.40^{***}	-0.23^{***}	-0.06^{**}	-0.17^{***}	-0.26^{***}
	(0.04)	(0.04)	(0.04)	(0.02)	(0.02)	(0.03)
R^2	0.71	0.66	0.61	0.91	0.90	0.88
No. of firms	747	840	887	785	872	907
No. of bonds	3,717	4,186	4,509	3,819	4,203	4,529
Observations	13,370	29,034	55,319	13,861	30,504	58,134
B. With time fixed effects						
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	-0.47^{***}	-0.38^{***}	-0.21^{***}	-0.06^{**}	-0.16^{***}	-0.25^{***}
	(0.04)	(0.03)	(0.04)	(0.02)	(0.02)	(0.03)
R^2	0.74	0.73	0.75	0.92	0.90	0.89
No. of firms	747	840	887	785	872	907
No. of bonds	3,717	4,186	4,509	3,819	4,203	4,529
Observations	13,370	29,034	55,319	13,861	30,504	58,134

Note: The dependent variable in all specifications is $\text{CS}_{i,j,t}$, the credit spread of bond j (issued by firm i) on business day t. The entries in the table denote the OLS estimates of coefficients associated with the specified explanatory variable: $\mathbb{1}[t \geq t^*] = 0/1$-indicator that equals one if date t is greater than or equal to the specified announcement date t^* and zero otherwise; and $\mathbb{1}[j = E] = 0/1$-indicator that equals one if bond j was eligible for purchase by the SMCCF as of March 22 and zero otherwise. All specifications include a vector of bond-specific controls (not reported) and issuer fixed effects. Asymptotic standard errors reported in parentheses are clustered at the issuer level: * $p < .10$; ** $p < .05$; and *** $p < .01$.

TABLE A-2: Effect of the SMCCF Announcements on Bid-Ask Spreads (Full-Sample Treatment and Control Groups)

	Mar-23 Announcement			Apr-9 Announcement		
	2-day	5-day	10-day	2-day	5-day	10-day
Explanatory Variable	(1)	(2)	(3)	(4)	(5)	(6)
A. Without time fixed effects						
$\mathbb{1}[t \geq t^*]$	−0.38***	−0.26***	−0.13***	0.03	−0.16***	−0.44***
	(0.04)	(0.03)	(0.02)	(0.04)	(0.02)	(0.02)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	−0.01	−0.17***	−0.14***	−0.05	−0.06*	−0.06**
	(0.05)	(0.03)	(0.03)	(0.05)	(0.03)	(0.02)
R^2	0.13	0.14	0.14	0.34	0.34	0.35
No. of firms	448	565	654	536	630	706
No. of bonds	2,117	2,657	3,133	2,383	2,893	3,323
Observations	5,830	12,824	24,821	6,301	14,544	28,393
B. With time fixed effects						
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	−0.02	−0.17***	−0.10***	−0.05	−0.06*	−0.06**
	(0.05)	(0.03)	(0.03)	(0.05)	(0.03)	(0.02)
R^2	0.15	0.18	0.25	0.34	0.35	0.38
No. of firms	448	565	654	536	630	706
No. of bonds	2,117	2,657	3,133	2,383	2,893	3,323
Observations	5,830	12,824	24,821	6,301	14,544	28,393

NOTE: The dependent variable in all specifications is $\ln \text{BAS}_{i,j,t}$, log of the bid-ask spread of bond j (issued by firm i) on business day t. The entries in the table denote the OLS estimates of coefficients associated with the specified explanatory variable: $\mathbb{1}[t \geq t^*] = 0/1$-indicator that equals one if date t is greater than or equal to the specified announcement date t^* and zero otherwise; and $\mathbb{1}[j = E] = 0/1$-indicator that equals one if bond j was eligible for purchase by the SMCCF as of March 22 and zero otherwise. All specifications include a vector of bond-specific controls (not reported) and issuer fixed effects. Asymptotic standard errors reported in parentheses are clustered at the issuer level: * $p < .10$; ** $p < .05$; and *** $p < .01$.

B The SMCFF Announcement Effects on Speculative-Grade Bonds

Recall that the SMCCF was expanded on April 9, 2020, to include outstanding bonds with remaining maturity of less than five years issued by fallen angels, companies rated as investment grade on March 22, 2020, but which were subsequently downgraded to the Ba/BB speculative-grade category; the facility was not extended to the maturity-eligible outstanding bonds issued by companies with a speculative-grade rating as of March 22, 2020. Table B-1 reports estimation results of our baseline specification (1), using narrow definitions of the treatment and control groups based on speculative-grade bonds.

TABLE B-1: Effect of the SMCCF Announcements on Speculative-Grade Bonds (Five-Day Window Bracketing Each Announcement)

	Credit Spreads		Bid-Ask Spreads	
	Mar-23	Apr-9	Mar-23	Apr-9
Explanatory Variable	(1)	(2)	(3)	(4)
A. Without time fixed effects				
$\mathbb{1}[t \geq t^*]$	0.52^{***}	-1.42^{***}	0.10	-0.22^{***}
	(0.09)	(0.08)	(0.07)	(0.05)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	−0.11	-0.33^{***}	0.01	0.02
	(0.08)	(0.09)	(0.10)	(0.09)
R^2	0.82	0.91	0.19	0.27
No. of firms	186	196	154	161
No. of bonds	483	492	458	480
Observations	3,072	3,242	1,395	1,526
B. With time fixed effects				
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	−0.11	-0.33^{***}	0.01	0.02
	(0.08)	(0.09)	(0.10)	(0.09)
R^2	0.87	0.91	0.19	0.27
No. of firms	186	196	154	161
No. of bonds	483	492	458	480
Observations	3,072	3,242	1,395	1,526

NOTE: The dependent variable in specifications (1) and (2) is $\mathrm{CS}_{i,j,t}$, the credit spread of bond j (issued by firm i) on business day t, while in specifications (3) and (4), the dependent variables is $\ln \mathrm{BAS}_{i,j,t}$, the log of the corresponding bid-ask spread. The entries in the table denote the OLS estimates of coefficients on the specified explanatory variable: $\mathbb{1}[t \geq t^*] = 0/1$-indicator that equals one if date t is greater than or equal to the specified announcement date t^* and zero otherwise; $\mathbb{1}[j = E] = 0/1$-indicator variable that equals one if the remaining maturity of bond j on March 22 was less than or equal to five years and zero otherwise. All specifications include a vector of bond-specific controls (not reported) and issuer fixed effects. Asymptotic standard errors reported in parentheses are clustered at the issuer level: * $p < .10$; ** $p < .05$; and *** $p < .01$.

C Residual Credit Spreads

This appendix provides details underlying the construction of credit spread residuals, our proxy for credit risk premia. Because we use information from equity markets to infer issuer-specific default risk, we restrict our sample of bonds to those issued by U.S. investment-grade publicly listed companies. To avoid any look-ahead bias when constructing credit risk premia, we use daily data between June 2002 and December 2019 to estimate the coefficients of specification (3) in the main text. Using these estimates, we then compute the predicted credit spreads, denoted by $\widehat{\mathrm{CS}}_{i,j,t}$, from January 2020 through the end of July 2020. The credit spread residual for a given bond is thus the difference between the actual credit spread $\mathrm{CS}_{i,j,t}$ and its predicted value $\widehat{\mathrm{CS}}_{i,j,t}$ (see Gilchrist and Zakrajšek, 2012, for details).

For each publicly listed firm in our sample, we measure its default risk by the standard "distance-to-default" (DD) framework developed in the seminal work of Merton (1974). Specifically, the daily firm-specific distance-to-default over the horizon of τ years is given by

$$\mathrm{DD}^{(\tau)} = \frac{\ln(V/D) + \left(\mu_V - 0.5\sigma_V^2\right)\tau}{\sigma_V\sqrt{\tau}}, \tag{C-1}$$

where V is the market value of the firm's assets, D is the face value of its debt—the so-called default point—and μ_V and σ_V denote the expected growth rate and the volatility of the firm's value, respectively. Following standard practice, we calibrate the default point D to the firm's current liabilities plus one-half of its long-term liabilities.

For each firm on each day, we infer V, μ_V, and σ_V using an iterative procedure proposed by Bharath and Shumway (2008). First, we initialize the procedure by letting $\sigma_V = \sigma_E\,[D/(E+D)]$, where E denotes the market value of the firm's equity and σ_E denotes the volatility of its equity. We estimate σ_E from historical daily stock returns using a 250-day moving window. Using this initial value of σ_V, we infer the market value of the firm for every day of the 250-day moving window based on the following equation for the value of the firm's equity implied by the Merton model:

$$E = V\Phi(\delta_1) - e^{-r\tau}D\Phi(\delta_2), \tag{C-2}$$

where r denotes the instantaneous risk-free interest rate (one-year U.S. Treasury yield), $\Phi(\cdot)$ is the cumulative standard normal distribution function, and

$$\delta_1 = \frac{\ln(V/D) + \left(r + 0.5\sigma_V^2\right)\tau}{\sigma_V\sqrt{\tau}} \quad \text{and} \quad \delta_2 = \delta_1 - \sigma_V\sqrt{\tau}.$$

Second, we calculate the implied daily log-return on assets (i.e., $\Delta\ln V$) and use the resulting series to generate new estimates of σ_V and μ_V. We then iterate on σ_V until convergence.

In addition to this firm-specific market-based measure of default risk ($\mathrm{DD}_{i,t}^{(\tau)}$), the bond-level credit-spread pricing regression (3) in the main text also includes the following bond-specific characteristics as controls: the bond's duration ($\mathrm{DUR}_{i,j,t}$), the par amount ($\mathrm{PAR}_{i,j}$), the bond's (fixed) coupon rate ($\mathrm{COUP}_{i,j}$), and the age of the issue ($\mathrm{AGE}_{i,j,t}$). As shown in Table C-1, the distance-to-default is a highly significant predictor of the (log) credit spreads: a decrease of one standard deviation in the distance-to-default $\mathrm{DD}_{i,t}^{(\tau)}$ leads to a widening of credit spreads of about 9 basis points. Moreover, this market-based indicator of default risk, together with other observable bond characteristics, explains a considerable portion of variation in daily (log) credit spreads over the June 2002 to December 2019 period.

To get a sense of how this measure of default risk reacted to the March 23 and April 9 announcements, Figure C-1 plots the kernel density estimates (i.e., smoothed histograms) of the

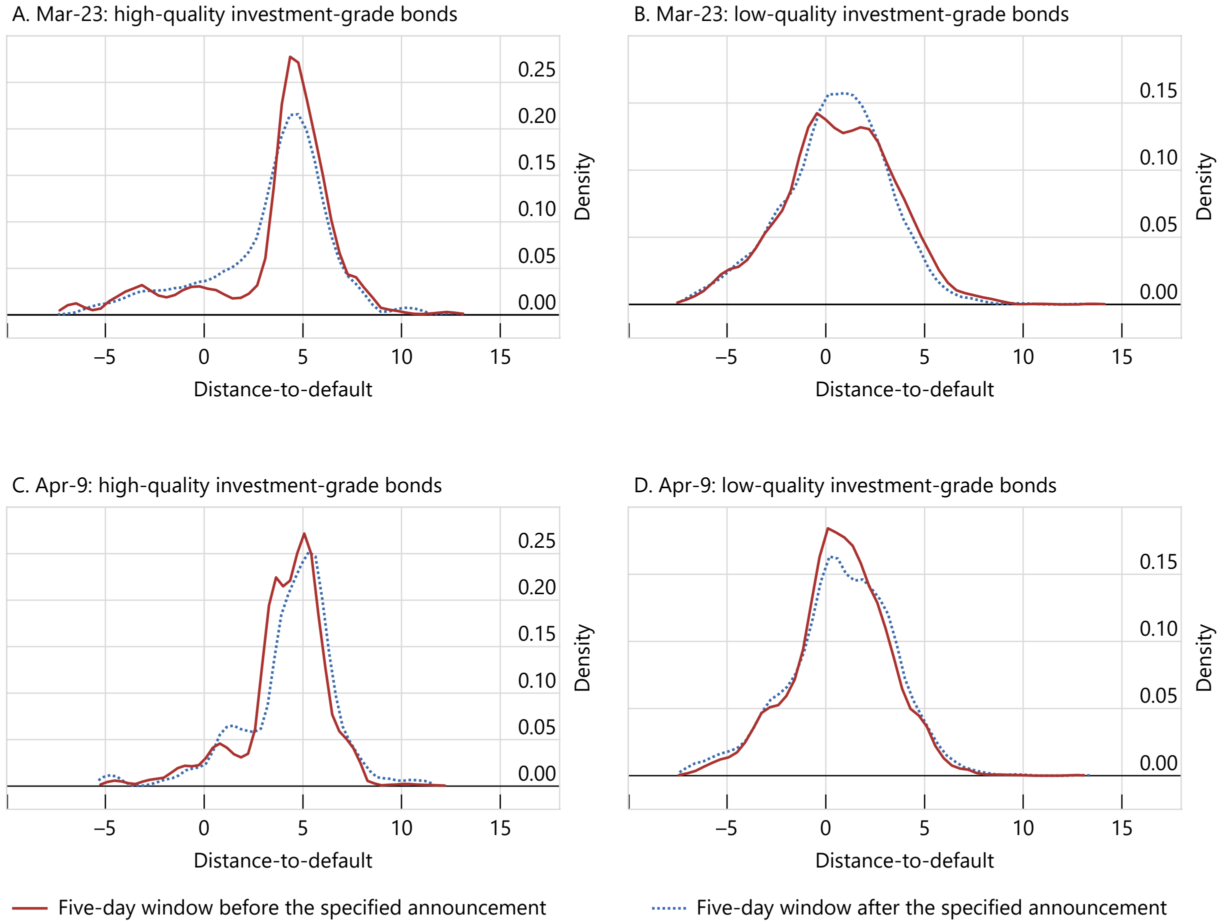
TABLE C-1: Credit Spreads and the Distance-to-Default

Explanatory Variable	*Coeff.*	*Std. Err.*
$-\text{DD}_{i,t}^{(\tau)}$	0.042***	0.003
$\ln \text{DUR}_{i,j,t}$	0.005	0.015
$\ln \text{PAR}_{i,j}$	−0.068***	0.015
$\ln \text{COUP}_{i,j}$	1.113***	0.029
$\ln \text{AGE}_{i,j,t}$	−0.073***	0.007
R^2	0.43	
No. of firms	1,648	
No. of bonds	18,730	
Observations	10,217,485	

NOTE: Sample period: daily data from June 1, 2002 to December 31, 2019. The dependent variable is $\ln \text{CS}_{i,j,t}$, the log of the credit spread on bond j (issued by firm i) on day t. Asymptotic standard errors are clustered in both the firm (i) and time (t) dimensions, according to Cameron et al. (2011).

distance-to-default $\text{DD}_{i,t}^{(\tau)}$ for our sample of investment-grade bonds. Panels A and B focus on the five-day windows bracketing the March 23 announcement, with panel A showing the pre- and post-announcement distributions of $\text{DD}_{i,t}^{(\tau)}$ for the high-quality investment-grade bonds and panel B showing the corresponding distributions for the low-quality investment-grade bonds; panels C and D contain the same information for the April 9 announcement.

Figure C-1: Default Risk Around the SMCCF Announcements

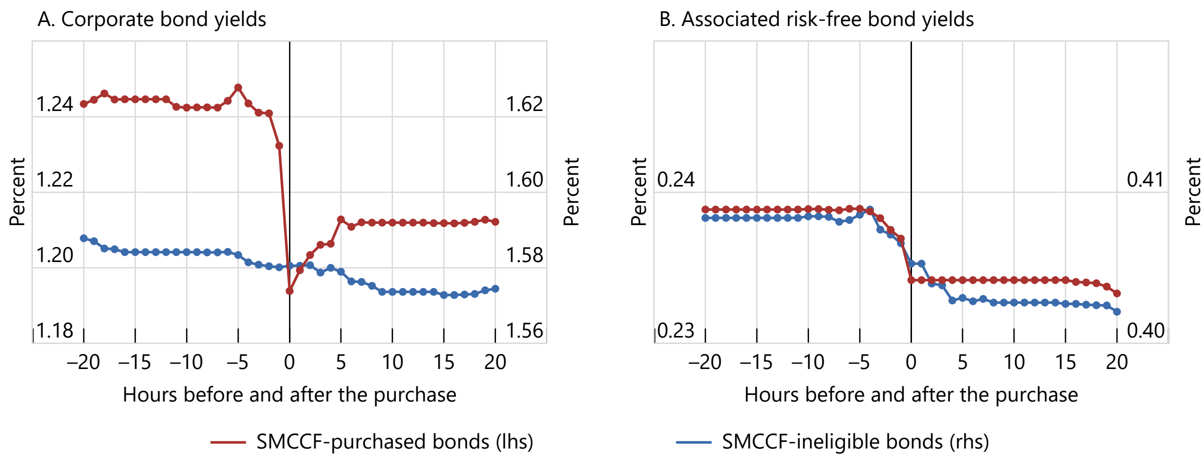


Note: The panels show the weighted kernel density estimates of the distance-to-default $DD_{i,t}^{(\tau)}$ (a measure of default risk on day t of a bond with the remaining maturity of τ years issued by an investment-grade company i) in the five-day windows bracketing the March 23 (Panels A and B) and April 9 (Panels C and D) SMCCF announcements. The five (business) day window before March 23 runs from March 16 to March 22, while the five (business) day window after March 23 runs from March 23 to March 30. The corresponding five (business) day windows bracketing the April 9 announcements are April 2 to April 8 and April 9 to April 17. High-quality investment-grade bonds are those rated as Aaa/AAA or Aa/AA, whereas low-quality investment-grade bonds are those rated A/A or Baa/BBB. The distance-to-default data are trimmed at P1 and P99 percentiles. In computing the kernel density estimates, the par amount outstanding of each bond issue are used as weights.

D The SMCCF Purchase Effects

In this appendix, we provide additional event-study results of the SMCCF's purchase effects. First, we decompose the average purchase effect on credit spreads into a portion attributable to a change in the corporate bond yields and a portion attributable to a change in the associated risk-free rates. Second, we estimate separate purchase effects for high- and low-quality investment-grade bonds.

FIGURE D-1: The SMCCF's Purchase Effects on Bond Yields



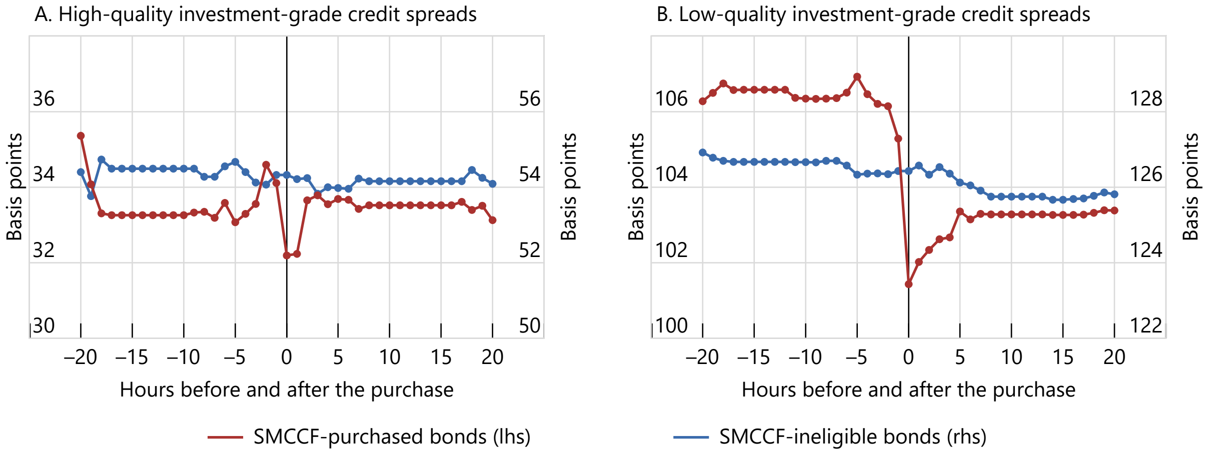
NOTE: The red line in Panel A shows the average yield-to-maturity on bonds purchased by the SMCCF in a symmetric 20-hour window bracketing their purchases, while the blue line shows the corresponding average yield-to-maturity on the SMCCF-ineligible bonds issued by the same set of issuers. The corresponding lines in Panel B show the average yield-to-maturity on the associated comparable-maturity synthetic Treasury securities.
SOURCE: Authors' calculations using TRACE date and data from the Federal Reserve Bank of New York.

The red line in Panel A in Figure D-1 shows the average yield on corporate bonds purchased by the SMCCF within the event window that spans 20 hours before and 20 hours after the purchase time, which, as in Figure 6 of the main text, is normalized to be equal to zero; the blue line shows the corresponding average yield in the control group of corporate bonds—that is, bonds issued by the same set of companies but whose remaining maturity is greater than five years. Panel B depicts the same time-series dynamics of the associated comparable-maturity (synthetic) Treasury yields, which are used to construct credit spreads for each security.

The results confirm that the difference in the average purchase effect on credit spreads between program-eligible and ineligible bonds is almost completely attributable to the differential purchase effect on the bonds' corporate bond yields. That is, the facility's purchases of corporate bonds lowered yields on the purchased bonds and, as expected, had a negligible effect on comparable-maturity risk-free yields.

Figure D-2 shows the average purchase effects for high-quality (Panel A) and low-quality (Panel B) investment-grade bonds. According to Panel A, spreads on high-quality bonds purchased by the facility declined about two basis point upon purchase, whereas those on low-quality purchased bonds declined nearly four basis points. A portion of these purchase-induced declines was reversed over the subsequent five hours, leaving credit spreads on high-quality purchased bonds about unchanged and those on low-quality purchased bonds down about three basis points, on net.

Figure D-2: The SMCCF's Purchase Effect on Credit Spreads by Issuer Credit Quality



Note: The red line in Panel A shows the average credit spread on high-quality investment-grade bonds purchased by the SMCCF in a symmetric 20-hour window bracketing their purchase, while the blue line shows the corresponding average credit spread on the SMCCF-ineligible bonds issued by the same set of issuers. The corresponding lines in Panel B shows the same information for low-quality investment-grade bonds.
Source: Authors' calculations using TRACE date and data from the Federal Reserve Bank of New York.

E Calibration Summary

In this appendix, we detail our calibration procedure, which, as noted in the main text, closely follows Vayanos and Vila (2021). Given the parallel structure of Treasury and corporate bond markets in our setup, the equilibrium conditions governing price dynamics in both markets are the same as those derived by Vayanos and Vila (2021), and we refer the interested reader to their paper for details.

TABLE E-1: Model Parameters and Targeted Moments

Parameter	Value	Empirical Moment	Value
Unconditional average of r_t		Average 1-year yield (level)	
$\bar{r}$	0.013	$\text{Ave}(y_t^{(1)})$	1.529
Mean reversion of r_t		Volatility of 1-year yield (level)	
κ_r	0.302	$\sqrt{\text{Var}(y_t^{(1)})}$	1.510
Diffusion of r_t		Volatility of 1-year yield (changes)	
σ_r	0.014	$\sqrt{\text{Var}(y_{t+12}^{(1)} - y_t^{(1)})}$	1.087
Mean reversion − demand factors		Average volatility of yield levels	
κ_β	0.189	$\frac{1}{20}\sum_{\tau=1}^{20}\sqrt{\text{Var}(y_t^{(\tau)})}$	1.169
$\kappa_{\tilde{\beta}}$	0.116	$\frac{1}{20}\sum_{\tau=1}^{20}\sqrt{\text{Var}(\tilde{y}_t^{(\tau)})}$	1.260
Risk aversion × demand intercept		Average volatility of yield changes	
$a \times \theta$	8462.6	$\frac{1}{20}\sum_{\tau=1}^{20}\sqrt{\text{Var}(y_{t+12}^{(\tau)} - y_t^{(\tau)})}$	0.734
$\tilde{a} \times \tilde{\theta}$	2542.8	$\frac{1}{20}\sum_{\tau=1}^{20}\sqrt{\text{Var}(\tilde{y}_{t+12}^{(\tau)} - \tilde{y}_t^{(\tau)})}$	0.685
Risk aversion × demand slope		Average correlation of yield changes	
$a \times \alpha$	39.5	$\frac{1}{20}\sum_{\tau=1}^{20}\text{Corr}(y_{t+12}^{(1)} - y_t^{(1)}, y_{t+12}^{(\tau)} - y_t^{(\tau)})$	0.590
$\tilde{a} \times \tilde{\alpha}$	17.1	$\frac{1}{20}\sum_{\tau=1}^{20}\text{Corr}(\tilde{y}_{t+12}^{(1)} - \tilde{y}_t^{(1)}, \tilde{y}_{t+12}^{(\tau)} - \tilde{y}_t^{(\tau)})$	0.615
Demand shock − short maturities		Relative volume − short maturities	
δ_α	0.351	$\frac{\sum_{0<\tau\le 2}\text{Volume}(\tau)}{\sum_{0<\tau\le 20}\text{Volume}(\tau)}$	0.199
$\tilde{\delta}_\alpha$	0.227	$\frac{\sum_{0<\tau\le 5}\text{Volume}(\tau)}{\sum_{0<\tau\le 20}\text{Volume}(\tau)}$	0.399
Demand shock − long maturities		Relative volume − long maturities	
δ_θ	0.361	$\frac{\sum_{11<\tau\le 20}\text{Volume}(\tau)}{\sum_{0<\tau\le 20}\text{Volume}(\tau)}$	0.094
$\tilde{\delta}_\theta$	0.237	$\frac{\sum_{10<\tau\le 20}\text{Volume}(\tau)}{\sum_{0<\tau\le 20}\text{Volume}(\tau)}$	0.309
Demand slope		Demand elasticity	
α	5.21	Estimate in KVJ (2012)	−0.746
$\tilde{\alpha}$	5.21	Estimate in KVJ (2012)	−0.746
Risk aversion × demand intercept		Average 5-year yield (level)	
$a \times \theta_0$	294.6	$\text{Ave}(y_t^{(5)})$	2.386
$\tilde{a} \times \tilde{\theta}_0$	208.9	$\text{Ave}(\tilde{y}_t^{(5)})$	3.190

NOTE: The entries in the table denote the calibrated values of the model parameters and the corresponding targeted moments in the data. All data are at monthly frequency (see the main text for details).

As shown in Table E-1, the parameters governing the dynamics of the short rate process, κ_r

and σ_r, are calibrated to match the volatility of one-year Treasury yield, $y_t^{(1)}$, and the volatility of its annual changes, $y_{t+12}^{(1)} - y_t^{(1)}$, respectively. As a normalization, we set the volatility of the demand-risk factors in both markets, σ_β and $\sigma_{\tilde{\beta}}$, to be equal to the volatility of the short rate σ_r.

The parameters κ_β, a, θ, $(a \times \alpha)$ are chosen to match the average of volatilities of Treasury yields and the average volatilities of their annual changes across maturities, as well as the average of correlations between annual changes in the one-year yield and annual changes in yields of other maturities. We pick values for the parameters δ_α and δ_θ to match the relative trading volumes of Treasuries with maturities of less than or equal to two years and with maturities of more than ten years, respectively.[35] And the parameters $\bar{r}$ and θ_0 are calibrated to match the average of the one- and five-year Treasury yields, respectively. We target the corresponding set of empirical moments from the corporate bond market when calibrating $\kappa_{\tilde{\beta}}$, $\tilde{a}$, $\tilde{\theta}_0$, $\tilde{\delta}_\alpha$, and δ_θ. Lastly, we set $\alpha = \tilde{\alpha} = 5.21$, which corresponds to the demand elasticity estimated by Krishnamurthy and Vissing-Jorgensen (2012) (KVJ).

The results of this calibration exercise imply a risk-aversion coefficient of $7.6 = 39.5/5.21$ for arbitrageurs in the Treasury bond market and a risk-aversion coefficient of $3.3 = 17.1/5.21$ for arbitrageurs in the corporate bond market. Consistent with the fact that credit spreads are typically positive, the parameter of the demand intercept in the Treasury market θ_0 is calibrated to be $38.8 = 294.6/7.6$, which is considerably lower than its counterpart in the corporate bond market, $\tilde{\theta}_0 = 63.7 = 210.3/3.3$. The calibration results also imply that the demand risk factor in the corporate bond market has a lower mean-reversion rate than its counterpart in the Treasury bond market (i.e., $\kappa_{\tilde{\beta}} = 0.116$ vs. $\kappa_\beta = 0.189$). This consistent with dislocations in the corporate bond market dissipating more slowly than in the Treasury bond market, owing to the fact that the former market is less liquid than the latter.

[35] The data on the trading volume of Treasuries and corporate bonds come from the Federal Reserve Bank of New York. We use data on the nominal U.S. Treasury bond trading volumes from April 2013 to January 2020 and data on the investment-grade corporate bond trading volumes from January 2015 to January 2020.