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

Simon Gilchrist, Bin Wei, Vivian Z. Yue, and Egon Zakrajšek

Abstract:
This paper evaluates the efficacy of the Secondary Market Corporate Credit Facility, a program designed to stabilize the U.S. corporate bond market during the COVID-19 pandemic. The program announcements on March 23 and April 9, 2020, significantly reduced investment-grade credit spreads across the maturity spectrum—irrespective of the program's maturity-eligibility criterion—and ultimately restored the normal upward-sloping term structure of credit spreads. The Federal Reserve's actual purchases reduced credit spreads of eligible bonds 3 basis points more than those of ineligible bonds, a sizable effect given the modest volume of purchases. A calibrated variant of the preferred habit model shows that a "dash for cash"—a selloff of shorter-term lowest-risk investment-grade bonds—combined with a spike in the arbitrageurs' risk aversion, can account for the inversion of the investment-grade credit curve during the height of 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.

JEL Classifications: E44, E58, G12, G14

Keywords: COVID-19, SMCCF, credit spreads, credit market support facilities, event study, purchase effects, preferred habitat

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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). As the acute risk-off sentiment swept through financial markets in early March 2020, prices of corporate bonds fell, and credit spreads increased sharply.[1]

The Federal Reserve reacted swiftly to the turmoil roiling financial markets by unveiling an 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 continued to deteriorate, and credit spreads widened further. In response, the Fed announced at 8 a.m. (EST) 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). 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 nearly $10 trillion market for outstanding corporate bonds.

The wide-ranging scope of the announcement, characterized by market participants as "whatever it takes" and "throwing the kitchen sink" at the markets, had an immediate effect, boosting stock prices, raising 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 the morning of April 9, 2020, announced updated terms for the two corporate bond-buying facilities. In particular, the Fed indicated that P/SMCCF-eligible issuers now included companies recently downgraded from investment grade to "junk," the so-called fallen angels.

This paper evaluates the efficacy of the SMCCF and analyzes the mechanism(s) through which it affected the pricing of investment-grade corporate bonds in the secondary market. We focus on the SMCCF because of its historic importance: It represents the first time the Fed directly supported corporate credit markets by signaling a willingness to purchase outstanding corporate debt and potentially assume a material amount of credit risk on its balance sheet. Understanding the efficacy of such programs and the channels through which they affect markets is thus critical for policy going forward.[2]

Formally evaluating the impact of the SMCCF on the corporate bond market is complicated by the fact that the Fed announced and expanded the SMCCF in conjunction with several other

[1] Several 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, 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—even though the underlying assets are significantly less liquid—the resulting "liquidity mismatch" made these funds especially vulnerable to runs (see Falato et al., 2021).

[2] 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. In the past, the Bank of Japan, the Bank of England, and the European Central Bank—when constrained by the effective lower bound on nominal interest rates—have all launched similar corporate bond-buying programs to ease broad financial conditions and stimulate their economies (see Suganuma and Ueno, 2018; D'Amico and Kaminska, 2019; Grosse-Rueschkamp et al., 2019; Todorov, 2020; Adelino et al., 2023).

emergency measures.[3] However, in announcing the SMCCF, the Fed imposed a specific eligibility criterion: outstanding investment-grade bonds had to have a remaining maturity of less than or equal to five years when purchased by the facility. In principle, this maturity-eligibility criterion could be used to isolate the price impact of the two SMCCF announcements in the secondary market.

A potential problem with this identification strategy is that in early March 2020, the relationship between investment-grade credit spreads and the bonds' remaining maturity—the so-called credit curve, which is normally upward sloping—inverted abruptly, with the long-short credit spread differential dropping deep into negative territory. The inversion was especially pronounced in the safest segment of the investment-grade portion of the market, as investors amid the panic rushed to sell first their most liquid securities, namely shorter-maturity high-quality investment-grade bonds. As shown below, the pandemic-induced inversion of the credit curve presents an important confounding factor in any analysis that relies on the SMCCF's maturity-eligibility criterion to estimate the effects of the program announcements on corporate bond prices. At the same time, the potential announcement-induced shifts in the credit curve highlight a powerful channel through which announcements of such policies can affect credit markets in times of widespread distress.

To identify the efficacy of the SMCCF amid these confounding factors, we construct a matched sample of program-eligible and -ineligible corporate bonds issued by the *same* firm. Using this panel data set, we estimate the effects of the March 23 and April 9 announcements by comparing post-announcement responses of credit spreads to those within a two-day pre-announcement window, a period that excludes the Fed's slew of early actions aimed at stabilizing financial markets.[4] Moreover, we control for the confounding effects associated with the inversion of the credit curve by estimating the price impact of the two announcements using a specification that relies on the program's maturity-eligibility criterion augmented with maturity interaction terms. These terms allow the slope of the curve to rotate in the post-announcement window and across the program-eligible and -ineligible bonds.

The baseline estimates, which do not control for the pandemic-induced inversion of the credit curve, indicate that the March 23 and April 9 announcements reduced credit spreads on program-eligible bonds significantly more than those of their ineligible counterparts. At first glance, this result suggests that program eligibility played an important role in the efficacy of the program. However, when we control for the announcement-induced shifts in the slope of the credit curve, there is no differential effect on the *level* of credit spreads around the five-year maturity cutoff. Rather, the analysis shows that the March 23 announcement induced a significant differential shift

[3]In announcing the two corporate bond-buying facilities on March 23, the Fed also revived the Term Asset-Backed Securities Loan Facility, expanded its quantitative easing (QE) program to include purchases of commercial mortgage-backed securities, and noted that it expected to announce shortly the Main Street Business Lending Program to support credit to small and medium-sized businesses. The April 9 announcement was accompanied by the establishment of the Municipal Liquidity Facility and the Paycheck Protection Program Liquidity Facility.

[4]Early in the pandemic, the Fed cut the federal funds rate to zero after its meetings on March 3 and March 15 and launched QE on March 15; on March 17, it revived the Commercial Paper Funding Facility and the Primary Dealer Credit Facility, followed by the Money Market Mutual Fund Liquidity Facility on March 18.

in the *slope* of the credit curve for program-eligible bonds compared with the slope of their ineligible counterparts.

The analysis also shows that the April 9 announcement led to a significant downshift and a uniform steepening of the entire investment-grade credit curve. In combination, these results strongly support the notion that the efficacy of the SMCCF reflected a broader mechanism, one that restored the normal upward slope of the investment-grade credit curve and had little to do with the program's eligibility criterion.

The announcement-induced downshift and steepening of the credit curve—irrespective of the maturity-eligibility criterion—argues against the presence of significant clientele effects, whereby in designing such programs eligibility criteria matter in the sense that they are part of the causal mechanism through which announcements of programs affect the prices of eligible assets. As emphasized by Hanson et al. (2020), the establishment and announcement of credit-support facilities provide 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 financial markets. The fact that the two announcements significantly rotated the entire investment-grade credit curve and ultimately restored the normal upward-sloping term structure of credit spreads—again, irrespective of the program's maturity-eligibility criterion—is consistent with this broader mechanism. And so are the results, which show 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.

The final empirical exercise quantifies the effect of the facility's actual purchases of individual corporate bonds on credit spreads. Using intraday transactions data that precisely identify the time of the facility's purchases of individual securities, we show that upon purchase, spreads of program-eligible bonds narrowed, on average, 11 basis points more than those of their ineligible counterparts. While some of this decline was reversed within the next hour, credit spreads of purchased bonds remained 3 basis points below those of ineligible bonds even six to eight hours after the purchase. Concentrated entirely at the lower end of the investment-grade quality spectrum, the differential of 3 basis points represents a sizable effect given the modest size—by QE standards—of the Fed's actual purchases.

Lastly, we synthesize the above empirical findings through the lens of the preferred habitat framework of Vayanos and Vila (2021). We extend the authors' 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. Quantitatively, 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 dash for cash—and a simultaneous jump in the arbitrageurs' risk aversion. In combination, these two shocks generate a localized clientele effect concentrated at the short end of the market that is sufficient to invert the credit curve as seen in the data.

Consistent with the empirical results, the Fed's subsequent announcements are modeled 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. All told, this paper speaks 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 can do less and still deliver on its promise.

Relation to literature: This paper contributes to the literature on the pandemic-induced dislocations in the U.S. corporate bond market and the Fed's response to the crisis. 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 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. Lastly, Boyarchenko et al. (2023) document a marked, though gradual, improvement in both primary and secondary market conditions after program announcements, attributing this to the increased willingness of dealers to underwrite issuances and intermediate in secondary markets.

Compared with the above papers, our paper is the first to use the SMCCF's eligibility criteria to identify how the program announcements affected pricing conditions in the secondary market. More importantly, in doing so, we control for the confounding effects associated with the pandemic-induced inversion of the investment-grade credit curve. By explicitly taking into account the announcement-induced rotations of the credit curve, our results argue against the importance of clientele effects and delineate a broader mechanism through which the two announcements affected the pricing of bonds in the secondary market. Furthermore, we use a calibrated preferred habitat framework to quantitatively illustrate the mechanism through which such credit-easing programs affect market prices, through both the announcements and the actual purchases. This novel exercise highlights the importance of a broad-based reduction in risk aversion as the main channel through which the SMCCF announcements stabilized conditions in the corporate bond market.

The paper is also related to the papers that study other aspects of the pandemic-induced market turmoil. Haddad et al. (2021) document that through mid-March of 2020, corporate bonds traded at a sizable discount compared with their corresponding credit default swap contracts. The discount was especially pronounced for higher-rated bonds, indicating a localized demand shock in the short-term investment-grade segment of the market: Investors, in need of cash, primarily sold safer and more liquid (that is, short-term) corporate bonds as opposed to relatively more illiquid credit derivatives. The extended preferred habitat framework allows us to capture and calibrate this adverse demand shock, as well as the sharp concomitant increase in the arbitrageurs' risk aversion, both of which are required to account for the observed shifts in the credit curve. The increase in the arbitrageurs' risk aversion is consistent with the findings of 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.

2 The SMCCF Announcement and Purchase Effects

On March 23, 2020, the Federal Reserve announced the establishment of the Primary Market Corporate Credit Facility and the Secondary Market Corporate Credit Facility, through which it committed to buying a potentially substantial amount of corporate debt in both the primary and secondary markets. The announcement stipulated that the two facilities were open to only investment-grade U.S. companies and that eligible bonds had to have a remaining maturity of five years or less. Less than two weeks later, on April 9, the Fed announced that it was opening the two facilities to 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 (see Appendix A for a detailed discussion).

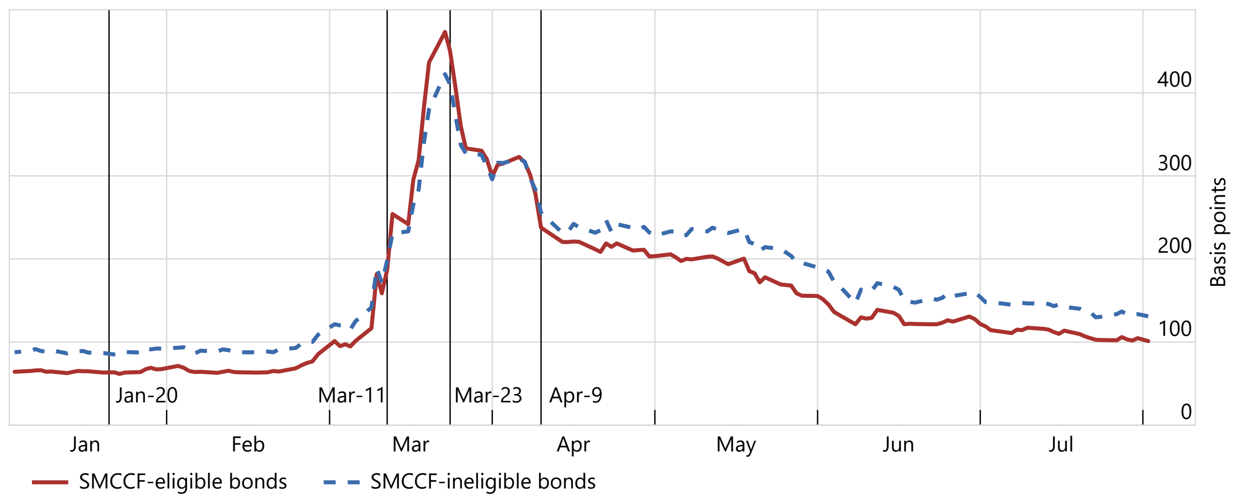
In the remainder of this section, we use daily transaction-level bond data to quantify how the March 23 and April 9 announcements affected credit spreads across the program-eligible and -ineligible segments of the corporate bond market. This analysis delineates the key way that the Fed's announcements affected the pricing of investment-grade corporate securities in the secondary market. In particular, it takes into account the impact of the two announcements on the slope of the credit curve within the program-eligible and -ineligible segments of the market. Furthermore, we examine the importance of the credit risk premium channel in these announcement effects, based on a decomposition of credit spreads by Gilchrist and Zakrajšek (2012). Lastly, we quantify the facility's actual purchases on credit spreads using an intraday event-study methodology.

2.1 Data Sources

The pricing data used in the analysis come from the regulatory version of Trade Reporting and Compliance Engine (TRACE) database, which is maintained by the Financial Industry Regulatory Authority (FINRA) and contains information about individual corporate bond transactions in the secondary market. After running the raw TRACE data through filters developed by Dick-Nielsen and Poulsen (2019), the resulting security-level transactions data are combined with the information from the Mergent's Fixed Income Securities Database (FISD) 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 coupon schedules that were issued by investment-grade U.S. companies and drop all transactions

Figure 1: Investment-grade Credit Spreads during the COVID-19 Pandemic



Note: The red solid line shows the daily average credit spread of SMCCF-eligible corporate bonds, while the dashed blue line shows the daily average credit spread of SMCCF-ineligible corporate bonds, with both types of securities issued by the same set of U.S. companies. 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/FINRA and Mergent FISD/WRDS.

involving bonds with a remaining maturity shorter that six months or longer than 12 years. From this sample, we construct a subsample of SMCCF-eligible bonds—that is, bonds issued by U.S. companies with an investment-grade rating as of March 22, 2020, and with remaining maturities of five years or less as of the March 23 announcement. The corresponding subsample of ineligible bonds consists of outstanding bonds from the same set of issuers but with maturities longer than five years, again as of the March 23 announcement. The daily price for a bond in either subsample is defined as the last transaction price recorded between 9 a.m. and 4 p.m. on a given business day.

Similar to Gilchrist and Zakrajšek (2012), we then 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). Table A-2 in Appendix A contains summary statistics of selected bond characteristics for our sample of investment-grade bonds.

2.2 Baseline Announcement Effects

Before we delve into estimation results, Figure 1 illustrates the dynamics of credit spreads of program-eligible and -ineligible bonds during the pandemic. In particular, the red line shows the daily average credit spread of SMCCF-eligible bonds, while the blue line shows the corresponding average credit spread of their ineligible counterparts, where the two types of bonds were issued by the same set of U.S. investment-grade firms.[5] Before the pandemic, SMCCF-ineligible bonds had, on average, higher credit spreads than eligible ones, mainly due to differences in liquidity and other characteristics. During the COVID-19 outbreak, the gap between the two spread series narrowed, disappearing completely during the turmoil that gripped markets in mid-March 2020.

After the Fed's announcement on March 23, both groups saw a significant drop in credit spreads. This suggests that the Fed's unequivocal commitment to support the real economy and the financial system—not the announcement of the bond-buying programs per se—eased market strains. The April 9 announcement, by contrast, had a differential effect, reducing credit spreads on eligible bonds to a greater degree compared with their ineligible counterparts, possibly due to the SMCCF's maturity-eligibility criterion.

Using the full matched sample of program-eligible and -ineligible corporate bonds issued by the same set of firms, we begin quantifying the price impact of the March 23 and April 9 announcements by estimating the following regression specification:

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

where $\mathrm{CS}_{i,j,t}$ denotes the credit spread on bond j (a liability of issuer i) on business day t. 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, whereas the 0/1-indicator $\mathbb{1}[j = E]$ equals one if bond j was eligible for purchase by the SMCCF. The regression also includes a vector of covariates, denoted by $\mathbf{X}_{i,j,t}$, comprising predetermined bond characteristics that can affect credit 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 dates. Issuer fixed effect η_i captures all (time-invariant) unobservable issuer characteristics, in effect differencing out all firm-specific effects of policy announcements within each estimation window.

In specification (1), the coefficient β_1 on the indicator $\mathbb{1}[t \geq t^*]$ measures the average change in credit spreads between the pre- and post-announcement segments of the windows bracketing each announcement. By exploiting the fact that U.S. investment-grade companies in our sample have outstanding bonds with a maturity shorter than or equal to five years—which were eligible for purchase by the SMCCF—as well as outstanding bonds with a maturity longer than five years that were ineligible, this approach should, in principle, identify the causal impact of the two announcements on credit spreads through the program's key eligibility criterion, the five-year maturity

[5] To minimize the maturity differential between the two groups of bonds, the SMCCF-eligible set of securities includes, for each issuer, a single eligible bond closest (from below) to the five-year maturity cutoff; similarly, the ineligible set of securities includes a single bond that is closest (from above) to the five-year cutoff. This approach resulted in 3,225 pairs of bonds, which were outstanding liabilities of 545 companies from January through August 2022; the average (median) difference in remaining maturities across these pairs of bonds was about 2.7 (2.3) years.

cutoff captured by the indicator $\mathbb{1}[j = E]$. 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 credit spreads between the program-eligible and -ineligible bonds of the same issuer in response to the specified announcement, a common metric used to judge the efficacy of the program.[6] To gauge the persistence of these announcement effects, we estimate specification (1) using a two-day pre-announcement window and two-, five-, and 10-day post-announcement windows bracketing the March 23 and April 9 announcements.

The two announcements' effects on secondary market pricing are reported in Table 1. Consider first the impact on credit spreads as reported in Panel A. The March 23 announcement (columns 1 through 3) led to a clear improvement in credit conditions, as evidenced by the steady narrowing of the average credit spread in the post-announcement windows. The estimates of coefficient β_1 imply that the average credit spread fell 26 basis points in the two-day window and nearly 70 basis points in the 10 days following the March 23 announcement. In response to the April 9 announcement (columns 4 through 6), credit spreads narrowed, on average, 50 basis points within the 10 days following the announcement. These estimates serve as useful benchmarks when assessing the additional impact of the two announcements through the maturity-eligibility criterion.

Turning to the estimates of β_2, the coefficient on the interaction term $\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$, the entries in columns (1) through (3) indicate that the March 23 announcement induced a persistent and significant—in both statistical and economic terms—narrowing of credit spreads on SMCCF-eligible bonds compared with their ineligible counterparts. Relative to ineligible bonds issued by the same set of companies, credit spreads on eligible bonds registered an additional decline of 48, 64, and 77 basis points within the first two, five, and 10 days following the March 23 announcement.

The April 9 announcement had a similar lasting effect on the level of credit spreads as the March 23 announcement. As shown in columns (4) through (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 of program-eligible bonds of 6 basis points, which increased to 11 basis points in the five-day window, and to 14 basis points in the 10-day window. Note that the differential effects of both announcements across program-eligible and ineligible bonds are estimated precisely—standard errors range from 2 to 4 basis points in all three estimation windows.

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 at slowing or containing the pandemic. To control for the myriad common shocks during the estimation windows, including news about actual or prospective fiscal and monetary policy actions, Panel B of Table 1 reports the estimation results of the baseline specification (1) augmented with a full set of time fixed effects.[7] Note that the estimates of coefficient β_2 are essentially the same

[6] The approach of identifying the price impact of the SMCCF's announcements in the secondary market vis-à-vis the maturity-eligibility criterion using a sample of program-eligible and -ineligible bonds issued by the same set of U.S. investment-grade firms was first proposed in our September 9, 2020, SSRN working paper (see Gilchrist et al., 2020). It was subsequently adopted by Nozawa and Qiu (2021) and Boyarchenko et al. (2023).

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

Table 1: Effect of the SMCCF Announcements on Credit Spreads

	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^{***}	-0.54^{***}	-0.69^{***}	-0.50^{***}	-0.50^{***}	-0.49^{***}
	(0.03)	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	-0.48^{***}	-0.64^{***}	-0.77^{***}	-0.06^{**}	-0.11^{***}	-0.14^{***}
	(0.04)	(0.04)	(0.04)	(0.02)	(0.02)	(0.02)
R^2	0.71	0.73	0.77	0.91	0.91	0.92
No. of firms	747	813	859	785	843	882
No. of bonds	3,717	4,059	4,298	3,819	4,068	4,307
Observations	13,370	21,845	35,717	13,861	22,238	35,804
B. With time fixed effects						
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	-0.47^{***}	-0.63^{***}	-0.76^{***}	-0.06^{**}	-0.11^{***}	-0.14^{***}
	(0.04)	(0.04)	(0.04)	(0.02)	(0.02)	(0.02)
R^2	0.74	0.77	0.81	0.92	0.91	0.92
No. of firms	747	813	859	785	843	882
No. of bonds	3,717	4,049	4,298	3,819	4,068	4,307
Observations	13,370	21,845	35,717	13,861	22,238	35,804

Note: The dependent variable in all specifications is $\mathrm{CS}_{i,j,t}$, the credit spread (in percentage points) 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 day t is greater than or equal to the specified announcement day t^*; 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. The n-day window corresponding to announcement on day t^* is defined as $[t^* - 2, t^* + n]$, for $n = 2, 5, 10$ days. All specifications include $\mathbb{1}[j = E]$, a vector of bond-specific controls, and issuer fixed effects. Asymptotic standard errors reported in parentheses are clustered at the issuer level: * $p < .10$; ** $p < .05$; and *** $p < .01$.

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 confound the estimates of the announcement effects.

All told, our baseline results indicate that the March 23 and April 9 announcements not only significantly lowered the average level of credit spreads in the investment-grade segment of the U.S. corporate bond market, but also induced a persistent and economically significant narrowing of credit spreads on SMCCF-eligible bonds. Across the two announcements, our 10-day window estimates imply a total decline of about 90 basis points in credit spreads for SMCCF-eligible bonds relative to their ineligible counterparts. By all measures, these are sizable program-specific effects, especially because the Fed has yet to purchase a single corporate bond (or bond ETF) in that time frame.

2.3 Announcements Effects and the Slope of the Credit Curve

The baseline estimation results reported in Table 1 emphasize the differential effect of the two SMCCF announcements on the average *level* of credit spreads across program-eligible and -ineligible bonds. Interpreting this as a causal impact of the program—identified vis-à-vis the maturity-eligibility criterion—assumes implicitly that the announcement effect is uniform across maturities within the program-eligible and -ineligible segments of the market. In this section, we show that this assumption is violated in the data, and we extend the analysis by considering how the two announcements also affected the *slope* of the term structure of investment-grade credit spreads.

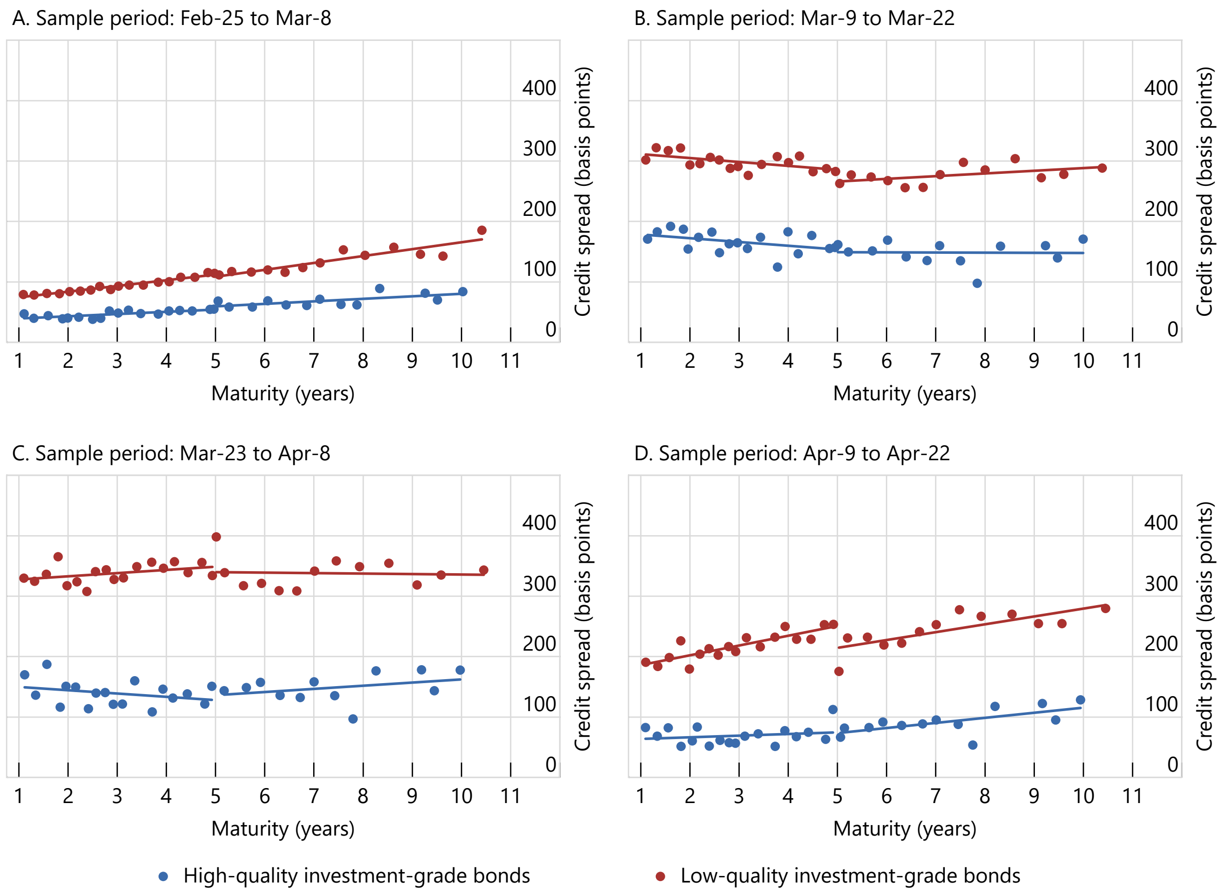
Motivation for this analysis is provided by Figure 2, which shows the cross-sectional relationship between credit spreads and maturity during the various phases of the pandemic. The figure distinguishes between "high-quality" and "low-quality" investment-grade bonds, with the former plotted in blue and the latter in red.[8]

Panel A of Figure 2 focuses on the early phase of the pandemic-induced turmoil in the corporate bond market. While credit spreads had widened during this period, the slope of the credit curve in both segments of the investment-grade market remained stable and upward. 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 the spreads of their longer-maturity counterparts (Panel B). The inversion was, on balance, more pronounced in the high-quality segment of the market, where the increase in spreads on shorter-maturity bonds was especially large in relative terms. These dynamics 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).

Panel C captures the period between the March 23 and April 9 announcements. During this period, credit spreads widened, on balance, while the inversion of the credit curve receded. Lastly,

[8]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.

FIGURE 2: Investment-grade Credit Curve during the COVID-19 Pandemic



NOTE: Each panel shows the binscatter plot of 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 (see the text for details).
SOURCE: Authors' calculations using data from TRACE/FINRA and Mergent FISD/WRDS.

Panel D focuses on the two 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 magnitudes comparable to those seen during the initial phase of the crisis shown in Panel A. These rotations suggest that one of the key results of the two announcements was the restoration of 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 announcement effects through the five-year eligibility cutoff.

To formally do so, we augment the baseline specification (1) with interaction terms that allow the slope of the credit curve to shift in the post-announcement window and across the program-eligible and -ineligible segments of the market. Specifically, we estimate:

$$\begin{aligned}\mathrm{CS}_{i,j,t} = {} & \beta_1 \mathbb{1}[t \geq t^*] + \beta_2\big(\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]\big) + \beta_3\big(\mathbb{1}[t \geq t^*] \times \tau_{i,j,t}\big) \\ & + \beta_4\big(\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E] \times \tau_{i,j,t}\big) + \theta' \mathbf{X}_{i,j,t} + \eta_i + \epsilon_{i,j,t},\end{aligned} \tag{2}$$

Table 2: 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^{***}	-1.05^{***}	-1.31^{***}	-0.67^{***}	-0.70^{***}	-0.68^{***}
	(0.11)	(0.10)	(0.10)	(0.06)	(0.05)	(0.05)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	-0.70^{***}	-0.74^{***}	-0.89^{***}	0.13	0.08	0.04
	(0.14)	(0.14)	(0.15)	(0.07)	(0.07)	(0.06)
$\mathbb{1}[t \geq t^*] \times \tau_{i,j,t}$	0.04^{**}	0.07^{***}	0.08^{***}	0.02^{***}	0.03^{***}	0.03^{***}
	(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^{***}	0.14^{***}	0.17^{***}	-0.03	-0.02	-0.02
	(0.03)	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)
R^2	0.72	0.74	0.78	0.91	0.91	0.92
No. of firms	747	813	859	785	843	882
No. of bonds	3,717	4,059	4,298	3,819	4,068	4,307
Observations	13,370	21,845	35,717	13,861	22,238	35,804

Note: The dependent variable in all specifications is $\text{CS}_{i,j,t}$, the credit spread (in percentage points) 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 day t is greater than or equal to the specified announcement day t^*; $\mathbb{1}[j = E] = 0/1$-indicator that equals one if bond j was eligible for purchase by the SMCCF as of March 22; and $\tau_{i,j,t}$ = bond j's remaining maturity (in years). The n-day window corresponding to announcement on day t^* is defined as $[t^* - 2, t^* + n]$, for $n = 2, 5, 10$ days. All specifications include $\mathbb{1}[j = E]$, $\mathbb{1}[j = E] \times \tau_{i,j,t}$, a vector of bond-specific controls, and issuer fixed effects. Asymptotic standard errors reported in parentheses are clustered at the issuer level: * $p < .10$; ** $p < .05$; and *** $p < .01$.

where $\tau_{i,j,t}$ denotes the remaining maturity (in years) of bond j on business day t. Table 2 summarizes the results of this exercise.

According to columns (1) through (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 curve by compressing credit spreads at the very short end of the curve by significantly more relative to their counterparts closer to the five-year eligibility cutoff. For example, two days after the announcement, credit spreads on bonds with a remaining maturity of one year fell 57 $(-0.70+0.13\times1)$ basis points, whereas spreads on bonds with a remaining maturity of five years narrowed a mere 5 $(-0.70 + 0.13 \times 5)$ basis points, according to our estimates.

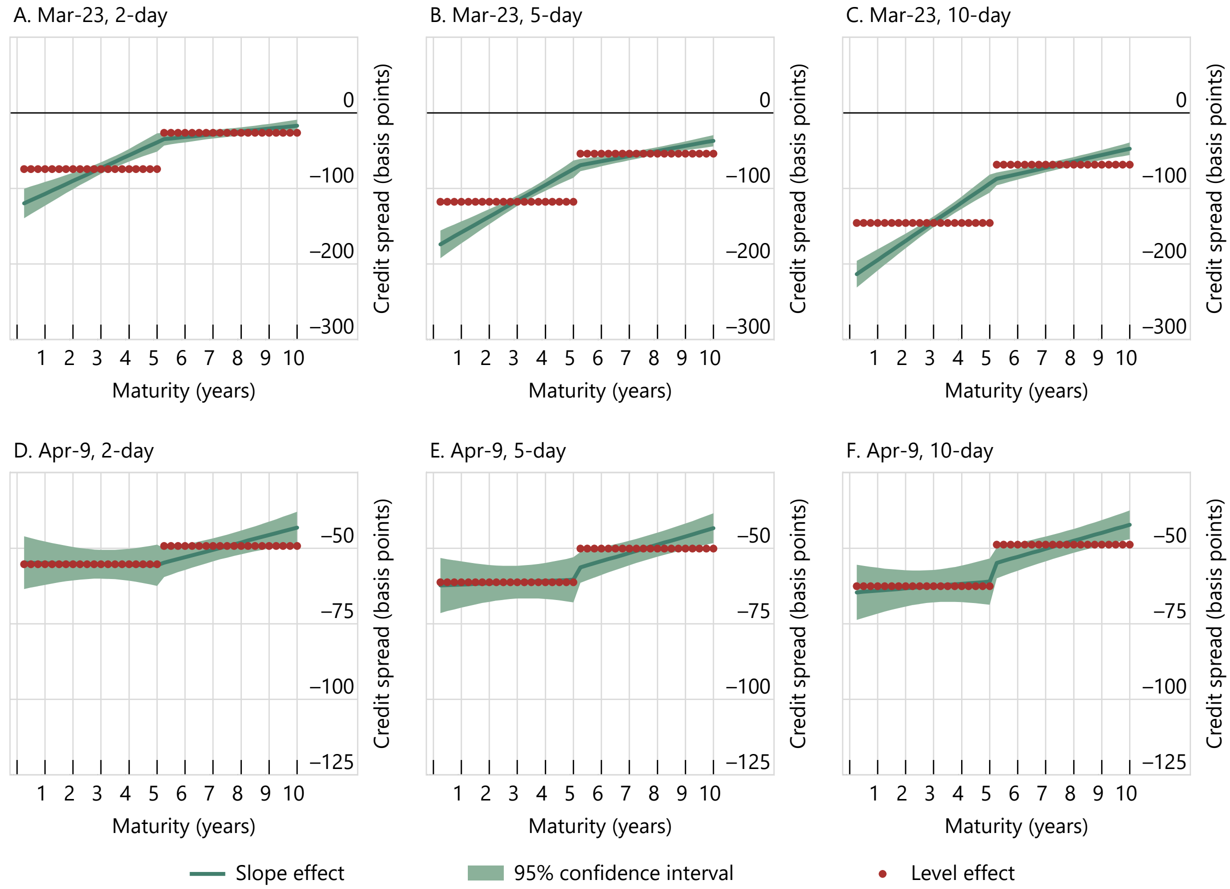
The latter effect, in addition to being economically small, is also statistically indistinguishable from zero and remains so 10 days after the announcement. The former effect, by contrast, is highly persistent and builds over time: 10 days after the announcement, credit spreads on bonds with a remaining maturity of one year narrowed by 72 $(-0.89 + 0.17 \times 1)$ basis points. As shown in column (1), the estimates of β_1 and β_3, the coefficients on the indicator $\mathbb{1}[t \geq t^*]$ and the interaction term $\mathbb{1}[t \geq t^*] \times \tau_{i,j,t}$, respectively, also indicate a mild steepening of the program-ineligible segment of the investment-grade credit in response to the March 23 announcement.

The estimates of β_2 and β_4 in columns (4) through (6), by contrast, are all statistically indistinguishable from zero, implying no such slope effect on the program-eligible segment of the market in response to the April 9 announcement. As evidenced by the statistically significant coefficients β_1 and β_3, the entire investment-grade credit curve steepened in response to the April 9 announcement—though less so than after the March 23 announcement—with no differential effect between the program-eligible and -ineligible segments of the curve.

Figure 3 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 the coefficients β_1, β_2, β_3, and β_4 reported in Table 2 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 specification (1), which does not control for the associated changes in the slope of the credit curve.

Turning first to the March 23 announcement (Panels A through C), the estimated rotations of the credit curve imply economically large differences in the announcement effect 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 10 years. And while the entire curve shifts noticeably lower as the estimation window lengthens, the slope effect induced by the March 23 announcement on program-eligible bonds remains considerably

Figure 3: Announcement-induced Shifts in the Slope of the Credit Curve



Note: The green solid lines in Panels 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. Specifically, using the estimates of specification (2) in Table 2, the green solid lines depict $(\hat{\beta}_1 + \hat{\beta}_2) + (\hat{\beta}_3 + \hat{\beta}_4) \times \tau$ for program-eligible bonds ($\tau \leq 5$) and $\hat{\beta}_1 + \hat{\beta}_3 \times \tau$ for program-ineligible bonds ($\tau > 5$). The red dotted lines show the baseline estimates of the two announcements—that is, $\hat{\beta}_1 + \hat{\beta}_2$ for program-eligible bonds and $\hat{\beta}_1$ for their ineligible counterparts, using the estimates of specification (1) reported in Panel A of Table 1.
Source: Authors' calculations.

stronger and statistically different from the slope effect on program-ineligible bonds.

These estimates also show that there is no differential level effect on credit spreads at the SMCCF's five-year maturity cutoff once the announcement-induced shifts in the slope of the credit curve are taken into account. The absence of a distinct level effect attributable to the program's key eligibility criterion is likely due to the sweeping nature of the March 23 announcement as reflected in its opening sentence, which stated that the Fed is "committed to using its full range of tools to support households, businesses and the U.S. economy overall." That the announcement also had a disproportionate effect on spreads of very short maturity program-eligible bonds—a segment of the market that experienced the greatest dislocation during the early stages of the pandemic—supports this interpretation.

In response to the April 9 announcement (Panels D through E), by contrast, the entire term structure of credit spreads steepens and moves somewhat lower as the estimation window lengthens. And while the March 23 announcement had a differential slope effect on the program-eligible segment of the market, the April 9 announcement had a uniform effect on the entire term structure of investment-grade credit spreads. When summed up across the two announcements, our estimates based on the 10-day window imply an announcement-induced differential of about 170 basis points between bonds with remaining maturities of one and 10 years.

To sum up, the results in Table 2 and Figure 3 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, by contrast, led to a uniform downshift and steepening of the entire term structure of credit spreads, ultimately restoring the investment-grade credit curve to its pre-pandemic shape and level. Whether the underlying securities were eligible or ineligible for purchase by the SMCCF played essentially no role in how the two announcements affected investment-grade corporate bond credit spreads.

Default risk versus credit risk premia: This section examines to what extent are the estimated announcement-induced declines in credit spreads due to a reduction in default risk as opposed to a decline in credit risk premia. Following Gilchrist and Zakrajšek (2012), we decompose investment-grade credit spreads into a component that captures issuer-specific time-varying 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. 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 (see Appendix B for details).

Table 3 reports the results from estimating specification (2), using the bond-specific residual credit spreads as the dependent variable. It is important to note that because the distance-to-default—our measure of firm-specific time-varying default risk inferred from equity prices—increases with the horizon (that is, the bond's remaining time-to-maturity), such variation in default risk is not automatically picked up by the inclusion of issuer fixed effects in the regression.[9] Nonetheless, the estimates of coefficients β_1, β_2, β_3, and β_4 using the residual credit spreads as the dependent variable are almost identical—in terms of both their magnitudes and temporal patterns—to those in Table 2 that use the actual credit spreads as the dependent variable.

These findings imply that the announcement-induced rotations of the investment-grade credit curve shown in Figure 3 are due primarily to a reduction in credit risk premia across the entire term structure, or a broad-based improvement in market sentiment, rather than to a reduction in

[9] Because measuring distance-to-default requires equity prices, the sample of bonds used in this exercise corresponds to a subset of investment-grade U.S. issuers in the TRACE database that are publicly listed.

TABLE 3: Effect of the SMCCF Announcements on Credit Risk Premia

	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)
$\mathbb{1}[t \geq t^*]$	-0.55^{***}	-1.09^{***}	-1.42^{***}	-0.56^{***}	-0.55^{***}	-0.56^{***}
	(0.12)	(0.12)	(0.12)	(0.06)	(0.06)	(0.06)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E]$	-0.68^{***}	-0.66^{***}	-0.72^{***}	0.03	-0.06	-0.08
	(0.17)	(0.16)	(0.16)	(0.08)	(0.07)	(0.06)
$\mathbb{1}[t \geq t^*] \times \tau_{i,j,t}$	0.03^{*}	0.07^{***}	0.09^{***}	0.01	0.01	0.01
	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
$\mathbb{1}[t \geq t^*] \times \mathbb{1}[j = E] \times \tau_{i,j,t}$	0.13^{***}	0.12^{***}	0.14^{***}	-0.01	0.00	0.01
	(0.04)	(0.04)	(0.03)	(0.02)	(0.01)	(0.01)
R^2	0.66	0.69	0.73	0.88	0.88	0.89
No. of firms	496	532	555	516	541	560
No. of bonds	2,555	2,737	2,846	2,596	2,716	2,833
Observations	9,889	16,164	26,369	10,037	16,090	26,006

NOTE: The dependent variable in all specifications is $\text{RCS}_{i,j,t}$, the residual credit spread (in percentage points) 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 day t is greater than or equal to the specified announcement day t^*; $\mathbb{1}[j = E] = 0/1$-indicator that equals one if bond j was eligible for purchase by the SMCCF as of March 22; and $\tau_{i,j,t}$ = bond j's remaining maturity (in years). The n-day window corresponding to announcement on date t^* is defined as $[t^* - 2, t^* + n]$, for $n = 2, 5, 10$ days. All specifications include $\mathbb{1}[j = E]$, $\mathbb{1}[j = E] \times \tau_{i,j,t}$, a vector of bond-specific controls, and issuer fixed effects. Asymptotic standard errors reported in parentheses are clustered at the issuer level: * $p < .10$; ** $p < .05$; and *** $p < .01$.

default risk, at least as perceived by equity markets.[10] As such, they provide further support for the theoretical framework of Hanson et al. (2020), in which the announcement of a corporate bond-buying facility affects the market through a reduction in uncertainty and the associated credit risk premia. The announcement-induced declines in the entire term structure of credit spreads net of default risk 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 is documented by O'Hara and Zhou (2021) and Kargar et al. (2021); we explore this mechanism quantitatively within a preferred habitat framework in Section 3.

Macroeconomic implications: We close with a brief discussion of the impact of the two SMCCF announcements on economic activity. According to our estimates, the EBP in the investment-grade segment of the corporate bond market fell a cumulative 170 basis points in response to the two announcements. To gauge the relevance of this decline for the real activity, we ask how much employment, industrial production, and unemployment rate would respond if this announcement effect is treated as a "shock" to the investment-grade EBP.

Appendix C outlines the local projections framework used to trace out the impact of such a shock on these monthly indicators of economic activity. Based on these estimates, such a counterfactual implies that absent the March 23 and April 9 announcements, the growth of industrial output would be nearly 6 percentage points lower, total employment growth would decline by an additional 1.5 percentage points, and the unemployment rate would increase by an additional 1.2 percentage points over the subsequent 12 months. While this counterfactual should be treated with caution, these estimates point to an economically significant role of the two SMCCF announcements in the Fed's overall effort to stabilize the real economy during the pandemic.

2.4 The SMCCF's Purchase Effects

The preceding section focused on the price impact of the two bond-buying program announcements in the corporate bond market. We now turn to quantifying the price impact of the Fed's actual purchases of eligible securities.

The SMCCF started buying corporate bond ETFs on May 12, 2020, and individual corporate bonds on June 16. In the latter half of June, the facility 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.[11] In total, the SMCCF purchased about \$5.4 billion corporate bonds in a span of five and a half months, with almost

[10] Nozawa and Qiu (2021) use an alternative decomposition of credit spreads into the expected default and risk premium components and find that the risk-premium channel and the default-risk channel are equally important.

[11] On July 28, 2020, 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. The SMCCF ceased purchases on November 27. The PMCCF, which was slated to commence purchases in the primary market on June 29, 2020, did not execute a single transaction during its operational phase.

70 percent of those purchases (about \$3.6 billion) taking place in the second half of June and July (see Figure D-1 in Appendix D).

To estimate the price impact of these purchases, we identify the Fed's purchases of individual bonds by using *intraday* TRACE transactions data from June 16 through July 31, a period during which the facility made most of its purchases. By matching the bond's CUSIP, purchase date and exact 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.[12] Using the exact time of each purchase, we begin with a simple event study to estimate an average purchase effect, the results of which are summarized in Figure 4.[13]

Panel A 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. Though the estimates are noisy, they indicate that the credit spread on an average purchased bond declined about 5 basis points upon the actual purchase. Over the subsequent five hours, the spread edged up about 2 basis points before stabilizing for the rest of the event window, resulting in a net decline of about 3 basis points. Panel B shows the corresponding average spread of bonds that were issued by the same set of issuers but had a remaining maturity longer than five years.[14] Notably, the actual purchases appear to have also had a delayed effect on the credit spreads of ineligible bonds, though this effect is very small, a mere basis point or so, and subject to considerable uncertainty.

Panels C and D show the same average purchase effects on credit spreads for high-quality investment-grade corporate bonds. As shown, spreads on high-quality bonds purchased by the facility declined, on average, 2 basis points upon purchase (Panel C), though the associated 95 percent confidence intervals are again quite wide. The average purchase effect is most visible for low-quality investment-grade bonds (Panels E and F). In that case, spreads on purchased bonds (Panel E) declined nearly 4 basis points. Although a portion of these purchase-induced declines was reversed over the subsequent five hours, credit spreads on purchased low-quality bonds remained lower by about 3 basis points, on net, even 20 hours after their purchase.

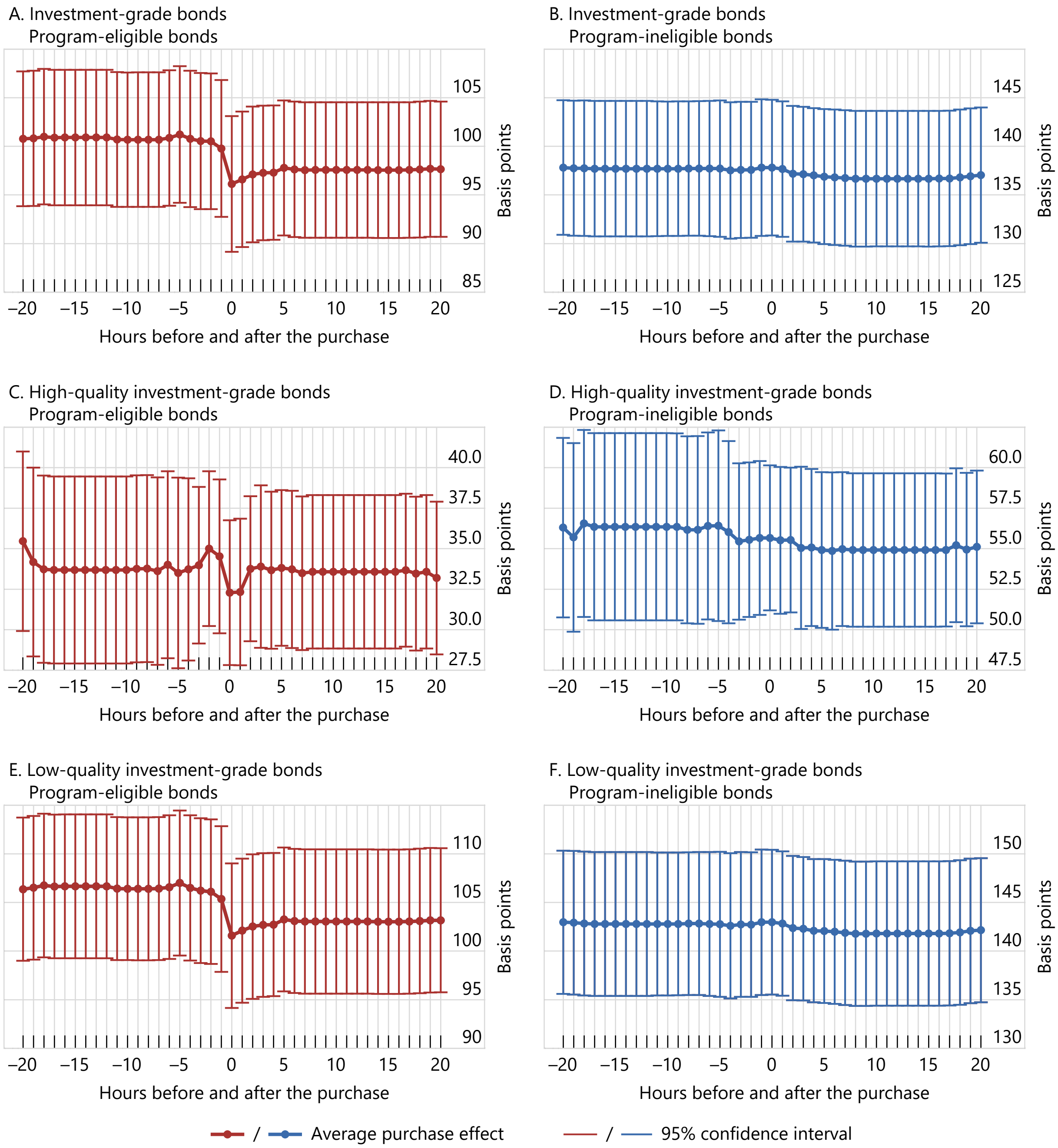
To improve on the power of this exercise and estimate the purchase effects more precisely, we expand the "narrow" matched sample used in Figure 4 to include all transactions of corporate

[12] 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.

[13] A potential concern with this analysis is that purchased bonds and their ineligible counterparts traded too infrequently around the times when the facility executed its purchases. As a result, one cannot obtain a sensible estimate of the average purchase effect because transactions are spaced too far apart in time. However, as shown in Appendix D, there were numerous transactions in active hourly trading sessions—for both purchased and ineligible bonds—even within three and six hours of the facility's purchase times. This suggests that the facility targeted the most liquid bonds, which can trade multiple times per day. The substantial number of intraday transactions for both purchased and ineligible bonds allows us to estimate an average purchase effect using intraday data.

[14] To minimize the maturity differential between the two groups, each bond purchased by the SMCCF is paired with a bond issued by the same company but whose remaining maturity is longer than five years. If an issuer has multiple bonds purchased by the SMCCF, the bond with a remaining maturity as close to five years as possible is chosen. Similarly, if there are multiple ineligible bonds that can be paired up with a given bond purchased by the SMCCF, the bond with a remaining maturity as close to five years as possible is selected. All told, there are 482 unique issuers in our sample.

Figure 4: The Impact of the SMCCF's Purchases on Credit Spreads



Note: Panel A shows the average credit spread on investment-grade bonds purchased by the SMCCF in one-hour windows within ±20 hours of their purchase between June 16 and July 31, 2020; Panel B shows the corresponding average credit spread on the SMCCF-ineligible bonds—with the remaining time-to-maturity that is closest to the purchased bonds—issued by the same set of issuers (see the text for details). The same information for high-quality investment-grade bonds is shown in Panels C–D and for low-quality investment-grade bonds in Panels E–F.
Source: Authors' calculations using data from TRACE/FINRA and FRBNY.

bonds issued by the same issuer within 24 hours of the purchase time for each bond purchased by the facility. We then use the resulting "broad" matched sample to estimate how program purchases affected credit spreads in the secondary market using the following specification:

$$\mathrm{CS}_{i,j,h} = \beta_1 \mathbb{1}[h \geq h^*] + \beta_2 \big(\mathbb{1}[h \geq h^*] \times \mathbb{1}[j = P]\big) + \theta' \mathbf{X}_{i,j,h} + \eta_i + \epsilon_{i,j,h}, \tag{3}$$

where h^* (*2020:mm:dd:mm:ss*) denotes the transaction time for a given purchase by the program, and $\mathrm{CS}_{i,j,h}$ denotes the credit spread on bond j (issued by firm i) at transaction time h (*2020:mm:dd:mm:ss*). The 0/1-indicator $\mathbb{1}[h \geq h^*]$ equals one if the transaction time h is greater than or equal to h^*, and the 0/1-indicator $\mathbb{1}[j = P]$ equals one if bond j was purchased by the SMCCF. The regression also includes our standard set of controls and issuer fixed effects.

Table 4 reports these regression results for various time intervals following the facility's purchases. Specifically, letting $(\underline{h}, \overline{h}) = (h^* - 00{:}05, h^* + 00{:}05)$ denote the 10-minute intervals bracketing h^*, the actual times at which the SMCCF executed its purchases of individual securities between June 16 and July 31, 2020, we consider the following time intervals: 5m $= [\underline{h} - 07{:}55, h^* + 00{:}05]$; 1h $= [\underline{h} - 07{:}55, \overline{h} + 00{:}55]$; 2h $= [\underline{h} - 07{:}55, \overline{h} + 01{:}55]$; 3h $= [\underline{h} - 07{:}55, \overline{h} + 02{:}55]$; 7h $= [\underline{h} - 07{:}55, \overline{h} + 06{:}55]$; and 8h $= [\underline{h} - 07{:}55, \overline{h} + 07{:}55]$. It is important to note that these time intervals exclude "overnight" hours.

The results in Panel A show that direct purchases reduced credit spreads of eligible bonds 11 basis points more than those of their ineligible counterparts within the first five minutes of the purchase (column 1). Although the purchase effect is estimated to shrink to 3 basis points within the first hour of the purchase (column 2), it remains statistically significant even after seven or eight hours (columns 5 and 6). Panels B and C contain the results for high- and low-quality investment-grade bonds, respectively. These estimates clearly indicate that economically and statistically significant overall purchase effects are due entirely to the facility's purchases of low-quality investment-grade bonds, a result consistent with the simple event study shown in Figure 4.

All told, the above analysis indicates that most of the impact of the SMCCF on the corporate bond market—as reflected in a substantially narrower credit spreads—is due to the announcement effects, which occurred well before the Fed intervened directly in the market. Nonetheless, we find evidence that the actual purchases of bonds in the secondary market also had an effect on their prices. In particular, for issuers at the bottom rungs of the investment-grade ladder, the Fed's actual purchases reduced credit spreads on purchased bonds, on average, 3 basis points relative to their program-ineligible counterparts. Given the modest—by QE standards—scale of these purchases, this is a sizable clientele effect, corroborating the overall efficacy of the program.

TABLE 4: Effect of the SMCCF Purchases on Credit Spreads

	5m	1h	2h	3h	7h	8h
Explanatory Variable	(1)	(2)	(3)	(4)	(5)	(6)
A. Investment-grade bonds						
$\mathbb{1}[h \geq h^*] \times \mathbb{1}[j = P]$	-0.11^{***}	-0.03^{**}	-0.02	-0.02	-0.03^{**}	-0.03^{**}
	(0.04)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
R^2	0.96	0.97	0.97	0.97	0.97	0.97
No. of firms	446	471	475	475	480	480
No. of bonds	1,982	2,042	2,071	2,087	2,135	2,140
Observations	47,833	54,828	61,782	68,285	87,388	92,002
B. High-quality investment-grade bonds						
$\mathbb{1}[h \geq h^*] \times \mathbb{1}[j = P]$	0.01	-0.02	-0.01	-0.00	-0.01	-0.01
	(0.03)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
R^2	0.84	0.83	0.83	0.82	0.81	0.81
No. of firms	21	21	21	21	21	21
No. of bonds	112	113	113	113	113	114
Observations	5,850	6,680	7,600	8,336	10,595	11,043
C. Low-quality investment-grade bonds						
$\mathbb{1}[h \geq h^*] \times \mathbb{1}[j = P]$	-0.12^{***}	-0.02	-0.02	-0.02	-0.03^{**}	-0.03^{**}
	(0.04)	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)
R^2	0.96	0.96	0.96	0.96	0.96	0.96
No. of firms	432	457	461	461	466	466
No. of bonds	1,869	1,928	1,957	1,973	2,021	2,025
Observations	41,982	48,147	54,181	59,948	76,792	80,958

NOTE: The dependent variable in all specifications is $\text{CS}_{i,j,h}$, the credit spread (in percentage points) at transaction time h (*2020:mm:dd:mm:ss*) of bond j, issued by firm i whose eligible bonds were purchased, in the specified time interval (see the text for details). The entries in the table denote the OLS estimates of coefficients associated with the interaction term $\mathbb{1}[h \geq h^*] \times \mathbb{1}[j = P]$, where $\mathbb{1}[h \geq h^*] = 0/1$-indicator that equals one if transaction time h is greater than or equal to the purchase time h^*; and $\mathbb{1}[j = P] = 0/1$-indicator that equals one if bond j was purchased by the SMCCF. All specifications include $\mathbb{1}[h \geq h^*]$, $\mathbb{1}[j = P]$, a vector of bond-specific controls, and issuer fixed effects. Asymptotic standard errors reported in parentheses are clustered at the issuer level: * $p < .10$; ** $p < .05$; and *** $p < .01$.

3 Inspecting the Mechanism

This section uses 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 include a 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 the March 23 and April 9 announcements.

The primary modeling contribution is to extend the Vayanos-Vila model, which focuses on the Treasury bond market, to include a parallel market for high-quality investment-grade corporate bonds. Within this extended framework, we are able to show that credit spreads (that is, the difference in yields on bonds of a given maturity between the two markets) depend on the differential impact of demand shocks and the difference in arbitrageurs' trading activities across these two markets. We then show how this natural extension of the Vayanos-Vila framework can be calibrated and characterized using a full set of empirical moments that capture pre-pandemic conditions in the Treasury and the high-quality investment-grade segment of the corporate bond market. While this extension is relatively straightforward, to our knowledge, it previously has not been considered.

3.1 The Setup

We posit a preferred habitat model with two markets: one for Treasury bonds and a parallel market for high-quality (that is, 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.[15] In this 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, the tilde symbol is used to refer to a variable or a process in the corporate bond market, while the counterparts without the tilde symbol are for 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), 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

[15] A significant inversion of the investment-grade credit curve also occurred during the Global Financial Crisis (GFC). At the height of the crisis (September to December 2008), credit spreads on one- to five-year investment-grade corporate bonds increased 200 basis points more than spreads on investment-grade bonds with maturity longer than 10 years. During the pandemic, the same differential was 150 basis points, roughly of similar magnitude. By contrast, the high-yield spread surged 1,400 basis points during the GFC but increased only 700 basis points during the pandemic. This suggests that the increase in the near-term default risk played a much more significant role in the inversion of the credit curve during the GFC than during the pandemic.

supply and with maturities ranging from 0 to T years. 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: $dr_t = \kappa_r(\bar{r} - r_t)dt + \sigma_r dB_{r,t}$, 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

$$\begin{aligned} \widetilde{Z}_t^{(\tau)} &= \tilde{\alpha}(\tau)\tau\tilde{y}_t^{(\tau)} - \tilde{\beta}_t^{(\tau)}, \\ \tilde{\beta}_t^{(\tau)} &= \tilde{\theta}_0(\tau) + \tilde{\theta}(\tau)\tilde{\beta}_t, \end{aligned}$$

where $\tilde{\alpha}(\tau)$ and $\tilde{\beta}_t^{(\tau)}$ denote the *slope* and *intercept* of their demand function at time t, respectively, and the demand risk factor $\tilde{\beta}_t$ follows an Ornstein-Uhlenbeck process: $d\tilde{\beta}_t = -\kappa_{\tilde{\beta}}\tilde{\beta}_t dt + \sigma_{\tilde{\beta}} dB_{\tilde{\beta},t}$.

As shown by Vayanos and Vila (2021), corporate bond yields in this framework are an affine function of the short rate r_t and the demand risk factor $\tilde{\beta}_t$

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

Similarly, Treasury bond yields have analogous expressions: $y_t^{(\tau)} = \frac{1}{\tau}\left[A_r(\tau)r_t + A_\beta(\tau)\beta_t + C(\tau)\right]$, 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, the credit spread of the maturity-τ corporate bond, that is, $s_t^{(\tau)} = \tilde{y}_t^{(\tau)} - y_t^{(\tau)}$, is given by

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

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.[16]

[16] This assumption seems reasonable because the default risk of high-quality investment-grade bonds is, in general, quite low, and as shown in Panels A and C of Figure B-1 in Appendix B, did not change materially in the narrow windows surrounding the March 23 and April 9 announcements.

Following Vayanos and Vila (2021), we parameterize the demand-side parameters in the corporate bond market as $\tilde{\alpha}(\tau) = \tilde{\alpha}e^{-\tilde{\delta}_\alpha \tau}$; $\tilde{\theta}_0(\tau) = \tilde{\theta}_0 \left(e^{-\tilde{\delta}_\alpha \tau} - e^{-\tilde{\delta}_\theta \tau}\right)$; and $\tilde{\theta}(\tau) = \tilde{\theta} \left(e^{-\tilde{\delta}_\alpha \tau} - e^{-\tilde{\delta}_\theta \tau}\right)$, for $0 \leq \tau \leq T$, and similarly for the Treasury market. Because a vast majority of outstanding corporate bonds have a remaining maturity shorter 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 high-quality investment-grade corporate bond yields and the corresponding yields of the duration-matched synthetic Treasuries constructed following Gilchrist and Zakrajšek (2012). We then use the Nelson-Siegel-Svensson framework to estimate a monthly yield curve from which we calculate the implied yields at maturities of $1, 2, \ldots, 20$ years in each market.

As do Vayanos and Vila (2021), parameters in the Treasury bond market (that is, κ_r, σ_r, κ_β, a, α, θ_0, θ, δ_α, and δ_θ) and the corresponding parameters in the corporate bond market (that is, $\kappa_{\tilde{\beta}}$, $\tilde{a}$, $\tilde{\alpha}$, $\tilde{\theta}_0$, $\tilde{\theta}$, $\tilde{\delta}_\alpha$, and $\tilde{\delta}_\theta$) are calibrated to match key moments or the values from the literature. Table E-1 in Appendix E summarizes the calibrated parameters and the empirical moments used to determine them. Figure E-1 shows the selected model-implied moments and their empirical analogs across the range of maturities considered in our calibration procedure.

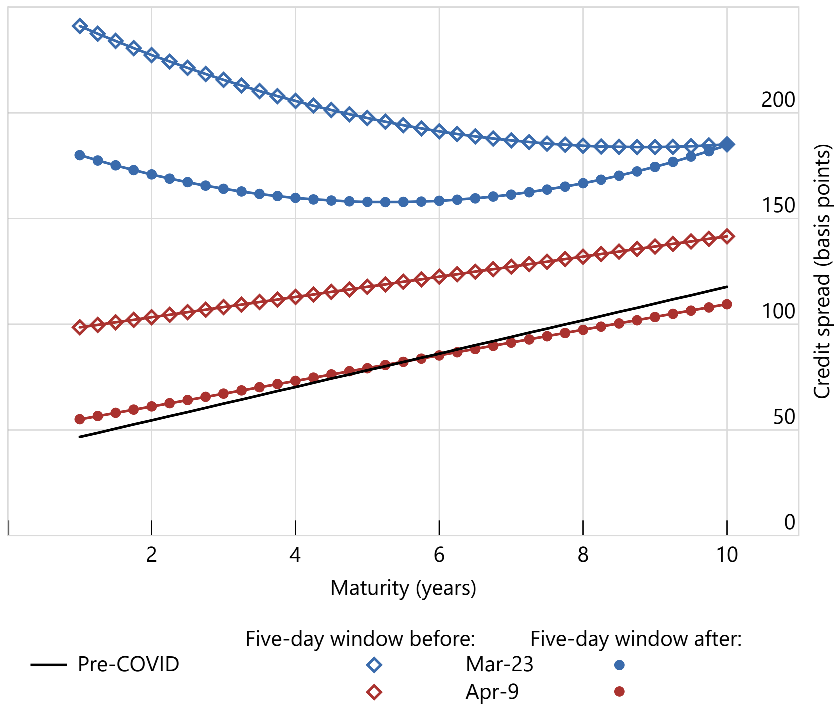
3.2 Understanding the Mechanism of the SMCCF

This section uses 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 5 highlights the key movements in the credit curve that are the focus of this exercise. Specifically, the figure shows the fitted credit curve at specified dates, obtained from a cross-sectional regression of credit spreads for high-quality investment-grade corporate bonds on a quadratic polynomial in the bonds' remaining time-to-maturity.

According to 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 on 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 has 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 has moved significantly lower, noticeably reducing 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, these dynamics are explained 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 (that is, shorter-maturity high-quality, investment-grade bonds). And second, faced with

FIGURE 5: 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 announcement 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 data from TRACE/FINRA and Mergent FISD/WRDS.

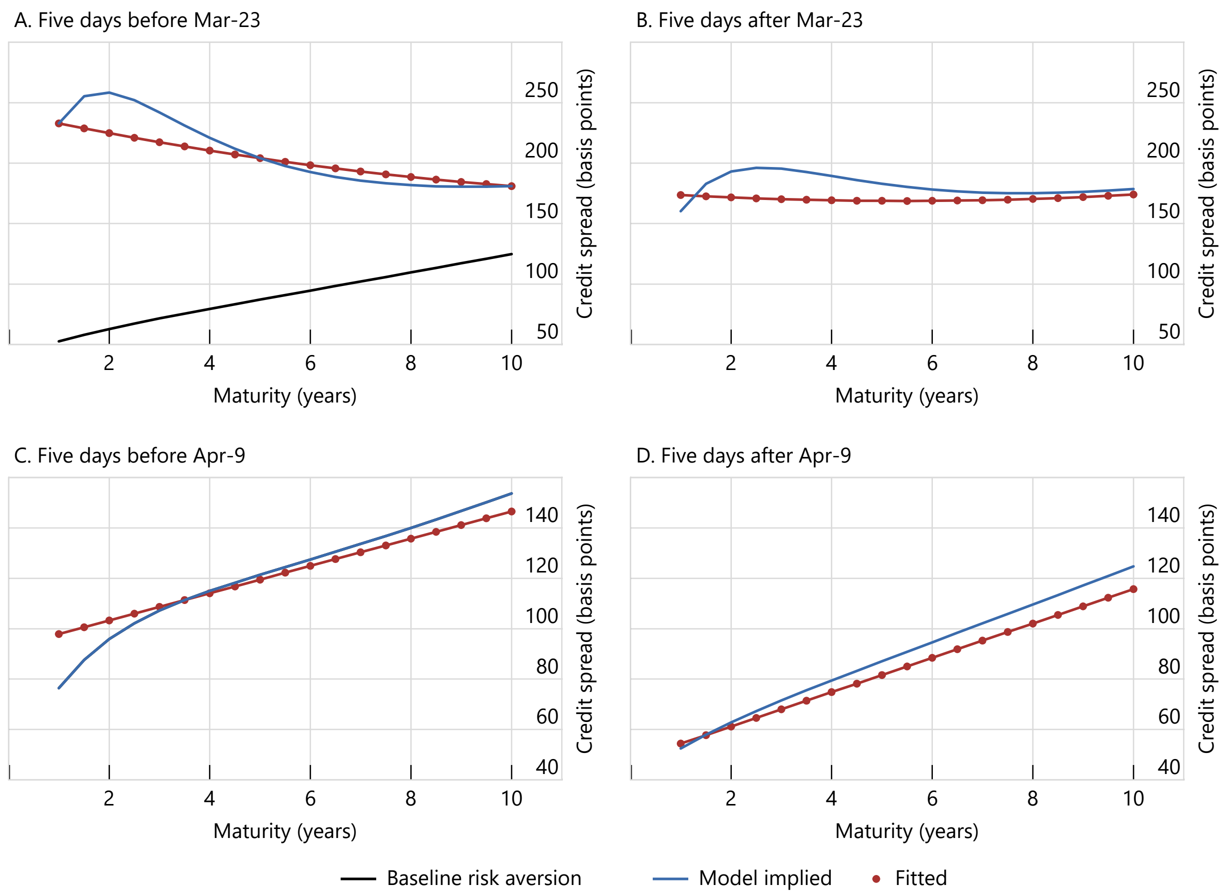
the massive selloff across fixed-income markets, broker-dealers became increasingly unwilling to take on inventory, a reduction in their risk-bearing capacity that is consistent with the spike in the average investment-grade credit risk premium shown in Figure B-2 of Appendix B.

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 5) to levels registered during the five-day window immediately preceding the March 23 announcement (the blue line with $\diamond$'s).

Figure 6: 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 high-quality investment-grade credit curves (see Figure 5).
Source: Authors' calculations.

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 10-year corporate bond yield upon the impact of the shock. To capture the concomitant reduction in the broker-dealers' risk-bearing capacity, 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. Note that the increase in $\tilde{a}$ by a factor of about 300 is commensurate with the increase in our estimate of the average credit risk premium shown in Figure B-2 of Appendix B.

The solid blue line in Panel A of Figure 6 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, we set the demand shock parameters $\Delta_1 = 0.41$ and $\Delta_2 = 1.11$.[17]

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 (that is, $(\Delta_1, \Delta_2) = (0.41, 1.11)$), 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 5. 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. 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.[18]

As noted above, we model the Fed's announcements as "calming" the market by reducing the arbitrageurs' risk aversion. Viewed through this lens, Panel B shows the value of $\tilde{a}$ that minimizes deviations of the model-implied one-, five-, and 10-year credit spreads from their empirical counterparts in the five-day window following the March 23 announcement. The result of this experiment implies a reduction in $\tilde{a}$ from 1,000 to 285.7. A 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 (for example, the one-year credit spread fell 75 basis points), resulting in a 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.[19]

[17] These values of Δ_1 and Δ_2 are about 14 and 38 times the unconditional volatility of the demand risk factor $\tilde{\beta}_t$ (that is, $\sqrt{\sigma^2_{\tilde{\beta}}/(2\kappa_{\tilde{\beta}})} = 0.029$). To match the increase in the 10-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 of 0.74 years.

[18] 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.

[19] In Appendix E, we also use the calibrated model to quantify the facility's actual purchase effects. As shown in Figure E-2, the model-implied purchase effects are essentially zero, consistent with the estimates of the average purchase effect for high-quality investment-grade bonds reported in Panel B of Table 4.

4 Conclusion

The aim of this paper is to quantify the announcement and purchase effects of the SMCCF on prices in the U.S. corporate bond market. Using a matched sample of program-eligible and -ineligible securities trading in the secondary market—with both types of securities issued by the same company—and a methodology that explicitly controls for the announcement-induced shifts in the credit curve, we show that the March 23 and April 9, 2020, announcements significantly reduced investment-grade credit spreads across the maturity spectrum. In particular, the two announcements progressively lowered and steepened the credit curve, ultimately restoring the pre-pandemic upward-sloping term structure of credit spreads.

The analysis also indicates that there is no differential announcement effect on credit spreads around the SMCCF's five-year maturity cutoff, once we control for the announcement-induced shifts in the slope of the credit curve. Moreover, the announcement-induced narrowing of credit spreads is due almost entirely to a decline in credit risk premia. In combination, these results support the notion that the efficacy of the program reflected a broader mechanism, one that restored the normal upward slope of the investment-grade credit curve and had little to do with the program's eligibility criterion.

Using an event-style methodology with precisely identified times of the Fed's purchases of individual corporate bonds, we document that these had economically sizable effects—given the size of the program—on credit spreads of bonds issued by firms on the lower rungs of the investment-grade ladder. Lastly, we show that the above empirical findings can be rationalized within the preferred habitat framework of Vayanos and Vila (2021). Overall, the results imply that the primary effect of the Fed's March 23 and April 9 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 Materials

— For Online Publications Only —

The supplementary materials contain five appendixes (A–E). Appendix A provides an overview of the Federal Reserve's corporate bond-buying programs and reports summary statistics for the key variables in our data set. In Appendix B, we provide details regarding the construction of the residual credit spreads, our proxy for credit risk premia. Appendix C describes the empirical methodology used to gauge the macroeconomic implications of the SMCCF reported in the main text. In Appendix D, we provide key background data details of our intraday event study of the facility's purchase effects. Lastly, Appendix E describes the calibration of the extended version of the Vayanos and Vila (2021) preferred habitat model.

A 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.[20]

Bonds eligible for purchase 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 10 percent of the issuer's maximum dollar amount of bonds outstanding on any day from March 22, 2019, through March 22, 2020. Importantly, the March 23 announcement stipulated that the two corporate bond-buying facilities were open to only investment-grade U.S. companies.

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.

As noted in the main text, 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. It is worth noting that the PMCCF, which was slated to commence purchases in the primary market on June 29, 2020, did not execute a single transaction during its operational phase.

[20]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.

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 (FRBNY) published on June 28, 2020, the initial Broad Market Listing (BML), a set of corporate bonds eligible for purchase by the SMCCF.[21] To shed light on which credits the facility was targeting, we report in Table A-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, was the largest issuer in the Consumer Non-Cyclical sector.[22]

TABLE A-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.
SOURCE: Authors' calculations using data from FRBNY.

Summary statistics: In Table A-2, we report summary statistics of selected bond characteristics for our sample of investment-grade bonds. 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 1 basis point or with credit spreads exceeding 2,000 basis points.

To get a sense of how the pandemic affected the U.S. corporate bond market, we focus on two sample periods: a pandemic period running from January through the end of July 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. Other than that, the remaining bond characteristics are similar across the two samples.

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

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

Table A-2: Summary Statistics of Selected Bond Characteristics

Variable	Mean	SD	P25	P50	P75
A. Pre-pandemic period (Jan–Jul 2019)[a]					
Credit spread (bps.)	99.4	63.2	56.1	84.7	128.0
Time-to-maturity (years)	4.8	2.5	2.7	4.4	6.8
Age (years)	4.3	4.2	1.7	3.4	5.9
Coupon rate (pct.)	3.7	1.1	3.0	3.5	4.1
Par amount ($ millions)	846.4	776.1	400.0	650.0	1000.0
B. Pandemic period (Jan–Jul 2020)[b]					
Credit spread (bps.)	158.8	152.9	61.4	109.3	204.9
Time-to-maturity (years)	4.5	2.6	2.3	4.1	6.5
Age (years)	4.6	4.4	1.8	3.7	6.3
Coupon rate (pct.)	3.6	1.1	2.9	3.5	4.0
Par amount ($ millions)	845.2	766.3	420.0	650.0	1000.0

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 separately for two non-overlapping sample periods of equal length: Panel A, the pre-pandemic period (Jan–Jul 2019) and Panel B, the pandemic period (Jan–Jul 2020).

[a] Panel dimensions: No. of bonds = 4,996; No. of firms = 957; and Observations =429,547.

[b] Panel dimensions: No. of bonds = 5,453; No. of firms = 975; and Observations = 451,917.

B Residual Credit Spreads

This appendix provides details underlying the construction of credit spread residuals, our proxy for credit risk premia. We follow Gilchrist and Zakrajšek (2012) and decompose investment-grade credit spreads into a component that captures issuer-specific time-varying 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.

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

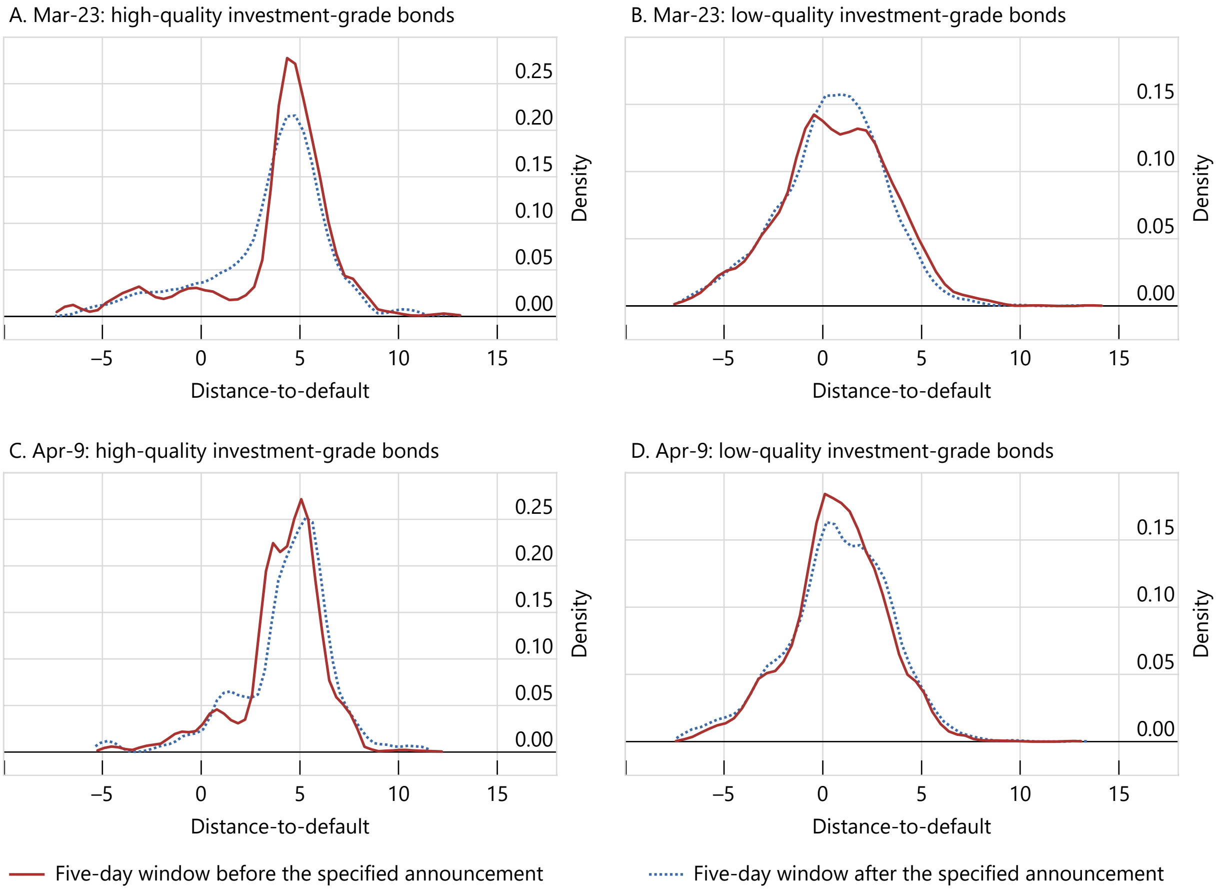
$$\mathrm{DD}^{(\tau)} = \frac{\ln(V/D) + \left(\mu_V - 0.5\sigma_V^2\right)\tau}{\sigma_V\sqrt{\tau}},$$

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:

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

Figure B-1: Default Risk around the SMCCF Announcements



Note: The panels show the weighted kernel density estimates of the distance-to-default $\mathrm{DD}_{i,t}^{(\tau)}$ (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 announcement 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 is used as a weight.
Source: Authors' calculations using data from TRACE/FINRA, Mergent FISD/WRDS, CRSP/WRDS, and S&P's Compustat/WRDS.

where r denotes the instantaneous risk-free interest rate (one-year constant-maturity 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 (that is, $\Delta \ln V$) and use the resulting series to generate new estimates of σ_V and μ_V. We then iterate on σ_V until convergence.

Figure B-1 plots the kernel density estimates (that is, smoothed histograms) of the distance-to-default $\mathrm{DD}_{i,t}^{(\tau)}$ for our sample of investment-grade bonds. Panels A and B focus on the five-

Table B-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).

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.

In the next step, we estimate the following regression:

$$\ln \text{CS}_{i,j,t}^{(\tau)} = \alpha_0 + \alpha_1 \text{DD}_{i,t}^{(\tau)} + \lambda' \mathbf{Z}_{i,j,t} + \nu_{i,j,t}^{(\tau)}, \tag{B-1}$$

where $\text{CS}_{i,j,t}^{(\tau)}$ is the credit spread of bond j (issued by firm i) with the remaining maturity of τ years, and $\text{DD}_{i,t}^{(\tau)}$ denotes the distance-to-default for issuer i over horizon τ. The bond-level credit-spread pricing regression also includes predetermined bond-specific characteristics—denoted by the vector $\mathbf{Z}_{i,j,t}$—that can influence credit spreads through liquidity premia. These include the bond's duration ($\text{DUR}_{i,j,t}$), the par amount outstanding ($\text{PAR}_{i,j}$), the bond's (fixed) coupon rate ($\text{COUP}_{i,j}$), and the age of the issue ($\text{AGE}_{i,j,t}$).

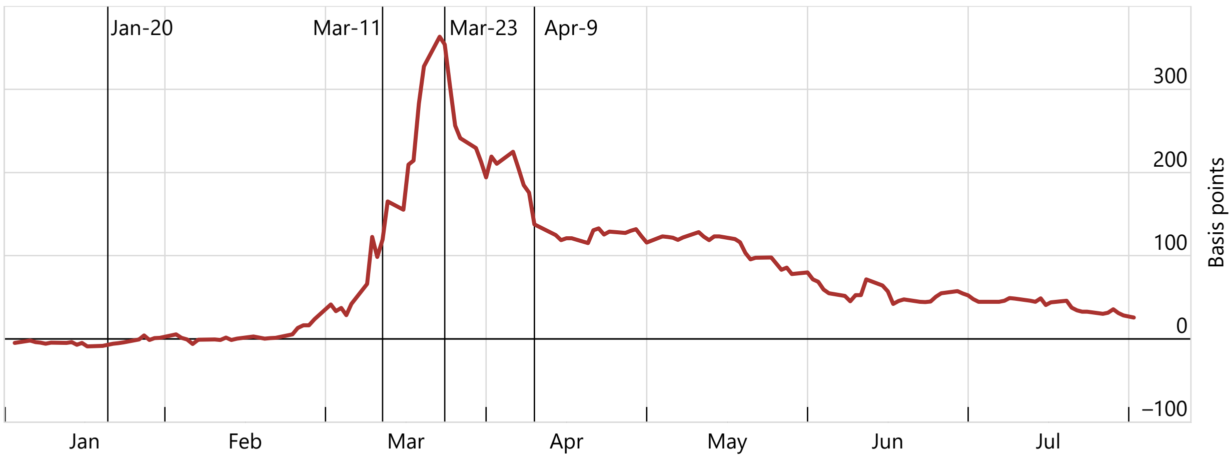
We estimate equation (B-1) by OLS using daily TRACE data on investment-grade corporate bonds issued by publicly listed U.S. companies over the sample period June 2002 to December 2019 and then use the resulting parameter estimates to predict credit spreads over the January–July 2020 period. We start 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.[23]

As shown in Table B-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 $\text{DD}_{i,t}^{(\tau)}$ 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–December 2019 period.

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.

[23] Using the entire sample yields similar results.

Figure B-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. 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/FINRA, Mergent FISD/WRDS, CRSP/WRDS, and S&P's Compustat/WRDS.

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 B-2 plots the daily estimate of the average 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 premium started to rise in late February and took off amid the 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.

C Macroeconomic Implications

This appendix outlines the empirical methodology used to gauge the macroeconomic effects of the March 23 and April 9 SMCCF announcements. In particular, we use the local projections approach of Jordà (2005) to trace out the impact of an announcement-induced (orthogonalized) shock to the investment-grade EBP, denoted by EBP_t^{IG}, on monthly indicators of economic activity.

Specifically, we estimate

$$y_{t+h} - y_t = \alpha^{(h)} + \beta^{(h)} \hat{s}_t + \mathbf{x}_t' \boldsymbol{\gamma}^{(h)} + \epsilon_{t+h}, \tag{C-1}$$

where y_t denotes an indicator of economic activity in month t and $\hat{s}_t$ is the estimated residual from

TABLE C-1: Local Projection Estimates of a Shock to EBP_t^{IG}

	$h = 6$ (months)			$h = 12$ (months)		
Estimate	IP	EMP	UR	IP	EMP	UR
$\beta^{(h)}$	−2.43	−0.40	0.39	−3.48	−0.87	0.72
	(0.56)	(0.23)	(0.15)	(0.74)	(0.35)	(0.23)
$1.7 \times \beta^{(h)}$	−4.13	−0.68	0.66	−5.92	−1.48	1.22

NOTE: Sample period: 1973:M1–2019:M12. The entries in the table denote the OLS estimates of the coefficient $\beta^{(h)}$ from the local projection approach described in the text. Asymptotic standard errors reported in parentheses are computed according to Newey and West (1987) with the "lag-length" parameter equal to h.

the following regression:

$$\text{EBP}_t^{IG} = \mu + \mathbf{z}_t'\boldsymbol{\theta} + s_t. \tag{C-2}$$

In specification (C-2), the vector $\mathbf{z}_t$ consists of contemporaneous and lagged values of economic and financial indicators that define our orthogonalization of shocks to the investment-grade EBP. Specifically, the vector $\mathbf{z}_t$ includes six lags of EBP_t^{IG}, along with the contemporaneous values and six lags of y_t, TS_t, and RFFR_t, where TS_t denotes the term spread and RFFR_t is the real federal funds rate.[24] Thus, s_t is a shock to EBP_t^{IG} that is orthogonal to the contemporaneous and lagged values of y_t and the stance of monetary policy as captured by the term spread and the real funds rate. As suggested by Plagborg-Møller and Wolf (2021), the vector $\mathbf{x}_t$ in specification (C-1) adds additional controls—consisting of six lags of $y_t, \text{TS}_t, \text{RFFR}_t$, and EBP_t^{IG}—to the local projection.

The original monthly excess bond premium estimated by Gilchrist and Zakrajšek (2012), which we denote simply by EBP_t, is available from 1973:M1 onward. The investment-grade EBP estimated in this paper is available starting in 2002:M6. To benchmark our results to the EBP_t^{IG} reported in this paper, we rely on a rescaled version of EBP_t as a proxy for EBP_t^{IG} before 2002:M6. Specifically, we construct the fitted value from regressing EBP_t^{IG} onto EBP_t over the sample period for which both series are available.[25] We then use this fitted value to proxy for EBP_t^{IG} over the 1973:M1–2002:M5 period.

This approach allows us to estimate the local projection specification (C-1) using data over the entire 1974:M1–2019:M12 period. We note that for the post-2002 period, we obtain very similar estimates of the coefficient of interest $\beta^{(h)}$ when using either EBP_t or EBP_t^{IG}. The advantage of using the longer sample period is that we obtain more precise estimates of $\beta^{(h)}$.

The first row of Table C-1 reports the estimates of $\beta^{(h)}$ for six-month ($h = 6$) and 12-month ($h = 12$) horizons for three monthly indicators of economic activity (y_t): the log of industrial production (IP), the log of payroll employment (EMP), and the unemployment rate (UR). These estimates imply that a 1 percentage point shock to EBP_t^{IG} in month t cuts almost 2.5 percentage points from the (annualized) growth of industrial production and 0.4 percentage point from the employment growth over the subsequent six months; the shock is estimated to also boost the unemployment rate about 0.40 percentage point over the same horizon. At the 12-month horizon, the effects of this shock rise to 3.5 percentage points for the growth of industrial production, about 0.90 percentage point for employment growth, and almost 0.75 percentage point for the unemployment rate.

[24] The term spread corresponds to the difference in yields on the 10-year and one-year nominal Treasury securities, and the real federal funds rate is defined as the nominal effective funds rate less 12-month CPI inflation lagged one month.

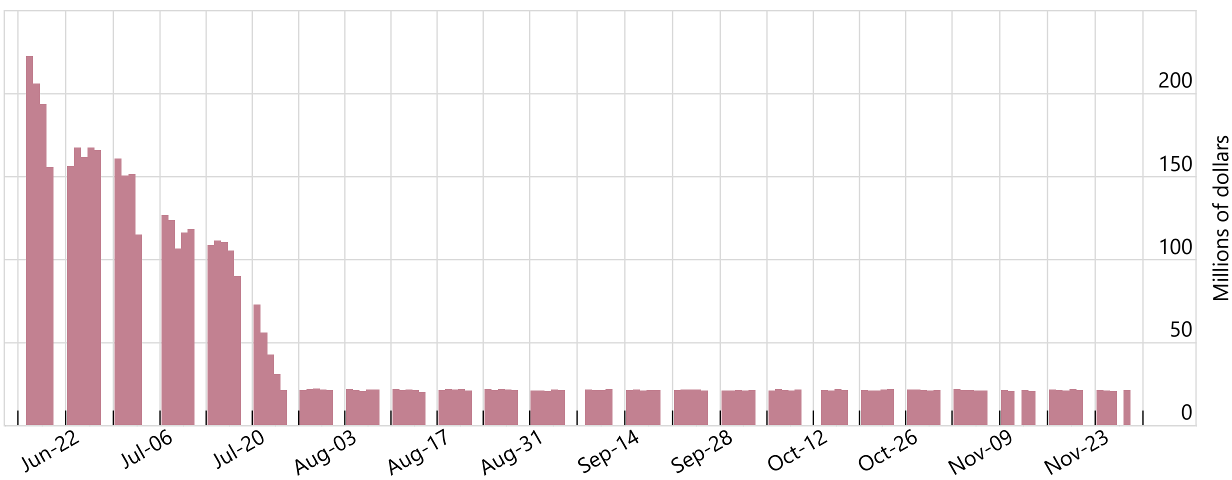
[25] The correlation between the two series is 0.94 during this period.

According to our estimates, EBP_t^{IG} fell 1.7 percentage points following the two SMCCF announcements. The bottom row of the table reports what we consider to be the relevant counterfactual, which is how much the three measures of economic activity would have changed if the two announcements had not been made. In essence, we consider the announcement-induced drop in EBP_t^{IG} to be equivalent to a shock in the local projection framework. Such a counterfactual implies that at the 12-month horizon, the growth of industrial output would be nearly 6 percentage points lower, payroll employment growth would register an additional decline of almost 1.5 percentage points, and the unemployment rate would increase by an additional 1.2 percentage points. While we treat this counterfactual with caution, these estimates, nevertheless, suggest economically sizable effects of the two SMCCF announcements on the path of future economic activity.

D Intraday Event Study of Purchase Effects

In this section, we provide a detailed analysis of the frequency and distribution of transactions in the secondary market around the facility's purchase times. Figure D-1 depicts the daily dollar volume of purchases from the second half of June through the end of November, 2020.

Figure D-1: 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' calculations using data from FRBNY.

In the tables below, we provide a detailed breakdown of the times when the facility executed purchases of individual corporate bonds. First, Table D-1 reports the distribution of execution times for the sample of 1,308 purchases from mid-June to the end of July, a period when the SMCCF purchased almost 70 percent of its total portfolio and which is the focus of our intraday event study. According to the entries, 81 percent of those purchases occurred between 9 a.m. and 1 p.m., that is, during main trading hours; and another 10 percent were executed between 2 p.m. and 3 p.m., again during main trading hours. The key takeaway from this table is that the facility's purchases were not concentrated at a specific time but were fairly spread out over the main trading hours.

In Table D-2, we report the distribution of the number of transactions within 24 hours of the purchase time for bonds purchased by the facility in the 1,308 identified transactions (column "Purchased Bonds"), as well as for their ineligible counterparts issued by the same company (column

"Ineligible Bonds"). The entries clearly show that there were numerous transactions in active hourly trading sessions for both purchased bonds and their ineligible counterparts. In fact, there were 21,537 transactions within 24 hours of the purchase time for purchased bonds and 73,345 transactions for their ineligible counterparts.

Table D-2 reports the total number of transactions for purchased bonds and their ineligible counterparts. Another way to document how frequently those bonds traded is to look at the number of transactions per bond within various trading windows. Table D-3 reports selected summary statistics for the number of transactions within 12, six, and three hours around the purchase times. As shown in Panel A, more than 75 percent of the facility's purchases have three or more transactions within three hours of the purchase time. The average (median) number of transactions is 10 (7) within a six-hour window. According to Panel B, similar trading frequencies are observed for ineligible bonds issued by the same issuers. All told, these results indicate that purchased bonds and their ineligible counterparts traded frequently within three or six hours of the times when the SMCCF executed its purchases.

Table D-1: SMCCF's Purchase Times

Purchase Time Interval (*hh:mm*)	Transactions	Freq. (%)
[08:00, 09:00)	3	0.2
[09:00, 10:00)	229	17.5
[10:00, 11:00)	288	22.0
[11:00, 12:00)	282	21.6
[12:00, 13:00)	259	19.8
[13:00, 14:00)	81	6.2
[14:00, 15:00)	129	9.9
[15:00, 16:00)	37	2.8
Memo: Total	1,308	100.0

Note: The table reports the distribution of transactions in one-hour intervals during which the SMCCF purchased individual fixed-coupon corporate bonds from June 16 through July 31, 2020.

Source: Authors' calculations using data from FRBNY.

Table D-2: Trading Activity around the SMCCF Purchase Times

Time Interval	Purchased Bonds[a]		Ineligible Bonds[b]	
	Transactions	Freq. (%)	Transactions	Freq. (%)
$[h^* - 24{:}00, h^* - 23{:}00)$	1,305	5.7	4,938	6.4
$[h^* - 23{:}00, h^* - 22{:}00)$	1,284	5.6	4,885	6.3
$[h^* - 22{:}00, h^* - 21{:}00)$	1,186	5.2	5,947	7.7
$[h^* - 21{:}00, h^* - 20{:}00)$	1,289	5.6	4,161	5.4
$[h^* - 20{:}00, h^* - 19{:}00)$	1,000	4.3	3,733	4.8
$[h^* - 19{:}00, h^* - 18{:}00)$	486	2.1	2,072	2.7
$[h^* - 18{:}00, h^* - 17{:}00)$	223	1.0	960	1.2
$[h^* - 17{:}00, h^* - 16{:}00)$	40	0.2	142	0.2
$[h^* - 16{:}00, h^* - 15{:}00)$	3	0.0	4	0.0
$[h^* - 15{:}00, h^* - 14{:}00)$	.	.	4	0.0
$[h^* - 14{:}00, h^* - 13{:}00)$	.	.	2	0.0
$[h^* - 13{:}00, h^* - 12{:}00)$	.	.	3	0.0
$[h^* - 12{:}00, h^* - 11{:}00)$	.	.	7	0.0
$[h^* - 11{:}00, h^* - 10{:}00)$	2	0.0	8	0.0
$[h^* - 10{:}00, h^* - 09{:}00)$	2	0.0	17	0.0
$[h^* - 09{:}00, h^* - 08{:}00)$	2	0.0	15	0.0
$[h^* - 08{:}00, h^* - 07{:}00)$	5	0.0	26	0.00
$[h^* - 07{:}00, h^* - 06{:}00)$	44	0.2	129	0.2
$[h^* - 06{:}00, h^* - 05{:}00)$	104	0.5	404	0.5
$[h^* - 05{:}00, h^* - 04{:}00)$	265	1.2	890	1.1
$[h^* - 04{:}00, h^* - 03{:}00)$	517	2.3	1,533	2.0
$[h^* - 03{:}00, h^* - 02{:}00)$	643	2.8	2,114	2.7
$[h^* - 02{:}00, h^* - 01{:}00)$	918	4.0	3,495	4.5
$[h^* - 01{:}00, h^* - 00{:}05]$	1,411	6.1	4,426	5.7
$(\mathbf{h^* - 00{:}05, h^* + 00{:}05})$	1,449	6.3	15	0.0
$[h^* + 00{:}05, h^* + 01{:}00)$	1,892	8.2	5,119	6.6
$[h^* + 01{:}00, h^* + 02{:}00)$	1,667	7.2	5,571	7.2
$[h^* + 02{:}00, h^* + 03{:}00)$	1,476	6.4	5,144	6.6
$[h^* + 03{:}00, h^* + 04{:}00)$	1,228	5.3	4,954	6.4
$[h^* + 04{:}00, h^* + 05{:}00)$	889	3.8	3,296	4.2
$[h^* + 05{:}00, h^* + 06{:}00)$	521	2.3	1,911	2.5
$[h^* + 06{:}00, h^* + 07{:}00)$	198	0.9	734	0.9
$[h^* + 07{:}00, h^* + 08{:}00)$	36	0.2	140	0.2
$[h^* + 08{:}00, h^* + 09{:}00)$	.	.	.	.
$[h^* + 09{:}00, h^* + 10{:}00)$	.	.	2	0.0
$[h^* + 10{:}00, h^* + 11{:}00)$	.	.	6	0.0
$[h^* + 11{:}00, h^* + 12{:}00)$	.	.	11	0.0
$[h^* + 12{:}00, h^* + 13{:}00)$	.	.	8	0.0
$[h^* + 13{:}00, h^* + 14{:}00)$	1	0.0	3	0.0
$[h^* + 14{:}00, h^* + 15{:}00)$	1	0.0	9	0.0
$[h^* + 15{:}00, h^* + 16{:}00)$	4	0.0	13	0.0
$[h^* + 16{:}00, h^* + 17{:}00)$	4	0.0	20	0.0
$[h^* + 17{:}00, h^* + 18{:}00)$	34	0.2	97	0.1
$[h^* + 18{:}00, h^* + 19{:}00)$	57	0.3	293	0.4
$[h^* + 19{:}00, h^* + 20{:}00)$	201	0.9	785	1.0
$[h^* + 20{:}00, h^* + 21{:}00)$	370	1.6	1,288	1.7
$[h^* + 21{:}00, h^* + 22{:}00)$	551	2.4	1,872	2.4
$[h^* + 22{:}00, h^* + 23{:}00)$	767	3.3	2,924	3.8
$[h^* + 23{:}00, h^* + 24{:}00)$	873	4.2	3,641	4.7
Memo: Total	21,537	100.0	73,345	100.0

Note: The table reports the distribution of the number of transactions in the specified time interval $h^* \pm$ *hh:mm*, where h^* (*2020:dd:mm:hh:mm:ss*) corresponds to the 1,308 transaction times at which the SMCCF purchased individual fixed-coupon corporate bonds from June 16 through July 31, 2020. (see Table D-1). The time interval in bold (that is, $(\underline{h}, \overline{h}) = (h^* - 00{:}05, h^* + 00{:}05)$) indicates the ±5-minute interval bracketing the SMCCF purchase times.

[a] A sample of 895 fixed-coupon bonds (issued by 485 firms) purchased by the SMCCF in 1,308 transactions from June 16 through July 31, 2020.

[b] A sample of 1,267 program-ineligible fixed-coupon bonds issued by 383 firms, whose (fixed-coupon) program-eligible bonds were purchased by the SMCCF from June 16 through July 31, 2020.

Table D-3: Summary Statistics of Trading Activity around the SMCCF Purchase Times

Transactions per Security	Mean	SD	Min	P25	P50	P75	Max
A. Purchased bonds[a]							
± 12 hours of h^*	17.6	23.2	1	5	11	21	324
± 6 hours of h^*	10.1	13.0	1	3	7	13	220
± 3 hours of h^*	9.9	12.8	1	3	7	12	220
B. Ineligible bonds[b]							
± 12 hours of h^*	18.5	35.2	1	4	10	23	1,476
± 6 hours of h^*	10.7	13.7	1	3	6	14	254
± 3 hours of h^*	10.5	13.5	1	3	6	13	254

Note: The table reports summary statistics for the number of transactions per security within 12, six, and three hours around h^*, the actual times (*2020:mm:dd:hh:mm:ss*) at which the SMCCF executed its purchases of individual securities from June 16 through July 31, 2020 (see Table D-2).

[a] A sample of 895 fixed-coupon bonds (issued by 485 firms) purchased by the SMCCF in 1,308 transactions from June 16 through July 31, 2020.

[b] A sample of 1,267 program-ineligible fixed-coupon bonds issued by 383 firms, whose (fixed-coupon) program-eligible bonds were purchased by the SMCCF from June 16 through July 31, 2020.

E Calibration Summary

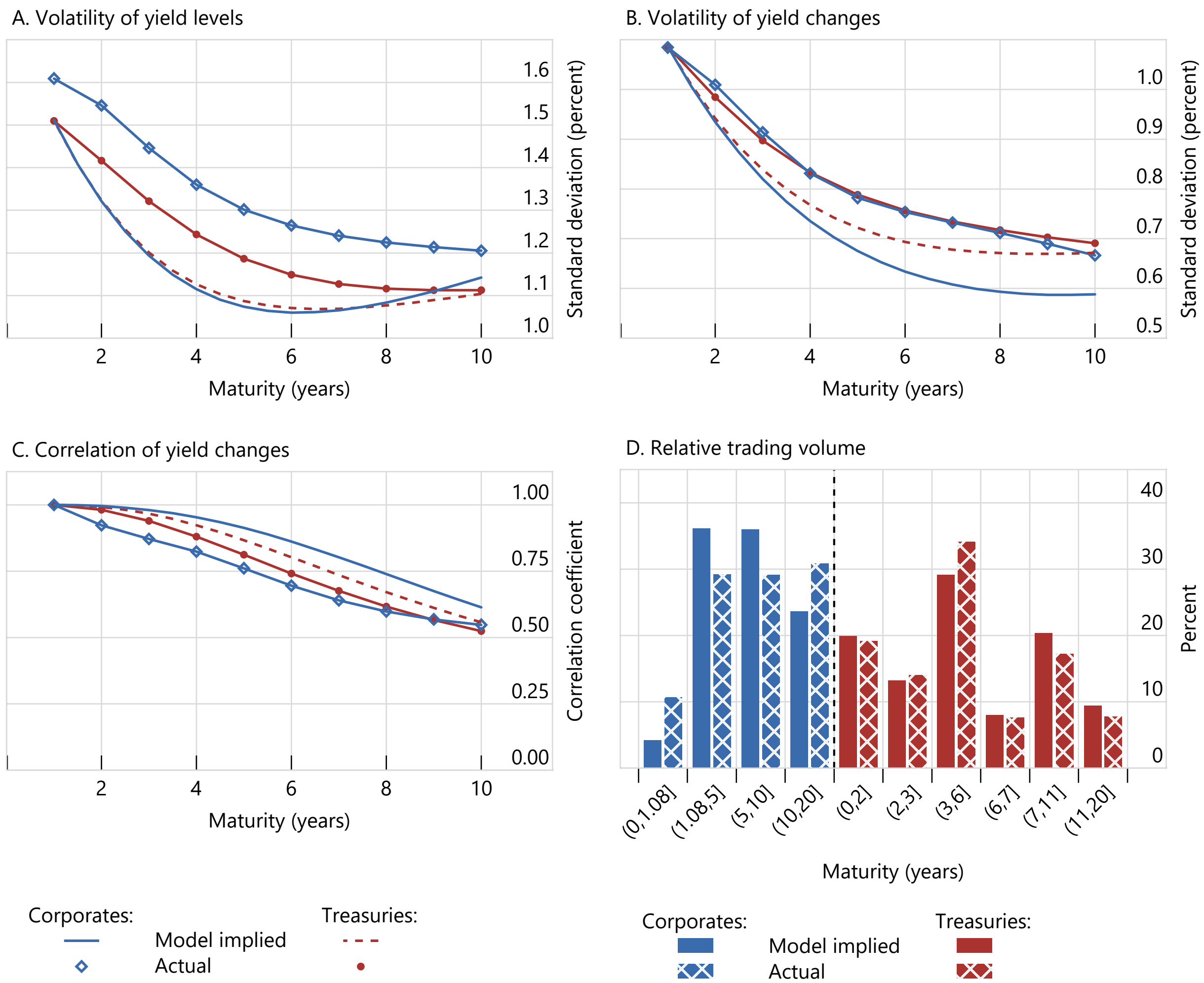
In this appendix, we detail our calibration procedure.

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	$\frac{\sum_{0<\tau\leq 2}\text{Volume}(\tau)}{\sum_{0<\tau\leq 20}\text{Volume}(\tau)}$	0.199
$\tilde{\delta}_\alpha$	0.227	$\frac{\sum_{0<\tau\leq 5}\text{Volume}(\tau)}{\sum_{0<\tau\leq 20}\text{Volume}(\tau)}$	0.399
Demand shock—long maturities		Relative volume—long maturities	
δ_θ	0.361	$\frac{\sum_{11<\tau\leq 20}\text{Volume}(\tau)}{\sum_{0<\tau\leq 20}\text{Volume}(\tau)}$	0.094
$\tilde{\delta}_\theta$	0.237	$\frac{\sum_{10<\tau\leq 20}\text{Volume}(\tau)}{\sum_{0<\tau\leq 20}\text{Volume}(\tau)}$	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.

Figure E-1: Selected Model-implied and Empirical Moments



Note: The figure compare the selected model-implied moments with their empirical counterparts. See Table E-1 for details.
Source: Authors' calculations using data from FRBNY.

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.

As noted in the main text, our calibration procedure 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.

The parameters κ_β, a, θ, $(a \times \alpha)$ are chosen to match the average of volatilities of Treasury yields and the average of 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 shorter than or equal to two years and with maturities longer than 10 years, respectively.[26] 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 (that is, $\kappa_{\tilde{\beta}} = 0.116$ vs. $\kappa_\beta = 0.189$). This is 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.

This appendix also contains a simulation exercise that uses the calibrated model to quantify the effects of purchases on corporate bond prices and compare the model-implied effects with the price impact of the Fed's actual purchases shown in Table 4 of the main text. We do so by assuming that a purchase of corporate bonds of maturity τ in the model corresponds to an unanticipated decrease in the parameter $\tilde{\theta}_0(\tau)$, denoted by $\Delta\tilde{\theta}_0(\tau) < 0$, of the stochastic demand intercept $\tilde{\beta}_t^\tau$. We further assume that the purchased bonds fall into 10 maturity bins: bonds in the first bin have a remaining maturity shorter than or equal to six months ($(0, 0.5]$), those in the second bin have a remaining maturity longer than six months, but shorter than or equal to one year ($(0.5, 1]$), and so on; those in the last bin have a remaining maturity longer than 4.5 years, but shorter than or equal five years (that is, $(4.5, 5]$). Using data on the program's actual purchases, we calculate the proportion of purchased bonds in each bin, which yields the following weights (in percentages): $(0, 0.5] = 0.3$; $(0.5, 1] = 8.2$; $(1, 1.5] = 5.9$; $(1.5, 2] = 10.3$; $(2, 2.5] = 10.0$; $(2.5, 3] = 14.9$; $(3, 3.5] = 7.6$; $(3.5, 4] = 14.8$; $(4, 4.5] = 7.3$; and $(4.5, 5] = 20.8$.

Letting $\overline{\Delta} = 0.25/21.4 = 0.0117$ denote the maximum size of the program (that is, \$250 billion) as a fraction of 2019 nominal GDP (that is, \$21.4 trillion), we then model purchases of bonds in the n-th maturity bin as

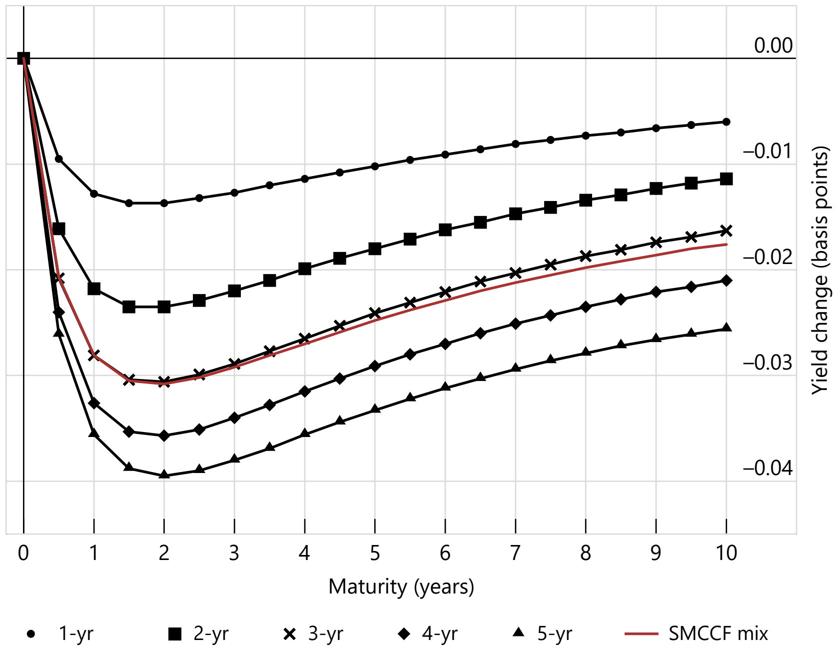
$$\Delta\tilde{\theta}_0(\tau_n) = -\overline{\Delta} \times w(\tau_n); \text{ for } n = 1, 2, \ldots, 10 \text{ and } \tau_n = 0.5, 1, \ldots, 5,$$

where τ_n denotes the end point of the maturity range in the n-th bin, and $w(\tau_n)$ denotes the corresponding weight.

Each patterned black line in Figure E-2 traces out the impact of the model-implied purchases in the specified narrow maturity bin τ_n on the entire investment-grade corporate bond yield curve. As shown, purchasing bonds in any SMCCF-eligible maturity bin τ_n lowers corporate bond yields across all maturities, though the effect is economically very small. The red solid line depicts the

[26] 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.

Figure E-2: Model-Implied Purchase Effects



Note: The figure plots yield changes of model-implied purchase effects on (high-quality) investment-grade corporate bonds. The patterned black lines depict the model-implied effects of purchasing bonds in the specified maturity bin (that is, 1-yr = $(0.5, 1]$; 2-yr = $(1.5, 2]$; 3-yr = $(2.5, 3]$; 4-yr = $(3.5, 4]$; and 5-yr = $(4.5, 5]$) on the entire corporate bond yield curve. The red solid line depicts the model-implied purchase effects across the maturity spectrum based on the actual mix of bonds purchased by the SMCCF.
Source: Authors' calculations using data from FRBNY.

overall impact of purchases on the corporate bond yield curve, based on the weighted combination of the 10 maturity bins actually purchased by the SMCCF. The average model-implied purchase effect on corporate bond yields in the program-eligible segment of the market is essentially zero, the same as the estimated average purchase effect for high-quality investment-grade bonds reported in Panel B of Table 4 in the main text.