The Impact of Monetary Policy on Long-term Liabilities of Households and Firms[*]

Jules van Binsbergen[†]   Marco Grotteria[‡]

May 29, 2023

Abstract

We examine the impact of monetary policy transmission on households' and firms' long-duration liabilities using high-frequency variation in 10-year swap rates around FOMC announcements. We find that mortgage rates respond about three weeks after monetary policy announcements, with a *larger* response to interest rate hikes compared to cuts in post-2010 data. The strong response of mortgage rates to 10-year swap rates leaves little explanatory power for mortgage concentration or bank market power. Corporate yields respond symmetrically to positive and negative shocks, with a higher sensitivity for firms with lower credit ratings. We show that expected future short rates and interbank market frictions play a significant role in explaining both mortgage rates and corporate bond yields. While the term premium correlates positively with mortgage rates, it has little comovement with corporate bond yields.

Keywords: Mortgage Lending; Firm Heterogeneity; Monetary Policy Transmission; Market Power; Banking Industry; United States

JEL codes: E44, E52, G21.

Increasing short-term interest rates are normally accompanied by a rise in longer-term yields. Historically, even ... distant forward rates have tended to rise in association with monetary policy tightening. Alan Greenspan

1 Introduction

The effectiveness of monetary policy by Central Banks depends on the extent that they have a meaningful macroeconomic impact on *financial prices that truly matter* to households and firms (Blinder, 1999). One of the key variables driving the investment/consumption decision is the long-term real interest rate (Stiglitz and Greenwald, 2003).[1] Yet, the importance of long-term rates for macroeconomic outcomes contrasts with two important empirical observations. First, the Federal Reserve has always viewed moderating long-term rates as a residual objective, achieved in a stable economy with low expected inflation (Feliz, 2021). Second, empirical macroeconomists tend to study how shocks computed from short-term rates propagate to macroeconomic outcomes (Bernanke and Gertler, 1995). Both facts are related to the conventional view that if the Federal Reserve changes short-term rates, that are under its control, longer-term yields will respond, as suggested by the epigraph by Alan Greenspan above.

Interestingly, changes in long-term rates on FOMC announcement days are, at best,

---

[1]The standard consumption Euler equation relates consumption to expected future nominal rates and change in prices (i.e., real rates):

$$c_t = E_t^{hh} c_{t+1} - \sigma \left(i_t - E_t^{hh} \pi_{t+1} \right) = -\sigma \sum_{j=0}^{\infty} E_t^{hh} i_{t+j} + \sigma \sum_{j=0}^{\infty} E_t^{hh} \pi_{t+j+1}, \tag{1}$$

where $E_t^{hh} x_{t+j}$ represents the household's expectation at time t for variable x at time $t+j$, i represents the nominal interest rate, and π represents inflation. Similarly, classical models of firm investment establish a linear relationship between the rate of investment $\frac{I_t}{K_t}$ and Tobin (1969) q, i.e., the value of capital relative to its replacement cost:

$$\frac{I_t}{K_t} = a + bq. \tag{2}$$

When these models imply an equivalence between Tobin's q and the ratio of the market-to-book value of capital $\frac{V}{K}$ (e.g., Hayashi, 1982), it is easy to see that, since V is the present value of all future cash-flows of a firm, long-term real interest rates (and not short-term rates) affect the investment rates of firms.
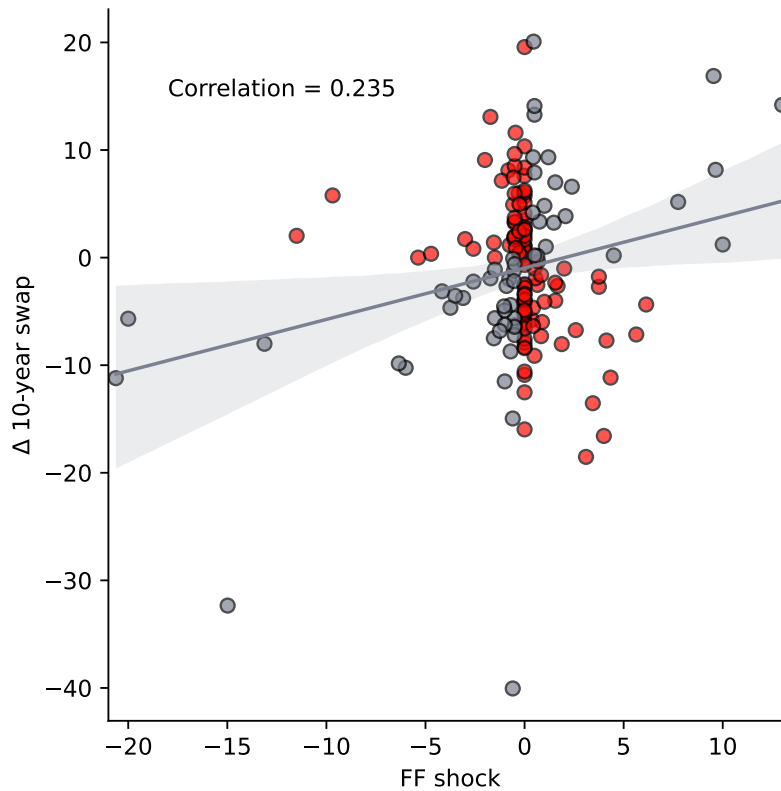
Fig. 1. *Notes*: The figure shows the scatterplot of the changes in 10-year swap rates on FOMC days against changes Kuttner (2001) Federal funds rate shock computed by Acosta (2022). Values are expressed in basis points. The grey dots represent FOMC events for which the changes in 10-year swap rates on FOMC and Kuttner (2001) shocks shared the same sign. The red dots are events in which the two shocks had opposite signs. The sample includes all scheduled FOMC meetings from February 2, 2000 to December 11, 2019.

weakly correlated with changes in short-term rates. For instance, for the 160 scheduled FOMC announcements between 2000 and 2019 listed in Table A.1, the correlation between changes in 10-year swap rates and Kuttner (2001) forecast revisions of the Federal Funds rate on the same days is a mere 23%. The corresponding scatterplot is in Figure 1. As eloquently stated by Bernanke (2015), "monetary policy is 98% talk and 2% action."

In this paper, we examine how monetary policy surprises, via long-term rates capturing bank funding costs, transmit to discount rates in both the long-term mortgage and corporate bond markets. We combine detailed data sets on 1) 30-year mortgages issued in the United States (US) from Corelogic, 2) interest rates for a range of mortgage products from RateWatch, 3) transactions of non-financial corporate bonds from TRACE, and 4) CDS spreads for non-financial companies from Markit, with a discontinuity-

based identification approach that exploits the lumpy way in which monetary news is communicated to investors around FOMC announcements (Cook and Hahn, 1989; Kuttner, 2001; Cochrane and Piazzesi, 2002; Bernanke and Kuttner, 2005; Nakamura and Steinsson, 2018). We construct monetary surprises using changes in long-term swap rates (10-year) over narrow windows around the FOMC statement releases and press conferences. We then estimate the response of long-term rates for consumers and firms in the 4 weeks after the monetary announcement. With regard to mortgage rates, we observe a symmetric one-to-one response to positive and negative monetary policy surprises. However, if one focuses only on the second half of the sample (post 2010), the response to positive surprises is statistically larger than the corresponding response to negative policy shocks. We confirm the robustness of this latter result using alternative data for swap rates and different windows for FOMC announcements.

The usual argument that the response to negative monetary policy surprises is more nuanced once rates are close to 0 cannot apply here, because swap rates in January 2010 were still 4% and thus substantially far away from a potential zero lower bound. The other common argument that banks exploit their market power when setting rates does not seem supported by the evidence either. When we test directly for this hypothesis using four standard measures of market power, we are unable to reject the hypothesis that the response is the same in high- and low-market power areas. The simplest explanation for our results is that variation in mortgage rates simply reflects variation in banks' funding costs as proxied for by long-duration interest-rate swaps. The difference between the loans proposed in different zip codes for a given bank appears slow-moving over time and infrequently revised. We then decompose changes in interest rate swap rates in expected future short interest rates, term premium (following Adrian, Crump, and Moench, 2013), and a residual component capturing interbank market frictions, we see that all three components add significant power when explaining the variation in mortgage rates.

To examine whether our results are driven by potential endogenous self-selection of borrowers after monetary policy announcements, we use rates from Ratewatch. Rate-Watch surveys bank branches throughout the US to collect data rates for ideal mortgages to the "best" borrowers, i.e., those with exceptional FICO scores, for particular constant loan volumes of $175K with 20% down-payment. We analyze the effect on rates for

3

both adjustable-rate and fixed-rate mortgages with maturities ranging from 1 year to 30 years. In all cases, with the exceptions of adjustable-rate mortgages whose fixed rates are set only for the first 1 or 3 years, we find that quoted rates respond as well to changes in swap rates on FOMC days with a direction and magnitude similar to the one from Corelogic.

Unlike mortgage rates, for corporate bonds we observe an immediate response to monetary surprises with similar magnitudes across the overall sample and the sample from 2010. When we decompose changes in interest rate swap rates in Adrian et al. (2013) expected future short interest rates, term premium, and a residual component capturing interbank market frictions, we see that only expected future short interest rates and the residual component add significant power when explaining the variation in bond yields, whereas results are insignificant for the term premium. Splitting the sample by credit rating, yields of speculative-grade bonds respond more strongly to negative monetary shocks and less to positive shocks than the yields of investment-grade bonds. The same is evident for CDS spreads, where we consistently observe a stronger response when we pass from a higher to a lower credit rating.

Our findings establish that bank's assets are strongly connected to monetary policy shocks affecting long-term rates. We then evaluate the effects of monetary policy on banks' net worth. Changes in long-term rates have mainly two effects on banks' values: discounting and cash-flow effects. When rates increase, future cash flows are discounted more heavily, leading to declining market values. However, if assets are repriced in the near term and funding comes from rather stable and sleepy sources (maybe because of market power in the deposit market, Drechsler, Savov, and Schnabl, 2017), banks can benefit from a larger difference between the rates they charge on their assets and their funding costs. We regress Fama-French 49 industry portfolios on the changes in 10-year swaps on FOMC days, controlling for Kuttner (2001) federal-funds shocks. We observe that, except for two days belonging to QE1 events, the banking industry shows the highest exposure to shocks in long-term rates, with a positive and significant coefficient of 7.91. This implies that bank stocks increase by 7.91% for every 1% positive shock to the 10-year swap rate. On the other hand, the exposure to short-term rates (fed funds shocks) is negative (-3.53) and not statistically significant, consistent with the estimate

of Drechsler, Savov, and Schnabl (2021).

We confirm these latter results using stock returns of individual bank holding companies. When we condition on the fractions of loans that get repriced within 1 year we see this variable to be the main determinant of banks' exposure to changes in 10-year swap rates. A larger fraction of loans repriced within 1 year is related to larger positive exposure to shocks to long-term rates. Similarly, banks enjoying higher equity-to-asset ratios are more positively exposed to changes in 10-year swap rates. This confirms the hypothesis that the response of banks' stock returns to long-term rates is explained by a cash-flow channel.[2]

Our results reject the following intermediary-based narrative, which has been a core argument of a large literature in finance and economics. Imagine intermediaries being constrained agents in the business of maturity transformation. As we have shown, higher long-term interest rates raise equity valuation for banks, so it's a positive net worth shock. The shock to net-worth increases intermediaries risk-bearing capacity and should result in lower borrowing costs for firms and households (Siriwardane, 2019). Yet, we observe the opposite response for both mortgage rates and bond yields. A possible interpretation could be that while intermediary frictions do explain some part of the variation in mortgage and bonds pricing, it makes up only a small fraction of the total variation (e.g., as discussed further in Section 2).

Our paper is related to two main strands of the literature. First, our results contribute to an extensive literature examining pass-throughs of monetary policy to interest rates (e.g., Scharfstein and Sunderam, 2016; Drechsler, Savov, and Schnabl, 2017; Benetton and Fantino, 2021; Benetton, Gavazza, and Surico, 2021; Wang, Whited, Wu, and Xiao, 2022). Relative to this literature, we show a) long-term and short-term rates have been

---

[2]Our results are important in light of the bank collapses of 2023. For a cash-flow effect to be present, depositors need to be sleepy, which happens when interest rates do not change too much too fast and depositors are guaranteed by the Federal Deposit Insurance Corporation (Jiang, Matvos, Piskorski, and Seru, 2023). In the case of Silicon Valley bank (SVB), the total withdrawal figure of $142 billion represented a staggering 81% of SVB's $175 billion in deposits as of year-end 2022. More importantly, SVB revealed they had over $150bn of uninsured deposits as of the end of last year, which made it prone to bank runs. In our sample period, from 2000 to 2019, SVB stock returns were also largely positively exposed to changes in 10-year swap rates. Figure A.25 shows the relation between stock returns in percentage and changes in 10-year swaps in bps for the three defaulted banks. The estimates for SVB imply that for every percentage point increase in swap rates the bank stock returns would be about 10 percentage points (with a t-stat of over 3 with robust standard errors).

disconnected in the recent decade due to the fact that short-term rates were close to 0, b) mortgages and long-term bonds are priced against long-term and not short-term rates, and c) an unexpected decrease (increase) in long-term rates is a negative (positive) surprise to banks' net worth.

Second, our results on the asymmetric response of mortgage rates to interest rate news relate our work to research documenting in various settings that output prices respond faster and to a larger extent to input increases than decreases (Borenstein, Cameron, and Gilbert, 1997; Peltzman, 2000; Benzarti, Carloni, Harju, and Kosonen, 2020; Butters, Sacks, and Seo, 2022). With regards to bank deposits, Neumark and Sharpe (1992) have shown that in markets where there are only a few banks dominating, interest rates on deposits slowly rise when market interest rates rise, but quickly decrease when market interest rates fall. We document that for mortgage rates the asymmetric response is not related to market power and concentration, but seems to reflect an asymmetric variation in banks' funding costs as proxied for by interest-rate swaps. If there is market power, it seems to be in the interbank/swap market.

2 Data description and motivating evidence

In this section, we offer some description of the data and motivating evidence for our identification approach.

Swap rates. To study the impact of monetary policy on long-term interest rates, we utilize daily data on 10-year fixed-to-floating swap rates denominated in U.S. dollars from Bloomberg and higher-frequency data from the Intercontinental Exchange (ICE). ICE swap rates are the primary worldwide benchmark for determining swap rates and spreads for interest rate swaps. They are extensively employed as the reference value for cash-settled swaptions, for final payments on the premature termination of interest rate swaps, for floating rate bonds, and more generally by lenders setting mortgage rates. Unfortunately, minute-level data from ICE are available only for the second half of our sample. So, for most of our analysis, we'll use daily changes in interest-rate swaps on the days of FOMC announcements, and we'll show the robustness of our results to intraday

changes in interest-rate swaps.

The swaps rate we use are set against LIBOR. Conceptually, a credit-sensitive interest rate benchmark such as LIBOR represents the interest paid by one bank to another for unsecured deposits, which for most of our sample period reflects well the marginal cost of funds to large financial institutions. The fixed rate on plain vanilla interest rate swaps where the floating payments are based on LIBOR can therefore be interpreted as the par rate against the LIBOR curve, capturing expectations on future rates and bank credit quality, i.e., the two major components of funding costs of banks. Therefore, the swap rate is designed to capture risks in the banking sector as well and is closely related to the bank's funding costs (Cooperman, Duffie, Luck, Wang, and Yang, 2023).[3] We annualize swap rates to reflect 365 days.

Figure A.1 shows the 10-year swap rate series against the 10-year government-bond par yield computed by Gürkaynak, Sack, and Wright (2007). The two series overlap almost perfectly after 2008. This can be interpreted as a) AA-rated banks have a similar credit risk to the US government in the long-term (because of expectations of being bailed out) and b) long-term government bonds do not enjoy the convenience yield documented by van Binsbergen, Diamond, and Grotteria (2022) for securities of less than 2.5-year maturity. Indeed swap rates appear good proxies for convenience-yield-free measures of interest rates available at longer maturities. In addition, the correlation between the two series is 99.41%. If one focused only on FOMC days, which we will do for our main analysis, changes in 10-year swap rates have a correlation of about 92% with changes in 10-year government-bond par yields.

How do swap rates co-move with shorter-term interest rates? To answer this question, Figure 2 shows the rolling-window correlation computed over 365 days between daily changes in 10-year swap rates (Δs) and daily changes in 1-year government bond yields (Δy^1) computed by Gürkaynak et al. (2007). Beyond the time variation, one can notice that after 2008, the relation between the daily changes of the two series is less strong,

---

[3]The idea that the 10-year swap rate should match the yield on a 10-year bond issued by a financially sound bank is incorrect. The 10-year swap is written against rolling three-month loans based on LIBOR (i.e., the three-month credit of banks on the polling list over time). Roughly speaking, LIBOR estimates the rate at which an AA-rated bank can obtain an unsecured short-term loan from another bank. Therefore, swap rates relative to LIBOR take into account updates in the bank poll to include only AA-rated banks.
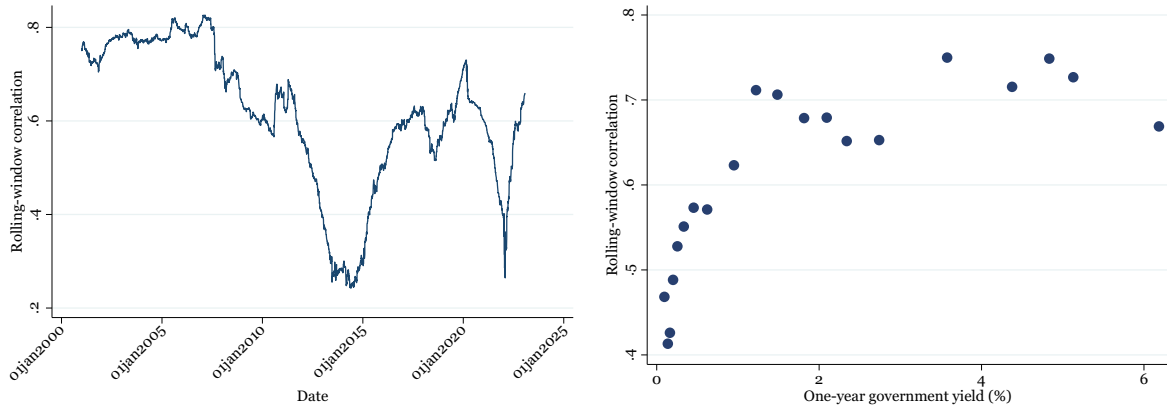
Fig. 2. *Notes*: The left plot shows the rolling-window correlation computed over 365 days between daily changes in 10-year swap rates (Δs) and daily changes in 1-year government bond yields computed by Gürkaynak et al. (2007) (Δy_1). The right plot is a binned scatterplot of the rolling window against the level of 1-year government bond yield.

reaching a minimum correlation of 0.24 in May 2014.

One potential cause could be that when short-term yields were stuck around 0, the Fed turned to forward guidance regarding the path of interest rates (sending signals on both the future of the economy and the monetary policy response function of the Fed) or large-scale asset purchases to steer interest rate expectations. The large usage of *unconventional* monetary policy tools in our sample and the zero lower bound for short-term rates potentially rationalizes a weak relation between the responses of short- and long-term rates around FOMC announcements. The plot on the right sheds some light on the mechanism. It shows a binned scatterplot of the same rolling-window correlation against the level of 1-year government bond yields. The relation is positive with periods of high 1-year yields experiencing on average higher correlations between the daily changes of 10-year swaps and 1-year bonds. Table A.2 formally tests the relationship between the rolling-window correlation and the level of short-term interest rates. We regress the rolling-window correlation computed using data from $t - 365$ to t on the level of 1-year government yields on $t - 365$. In some specifications, we also control for the current level of yields, but results are almost indistinguishable, supporting a strong relationship between the correlation of long- and short-term rates and the level of short-term rates.

Another piece of evidence comes from the dot plots. From January 25, 2012, the Federal Reserve started revealing individual forecasts made by all FOMC meeting

participants about the federal funds rate in the short and long term. The dot plot, which is a chart revealing these individual forecasts for the federal funds rate, is what the market and financial press refer to as the rate forecasts. Our sample includes 32 subsequent dot plot observations from January 2012 to December 2019. Following Hillenbrand (2021), we estimate the following equation

$$\Delta s_t = \alpha + \beta \Delta \mathbb{E}[\text{Long-term fed funds rate}] + \epsilon_t, \tag{3}$$

where Δs_t is the daily change in swap rates on FOMC days and $\Delta \mathbb{E}[\text{Long-term fed funds rate}]$ is the change in the median forecast of the long-term fed funds rate relative to the previous dot plot. Column (1) shows a positive and statistically significant relation between the median forecast revision by FOMC members and the swap rate changes. When we control for the level of disagreement in the revision (the standard deviation of the forecasts weighted by the number of people forecasting the same value) we find that higher disagreement is related to positive changes in 10-year rates. Finally, we condition on the level of 1-year government bond yields and split the observations in terciles of 1-year government bond yields. We find that the sensitivity of 10-year swap rates to forecast revision of long-term fed funds rate is higher when 1-year government bond yields are in the lowest tercile. The sensitivity goes down monotonically as 1-year government bond yields increase. This evidence again supports the hypothesis that when short-term yields are stuck around 0, expectation management through channels other than changes in the policy rate becomes important for monetary policy transmission.

Mortgages. Mortgage information is from Corelogic LLMA and Corelogic Deeds Mortgages. The LLMA data contain detailed information on mortgage and borrower characteristics at origination — the interest rate, loan-to-value (LTV) ratio, sale price, credit score, whether the mortgage was GSE-eligible, insured at origination, or whether it was prime or subprime — for a large sample of anonymized borrowers. CoreLogic collects these data from 25 of the largest mortgage servicers in the US. The LLMA data track approximately 5.7 million mortgages each year including on average about 45% of mortgages originated in the US over the sample period. We restrict the sample to

9

Table 1. Summary statistics mortgage sample

Panel A

	N	average	st.dev.	p10	p25	p50	p75	p90
Initial interest rate	6,307,329	5.400	1.177	3.875	4.375	5.5	6.375	6.875
FICO score at origination	5,490,103	742.946	51.759	669	710	754	784	800
original LTV	6,293,105	80.393	14.018	62.8	77.1	80	90	95
original term	6,307,329	360	0	360	360	360	360	360

Panel B

GSE-eligible	Non-conforming (5.15%)	Conforming (92.13%)	Jumbo conforming (2.72%)
Inferred Collateral type	Prime (92.12%)	Subprime (4.02%)	No info (3.86%)
Mortgage insurance	No (64.24%)	Yes (29.28%)	No info (6.48%)

Notes: The table reports the summary statistics for the sample of mortgages. Data span January 2000 to December 2019.

30-year conventional loans (i.e., not originated under a government program) where the borrower's stated purpose was to purchase (e.g., not refinance, education, vehicle purchase, or medical loan) single-family residences or residential condominiums and there was no buy-down. We remove mortgage rates in the bottom and top 1% by year-quarter.

From the Deeds Mortgages, we only use the mortgage origination date, the original balance, the maturity date, the state, and the property zip. All these variables are also present in the LLMA dataset. We exclude all other variables.[4] The only information we need from the Deeds Mortgages is an accurate origination date, adding the day of origination to the year-month in LLMA. So, we keep only the observations in the LLMA data where it is possible to uniquely identify the exact mortgage origination date. Summary statistics are reported in Table 1. Both the median and the average LTV ratio are about 80%. We have restricted the original term to 30 years, so there is no variation there. Among all mortgages, 92% are GSE-eligible and 92% are prime mortgages. Only 29% of mortgages are insured at origination.

---

[4]By doing so, it is impossible for us to (i) determine any individual personally identifiable consumer information or the servicer of any individual loan included in the LLMA Data; or (ii) identify loan or location information more granular than the 5-digit zip code level for any individual loan included in the LLMA Data.

Fig. 3. *Notes*: The figure reports individual-level mortgage rates for 30-year fixed-rate mortgages across the country (the date assigned to a mortgage is the borrower's signature date on the mortgage) against the daily 10-year swap rate from January 2000 to July 2020.

Figure 3 shows all the mortgage rates in our sample by the date on which the deed of the mortgage was signed by the borrower (blue dots) against the 10-year swap rate series from Bloomberg, USSA10 (red solid line).[5] The relation between the level of swap rates and the level of mortgage rates is clear from the figure. An increase in the swap rate appears followed by a rapid rise in mortgage rates, and a decline in the swap rate series is accompanied by a drop in mortgage rates.

We test this relation formally in Table A.4 in the appendix. The table reports the R^2 for different specifications where we regress all our mortgage rates in our sample against different sets of controls. In column 1, we use the (4-week lagged) 10-year swap rate as our unique regressor. The swap rate series explains already about 86% of the variation in mortgage rates (column 1). Including date fixed effects rather than the swap rate series marginally increases the R^2 to 88% (column 2), suggesting that the average variation in

---

[5]This is the par rate paid annually on the swap fixed leg.

11

a day is already quite well-captured by swap rates. Including borrowers' characteristics to the specification of column 2 leads to a small improvement of the R^2 to 89%. Finally, including lender-by-date (rather than date) fixed effects and both lender-by-date and metropolitan statistical area (MSA)-by-date fixed effects increases the R^2 to about 92%. Results are similar in panel B, where we include lender-by-msa-by-date fixed effects. Evidence from Table A.4 suggests that the majority of the variation is explained by the 10-year swap rate alone and so the variation in that series is key to understanding the variation over time of mortgage rates.

Moreover, the use of swap rates as a benchmark for 30-year mortgages is also largely driven by their popularity among institutions that hedge MBS, including Fannie Mae and Freddie Mac. These two agencies play a significant role in issuing and guaranteeing credit for a large portion of pass-through MBS. They also hold a substantial amount of mortgage loans and MBS in their portfolios. Managing the interest rate risk of their retained portfolio requires them to engage in interest rate swaps, whereby they exchange fixed-rate interest payments for floating-rate payments that more closely reflect their short-term borrowing costs. It is standard industry practice to average the five-year and ten-year swap rates to approximate the relevant swap yield since these maturities enjoy much greater liquidity than other swaps with different maturities. Hedging strategies typically rely on these widely-traded maturities, hence their widespread adoption as a reference point (Hancock and Passmore, 2012; Malkhozov, Mueller, Vedolin, and Venter, 2016).[6]

Corporate bonds. The Enhanced TRACE data consists of transaction-level information from dealers trading corporate bonds. This information includes the identity of the bonds traded, the date and time of execution, price, and volume.[7] We keep regular secondary market trades. We combine these data with Mergent/FISD (*issues* and *issuers* files). From Mergent/FISD we obtain information including the bonds' initial terms for offering, the offering date, maturity, and outstanding principal amount (Seltzer,

---

[6]According to Fannie Mae's 10-K, "in measuring the estimated impact of changes in the level of interest rates, we assume a parallel shift in all maturities of the U.S. LIBOR interest-rate swap curve." It follows that a key metric is the duration of the MBS. As an example, as of March 22, 2023, the duration of the 30-year MBS FN MA4993 issued on March 1, 2023, with a coupon of 4% was 7.09, below the
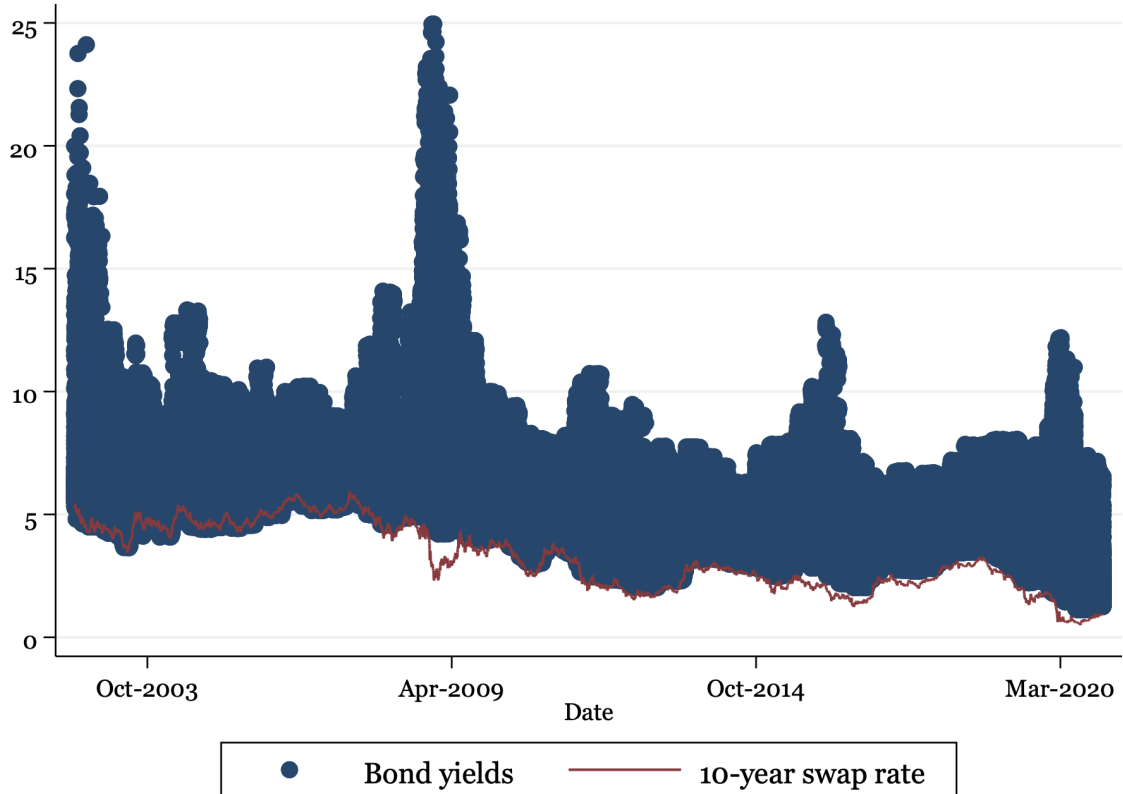
Fig. 4. *Notes*: The figure reports individual-level bond yields across the country (the date assigned to a mortgage is the borrower's signature date on the mortgage) against the daily 10-year swap rate from January 2000 to July 2020.

Starks, and Zhu, 2022). We restrict the sample to US corporate debentures, corporate medium-term notes, and US Corporate Bank Notes. We keep senior unsecured bonds with a fixed coupon rate. We drop the observations in which the interest on the issue may be paid in more of the same security or other securities (pay-in-kind). We drop if the issuer was a foreign agent or Canadian keeping only if the country of domicile was the US and the bond was not issued in a foreign currency. We drop bonds that were privately placed or fell under rule 144a. We drop defaulted bonds and preferred, perpetual, exchangeable, or putable securities. We keep bonds where the remaining time to maturity is between 9 and 11 years, matching the tenor of the swaps.[8] We remove the

---

duration of the 10-year swap (8.185) and above the duration of the 5-year swap (4.264).

[7]Trace enhanced has been cleaned using the code by Qingyi (Freda) Song Drechsler available on WRDS. The code follows the suggestions by Dick-Nielsen (2009) and Dick-Nielsen (2014)

[8]To study the monetary policy response of bonds with issuers in the non-banking sector, we drop bonds where the issuer was in the banking sector (as defined for sector 44 in the definition of the 48 Fama-French industry portfolios, available from Ken French's website) and where the SIC code of the

13

top and bottom 1% of observations by year-quarter.

Figure 4 shows as blue dots all bond yields aggregated at the bond issue-daily level, with the aggregation of intraday transactions weighted by transaction size. The red solid line is the 10-year swap rate series from Bloomberg, USSA10. The relation between the level of swap rates and the level of bond yields is clear from the figure. An increase in the swap rate appears to be followed by a rapid rise in mortgage rates, and a decline in the swap rate series is accompanied by a drop in mortgage rates. However, there are instances, such as during the financial crisis of 2008-2009, where the swap rate series declined, whereas bond yields (especially at the top of the distribution) rose.

We test this relation formally in Table A.5 in the appendix. The table reports the R^2 for different specifications where we regress all our corporate bond yields in our sample against different sets of controls. In column 1, we use the 10-year swap rate as our unique regressor. The swap rate series explains about 43% of the variation in corporate bond yields (column 1). Including date fixed effects rather than the swap rate series marginally increases the R^2 to 55% (column 2), suggesting that the average variation in a day is already quite well-captured by swap rates. However, unlike for mortgages, including time-varying borrowers' characteristics leads to large improvements: adding borrower-by-year-month fixed effects leads to an R^2 of almost 99%. The hypothesis is that credit risk plays for our bond panel sample a larger role than for mortgages. Indeed, mortgages are collateralized loans, whereas here we are focusing on unsecured bonds. To provide supporting evidence for the larger role of credit risk premia, in Panel B of Table A.5 we consider only AA-rated firms. In this sub-sample, the R^2 computed using swap rates as the only regressor is already 86%, which indeed show that when credit risk is minimal, the 10-year swap rates captures already very well the variation of corporate bond yields.

Credit default swaps. Credit default swaps (CDS) contracts can be viewed as agreements for credit protection, involving periodic payments of the "insurance premium" until a default or credit event. We obtain the CDS data from Markit Group Limited, a company founded in 2001 that collects daily CDS spread quotes from a network of

───────────────────────

issuer is missing.

partner banks. Our dataset covers the period from January 2001 to December 2019. We restrict our sample to observations in which the underlying currency is USD, the underlying company is a non-financial company, and where the country of the issuing organization is the US. The number of underlying companies with available data increased from nearly 204 in 2001 to approximately 912 in 2011 before stabilizing at that level and then decreasing to 710 in 2019. We focus on 10-year contracts, which are closer to the tenor of interest-rate swaps.

3 Identification and methodology

3.1 Identification

In studies that focus on identifying monetary policy using high-frequency data, it is typical to examine variation in interest rates in a timeframe of one or two days before and after FOMC announcements. This approach, adopted, among others, by Cook and Hahn (1989), Kuttner (2001), Cochrane and Piazzesi (2002), Bernanke and Kuttner (2005) and Hanson and Stein (2015), assumes that no other factor affects the policy indicator during this period. For all scheduled FOMC days from 2000, we use the days when monetary policy decisions after scheduled meetings became known to the public as listed by Hillenbrand (2021) and reported in Table A.1. For all scheduled FOMC days, we compute daily changes in the level of swap rates.

Nakamura and Steinsson (2018) propose to use shorter time windows surrounding Federal Reserve announcements. Given that between 2011 and 2018 in about half of the FOMC dates the statement release has been followed by a press conference, and that from 2019 all FOMC statement releases have been followed by a press conference, we decided to use a longer window than 30 minutes. Our high-frequency monetary policy surprise is the change in the 10-year swap rate from 10 minutes before the statement release to 3:20 pm on the same day.[9] This method is consistent with recent literature highlighting a link between the policy statement news and the press conference and the importance of the press conference as a channel to communicate monetary policy news

---

[9]We follow Nakamura and Steinsson (2018) and take the difference between the last price observed more than 10 minutes before the FOMC announcement and the first price observed after 3:20 pm.
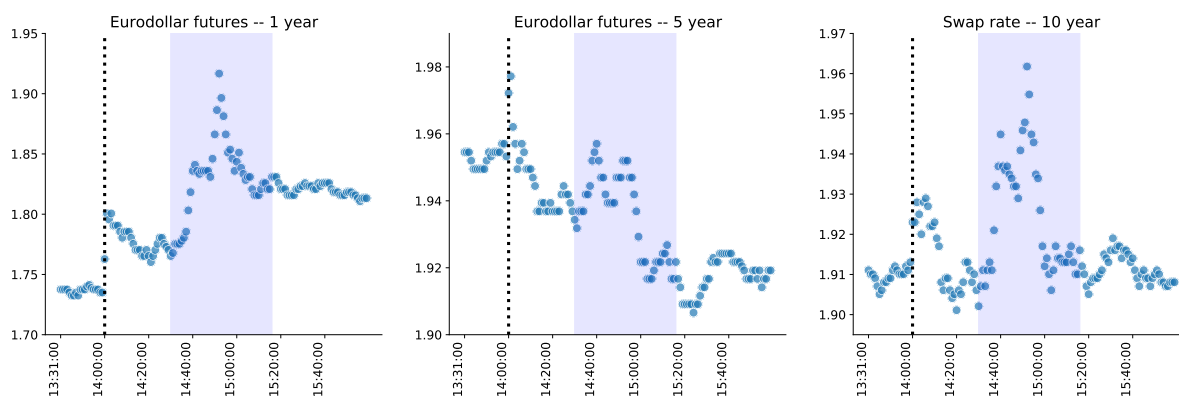
Fig. 5. *Notes*: The Figure shows the intraday evolution of the implied rate from the 12-month Eurodollar futures, the 5-year Eurodollar futures, and the 10-year swap rate on July 31, 2019. The black dashed vertical line highlights the time at which the FOMC statement was released (14:00). The shaded area denotes the FOMC press conference. The conference started at 14:30 and lasted for about 45 minutes. All rates are continuously compounded.

and, in particular, forward guidance to investors (Gómez-Cram and Grotteria, 2022).[10]

To support the choice of a slightly longer window and the usage of 10-year swap rates as a measure of the effects of monetary policy on long-term rates, Figure 5 shows, as an example, the intraday evolution of the implied rate from the 12-month Eurodollar futures, the 5-year Eurodollar futures, and the 10-year swap rate on July 31, 2019. The black dashed vertical line highlights when the FOMC statement was released (14:00). The shaded area denotes the FOMC press conference. The conference started at 14:30 and lasted for about 45 minutes. Two important messages must be taken from the figure.

First, the response of interest rates at the long-term end of the curve to FOMC announcements is quite different from the variation for short- or medium-term rates: while the implied rate from the eurodollar futures 1-year contract increased around the FOMC announcement (both the statement release at 2 pm and the press conference), the rate implied from the 5-year contract went down, and the par rate on the 10-year swap was almost unchanged. This observation is consistent with Gürkaynak, Sack, and Swanson (2005), who find a two- rather than a one-factor structure of monetary policy surprises, where the second factor is a "future path of policy" factor.[11] Considering

---

[10] We can compute high-frequency monetary policy surprises for 62 out of the 79 FOMC dates from March 2010 to December 2019. We miss data for 10-Aug-10, 21-Sep-2011, 25-Apr-2012, 20-Mar-13, 01-May-2013, 31-Jul-13, 30-Oct-13, 29-Apr-15, 17-Jun-15, 28-Oct-2015, 27-Apr-2016, 15-Jun-2016, 27-Jul-2016, 02-Nov-16, 01-Nov-17, 19-Dec-2018, 01-May-19.

[11] A similar point can be noted for other FOMC events, e.g., Figure A.5 shows the case of January 03, 2010.

all the 160 FOMC announcements in Table A.1, the correlation between the change in 10-year swap rates on FOMC days and the policy news by Nakamura and Steinsson (2018) as computed by Acosta (2022) is about 46%, whereas the correlation between the change in 10-year swap rates on FOMC days and the forecast revisions of the Federal Funds rate immediately after the FOMC meeting is only 23%. The scatterplot showing changes in 10-year swap rates on FOMC days against Nakamura and Steinsson (2018) shocks is shown in Figure A.2. The overall observation that short-term and long-term rates are disconnected is not new: Backus and Wright (2007) talked about a "conundrum," where from 2004 to 2006 the FOMC raised the target federal funds rate by 4.25 percentage points, yet long-maturity yields and forward rates fell. Second, press conferences are important events where substantial monetary policy information gets communicated to investors leading to observable variation in asset prices. Press conferences are an integral part of the learning process of investors around monetary policy events.

3.2   Methodology

To investigate how monetary policy surprises transmit into the mortgage markets, we estimate the response of mortgage rates to the high-frequency policy news using panel local projections á la Jordà (2005). Unlike other asset classes, where it's possible to compute price changes over short-time windows and then regress those changes onto the monetary policy news, each mortgage is issued only once. So, we run the following regression:

$$
\begin{aligned}
m_{ijcf,h} = \alpha_{jcf} + \delta X_i &+ \sum_{j=-5}^{-2} \gamma_{Nj} \mathbb{1}_{h=j} \\
&+ \sum_{j=0}^{28} \gamma_{Nj} \mathbb{1}_{h=j} + \sum_{j=-5}^{-2} \gamma_{Pj} \mathbb{D} \, \mathbb{1}_{h=j} + \sum_{j=0}^{28} \gamma_{Pj} \mathbb{D} \mathbb{1}_{h=j} + \sum_{j=-5}^{-2} \beta_{Nj} \mathbb{1}_{h=j} \Delta s_f \quad (4) \\
&+ \sum_{j=0}^{28} \beta_{Nj} \mathbb{1}_{h=j} \Delta s_f + \sum_{j=-5}^{-2} \beta_{Pj} \Delta s_f \, \mathbb{1}_{h=j} \, \mathbb{D} \, + \sum_{j=0}^{28} \beta_{Pj} \Delta s_f \, \mathbb{1}_{h=j} \, \mathbb{D} + \epsilon_{ijcf,h}
\end{aligned}
$$

where $m_{ijcf,h}$ is the the mortgage rate for borrower i in metropolitan area code c for a 30-year mortgage issued by lender j on date h around FOMC event f. X_i is a set of borrower characteristics, including the level and the square of the FICO score at origination and

17

of the loan-to-value (LTV) ratio at origination, whether the mortgage was GSE-eligible, insured at origination, or whether it was prime or subprime. Δs_f is the absolute value of the change in par swap rate for a 10-year tenor around the FOMC announcement f, $\mathbb{1}_{h=j}$ is a dummy variable taking value 1 if h is equal to j and zero otherwise, $\mathbb{D}$ is a dummy variable taking a value of 1 if Δs_f is positive and zero otherwise, and ϵ is the error term. All regressions control for lender-by-metropolitan-area-by-FOMC-event fixed effects. Standard errors are clustered at the MSA by origination year-month level.

To investigate how monetary policy surprises instead transmit into the corporate debt market (corporate bond yields and credit spreads), we can again use panel local projections á la Jordà (2005). However, now we can exploit the fact that we observe the same asset before and after an FOMC event. Therefore, we modified the specification to include security-by-FOMC-event fixed effects:

$$
\begin{aligned}
y_{if,h} = \alpha_{if} &+ \sum_{j=-5}^{-2} \gamma_{Nj}\mathbb{1}_{h=j} \ + \sum_{j=0}^{28} \gamma_{Nj}\mathbb{1}_{h=j} + \sum_{j=-5}^{-2} \gamma_{Pj}\mathbb{D} \ \mathbb{1}_{h=j} + \sum_{j=0}^{28} \gamma_{Pj}\mathbb{D}\mathbb{1}_{h=j} \\
&+ \sum_{j=-5}^{-2} \beta_{Nj}\mathbb{1}_{h=j}\Delta s_f \\
&+ \sum_{j=0}^{28} \beta_{Nj}\mathbb{1}_{h=j}\Delta s_f + \sum_{j=-5}^{-2} \beta_{Pj}\Delta s_f \ \mathbb{1}_{h=j} \ \mathbb{D} \ + \sum_{j=0}^{28} \beta_{Pj}\Delta s_f \ \mathbb{1}_{h=j} \ \mathbb{D} + \epsilon_{if,h}
\end{aligned}
\tag{5}
$$

where, depending on the analysis, $y_{if,h}$ is the yield on bond i or the par spread on the CDS i on date h around FOMC event f. As before, Δs_f is the absolute value of the change in par swap rate for 10-year tenor around the FOMC announcement f, $\mathbb{1}_{h=j}$ is a dummy variable taking value 1 if h is equal to j and zero otherwise, $\mathbb{D}$ is a dummy variable taking a value of 1 if Δs_f is positive and zero otherwise, and ϵ is the error term. Standard errors are clustered at the transaction year-month level.

4 Results

4.1 The impact of monetary policy on mortgage rates

Results. Our benchmark measure of interest rate shocks reflecting monetary policy news surrounding FOMC events uses the daily change in 10-year swap rates on FOMC

days.[12] We estimate the response of mortgage rates to monetary policy news in the 28 days following an FOMC announcement. Results are reported in Figure 6. We analyse separately the whole sample including all events from 2000 to 2019 and a sample starting only in 2010. In the sample from 2000, the average response to positive shocks is statistically indistinguishable from the response to negative shocks. So, we repeat the analysis focusing on the more recent subsample, where we observe an average response to positive monetary policy shocks that is larger than the response to negative shocks by about 54 basis points per 100 basis points of the shock with a t-statistics of 3.65.

Our second measure of interest rate shocks uses the intra-daily change in 10-year swap rates on FOMC days. Again, we estimate the response of mortgage rates to monetary policy news in the 28 days following an FOMC announcement. Results are reported in Figure 7. Note that the *intra*-daily change in 10-year swap rates on FOMC days is smaller in magnitude compared to the inter-daily change. This different magnitude explains why the estimated coefficient in the regression is larger in the second case compared to the first. Focusing on the sample after 2010 only, we observe an average response to positive monetary policy shocks that is larger than the response to negative shocks by about 124 basis points per 100 basis points of the shock with a t-statistic of 3.[13]

---

[12]For the days for which we can also compute the higher-frequency shock on 10-year swap rates, Figure A.6 shows the scatterplot of daily changes against intraday changes. The correlation between the 2 shocks is 0.81.

[13]Figure A.20 shows the results with respect to Nakamura and Steinsson (2018) shocks. Surprises in rates up until 1 year do not necessarily transmit to the mortgage market.
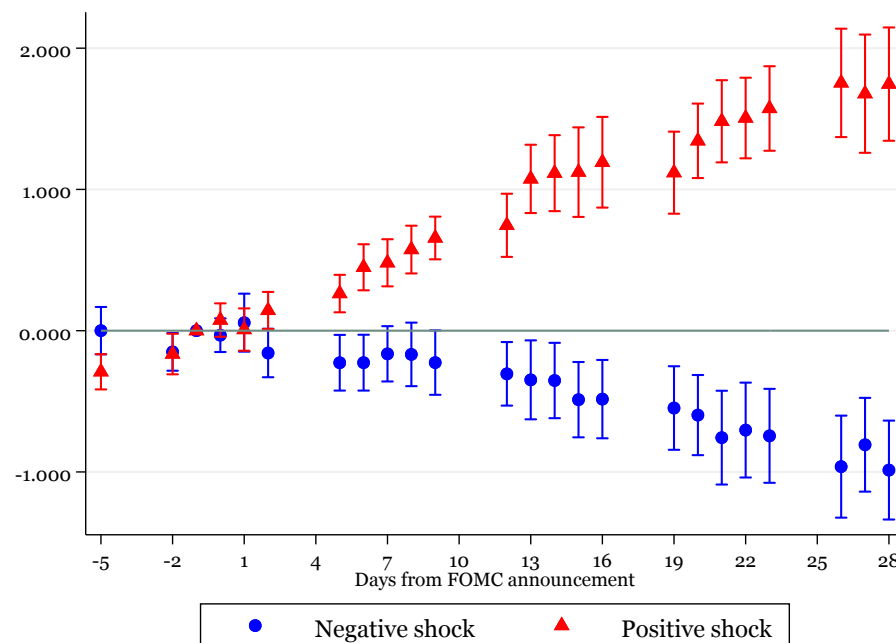
19

Fig. 6. **Response of mortgage interest rates to daily FOMC shocks**

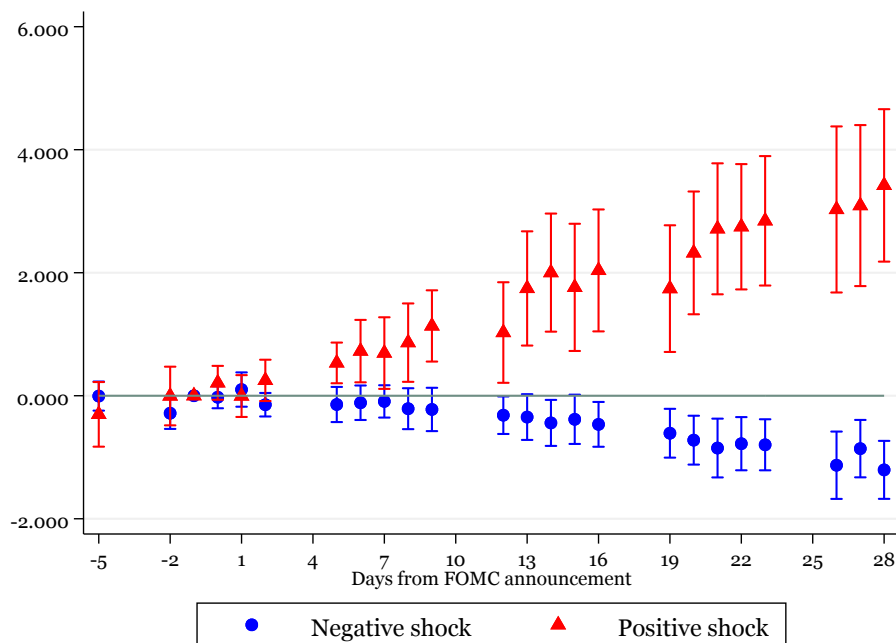Panel A: January 2000 to December 2019

Panel B: January 2010 to December 2019

Notes: The figure reports the slope coefficient β_h and 95%-confidence interval from Equation 4. The regression controls for lender-by-metropolitan-area-by-FOMC-event fixed effects as well as borrower's characteristics. Standard errors are clustered at MSA $\times$ origination year-month level. The sample consists of all the conventional loans (not originated under a government program) where the borrower's stated purpose is to purchase a property and the property type is either a condominium or single-family residence. Data span from January 2000 to December 2019.

20

Fig. 7. **Response of mortgage interest rates to high-frequency FOMC surprises.**

January 2010 to December 2019

Notes: The figure reports the slope coefficient β_h and 95%-confidence interval from Equation 4. The regression controls for lender-by-metropolitan-area-by-FOMC-event fixed effects as well as borrower's characteristics. Standard errors are clustered at MSA × origination year-month level. The sample consists of all the conventional loans (not originated under a government program) where the borrower's stated purpose is to purchase a property and the property type is either a condominium or single-family residence. Data span from January 2000 to December 2019.

Potential explanations Can the asymmetric response to monetary policy news be a consequence of market power in mortgage lending? A large literature arguing for the presence of market power in banks' lending markets ((Scharfstein and Sunderam, 2016; Crawford, Pavanini, and Schivardi, 2018, among others)) suggests that it could. However, We find that the response of mortgage rates to monetary policy does not depend on mortgage market concentration at the county or zip3 level. Further, our results are robust to alternative measures of concentration. In particular, we compute four proxies of market power: a) Herfindahl-Hirschman index based on all loans approved in a FIPS county; b) the market share of the top 4 lenders in a county; c) a county-level measure of excess demand, i.e., the number of loans approved plus the number of loans rejected over the number of loans approved; d) the component of interest rate above and beyond what can be explained by borrower's and loan's characteristics aggregated at the zip3 level.

21

For each market power proxy, we sort geographical areas into quintiles by year-quarter creating a vector of dummies Q whose jth observation takes a value of 1 if the loan was originated in an MSA belonging to the jth quintile and 0 otherwise. We interact Q with the dummies representing the days in the event window surrounding an FOMC announcement and estimate the following equation:

$$
\begin{aligned}
m_{ijcf,h} = \alpha_{jcf} + \delta X_i + \sum_{j=-5}^{-2} \gamma_{Nj} \mathbb{1}_{h=j} Q \\
+ \sum_{j=0}^{28} \gamma_{Nj} \mathbb{1}_{h=j} Q + \sum_{j=-5}^{-2} \gamma_{Pj} \mathbb{D} \ \mathbb{1}_{h=j} Q + \sum_{j=0}^{28} \gamma_{Pj} \mathbb{D} \mathbb{1}_{h=j} Q + \sum_{j=-5}^{-2} \beta_{Nj} \mathbb{1}_{h=j} \Delta s_f Q \\
+ \sum_{j=0}^{28} \beta_{Nj} \mathbb{1}_{h=j} \Delta s_f Q + \sum_{j=-5}^{-2} \beta_{Pj} \Delta s_f \ \mathbb{1}_{h=j} \ \mathbb{D} Q + \sum_{j=0}^{28} \beta_{Pj} \Delta s_f \ \mathbb{1}_{h=j} \ \mathbb{D} Q + \epsilon_{ijcf,h}.
\end{aligned}
$$

(6)

This allows us to compute and compare the monetary policy pass-through in areas of low market power against the pass-through in areas of high market power.

1. Herfindahl-Hirschman index. The Home Mortgage Disclosure Act (HMDA) mandates that the vast majority of mortgage lenders in the United States furnish information to regulatory bodies regarding the loan, property, and borrower attributes of every mortgage application. Among the data that must be reported are the specifics of the loan including loan size, type, and action taken. Additionally, borrower characteristics such as income, race, ethnicity, and gender, as well as property characteristics including property type, occupancy status, state, county, and census tract, must also be reported.

We focus on all loans originated (i.e., the variable *action taken* equals 1) and link them to the parent company using the HMDA panel files by year. We sum all loans by parent company and county FIPS and compute the HHI at the county level for each year. Panel A of Figure A.9 reports the histogram of county-level HHI in our sample, whereas panel B shows the spatial variation of average HHI over time.

Each year, we sort all counties by their HHI, creating the vector of quintile dummies Q, and then estimate (6). In either sample (i.e., in the whole sample from 2000 or starting from 2010) we do not observe a significantly different response to monetary policy news

in high-HHI areas relative to low-HHI areas (5th vs 1st quintile). Figure A.10 shows the results for the whole sample.

2. Share of the top 4 lenders by county. We follow Scharfstein and Sunderam (2016) and compute from HMDA data the market share of the top 4 lenders in a county as a measure of concentration. For each county-year we sort all lenders based on the values of loans originated and compute the ratio between the total amount of mortgages originated by the top 4 lenders and the total amount of mortgages originated by all lenders in that geographical area.

Each year, we sort all counties based on this measure of concentration, creating the vector of quintile dummies Q. We then estimate the specification in (6). In either sample (i.e., in the whole sample from 2000 or starting from 2010) we do not observe a significantly different response to monetary policy news in high-concentration areas relative to low-concentration areas (5th vs 1st quintile). Figure A.11 shows the results for the whole sample.

3. Excess demand. Again from HMDA data, we sum all loans originated and the application approved but not accepted by the borrower as loans accepted (*action taken* equal to 1 and 2). We compare them with the sum of applications denied by financial institutions and files closed for incompleteness (*action taken* equal to 3 and 5). We define excess demand as the sum of applications accepted and denied over the applications accepted in a given FIPS county and year. Each year, we sort all counties by the values of excess demand, creating the vector of quintile dummies Q, and then estimate the specification in (6). In either sample (i.e., in the whole sample from 2000 or starting from 2010) we do not observe a significantly different response to monetary policy news in high-excess-demand areas relative to low-excess-demand areas (5th vs 1st quintile). Results for the whole sample are shown in Figure A.12.

4. Interest rate residual by zip3. To eliminate the influence of borrower and loan characteristics on mortgage rates, we follow Hurst, Keys, Seru, and Vavra (2016) and use loan-level microdata from the Federal Home Loan Mortgage Corporation (Freddie

Mac) to estimate the following equation:

$$r_{ikt} = \alpha_0 + \alpha_1 X_{it} + \alpha_2 D_t + \alpha_3 D_t \cdot X_{it} + \eta_{ikt}, \tag{7}$$

where r_{ikt} is the mortgage rate for borrower i in MSA k in year-quarter t. X_{it} is a set of control variables for borrower i in period t including the level and square of the FICO score and LTV ratio. D_t is a vector of time dummies representing the quarter of origination. The residuals obtained from these equations represent the spatially adjusted mortgage rates for a borrower in an MSA for a given quarter.

We want to compute a measure of how expensive is the average loan in an area after adjusting for borrowers', loans' characteristics, and the time of origination. Using the residuals from the previous regression η_{ikt}, we compute

$$R_{kt} = \frac{1}{N_{kt}} \sum_{i=1}^{N_{kt}} \eta_{ikt}, \tag{8}$$

for an MSA k and year-quarter t. R_{kt} represents the average difference between the observed mortgage rate for loans made in that MSA and the mortgage rate predicted by the borrower and loan characteristics and time fixed effects. N_{kt} is the number of loans originated in MSA k at time t. Figure A.13 shows the spatial variation of R_{kt} averaged over time.

Each quarter we then sort MSAs into 5 quintiles based on the value of R_{kt}, and create the vector of dummy variables Q so as to estimate (6). Results are in Figure A.14. Again, in either sample (i.e., in the whole sample from 2000 or starting from 2010) we do not observe a significantly different response to monetary policy news in high-interest-rate-residual areas relative to low-interest-rate-residual areas (5th vs 1st quintile).

General movements in swap rates. We now explore an alternative explanation. Can the response of mortgage rates be justified by a general movement of interest rates, e.g., in the level of 10-year swap rates?

Figure 8 shows the estimated response of 10-year swap rates to monetary policy surprises (daily changes in the 10-year swap rate in FOMC days) and corresponding 95%

Fig. 8. **Response of 10-year swap rates to daily FOMC surprises.**
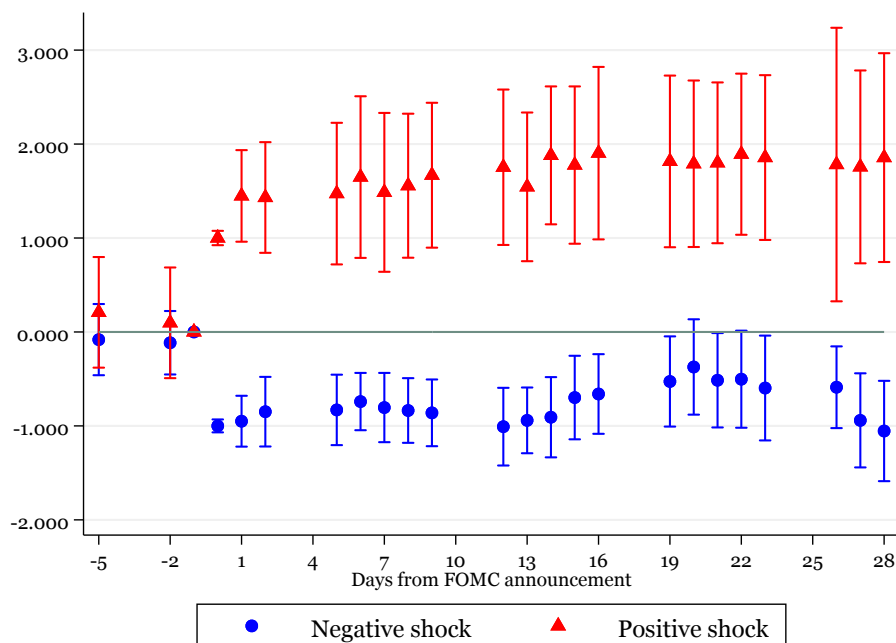
Notes: The figure shows the estimated response of 10-year swap rates to monetary policy surprises in the 10-year swap rates and corresponding 95% confidence interval from (9). Standard errors are clustered at the year-month level. Data span January 2000 to December 2019.

confidence interval from the following regression:

$$
\begin{aligned}
c_{hf} = a_f &+ \sum_{j=-5}^{-2} \gamma_{Nj} \mathbb{1}_{h=j} \ + \sum_{j=0}^{28} \gamma_{Nj} \mathbb{1}_{h=j} + \sum_{j=-5}^{-2} \gamma_{Pj} \mathbb{D} \ \mathbb{1}_{h=j} \\
&+ \sum_{j=0}^{28} \gamma_{Pj} \mathbb{D} \mathbb{1}_{h=j} + \sum_{j=-5}^{-2} \beta_{Nj} \mathbb{1}_{h=j} \Delta s_f \\
&+ \sum_{j=0}^{28} \beta_{Nj} \mathbb{1}_{h=j} \Delta s_f + \sum_{j=-5}^{-2} \beta_{Pj} \Delta s_f \ \mathbb{1}_{h=j} \ \mathbb{D} \ + \sum_{j=0}^{28} \beta_{Pj} \Delta s_f \ \mathbb{1}_{h=j} \ \mathbb{D} + \epsilon_{hf},
\end{aligned}
\tag{9}
$$

where h represents the number of days from the FOMC announcement day f, c is the 10-year swap rate with annual payments for the fixed leg against 3-month LIBOR, Δs_f is the absolute value of the change in par swap rate for 10-year tenor around the FOMC announcement f, $\mathbb{1}_{h=j}$ is a dummy variable taking value 1 if h is equal to j and zero otherwise, $\mathbb{D}$ is a dummy variable taking a value of 1 if Δs_f is positive and zero otherwise, and ϵ is the error term. Standard errors are clustered at the year-month level. Data

25

span January 2000 to December 2019. The response is 1 (by construction) on FOMC days, but, more importantly, it becomes larger than 1 immediately on the day after the FOMC announcement and then stabilizes: the response of swap rates is larger than the response of mortgage rates in the same period (panel A Figure 6).[14]

Now, to assess whether the changes in mortgage rates observed in the previous section are just a response to a change in swap rates, we run the following 2-step procedure. First, we compute the fitted values of the 10-year swap rates $\hat{c}_{hf}$ from (9). Second, we use these fitted values $\hat{c}_{hf}$ as an additional control in (4): we want to study the variation in mortgage rates above and beyond what can be explained by the change in the level of the bank's cost of funding alone which naturally follows monetary policy surprises. Results are shown in Figure 9. The swap rates respond to monetary policy news much faster than the corresponding mortgage. This asynchronicity is what causes the coefficient to switch. Regardless, in the 28-day period, there is no evidence of a movement in mortgage rates above and beyond the movements of swaps.

---

[14]As a comparison we estimate the response of 10-year nominal government par rates and 10-year real government par rates using (9) and show it in Figure A.7 and Figure A.8, respectively.

Fig. 9. **Response of mortgage interest rates to daily FOMC shocks controlling for changes in USSA10**

Panel A: January 2000 to December 2019
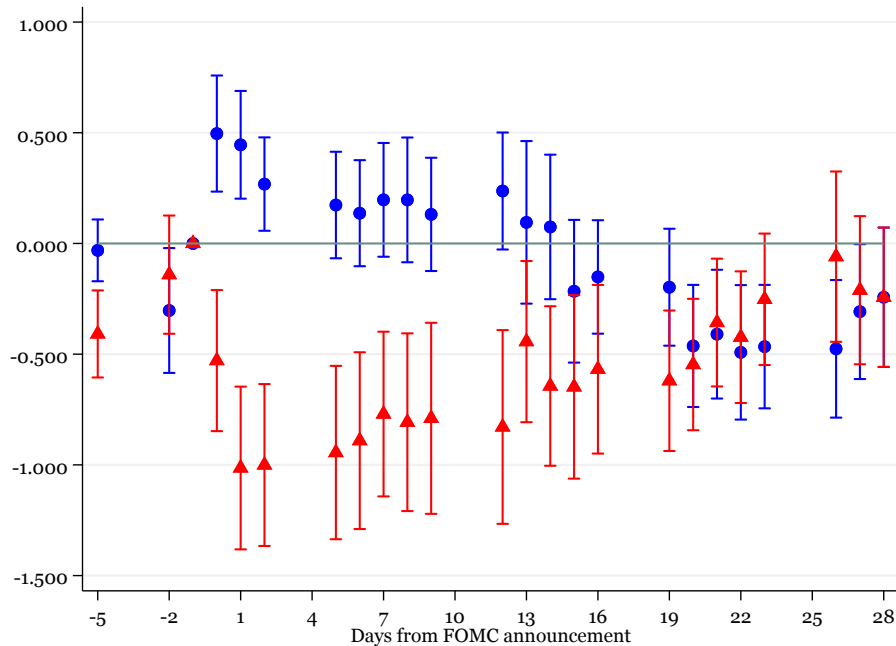
Panel B: January 2010 to December 2019
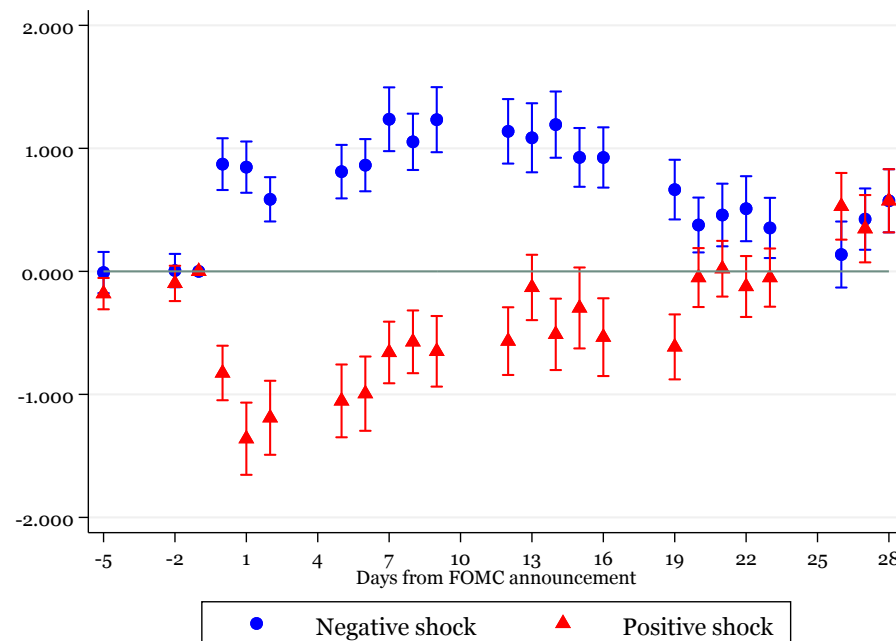
Notes: We estimate 4 adding as a control the predicted swap rate from 9. The figure reports the slope coefficient β_h and 95%-confidence interval from the estimation. The regression controls for lender-by-metropolitan-area-by-FOMC-event fixed effects as well as borrower's characteristics. Standard errors are clustered at MSA × origination year-month level. Data span from January 2000 to December 2019.

27

Table 2. Sensitivity of different loan products to monetary policy

Product name	Negative shock		Positive shock	
	0-14 days	15-23 days	0-14 days	15-23 days
1 Year ARM @ 175K - Rate	0.17	-0.493***	0.037	-0.07
	(0.256)	(0.148)	(0.405)	(0.219)
3 Year ARM @ 175K - Rate	-1.035***	-0.95***	0.506	0.062
	(0.25)	(0.114)	(0.347)	(0.157)
5 Year ARM @ 175K - Rate	-1.129***	-1.096***	0.442	0.616***
	(0.233)	(0.121)	(0.286)	(0.213)
7 Year ARM @ 175K - Rate	-1.121***	-1.394***	0.981	1.005***
	(0.268)	(0.197)	(0.697)	(0.345)
10 Yr Fxd Mtg @ 175K - Rate	-1.79***	-0.985***	2.053***	0.222
	(0.412)	(0.214)	(0.615)	(0.25)
15 Yr Fxd Mtg @ 175K - Rate	-1.105***	-0.958***	0.523***	0.57***
	(0.155)	(0.116)	(0.2)	(0.146)
20 Yr Fxd Mtg @ 175K - Rate	-1.284***	-1.175***	1.886***	1.036***
	(0.316)	(0.161)	(0.528)	(0.341)
30 Yr Fxd Mtg @ 175K - Rate	-1.262***	-1.039***	0.783***	0.776***
	(0.163)	(0.098)	(0.221)	(0.197)

4.2 The impact of monetary policy on other mortgage products – Ratewatch

One may wonder to what extent endogenous self-selection of borrowers after monetary policy announcements influences our results. In particular, it could be the case that riskier borrowers borrow more after an increase in rates (moral hazard). To examine whether our results are driven by such a potential endogenous self-selection mechanism, we use rates from Ratewatch. RateWatch surveys bank branches throughout the US to collect data on a broad range of consumer loan products. Their data go back to 2001 and contain details such as the date the survey was conducted, the particulars of various loan agreements (including interest rates), and the branch responsible for determining the interest rate. We group by product category rates for ideal mortgages to the "best" borrowers, i.e., those with exceptional FICO scores, for a particular constant loan volume of $175K with 20% down-payment.[15] We consider both adjustable-rate mortgages (ARM)

---

[15]The credit score cutoff is for most banks 740 or higher, e.g., Bank of America.

and fixed-rate mortgages. In our sample, ARMs all have a 30-year tenor with an initial rate fixed for a certain number of years and a variable rate for the remaining years.

We estimate (4) while including in all specifications account-number-by-MSA-by-FOMC-event fixed effect. However, rather than doing the analysis at a daily frequency, because the survey is conducted monthly for each branch level, we group together observations in 3 separate windows. The first window includes dates between the day of the FOMC announcement and 14 days after the announcement. The second window includes dates between 15 and 23 days after the FOMC announcement. Finally, the third window is the pre-period that goes from 10 days before the FOMC announcement to 1 day before. Table 2 reports our estimates. In all cases, with the exception of the mortgages of the shortest maturities, we find that quoted rates respond to changes in swap rates on FOMC days. Our findings provide evidence against the hypothesis that endogenous self-selection of borrowers after monetary policy announcements, with riskier borrowers desiring to borrow more after an increase in rates, are an important driver of our results.

4.3   The impact of monetary policy on corporate funding costs

In this section, we study the role of financial frictions and firm heterogeneity in the transmission of monetary policy surprises to the cost of firms' external financing: corporate bonds. Evaluating the response of corporate bonds separately from bank loans is important because the two assets are not perfect substitutes. Among others, the main differences are: a) corporate bonds are less flexible, and their terms are harder to renegotiate than bank loans; b) bank loans are extended by highly-leveraged intermediaries with significant liquidity mismatches; and c) more generally, bonds and loans have different contractual features. It's therefore unclear from the previous results how bond pricing contributes to transmitting aggregate shocks such as monetary policy. We use secondary market prices on corporate bonds and CDS to shed light on this question.

We first estimate (5) using corporate bond yields from TRACE. The equation includes bond-cusip-by-FOMC fixed effects to control for all unobserved characteristics of the bond in the 1-month window surrounding an FOMC announcement. Unlike

29

mortgages, secondary market yields respond immediately to monetary policy shocks. More importantly, the asymmetry is negative, with a larger response of corporate bond yields to negative shocks. The response to positive shocks is 1-to-1 and stabilizes already after 4 days. On the other hand, the response to negative shocks is 1-to-1 only in the first week after the announcement and then slowly converges to a 2-to-1 response. Yet, the difference between the absolute magnitude of positive and negative responses is statistically insignificant with a t-statistic of -1.14. Focusing on the second half of the sample, we find similar responses to negative monetary policy surprises and a response to positive shocks which is of the same magnitude as the overall sample (confidence intervals are larger).[16]

We now test whether the response depends on the bond's credit ratings. Using the complete cusip, (issue and issuer cusip), issue name, issuer id, maturity, and offering date, we merge the universe of bonds in Mergent/FISD with the rating file that Mergent provides. We separate all bonds for which we have a rating into investment-grade and speculative-grade bonds using the most recent credit rating issued before the transaction date. Figure A.15 shows the results. Speculative grade bonds appear to respond more strongly to negative news in long-term rates with their yields dropping by a larger amount. On the other hand, the response to positive shocks is weaker for speculative-grade bonds with yields increasing only in the few days after the FOMC announcements and then becoming statistically indistinguishable from 0.

Finally, we test the response of Credit Default Swaps (CDS) to monetary policy surprises. Krishnamurthy and Vissing-Jorgensen (2011) have shown an effect of quantitative easing in lowering the default risk of companies as measured by CDS spreads. We study this relation for all monetary policy events from 2000, and (as before) separate positive from negative monetary policy news for long-term rates. In the overall sample, we find that a drop in rates has been accompanied by a drop in credit risk, while we do not observe an increase in CDS spreads after a positive monetary policy surprise (Figure 11). Panel B Figure 11 shows that the negative response is more pronounced for the Credit Default Swaps of B-rated non-financial firms. More generally, we consistently observe a

---

[16]Figure A.21 shows the results with respect to Nakamura and Steinsson (2018) shock. Surprises in rates up until 1 year do not necessarily transmit to long-term bond yields.

stronger response when we pass from a higher to a lower credit rating grade. Nonetheless, and perhaps surprisingly inconsistent with the results on bonds, we only observe a drop in CDS spread in the period of the financial crisis consistent with the evidence by Krishnamurthy and Vissing-Jorgensen (2011) on the effects of unconventional monetary policy on firms' credit default spreads. The response of CDS spreads to either positive or negative shocks in the more recent sample from 2010 is statistically indistinguishable from 0.

Overall, we show that the corporate bonds of firms with low ratings were the most responsive to monetary shocks and that most of the effect went through a change in the credit risk of these firms. These results are complementary to Ottonello and Winberry (2020), who document that firms with low default risk invest more in response to monetary shocks. They highlight that highly-rated firms invest more in response to monetary policy surprises because they face a flatter marginal cost curve for financing investment, which is indeed consistent with what we observe.

5 Decomposing long-term monetary policy news: Expected future rates vs term premia

In this section, we explore the drivers of our results, namely, how the factors underlying our shocks in 10-year swap rates get transmitted to long-term mortgage and corporate bond markets. Call i_t the 1-year zero rate between year t and $t+1$ and i_t^m the zero rate at time t for m years, we can decompose i_t^m as

$$i_t^m = \underbrace{E_t \frac{1}{m} \left\{ \sum_{j=0}^{m-1} i_{t+j} \right\}}_{EI} + \phi_t^m, \tag{10}$$

where EI stands for expected future interest rates, and ϕ_t^m is the annualized term premium. The term premium compensates investors in long-term bonds for interest rate risk. We will use zero rates computed from government bonds, and so the corresponding measures of EI and ϕ.

Table 3. Estimates from regressing change in 10-year swap rates on expected short-term interest rates and term premia

	Δs
Term premium	0.789***
	(0.083)
Expected short-term interest rates	1.195***
	(0.074)
Constant	-0.005*
	(0.002)
Observations	160
Adjusted R^2	0.841

Notes: This table presents the regression coefficient estimates from (11). Standard errors are in parentheses. ***, **, * denote significance at the 1%, 5%, and 10% level, respectively. The sample is from January 2000 to December 2019.

To shed light on whether our results stem from variation in a) expected future interest rates, b) term premia, or c) specific features of 10-year swaps relative to 10-year treasury bond zero yields, we use the decomposition proposed by Adrian et al. (2013). We first regress the daily change in the 10-year swap rate on FOMC days onto the daily change in expected future interest rates and term premia, estimated for 10-year Treasury zero coupon yields by Adrian et al. (2013):

$$\Delta s_t = \alpha + \beta_{\text{EI}} \Delta \text{EI}_t + \beta_\phi \Delta \phi_t^{10} + \eta_t, \tag{11}$$

where t is the FOMC event and η is the regression residual. Residuals, which are orthogonal to the two regressors by construction, represent all other factors affecting swaps but not captured by EI or ϕ. The residual term can capture the fact that 10-year swap rates are par yields whereas the two factors are constructed from zero-coupon yields, but it can also capture frictions specific to the banking sector. Table 3 reports the estimates from (11). The two factors explain about 84% of the variation in 10-year swaps on FOMC days. The decomposition of the time series of changes in 10-year swap rates in expected short-term rates and term premia is plotted in Figure A.22.

We then extend (4) and (5) to include together ΔEI, $\Delta \phi$ and η. Results are in

Figure A.18 and Figure A.19. Mortgage rates appear to respond consistently to all three components, although the most significant effects can probably be observed for the residual component η. Instead, corporate bond yields respond mostly to future short-term rates. We do not see any significant relation between corporate bond yields and variation in term premia, whereas the residual component appears to account for drops in yields mostly.

6   Implications for bank net worth

Our research demonstrates that monetary policy affects long-term rates differently from short-term rates and exploits this observation to assess the response of long-term mortgage rates and bond yields to the monetary policy news that *really matter* for their pricing. The same insight also has important implications for banks' net worth in light of the conventional narrative suggesting that banks borrow funds on a short-term basis and lend them out to borrowers on a longer-term basis (maturity transformation).

Banks' profits are influenced by various interest rates rather than just one market interest rate. Different assets and liabilities on a bank's balance sheet have different degrees of liquidity, market and credit risk, and, most importantly, maturity, making it impossible to rely solely on a single market interest rate to evaluate a bank's exposure to interest rate changes (Hancock, 1985). While previous studies have used a single short-term interest rate to estimate banks' sensitivity to interest rates (Samuelson, 1945; Drechsler et al., 2021), we recommend distinguishing between short-term and long-term rates and considering both sets of rates when evaluating how a bank's wealth responds to monetary surprises.[17]

Figure 12 shows the results of regressing Fama-French 49 industry portfolios on the changes in 10-year swaps on FOMC days (controlling for Kuttner (2001) federal-funds

---

[17]Both short-term and long-term monetary policy surprises have two distinct effects on bank stock prices: discounting and cash-flow effects. When rates increase, future dividends are discounted more heavily, leading to declining market values. An increase in the term premium leads to higher net interest margins for banks, while non-financial firms are unlikely to experience such an effect (Paul, 2023): most firms face increased interest expenses due to higher term premia, causing a decline in cash flows. As a result, bank stock returns tend to respond more positively than those of non-financial companies following an increase in the term premium via the cash-flow channel.

shocks). The equation used for the regression is:

$$R_{jt} = \alpha + \beta_j \Delta s_t + \gamma_j \Delta \mathrm{FF}_t + \epsilon_{jt}, \tag{12}$$

where R_{jt} us the daily return for industry j on FOMC day t, Δs is the change in 10-year swap rates in FOMC days, and $\Delta \mathrm{FF}$ is the federal-funds shock. Considering all dates in our analysis, we find that banks are positively exposed to an increase in long-term rates, although the coefficient is small in magnitude and statistically insignificant. However, the results are remarkably large in magnitude and significance when we exclude the three Quantitative Easing 1 (QE1) scheduled FOMC announcements from the sample. These three dates account for the two largest declines in 10-year swaps, but they also indicate substantial protection for the financial sector during times of banking distress (definitely, positive news for banks).[18] The banking industry shows the highest exposure to shocks in long-term rates, with a positive and significant coefficient of 7.91. This implies that bank stocks increase by 7.91% for every 1% positive shock to the 10-year swap rate. On the other hand, the exposure to short-term rates (fed funds shocks) is negative (-3.527), consistent with the estimate of Drechsler et al. (2021).

We now use the same approach employed in analyzing the bank industry portfolio to compute the exposure of publicly traded commercial banks to changes in interest rates. For all FOMC days excluding the QE1 events, we regress individual bank daily stock returns onto the change in the swap rates on the same days and the Kuttner (2001) fed fund futures shock. We model the individual bank's exposure to swap changes as a linear function of the bank's characteristics. In all specifications, we directly use as control the same characteristics and include bank-level fixed effects. Standard errors are clustered at the FOMC level. Table 4 reports the estimates for the following equation

$$R_{it} = \beta_{0i} + \beta_{FFi} \times \Delta FF + \beta_x \times X_{it} + \beta_s \times \Delta s + \beta_{sx} \times X_{it} \times \Delta s + \epsilon_{it}. \tag{13}$$

both in the case of no weight (columns 1, 3, and 5) and for a WLS using the natural

---

[18]Figure A.24 shows the scatterplot of the bank's daily returns on FOMC days against the daily change of 10-year swap rates for all dates, including QE1 events. The two top-left points refer both to QE1 scheduled FOMC announcements.

logarithm of the bank's assets as weight (columns 2, 4, and 6). We confirm the positive relationship between changes in swap rates and bank stock returns when we do not specify the exposure as a function of banks' characteristics. Once we include the fraction of loans repricing in the next year as a determinant of the bank's exposure to rates, we see that banks with a larger fraction of loans repricing in the short term benefit the most from increased rates. Results with respect to the fraction of government securities with a remaining maturity or next repricing date of 1 year or less are qualitatively similar but statistically insignificant. When we also control for the bank's equity ratio we see that banks with a higher equity ratio benefit more from increases in long-term interest rates.

Our findings have significant implications in the context of existing research that highlights how monetary policy surprises can impact the real economy through their effects on banks' net worth (Gertler and Kiyotaki, 2010; He and Krishnamurthy, 2013; Brunnermeier and Sannikov, 2014; Ottonello and Song, 2022). Specifically, our results shed light on the positive impact of news about long-term rates on the banking sector and banks' shareholders. In contrast to changes in short-term rates that do not always translate into equivalent changes in banks' funding costs, particularly when banks have significant market power in deposits markets (Hannan and Berger, 1991; Neumark and Sharpe, 1992; Drechsler et al., 2017), we have documented monetary policy shocks affecting the long end of the term structure and banks' assets.

Our results may be consistent with an intermediary asset pricing hypothesis (especially the results decomposing changes in swap rates) but would reject the following standard intermediary-based narrative, which has been a core argument for an extensive literature in finance and economics. Imagine intermediaries being constrained agents in the business of maturity transformation. Higher long-term interest rates raise equity valuation for banks, so it's a positive net worth shock. The shock to net-worth increases intermediaries risk-bearing capacity and results in lower borrowing costs for firms (Siriwardane, 2019). The heterogeneous (high and low leverage firms) bond response may be driven by Value-at-Risk constraints (Adrian and Shin, 2013). Yet, this story implies a response exactly opposite to what we have documented in Section 4. This is probably because the intermediary frictions can explain only a small fraction of the variation in mortgages and bonds (e.g., as Table A.4 shows, lender-date specific factors explain only a small fraction

35

of the variation in mortgage rates). As we document in Section 5, interbank frictions are a significant predictor of variation in corporate and household liability rates, but at least on FOMC days, most of the variation is a monetary policy news and not a net-worth shock.

7   Conclusion

Much of the academic and practitioner literature implicitly assumes that the Federal Reserve's monetary policy impact is limited to its short-term policy rate. This historical perspective has prompted academic researchers to use the changes in expected short-term interest rates computed in a narrow time window surrounding FOMC announcements as a proxy for monetary policy shocks. However, in the more recent period, with short-term interest rates close to zero, the Fed had limited possibilities to surprise the market by changing the target on the Fed funds rate, and decided to rely more heavily on investors' expectation management through forward guidance and large-scale asset purchases.

In this paper, we examine the impact of monetary policy transmission on the long-term liabilities of households and firms, using high-frequency changes in 10-year swap rates surrounding FOMC announcements. We find that mortgage rates respond to monetary policy announcements in the three weeks after an FOMC announcement and, more importantly, symmetrically to positive and negative monetary policy shocks. On the other hand, in post-2010 data, interest rate hikes have had a greater impact on mortgage rates than cuts did. We explore several hypotheses underlying the stronger response to rate increases after 2010, and we reject hypotheses based on mortgage market concentration or bank local market power. We instead find that the asymmetric response in mortgage rates can be fully explained by an asymmetric response in 10-year swap rates in the days after the FOMC announcements. We conclude that understanding mortgage rates is tantamount to understanding the drivers of 10-year swap rate, which seems to best proxy for banks' funding costs.

When we look at the impact of monetary policy on corporate yields, we observe they respond symmetrically to positive and negative shocks, and we show a greater sensitivity for firms with lower credit ratings. Finally, we study the implications of our findings for

banks' net worth. The banking industry is positively exposed to shocks in long-term rates, with bank stocks increasing by 7.91% for every 1% positive surprise to the 10-year swap rate, outside of unconventional monetary policy interventions.

Fig. 10. **Response of corporate bond yields to daily FOMC shocks**

Panel A: January 2000 to December 2019

Panel B: January 2010 to December 2019

Notes: The figure reports the slope coefficient β_h and 95%-confidence interval from (5) for corporate bond yields. The regression controls for issue-cusip-by-FOMC-event fixed effects as well as borrower's characteristics. Standard errors are clustered at the year-month level. Data span from January 2000 to December 2019.

38
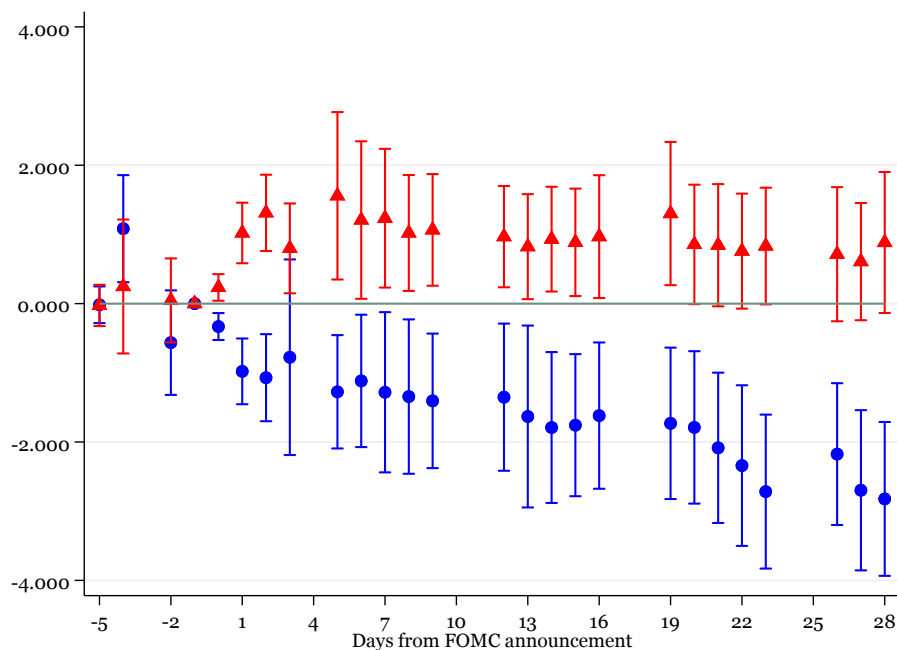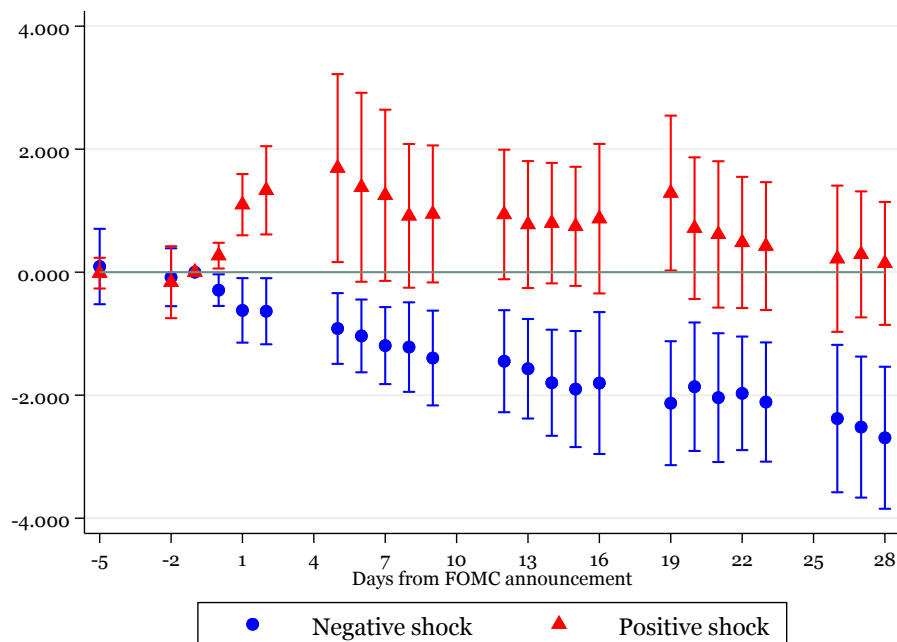
Fig. 11. **Response of 10-year CDS of nonfinancial firms to daily FOMC shocks**

Panel A: January 2000 to December 2019
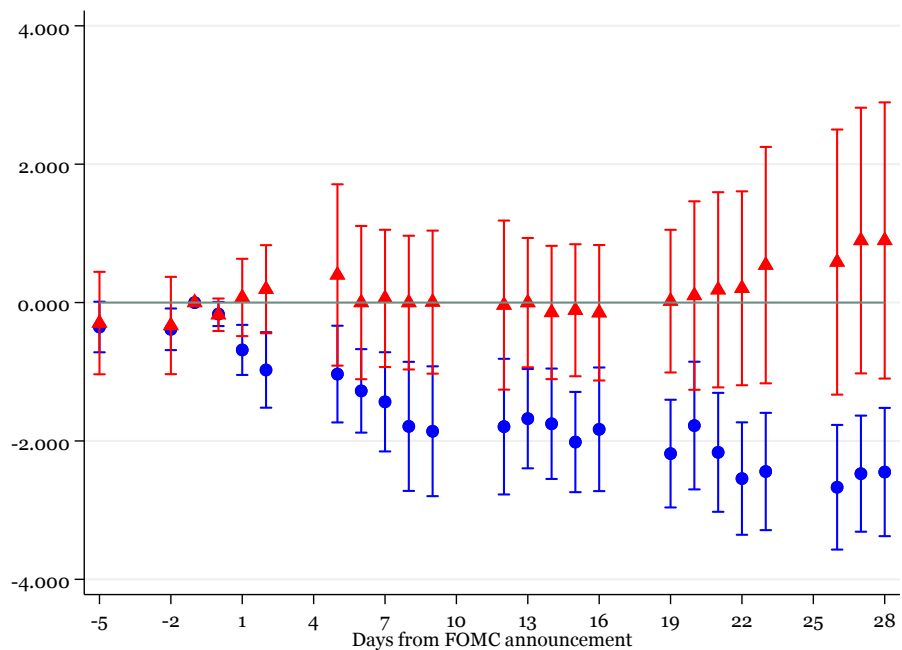
Panel A: January 2000 to December 2019 – B-rated CDS

● Negative shock ▲ Positive shock
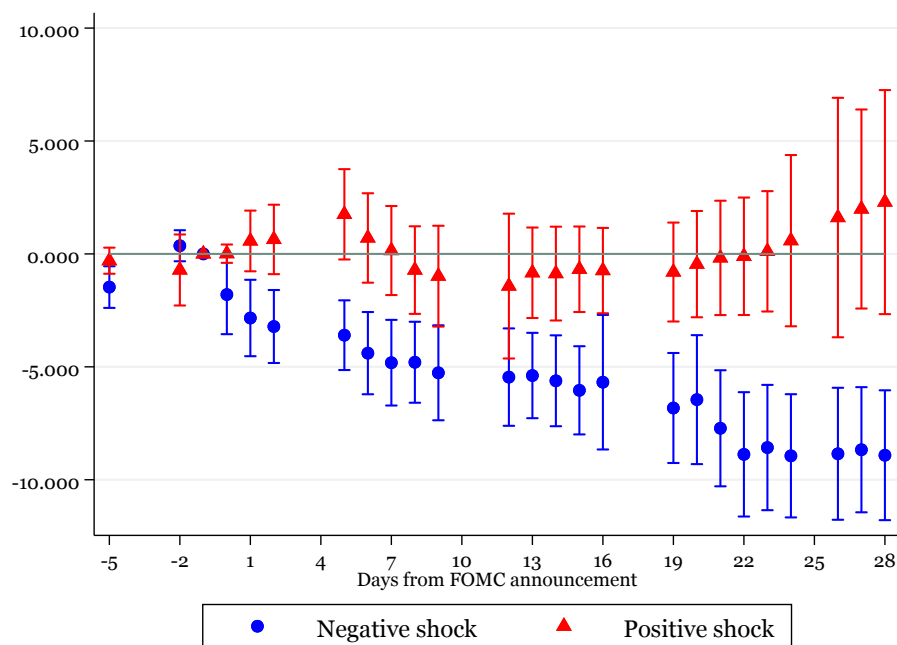
Notes: The figure reports the slope coefficient β_h and 95%-confidence interval from (5) for credit default swaps. The regression controls for underlying company-by-FOMC-event fixed effects. Standard errors are clustered at the year-month level. Data span from January 2000 to December 2019.

Fig. 12. **Banking industry stock returns on the 10-year interest swap rate changes.**

Panel A: β_j estimates from $R_{jt} = \alpha + \beta_j \Delta s_t + \gamma_j \Delta \text{FF}_t + \epsilon_{jt}$ – All dates
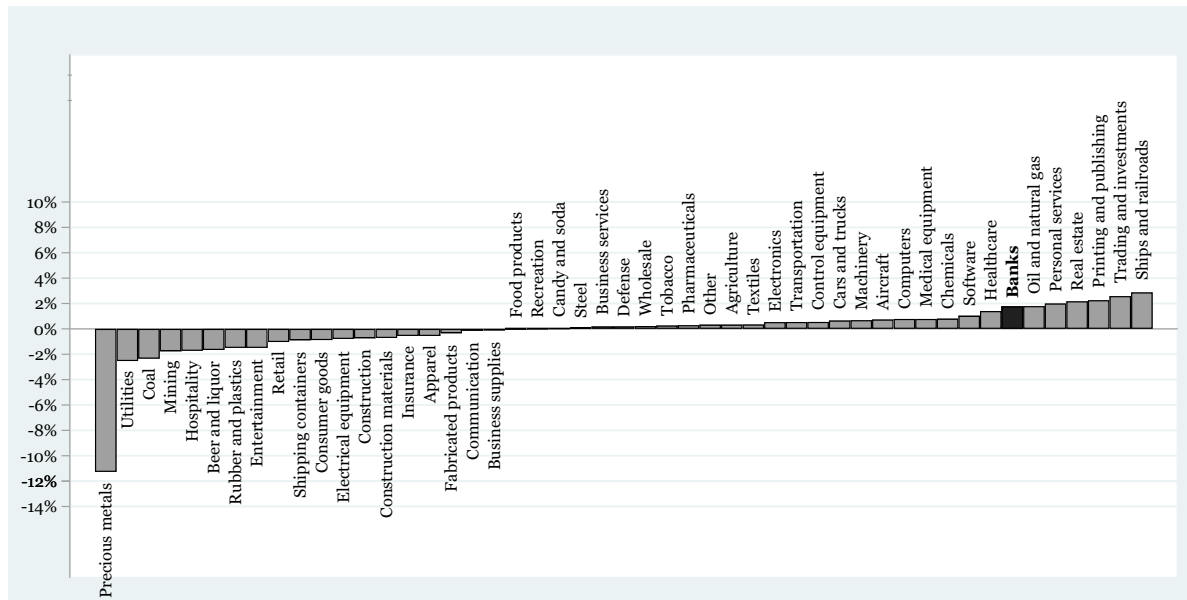
Panel B: β_j estimates from $R_{jt} = \alpha + \beta_j \Delta s_t + \gamma_j \Delta \text{FF}_t + \epsilon_{jt}$ – Excluding the 3 QE1 scheduled FOMC dates, i,e., December 16, 2008, January 28, 2009, and March 18, 2009.
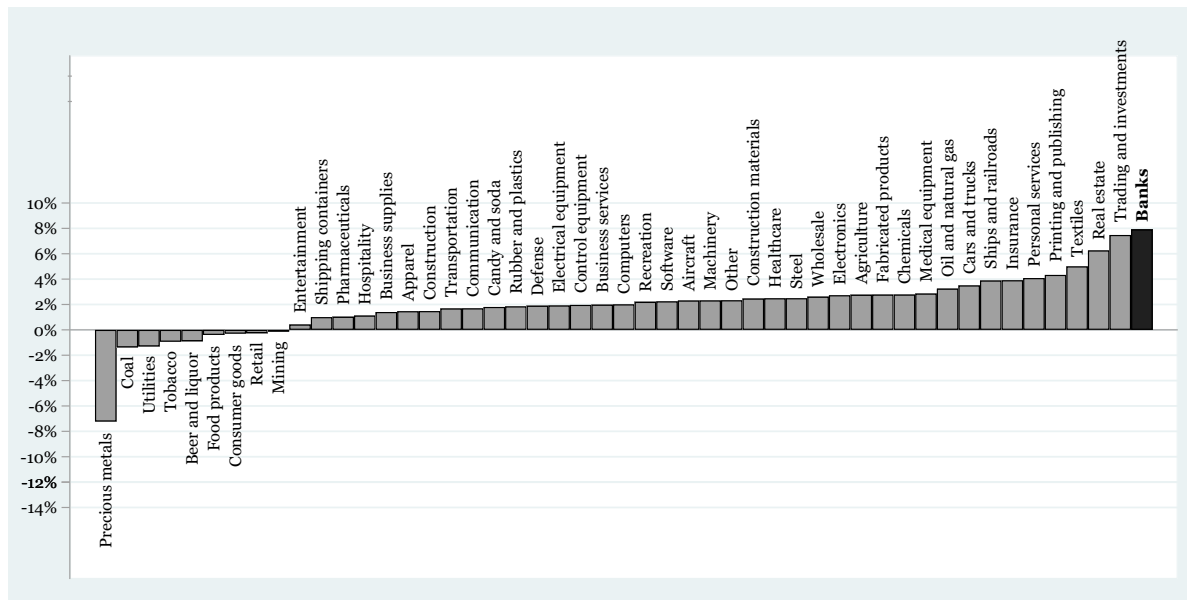
Notes: The figure shows the sensitivity of industry stock portfolios to FOMC rate changes. Industry data are the returns of the Fama-French 49 industry portfolios, downloaded from Ken French's website. The figure plots the coefficients from regressing daily industry returns on the daily changes in 10-year swap rates controlling for Kuttner (2001) federal-funds shocks. Panel A shows the results for all dates, whereas Panel B excludes the three scheduled FOMC announcements listed as QE1 dates by Krishnamurthy and Vissing-Jorgensen (2011) and van Binsbergen et al. (2022). Values are expressed as the drop in the industry portfolio for every 1% unexpected positive shock in 10-year swap rates.

Table 4. Individual Bank holding company stock returns on the 10-year interest rate swap changes

This table presents the sensitivity of individual bank stock returns to changes in 10-year swap rates on FOMC days excluding the 3 scheduled QE1 dates listed by Krishnamurthy and Vissing-Jorgensen (2011) and van Binsbergen et al. (2022):

$$R_{it} = \beta_{0i} + \beta_{FFi} \times \Delta FF + \beta_x \times X_{it} + \beta_s \times \Delta s + \beta_{sx} \times X_{it} \times \Delta s + \epsilon_{it}.$$

All regression control for Kuttner (2001) federal-funds shocks. For each column we also control for the same variables interacted with Δs. Columns (2), (4), (6) show the results for WLS using market capitalization as weight. All bank characteristics refer to 1 quarter before the FOMC announcement. Standard errors are clustered at the FOMC-day-level and are robust to heteroscedasticity. ***, **, * denote significance at the 1%, 5%, and 10% level, respectively.

Dependent variable:	Stock returns					
	(1)	(2)	(3)	(4)	(5)	(6)
Δs	4.749***	5.191***	1.427	1.612	0.070	0.089
	(1.790)	(1.838)	(1.302)	(1.318)	(1.302)	(1.252)
ΔFF	-3.127	-3.289	-3.329	-3.487	-3.315	-3.468
	(3.487)	(3.676)	(3.094)	(3.399)	(3.097)	(3.405)
Loans repricing in 1 year $\times \Delta s$			5.980**	6.417**	5.751**	6.196**
			(2.904)	(2.910)	(2.846)	(2.865)
Gov. sec repricing in 1 year $\times \Delta s$			1.879	2.149*	1.614	1.907
			(1.241)	(1.232)	(1.157)	(1.160)
Equity-ratio $\times \Delta s$					14.468*	16.068*
					(7.545)	(9.073)
Control	No	No	Yes	Yes	Yes	Yes
Permno fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Weighted by log(Assets)	No	Yes	No	Yes	No	Yes
R^2	0.03	0.03	0.03	0.03	0.03	0.03
Observations	95,720	95,720	88,818	88,818	88,818	88,818

References

Acosta, M., 2022. The Perceived Causes of Monetary Policy Surprises. Working paper, Federal Reserve Board.

Adrian, T., Crump, R. K., Moench, E., 2013. Pricing the term structure with linear regressions. Journal of Financial Economics 110, 110–138.

Adrian, T., Shin, H. S., 2013. Procyclical Leverage and Value-at-Risk. The Review of Financial Studies 27, 373–403.

Backus, D. K., Wright, J. H., 2007. Cracking the Conundrum. Brookings Papers on Economic Activity 38, 293–329.

Benetton, M., Fantino, D., 2021. Targeted monetary policy and bank lending behavior. Journal of Financial Economics 142, 404–429.

Benetton, M., Gavazza, A., Surico, P., 2021. Mortgage pricing and monetary policy. Bank of England working papers 936, Bank of England.

Benzarti, Y., Carloni, D., Harju, J., Kosonen, T., 2020. What goes up may not come down: Asymmetric incidence of value-added taxes. Journal of Political Economy 128, 4438–4474.

Bernanke, B. S., 2015. Inaugurating a New Blog. Brookings Institute .

Bernanke, B. S., Gertler, M., 1995. Inside the black box: The credit channel of monetary policy transmission. Journal of Economic Perspectives 9, 27–48.

Bernanke, B. S., Kuttner, K. N., 2005. What explains the stock market's reaction to federal reserve policy? The Journal of Finance 60, 1221–1257.

van Binsbergen, J. H., Diamond, W. F., Grotteria, M., 2022. Risk-free interest rates. Journal of Financial Economics 143, 1–29.

Blinder, A. S., 1999. Central banking in theory and practice. Mit press.

Borenstein, S., Cameron, A. C., Gilbert, R., 1997. Do gasoline prices respond asymmetrically to crude oil price changes? The Quarterly journal of economics 112, 305–339.

Brunnermeier, M. K., Sannikov, Y., 2014. A macroeconomic model with a financial sector. American Economic Review 104, 379–421.

Butters, R. A., Sacks, D. W., Seo, B., 2022. How do national firms respond to local cost shocks? American Economic Review 112, 1737–1772.

Cochrane, J. H., Piazzesi, M., 2002. The fed and interest rates - a high-frequency identification. American Economic Review 92, 90–95.

Cook, T., Hahn, T., 1989. The effect of changes in the federal funds rate target on market interest rates in the 1970s. Journal of monetary economics 24, 331–351.

Cooperman, H. R., Duffie, D., Luck, S., Wang, Z. Z., Yang, Y., 2023. Bank funding risk, reference rates, and credit supply. Working Paper 30907, National Bureau of Economic Research.

Crawford, G. S., Pavanini, N., Schivardi, F., 2018. Asymmetric information and imperfect competition in lending markets. American Economic Review 108, 1659–1701.

Dick-Nielsen, J., 2009. Liquidity biases in trace. The Journal of Fixed Income 19, 43–55.

Dick-Nielsen, J., 2014. How to clean enhanced trace data. Available at SSRN 2337908 .

Drechsler, I., Savov, A., Schnabl, P., 2017. The Deposits Channel of Monetary Policy*. The Quarterly Journal of Economics 132, 1819–1876.

Drechsler, I., Savov, A., Schnabl, P., 2021. Banking on deposits: Maturity transformation without interest rate risk. The Journal of Finance 76, 1091–1143.

Feliz, R. A., 2021. The Fed Explained: What the Central Bank Does. Reports and Studies 4860, Board of Governors of the Federal Reserve System (U.S.).

Gertler, M., Kiyotaki, N., 2010. Chapter 11 - financial intermediation and credit policy in business cycle analysis. Elsevier, vol. 3 of *Handbook of Monetary Economics*, pp. 547–599.

Gürkaynak, R. S., Sack, B., Wright, J. H., 2010. The tips yield curve and inflation compensation. American Economic Journal: Macroeconomics 2, 70–92.

Gómez-Cram, R., Grotteria, M., 2022. Real-time price discovery via verbal communication: Method and application to fedspeak. Journal of Financial Economics 143, 993–1025.

Gürkaynak, R., Sack, B., Swanson, E., 2005. Do actions speak louder than words? the response of asset prices to monetary policy actions and statements. International Journal of Central Banking Vol. 1, 55–93.

Gürkaynak, R. S., Sack, B., Wright, J. H., 2007. The U.S. treasury yield curve: 1961 to the present. Journal of Monetary Economics 54, 2291–2304.

Hancock, D., 1985. Bank profitability, interest rates, and monetary policy. Journal of Money, Credit and Banking 17, 189–202.

Hancock, D., Passmore, W., 2012. The Federal Reserve's portfolio and its effects on mortgage markets. Finance and Economics Discussion Series 2012-22, Board of Governors of the Federal Reserve System (U.S.).

Hannan, T. H., Berger, A. N., 1991. The rigidity of prices: Evidence from the banking industry. The American Economic Review 81, 938–945.

Hanson, S. G., Stein, J. C., 2015. Monetary policy and long-term real rates. Journal of Financial Economics 115, 429–448.

Hayashi, F., 1982. Tobin's marginal q and average q: A neoclassical interpretation. Econometrica 50, 213–224.

He, Z., Krishnamurthy, A., 2013. Intermediary asset pricing. American Economic Review 103, 732–70.

Herfindahl, O. C., 1950. Concentration in the U.S. steel industry. Tech. rep., Unpublished doctoral dissertation, Columbia University.

Hillenbrand, S., 2021. The fed and the secular decline in interest rates. Available at SSRN 3550593 .

Hirschman, A. O., 1964. The paternity of an index. The American Economic Review 54, 761–762.

Hurst, E., Keys, B. J., Seru, A., Vavra, J., 2016. Regional redistribution through the us mortgage market. American Economic Review 106, 2982–3028.

Jiang, E. X., Matvos, G., Piskorski, T., Seru, A., 2023. Monetary tightening and us bank fragility in 2023: Mark-to-market losses and uninsured depositor runs? Tech. rep., National Bureau of Economic Research.

Jordà, O., 2005. Estimation and inference of impulse responses by local projections. American Economic Review 95, 161–182.

Krishnamurthy, A., Vissing-Jorgensen, A., 2011. The Effects of Quantitative Easing on Interest Rates: Channels and Implications for Policy. Brookings Papers on Economic Activity 42, 215–287.

Kuttner, K. N., 2001. Monetary policy surprises and interest rates: Evidence from the fed funds futures market. Journal of Monetary Economics 47, 523–544.

Malkhozov, A., Mueller, P., Vedolin, A., Venter, G., 2016. Mortgage risk and the yield curve. The Review of Financial Studies 29, 1220–1253.

Nakamura, E., Steinsson, J., 2018. High-Frequency Identification of Monetary Non-Neutrality: The Information Effect. The Quarterly Journal of Economics 133, 1283–1330.

Neumark, D., Sharpe, S. A., 1992. Market structure and the nature of price rigidity: Evidence from the market for consumer deposits. The Quarterly Journal of Economics 107, 657–680.

Ottonello, P., Song, W., 2022. Financial intermediaries and the macroeconomy: Evidence from a high-frequency identification. Working Paper 29638, National Bureau of Economic Research.

Ottonello, P., Winberry, T., 2020. Financial heterogeneity and the investment channel of monetary policy. Econometrica 88, 2473–2502.

Paul, P., 2023. Banks, maturity transformation, and monetary policy. Journal of Financial Intermediation 53, 101011.

Peltzman, S., 2000. Prices rise faster than they fall. Journal of Political Economy 108, 466–502.

Samuelson, P. A., 1945. The effect of interest rate increases on the banking system. The American Economic Review 35, 16–27.

Scharfstein, D., Sunderam, A., 2016. Market power in mortgage lending and the transmission of monetary policy. Unpublished working paper. Harvard University 2.

Seltzer, L. H., Starks, L., Zhu, Q., 2022. Climate regulatory risk and corporate bonds. Working Paper 29994, National Bureau of Economic Research.

Siriwardane, E. N., 2019. Limited investment capital and credit spreads. The Journal of Finance 74, 2303–2347.

Stiglitz, J., Greenwald, B., 2003. Towards a New Paradigm in Monetary Economics. Raffaele Mattioli Lectures, Cambridge University Press.

Tobin, J., 1969. A general equilibrium approach to monetary theory. Journal of Money, Credit and Banking 1, 15–29.

Wang, Y., Whited, T. M., Wu, Y., Xiao, K., 2022. Bank market power and monetary policy transmission: Evidence from a structural estimation. The Journal of Finance 77, 2093–2141.

Appendix A Additional results

Table A.1. Dates of scheduled FOMC meetings since 2000

		Scheduled FOMC meetings							
Year	N	1.	2.	3.	4.	5.	6.	7.	8.
2000	8	02-Feb	21-Mar	16-May	28-Jun	22-Aug	03-Oct	15-Nov	19-Dec
2001	8	31-Jan	20-Mar	15-May	27-Jun	21-Aug	02-Oct	06-Nov	11-Dec
2002	8	30-Jan	19-Mar	07-May	26-Jun	13-Aug	24-Sep	06-Nov	10-Dec
2003	8	29-Jan	18-Mar	06-May	25-Jun	12-Aug	16-Sep	28-Oct	09-Dec
2004	8	28-Jan	16-Mar	04-May	30-Jun	10-Aug	21-Sep	10-Nov	14-Dec
2005	8	02-Feb	22-Mar	03-May	30-Jun	09-Aug	20-Sep	01-Nov	13-Dec
2006	8	31-Jan	28-Mar	10-May	29-Jun	08-Aug	20-Sep	25-Oct	12-Dec
2007	8	31-Jan	21-Mar	09-May	28-Jun	07-Aug	18-Sep	31-Oct	11-Dec
2008	8	30-Jan	18-Mar	30-Apr	25-Jun	05-Aug	16-Sep	29-Oct	16-Dec
2009	8	28-Jan	18-Mar	29-Apr	24-Jun	12-Aug	23-Sep	04-Nov	16-Dec
2010	8	27-Jan	16-Mar	28-Apr	23-Jun	10-Aug	21-Sep	03-Nov	14-Dec
2011	8	26-Jan	15-Mar	27-Apr	22-Jun	09-Aug	21-Sep	02-Nov	13-Dec
2012	8	25-Jan	13-Mar	25-Apr	20-Jun	01-Aug	13-Sep	24-Oct	12-Dec
2013	8	30-Jan	20-Mar	01-May	19-Jun	31-Jul	18-Sep	30-Oct	18-Dec
2014	8	29-Jan	19-Mar	30-Apr	18-Jun	30-Jul	17-Sep	29-Oct	17-Dec
2015	8	28-Jan	18-Mar	29-Apr	17-Jun	29-Jul	17-Sep	28-Oct	16-Dec
2016	8	27-Jan	16-Mar	27-Apr	15-Jun	27-Jul	21-Sep	02-Nov	14-Dec
2017	8	01-Feb	15-Mar	03-May	14-Jun	26-Jul	20-Sep	01-Nov	13-Dec
2018	8	31-Jan	21-Mar	02-May	13-Jun	01-Aug	26-Sep	08-Nov	19-Dec
2019	8	30-Jan	20-Mar	01-May	19-Jun	31-Jul	18-Sep	30-Oct	11-Dec

Table A.2. Rolling-window correlation between daily changes in 10-year swap rates and 1-year bond yields against the level of 1-year bond yields

Dependent variable:	Roll.-wind. corr. b/w Δs and Δy^1 from $t-365$ and t				
	(1)	(2)	(3)	(4)	
y_{t-365}	0.050***		0.036***	0.322***	0.272***
	(0.001)		(0.001)	(0.007)	(0.005)
y_t		0.048***	0.020***		0.056***
		(0.001)	(0.001)		(0.006)
y_{t-365}^2				-0.087***	-0.077***
				(0.003)	(0.002)
y_t^2					-0.000
					(0.002)
y_{t-365}^3				0.007***	0.006***
				(0.000)	(0.000)
y_t^3					-0.001***
					(0.000)
R^2	0.36	0.30	0.38	0.60	0.66
N	4,660	4,703	4,660	4,660	4,660

Notes: We first compute the rolling-window correlation between daily changes in 10-year swap rates and 1-year bond yields over 365 days. This table presents the estimates from regressing this rolling-window correlation against the level of 1-year bond yields (current or lagged 365 days), its square and cube. The sample is from January 1997 to January 2023.

Table A.3. The relation between 10-year swap rates and the long-run dots

Dependent variable:	Δs		
	(1)	(2)	(3)
$\Delta\mathbb{E}$[Long-term fed funds rate]	0.424**	0.411**	2.112***
	(0.173)	(0.183)	(0.603)
Uncertainty		0.272*	
		(0.144)	
Tercile(2)			-0.074*
			(0.039)
Tercile(3)			-0.130***
			(0.036)
Tercile(2) $\times$ $\Delta\mathbb{E}$[Long-term fed funds rate]			-1.663**
			(0.643)
Tercile(2) $\times$ $\Delta\mathbb{E}$[Long-term fed funds rate]			-2.164***
			(0.678)
R^2	0.14	0.24	0.45
Observations	32	32	32

Notes: The table shows the estimates from regressing the daily change in 10-year swap rates on the FOMC meeting participants' median forecast for the long-term level of the federal funds rate. Uncertainty is the standard deviation of forecasts for the meeting with each value weighted by the number of people forecasting that value. Tercile represents the tercile of the 1-year government bond yield levels on the 32 dates in the sample. Robust standard errors are shown in parentheses. The sample is from January 2012 to December 2019.

Table A.4. Variation in mortgage rates

Panel A: Balanced panel – MSA-year-month and Lender-date

Dependent variable:	Mortgage rates				
	(1)	(2)	(3)	(4)	(5)
4-week lagged 10-year swap	Yes	No	No	No	No
Date fixed effects	No	Yes	Yes	No	No
Borrower's characteristics	No	No	Yes	Yes	Yes
Lender−date fixed effects	No	No	No	Yes	Yes
MSA−year−month fixed effects	No	No	No	No	Yes
R^2	86.55	88.42	89.87	92.09	92.38
Observations	4,483,508	4,483,508	4,483,508	4,483,508	4,483,508

Panel B: Balanced panel – MSA-Lender-date

Dependent variable:	Mortgage rates				
	(1)	(2)	(3)	(4)	(5)
4-week lagged 10-year swap	Yes	No	No	No	No
Date fixed effects	No	Yes	Yes	No	No
Borrower's characteristics	No	No	Yes	Yes	Yes
Lender−date fixed effects	No	No	No	Yes	No
Lender−MSA−date fixed effects	No	No	No	No	Yes
R^2	87.10	88.98	90.35	92.17	93.78
Observations	2,993,199	2,993,199	2,993,199	2,993,199	2,993,199

Notes: This table presents the coefficient of determination (R^2) for different specifications of mortgage rates. Borrower's characteristics include the level and the square of the FICO score at origination and of the loan-to-value (LTV) ratio at origination, whether the mortgage was GSE-eligible, insured at origination, or whether it was prime or subprime. The sample is from January 2000 to July 2020.

Table A.5. Variation in corporate bond yields

Panel A: Whole sample

Dependent variable:	Corporate bond yields				
	(1)	(2)	(3)	(4)	(5)
10-year swap	Yes	No	Yes	No	No
Date fixed effects	No	Yes	No	Yes	No
Borrower fixed effects	No	No	Yes	Yes	No
Borrower−year−month fixed effects	No	No	No	No	Yes
R^2	42.59	54.71	74.14	86.45	98.59
Observations	770,878	770,878	770,878	770,878	770,878

Panel B: AA-rated companies

Dependent variable:	Corporate bond yields – AA-rated firms				
	(1)	(2)	(3)	(4)	(5)
10-year swap	Yes	No	Yes	No	No
Date fixed effects	No	Yes	No	Yes	No
Borrower fixed effects	No	No	Yes	Yes	No
Borrower−year−month fixed effects	No	No	No	No	Yes
R^2	85.78	95.40	92.13	98.01	99.22
Observations	48,348	48,266	48,348	48,266	48,348

Notes: This table presents the coefficient of determination (R^2) for different specifications of corporate bond yields. Panel B restricts the sample to AA-rated firms. The sample for both panels is from January 2000 to December 2019.
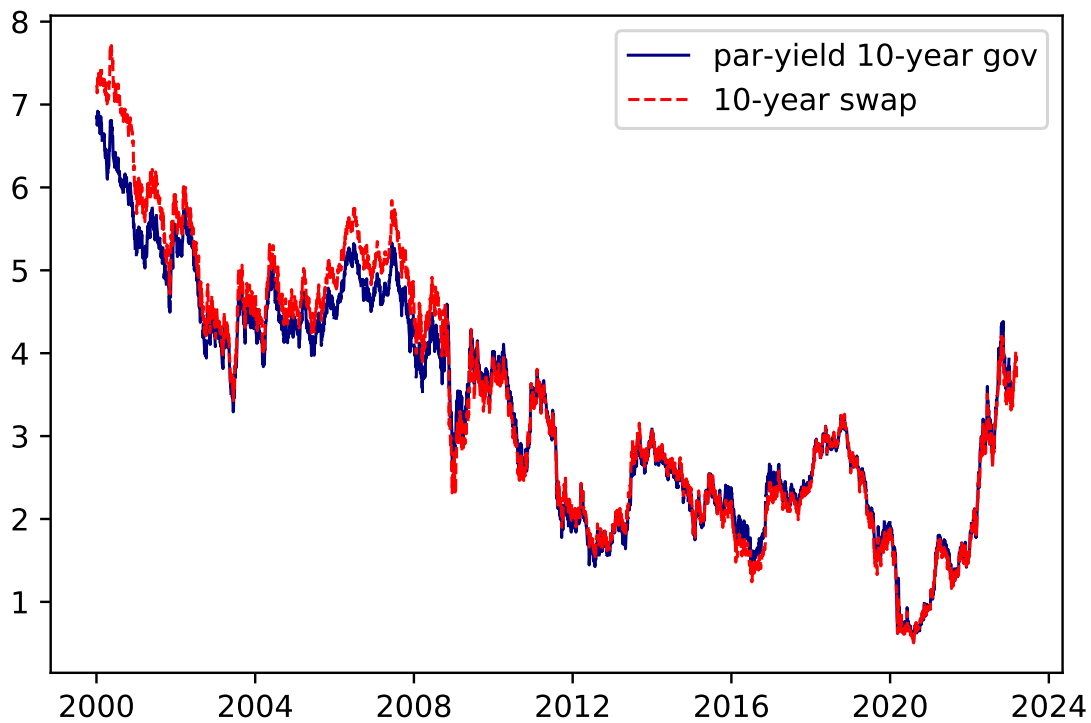
Fig. A.1. *Notes*: The Figure shows the 10-year swap rate (annualized to reflect 365 days) against the 10-year government-bond par-yield as computed by Gürkaynak et al. (2007). All rates are continuously compounded.
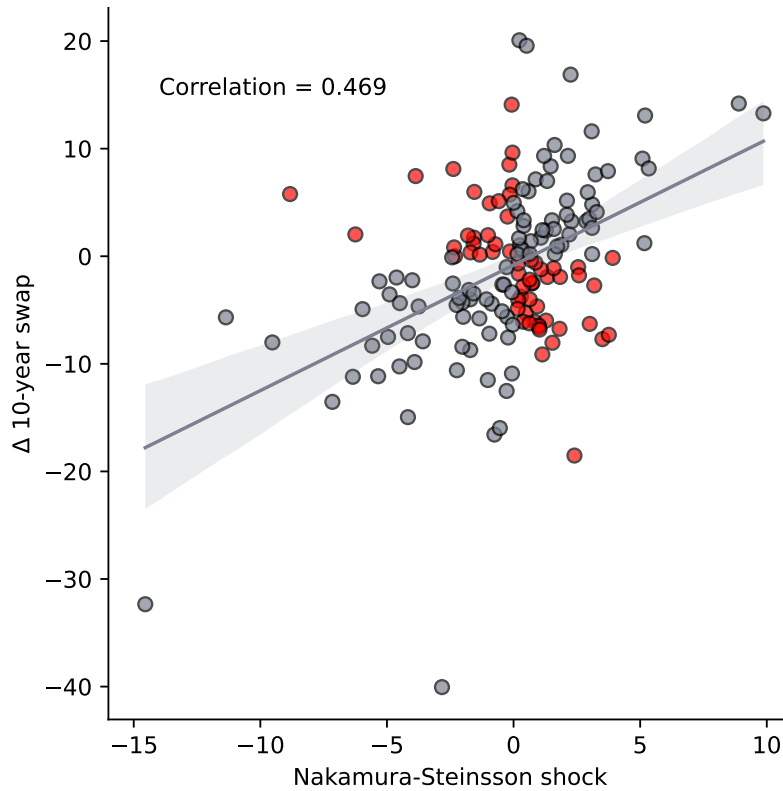
Fig. A.2. *Notes*: The figure shows the scatterplot of the changes in 10-year swap rates on FOMC days against Nakamura and Steinsson (2018) shocks computed by Acosta (2022). Values are expressed in basis points. The grey dots represent FOMC events for which the changes in 10-year swap rates on FOMC and Nakamura and Steinsson (2018) shocks shared the same sign. The red dots are events in which the two shocks had opposite signs. The sample includes all scheduled FOMC meetings from February 2, 2000 to December 11, 2019.
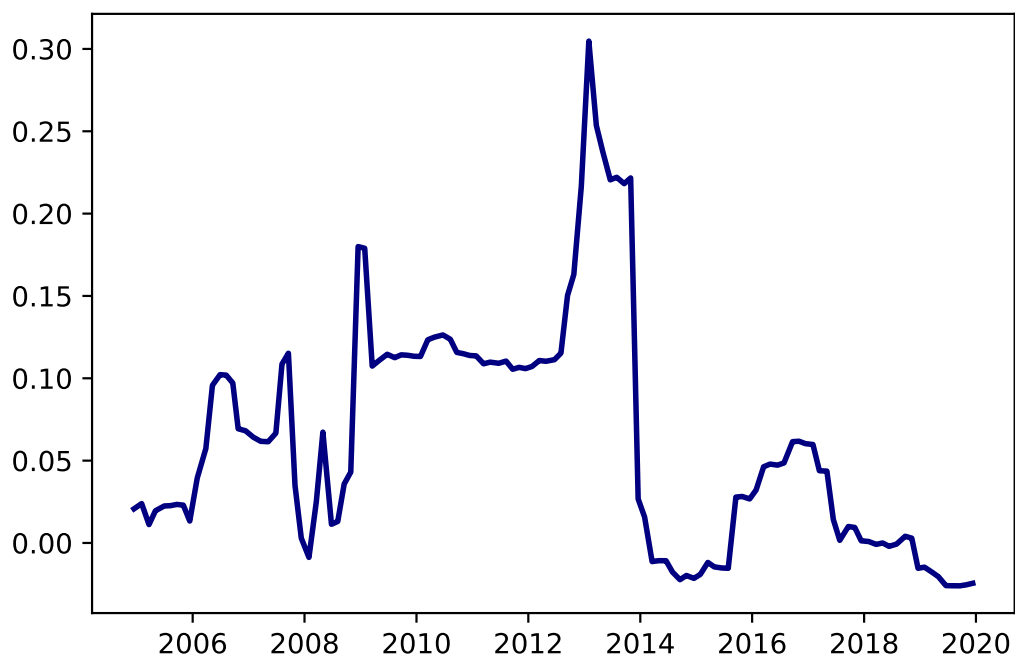
Fig. A.3. *Notes*: The figure shows the 5-year rolling window of the adjusted R^2 from regressing changes in 10-year swap rates on FOMC days against changes Kuttner (2001) Federal funds rate shock computed by Acosta (2022). The sample includes all scheduled FOMC meetings from February 2, 2000 to December 11, 2019.

Fig. A.4. *Notes*: The figure shows the 5-year rolling window of the adjusted R^2 from regressing changes in 10-year swap rates on FOMC days against Nakamura and Steinsson (2018) shocks computed by Acosta (2022). Values are expressed in basis points. The sample includes all scheduled FOMC meetings from February 2, 2000 to December 11, 2019.

56
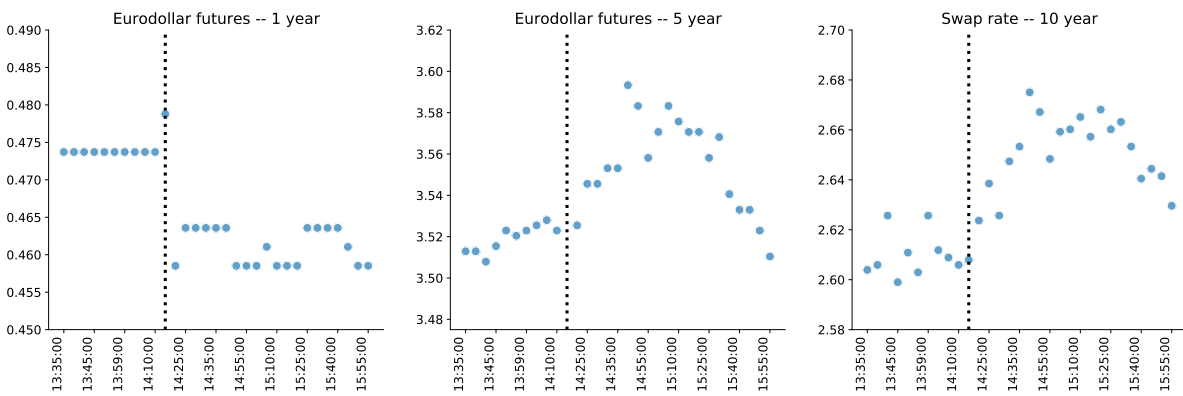
Fig. A.5. *Notes*: The Figure shows the intraday evolution of the implied rate from the 12-month Eurodollar futures, the 5-year Eurodollar futures, and the 10-year swap rate on November 03, 2010. The black dashed vertical line highlights the time in which the FOMC statement was released (14:16). All rates are continuously compounded.

57
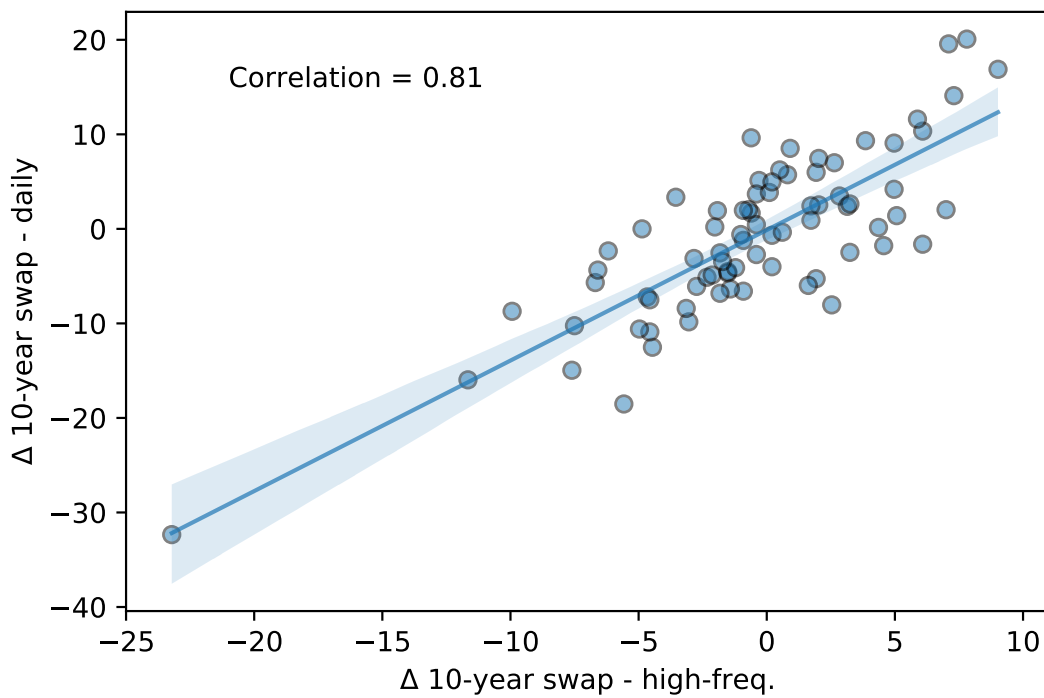
Fig. A.6. *Notes*: The figure shows the scatterplot of the intradaily changes in 10-year swap rates on FOMC days against the daily changes in 10-year swap rates on the same days. Values are expressed in basis points. The sample includes all scheduled FOMC meetings from February 2, 2000 to December 11, 2019.

Fig. A.7. **Response of 10-year government bond yields to daily FOMC surprises.**

Notes: The figure shows the estimated response of 10-year government bond yields to monetary policy surprises and corresponding 95% confidence interval from the following regression:

$$c_{hf} = a_f + \sum_{j=-5}^{-2} \gamma_{Nj} \mathbb{1}_{h=j} + \sum_{j=0}^{28} \gamma_{Nj} \mathbb{1}_{h=j} + \sum_{j=-5}^{-2} \gamma_{Pj} \mathbb{D}\ \mathbb{1}_{h=j} + \sum_{j=0}^{28} \gamma_{Pj} \mathbb{D}\mathbb{1}_{h=j} + \sum_{j=-5}^{-2} \beta_{Nj} \mathbb{1}_{h=j} \Delta s_f$$
$$+ \sum_{j=0}^{28} \beta_{Nj} \mathbb{1}_{h=j} \Delta s_f + \sum_{j=-5}^{-2} \beta_{Pj} \Delta s_f\ \mathbb{1}_{h=j}\ \mathbb{D}\ + \sum_{j=0}^{28} \beta_{Pj} \Delta s_f\ \mathbb{1}_{h=j}\ \mathbb{D} + \epsilon_{hf}$$

where h represents the number of days from the FOMC announcement day f, c is the par-yield on 10-year nominal government bonds as computed by Gürkaynak et al. (2007), Δs_f is the absolute value of the change in par swap rate for 10-year tenor around the FOMC announcement f, $\mathbb{1}_{h=j}$ is a dummy variable taking value 1 if h is equal to j and zero otherwise, $\mathbb{D}$ is a dummy variable taking a value of 1 if Δs_f is positive and zero otherwise, and ϵ is the error term. Standard errors are clustered at the year-month level. Data span January 2000 to December 2019.

59

Fig. A.8. **Response of 10-year TIPS yields to daily FOMC surprises.**

Notes: The figure shows the estimated response of 10-year TIPS yields to monetary policy surprises and corresponding 95% confidence interval from the following regression:

$$c_{hf} = a_f + \sum_{j=-5}^{-2} \gamma_{Nj} \mathbb{1}_{h=j} + \sum_{j=0}^{28} \gamma_{Nj} \mathbb{1}_{h=j} + \sum_{j=-5}^{-2} \gamma_{Pj} \mathbb{D}\ \mathbb{1}_{h=j} + \sum_{j=0}^{28} \gamma_{Pj} \mathbb{D} \mathbb{1}_{h=j} + \sum_{j=-5}^{-2} \beta_{Nj} \mathbb{1}_{h=j} \Delta s_f$$
$$+ \sum_{j=0}^{28} \beta_{Nj} \mathbb{1}_{h=j} \Delta s_f + \sum_{j=-5}^{-2} \beta_{Pj} \Delta s_f\ \mathbb{1}_{h=j}\ \mathbb{D}\ + \sum_{j=0}^{28} \beta_{Pj} \Delta s_f\ \mathbb{1}_{h=j}\ \mathbb{D} + \epsilon_{hf}$$

where h represents the number of days from the FOMC announcement day f, c is the par-yield on 10-year TIPS as computed by Gürkaynak, Sack, and Wright (2010), Δs_f is the absolute value of the change in par swap rate for 10-year tenor around the FOMC announcement f, $\mathbb{1}_{h=j}$ is a dummy variable taking value 1 if h is equal to j and zero otherwise, $\mathbb{D}$ is a dummy variable taking a value of 1 if Δs_f is positive and zero otherwise, and ϵ is the error term. Standard errors are clustered at the year-month level. Data span January 2000 to December 2019.

60

Fig. A.9. **Herfindahl-Hirschman Index distribution.**

Panel A: Histogram of HHI from 2000 to 2017
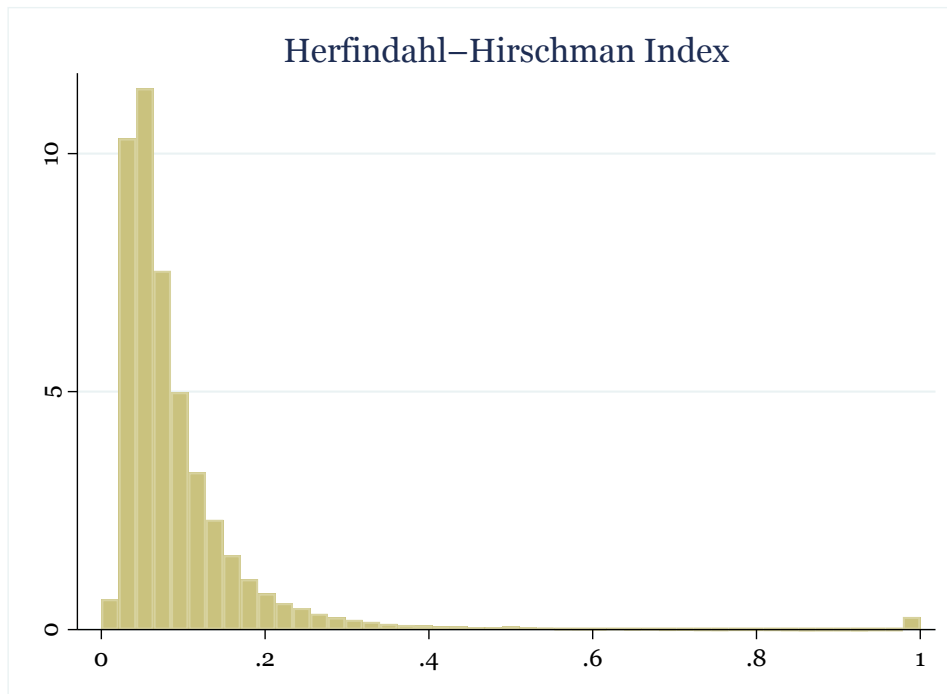
Panel B: Spatial variation HHI
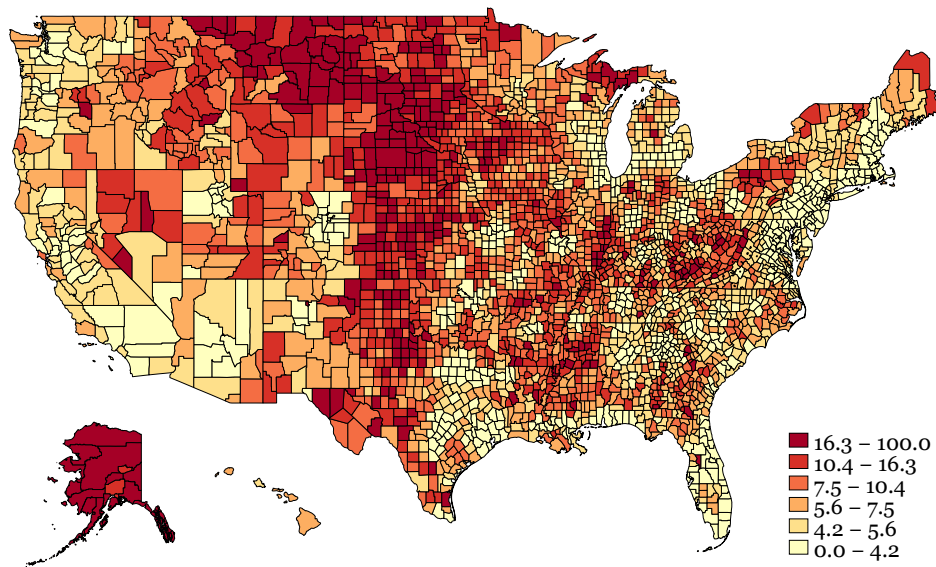
Notes: The figure shows the histogram of the Herfindahl-Hirschman Index in our sample and the spatial distribution of the average Herfindahl-Hirschman Index in our sample.

61

Fig. A.10. **Response of mortgage interest rates to daily FOMC shocks by Herfindahl-Hirschman index (HHI)**

Panel A: Response to a negative shock
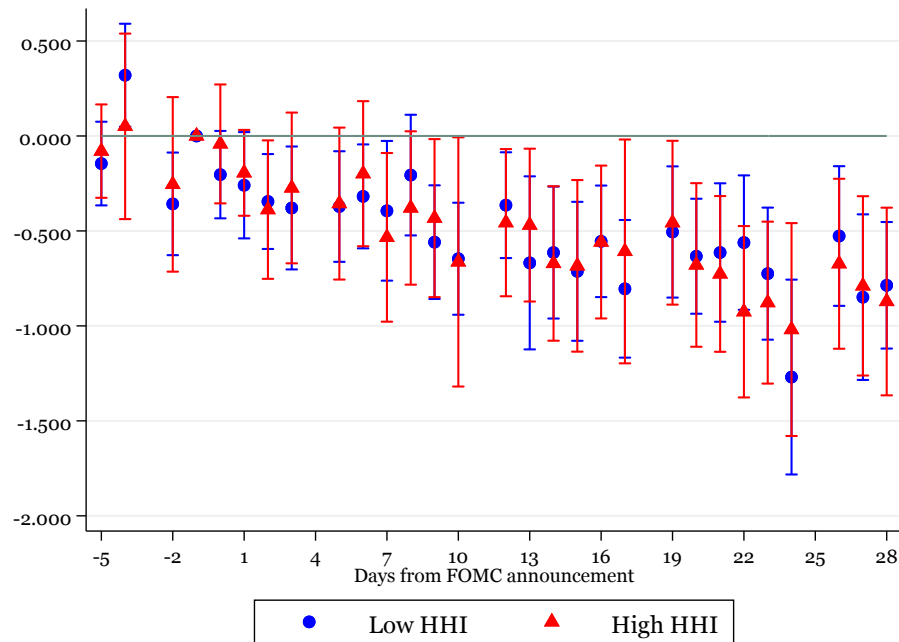
Panel B: Response to a positive shock
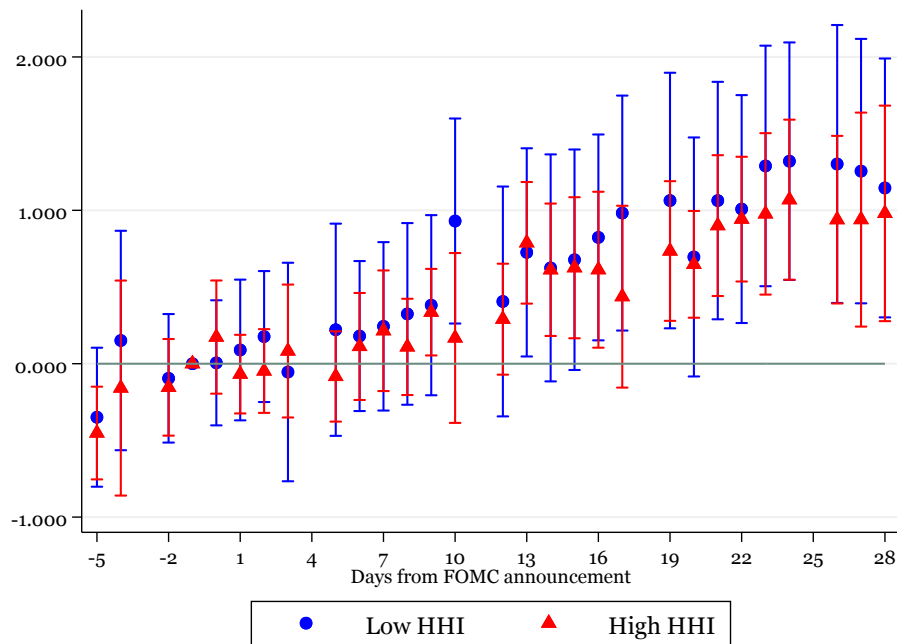
Notes: For each year-quarter we create quintiles based on the HHI computed from HMDA. We interact the response of mortgage rates to swap rates by the quintile. Data span from January 2000 to December 2019.

62

Fig. A.11. **Response of mortgage interest rates to daily FOMC shocks by share of top 4 lenders in a FIPS county**

Panel A: Response to a negative shock
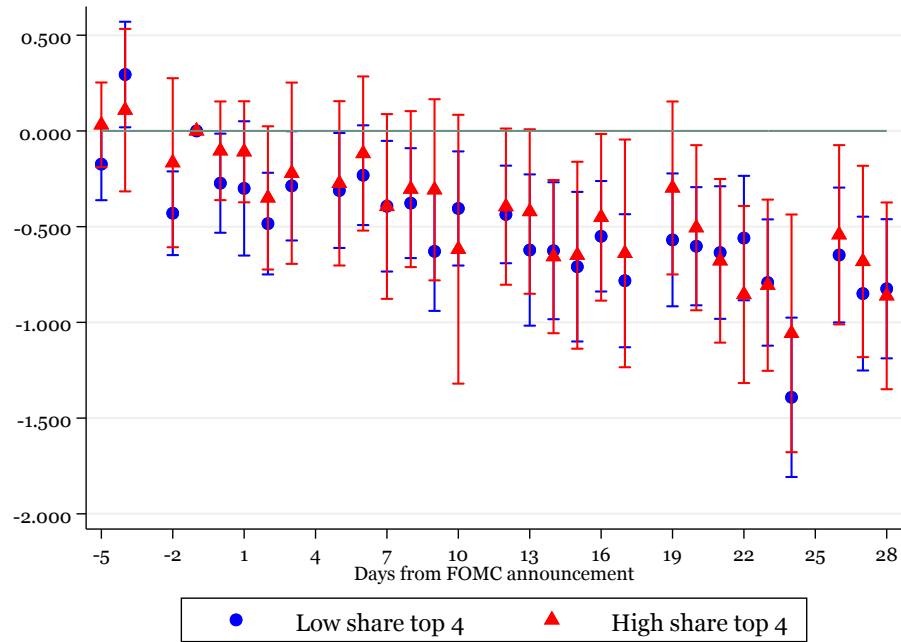
Panel B: Response to a positive shock
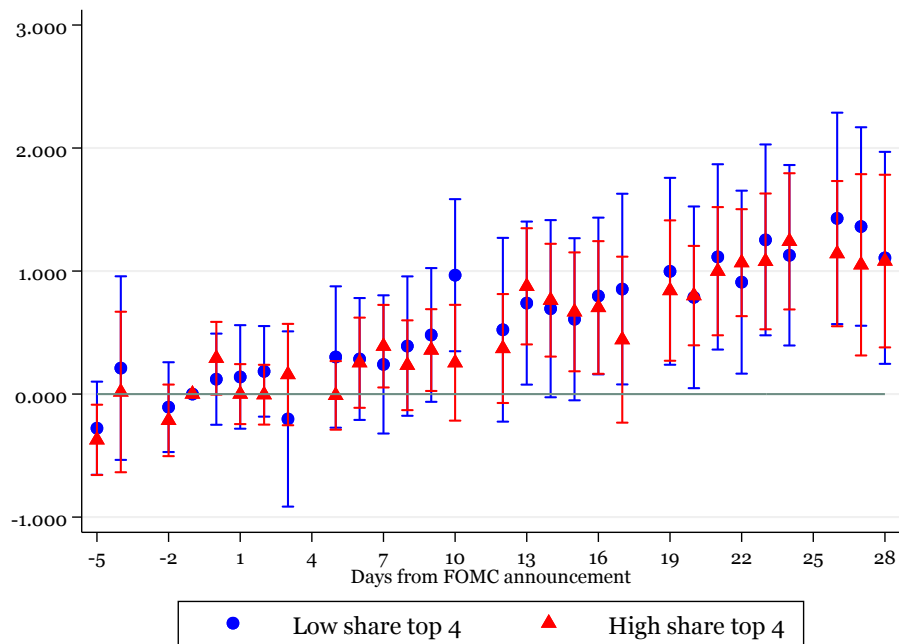
Notes: For each year-quarter we create quintiles based on the share of the top 4 lenders in a FIPS county computed from HMDA. We interact the response of mortgage rates to swap rates by the quintile. Data span from January 2000 to December 2019.

63

Fig. A.12. **Response of mortgage interest rates to daily FOMC shocks by mortgage excess demand**

Panel A: Response to a negative shock
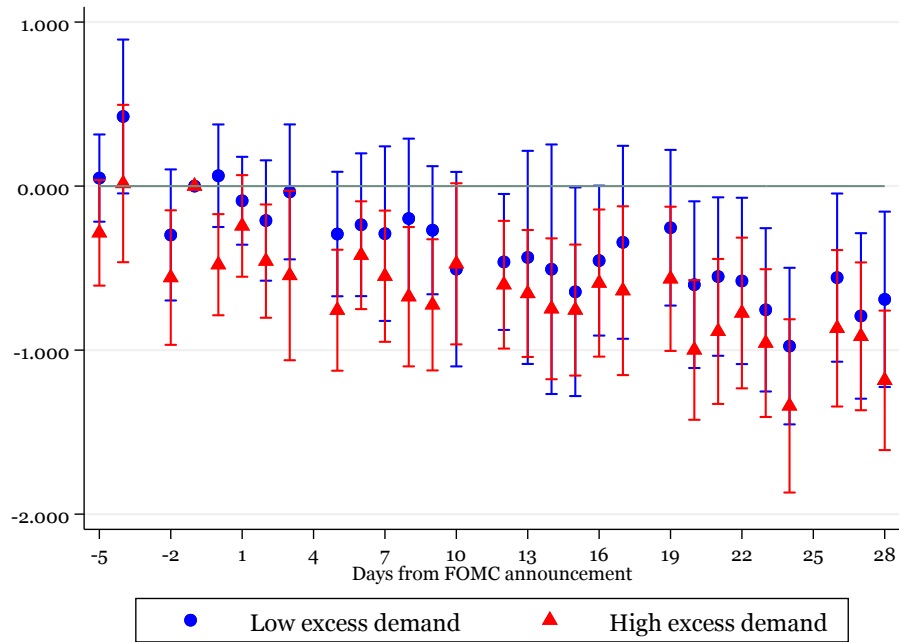
Panel B: Response to a positive shock
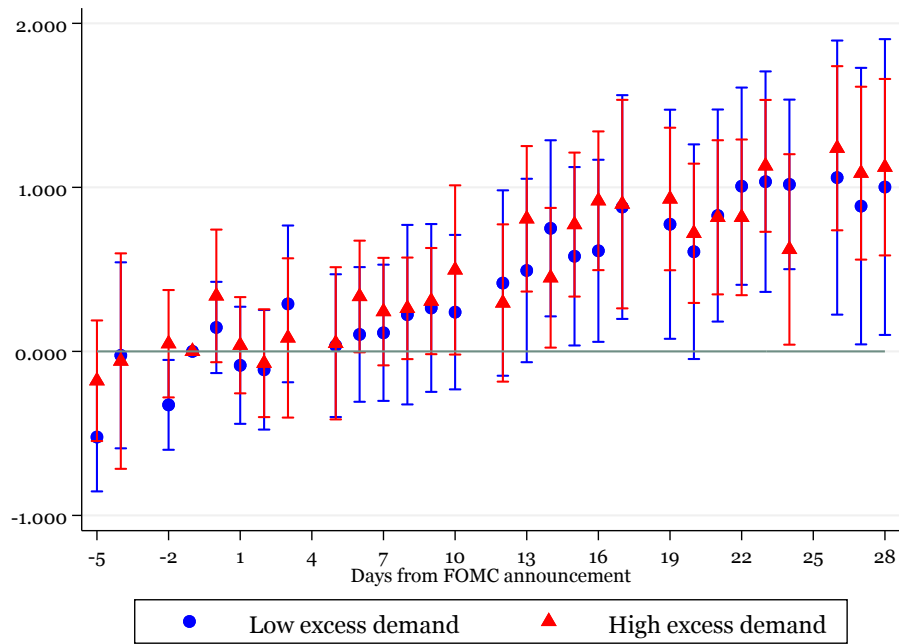
Notes: For each year-quarter we create quintiles based on the loan excess demand (the amount of loans approved plus the amount of loans rejected over the amount of loans approved by county FIPS) computed from HMDA. We interact the response of mortgage rates to swap rates by the quintile. Data span from January 2000 to December 2019.

64

Fig. A.13. **Spatial variation in mortgage rates controlling for borrower and loan characteristics.**

■	0.058 − 0.278
■	0.037 − 0.058
■	0.010 − 0.037
■	-0.011 − 0.010
■	-0.038 − -0.011
■	-0.191 − -0.038
■	No data
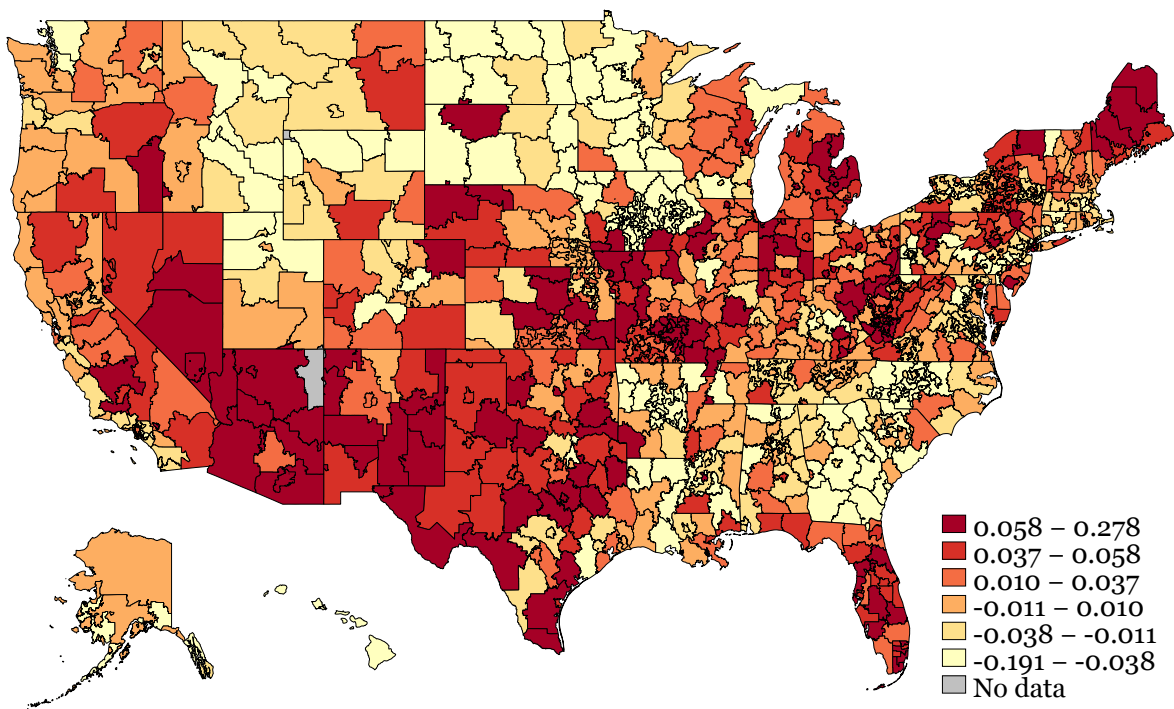
Notes: The figure shows the spatial variation in the residualized mortgage rates from Freddie mac dataset after controlling for borrower and loan characteristics following Hurst et al. (2016).

65

Fig. A.14. **Response of mortgage interest rates to daily FOMC shocks by residualized mortgage rates (intresid)**

Panel A: Response to a negative shock
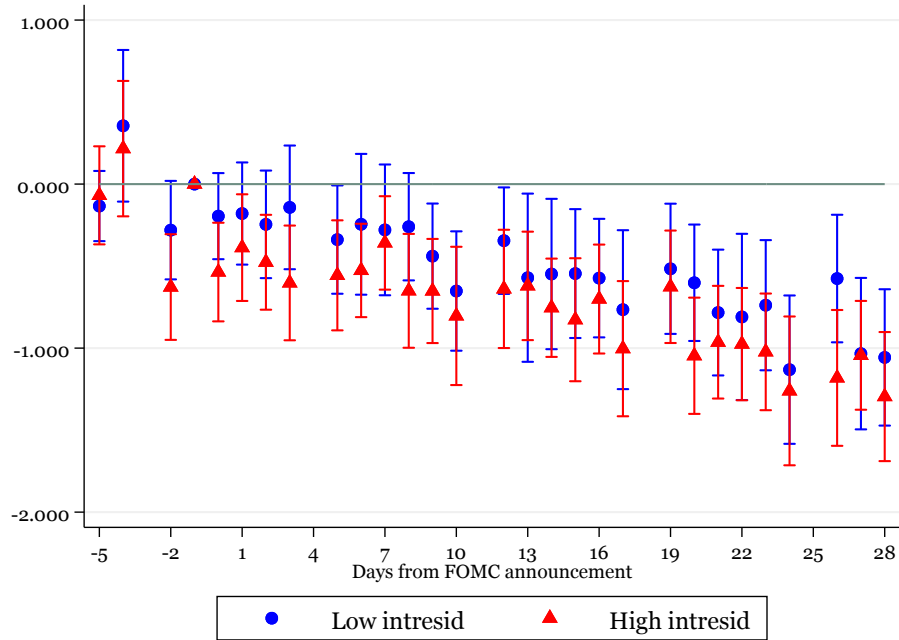
Panel B: Response to a positive shock
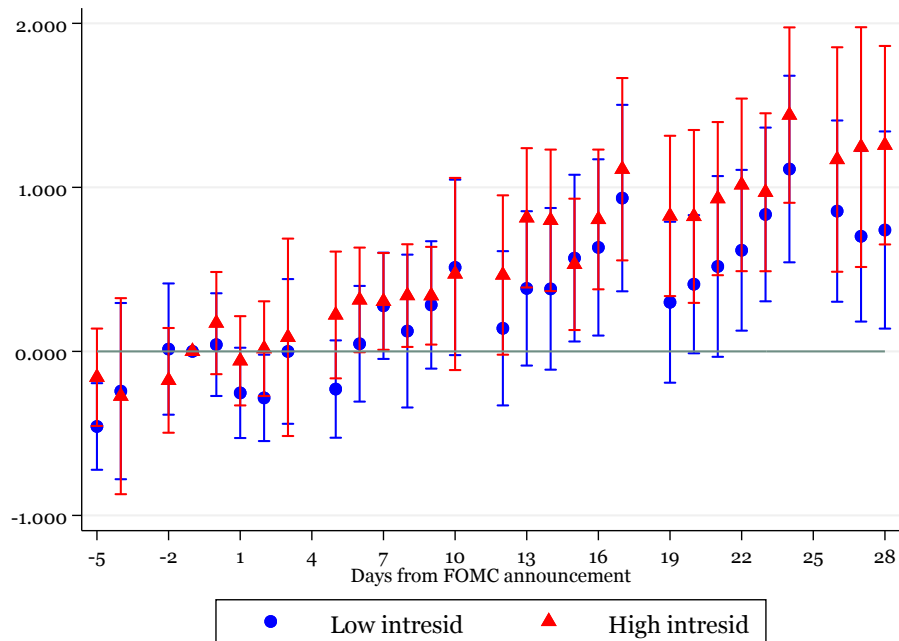
Notes: For each year-quarter we create quintiles based on the residualized mortgage rates from Freddie mac dataset after controlling for borrower and loan characteristics computed following Hurst et al. (2016) (intresid). We interact the response of mortgage rates to swap rates by the quintile. Data span from January 2000 to December 2019.

66

Fig. A.15. **Response of corporate bond yields to daily FOMC shocks by ratings**

Panel A: January 2000 to December 2019 – Investment grade
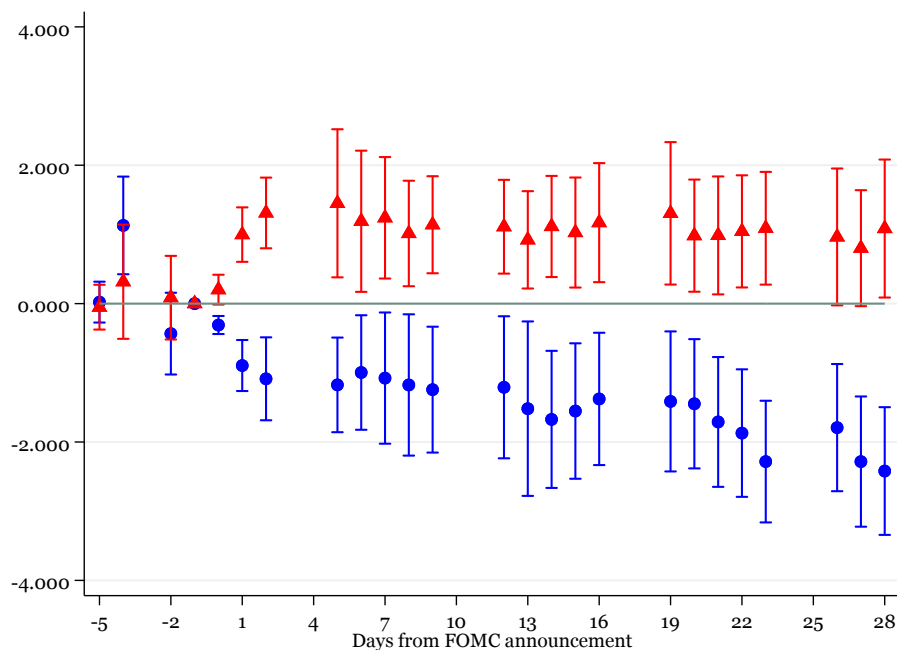
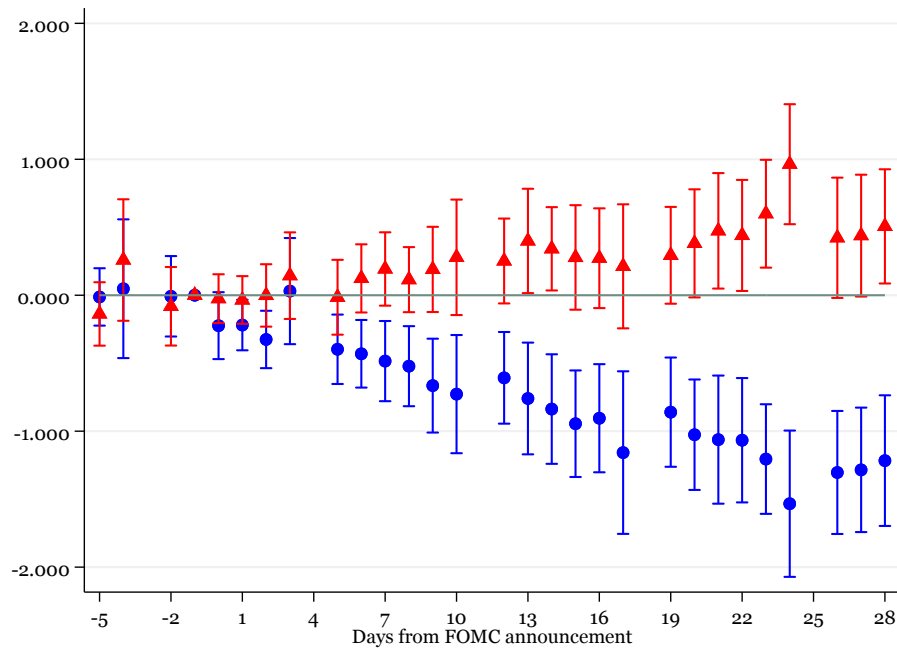Panel B: January 2000 to December 2019 – Speculative grade
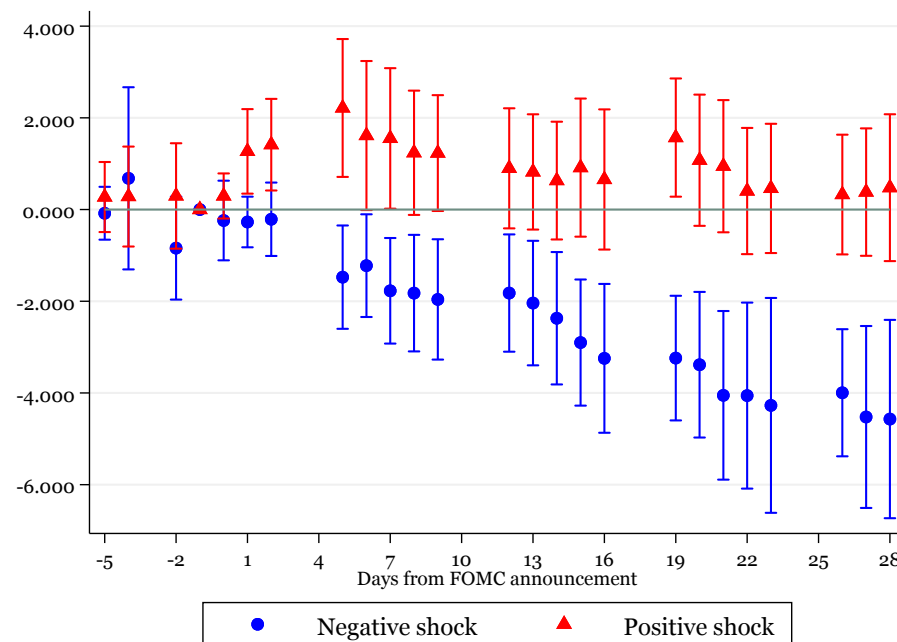
Notes: The figure reports the slope coefficient β_h and 95%-confidence interval from (5) for corporate bond yields. The regression controls for issue-cusip-by-FOMC-event fixed effects as well as borrower's characteristics. Standard errors are clustered at year-month level. Data span from January 2000 to December 2019.

67

Fig. A.16. **Response of mortgage interest rates to daily government par-yield shocks and the difference between swap rate and gov. par-yield**

Panel A: Response to svenpy10

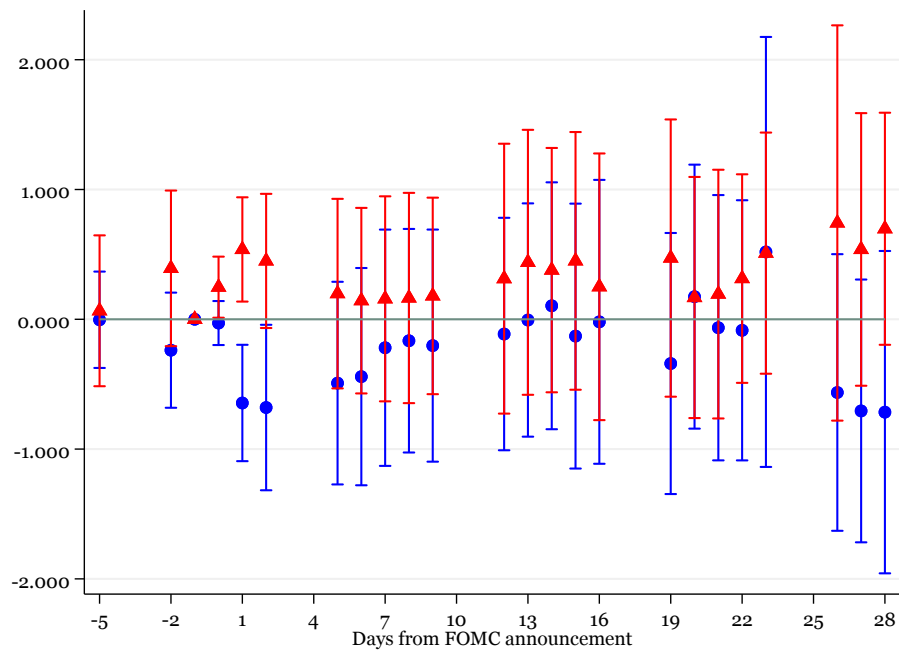Panel B: Response to the difference between swap rate and svenpy10

Notes: The figure reports the slope coefficient β_h and 95%-confidence interval from Equation 4. The regression controls for lender-by-metropolitan-area-by-FOMC-event fixed effects as well as borrower's characteristics. Standard errors are clustered at MSA $\times$ origination year-month level. The sample consists of all the conventional loans (not originated under a government program) where borrower's stated purpose is to purchase a property and the property type is either condominium or single family residence. Data span from January 2000 to December 2019.

Fig. A.17. **Response of corporate bond yields to daily government par-yield shocks and the difference between swap rate and gov. par-yield**

Panel A: Response to svenpy10

Panel B: Response to the difference between swap rate and svenpy10
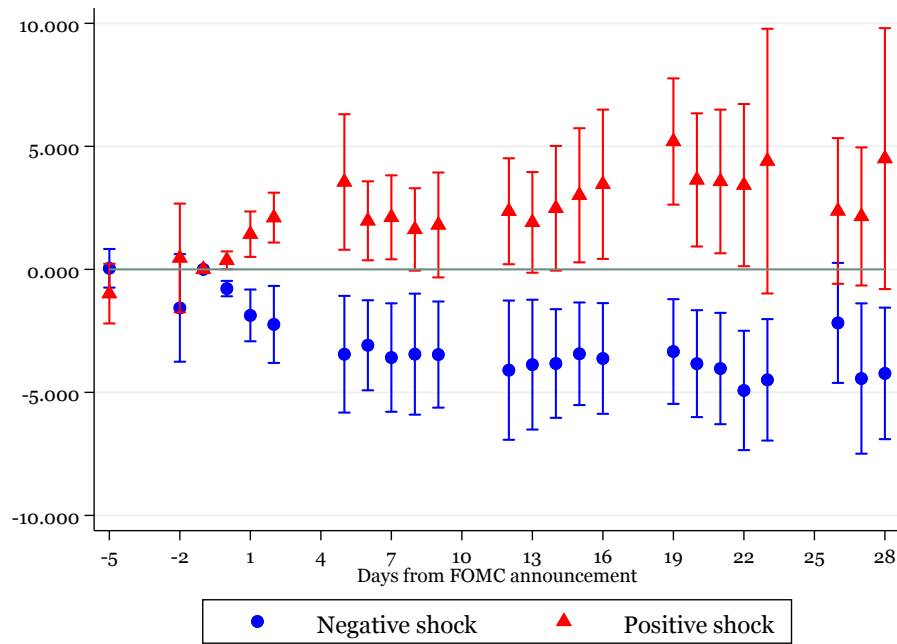
Notes: The figure reports the slope coefficient β_h and 95%-confidence interval from (5) for corporate bond yields. The regression controls for issue-cusip-by-FOMC-event fixed effects as well as borrower's characteristics. Standard errors are clustered at year-month level. Data span from January 2000 to December 2019.

69

Fig. A.18. **Response of mortgage interest rates to to news about expected future rates and term premia, Adrian et al. (2013)**

Panel A: Response to future short-term rates
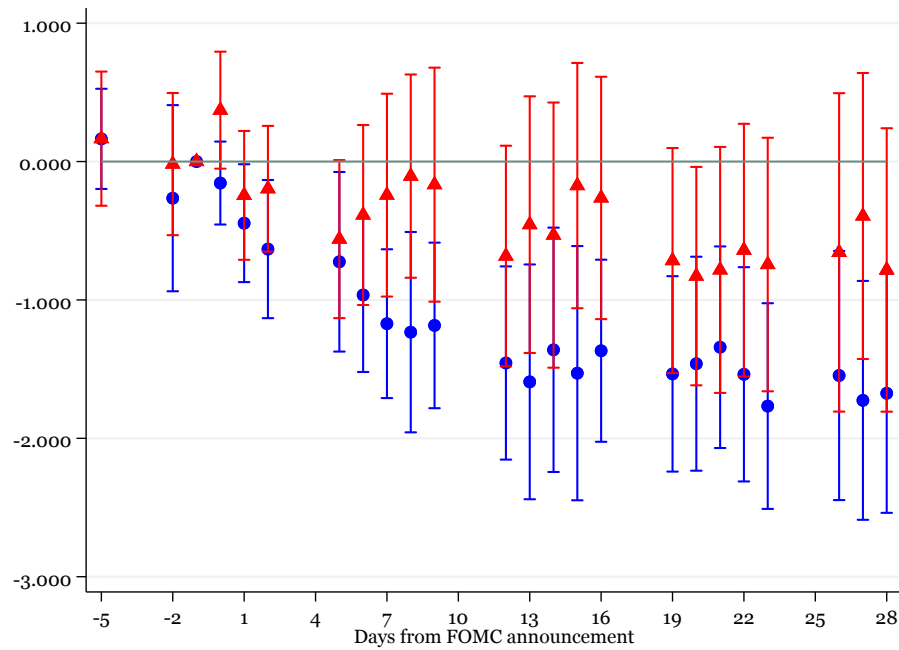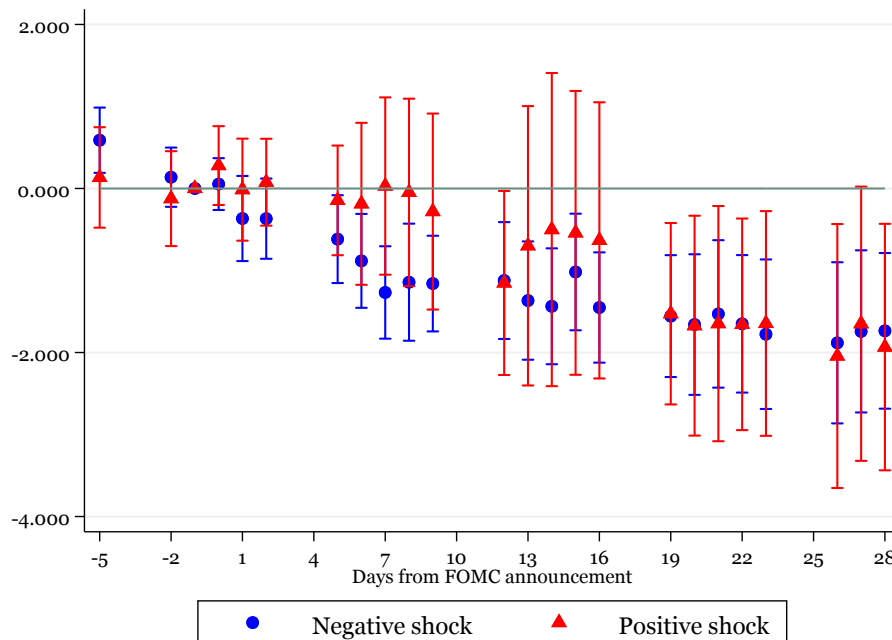

Panel B: Response to term premia

70

Panel C: Response to swap rate change residualized by future short-term rates and term premia

Notes: The figure reports the slope coefficient β_h and 95%-confidence interval from Equation 4. The regression controls for lender-by-metropolitan-area-by-FOMC-event fixed effects as well as borrower's characteristics. Standard errors are clustered at MSA × origination year-month level. The sample consists of all the conventional loans (not originated under a government program) where borrower's stated purpose is to purchase a property and the property type is either condominium or single family residence. Data span from January 2000 to December 2019.

71

Fig. A.19. **Response of corporate bond yields to news about expected future rates and term premia, Adrian et al. (2013)**

Panel A: Response to future short-term rates
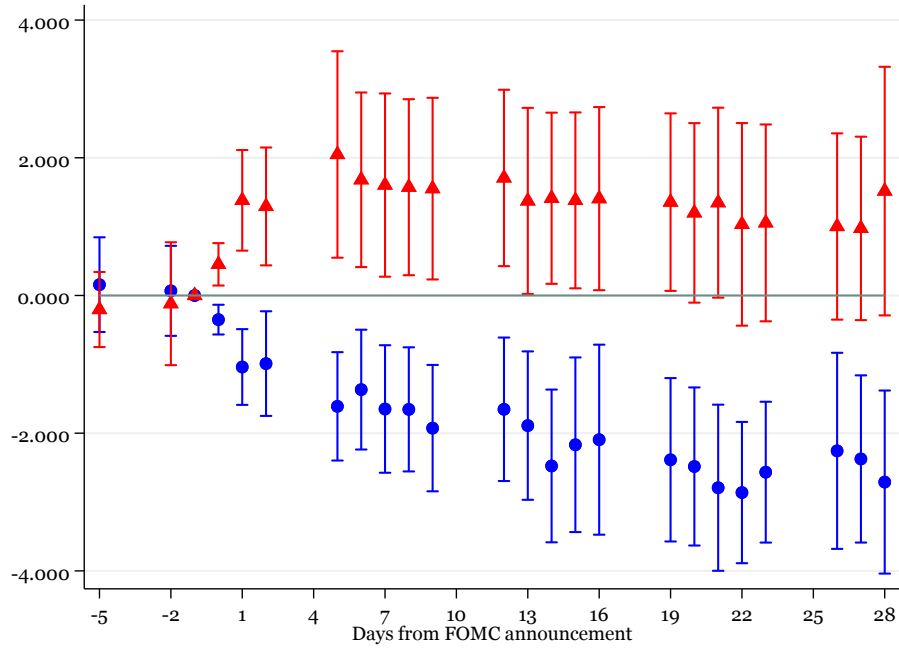
Panel B: Response to term premia

Panel C: Response to swap rate change residualized by future short-term rates and term premia

Notes: The figure reports the slope coefficient β_h and 95%-confidence interval from (5) for corporate bond yields. The regression controls for issue-cusip-by-FOMC-event fixed effects as well as borrower's characteristics. Standard errors are clustered at year-month level. Data span from January 2000 to December 2019.

73

Panel A: January 2000 to December 2019

Panel B: January 2010 to December 2019

Notes: The figure reports the slope coefficient β_h on Nakamura and Steinsson (2018) shocks and 95%-confidence interval from Equation 4. The regression controls for lender-by-metropolitan-area-by-FOMC-event fixed effects as well as borrower's characteristics. Standard errors are clustered at MSA × origination year-month level. The sample consists of all the conventional loans (not originated under a government program) where borrower's stated purpose is to purchase a property and the property type is either condominium or single family residence. Data span from January 2000 to December 2019.

74
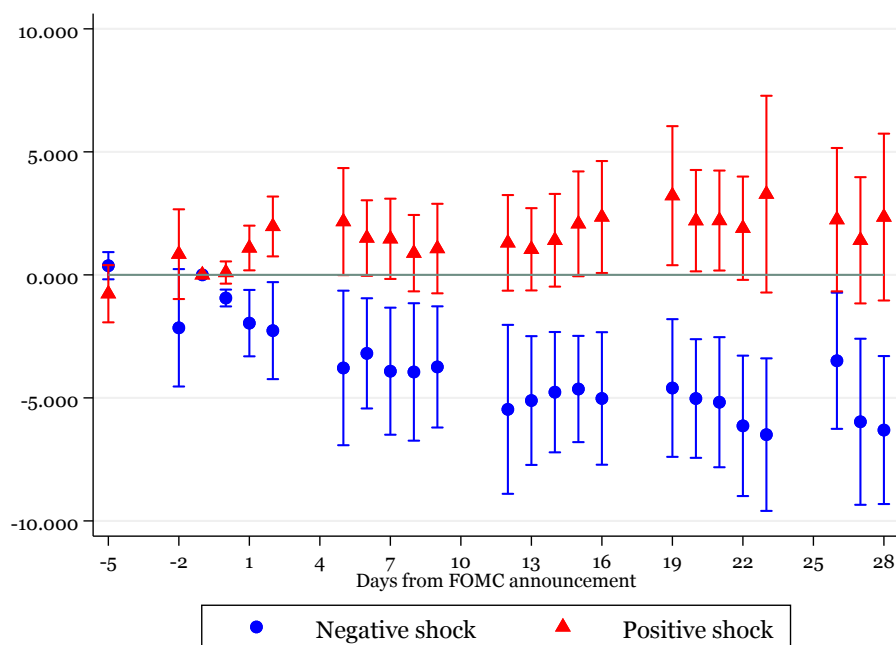
Fig. A.21. **Response of corporate bond yields to Nakamura and Steinsson (2018) shocks**

Panel A: January 2000 to December 2019
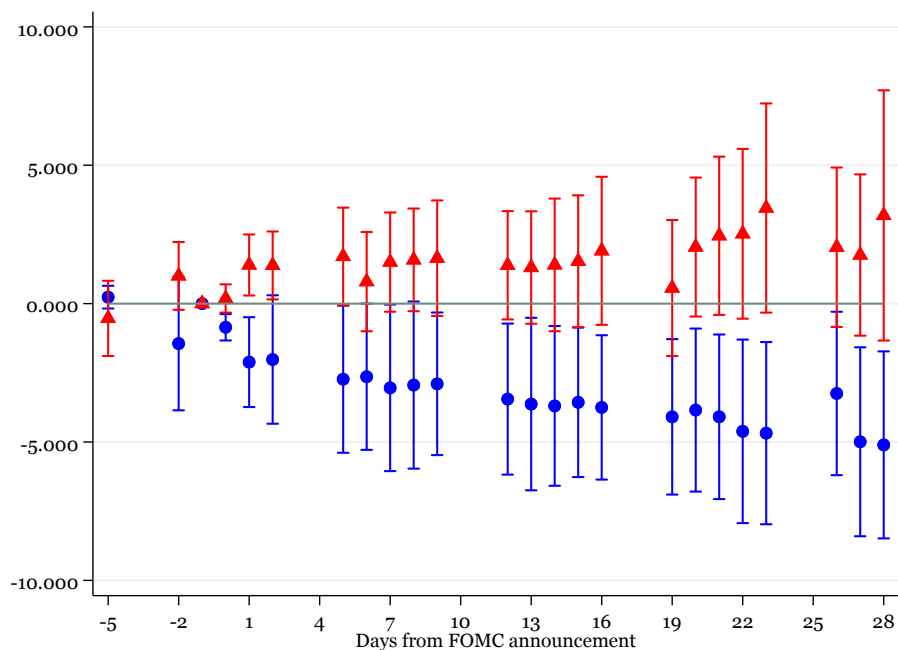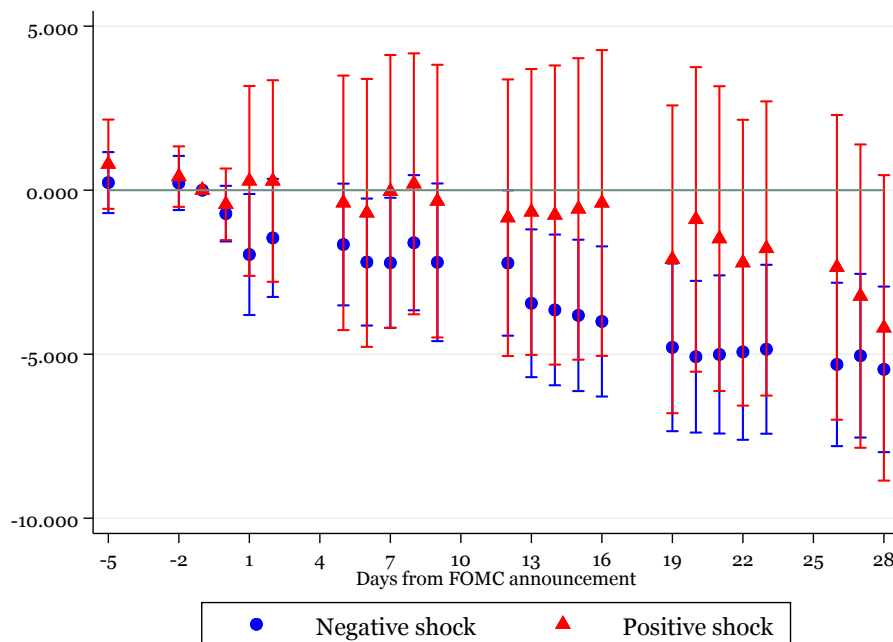
Panel B: January 2010 to December 2019

Notes: The figure reports the slope coefficient β_h on Nakamura and Steinsson (2018) shocks and 95%-confidence interval from (5) for corporate bond yields. The regression controls for issue-cusip-by-FOMC-event fixed effects as well as borrower's characteristics. Standard errors are clustered at year-month level. Data span from January 2000 to December 2019.

75

Fig. A.22. **Decomposition of changes in 10-year swap rates in future expected short rates and term premium**

Notes: The figure shows the decomposition of changes in 10-year swap rates into Adrian et al. (2013) future expected short rates and term premium and a residual component.

Fig. A.23. **Banking industry stock returns on the Adrian et al. (2013) expected future rates and term premia**

Panel A: β_{EI} estimates from $R_{jt} = \alpha + \beta_{\text{EI}}\Delta\text{EI}_t + \beta_\phi\Delta\phi_t^{10} + \gamma_j\Delta\text{FF}_t + \epsilon_{jt}$ – Excluding the 3 QE1 scheduled FOMC dates, i,e., December 16, 2008, January 28, 2009 and March 18, 2009.
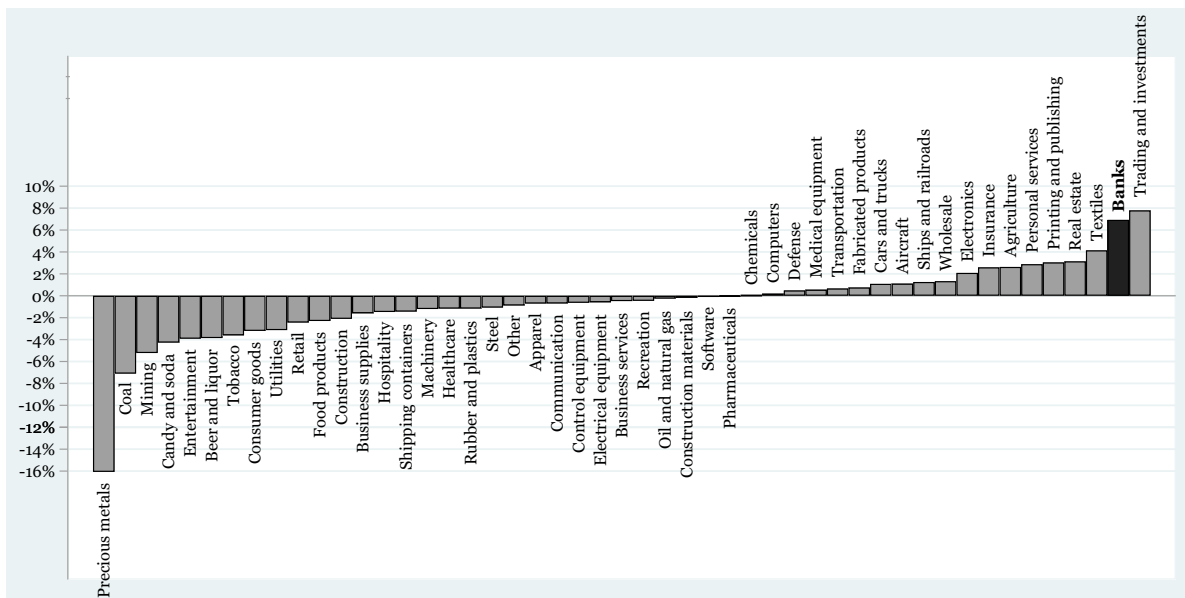

Panel B: β_ϕ estimates from $R_{jt} = \alpha + \beta_{\text{EI}}\Delta\text{EI}_t + \beta_\phi\Delta\phi_t^{10} + \gamma_j\Delta\text{FF}_t + \epsilon_{jt}$ – Excluding the 3 QE1 scheduled FOMC dates, i,e., December 16, 2008, January 28, 2009 and March 18, 2009.
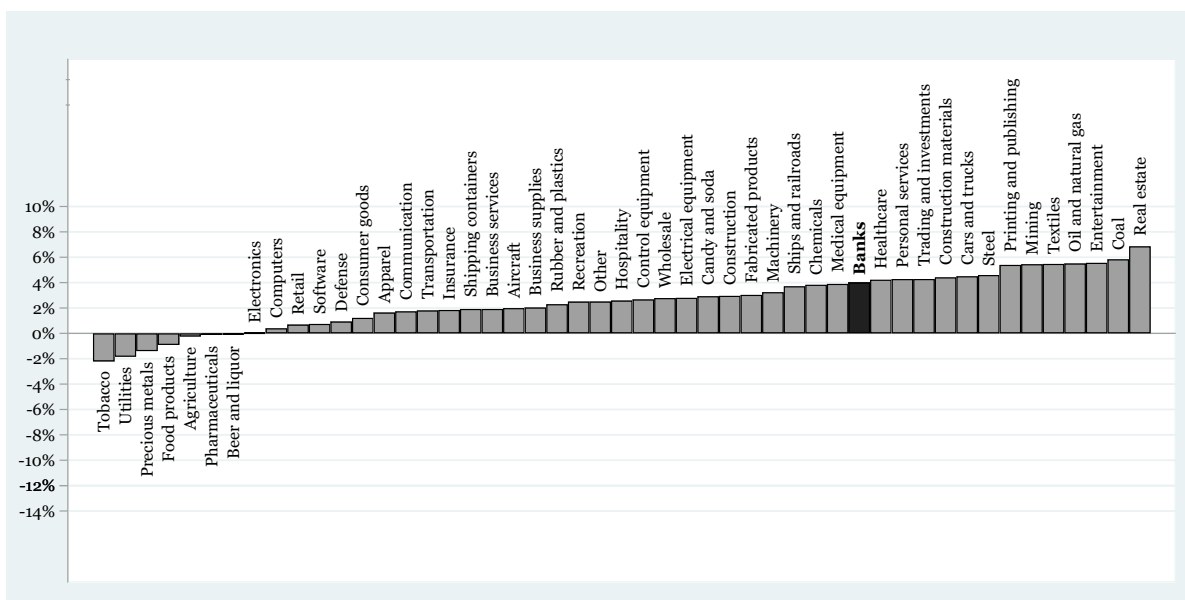

Notes: The figure shows the sensitivity of industry stock portfolios to changes in expected future rates and term premia by Adrian et al. (2013) on FOMC days. Industry data are the returns of the Fama-French 49 industry portfolios, downloaded from Ken French's website. All regressions control for Kuttner (2001) federal-funds shocks. Panel A shows the results for β_{EI}, whereas Panel B for β_ϕ. Values are expressed as the drop in the industry portfolio for every 1% unexpected positive shock in 10-year swap rates.

77

Fig. A.24. **Banking industry stock returns on the 10-year interest swap rate changes.**
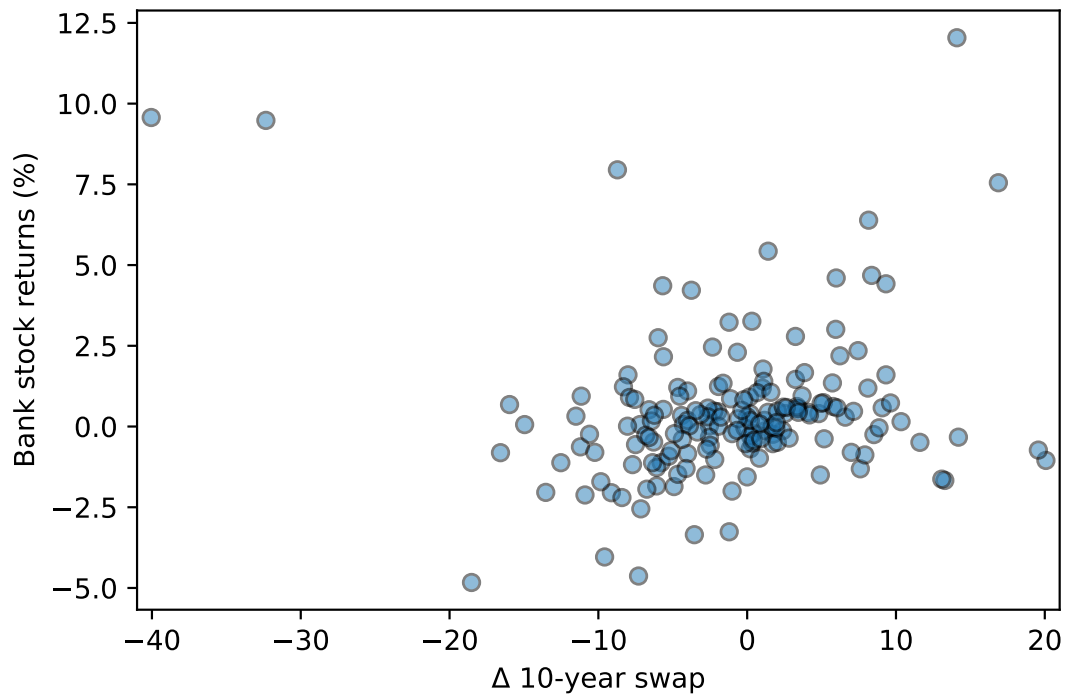
Notes: The figure shows the scatterplot of the banking industry stock returns on FOMC days against changes in the 10-year swap rates for the same dates. Industry data are the returns of the Fama-French 49 industry portfolios, downloaded from Ken French's website. Values for Δ 10-year swap rates are expressed in basis points, while bank stock returns are in percentage. The sample includes all scheduled FOMC meetings from February 2, 2000 to June 16, 2021.

Fig. A.25. **Scatterplot of stock returns of 3 individual banks – Excluding the 3 QE1 scheduled FOMC dates, i,e., December 16, 2008, January 28, 2009 and March 18, 2009.**
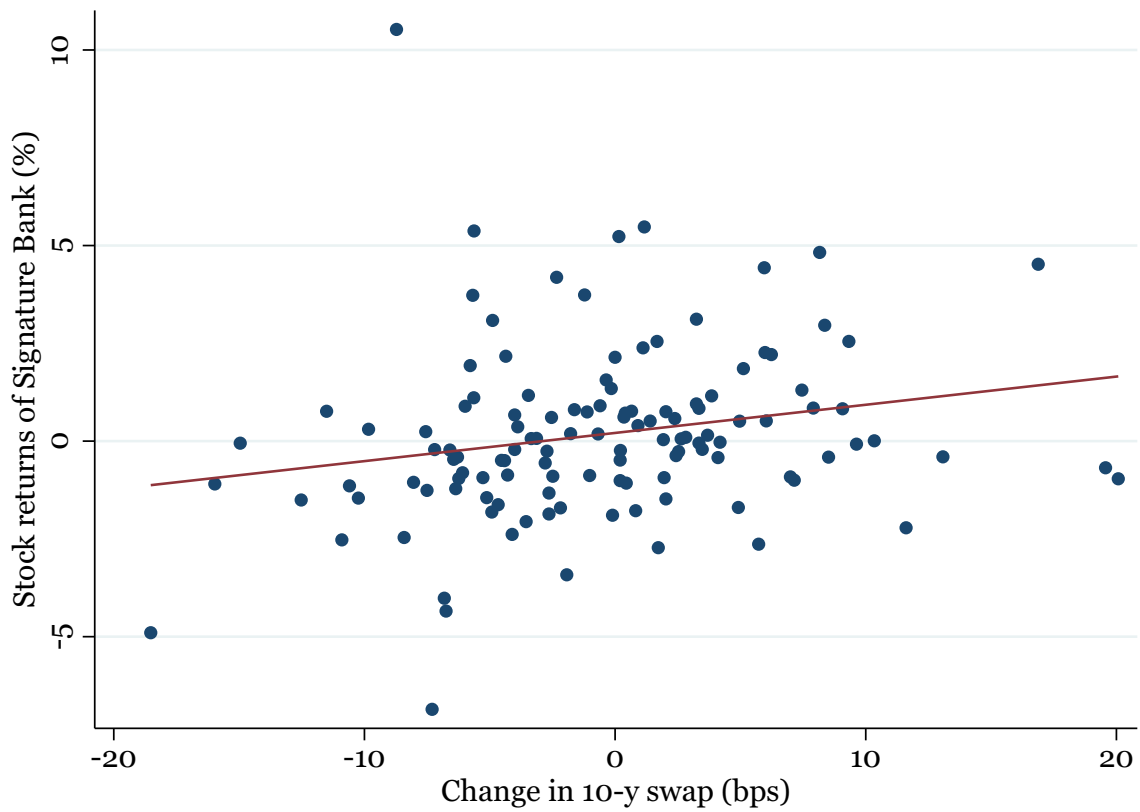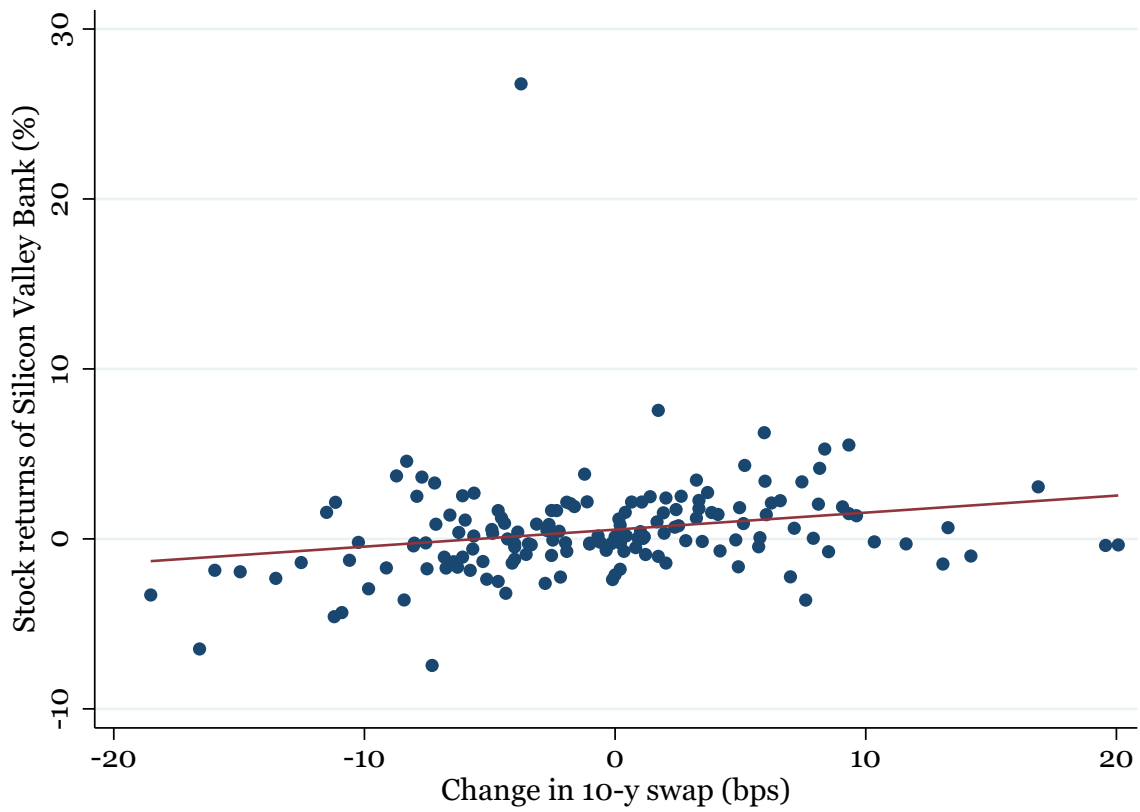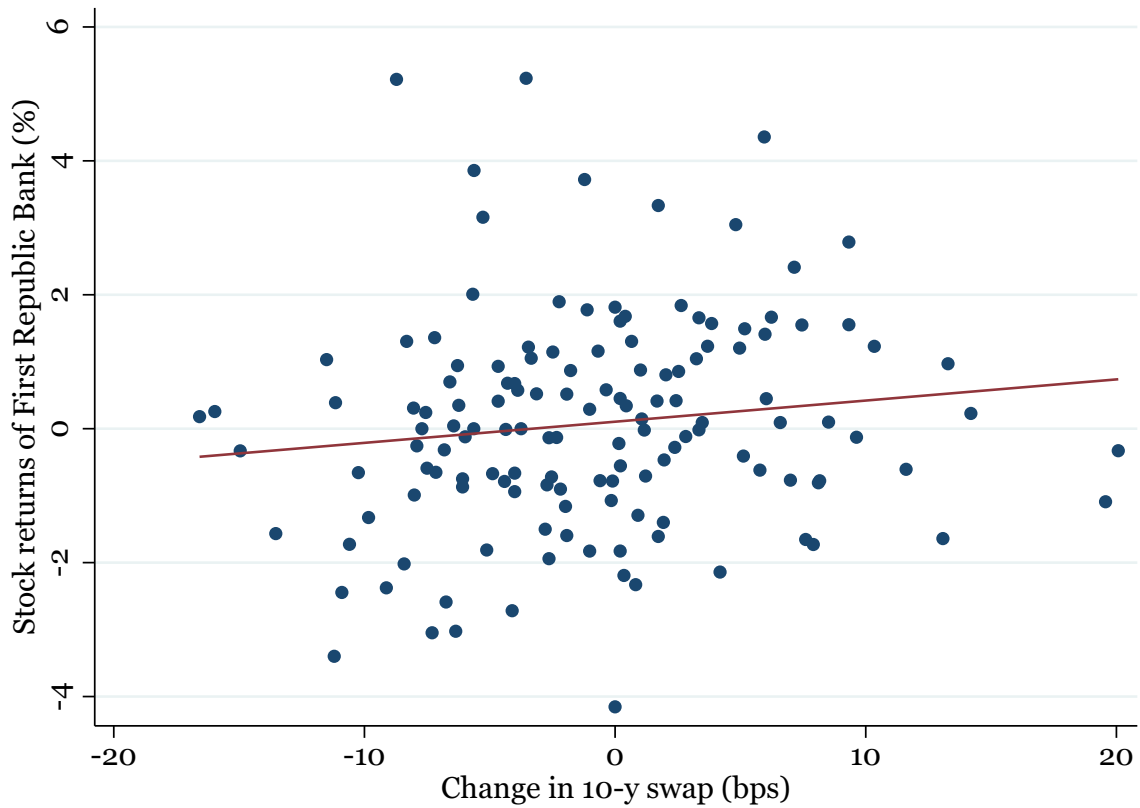
Notes: The figure shows the scatterplot of the stock returns 3 individual banks on FOMC days against changes in the 10-year swap rates for the same dates. Values for Δ 10-year swap rates are expressed in basis points, while bank stock returns are in percentage. The sample includes all scheduled FOMC meetings from February 2, 2000 to December, 2019.