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October 7, 2026

WP 2026-20

<https://doi.org/10.21033/wp-2026-20>



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Monetary-Policy Risk and Equilibrium Asset Prices

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Abstract

We quantify the effects of uncertainty about the conduct of monetary policy on asset prices. Nominal yields, real yields, stock returns, and implied volatilities all vary more on days with FOMC communication than on FOMC “blackout” days, consistent with policy communication being a source of risk. Furthermore, these asset prices all respond significantly to changes in interest-rate uncertainty around FOMC announcements, consistent with that risk being priced. A New Keynesian model with stochastic variation in the policy rule quantitatively matches this evidence. Our model shows that monetary-policy risk contributes to excess returns on equities and nominal bonds and accounts for much of the risk premium on inflation-indexed bonds. Transitory increases in monetary-policy risk lower stock and bond prices, reduce inflation-risk premia, and generate macroeconomic weakness similar to that produced by aggregate-demand shocks. The effective lower bound on interest rates reduces the size of monetary-policy risk in the model but magnifies its asset-pricing consequences, consistent with the data. We use our model to assess the effects of the increase in monetary-policy risk that occurred following the rapid tightening of policy in 2022.

JEL Codes: E43, E44, E52, E58, G12

Keywords: Monetary-policy risk; monetary-policy uncertainty; asset pricing; risk premia; interest-rate volatility; monetary-policy communication

*Corresponding author, thomas.king@chi.frb.gov. For helpful comments and discussions, we thank Gadi Barlevy, Brent Bundick, Stefania D’Amico, Anthony Diercks, Francois Gourio, Carolin Pflueger, and seminar participants at the Federal Reserve Bank of Chicago. Albert Chen provided magnificent research assistance. The views expressed here do not reflect official positions of the Federal Reserve.

1 Introduction

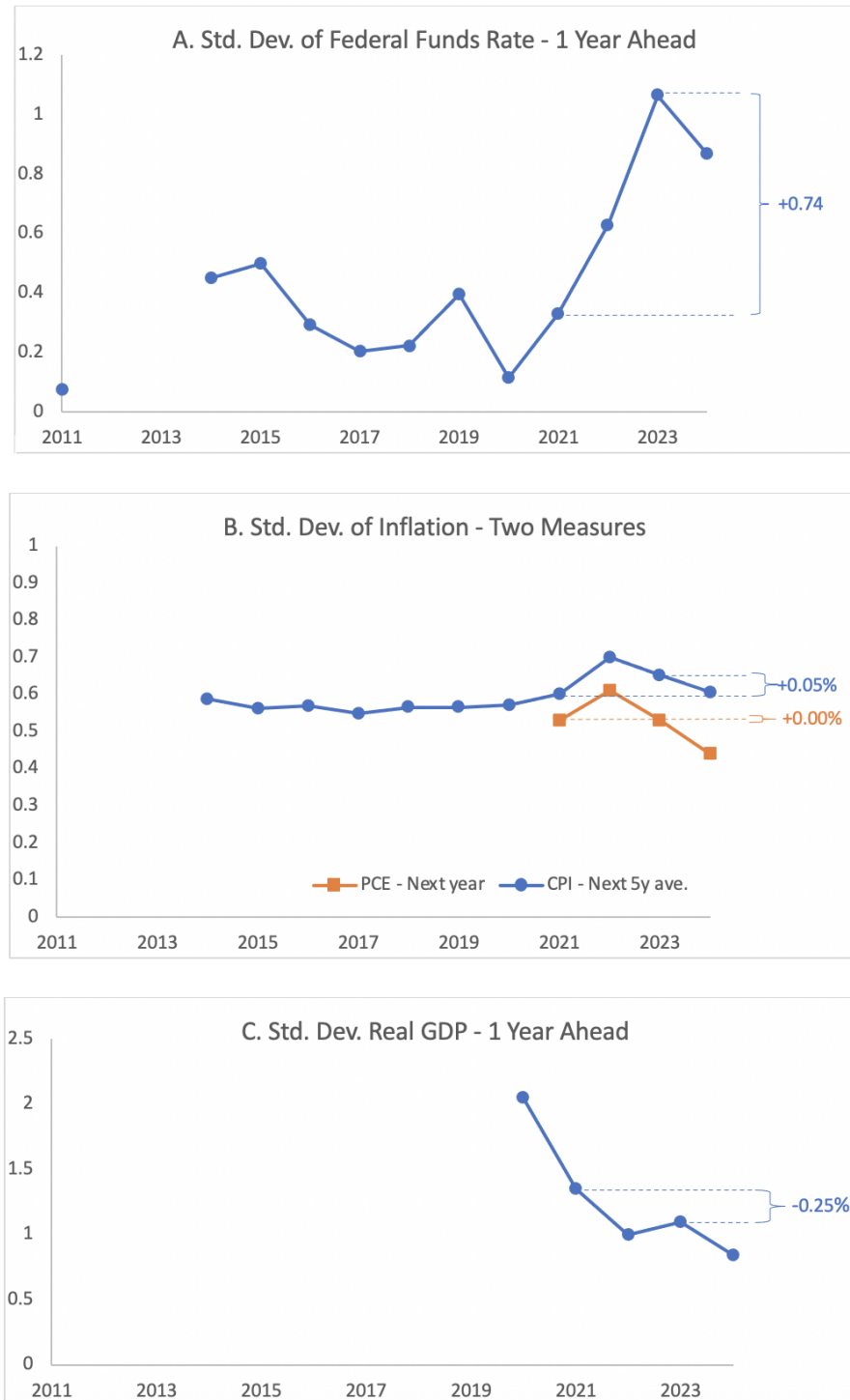
Monetary policy moves markets. See Rigobon and Sack (2004), Bernanke and Kuttner (2005), Gertler and Karadi (2015), and Bauer and Swanson (2023), among many others. When investors become unsure about how monetary policy will be conducted, therefore, we would expect them to incorporate that uncertainty, and any effects it might have on the macroeconomy, into the expected cash flows and discount rates of financial assets. In this paper, we study the implications of such uncertainty about the central bank reaction function—what we call “monetary-policy risk”—for the pricing of equities, the term structures of nominal and real interest rates, and their volatilities.

To take a concrete example, monetary-policy risk (MPR) appears to have risen dramatically during the post-COVID tightening of policy in the United States. Panel A of Figure 1 plots the standard deviation of the one-year-ahead federal funds rate calculated from subjective probability distributions reported by large dealer banks.¹ Between year-end 2021 and year-end 2023, this measure of uncertainty about the path of monetary policy more than tripled. Meanwhile, Panels B and C show that uncertainty about short- and medium-term inflation was essentially unchanged between year-end 2021 and 2023, while uncertainty about one-year-ahead GDP declined. Thus, the increase in uncertainty about interest rates was not a mechanical passthrough of uncertainty about the macroeconomy via the monetary-policy reaction function; rather, it reflected an increase in uncertainty about the course of monetary policy *conditional* on the state of the economy.²

¹Federal Reserve Bank of New York, Survey of Primary Dealers. Data available at <https://www.newyorkfed.org/markets>. Dealers report probabilities of outcomes within bins, and we calculated the standard deviation using the midpoint of each bin, averaging the responses across dealers. Collection of these data has been inconsistent over the years, resulting in the missing observations of the fed funds rate uncertainty in 2012 and 2013 and shorter series lengths for inflation and GDP.

²Contemporary commentary was consistent with this interpretation. See, for example, Greg Ip, “Fed’s Rate Increases Defy all the Rules,” *The Wall Street Journal*, Oct. 5, 2022. In principle, a perceived increase in the magnitude of reaction-function coefficients, such as that estimated by Bauer et al. (2024a), could also raise interest-rate uncertainty without an increase in macro uncertainty. However, we show in Section 6 that this effect cannot account for the magnitudes shown in the figure.

Figure 1: **Uncertainty Measures from December Survey of Primary Dealers**



Notes: The figures show uncertainty about the federal funds rate, inflation, and GDP at various horizons, as computed from the New York Fed's Survey of Primary Dealers. Standard deviations are calculated by multiplying the midpoint of each bin for which dealers report probabilities by their assigned probabilities. We then average the responses across dealers. Missing data reflect years in which the survey did not collect data for particular variables.

We show empirically and theoretically how shocks to MPR like this one can be important drivers of financial returns. Heuristically, there are three channels through which we should expect MPR shocks to pass through to asset prices. First, because they increase volatility and downside risk, such shocks lead to higher risk premia across assets. Second, because they induce macroeconomic weakness through precautionary behavior (and perhaps through other mechanisms), such shocks lower equity prices by reducing expected cash flows, and raise nominal bond prices by lowering the compensation required for expected inflation. Third, if monetary policy itself responds to the macroeconomic deterioration, positive MPR shocks may be associated with a lower expected path of risk-free rates, and the resulting discounting effects may offset part of the declines in stock and bond prices. The jump in MPR in 2022 indeed coincided with increases in estimates of real and nominal term premia, a narrowing of Treasury-based inflation compensation, and a significant decline in the stock market.³ Although the FOMC was raising the federal funds rate during this period to fight inflation, it likely did so by less than it would have in the absence of elevated MPR because the increase in MPR was itself exerting some restraint on the economy.

Over longer periods, we quantify how MPR contributes to the general levels of risk premia on stocks and bonds, helping to address long-standing questions about why the excess returns on those assets are so large (Mehra and Prescott (1985), Backus et al. (1989)). The intuition is straightforward: shocks to monetary policy move the payoffs on both stocks and bonds in the same direction as real activity. More uncertainty about such shocks—i.e., a higher level of MPR—therefore increases the likelihood that these assets will experience bad returns in bad states of the world, and risk-averse investors require greater compensation if that risk is high. Of particular note, because monetary-policy shocks are transitory, MPR can also explain the positive average premium on inflation-indexed bonds, a feature of the data that consumption-based asset-pricing models have struggled to match (Hsu et al. (2021)).

We begin by quantifying the size and behavior of MPR in the data since 1994. Our measurement relies on two separate analyses. First, we gauge the size of MPR by comparing asset-price volatility on days with and without FOMC communication. Specifically, during FOMC “black-

³Cieslak et al. (2026) also tie changes in the nominal term premium over this period to “Fed-induced uncertainty” about the course of policy.

out periods,” Fed officials and staff refrain from speaking to the public about matters having to do with monetary policy. Thus, effectively, there is no monetary-policy risk on these days. We show that asset prices are uniformly more volatile on non-blackout days, implying that MPR adds a nontrivial amount to risk across markets. For example, based on the differences in volatility between blackout and non-blackout days, we estimate that MPR contributes about 1.7 percentage points (compared to a total of 8.5%) to the quarterly standard deviation of stock prices and 12 basis points (compared to 41 total) to the quarterly standard deviation of the 10-year nominal yield. Second, we expand the event-study evidence of Bauer et al. (2022), who computed the responses of various financial variables to changes in short-term interest-rate uncertainty on the days of FOMC announcements. This exercise identifies the passthrough of changes in MPR to stock prices, real and nominal bond yields, and the implied volatilities of equities and interest rates. We confirm the Bauer et al. result that these multipliers are economically large and statistically significant.

We make quantitative sense of these results through the lens of a nonlinear New Keynesian (NK) asset-pricing model. We allow for MPR in the model by adding persistent shocks with stochastic volatility to the monetary-policy reaction function. The shock volatility captures uncertainty about random deviations from the policy rule (i.e., uncertainty about standard NK “monetary-policy shocks”) and also approximates to a first order uncertainty about transitory changes in the parameters of the rule itself. We calibrate the size and persistence of the shocks, and the size and persistence of their volatility, to match the blackout-period and event-study evidence. Notably, the model can quantitatively account for the empirical responses of asset prices to standard monetary-policy shocks as well as to MPR shocks.

With parameters that thus match the data, the level and variation in MPR in the model contributes about 40 basis points to the average nominal ten-year term premium and about 120 basis points to the average equity premium. It implies an average real ten-year term premium of 71 basis points, roughly matching the average level in the data. Near the steady-state, a transitory increase in MPR generates the following outcomes: (1) implied and realized volatilities rise across asset classes; (2) output and inflation fall due to precautionary behavior; (3) the equity premium rises, which, together with the lower expected nominal output, causes stock

prices to fall; (4) real and nominal term premia rise, and long-term bonds realize negative excess returns; (5) at shorter horizons, the nominal term-premium increase is smaller than the decline in expected policy interest rates, so the nominal yield curve steepens; (6) the inflation risk premium falls.

When we allow for an effective lower bound (ELB) in the model, the impact of MPR on interest-rate volatility is damped, so shocks to MPR generally have smaller effects on risk premia. However, for a given increase in volatility, the effects on realized stock and real-bond returns are larger at the ELB because the inability of nominal interest rates to respond amplifies the drag on output and inflation. To test these predictions in the data, we return to the event-study, dividing the sample into observations when policy was constrained by the ELB and observations when it was not. We show that, as the model predicts, the relationships between short-rate uncertainty and asset prices are larger at the ELB.

Finally, as a case study, we use the model to examine the rise in MPR that occurred around the rapid policy tightening of 2022. Introducing shocks to replicate the increase in interest-rate uncertainty shown in Figure 1, and accounting for the effects of liftoff from the ELB in 2022, the model implies that shocks to MPR added about 35 bp to the 10-year real term premium, caused the 10-year inflation-risk premium to fall by 12 bp, pushed up the VIX by about 5 points, and caused a cumulative drop in equity prices of about 4 percent during this period. Various sources of measurement error imply that these estimates are likely lower bounds on the actual effects. The model also suggests that the FOMC tightened monetary policy less aggressively than it would have in the absence of MPR shocks; this extra accommodation avoided significantly greater losses in the stock and bond markets.

In addition to clarifying how uncertainty about monetary policy can help to explain risk premia and volatility in financial markets, our results also underscore the importance of clarity in monetary-policy communication. When MPR is high and variable, it can weigh on asset prices, cause financial conditions to tighten, and obscure the signals about the economic trajectory that financial markets can provide. While the central bank may in principle be able to offset the effects of MPR on the economy by adjusting the policy rate in the opposite direction, such adjustments are likely to be difficult to calibrate in real time and risk confusing the public

further.

Relationship to the literature. Our paper connects to three broad strands of literature. First, recent work by Husted et al. (2020), Bauer et al. (2022), and Bundick et al. (2022) has shown that uncertainty about interest rates and monetary policy varies over the business cycle and even over the FOMC cycle.⁴ These empirical measures of uncertainty capture the *total* variance of policy rates rather than isolating the exogenous component of policy that we are interested in, but it is likely that some portion of the fluctuations documented in these studies is due to changes in the public’s uncertainty about the exogenous component. Ochoa and Wiegand (2026) decompose interest-rate volatility into uncertainty about the macroeconomy and uncertainty stemming from the policy reaction function and find that the latter has effects on asset prices broadly similar to what our model predicts. Several other recent empirical studies, including Cieslak and Schrimpf (2019), Cieslak and Pang (2021), Kroencke et al. (2021), Boehm and Kroener (2024), and Bauer et al. (2026), provide evidence that much of the movement in financial markets around FOMC communications cannot be ascribed to changes in the expected path of interest rates alone and argue that it must instead be accounted for by changes in perceptions of risk or risk appetite. Our model provides one mechanism that can help to make sense of these results—policy communications involve changes in uncertainty about the reaction function that have significant effects on asset prices.

Second, our results are related to an expansive empirical and theoretical literature that attempts to characterize the effects of uncertainty on business cycles, including Bloom (2009), Gilchrist et al. (2014), Jurado et al. (2015), Basu and Bundick (2017), Bloom et al. (2018), and many others. Some of these papers consider macroeconomic or financial uncertainty in reduced form, rather than homing in on specific structural sources; those that are specific tend to focus on fiscal or supply-side sources, rather than on monetary policy. One exception is Mumtaz and Zanetti (2013), who embed stochastic volatility in the policy reaction function in a structural model in a way that is similar to ours. However, like most of this literature, they do not examine asset-pricing implications. While uncertainty shocks of different structural types tend to have

⁴Tilmann (2020) and de Pooter et al. (2021) show that the levels of interest-rate uncertainty measured like this affect the impact of monetary-policy shocks themselves, a phenomenon we do not explore here.

similar macroeconomic effects in theoretical models, they can have much different consequences for asset prices, as we discuss below.

The third set of papers we build on are those that seek to understand asset prices in the context of general-equilibrium models, and in sticky-price frameworks in particular. Papers such as Hordahl et al. (2008), Bekaert et al. (2010), Li and Palomino (2014), Rudebusch and Swanson (2012), van Binsbergen et al. (2012), Dew-Becker (2014), Kung (2015), and Swanson (2021) have closed much of the gap between asset pricing and modern macroeconomic theory. Yet, in all of these models, monetary-policy shocks are treated as homoskedastic or are absent altogether. Within this literature, the papers that are perhaps methodologically closest to ours are Gourio and Ngo (2020) and Pflueger (2025). Gourio and Ngo show how the proximity of the ELB endogenously lowers term premia and changes the correlation of stock returns and inflation from negative to positive in the presence of productivity shocks. Pflueger identifies changes in the monetary policy rule, in conjunction with a declining role for technology shocks, as the reason for the decline in the stock-bond correlation over time. Both papers price assets in a New Keynesian (NK) framework and attempt to explain changes in asset-price behavior through changes in the mix of structural risks, as we do. However, neither focuses on monetary policy as a source of risk.

Outline. The paper proceeds as follows. In Section 2, we present our empirical evidence measuring differences in asset-price volatility across FOMC blackout and non-blackout periods and responses to MPR shocks on FOMC announcement days. In Section 3, we explain our basic intuition by showing how risk premia on stocks and bonds should be related to MPR in a two-period linear-Gaussian economy, where simple closed-form solutions are available, and how these relationships compare to the effects of other kinds of risk. In Section 4 we develop our quantitative, nonlinear NK model and explain the solution method and parameterization. Section 5 presents the baseline results, calibrated to our empirical evidence. Section 6 discusses our case study of the experience of 2021 - 2023, and Section 7 concludes.

2 Quantifying Monetary-Policy Risk

2.1 A Simple Framework

We define monetary-policy risk as uncertainty about the policy rate, conditional on the state of the economy. To fix ideas, let i_t be the interest rate in period t and $\mathbf{x}_t$ be a vector representing the economic state. Suppose the joint evolution of monetary policy and the economy is described by the system

$$i_t = \phi \mathbf{x}_t + v_t \tag{1}$$

$$\mathbf{x}_t = \mathbf{k} v_t + \mathbf{A} \epsilon_{x,t} \tag{2}$$

where v_t and $\epsilon_{x,t}$ are mean-zero stochastic processes that are exogenous with respect to $\mathbf{x}_t$ and i_t at all leads and lags. For expositional simplicity, we suppress intercept terms, abstract from possible inertia in the policy rule, and ignore the ELB on nominal interest rates. These features, which we will include in our quantitative model below, do not affect the qualitative points we wish to make in this section. In this framework, MPR at horizon h is given by

$$\text{std}_t[i_{t+h}|\mathbf{x}_{t+h}] = \text{std}_t[v_{t+h}] \tag{3}$$

where std_t indicates the standard deviation conditional on information at time t .

Our assumption that interest rates embed an exogenous stochastic component v_t follows extensive theoretical and empirical literatures that motivate such a term in a variety of ways.⁵ The simplest case is the possibility of random departures from a central bank’s fixed policy rule, as in the standard monetary-policy shocks that appear in NK models such as Rotemberg and Woodford (1997). Such shocks are isomorphic to transitory changes in the intercept of the rule—specifically, to possible variation in the central bank’s inflation target. They may also be thought of as approximating, to a first order, changes in the other parameters of the rule. For

⁵Here and below, we associate the interest rate with a short-term nominal bond yield, rather than the overnight interest rate, to allow expectations about future monetary policy to matter. v_t may thus be thought of as reflecting central-bank forward guidance about the near-term path of policy as well as actual policy surprises in overnight rates.

example, one interpretation of the events of 2022 was that the public became less sure about how aggressively the Fed would respond to inflation.⁶ Another possibility is that monetary-policy “shocks” may reflect differences between market and central-bank beliefs about the economy, as in Caballero and Simsek (2022). If so, then MPR may rise during periods of higher disagreement about macroeconomic fundamentals.

We remain agnostic about the microfoundations of v_t , but we assume that this component of the policy rule is stationary and potentially heteroskedastic. To be concrete, we specify

$$v_t = \rho_v v_{t-1} + \zeta_{t-1} \epsilon_{v,t} \quad (4)$$

where $\epsilon_{v,t} \sim Niid(0, 1)$, and

$$\zeta_t = (1 - \rho_\zeta) \bar{\zeta} + \rho_\zeta \zeta_{t-1} + \eta_t \quad (5)$$

where η_t is iid and independent of $\epsilon_{v,t}$ with $\text{var}[\eta_t] = \omega^2$. With this structure, in period t , MPR at horizon $h = 1$ is simply equal to ζ_t , and MPR at all other horizons is proportional to this value.⁷

Let p_t be the log price of a financial asset. Assume that all such asset prices are functions of the economic and policy state variables:

$$p_t = f(v_t, \epsilon_{x,t}, \zeta_t) \quad (6)$$

Within this framework, we ask two questions. First, what is the overall contribution of MPR and its volatility to the variation of asset prices, i.e., what is the value of $\text{var}[p_t | \epsilon_{x,t} = 0]$? Second, what is the reaction of asset prices to a shock to MPR, $\partial f / \partial \zeta_t$?

⁶Bauer et al. (2024b) and Bocola et al. (2024) have recently presented evidence of medium-frequency shifts in coefficients of the FOMC’s reaction function that could motivate such uncertainty. Expanding the model above to allow for stochastic parameters ϕ_t in a linear approximation around the steady state produces $\text{var}_t[i_{t+h} | x_{t+h}] = \text{var}_t[\phi_{t+h}] + \text{var}_t[v_{t+h}]$. Thus, an increase in MPR as we have defined it is observationally equivalent, to a first-order, to an increase in uncertainty about ϕ .

⁷For the moment, we ignore the positivity constraint on ζ_t , assuming that $\bar{\zeta}$ is far enough above zero that the process is essentially linear; this assumption will be relaxed in Section 4.

2.2 Measuring the Contribution of MPR to Return Variation

To measure the overall contribution of MPR to asset-price volatility, we compare days when monetary-policy and MPR shocks are shut down (i.e., when $\epsilon_{v,t} = \eta_t = 0$) to days when these shocks are operative. Specifically, we exploit FOMC “blackout periods” to provide a measure of the background volatility in asset prices that is not related to news about monetary policy. During blackout periods, Fed officials and staff refrain from communication with the public about matters having to do with monetary policy. Thus, variation in asset prices during these periods must come entirely from news about shifting macroeconomic and financial conditions, rather than from news about the policy rule, possible deviations from the rule, or fluctuations in uncertainty about the rule. Comparing this volatility to the volatility on non-blackout days allows us to back out the contribution of monetary factors to the total variation of asset prices.⁸

Specifically, we can show

Proposition 1. In the economy described in equations (1) through (6), if ρ_ζ and ρ_v are close to 1, then

$$\underbrace{\text{var}[\Delta p_t | \epsilon_{x,t} = 0]}_{\text{contribution of MPR}} \approx \underbrace{\text{var}[\Delta p_t]}_{\text{non-blackout days}} - \underbrace{\text{var}[\Delta p_t | \epsilon_{v,t} = \eta_t = 0]}_{\text{blackout days}} \quad (7)$$

Proof. See Appendix A.1.

The quantity on the left-hand side of (7) is the measure we are interested in: the volatility of asset prices when all sources of risk other than monetary policy have been shut down. The assumption that ρ_ζ and ρ_m are close to 1 is likely to hold in the daily data we work with. Indeed, when we calibrate to match the data in Section 4, we find values of both of these parameters that are equivalent to more than 0.99 at a daily frequency. To the extent that the true values of ρ_ζ and ρ_m are less than one, our calculation will understate the size of MPR, so in this sense the assumption is conservative.

The timing of blackout dates has shifted a bit over the years. The FOMC appears to have adopted a voluntary quiet period around each meeting sometime in the early 1980s. It

⁸The Fed’s communications policy is available at https://www.federalreserve.gov/monetarypolicy/files/FOMC_ExtCommunicationParticipants.pdf. If, despite the policy, some communication about monetary policy still takes place during blackout dates, our measurement provides a lower bound on the contribution of MPR.

is discussed as an established practice, though only in the form of “informal guidelines,” in the transcripts of the January 1995 meeting, where participants debated its merits.⁹ At that point, the norm was that FOMC participants should refrain from policy-related communication from one week before each meeting to one week after. At the conclusion of the discussion at that meeting, Chairman Greenspan proposed, without objection, that the blackout period be established from one week prior to each FOMC meeting through “the balance of the week of the meeting.” The next change to the practice—and apparently the first official public communication about it—came in June 2011, when the FOMC formally adopted its Policy on External Communications of Committee Participants, which set the blackout period from one week prior to each FOMC meeting to one day after.¹⁰ Then, in February 2017, the period was extended three days earlier, bringing it to its current starting time of two Saturdays before each FOMC meeting.¹¹

We use this chronology to identify blackout and non-blackout days over the period January 1994 - May 2026. We drop observations from the crisis period between September 2008 and March 2009, as well as those from March 2020, which reflected the most acute stress associated with the Covid-19 pandemic. During these periods, markets were extremely volatile and sometimes dysfunctional, and the Fed took a number of extraordinary actions, some announced during blackout periods. Our main results are robust to including these dates, as shown in Appendix B. In addition, to reduce the influence of outliers, we trim the data at three standard deviations. The main effect of this restriction is to exclude several days after the failure of Silicon Valley Bank in March 2023, which exhibited extraordinarily large movements in bond yields (and happened to overlap with a blackout period).

We include FOMC announcement days in the “non-blackout” category, since they clearly involve important communication to the public. One complication is that the market reactions to the information revealed at the FOMC may spill over into the following day. The chair’s press conference, for example, often runs until 3:30 PM, just half an hour before U.S. stock markets

⁹January 1995 FOMC transcript, pp. 34 - 38, <https://www.federalreserve.gov/monetarypolicy/files/FOMC19950201meeting.pdf>

¹⁰https://fraser.stlouisfed.org/files/docs/historical/FOMC/FOMC_ExtCommunicationParticipants.pdf

¹¹Minutes of the January 2017 FOMC meeting, <https://www.federalreserve.gov/monetarypolicy/fomcminutes20170201.htm>.

close. If it takes investors longer to digest the incoming information, their reaction will not show up in pricing until the following morning. Indeed, in our data, the volatility of the 10-year yield and the S&P500 are both higher on the days *after* FOMC announcements than they are on the announcement days themselves. Because of these concerns, we err on the side of including the days after FOMC meetings in the non-blackout sample for our baseline measurement.¹² In Appendix B, we show that our results are robust to simply dropping the days after FOMC meetings altogether.

In the top panel of Table 1, we report the daily standard deviation of various asset prices and related financial variables on blackout and non-blackout days.¹³ When calibrating our model below, we rely on the comparison of blackout days to *all* non-blackout days, as suggested by equation (7), but in the table we also break out the response on FOMC and non-FOMC days separately for comparison. The calculations show that, across all assets, there is more volatility outside of blackout days than during blackout days. The differences are statistically significant with p values of at most 0.09 for all variables except the VIX. Although FOMC announcement days are particularly volatile, non-blackout days *other* than FOMC days also exhibit uniformly more asset-price volatility than blackout days do.

In the final column of the table we report the amount of volatility of each asset that is due to MPR. Specifically, using equation (7) we calculate $\sqrt{50.3 \times \text{var}[\Delta p_t | \epsilon_{x,t} = 0]}$, where the multiplier 50.3 scales the daily volatility to a quarterly frequency using the average number of non-blackout business days in each quarter (and approximating the asset prices as random walks over these short horizons). These results show that volatility associated with FOMC communications injects a substantial amount of risk into asset prices. For example, it contributes 1.7 percentage points to the quarterly standard deviation of stock prices and 12 basis points to the quarterly standard deviation of the 10-year nominal yield. These contributions compare to overall quarterly standard deviations of 8.5% and 41 bp, respectively. In terms of the stylized model above, they reflect both the effects of shocks to monetary policy ($\epsilon_{v,t}$) and shocks to MPR

¹²This choice follows previous papers, such as Hanson and Stein (2015) and de Pooter et al. (2021) that, for similar reasons, consider two-day windows when conducting event studies around FOMC meetings.

¹³Data are obtained from and used with permission of Bloomberg. We include data from Treasury Inflation-Protected Securities (TIPS) only beginning in 2003 because market liquidity and data reporting were imperfect prior to that time.

Table 1: **Comparison of Volatility on Blackout and Non-Blackout Days: Full Sample**

	Blackout	Non-blackout			Implied qtrly std. dev. due to MPR
		All non- blackout days	FOMC days	Non-FOMC days	
5Y nom. yield	5.2	5.4***	6.3***	5.4**	12.4
10Y nom. yield	5.0	5.3***	5.9***	5.3**	12.1
5Y real yield	4.6	5.0***	6.4***	4.9**	13.2
10Y real yield	4.2	4.5***	5.3***	4.4**	11.2
SPX	90	93*	104***	92	166
VIX	120	122	129*	121	156
MOVE	3.1	3.4***	4.1***	3.4***	9.9
Sample Size	1663	6189	496	5693	

Notes: The table compares the standard deviations of daily changes in various asset prices on days that fall during blackout periods to days that do not over the period 1994 - 2026, except for real yields, which begin in 2003. All values are reported in units of basis points. Data-cleaning restrictions are described in the text. Asterisks indicate statistically significant differences between blackout and non-blackout days at the 10% (*), 5% (**), and 1% (***) levels, based on one-sided bootstrapped confidence bounds. The final column shows the implied contribution of monetary-policy risk to the quarterly standard deviation of each asset, calculated by scaling the variances up by the average number of non-blackout business days per quarter.

itself (η_t).

In our model, MPR will behave differently when policy rates are constrained by the ELB. Therefore, we also break out the empirical results into periods at the ELB and away from the ELB, as shown in Table 2. Likely reflecting the compression of the distribution of interest rates in this state of the world, the differences between volatility on blackout and non-blackout days are generally smaller and less statistically significant in the ELB sample, and the implied quarterly contributions of MPR are thus also smaller. The notable exception is the 10-year nominal yield, which may have been disproportionately affected by quantitative-easing announcements during this sub-sample. When we do our calibration, we focus on the moments of the model near the steady state and match to the results above the ELB, although our model will result in volatility compression in ELB states, as these data suggest.

2.3 Measuring the Effects of MPR Shocks

Our second empirical question is: what is the effect of a shock to MPR, η_t , on asset prices? Here, we follow Bauer et al. (2022), who conducted event studies to measure the reaction of

Table 2: **Comparison of Volatility: Split at Effective Lower Bound***A. Short rate above ELB*

	Blackout	Non-blackout			Implied qtrly std. dev. due to MPR
		All non- blackout days	FOMC days	Non-FOMC days	
5Y nom. yield	5.3	5.7***	6.5***	5.6**	13.7
10Y nom. yield	5.1	5.3**	5.8***	5.3*	11.0
5Y real yield	4.6	5.1***	6.1***	5.0***	15.6
10Y real yield	4.1	4.5***	5.3***	4.4**	13.1
SPX	89	92*	101***	91*	165
VIX	114	117	118	117	187
MOVE	3.2	3.5***	4.1***	3.5***	10.7
Sample Size	796	3741	185	3556	

B. Short rate at ELB

	Blackout	Non-blackout			Implied qtrly std. dev. due to MPR
		All non- blackout days	FOMC days	Non-FOMC days	
5Y nom. yield	4.7	4.8	5.9***	4.7	8.2
10Y nom. yield	4.8	5.3**	6.0***	5.2**	14.6
5Y real yield	4.7	4.9	6.8***	4.7	8.8
10Y real yield	4.3	4.4	5.5***	4.3	7.8
SPX	94	95	111**	93	98
VIX	136	135	156*	133	—
MOVE	2.9	3.1	3.9***	3.0	7.3
Sample Size	359	1526	73	1543	

Notes: The table compares the standard deviations of daily changes in various asset prices on days that fall during blackout periods to days that do not over the period 1994–2026, except for real yields, which begin in 2003. We split the sample into observations where the federal funds rate was above its effective lower bound (panel A) and observations where it was constrained by the ELB (panel B). All values are reported in units of basis points. Data-cleaning restrictions are described in the text. Asterisks indicate statistically significant differences between blackout and non-blackout days at the 10% (*), 5% (**), and 1% (***) levels, based on one-sided bootstrapped confidence bounds. The final column shows the implied contribution of monetary-policy risk to the quarterly standard deviation of each asset, calculated by scaling the variances up by the average number of non-blackout business days per quarter.

financial-market variables to changes in near-term interest rates and interest-rate uncertainty on days of FOMC announcements, over the period 1994 – 2019. Under the assumption that macroeconomic news is negligible or uncorrelated with monetary-policy news on these days, regression coefficients of asset prices on measures of the policy stance and its uncertainty identify the effects of these variables on the asset price itself. This is because, in our notation, if there

are no $\epsilon_{x,t}$ shocks, equations (1), (4), and (5) imply

$$\Delta i_t \approx \zeta_{t-1} \epsilon_{v,t} \quad (8)$$

$$\Delta \text{var}_t[i_{t+1}] \approx \eta_t \quad (9)$$

where, again, the equalities are exact when there is no predictability in returns (i.e., $\rho_\zeta = 1$ and $\rho_m = 1$ —a close approximation at the daily frequency). Thus, a regression of Δp_t on Δi_t and $\Delta \text{var}_t[i_{t+1}]$, if feasible, would recover the values of $\partial f / \partial v_t$ and $\partial f / \partial \zeta_t$ local to the sample mean.

Following the other recent event-study literature, Bauer et al. (2022) use one-year interest rates, rather than the overnight federal funds rate, to capture the policy stance. This is appropriate in an environment where strong policy signals are sent in advance of movements in the actual policy rate, as was the case for most of the period covered by the data. As the analogue of $\text{var}_t[i_{t+1}]$, they use the first principal component of implied volatilities of short-term interest rates up to one year ahead—which they refer to as “short-rate uncertainty”—as calculated from eurodollar options. The factor is scaled such that a change of 0.01 implies a 1 percentage-point change in the one-year forward volatility. We follow the same approach.¹⁴ We also expand the regressions to examine some variables that Bauer et al. did not consider. Specifically, we include the five-year TIPS yield and the MOVE index of implied Treasury-yield volatility, which we collect from Bloomberg. All data cover the same time period as in the blackout exercise above.

The results for the full sample appear in Table 3. With respect to conventional monetary-policy shocks, we confirm other studies, such as Bernanke and Kuttner (2005), Hanson and Stein (2015), Gertler and Karadi (2015), Bauer et al. (2022), and Bauer and Swanson (2023) that document strong responses to such shocks across both equities and bonds, and we show that these results continue hold quantitatively even when changes in MPR are taken into account. For example, a 100 bp increase in the one-year policy rate raises the 5-year nominal yield by 43 bp and causes a 2.8% decline in the stock market. On the other hand, monetary-policy shocks do not have significant effects on implied volatilities of equities or interest rates.

¹⁴We obtain the Bauer et al. data from Michael Bauer’s website updated through 2026. <https://www.michaeldbauer.com>.

Table 3: **Responses around FOMC Announcements: Pooled Sample**

	5Y nom. yield	10Y nom. yield	5Y real yield	10Y real yield	SPX	VIX	MOVE
Monetary-policy shock	0.43*** (0.07)	0.21*** (0.06)	0.59*** (0.10)	0.39*** (0.08)	-2.80** (1.22)	1.66 (1.86)	-0.08* (0.04)
Short-rate uncertainty (MPR)	0.64*** (0.16)	0.61*** (0.14)	1.00*** (0.23)	0.83*** (0.19)	-7.16** (2.88)	20.8*** (4.37)	0.92*** (0.10)
Adj. R^2	0.26	0.16	0.28	0.23	0.06	0.10	0.27
# obs.	239	239	165	165	239	239	238

Notes: Daily data on days of FOMC announcements, 1994 – 2026, except for real yields, which begin in 2003. We exclude observations between September 2008 and March 2009, as well as March 2020. Monetary-policy shocks are measured as changes in one-year eurodollar rates and MPR shocks are measured as changes in one-year federal-funds implied volatility. All values are reported in basis points. Standard errors in parentheses. Asterisks indicate statistical significance at the 10% (*), 5% (**), and 1% (***) level.

Meanwhile, similarly to the results in Bauer et al. (2022), an increase in MPR, controlling for standard shocks to the level of the monetary-policy stance, is associated with a statistically significant fall in stock prices, increases in both real and nominal bond yields, and increases in the implied volatilities of both equities and longer-term Treasuries. Inflation compensation—the difference between the response of nominal and real yields—also declines in response to such shocks at both the five- and ten-year horizon. The magnitudes are economically large. For example, an increase in MPR that causes short-rate uncertainty to rise by 10 bp at the one-year horizon results in an 6.1 bp increase in the ten-year yield and a 0.72% decline in the S&P500. The fact that both the coefficients on TIPS yields are large and positive is strongly suggestive of an increase in real term premia in response to MPR shocks, since it is unlikely that monetary policy would be expected to tighten in response to these shocks and investors would have to expect such a tightening to be very persistent to explain a large rise in the 5-year-ahead forward real rate.¹⁵

We also extend the Bauer et al. (2022) analysis by examining differences across ELB and non-ELB periods. Specifically, we interact the monetary-policy and short-rate-uncertainty shocks with an indicator variable that takes a value of 1 on dates when the federal funds rate was at the ELB. Table 4 presents the results of the regressions with these interaction terms. We find little difference in the effects of conventional monetary-policy shocks at the ELB. In contrast,

¹⁵As above, the results are quantitatively similar if we include outlier observations; see Appendix B.

Table 4: **Responses around FOMC Announcements: Allowing Break at ELB**

		5Y nom. yield	10Y nom. yield	5Y real yield	10Y real yield	SPX	VIX	MOVE
Monetary -policy shock	Above ELB	0.42*** (0.07)	0.22*** (0.06)	0.55*** (0.11)	0.36*** (0.09)	-3.68*** (1.28)	2.6 (1.9)	-0.08* (0.04)
	Marginal Effect of ELB	0.06 (0.21)	-0.03 (0.19)	0.14 (0.26)	0.05 (0.21)	7.89** (3.82)	-9.1 (5.7)	-0.01 (0.13)
Short-rate uncertainty (MPR)	Above ELB	0.52*** (0.16)	0.50*** (0.15)	0.79*** (0.25)	0.63*** (0.21)	-4.99* (3.01)	14.6*** (4.5)	0.91*** (0.11)
	Marginal Effect of ELB	0.68* (0.35)	0.65** (0.33)	0.67 (0.45)	0.70* (0.36)	-16.72*** (6.40)	41.1*** (9.5)	0.05 (0.22)
Adj. R^2		0.27	0.17	0.29	0.24	0.08	0.16	0.26
# obs.		239	239	165	165	239	239	238

Notes: Daily data on days of FOMC announcements, 1994 – 2026, except for real yields, which begin in 2003. We exclude observations between September 2008 and March 2009, as well as March 2020. Monetary-policy shocks are measured as changes in one-year eurodollar rates and MPR shocks are measured as changes in one-year federal-funds implied volatility. The regressions use interactive dummy variables to allow for changes in the coefficients during periods when the federal funds rate is constrained by its effective lower bound. All values are reported in basis points. Standard errors in parentheses. Asterisks indicate statistical significance at the 10% (*), 5% (**), and 1% (***) level.

we find that the ELB substantially enhances the effects of MPR on most asset prices. The effect of changes in MPR on the ten-year nominal yield, for example, are more than twice as large for the ELB observations than in the rest of the sample, and for the stock market the increase in the coefficient is more than four-fold. Most of these increases in magnitude at the ELB are statistically significant.

3 Intuition under Log Normality

In our empirical exercises above, we demonstrated (1) that uncertainty about monetary policy, conditional on the state of the economy, contributes importantly to return volatility across asset classes; and (2) that positive shocks to this kind of uncertainty cause significant price declines across asset classes. Our structural model of Section 4 will match these qualitative results, as well as most of the quantitative estimates in Tables 1 and 2. Because that model does not have a closed-form solution, however, we first explain in a simpler context how we would expect MPR to translate into asset prices and how those effects compare to the effects of other possible types

of risk in the economy. In addition to setting up the intuition of our model, the discussion of this section also illustrates that the qualitative results of that model do not hinge on its specific structural assumptions.

3.1 Reduced-Form Assumptions

Building on the linear macroeconomic framework developed above, let the economic state be $\mathbf{x}_t = [y_t, \pi_t]$, where y_t is the output gap and π_t is log of the gross change in the price level between periods t and $t-1$; let ϕ_y and ϕ_π be the coefficients on output and inflation in the monetary-policy rule; let $\mathbf{k} = [k_y, k_\pi]$ be the reduced-form macroeconomic effects of the monetary-policy shock; and let $\mathbf{A} = \begin{pmatrix} a_y^d & a_y^s \\ a_\pi^d & a_\pi^s \end{pmatrix}$ be the reduced-form effects of macroeconomic “demand” and “supply” shocks $\epsilon_t = [\epsilon_{d,t}, \epsilon_{s,t}]$. We assume that $\epsilon_{d,t}$, $\epsilon_{s,t}$, and $\epsilon_{v,t}$ are mean zero, normally distributed, independent, and serially uncorrelated. We assume $\phi_\pi > 1$ and $\phi_y > 0$, consistent with a typical policy rule that satisfies the Taylor principle. Finally, we assume that the reduced-form coefficient values determining the macroeconomic variables are such that the three shocks have the typical properties of aggregate-demand, aggregate-supply, and monetary-policy shocks in models with monetary non-neutrality. Specifically, we assume $a_\pi^s < 0$, $a_y^s > 0$, $a_\pi^d > 0$, and $a_y^d > 0$, and we assume that $\mathbf{k}$ satisfies $k_\pi < 0$, $k_y < 0$, and $\phi' \mathbf{k} > -1$. This last restriction ensures that positive monetary-policy shocks generate a contemporaneous rise in both real and nominal short-term interest rates.

For the purposes of this section, we shut down the stochastic volatility of $\epsilon_{v,t}$ by assuming constant conditional standard deviation $\zeta_t = \zeta$, and we explore the effects of MPR via comparative statics over values of ζ . Adopting the stochastic process for ζ_t , as we do in the next section, introduces second-order effects that prevent analytical tractability but do not qualitatively change the results. We also assume constant variances σ_d^2 and σ_s^2 for the macroeconomic shocks.

In order to be more explicit about the pricing function in (6), we now require the one-period real stochastic discount factor (SDF) $M_{t+1} = \exp(m_{t+1})$ in period t . An absence of equilibrium arbitrage opportunities and non-pecuniary motives for holding financial assets implies that M_{t+1} prices all real assets in the economy. In this section, for analytical tractability, we assume that

Table 5: **Assumptions about Responses to Shocks in the Log-Linear Model**

	Monetary-policy shocks	Supply shocks	Demand shocks
Output	-	+	+
Inflation	-	-	+
Nominal short rate	+	?	+
Stochastic discount factor	+	-	-

Notes: The table shows the assumed signs of the responses of the endogenous variables to each of the structural shocks in the stylized log-normal, homoskedastic model presented in this section.

the log SDF, reflecting investors' relative marginal utility in successive periods, is given by

$$m_t = -r_{t-1} + \mu^d \epsilon_{d,t} + \mu^s \epsilon_{s,t} + \mu^v \zeta \epsilon_{v,t} - \frac{\sigma_m^2}{2} \quad (10)$$

where r_t is the real one-period interest rate and σ_m^2 is the (constant) conditional variance of m_t . We assume that the effect of each of the shocks on the SDF is directionally the opposite as its effect on output, so that $\mu^v > 0$, $\mu^d < 0$, and $\mu^s < 0$. We summarize the coefficient sign assumptions in Table 5.

3.2 Asset Prices in the Log-Normal Model

In this setting, the one-period interest rate carries an inflation risk premium (IRP) equal to $\text{cov}_t[m_{t+1}, \pi_{t+1}]$. (See Appendix A.2 for derivations of standard asset pricing equations used here.) This premium represents the compensation that holders of the one-period bond demand for uncertainty about their real payoff one period hence. The IRP is higher when inflation tends to occur in bad states of the world, because in these cases the real payoffs on nominal bonds are lower in bad times.¹⁶ Under the linearity and homoskedasticity assumptions of this section, this term, like all risk premia, is constant. In the nonlinear model presented below, it will fluctuate with the economic state.

In particular, we can expand the one-period inflation-risk premium as follows:

$$IRP_t^{(1)} = a_\pi^s \mu^s \sigma_s^2 + a_\pi^d \mu^d \sigma_d^2 + \kappa_\pi \mu^v \zeta^2 \quad (11)$$

¹⁶Note that, because of this term, the Fisher equation does not generally hold. Again, this is a standard result under no-arbitrage.

Given the signs of the parameters assumed in Table 3, it follows that monetary-policy risk, as captured by the last of these three additive terms, contributes negatively to this premium. Thus, when MPR (ζ) increases, the effect on the inflation-risk premium is negative. Intuitively, monetary-policy shocks move inflation in the same direction as output and therefore in the opposite direction of marginal utility. When such shocks are more prevalent—i.e., MPR increases—the inflation risk of nominal bonds falls. The effect of uncertainty about demand shocks is the same, as demand shocks themselves send output and inflation in the same direction. In contrast, the impact of uncertainty about supply shocks is positive.

The real one-period bond has no risk premium, but the real two-period bond yield carries a real term premium (RTP), which, as Appendix A.2 shows, is given by

$$RTP^{(2)} \equiv \text{cov}_t[m_{t+1}, r_{t+1}] = [\phi_y a_y^s + (\phi_\pi - 1) a_\pi^s] \mu^s \sigma_s^2 + [\phi_y a_y^d + (\phi_\pi - 1) a_\pi^d] \mu^d \sigma_d^2 + (1 + \phi' \mathbf{k} - \kappa_\pi) \mu^v \zeta^2. \quad (12)$$

The real term premium reflects compensation for comovement between the SDF and the payoffs on real bonds (which are inversely related to the $t + 1$ real short rate). Shocks that move payoffs on real bonds and marginal utility in opposite directions increase the real term premium. In particular, from (12), a higher level of MPR results in a higher real term premium, since, under our assumptions, $(1 + \phi' \mathbf{k} - \kappa_\pi) \mu^v$ is strictly positive. The reason is straightforward. In economies with nominal rigidities, monetary-policy shocks move payoffs on real bonds and marginal utility in opposite directions. When such shocks are more prevalent—i.e., MPR increases—the real risk of real bonds rises.

A higher level of demand-side risk σ_d^2 lowers the real term premium (bonds are a hedge to bad demand shocks). The observation that demand-side risk implies negative real term premia has presented a puzzle when it comes to explaining empirical real term premia that appear to be positive on average.¹⁷ Meanwhile, the theoretical effects of supply-side risk on the real term premium are ambiguous. With the additional restriction that real rates must rise in response to a supply shock ($\phi_y a_y^s + \phi_\pi a_\pi^s > 1$), supply side uncertainty raises the RTP, but models that

¹⁷E.g., Li and Palomino (2014). D’Amico et al. (2018) estimate a 10-year real term premium of over 60 basis points over our sample period. A model free statistic is that the 10-year TIPS yield exceeded the 5-year TIPS yield by 42 on average between 2003 and 2025.

explicitly incorporate this channel (e.g., Swanson, 2021) produce real term premia that are quantitatively small. MPR presents a possible resolution to this puzzle. As we will see below, our quantitative results suggest that its contribution is substantial.

The two-period nominal term premium includes the two-period real term premium and the one-period inflation-risk premium, and it also contains additional terms reflecting the covariance between inflation and the SDF across the second period.

$$NTP_t^{(2)} \equiv RTP_t^{(2)} + \underbrace{IRP_t^{(1)} + \text{cov}_t[\pi_{t+1}, m_{t+2}] + \text{cov}_t[\pi_{t+2}, m_{t+1}] + \text{cov}_t[\pi_{t+2}, m_{t+2}]}_{IRP_t^{(2)}}. \quad (13)$$

What matters for the IRP on two-period bonds is the covariance between the cumulative values of inflation and the SDF, which involves their dynamic cross-covariances. Even in our simple model, these terms are somewhat tedious to work out algebraically. But, for similar reasons to those we discussed in the case of the one-period inflation-risk premium, those calculations show that the effect of MPR on $IRP^{(2)}$ are negative. It turns out that the positive effects of MPR on the RTP are always larger than the negative effects on the IRP, so that the nominal term premium rises with greater MPR. (See Proposition 2 below.)

Turning to equities, consider the 2-period equity claim that pays a dividend equal to output $\exp[y_{t+1}]$. The expected log nominal return on this asset is given by

$$E_t[r_{t+1}^{eq} + \pi_{t+1}] = i_t - \underbrace{\text{cov}_t[m_{t+1}, y_{t+1}]}_{ERP_t} + \underbrace{\text{cov}_t[\pi_{t+1}, y_{t+1}] - \frac{1}{2}\text{var}_t[y_{t+1}]}_{\text{convexity}} \quad (14)$$

The first term is nominal risk-free rate, and the second term, reflecting the covariance between dividends and the SDF, is the equity risk premium (ERP). In our setting, the ERP is:

$$ERP_t = -\text{cov}_t[m_{t+1}, y_{t+1}] = -a_y^s \mu^s - a_y^d \mu^d - \kappa_y \mu^v \zeta^2 \quad (15)$$

Monetary-policy shocks cause nominal dividends to fall in bad states of the world, so an increase

in MPR pushes the equity premium higher.¹⁸

The following proposition summarizes the results of this section:

Proposition 2. In the log-linear Gaussian model described above, an increase in monetary-policy risk ζ results in:

- higher variance of real and nominal interest rates and bond yields;
- higher variance of returns on real bonds, nominal bonds, and equities;
- a higher level of real and nominal term premia;
- a lower level of inflation risk premia;
- a higher level of the equity premium.

Proof. See Appendix A.3.

The effects of MPR on risk premia are summarized in Table 6. Although we will not focus on the effects of changes in aggregate-demand and aggregate-supply risk, these can be calculated in the same manner and are also reported for comparison. It is instructive to note the ways in which different structural sources of risk have similar effects and the ways in which they differ. For example, although all three sources of risk increase the equity premium, they have different directional effects on the components of the nominal term premium. Supply-side risk, like monetary-policy risk, makes nominal and real bonds riskier, since supply shocks cause interest rates to rise in bad states of the world. Demand shocks, in contrast, are associated with *lower* short-term interest rates in bad states, and thus a higher variance of demand shocks makes bonds more of a hedge. Thus, supply and monetary-policy risk lead to higher real and nominal term premia, while demand risk reduces them. On the other hand, supply and monetary-policy risk differ in their implications for the inflation-risk premium. This is because supply shocks move inflation and real activity in opposite directions, while monetary-policy shocks move them in the same direction.

The results of this section have focused on the effects of MPR on risk premia. But asset prices are also affected by the expected path of risk-free rates, inflation, and, in the case of

¹⁸The last two terms in (14) reflect an additional effect stemming from the convexity of the log return. The net effect of MPR on these terms depends on monetary policy shocks' relative impact on inflation and dividends.

Table 6: **Effects of Structural Risks on Risk Premia**

	Monetary-policy risk	Supply risk	Demand risk
Nominal term premium	+	+	-
Real term premium	+	+	-
Inflation risk premium	-	+	-
Equity premium	+	+	+

Notes: The table shows the signs of the changes of risk premia with respect to a change in the variance of each of the structural shocks in the stylized log-normal, homoskedastic model presented in this section.

equities, expected dividends. In general, risk shocks (from any source) reduce economic activity, implying that both risk-free rates and dividends will typically fall in the face of MPR shocks. In simple New Keynesian models, like the one we present in the next section, this occurs via an increase in precautionary savings, but in reality there are likely a variety of channels through which increases in macroeconomic uncertainty have economically depressing and disinflationary effects (see Bloom (2014)). All else equal, such outcomes induce an endogenous monetary-policy easing. In other words, increases in risk behave like negative aggregate demand shocks with respect to macroeconomic variables (Leduc and Liu (2016)). As a result, the overall effects on stock returns and on the real and nominal yield curves are complex and can only be determined in the context of a fully specified general-equilibrium model. We turn to such a model next.

4 Monetary-Policy Risk in a New Keynesian Model

We now develop a formal structural model to micro-found and quantify the intuitive effects discussed above. For monetary policy and its risk to be theoretically interesting, we require nominal rigidities, so we work within the New Keynesian paradigm. In order to isolate the effects we are interested in, we keep the model as simple as possible. However, we add three somewhat non-standard features that help us understand the effects of MPR on asset prices. First, we endow households with recursive preferences over consumption (Epstein and Zin (1991)), which helps to deliver realistic behavior of risk premia. Second, we introduce stochastic volatility into the shocks in the monetary-policy rule. This is how we model transitory fluctuations in MPR. Third, we allow for the effective lower bound on nominal rates, a feature that has been

a significant constraint on policy in recent years. As suggested above, including the ELB is important for explaining changes in asset-price behavior over time, as the proximity of the ELB changes the quantity and nature of interest-rate risk and thereby affects term premia (King (2019), Gourio and Ngo (2020)).

4.1 Model specification

4.1.1 Monetary policy and MPR

Expanding equation (1), we assume that the central bank follows an inertial policy rule for the short-term interest rate, subject to an effective lower bound. We model the ELB using a shadow-rate process, as in Fernandez-Villaverde et al. (2015); Gourio and Ngo (2020), and others:

$$\hat{i}_t = (1 - \varphi_i) [\pi^* + r^* + \phi_y \tilde{y}_t + \phi_\pi (\pi_t - \pi^*)] + \varphi_i \hat{i}_{t-1} + v_t \quad (16)$$

$$i_t = \max[\hat{i}_t, 0] \quad (17)$$

where π^* is the long-run inflation target, r^* is the steady-state real interest rate, $\tilde{y}_t$ is the log deviation of output from its steady-state value, and v_t is the monetary-policy shock. We continue to assume that v_t follows the dynamics given in equation (4) and that its conditional standard deviation, given by the state variable ζ_t , follows the AR(1) dynamics in (5). However, in order to rule out negative variance, we now assume that the distribution of η_t is subject to truncation below at $-E_{t-1}[\zeta_t]$.¹⁹ We specifically choose to model the stochastic volatility term as a linear, rather than log-linear or square-root, process to ensure that the conditional variance of MPR, $\text{var}_t[\zeta_{t+1}]$, is nearly constant. Under a log process, this variance increases with η_t . Increases in the conditional variance of η_t have qualitatively different effects on asset prices than increases in the conditional variance of v_t , and nonlinear processes for ζ_t confound these two effects. The linear process, in contrast, allows us to isolate the effects of $\text{var}_t[v_{t+1}]$ that we are interested in.²⁰

¹⁹Specifically, the PDF of η_t is $\frac{1}{\sqrt{2\pi}\omega^{3/2}} \exp\left[-\frac{\eta_t}{2\omega^2}\right] \left(1 - \text{erf}\left[\frac{(1-\rho_\zeta)\tilde{\zeta} + \rho_\zeta\zeta_{t-1}}{\sqrt{2}}\right]\right)$.

²⁰The linear specification follows Mumtaz and Zanetti (2013) and Basu and Bundick (2017), although, because those papers solve by perturbation they sidestep the problem of the non-negativity constraint.

4.1.2 Households

We want to exploit the attractive asset-pricing properties of recursive preferences over consumption (C_t) within the context of a New Keynesian model where households also receive disutility from labor (H_t). The question of how best to extend recursive preferences to include leisure is not settled in the literature, and the few other papers that have worked with such models have done it in different ways.²¹ Our approach is to assume that households' lifetime utility is given by preferences of the following form:

$$U_t = \left[(1 - \beta)(C_t^\rho - \theta_t H_t^{\chi\rho}) + \beta \left(\mathcal{R}_t (V_{t+1}^C)^\rho + \mathcal{R}_t (V_{t+1}^H)^\rho \right) \right]^{1/\rho} \quad (18)$$

where θ_t is an exogenous process governing the consumption-leisure trade-off, the certainty-equivalence function $\mathcal{R}_t(\cdot)$ is defined by

$$\mathcal{R}_t(X_{t+1}) \equiv E_t[X_{t+1}^\alpha]^{1/\alpha}$$

for any random variable X_{t+1} , and the continuation values of consumption and leisure are given by the recursions

$$\begin{aligned} V_t^C &= \left[(1 - \beta)C_t^\rho + \beta \mathcal{R}_t (V_{t+1}^C)^\rho \right]^{1/\rho} \\ V_t^H &= \left[(1 - \beta)\theta_t H_t^{\chi\rho} + \beta \mathcal{R}_t (V_{t+1}^H)^\rho \right]^{1/\rho} \end{aligned}$$

This specification nests both Epstein-Zin preferences over consumption (if $\theta_t = 0$ for all t) and the additively separable preferences over consumption and leisure that are standard in the NK literature (if $\alpha = \rho$). It has the convenient feature that the intra-temporal optimality condition will not involve the recursive terms, preserving the usual tractability of general-equilibrium business-cycle models, but that the *inter*-temporal optimality condition will reproduce the Epstein-Zin discount factor familiar from consumption-based asset pricing.

Letting P_t be the price of the consumption good and W_t be the nominal wage, the household

²¹See, for example, the various functional forms adopted by Dew-Becker (2014), Basu and Bundick (2017), and Swanson (2021)

maximizes lifetime utility subject to the flow budget constraint:

$$P_t C_t + P_t^{\S(1)} B_t \leq B_{t-1} + W_t H_t + \mathcal{P}_t \quad (19)$$

where $\mathcal{P}_t$ is firm profits. Since the model will feature a trend growth rate g in log real wages and log consumption, to keep hours stationary, we assume that θ_t grows at a constant rate $-\rho g$, following Rudebusch and Swanson (2012).

4.1.3 Firms

The production side of our economy is completely standard. Firms produce differentiated goods, indexed by i , according to the Cobb-Douglas production function

$$y(i)_t = a_t^* + \psi \log H(i)_t \quad (20)$$

where the steady-state level of log productivity a_t^* follows deterministic growth

$$a_t^* = a_{t-1}^* + g$$

Firms face iso-elastic demand from households and are subject to Calvo pricing. Aggregate output is given by the Dixit-Stiglitz aggregate of $e^{y(i)}$.

4.2 Equilibrium and solution

Appendix C shows that the optimality conditions for the household's problem can be written as

$$\chi \theta_t H_t^{\chi \rho - 1} C_t^{1 - \rho} = \frac{W_t}{P_t} \quad (21)$$

$$\tilde{C}_t = \exp[g] \left(\beta \exp[i_t] E_t \left[\frac{\tilde{C}_{t+1}^{\rho-1}}{\Pi_{t+1}} \left(\frac{\tilde{V}_{t+1}^C}{\mathcal{R}_t(\tilde{V}_{t+1}^C)} \right)^{\alpha - \rho} \right] \right)^{\frac{1}{\rho-1}} \quad (22)$$

$$\tilde{V}_t^C = \left[(1 - \beta \exp[g\rho]) \tilde{C}_t^\rho + \beta \exp[g\rho] \mathcal{R}_t \left(\tilde{V}_{t+1}^C \right)^\rho \right]^{1/\rho} \quad (23)$$

where tildes denote percentage deviations from the balanced-growth path. After log-linearizing, the firm's problem, combined with the labor-supply curve (21), results in the New Keynesian Phillips Curve:

$$\pi_t = \pi^* + \beta \exp[g(\rho - 1)] E_t[\pi_{t+1} - \pi^*] + \kappa \tilde{y}_t \quad (24)$$

where $\kappa > 0$ is reduced-form parameter. Note that, in order to obtain meaningful risk premia, we do not linearize the Euler equation.

A solution to the model consists of values of C_t , H_t , W_t/P_t , π_t , and i_t such that equations (21) through (24) are satisfied and the goods, bond, and labor markets clear ($C_t = \exp y_t$, $B_t = 0$, and $\int H(i)_t di = H_t$) for all t .

4.3 Asset Prices

The log SDF in this model is given by

$$m_{t+1} = \log \beta + (\rho - 1)(g + \tilde{c}_{t+1} - \tilde{c}_t) + (\alpha - \rho)(\log \tilde{V}_{t+1}^C - \log \mathcal{R}_t(\tilde{V}_{t+1}^C)) \quad (25)$$

and the prices of nominal and real bonds of maturity N are given by

$$P_t^{\$(N)} = E_t \left[\frac{M_{t+1}}{\Pi_{t+1}} P_{t+1}^{\$(N-1)} \right] \quad (26)$$

$$P_t^{(N)} = E_t \left[M_{t+1} P_{t+1}^{(N-1)} \right] \quad (27)$$

The yield on a nominal bond is defined as

$$y_t^{\$(N)} \equiv -\frac{1}{N} \log P_t^{\$(N)} \quad (28)$$

and similarly for real yields. "Inflation compensation" is the difference between nominal yields and real yields. One-period yields on real and nominal bonds are the short-term interest rates; $r_t = y_t^{(1)}$ and $i_t = y_t^{\$(1)}$.

The real term premium is the difference between the yield on an inflation-indexed bond and the value that the yield would take if agents were risk-neutral. Under risk neutrality, assets

are priced as if M_{t+n} , for all n , is always equal to its time- t expectation with certainty. In this case, the term structure of interest rates is given by the geometric average of expected future short-term rates. In Appendix A.2, we define a geometric expectation operator $\mathcal{E}_t^{(N)}[\cdot]$, which subsumes Jensen's inequality terms into physical expectations of future prices and cash flows and leave risk premia as clean measures of deviations from risk neutrality. Thus, the real term premium (RTP) is

$$RTP_t^{(N)} = y_t^{(N)} - \mathcal{E}_t^{(N)}[r_t] \quad (29)$$

The inflation risk premium (IRP), representing the extra yield that nominal bonds must pay over real bonds for bearing the risk of nominal payoffs, is defined as inflation compensation less expected inflation:

$$IRP_t^{(N)} = y_t^{\$(N)} - y_t^{(N)} - \mathcal{E}_t^{(N)}[\pi_t] \quad (30)$$

We then define the nominal term premium (NTP) as

$$NTP_t^{(N)} = RTP_t^{(N)} + IRP_t^{(N)} \quad (31)$$

$$= y_t^{\$(N)} - \left(\mathcal{E}_t^{(N)}[r_t] + \mathcal{E}_t^{(N)}[\pi_t] \right) \quad (32)$$

Note that, although we have defined the RTP in the usual intuitive way as the difference between real yields and real rate expectations, the NTP is not defined analogously (since $\mathcal{E}_t^N[i_t] \neq \mathcal{E}_t^N[r_t] + \mathcal{E}_t^N[\pi_t]$). This asymmetry is necessary to ensure that the IRP and RTP add up to the NTP. Although this definitional choice is not important for our results, one implication is that short-term nominal bonds carry a term premium, consisting entirely of $IRP_t^{(1)}$. In other words, as in the model of Section 3, the Fisher equation does not generally hold because even one-period nominal bonds bear some amount of inflation risk, which requires compensation. Again, this is a standard no-arbitrage result.

The nominal return on any bond is its real return plus inflation. One-period log real returns

on nominal bonds and real bonds are thus as follows:

$$r_{t+1}^{\$(N)} = \log P_{t+1}^{\$(N-1)} - \log P_t^{\$(N)} - \pi_{t+1} \quad (33)$$

$$r_{t+1}^{(N)} = \log P_{t+1}^{(N-1)} - \log P_t^{(N)} \quad (34)$$

To price equities, first consider a hypothetical contract that pays a stream of cash flows equal to aggregate nominal output. The price of such a contract can be found recursively as

$$P_t^y = E_t[\exp(m_{t+1} - \pi_{t+1})(P_{t+1}^y + \exp(y_{t+1} + \pi_{t+1}))] \quad (35)$$

and its log real returns are

$$r_{t+1}^y = \log [P_t^y + \exp(y_t + \pi_t)] - \log P_{t-1}^y - \pi_{t+1} \quad (36)$$

Following Campbell (1986), we model equity contracts as *levered* claims on output. To fund levered positions, firms are assumed to borrow a multiple $1/\delta$ of their equity value at the short-term nominal rate i_t , rolling over this debt every period. In this case, Campbell shows that the real return on equities is given by

$$r_{t+1}^{eq} = \log \left[\frac{1}{\delta} \exp(r_t^y + \pi_t) - \frac{1-\delta}{\delta} \exp(i_{t-1}) \right] - \pi_{t+1} \quad (37)$$

The equity-risk premium is the expectation of the nominal equity return, less the value that this expectation would take under risk neutrality, implying

$$ERP_t = E_t[r_{t+1}^{eq} + \pi_{t+1}] - i_t \quad (38)$$

Because of convexity terms, this is slightly different from the expected excess returns on equities, which are given by $E_t[r_{t+1}^{eq} + \pi_{t+1}] - i_t$.²²

²²Note that, while this is typically how excess returns are measured in the data, some theoretical treatments focus on what one might call the “real excess returns,” $E_t[r_{t+1}^{eq}] - r_t$. These are not generally the same because the Fisher equation does not hold. Unless otherwise noted, we work with the nominal returns.

Table 7: **Parameter Values in New Keynesian Model**

Parameter	Interpretation	Value	Source
β	Time preference	0.99	Standard
ρ	Inverse IES	0.5	Standard
ψ	Labor share in production	0.35	Standard
κ	Philips Curve slope	0.05/4	Standard
ϕ_π	Policy response to inflation	1.5	Standard
ϕ_y	Policy response to output gap	0.5/4	Standard
r^*	Steady-state real rate	0	Average of data
π^*	Steady-state inflation	0.02/4	Average of data
α	Risk aversion	-254	Match empirical results
ϕ_i	Policy rule inertia	0.63	Match empirical results
ρ_ν	Persistence of monetary policy shocks	0.75	Match empirical results
$\bar{\zeta}$	Steady-state MPR	0.0048/4	Match empirical results
ρ_ζ	Persistence of MPR shocks	0.62	Match empirical results
ω	Std. dev. MPR shocks	0.0026/4	Match empirical results
δ	Equity leverage	0.84	Match empirical results

Note: The table shows the calibrated parameter values used to produce quantitative results from the New Keynesian model described in the text. Time interval is taken to be quarterly.

4.4 Calibration and Solution

Table 7 presents the parameter values we use in our baseline model. Time is quarterly. We calibrate a subset of our parameters to standard values. Specifically, we set the quarterly rate of time preference β to 0.99, the intertemporal elasticity of substitution to 2, the labor share ψ to 0.35, and the slope of the Phillips curve to 0.05, when inflation is expressed in annualized terms.²³ We calibrate the response coefficients in the policy rule to the Taylor (1993) values of $\phi_\pi = 1.5$ and $\phi_y = 0.5$, and we set $\pi^* = 2.0\%$ and $r^* = 0\%$, annualized, to match the average values of inflation and short-term interest rates since the mid-1990s.²⁴

The remaining values we calibrate to match the empirical results presented earlier, via simulated method of moments. Specifically, we choose parameters that minimize the deviations from the implied quarterly volatilities due to MPR across all assets and the responses of all asset prices to monetary-policy and MPR shocks, as estimated above, normalizing by the respective standard errors. We weight the equity moments twice as heavily as the bond moments because

²³The frequency of Calvo price adjustments and the parameter governing the elasticity of labor supply (χ) both enter into the Phillips Curve slope. Since we do not focus on any results concerning wages, these parameters enter nowhere else in our model. Thus, we calibrate the Phillips Curve slope directly, rather than taking a stand on the values of these two structural parameters.

²⁴Given the other parameters, calibrating r^* is equivalent to calibrating the steady-state growth rate g .

we have more of the latter. In doing these comparisons, we evaluate the model at the steady state and match to the empirical results in the samples that exclude periods at the ELB.

Conditional on the other parameters, the variances of the monetary-policy and MPR shocks, ζ and ω , are identified from the differences in asset-price volatilities between blackout and non-blackout days reported in Table 2. In particular, these data lead us to calibrate a steady-state conditional standard deviation of monetary-policy shocks, ζ , of 48 bp per quarter and a conditional standard-deviation of MPR shocks of 26 bp (with interest rates expressed as annual rates).²⁵ Meanwhile, the relative responses of real and nominal yields of different maturities to the two shocks largely pin down the persistence parameters ϕ_i , ρ_z , and ρ_v . All three of these take relatively modest values between 0.60 and 0.75 at a quarterly frequency, but in combination they will imply somewhat persistent dynamics for MPR shocks.

Finally, in order to match the empirical responses to monetary-policy and MPR shocks in Table 4, the model requires $\alpha = -254$ and $\delta = 0.84$. As is well known, matching equity and bond returns requires a high level of risk aversion, and indeed the principal advantage of Epstein-Zin preferences are that they allow us to set α to a high value while maintaining an empirically realistic IES. While $\alpha = -254$ is unrealistic at the level of an individual consumer, we view the representative-household preferences used here as an aggregate across many individuals and institutions in the economy, which need not preserve microeconomic parameter values. (Gourio and Ngo (2020) make a similar point.)

Previous papers that have studied asset prices in DSGE models have often relied on higher-order perturbation methods to solve them.²⁶ Although this approach is relatively efficient and can give accurate approximations for macroeconomic dynamics near the steady state, it is unlikely to have good properties for asset prices in the presence of significant nonlinearities. This is because asset prices can be sensitive to the magnitudes of outcomes with low unconditional probabilities, which are necessarily far away from the steady-state perturbation point. If the

²⁵The value of ζ is somewhat high compared to the conditional variance of short-term interest rates in the data. However, shocks to monetary policy do not pass through one-to-one to short-term interest rates, because they have effects on output and inflation that negatively feed back through the policy rule. In addition, ζ needs to capture uncertainty about the entire path of policy, and, in the data, that uncertainty is considerably larger beyond the very short term.

²⁶Obtaining time-varying risk premia requires perturbation of at least the third order. Examples include Rudebusch and Swanson (2012), Swanson (2021), and Mumtaz and Zanetti (2013).

policy functions are not sufficiently smooth, polynomial approximations far from this point can be quite poor.²⁷ These problems may be compounded if the model involves non-differentiable functions, such as those induced by the ELB. To avoid these issues, we rely on quadrature and interpolation across the state space, a method that is computationally intensive but provides a globally accurate solution, even in the presence of occasionally binding constraints. Appendix D describes the numerical procedure in detail.

5 Model Results

5.1 Matching empirical moments

Table 8 shows how the model performs in matching the evidence on the contribution of MPR to asset-price volatility that we derived from blackout-period data in Table 2, focusing on the subsample with rates above the ELB. Specifically, we calculate the unconditional standard deviations of various asset prices in the model, discarding observations where the ELB is binding. Since MPR and its volatility are the only sources of variation in the model, we can compare these numbers to the contributions of MPR and its volatility that we isolated in the data using the blackout-period evidence. Although we calibrated the parameters of the model to roughly match these results, as well as those in Table 4, we have many more empirical moments than model parameters so a good overall fit is not guaranteed. Even so, the model is generally able to match the difference in variance between blackout and non-blackout days of real and nominal bond prices, stock prices, and their implied volatilities.

Table 9 shows how the model matches the evidence on the responses of asset prices to both MPR shocks and monetary-policy shocks, based on the event studies around FOMC announcements that we presented in Table 4. We compare to the empirical results in the sample away from the ELB. To construct these values in the model, we simulated 10,000 observations and ran the same regressions that we ran on the actual data, discarding any observations for which the shadow rate was negative. The case of shocks at the ELB will be examined below.

For both types of shocks, the model performs well in matching the regression results. For

²⁷See Aldrich and Kung (2019).

Table 8: **MPR Contribution to Quarterly Volatility of Asset Prices**

	Stock return	5y nom. yield (bps)	5y real yield (bps)	10y nom. yield (bps)	10y real yield (bps)	VIX (ann. %)	MOVE (ann. bps)
Data	1.6 (1.7)	14 (5)	16 (5)	11 (6)	13 (5)	1.9 (2.2)	11 (3)
Model	1.1	15	23	7	11	0.9	8

Notes: The table shows the quarterly standard deviations of asset prices calculated in the New Keynesian model, compared to the values in the data. Data values are calculated from daily market volatilities during blackout periods, relative to non-blackout dates, scaled to a quarterly frequency, excluding observations at the ELB. (See Table 2.) Bootstrapped standard errors are reported in parentheses. For the purposes of comparison, the MOVE index is compared to the volatility of the model-based five-year nominal yields, and variance risk premia are assumed constant.

Table 9: **Asset-Price Responses to Shocks**

A. Response to monetary-policy shock on impact, holding 1-year uncertainty constant

	5y nom. yield (bps)	5y real yield (bps)	10y nom. yield (bps)	10y real yield (bps)	Stock return (%)	VIX (ann. %)	MOVE (ann. bps)
Data	+42 (7)	+55 (11)	+22 (6)	+39 (9)	-3.7 (1.3)	+2.6 (1.9)	-7 (4)
Model	+34	+53	+17	+27	-3.5	-0.1	-1

B. Response to MPR shock on impact, holding 1-year short rate constant

	5y nom. yield (bps)	5y real yield (bps)	10y nom. yield (bps)	10y real yield (bps)	Stock return (%)	VIX (ann. %)	MOVE (ann. bps)
Data	+52 (16)	+79 (26)	+50 (15)	+63 (21)	-5.0 (3.0)	+14.6 (4.5)	+91 (11)
Model	+33	+76	+17	+38	-5.2	+9.2	+85

Notes: The table shows the coefficients in regressions of asset-price changes on one-year-forward short-term interest rates (panel A) and one-year-forward short-rate volatility (panel B), based on data simulated from the New Keynesian model described in the text, compared to the corresponding coefficients in the data. Model-based values are generated from a simulated sample of 10,000 observations, dropping observations where short rate is constrained by the ELB. The data values are calculated from the coefficients estimated based on data excluding observations at the ELB (see Table 4). Standard errors are reported in parentheses. For the purposes of comparison, the MOVE index is matched to the volatility of the model-based five-year nominal yields, and variance risk premia are assumed constant.

Table 10: **Model Contributions to Selected Risk Premia**

	10y nom. term prem. (bps)	10y real term prem. (bps)	10y inflation risk prem. (bps)	Equity premium (bps, annualized)
MPR ($\bar{\zeta}$)	43	65	-22	104
MPR volatility (ω)	0	6	-6	12
Total	43	71	-28	116

Notes: The table shows the average values of risk premia on long-term bonds and equities produced by the New Keynesian model. Premia are decomposed into the contributions of constant monetary-policy risk and variability in monetary-policy risk by recomputing values in a version of the model where ω is set to zero.

monetary-policy shocks, it matches every coefficient to within one standard error. For MPR shocks, all of the model-implied moments have the same sign as we find in the data, and the magnitudes of the responses are mostly within two standard errors of the regression point estimates. Quantitatively, the model does err somewhat to the downside when it comes to the responses of long-term yields to MPR shocks. This likely reflects the relatively low persistence of the shocks in our calibration. Even so, given the relatively small number of free parameters, the model does remarkably well at explaining the the overall behavior of the data with respect to monetary and MPR shocks.

5.2 Contributions to risk premia

Table 10 reports unconditional means of the nominal ten-year term premium, inflation risk premium, real term premium, and equity premium and the contribution to these overall values that come from MPR and from the volatility of MPR. We measure the contribution of the volatility of MPR shocks as the difference between the model with homoskedastic policy shocks and the baseline model that allows for stochastic variance.²⁸

In total, stochastic variation in the policy rule implies an average 10-year nominal term premium of 43 bp and an average equity premium of 116 bp (expressed annually). The nominal term premium reflects an inflation-risk premium of -28 bp and a real term premium of 71 bp, on average. The inflation-risk premium is negative for reasons described above. The average slope

²⁸In the homoskedastic model, $\bar{\zeta}$ is adjusted such that the policy shocks have exactly the same unconditional variance as the shocks in the baseline model.

of the real yield curve is determined by real term premia. Thus, compensation for risk arising from monetary policy is one potential explanation for why the real yield curve slopes upward on average, and this effect passes through to nominal yields as well. The effect is roughly on par with estimates of the total size of the real term premium in the data. For example, over the period we cover, the model of D’Amico et al. (2018) estimates an average 10-year real term premium of 63 bp.²⁹

The effects on average risk premia can be decomposed into the contribution of the average level of MPR in the model (as determined by $\bar{\zeta}$) and the variation in MPR (as determined by ω). Most of the contribution comes from the level. Shutting down the shocks to ζ_t , results in a real term premium of 65 bp, an inflation-risk premium of -22 bp, and an equity premium of 104 bp. Meanwhile, volatility in MPR itself contributes 6 bp to the RTP, subtracts 6 basis points from the average value of the IRP, and makes a small positive contribution to the equity premium. (The variance of MPR also contributes to the existence of negative variance risk premia, not shown, though these effects are relatively small.)

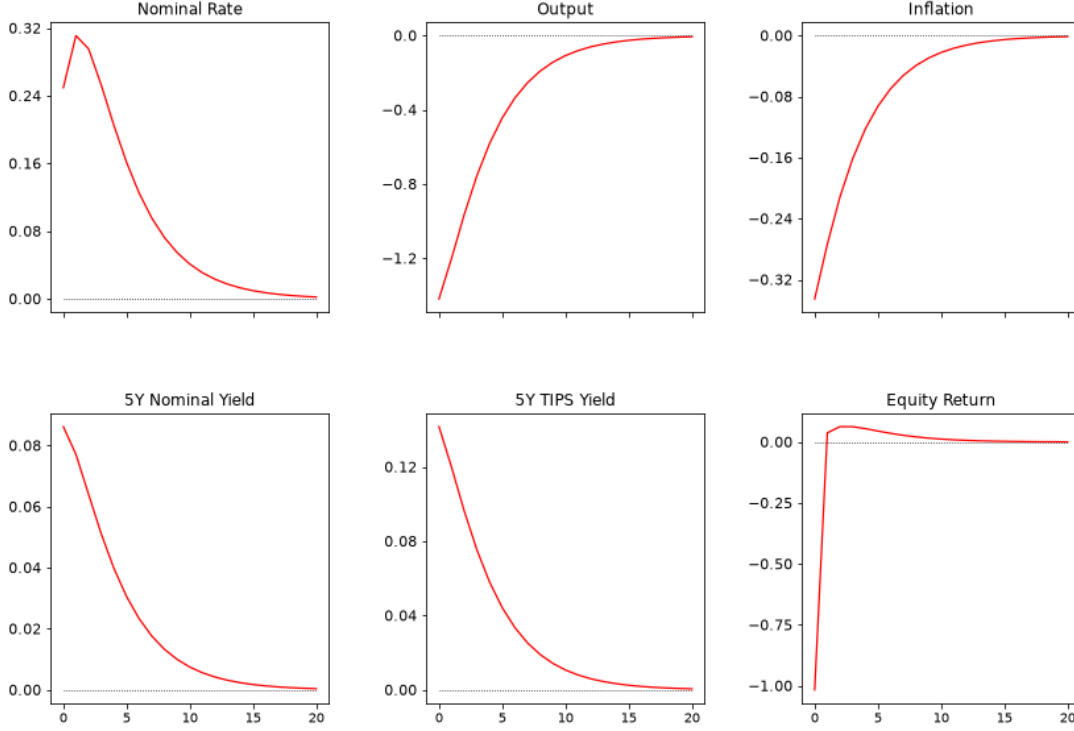
We note that the contributions of $\bar{\zeta}$ to risk premia all have the signs that we hypothesized based on the reduced-form model (see Table 6). In most respects, η_t shocks act like demand shocks, both for asset prices and the macroeconomy. (See Leduc and Liu (2016).) However, a positive shock to MPR causes real yields to rise through its effect on real term premia. Consequently, when MPR shocks are volatile, real bonds are more likely to experience losses in bad states of the world, pushing the average level of real term premia higher. In this respect, they differ from typical aggregate-demand shocks, for which real bonds are a hedge and command negative risk premia.

5.3 Effects of shocks at the steady state

As shown in Figure 2, monetary-policy shocks in our model have effects on the macroeconomy and asset prices that are fairly standard. When the economy is at the steady state, a 25 bp exogenous increase in the short-term interest rate reduces the output gap by about 0.9% and

²⁹We do not, of course, claim that MPR is the *only* determinant of real term premia. As noted in Section 3, other types of risk may move the real term premium in either direction. In particular, risk to the demand side of the economy generally has a negative effect on real term premia.

Figure 2: **Impulse-Responses to Monetary-Policy Shock at Steady State**

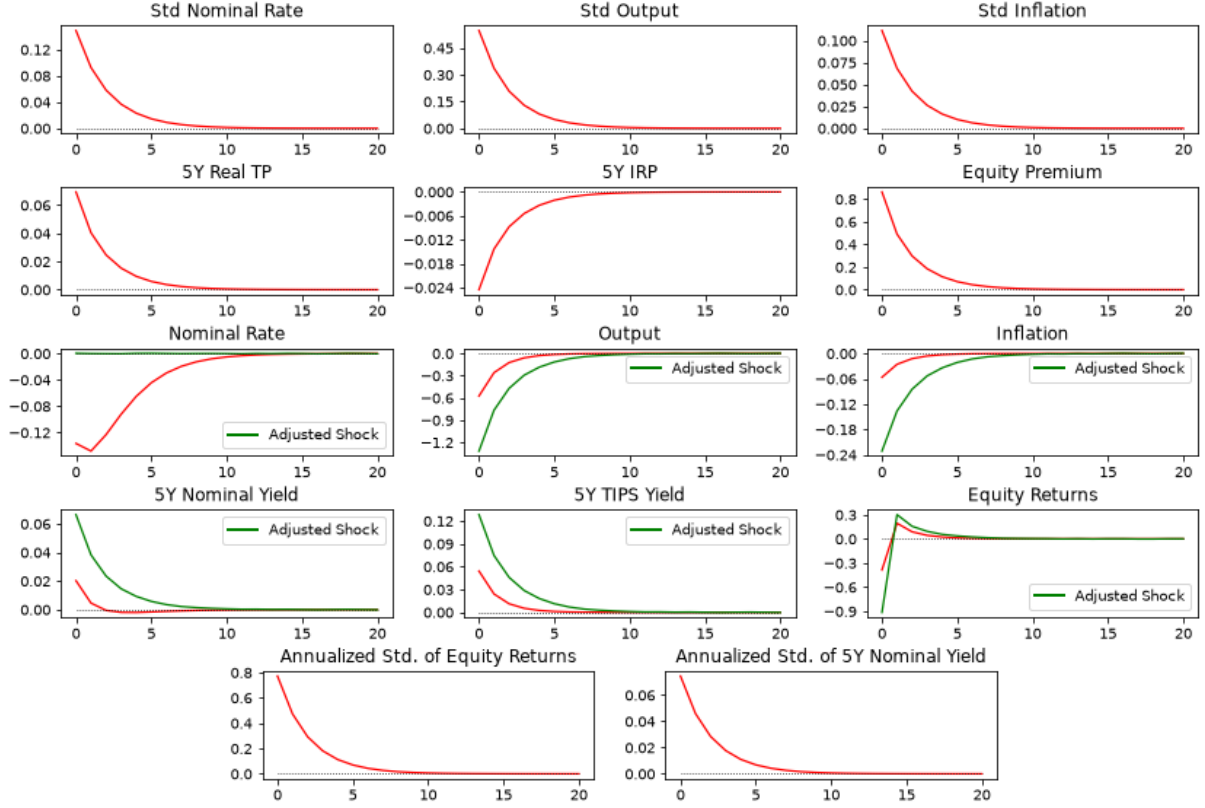


Notes: The figure shows the responses of the indicated variables to a monetary-policy shock ($\epsilon_{v,t}$) in the baseline model, initialized at the steady state. The shock is scaled to produce a response of 25 bp in the nominal short-term interest rate on impact. The horizontal scale shows time in quarters. All vertical scales are in percentage points. Interest rates and inflation are reported at annualized rates.

inflation by about 0.2% on average over the subsequent year. On impact, it increases the 5-year nominal yield by 7 basis points, increases the 5-year real yield by 14 basis points (so, 5-year inflation compensation falls by 7 basis points), and causes stock returns to fall by about 1%. As we showed in Table 7, these asset-price responses are all quantitatively very similar to those estimated from the event-study data. Near the steady state, monetary-policy shocks do not have appreciable effects on volatilities or risk premia.

Our main objects of interest are the responses to the MPR shocks. The top row of Figure 3 shows that a one-standard-deviation positive shock to MPR causes the conditional (one-period-ahead) standard deviation of the nominal interest rate to increase by 14 bp, with corresponding increases in the conditional standard deviations of output and inflation. As shown in the second row, this increase in uncertainty raises real risk premia for both bonds and equities, with the 5-year real term premium increasing 7 bp and the equity premium rising 85 bp (at an annualized

Figure 3: **Impulse-Responses to MPR Shock at Steady State**



Notes: The red lines in the figure show the responses of the indicated variables to a one standard-deviation monetary-policy risk shock (η_t) in the baseline model, initialized at the steady state. The green lines show the responses when the effect of this shock on the nominal short rate is neutralized by introducing a series of offsetting monetary-policy shocks (η_t). The horizontal scale shows time in quarters. All vertical scales are in percentage points. Interest rates and inflation are reported at annualized rates.

rate). For reasons discussed above, the inflation risk premium declines slightly, but the magnitude is less than the increase in the real term premium, so that the nominal term premium (not shown separately) rises.

In addition to the effects on risk premia, the MPR shocks also have effects on risk-free rates and cash flows through macroeconomic channels. As shown in the third row of the figure, the MPR shock by itself reduces both output and inflation modestly in the short run. These effects are damped by the endogenous monetary-policy response—the shock triggers a short-run decline of about 15 bp in the policy rate. To isolate the effect of MPR, it is also helpful to consider the case in which the central bank does not ease its current policy in response to uncertainty about its future policy. We can construct such a scenario by introducing a sequence of additional

monetary-policy shocks to keep the policy rate unchanged over time. This scenario is depicted by the green lines.³⁰

The final two rows of the figure show how these effects net out for asset prices. Holding monetary policy constant, the MPR shock raises 5-year nominal and real yields by 7 and 13 bp, respectively, on impact (so inflation compensation falls by 6 bp), and it reduces stock returns by about 0.9%. For comparison, these asset-price responses are quantitatively similar to those that would be induced by a monetary-policy shock of about 25 bp (see Figure 2). In the quarters after the shock, equities post excess returns slightly above their steady-state values, reflecting the persistently elevated equity premium. In addition, the MPR shock causes the annualized conditional volatility of the stock market (our counterpart to the VIX index) to rise by about 80 bp and that of the 5-year yield (our counterpart to the MOVE) to rise by about 8 bp on impact.³¹

Finally, Figure 4 shows how the term structures of volatility and interest rates react in the period when the MPR shock occurs. In the top panel, we show the response of the standard deviation of the one-period interest rate h periods ahead ($\text{std}_t[i_{t+h}]$). It decreases monotonically in h , with no substantial impact beyond a horizon of about three years. Even so, risk premia in the bond market respond non-monotonically across maturities. As shown in the lower panel, the response of the real term premium, which is generally the component of the yield curve that moves the most in response to MPR shocks, peaks at about 14 basis points at the two-year maturity, but it is elevated out to at least ten years. This response of real term premia contributes to the substantial steepening of the yield curve, as shown by the black line.³²

5.4 Effects of MPR at the effective lower bound

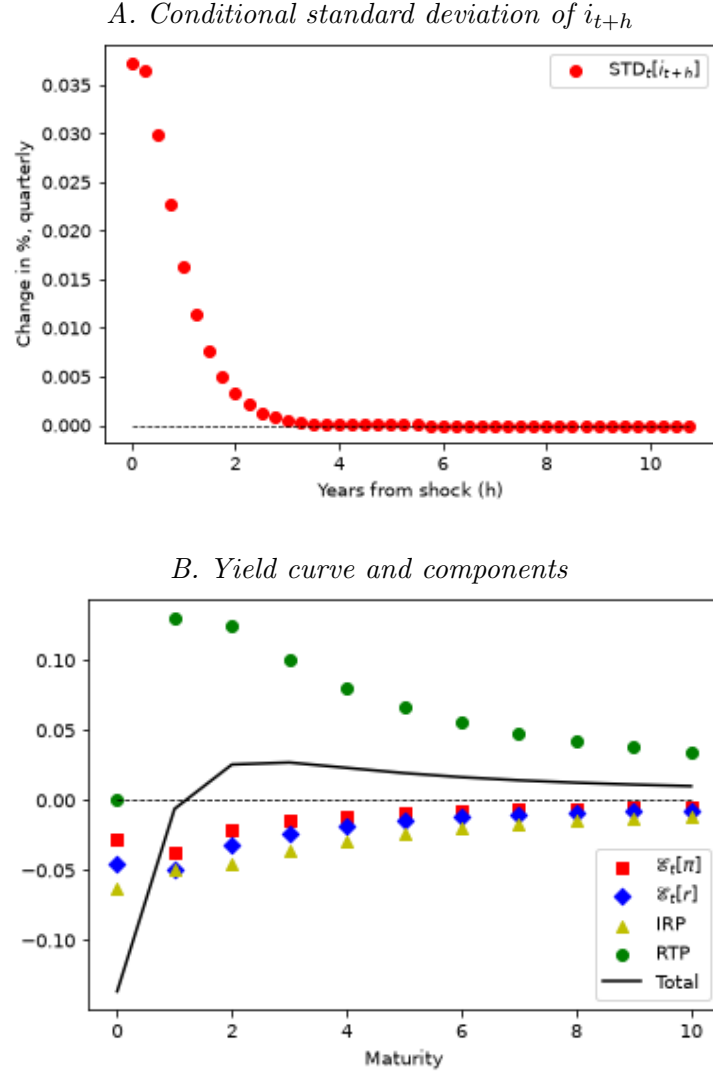
The effective lower bound on nominal interest rates can have significant consequences for the asset-pricing effects of MPR. Because policy rates are essentially frozen in the near term, interest-

³⁰Again, near the steady state, these monetary-policy shocks have negligible effects on volatilities or risk premia, so the IRFs shown in the first two rows of the figure are essentially the same.

³¹Our model also generates analogues to option-implied volatilities, but variance risk premia are small and respond only by negligible amounts to shocks. Therefore, for simplicity, we report the asset-price volatilities under the physical measure.

³²In the figure, we show the response of the yield curve and its components without offsetting the effect of the MPR shock on the path of interest rates. Although that modification would affect the expectations components of real rates and inflation, it has little effect on term premia at the steady state.

Figure 4: **Response of rate uncertainty and yield curve to MPR shock on impact**



Notes: The figures show the time-0 responses of the indicated variables to a one-standard-deviation monetary-policy risk shock (ηt) in the baseline model, initialized at the steady state. Panel A shows the change in the conditional standard deviation of the short-term interest rate across horizons. The horizontal scale is in quarters. Panel B shows the change in the yield curve (black line) and its components: expected inflation (red dots), expected real short rates (blue dots), the inflation risk premium (yellow dots), and the real term premium (green dots). Maturities are indicated in years.

rate volatility is lower at the ELB, and, all else equal, this has a dampening effect on risk premia. For the same reason, MPR shocks pass through less to risk premia across assets. In addition, because short-term rates cannot fall to offset the economic deterioration that occurs in response to greater uncertainty, output and inflation generally fall by more at the ELB than they do

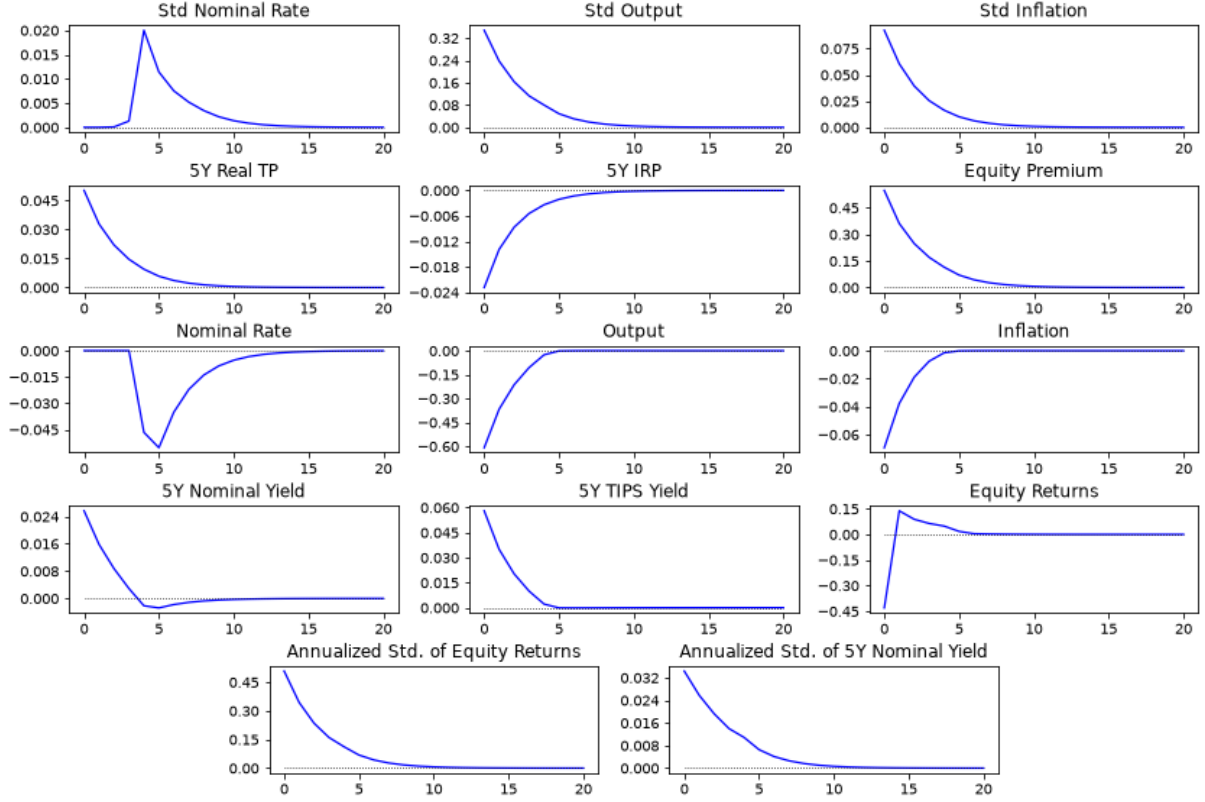
at the steady state. (This mirrors the standard result for aggregate-demand shocks in New Keynesian models.) The net effects of changes in risk premia, risk-free rates, and cash flows are ambiguous.

Figure 5 shows the key impulse-responses when initiated at the ELB. Specifically, for the purposes of these plots, we initialize the IRF with values of i_{t-1} that would result in a short-term rate that would remain at the ELB for four quarters in the absence of shocks. All of the asset-price responses have the same sign at the steady state and at the ELB, but the magnitudes differ somewhat. For example, the impact of the MPR shock on contemporaneous stock returns is about 0.30% at the steady state (when we do not offset the endogenous monetary-policy response), but it is about 0.45% at the ELB, even though the equity premium rises by a *smaller* amount in that state. This reflects the larger multipliers associated with the absence of a near-term policy response at the ELB.

As shown in Figure 6, the effects on the long end of the nominal yield curve are similar to the steady-state case, but the short end of the curve cannot decline in this environment. The constrained nominal short rate, together with the amplified effect of MPR on inflation at the ELB, imply that the path of real rates rises significantly after the shock, in contrast to the steady-state case. For this reason, medium-term nominal yields respond more to MPR shocks at the ELB than at the steady state.

Table 4 showed that, at the ELB, a given change in near-term interest-rate uncertainty around FOMC meetings is generally associated with a larger changes in asset prices than when policy is not constrained. In Table 9, we showed that, at the steady-state, the model comes close to quantitatively matching the empirical responses to shocks, based on regressions on data simulated from the model. We cannot run the corresponding regressions to check for coefficient changes at the ELB for three reasons. First, changes in one-year rates and one-year-ahead volatilities become nearly colinear as the shadow rate becomes deeply negative (unlike at the steady state, where they are effectively independent). Second, the extreme nonlinearity associated with these states of the world imply that the results are highly sensitive to the initial conditions and the magnitude and direction of the shocks, which we do not observe in the data. Third, our model omits macroeconomic shocks on the demand and supply side that would

Figure 5: **Impulse-Responses to MPR Shock at ELB**

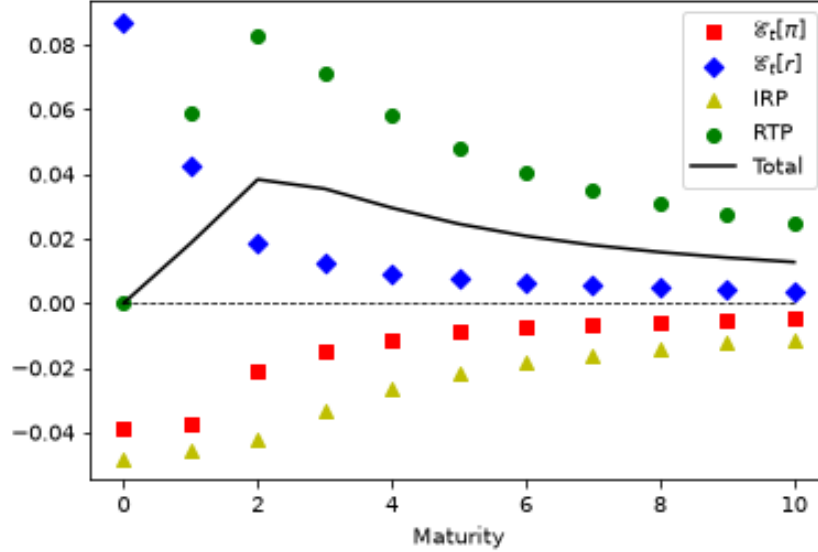


Notes: The figure shows the responses of the indicated variables to a one standard-deviation monetary-policy risk shock (η_t) in the baseline model, initialized at a value of i_{t-1} that would keep the short rate at the ELB for four quarters in the absence of shocks. The horizontal scale shows time in quarters. All vertical scales are in percentage points. Interest rates and inflation are reported at annualized rates.

endogenously affect uncertainty about the interest-rate path. Ignoring these other sources of uncertainty makes little difference away from the ELB because the effects of the shocks are roughly additively separable. But when the shadow rate is negative the shocks interact, and the omitted sources of uncertainty matter for the results.

To examine how the ELB matters for the transmission of MPR on asset prices, we instead compute the contemporaneous effects of a one-standard-deviation MPR shock, divided by its effect on the conditional standard deviation of i_{t+4} (corresponding to how we measured short-rate uncertainty in the data). We isolate the effects of MPR by neutralizing the contemporaneous endogenous change in i_t with an offsetting monetary policy shock, as we did in the green lines in Figure 3.

Figure 6: **Response of Yield Curve to MPR on impact, at ELB**



Notes: The figure shows the time-0 responses to a one-standard-deviation monetary-policy risk shock (η_t) in the baseline model, initialized at the ELB of the yield curve (black line) and its components: expected inflation (red dots), expected real short rates (blue dots), the inflation risk premium (yellow dots), and the real term premium (green dots). Maturities are indicated in years.

Table 11 reports the results of this calculation at the steady state and for two initial values of the shadow rate below zero. The steady-state responses broadly line up with those estimated in the data above the ELB and those that we found using the simulated regressions in Table 9. Below the ELB, the responses are non-monotonic in the shadow rate. For modestly negative shadow-rate values, they are slightly smaller than in the steady state. But once the policy rate is expected to remain at zero for more than a few quarters, the responses increase dramatically. For a value of the shadow rate that roughly corresponds to an expected time until liftoff of six quarters, the responses of bond yields are 50% to 100% higher than their steady-state values, and the response of equity returns nearly triples, approaching the effects we observed in the data in Table 4. The response of equity implied volatility is also higher in this environment. For even more negative values of the shadow rate (not shown in the table) these coefficients rapidly explode in magnitude. Meanwhile, the response of 5-year bond volatilities (our proxy of the MOVE index) is not much different from the steady state, which again matches what we found in the data in Table 4.

Table 11: **Responses to One-Year Rate Volatility in Different Environments**

Initial i_t	5y nom. yield	5y real yield	10y nom. yield	10y real yield	Equity return	Std. eq. ret. (ann.)	Std. 5Y yld (ann.)
At steady state	1.0	2.0	0.5	1.0	-13.9	11.6	1.1
Below ELB 4 quarters	0.8	1.6	0.4	0.8	-12.1	7.6	0.5
Below ELB 6 quarters	1.5	4.2	0.8	2.1	-33.3	19.8	1.1

Notes: The table shows the model-implied responses of asset prices to a one-standard-deviation positive shock in MPR, divided by the response of the four-quarter-ahead conditional standard deviation of the short rate. The effects of the shocks on the shadow rate are neutralized with an offsetting contemporaneous monetary-policy shock. The model is initialized at values of the shadow rate at the steady state and at negative values sufficient to keep the short-term interest rate at zero for four or six quarters in the absence of shocks.

Intuitively, what drives these results is that, as the ELB is expected to bind for a longer period of time, there is less probability mass for the policy rate above zero at any given horizon. Thus, the conditional uncertainty about the one-year-ahead interest rate—the denominator in the ratios reported in the table—falls rapidly. Meanwhile, as shown in Figure 5, asset-price responses to a given-sized MPR shock (the numerators in the ratios) are generally larger at the ELB, reflecting the absence of a near-term policy response. Although we control for the latter effect by neutralizing changes in the shadow rate in this exercise, asset prices still only move a finite amount, while interest-rate volatilities go to zero as the shadow rate becomes deeply negative.

6 Case Study: MPR around the 2022 Liftoff

In this section, we present a recent case study of how changes to MPR can be important for asset prices. As we showed in Figure 1 in the Introduction, surveys of market participants imply that interest-rate uncertainty increased dramatically between year-end 2021 and year-end 2023 and that this increase could not be explained by an increase in macroeconomic uncertainty, strongly suggesting an increase in MPR. To gauge the effects of this episode on asset prices and the economy, we simulate a series of shocks that match this change in our model. This exercise complements recent work by Cieslak et al. (2026), who document an increase in what they call “reaction-function uncertainty” around this period using text analysis of newspaper articles.

Although they do not quantify the magnitude of the increase in risk or its exact contributions to asset returns, they estimate that the common component of risk premia in equities and nominal bonds rose considerably during this time, and they tie those movements to FOMC communications.

We begin by setting initial conditions to match the level of the federal funds rate and the one-year-ahead uncertainty about its value (based on Figure 1), as of Q4 2021. At this point, the fed funds rate was still at the ELB, so the shadow rate was negative. We set it to a value that implies an expected policy liftoff one quarter ahead, consistent with market expectations at the end of 2021. For simplicity, we assume that the value of the monetary-policy shock $\epsilon_{v,t}$ was initially at its steady-state value. We then simulate a series of eight quarterly shocks to η_t to generate the observed paths of the one-year-ahead uncertainty in the federal funds rate over the subsequent two years. Recall from Figure 1 that this uncertainty rose by 74 bp during the two-year period. (Since we only have the survey data on interest-rate uncertainty at year ends, we interpolate to obtain quarterly series.) We compare the outcomes in this simulation to a counterfactual economy in which there were no η_t shocks and MPR simply drifted back to its steady state from its Q4 2021 value.³³

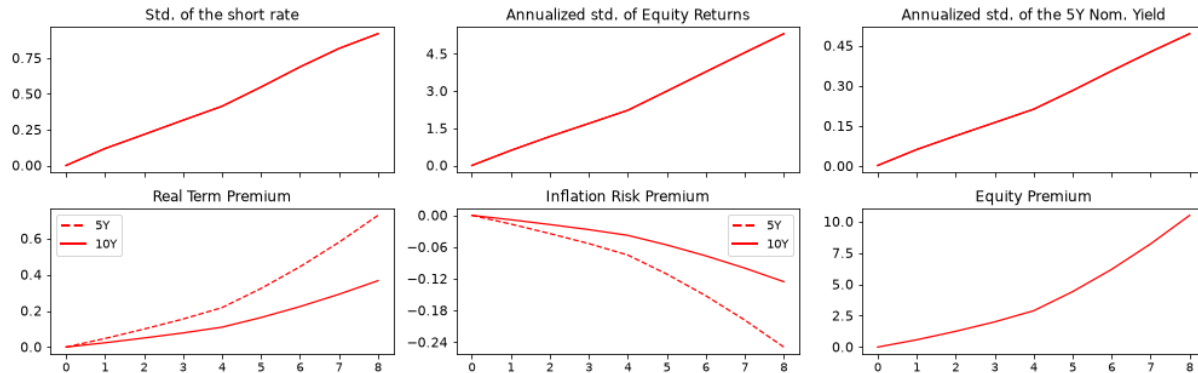
Figure 7 shows the results. The conditional volatilities of interest rates, bond yields, and equity prices rise monotonically over the period. For example, our proxy for the VIX rises by about 5 percentage points. These changes drive significant increases in risk premia. 5- and 10-year real term premia rise by about 70 and 35 bp in total, and the equity premium rises by about 10 percentage points at an annualized rate. The inflation risk premium falls by 25 and 12 bp at the 5- and 10-year maturities, offsetting part of the effect of real term premia on the nominal yield curve.³⁴

The overall effects of the rise in MPR on stock and bond prices depend not only on these changes in risk premia but also on the path of cash flows and short-term interest rates. As shown in Figure 8, real bond yields rise about 50 bp and 24 bp at the 5- and 10-year maturities,

³³In doing this simulation, we re-calibrate the steady-state values π^* and r^* to 2.5% and 1%, corresponding roughly to their post-2020 averages. This allows the path of short-term rates in the model to better match the observed upward trajectory and prevents the MPR shocks from counterfactually sending the economy back to the ELB.

³⁴The net effects on nominal yields appear to be in the ballpark of those estimated by Cieslak et al. (2026), although differences in methodology and the timing of observations makes it difficult to compare precisely.

Figure 7: **Simulated Responses to MPR Rise in 2022-23: Volatility and Risk Premia**

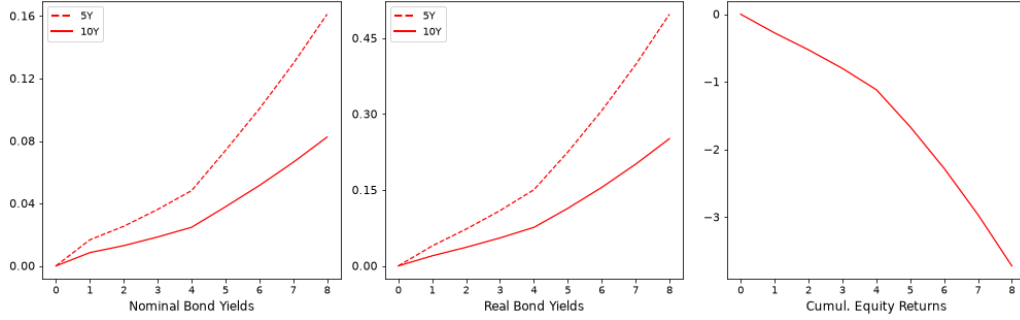


Notes: The figure shows the simulated cumulative changes in conditional volatilities and risk premia when the model is subjected to MPR shocks that match the increase in one-year-ahead interest-rate volatility between December 2021 and December 2023, as depicted in Figure 1. The horizontal axes measure time in quarters since the initial period.

and nominal yields rise about 17 and 8 bp. Meanwhile, stock prices fall about 4% cumulatively. As noted earlier, our baseline model implies a significant amount of monetary-policy easing in response to a positive MPR shock, and that response offsets a considerable portion of the effect of increases in risk premia on asset prices. (Here, the policy response manifests as a path of short-term rates that rises less aggressively than would have been the case in the absence of the increase in MPR.) The model suggests that, without this response, 10-year real yields would have risen an additional 35 bp and stock returns would have declined an additional 8% over these two years (not shown in the figure).

We note that this exercise is likely to *understate* the size and effects of the MPR shock during this period. This is because, in the model, uncertainty about output and inflation rise in response to the MPR shock, as shown in the second row of the table. In order for macroeconomic uncertainty to remain approximately constant, as Figure 1 suggests it did, some other structural source of risk must have declined at the same time the MPR was rising. In particular, as the disruptions resulting from the COVID-19 pandemic receded and supply chains were repaired, a number of contemporary reports suggested a decline in supply-side uncertainty. Since these forces were also endogenously reducing uncertainty about interest rates, it must be the case that the increase in MPR was even larger than Figure 1 suggests. Thus, our results may be

Figure 8: **Simulated Responses to MPR Rise in 2022-23: Asset Prices**



Notes: The figure shows the simulated cumulative changes in bond yields and equity values when the model is subjected to MPR shocks that match the increase in one-year-ahead interest-rate volatility between December 2021 and December 2023, as depicted in Figure 1. The horizontal axis measures time in quarters since the initial period.

interpreted as a lower bound on the effects of the MPR shocks during this period.

It is possible to formulate alternative explanations for the rise in interest-rate uncertainty during this period. In particular, Bauer et al. (2024a) argue that the market's perception of the inflation coefficient in the policy rule (ϕ_π in our notation) rose after the pandemic. If so, that would imply a higher level of interest rate uncertainty for any given amount of uncertainty in inflation, all else equal. However, the changes documented by Bauer et al. are not quantitatively large enough to explain the increase in uncertainty shown in Figure 1. Based on their estimates, we can perform the following calculation. As of the end of 2021, both policy-rule coefficients were approximately zero (see their Figure 2). Thus, MPR must have accounted for the entirety of interest-rate uncertainty, which our survey evidence shows was 33 bp at the one-year horizon. At the end of 2023, Bauer et al. estimate approximately $\phi_\pi = 1$ and $\phi_y = 0.3$, using annualized rates. At the same time, the survey measure of interest-rate uncertainty was 1.07%, one-year inflation uncertainty was 0.53%, and output uncertainty was 1.1%. Approximating the covariance between inflation and output as zero and disregarding

inertial terms (which are small at the one-year horizon), we have³⁵

$$\text{var}_t[i_{t+4}] \approx 1.07^2 = 1^2 0.53^2 + 0.3^2 1.1^2 + MPR_{t+4}^2 \quad (39)$$

This produces a contribution of MPR to the standard deviation of the one year ahead interest rate of 0.99%. In other words, even after accounting for a plausible increase in the perceived policy-rule coefficients, we still attribute $99 - 33 = 66$ basis points of the increase in one-year interest-rate uncertainty to higher MPR. Again, this value would be even larger if we were able to properly account for declining uncertainty about other shocks in the economy during this time.

We conclude that an increase in uncertainty about the monetary-policy reaction function contributed to the substantial decline in asset prices during the post-COVID policy tightening. For comparison, in the data, 10-year nominal bond yields rose about 240 basis points over this period, while 10-year TIPS yields rose about 270 basis points. Although some of that rise certainly had to do with the realized and expected tightening of monetary policy, measures of term premia, including those of Kim and Wright (2005), Adrian et al. (2013), and D’Amico et al. (2018) also rose significantly, as our model suggests they should have. The S&P 500 index of equity prices trailed its average two-year return by about 25 percentage points. Of course, many other shocks affected the economy during this period, so we cannot be more precise in comparing the simulation to the data. But, the analysis shows that the increase in MPR is directionally consistent with, and quantitatively accounts for a significant portion of, the net changes in asset prices that we observed in 2022 and 2023.

7 Conclusion

We have shown how risk surrounding the exogenous component of monetary policy affects premia, volatility, and returns in equities and real and nominal bonds. Empirically, this kind of

³⁵We use headline PCE inflation, which is all that is reported in the primary dealer survey at the one-year horizon. The calculation conservatively assumes that all uncertainty about GDP one year ahead is due to the output gap, not to potential output. If inflation and GDP were viewed as negatively correlated, as seems plausible during this period with prevalent supply-side dynamics, that would increase the calculated contribution of MPR even further.

risk contributes significantly to the observed return volatility across asset classes and generates sizable price movements in event studies. The qualitative effects of MPR on risk premia, asset prices, and macroeconomic aggregates can be rationalized in a broad class of models where money is not neutral. Quantitatively, in a benchmark New Keynesian model with Epstein-Zin preferences, MPR explains significant portions of the empirical values of the equity premium and nominal term premium, largely accounts for the real term premium, and contributes negatively to the inflation-risk premium. At business-cycle frequencies, an increase in monetary-policy risk leads to declines in output and inflation; a rise in the equity premium; a fall in stock prices; an increase in real and nominal term premia increase; a steeper nominal yield curve; positive excess returns on nominal bonds and negative abnormal returns on real bonds; and a lower inflation-risk premium. Most of these effects are quantitatively significant. The ELB increases the magnitudes of these responses, also in line with the data.

We have not analyzed the welfare implications of MPR, but, in our model, the optimal size of monetary-policy shocks is necessarily zero. Uncertainty has a direct negative effect on utility and distorts consumption relative to the risk-free case. From a financial-market perspective, a high and variable level of MPR weighs on asset prices, causes financial conditions to tighten, and obscures the signals that asset prices convey about the economy. Our results thus highlight several ways in which clarity in monetary-policy communication may be of value.

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APPENDICES

A Proofs of Propositions

A.1 Proof of proposition 1

Taking a first-order approximation of asset returns around the steady state,

$$R_t = f(v_t, \epsilon_{x,t}, \zeta_t) - f(v_{t-1}, \epsilon_{x,t-1}, \zeta_{t-1}) \quad (40)$$

$$\approx \frac{\partial f}{\partial v_t}[(\rho_v - 1)v_{t-1} + \zeta_{t-1}\epsilon_{v,t}] + \frac{\partial f}{\partial \epsilon_{x,t}}\epsilon_{x,t} + \frac{\partial f}{\partial \zeta_t}[(\rho_\zeta - 1)(\zeta_{t-1} - \bar{\zeta}) + \eta_t] \quad (41)$$

Taking macroeconomic shocks to be zero and assuming $\rho_v \approx \rho_\zeta \approx 1$, this simplifies to

$$R_t|_{\epsilon_{x,t}=0} \approx \frac{\partial f}{\partial v_t}(\zeta_{t-1}\epsilon_{v,t}) + \frac{\partial f}{\partial \zeta_t}\eta_t \quad (42)$$

so

$$\text{var}[R_t|_{\epsilon_{x,t}=0}] \approx \left(\frac{\partial f}{\partial v_t}\right)^2 \text{var}[\zeta] + \left(\frac{\partial f}{\partial \zeta_t}\right)^2 \sigma_t^2 \quad (43)$$

Meanwhile, on blackout days,

$$\text{var}[R_t|_{\epsilon_{v,t}=\eta_t=0}] \approx \left(\frac{\partial f}{\partial \epsilon_{x,t}}\right)^2 \sigma_x^2 \quad (44)$$

where σ_x^2 is the variance of $\epsilon_{x,t}$. Since the three shocks are independent, equation (7) follows.

A.2 Asset Pricing Derivations

We define the operator $\mathcal{E}_t^{(N)}$ as the time- t log expectation of the N -period geometric mean. That is, for any random variable x_t ,

$$\mathcal{E}_t^{(N)}[x_s] \equiv \log E_t \left[\exp \left(\frac{\sum_{n=1}^N x_{s+n-1}}{N} \right) \right]. \quad (45)$$

We work with geometric expectations, rather than the arithmetic expectations because they subsume Jensen's inequality terms, which are not of primary interest, into physical expectations of future prices and cash flows and leave risk premia as clean measures of deviations from risk neutrality. Under the assumption of log normality that we maintain in this section,

$$\mathcal{E}_t^{(N)}[x_s] = E_t \left[\frac{1}{N} \sum_{n=1}^N x_{s+n-1} \right] + \frac{1}{2} \text{var}_t \left[\frac{1}{N} \sum_{n=1}^N x_{s+n-1} \right]. \quad (46)$$

From equation (10), the short-term real interest rate is equal to the negative log expectation of the one-period SDF, as is standard:

$$r_t = -E_t[m_{t+1}] - \frac{\sigma_m^2}{2} = -\mathcal{E}_t^{(1)}[m_{t+1}] \quad (47)$$

Similarly, the nominal interest rate satisfies

$$i_t = -\mathcal{E}_t^{(1)}[m_{t+1} - \pi_{t+1}] \quad (48)$$

$$= r_t + \mathcal{E}_t^{(1)}[\pi_{t+1}] + \underbrace{\text{cov}_t[m_{t+1}, \pi_{t+1}]}_{IRP^{(1)}}. \quad (49)$$

In equilibrium, the real interest rate adjusts such that both the policy rule (1) and the no-arbitrage condition (49) are always satisfied.

The two period real yield is given by:

$$y_t^{(2)} = -\frac{1}{2} \log E_t[\exp(m_{t+1} + m_{t+2})] \quad (50)$$

$$= \mathcal{E}_t^{(2)}[r_t] + \underbrace{\text{cov}_t[m_{t+1}, r_{t+1}]}_{RTP_t^{(2)}}. \quad (51)$$

The two-period nominal yield can be written as

$$y_t^{(2)} = \mathcal{E}_t^{(2)}[i_t] + \text{cov}_t[m_{t+1} - \pi_{t+1}, i_{t+1}] \quad (52)$$

$$= \mathcal{E}_t^{(2)}[r_t] + \mathcal{E}_t^{(2)}[\pi_t] + RTP_t^{(2)} + \underbrace{IRP_t^{(1)} + \text{cov}_t[\pi_{t+1}, m_{t+2}] + \text{cov}_t[\pi_{t+2}, m_{t+1}] + \text{cov}_t[\pi_{t+2}, m_{t+2}]}_{IRP_t^{(2)}} \quad (53)$$

The stock price is

$$P_t^{eq} = E_t[\exp(m_{t+1} + y_{t+1})] \quad (54)$$

The log nominal return on this asset is given by

$$r_{t+1}^{eq} + \pi_{t+1} = y_{t+1} - \mathcal{E}_t^{(1)}[m_{t+1} - \pi_{t+1} + y_{t+1}] \quad (55)$$

A.3 Proof of proposition 2

The variance of nominal interest rates is given by

$$\begin{aligned} \text{var}[i_{t+1}] &= \phi' \text{var}[\mathbf{x}_{t+1}] \phi + \zeta \phi' \text{cov}_t[\mathbf{x}_{t+1}, \epsilon_{v,t+1}] + \zeta^2 \\ &= \phi' (\zeta^2 \mathbf{k}' \mathbf{k} + \mathbf{A}' \mathbf{A}) \phi + \zeta^2 (2\phi' \mathbf{k} + 1) \end{aligned}$$

making use of the normalizations that all shocks are uncorrelated and have unit variance. Because we have assumed $\phi' \mathbf{k} > -1$ (ensuring that the policy rate moves in the same direction as the monetary shock), the variance of nominal short rates unambiguously rises with MPR. Meanwhile, from equation (49), the real short rate is

$$r_t = i_t - \mathcal{E}_t^{(1)}[\pi_{t+1}] - \text{cov}_t[m_{t+1}, \pi_{t+1}] \quad (56)$$

$$= i_t - E_t[\pi_{t+1}] + \frac{1}{2} \text{var}[\pi_{t+1}] - \text{cov}_t[m_{t+1}, \pi_{t+1}] \quad (57)$$

So $\text{var}_t[r_{t+1}] = \text{var}_t[i_{t+1}] + \text{var}_t[E_{t+1}[\pi_{t+2}]] - \text{cov}_t[i_{t+1}, E_{t+1}[\pi_{t+2}]]$. This gives

$$\text{var}_t[r_{t+1}] = \phi' (\zeta^2 \mathbf{k}' \mathbf{k} + \mathbf{A}' \mathbf{A}) \phi + \zeta^2 (2\phi' \mathbf{k} + 1) + k_\pi^2 \rho^2 \zeta^2 - k_\pi \zeta^2 \quad (58)$$

Since k_π is negative by assumption (positive monetary-policy shocks reduce inflation), higher levels of MPR are associated with a higher variance of short-term real interest rates.

Similarly, with constant second moments the variance of nominal bond yields is

$$\text{var}_t[y_{t+1}^{(2)}] = \text{var}_t[i_{t+1}] + \text{var}_t[E_{t+1}[i_{t+2}]] + 2\text{cov}_t[i_{t+1}, i_{t+2}] \quad (59)$$

$$= \phi' (\zeta^2 ((1 + \rho)^2 + 1) \mathbf{k}' \mathbf{k} + \mathbf{A}' \mathbf{A}) \phi + \zeta^2 ((1 + \rho)^2 + 1) (\phi' \mathbf{k} + 1) + 2\zeta^2 \rho \quad (60)$$

and the variance of real yields is

$$\text{var}_t[y_{t+1}^{(2)}] = \text{var}_t[r_{t+1}] + \text{var}_t[r_{t+2}] + 2\text{cov}_t[r_{t+1}, r_{t+2}] \quad (61)$$

$$= \text{var}_t[i_{t+1}] + \text{var}_t[E_{t+1}[i_{t+2}]] + \text{var}_t[E_{t+1}[\pi_{t+2}]] + \text{var}_t[E_{t+1}[\pi_{t+3}]] \quad (62)$$

$$+ 2(\text{cov}_t[i_{t+1}, E_{t+1}[i_{t+2}]] - \text{cov}_t[i_{t+1}, E_{t+1}[\pi_{t+2}]] - \text{cov}_t[E_{t+1}[i_{t+2}], E_{t+1}[\pi_{t+3}]]) \quad (63)$$

$$+ 2(-\text{cov}_t[i_{t+1}, E_{t+1}[\pi_{t+3}]] - \text{cov}_t[E_{t+1}[i_{t+2}], E_{t+1}[\pi_{t+2}]] + \text{cov}_t[E_{t+1}[\pi_{t+2}], E_{t+1}[\pi_{t+3}]])) \quad (64)$$

$$= \phi' (\zeta^2 \mathbf{k}' \mathbf{k} + \mathbf{A}' \mathbf{A}) \phi + \zeta^2 (2\phi' \mathbf{k} + 1) + \rho^2 \zeta^2 \quad (65)$$

$$+ k_\pi^2 \rho^2 \zeta^2 + k_\pi^2 \rho^4 \zeta^2 + 2(\rho \zeta^2 - k_\pi \rho \zeta^2 - k_\pi \rho^3 \zeta^2 - 2k_\pi \rho^2 \zeta^2 + k_\pi^2 \rho^3 \zeta^2) \quad (66)$$

Again, recalling that $k_\pi < 0$ and $\phi' k_\pi > -1$, both variances are strictly increasing in ζ .

Turning to the variances of asset returns, for an N -period nominal bond, the log price is, by definition, $p_t^{(N)} = -N y_t^{(N)}$, and the one-period nominal return is $p_{t+1}^{(N-1)} - p_t^{(N)}$. Thus, the variance of the return is

$$\text{var}_t[-i_{t+1} + 2y_t^{(2)}] = \phi' (\zeta^2 \mathbf{k}' \mathbf{k} + \mathbf{A}' \mathbf{A}) \phi + \zeta^2 \quad (67)$$

which is increasing in ζ . For inflation-indexed bonds, the analogous calculation is

$$\text{var}_t[-r_{t+1} + 2y_t^{(2)} + \pi_{t+1}] = \phi' (\zeta^2 \mathbf{k}' \mathbf{k} + \mathbf{A}' \mathbf{A}) \phi + 2\phi' \mathbf{k} (1 - \rho) k_\pi \zeta^2 \quad (68)$$

$$+ k_\pi^2 (\rho - 1)^2 \zeta^2 \quad (69)$$

Again, this is increasing in ζ under our parameter restrictions.

From equation (55), the conditional variance of the return on the two-period equity claim is simply

$$\text{var}_t[y_{t+1}] = k_y^2 \zeta^2 + (a_y^d \sigma_d)^2 + (a_y^s \sigma_s)^2 \quad (70)$$

which is clearly increasing in ζ .

From equation (12), the effect of MPR on the real term premium is $(1 + \phi' \mathbf{k} - \kappa_\pi) \mu^v$, which is strictly positive. From equation (53), the two-period inflation risk premium is

$$IRP_t^{(2)} = IRP_t^{(1)} + \text{cov}_t[\pi_{t+1}, m_{t+2}] + \text{cov}_t[\pi_{t+2}, m_{t+1}] + \text{cov}_t[\pi_{t+2}, m_{t+2}]$$

and, from equation (11), we have $IRP_t^{(1)} = a_\pi^s \mu^s + a_\pi^d \mu^d + \kappa_\pi \mu^v \zeta_{t-1}^2$. Meanwhile it is straightforward to show

$$\text{cov}_t[\pi_{t+1}, m_{t+2}] = 0 \quad (71)$$

$$\text{cov}_t[\pi_{t+2}, m_{t+1}] = \kappa_\pi \rho \mu^v \zeta^2 \quad (72)$$

$$\text{cov}_t[\pi_{t+2}, m_{t+2}] = \mu^d a^d + \mu^s a^s + \kappa_\pi \mu^v \zeta^2 \quad (73)$$

Consequently, the effect of an increase in ζ^2 on $IRP_t^{(2)}$ is $\kappa_\pi \mu^v (\rho + 2)$, which is strictly negative.

The two-period nominal term premium is

$$NTP_t^{(2)} = \text{cov}_t[m_{t+1} - \pi_{t+1}, i_{t+1}] \quad (74)$$

$$= (\phi_y a_y^d + \phi_\pi a_\pi^d)(\mu^d - a_\pi^d) + (\phi_y a_y^s + \phi_\pi a_\pi^s)(\mu^s - a_\pi^d) \quad (75)$$

$$+ (1 + \phi' \mathbf{k})(\mu^v - k_\pi) \zeta_t^2 \quad (76)$$

Since $\mu^v > 0$ and $k_\pi < 0$, the effect of ζ is strictly positive.

Finally, the equity premium in the two-period log-normal model is given in equation (15). Since $\mu^v > 0$ and $k_y < 0$, the effect of ζ in this equation is strictly positive.

B Robustness Checks of Empirical Results

This appendix reports the robustness of the empirical results to various changes in the sample construction. First, in the baseline results, we drop all observations between September 2008 and March 2009, corresponding to the most acute phase of the global financial crisis and the Fed's interventions, and March 2020, the month with the extraordinary financial turmoil (and Fed response) associated with the Covid-19 pandemic. Here, we include those observations and reproduce Tables 1 and 3.

Table 12: **Comparison of Volatility on Blackout and Non-Blackout Days: Including Crisis Periods**

	Blackout	Non-blackout			Implied qtrly std. dev. due to MPR
		All non- blackout days	FOMC days	Non-FOMC days	
5Y nom. yield	5.3	5.6***	6.5***	5.5**	13.8
10Y nom. yield	5.2	5.4***	6.3***	5.4*	11.3
5Y real yield	5.2	5.4***	6.9***	5.2*	14.4
10Y real yield	4.4	4.6***	5.5***	4.5	9.7
SPX	94	98***	111***	96	197
VIX	125	129	140	128	226
MOVE	3.2	3.5***	4.2***	3.5***	10.4
Sample Size	1663	6189	496	5693	

Notes: The table compares the standard deviations of daily changes in various asset prices on days that fall during blackout periods to days that do not over the period 1994 - 2026, except for real yields, which begin in 2003. All values are reported in units of basis points. This sample includes data from September 2008 - March 2009 and March 2020. Data-cleaning restrictions are described in the text. Asterisks indicate statistically significant differences between blackout and non-blackout days at the 10% (*), 5% (**), and 1% (***) levels, based on one-sided bootstrapped confidence bounds. The final column shows the implied contribution of monetary-policy risk to the quarterly standard deviation of each asset, calculated by scaling the variances up by the average number of non-blackout business days per quarter.

Table 13: **Responses around FOMC Announcements: Including Crisis Periods**

	5Y nom. yield	10Y nom. yield	5Y real yield	10Y real yield	SPX	VIX	MOVE
Monetary-policy shock	0.43*** (0.07)	0.23*** (0.07)	0.65*** (0.11)	0.39*** (0.10)	-2.45* (1.27)	0.90 (1.92)	-0.05 (0.04)
Short-rate uncertainty (MPR)	0.82*** (0.15)	0.83*** (0.14)	1.52*** (0.22)	1.29*** (0.20)	-11.8** (2.73)	25.6*** (4.13)	0.90*** (0.10)
Adj. R^2	0.30	0.22	0.40	0.30	0.12	0.16	0.27
# obs.	255	255	178	178	255	255	254

Notes: Daily data on days of FOMC announcements, 1994 – 2026, except for real yields, which begin in 2003. This sample includes data from September 2008 - March 2009 and March 2020. Monetary-policy shocks are measured as changes in one-year eurodollar rates and MPR shocks are measured as changes in one-year federal-funds implied volatility. All values are reported in basis points. Standard errors in parentheses. Asterisks indicate statistical significance at the 10% (*), 5% (**), and 1% (***) level.

Table 14: **Comparison of Volatility on Blackout and Non-Blackout Days: Dropping Days After FOMC**

	Blackout	Non-blackout			Implied qtrly std. dev. due to MPR
		All non- blackout days	FOMC days	Non-FOMC days	
5Y nom. yield	5.2	5.4**	6.2***	5.4**	10.5
10Y nom. yield	5.0	5.3**	5.6***	5.3**	10.9
5Y real yield	4.6	4.9***	6.5***	4.9**	12.2
10Y real yield	4.2	4.4**	5.0**	4.4**	10.2
SPX	90	92	103**	92	135
VIX	120	122	124	121	156
MOVE	3.1	3.4***	4.0***	3.4***	9.6
Sample Size	1663	5945	252	5693	

Notes: The table compares the standard deviations of daily changes in various asset prices on days that fall during blackout periods to days that do not over the period 1994 - 2026, except for real yields, which begin in 2003. All values are reported in units of basis points. We drop all days immediately following FOMC announcements. Other data-cleaning restrictions are described in the text. Asterisks indicate statistically significant differences between blackout and non-blackout days at the 10% (*), 5% (**), and 1% (***) levels, based on one-sided bootstrapped confidence bounds. The final column shows the implied contribution of monetary-policy risk to the quarterly standard deviation of each asset, calculated by scaling the variances up by the average number of non-blackout business days per quarter.

Next, in the baseline statistics calculated in Tables 1 and 2, we include the days after FOMC statement days in our “FOMC day” subsample. As noted in the text, this choice reflects the idea that it may take markets more than a couple of hours to digest the information released by the FOMC, and it is consistent with the observation that some asset classes are more volatile on days *after* FOMC announcements than they are on the days *of*. Still, one may question whether this window is too large to accurately capture the relevant volatility. Here, we re-calculate the numbers eliminating the days after FOMC meetings from the sample. Note that these criteria only apply to the blackout/non-blackout calculations, not to the event-study regressions in Tables 3 and 4.

C Equilibrium Conditions in the NK Model

C.1 Households

Denoting Lagrange multipliers by Λ_t , the first-order conditions for the household's problem of maximizing (18), subject to (19), are

$$\frac{\partial U_t}{\partial C_t} = (1 - \beta)C_t^{\rho-1}U_t^{1-\rho} = -P_t\Lambda_t \quad (77)$$

$$\frac{\partial U_t}{\partial H_t} = -(1 - \beta)\theta_t\chi H_t^{\chi\rho-1}U_t^{1-\rho} = W_t\Lambda_t \quad (78)$$

$$Q_t = E_t \left[\frac{\Lambda_{t+1}}{\Lambda_t} \right] \quad (79)$$

These imply the standard relationships

$$Q_t = E_t \left[M_{t+1} \frac{P_t}{P_{t+1}} \right] \quad (80)$$

$$\frac{W_t}{P_t} = -\frac{\partial U_t / \partial H_t}{\partial U_t / \partial C_t} \quad (81)$$

Where $M_{t+1} = \frac{\partial U_t / \partial C_{t+1}}{\partial U_t / \partial C_t}$ is the real stochastic discount factor. Equation (21) follows immediately from (81).

Let C^* denote the trend value of consumption, which follows a deterministic log-linear process with growth rate g . Let $\tilde{C}_t = C_t / C_t^*$ denote the percentage deviation from this trend. Equation (22) then follows immediately from (77) and (80).

Adopting similar notation, define $\tilde{V}_t^C = V_t^C / V_t^{C*}$. Expanding V_t^C using this decomposition,

$$V_t^C = \tilde{V}_t^C V_t^{C*} = \left[(1 - \beta)(\tilde{C}_t C_t^*)^\rho + \beta \mathcal{R}_t \left(\tilde{V}_{t+1}^C V_{t+1}^{C*} \right)^\rho \right]^{1/\rho} \quad (82)$$

Note that in the steady state this simplifies to

$$(V_t^{C*})^\rho = (1 - \beta)(C_t^*)^\rho + \beta(V_{t+1}^{C*})^\rho \quad (83)$$

Iterating forward one period,

$$(V_t^{C*})^\rho = (1 - \beta)(C_t^*)^\rho + \beta \left[(1 - \beta)(C_t^* \exp[g])^\rho + \beta(V_{t+2}^{C*})^\rho \right]$$

Continuing in this way to the limit,

$$(V_t^{C*})^\rho = (1 - \beta)(C_t^*)^\rho \sum_{h=0}^{\infty} \beta^h \exp[hg\rho]$$

or, calculating the infinite sum,

$$V_t^{C*} = \left(\frac{1 - \beta}{1 - \beta \exp[g\rho]} \right)^{1/\rho} C_t^*$$

This implies that $\log V_t^{C*}$ has the same growth rate as c_t^* , namely g . So we can rewrite (83) as

$$(V_t^{C*})^\rho = (1 - \beta)(C_t^*)^\rho + \beta \exp[g\rho](V_t^{C*})^\rho$$

or

$$V_t^{C*} = \left[\frac{(1 - \beta)(C_t^*)^\rho}{1 - \beta \exp[g\rho]} \right]^{\frac{1}{\rho}}$$

Substituting this back into (82) and rearranging produces equation (23).

C.2 Firms

Output is given by the aggregation of differentiated goods indexed on the unit interval:

$$Y_t = \left(\int_0^1 Y(i)_t^{1-1/\epsilon} di \right)^{\epsilon/(\epsilon-1)} \quad (84)$$

with corresponding price index

$$P_t = \left(\int_0^1 P(i)_t^{1-\epsilon} di \right)^{1/(1-\epsilon)} \quad (85)$$

Demand functions for each good are given by

$$C(i)_t = \left(\frac{P(i)_t}{P_t} \right)^{-\epsilon} Y_t \quad (86)$$

As is standard, the firm's problem, log-linearized around the non-stochastic steady state, results in the following recursive solution for inflation:

$$\pi_t = (1 - \vartheta)(\log P_t^{opt} - \log P_{t-1}) \quad (87)$$

where $1 - \vartheta$ is the probability of price adjustment and P_t^{opt} the firm's optimal reset price. Firms choose $P(i)_t^{opt}$ to maximize the expected discounted value of profits, taking demand as given. For a firm resetting its price in period t , this problem is

$$\max_{P_t^{opt}} \sum_{s=1}^{\infty} \vartheta^s E_t \left[\frac{M_{t+s}}{P_{t+s}} \left(P_t^{opt} Y(i)_{t+s} - \frac{W_{t+s}}{(1 - \psi) A_{t+s} H(i)_{t+s}^{-\psi}} \right) \right] \quad (88)$$

subject to

$$Y(i)_{t+s} = \left(\frac{P_t^{opt}}{P_{t+s}} \right)^{-\epsilon} Y_{t+s} \quad (89)$$

In the non-stochastic steady state with Epstein-Zin preferences and constant trend growth, the one-period SDF is equal to $M^* = \beta \exp g(\rho - 1)$. The firm's problem thus results in the following log-linear condition for inflation:

$$\pi_t = \beta \exp[g(\rho - 1)] E_t[\pi_{t+1}] - \lambda \hat{\mu}_t,$$

where

$$\hat{\mu}_t = -\log\left(\frac{W_t}{P_t}\right) + a_t - \psi \log H_t + \log(1 - \psi) - \log(M^*)$$

We can thus write steady-state inflation as

$$\pi^* = -\frac{\lambda \hat{\mu}^*}{1 - \beta \exp[g(\rho - 1)]}$$

This assumes that markups have a constant steady-state value (no trend). It is straightforward to verify this claim by noting that $\hat{\mu}_t$ in the steady-state is given by:

$$\hat{\mu}_t^* = \left(\frac{(\chi\rho)}{(1-\psi)} \right) a_t - \left(\frac{(\chi\rho)}{(1-\psi)} - \rho \right) c_t^* - \log\left(\frac{\theta_t \chi M}{1-\psi}\right)$$

Given the trend growth rates of a , c , and $\log \theta$ we derived earlier, the trend growth rate of $\hat{\mu}_t$ is

$$\hat{\mu}_{t+1}^* - \hat{\mu}_t^* = \left(\frac{(\chi\rho)}{(1-\psi)} \right) g - \left(\frac{(\chi\rho)}{(1-\psi)} - \rho \right) g - \rho g = 0$$

So $\hat{\mu}^*$ is indeed constant (and therefore so is π^*).

Finally, we can rewrite the inflation equation as

$$\pi_t = \beta \exp[g(\rho - 1)] E_t[\pi_{t+1}] - \lambda \left(\frac{(\chi\rho)}{(1-\psi)} - \rho \right) \tilde{y}_t$$

or, linearizing around π^* ,

$$\pi_t = \pi^* + \beta \exp[g(\rho - 1)] E_t[\pi_{t+1} - \pi^*] + \kappa \tilde{y}_t$$

where κ is a reduced-form parameter.

D Solution Method

A solution in our model consists of a set of rules relating the state of the economy to each of our recursively-determined endogenous variables (consumption, inflation and lifetime utility) which satisfy the equilibrium conditions. Specifically, let N_s be the number of continuous state variables summarizing the state of the economy and $\mathbf{S}_t \in \mathbf{R}^{N_s}$ denote the state of the economy. Then a solution consists of mappings

$$f_V : \mathbf{R}^{N_s} \rightarrow \mathbf{R} \text{ s.t. } \tilde{V}_t^C = f_V(\mathbf{S}_t), \quad (90)$$

$$f_y : \mathbf{R}^{N_s} \rightarrow \mathbf{R} \text{ s.t. } \tilde{y}_t = f_y(\mathbf{S}_t), \quad (91)$$

$$f_\pi : \mathbf{R}^{N_s} \rightarrow \mathbf{R} \text{ s.t. } \pi_t = f_\pi(\mathbf{S}_t) \quad (92)$$

that satisfy the equilibrium conditions (21), (22) and (23), given the interest-rate rule (16), the Philips Curve (24), and the assumed dynamics of the exogenous variables η_t and $\epsilon_{v,t}$.

To find the functions f_V , f_y , and f_π , we employ time iteration (Judd, 1996), computing expectations via quadrature, following Tauchen and Hussey (1991). To do this, we define uniformly spaced grids over possible values of ν_t and ζ_t , $\mathbf{g}_\nu = \{\nu^1, \nu^2, \dots, \nu^{N_\nu}\}$ and $\mathbf{g}_\zeta = \{\zeta^1, \zeta^2, \dots, \zeta^{N_\zeta}\}$, and we build our discrete state space $\hat{\mathbf{S}}$ as their Cartesian product:

$$\hat{\mathbf{S}} = \{(\nu^i, \zeta^j) | \nu^i \in \mathbf{g}_\nu, \zeta^j \in \mathbf{g}_\zeta\}$$

Because ζ is a standard deviation, the nodes of $\mathbf{g}_\zeta$ are all positive values. Meanwhile, $\mathbf{g}_\nu$ is centered at zero. The set $\hat{\mathbf{S}}$ has length $N_\nu \times N_\zeta$.

Discrete transition probabilities from an arbitrary node $\hat{\mathbf{S}}, (\nu_i, \zeta_j) \in \hat{\mathbf{S}}$ to the node $(\nu_k, \zeta_l) \in \hat{\mathbf{S}}$ are approximated as

$$q_{i,j}(k,l) = \Pr \left[\mathbf{S}_{t+1} = (\nu^k, \zeta^l) | \mathbf{S}_t = (\nu^i, \zeta^j) \right] \quad (93)$$

$$\approx \frac{1}{\Phi_{i,j}} \varphi_\nu(\nu_k | (\nu^i, \zeta^j)) \varphi_\zeta(\zeta^l | \zeta^j) \quad (94)$$

where $\varphi_\nu(\cdot | \cdot)$ is the continuous probability density function of ν_t conditional on ν_{t-1} and ζ_{t-1} (implied by the heteroskedastic autoregressive process (4)), $\varphi_\zeta(\cdot | \cdot)$ is the continuous probability density function of ζ_t conditional on ζ_{t-1} (given by the truncated autoregressive process stated in footnote 19), and $\Phi_{i,j}$ is the normalizing constant

$$\Phi_{i,j} = \sum_{k,l} \varphi_\nu(\nu_k | (\nu^i, \zeta^j)) \varphi_\zeta(\zeta^l | \zeta^j)$$

Equation (94) follows from the independence of η_t and $\epsilon_{v,t}$. The approximation error converges to zero as $N_\nu \rightarrow \infty$ and $N_\zeta \rightarrow \infty$. We collect the transition probabilities in the vectors $\mathbf{q}_{i,j} = (q_{i,j}(1,1), \dots, q_{i,j}(N_\nu, N_\zeta))$.

Because we have inertia in the policy rule (equation (16)) our exogenous state space $\hat{\mathbf{S}}$ does not fully summarize the state of the economy, which depends critically on $\hat{i}_{t-1}$. We therefore extend the state space by taking the Cartesian product of $\hat{\mathbf{S}}$ and a grid over possible values of the lagged shadow rate $\mathbf{g}_{\hat{i}} = \{\hat{i}^1, \hat{i}^2, \dots, \hat{i}^{N_{\hat{i}}}\}$. Note that when the grids for ν and ζ are associated with time- t values, the grid for $\hat{i}$ will stand in for $t-1$ values of the shadow rate. Also note that the transition probabilities $q_{i,j}$ are independent of the value of $\hat{i}$.

We begin the iteration by guessing an initial value for each of the f functions across all values of $\hat{\mathbf{S}}$ for each value of $\mathbf{g}_{\hat{i}}$, denoted $\hat{f}_y^n(\nu^i, \zeta^j, \hat{i}^k)$, $\hat{f}_V^n(\nu^i, \zeta^j, \hat{i}^k)$, and $\hat{f}_\pi^n(\nu^i, \zeta^j, \hat{i}^k)$. We linearly interpolate across these values to obtain continuous approximating functions, and we stack these functions in the vectors $\mathbf{F}_y^0(\hat{i})$, $\mathbf{F}_V^0(\hat{i})$, and $\mathbf{F}_\pi^0(\hat{i})$, each of which has dimension $N_\nu \times N_\zeta$.

For any iteration n , we first use the policy rule (1) to calculate

$$\iota_{i,j,k}^n = (1 - \varphi_i) \left[\pi^* + r^* + \phi_y \hat{f}_y^{n-1}(\nu^i, \zeta^j, \hat{i}^k) + \phi_\pi (\hat{f}_\pi^{n-1}(\nu^i, \zeta^j, \hat{i}^k) - \pi^*) \right] + \varphi_i \hat{i}^k + \nu_i \quad (95)$$

at each value of ν , ζ , and $\hat{i}$ in the grid. Using our interpolated functions, we can then compute state-dependent expectations of the other endogenous variables, or of functions of them, at any value of the current exogenous and endogenous states. For example, in iteration n , our approximation to the Phillips Curve, given by equation (24), when the lagged shadow rate takes a value of $\hat{i}^k$, is

$$\hat{f}_\pi^n(\nu^i, \zeta^j, \hat{i}^k) = \pi^* + \beta \exp[g(\rho - 1)] (\mathbf{q}_{i,j}' \hat{\mathbf{F}}_\pi^{n-1}(\iota_{i,j,k}^n)) - \pi^* + \kappa \hat{f}_y^{n-1}(\nu^i, \zeta^j, \hat{i}^k) \quad (96)$$

where the dot product $\mathbf{q}_{i,j}' \hat{\mathbf{F}}_\pi^{n-1}(\iota_{i,j,k}^n)$ computes the one-period-ahead expectation of π , conditional on being at node (i, j, k) in the current period. We apply a similar computation to the other forward-looking equations of our model, specifically (22) and (23), to solve for $\hat{\mathbf{F}}_y^n(\hat{i}^k)$ and $\hat{\mathbf{F}}_V^n(\hat{i}^k)$ in each iteration. (Computing equity returns requires a similar iteration for price-dividend ratios, which we omit here as it is not necessary for the model solution.)

We iterate n to convergence.