Monetary-Policy Risk and Equilibrium Asset Prices

Corey M. Feldman Thomas B. King*

January 18, 2024

Abstract

We study how the variance of the exogenous component of monetary policy contributes to risk premia and returns across asset classes. Monetary-policy risk (MPR) helps explain why the equity premium and real and nominal term premia are all positive on average. A secular reduction in MPR, such as likely resulted from the changes in Federal Reserve operations and communications over the last three decades, is directionally consistent with the observed long-term declines in interest-rate volatility, the nominal term premium, and the correlation between bond and stock returns. These trends are compounded by the policy rate being close to its effective lower bound and by greater cyclical variation in MPR itself. Transitory shocks to MPR generate countercyclical risk premia on stocks and bonds and contribute to negative variance-risk premia. We quantify these results in a nonlinear New Keynesian model that matches both the long-run empirical features of asset prices and the responses of asset prices to MPR shocks in event studies.

PRELIMINARY DRAFT

*Federal Reserve Bank of Chicago. Contacts: Corey Feldman (corey.feldman@chi.frb.org) and Thomas King (thomas.king@chi.frb.org). The views expressed here do not reflect official positions of the Federal Reserve.

1 Introduction

At any point in time, the future course of the short-term interest rate set by monetary policymakers is uncertain. This uncertainty can be thought of as consisting of two parts—an endogenous component that reflects the passthrough of macroeconomic uncertainty via the policy rule, and an exogenous component that reflects policymaker discretion. The theoretical and empirical effects of exogenous, discretionary policy shocks have been widely studied in the macroeconomic literature (see Christiano, Eichenbaum, and Evans (1996), Rotemberg and Woodford (1996), and countless subsequent papers). Yet research on the effects of the risk attending these shocks is sparse. This is somewhat surprising since, if monetary-policy shocks are important for the economy, it stands to reason that changes in the *variance* of those shocks could also matter. Asset prices, which are especially sensitive to fluctuations in risk, may be where such effects are most manifest. With that motivation in mind, we consider the effects of such "monetary-policy risk" on nominal bonds, real bonds, and equities through the lens of standard macro-finance theory.

The qualitative intuition behind our main results is straightforward and not model-specific. In the presence of nominal rigidities, monetary-policy shocks move the payoffs on stocks, nominal bonds, and inflation-indexed bonds in the same direction as real activity. Higher monetary-policy risk (MPR) therefore increases the likelihood that these assets will experience bad returns in bad states of the world, and risk-averse investors require compensation for that risk. Thus, MPR contributes to risk premia on these assets, shedding light on long-standing questions about why their excess returns are all positive on average (Mehra and Prescott (1985); Backus, Gregory, and Zin (1989); Hsu, Li, and Palomino (2021)). On the other hand, policy shocks cause inflation to increase in *good* states of the world, and therefore, unlike most other sources of interest-rate risk, MPR commands a negative inflation-risk premium.

Long-run changes in MPR are also consistent with trends in the asset-price data over time. For example, Figure 1 uses daily data to construct quarter-by-quarter measures of some key asset-price moments. It shows that between the early 1980s and the early 2010s, the realized volatilities of both short- and long-term nominal interest rates trended markedly downward. (In a shorter sample, Bauer, Lakdawala, and Mueller (2022) show that a similar trend applies to risk-neutral rate volatilities implied by options.) Meanwhile, nominal term premia, according to two popular models, declined at a steady pace of about 80 basis points per decade. And, as emphasized by Campbell, Pflueger, and Viceira (2020), the correlation between returns on equities and nominal 10-year Treasuries dropped from about +40% to about -40%. We argue that all three phenomena are consistent with a secular decline in monetary-policy risk.

Unlike other potential explanations for these trends, like a possible decline in supply-side volatility, a downward drift in MPR has a clear concomitant in the historical record—the steady evolution of a more predictable and transparent monetary policy. The date at which we begin Figure 1 corresponds to the introduction of the new operating regime introduced by Paul Volcker, which, as a stark break from the passive regime of the 1970s, injected a great deal of uncertainty and volatility into the monetary-policy process. Over time, the FOMC adhered more closely to a consistent rule, and it gradually revealed this rule to the market through actions and communications. Beginning in the early 90s it undertook a series of initiatives specifically intended to enhance its transparency and credibility, including releasing meeting minutes, issuing increasingly detailed FOMC statements, adopting a formal inflation target, engaging in forward guidance, introducing the Summary of Economic Projections, conducting press conferences, and promulgating the Statement of Longer-run Goals and Policy Strategy. All of these measures likely reduced the size of policy shocks and investors' uncertainty about policymaker behavior.[1]

That policy has become more consistent and transparent on average does not imply that MPR is always low, however. Recent work by Husted, Rogers, and Sun (2020), Bauer et al. (2022), and Bundick, Herriford, and Smith (2022) has shown that uncertainty about interest rates and monetary policy varies over the business cycle and even over the FOMC cycle. 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 whether monetary policymakers might at some point deviate from their rule or change the rule itself. Thus, MPR may help to explain variation in stock and bond returns across a range of frequencies.[2] In particular, it provides one potential explanation for why both the equity premium and the nominal term premium appear to be countercyclical in the data.

We formalize and quantify these intuitive claims using a workhorse New Keynesian model, which we expand in a few ways to generate realistic behavior of asset prices. First, as in many other papers in this literature, we introduce recursive preferences. Although our model does not feature long-run risk, the Epstein-Zin pricing kernel nonetheless helps us achieve empirically realistic values for risk premia. Second, we introduce stochastic volatility in the monetary-policy shock. Depending on its persistence, this term can help us to understand MPR fluctuations over short or long horizons. Third, we allow for an effective lower bound (ELB) on nominal interest rates, using a "shadow rate" specification for the policy rule. This

[1] See Swanson (2006) for an early assessment of how these measures were reducing the volatility and uncertainty associated with monetary policy. The trendlines we plot in the figure end in 2012, by which point this full suite of communication tools had been deployed. Indeed, subsequent to about 2012, all three trends slowed or reversed.

[2] Justiniano and Primiceri (2008) find significant low- and high-frequency variation in the variance of monetary-policy shocks in an estimated DSGE model with stochastic volatility.

is important because the effects of structural shocks in New Keynesian models can change dramatically, even switching sign, at the ELB; consequently, the mapping from structural sources of risk to risk premia can also change. All three of these extensions add significant nonlinearity to the model. Because it is important for us to obtain accurate, state-dependent values for risk premia, we solve numerically using a global solution method. The model broadly matches the key unconditional moments of asset prices and macro variables in the data.

On average in our model, MPR explains 30 to 40 basis points of the nominal ten-year term premium, about one-third of its average value, and essentially all 44 bp of the real term premium. Meanwhile, it contributes about 3.2 percentage points to the equity premium on average. Uncertainty about MPR—that is, the shock to the variance of policy shocks—has relatively minor effects, but it contributes a few basis points of negative pressure to the inflation-risk premium and to the variance risk premia on both stocks and bonds on average. Since it acts like a demand shock on the macroeconomy, MPR pushes term premia and expected short-term rates in opposite directions, consistent with the observation that the correlation between these two components of bond yields is typically negative (Cieslak and Povala (2016)).

Near the steady-state, a transitory increase in MPR generates the following outcomes in our model: (1) implied and realized volatilties 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 increase, but these effects are smaller than the decline in expected short-term interest rates in response to the deteriorating economy; thus, (5) the nominal yield curve shifts lower and steepens, and nominal bonds realize positive excess returns; (6) on the other hand, the real yield curve moves higher; (7) the inflation risk premium falls. Most of these effects are quantitatively significant. For example, an increase in MPR that temporarily raises the conditional volatility of the short rate by 50 bp produces a contemporaneous decline in stock returns of 2% and an increase in 10-year bond returns of 0.5%. At the ELB, the impact of MPR on interest-rate volatility is dampened, so shocks to MPR generally have smaller effects on risk premia. Nonetheless, the effects on realized stock and real-bond returns are larger at the ELB because, as with demand shocks, the absence of a monetary-policy response amplifies the drag on output and inflation.[3]

As an external validation of the model, we compare our results to the event-study evidence of Bauer et al. (2022), who estimate the impact of changes in MPR on asset prices on FOMC-announcement days. The model is generally consistent with the signs and magnitudes of the

[3]The higher effect of risk on the macroeconomy at the ELB is consistent with the empirical findings of Caggiano, Castelnuovo, and Pellegrino (2017) and similar to the theoretical results of Nakata (2017) and Basu and Bundick (2017), although those papers do not address MPR specifically.

responses of equity prices, stock-market volatility, and term premia to such shocks. We also extend this analysis by dividing the Bauer et al. sample into observations where policy was constrained by the ELB and observations where it was not. We show that the relationships between short-rate uncertainty and asset prices are larger and more statistically significant at the ELB, consistent with our model's predictions. Indeed our model broadly matches the quantitative responses of most asset prices in both the ELB and non-ELB samples.

As noted above, over the longer run it may be that improvements in the predictability of monetary policy have effectively altered the parameters that govern MPR in our model. We therefore consider how the unconditional moments of asset prices change when these parameters change. We show that a shift in parameters that reduces the average level of MPR leads to lower implied and realized interest-rate volatilities, a lower nominal term premium, and a lower correlation between stock and bond returns—all qualitatively consistent with the patterns documented in Figure 1. Although the effects of MPR are not big enough to fully account for the magnitudes of the observed changes on their own, two further phenomena explain much of the discrepancy. First, proximity to the ELB has been a key feature of the post-2000 U.S. economy. Second, even though we have argued that the average level of MPR has fallen, it is also plausible that the increased frequency and volume of FOMC communication has caused MPR to vary more, resulting in an increase in the volatility of MPR shocks themselves. These mechanisms work in the same direction as a decrease in the average level of MPR with respect to the asset-pricing regularities just mentioned.

Our paper connects to three broad strands of literature. The first is the recent research on interest-rate and policy uncertainty mentioned above. (Husted et al., 2022, Bauer et al., 2022, Bundick et al., 2023). That work is largely empirical and descriptive, and our model helps to make sense of some results that it documents, as well as to connect these facts to broader asset-pricing theory. 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, Sim, and Zakrajsek (2014), Jurado, Ludvigson, and Ng (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 MPR shocks 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 similar macroeconomic effects in theoretical models, they can have much different consequences for asset prices, as we show.

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, Tristani, and Vestin (2008), Bekaert, Cho, and Moreno (2010), Li and Palomino (2014), Rudebusch and Swanson (2012), 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 (2023). Gourio and Ngo show how the proximity of the ELB in a New Keynesian model endogenously lowers term premia and changes the correlation of stock returns and inflation from negative to positive in the presence of productivity shocks (a result that also occurs in our model). 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. Although neither paper studies monetary-policy risk, both match risk premia across assets in a New Keynesian framework and attempt to explain long-term changes in asset-price behavior through changes in the structural environment, as we do.[4]

The paper proceeds as follows. In Section 2, we lay out some basic asset-pricing concepts we will need. In Section 3, we explain our basic intuition by showing how risk premia on stocks and bonds are related to structural sources of risk—including monetary-policy risk—in a two-period linear-Gaussian economy, where simple closed-form solutions are available. In Section 4 we develop our quantitative, nonlinear NK model and explain the solution method and parameterization. Section 5 presents the baseline results, calibrated roughly to asset prices over since the 1970s. Section 6 presents the comparison of our model results to the event-study evidence. Section 7 explores how the results change with the underlying policy-risk process and proximity to the ELB. Section 8 concludes.

2 Asset-Pricing Preliminaries

To establish some notation and definitions, we begin by reviewing some basic no-arbitrage asset-pricing equations that will hold throughout the paper. Let $M_{t+1} = \exp(m_t)$ be the the one-period real stochastic discount factor (SDF) in period t. An absence of equilibrium arbitrage opportunities and non-pecuniary motives for holding financial assets implies that $E_t[M_{t+1}R^x_{t+1}] = 1$, where $R^x_{t+1} = \exp(r^x_t)$ is the real gross return on any asset x. The price of a one-period bond that has a nominal payoff of 1 dollar in $t + 1$ is thus

$$P_t^{\$(1)} = E_t[M_{t+1}/\Pi_{t+1}] \tag{1}$$

where $\Pi_{t+1} = \exp(\pi_{t+1})$ is the gross change in the consumer price level between periods t and $t + 1$ and the \$ superscript indicates that the bond has nominal payoffs.

[4]Pflueger's results indicate a reduction in the volatility of monetary-policy shocks between the pre- and post-2000 periods, consistent with our argument, but this is not a change in parameter values she emphasizes.

The price of an N-period zero-coupon nominal bond can be found recursively as

$$P_t^{\$(N)} = E_t \left[\frac{M_{t+1}}{\Pi_{t+1}} P_{t+1}^{\$(N-1)} \right] \tag{2}$$
$$= E_t \left[\exp \left(\sum_{n=1}^{N} m_{t+n} - \pi_{t+n} \right) \right] \tag{3}$$

The yield on a nominal bond is defined as

$$y_t^{\$(N)} \equiv -\frac{1}{N} \log P_t^{\$(N)} \tag{4}$$

Similarly, the price of a real (inflation-indexed) bond can be written as

$$P_t^{(N)} = E_t \left[M_{t+1} P_{t+1}^{(N-1)} \right] \tag{5}$$
$$= E_t \left[\exp \left(\sum_{n=1}^{N} m_{t+n} \right) \right] \tag{6}$$

and the real yield is given by $y_t^{(N)} \equiv -\frac{1}{N} \log P_t^{(N)}$. "Inflation compensation" is the difference between nominal yields and real yields. One-period yields on real and nominal bonds are "short-term interest rates," which we denote by $r_t = y_t^{(1)}$ and $i_t = y_t^{\$(1)}$ throughout the paper.

In our model of Section 4, following Gourio and Ngo (2020), Pflueger (2023), and others, we allow for the possibility that investors receive additional value from holding government bonds—say, for safety or liquidity reasons—and that this benefit may vary over time. In this case, equations (3) and (6) generalize to

$$P_t^{\$(N)} = \exp(\xi_t) E_t \left[\frac{M_{t+1}}{\Pi_{t+1}} P_{t+1}^{\$(N-1)} \right] \tag{7}$$
$$P_t^{(N)} = \exp(\xi_t) E_t \left[M_{t+1} P_{t+1}^{(N-1)} \right] \tag{8}$$

where ξ_t is a random variable governing the demand for bonds. All else equal, such shocks result in equal parallel shifts of the real and nominal yield curves, though in general equilibrium the effects may be more complicated.

Term premia are effectively the average expected excess returns on bonds. Specifically, we define the real term premium as 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. Thus, the real term premium (RTP) is

$$RTP_t^{(N)} = y_t^{(N)} - \mathscr{E}_t^{(N)}[r_t] \tag{9}$$

where we define the operator $\mathscr{E}_t^{(N)}$ as the time-t log expectation of the N-period geometric mean:

$$\mathscr{E}_t^{(N)}[r_s] \equiv \log E_t \left[\exp \left(\frac{\sum_{n=1}^{N} r_{s+n-1}}{N} \right) \right] \tag{10}$$

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:

$$IRP_t^{(N)} = y_t^{\$(N)} - y_t^{(N)} - \mathscr{E}_t^{(N)}[\pi_t] \tag{11}$$

We then define the nominal term premium (NTP) as

$$\begin{aligned} NTP_t^{(N)} &= RTP_t^{(N)} + IRP_t^{(N)} && (12) \\ &= y_t^{\$(N)} - \left(\mathscr{E}_t^{(N)}[r_t] + \mathscr{E}_t^{(N)}[\pi_t] \right) && (13) \end{aligned}$$

We note two ways in which these definitions are not entirely standard. First, we have defined term premia relative to geometric expectations of interest rates, rather than the arithmetic expectations that are sometimes used in the term-structure literature. This subsumes Jensen's inequality terms, which are not of primary interest, into the expectations component of yields and leaves term premia as clean measures of deviations from risk neutrality. Second, 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 $\mathscr{E}_t^N[i_t] \neq \mathscr{E}_t^N[r_t] + \mathscr{E}_t^N[\pi_t]$). This asymmetry is necessary to ensure that the IRP and RTP add up to the NTP. One implication is that short-term nominal bonds carry a term premium, consisting entirely of $IRP_t^{(1)}$. In other words, the Fisher equation does not generally hold because even one-period nominal bonds bear some amount of inflation risk, which requires compensation. Neither of these definitional choices is important for our results.

Meanwhile, the price on an equity contract P_t^{eq} is the stream of future nominal dividends $\{D_t\}$, adjusted for inflation:

$$\begin{aligned} P_t^{eq} &= \mathrm{E}_t \left[\sum_{n=1}^{\infty} \frac{M_{t+n}}{\Pi_{t+n}} D_{t+n} \right] && (14) \\ &= \mathrm{E}_t \left[\frac{M_{t+1}}{\Pi_{t+1}} (P_{t+1}^{eq} + D_{t+1}) \right] && (15) \end{aligned}$$

The nominal return on any asset x is its real return times inflation $R_{t+1}^x \Pi_{t+1}$. One-period log nominal returns on nominal bonds, real bonds, and equities are as follows:

$$r_{t+1}^{\$(N)} + \pi_{t+1} = \log P_{t+1}^{\$(N-1)} - \log P_t^{\$(N)} \tag{16}$$

$$r_{t+1}^{(N)} + \pi_{t+1} = \log P_{t+1}^{(N-1)} - \log P_t^{(N)} + \pi_{t+1} \tag{17}$$

$$r_{t+1}^{eq} + \pi_{t+1} = \log(P_{t+1}^{eq} + D_{t+1}) - \log P_t^{eq} \tag{18}$$

The equity-risk premium (ERP) is the expectation of (18), less the value that this expectation would take under risk neutrality:

$$ERP_t = \log \mathrm{E}_t \left[\frac{P_{t+1}^{eq} + D_{t+1}}{\Pi_{t+1}} \right] - \log[P_t^{eq}] - r_t \tag{19}$$

Because of convexity terms, this is slightly different from the expected excess returns on equities, which are given by $\mathrm{E}_t \left[r_{t+1}^{eq} + \pi_{t+1} \right] - i_t$.[5]

One-period implied volatilities on stocks and nominal bonds are given by the risk-neutral standard deviations of log returns:

$$IV_t^{eq} = \sqrt{ \frac{\mathrm{E}_t \left[\frac{M_{t+1}}{\Pi_{t+1}} (r_{t+1}^{eq} + \pi_{t+1})^2 \right]}{P_t^{\$(1)}} - \left(\frac{\mathrm{E}_t \left[\frac{M_{t+1}}{\Pi_{t+1}} (r_{t+1}^{eq} + \pi_{t+1}) \right]}{P_t^{\$(1)}} \right)^2 } \tag{20}$$

and similarly for bonds. Similarly, we can price claims on asset volatility, such as may be constructed from options strategies or variance swaps. In particular, a contract that pays the level of realized equity volatility as if period $t+1$ has a time-t price given by

$$VSWAP_t^{eq} = \mathrm{E}_t \left[\frac{M_{t+1}}{\Pi_{t+1}} \mathrm{var}_{t+1}[r_{t+2}^{eq} + \pi_{t+2}] \right] \tag{21}$$

The "variance risk premium" (VRP) on an asset is defined as the difference between the price of the variance swap and the physical variance of equity returns, conditional on time-t information:

$$VRP_t^{eq} = VSWAP_t^{eq} - \mathrm{var}_t[r_{t+2}^{eq} + \pi_{t+2}] \tag{22}$$

The vareiance-swap price and variance risk premium are calculated analogously for nominal bonds $VRP_t^{\$(N)}$. We could also compute these quantities for real bonds, but there are no risk-neutral volatility series available for comparison in the data.

[5]Note 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," $\mathrm{E}_t[r_{t+1}^{eq}] - r_t$. These are not generally the same because the Fisher equation does not hold. Unless otherwise noted, we work with the nominal returns.

3 Analytical Intuition

The basic qualitative relationships we wish to explore can be illustrated easily the case of assets that live for only two periods when m_t, π_t, and $\log D_t$ are jointly normal. Most of these results are standard in the asset-pricing literature, but we highlight the aspects of them that are relevant for thinking about monetary-policy risk. The New Keynesian model we develop in Section 4 is not log-normal and does not have closed-form solutions, but qualitatively the same relationships hold. In this section, we develop intuition by first pointing out heuristically how MPR should feed into the covariances that determine risk premia and then by making these arguments precise in the context of a reduced-form linear model where we can obtain exact solutions.

We defer the discussion of variance risk premia because one can show that those premia are always zero in the case of log-normality. The structural model we develop in Section 4 will feature both nonlinearity and heteroskedasticity and will produce nonzero variance risk premia on both stocks and bonds.

3.1 Risk premia in the log-normal, two-period case

Under log-normality, one-period real and nominal bond yields (i.e., short-term interest rates) are given by

$$r_t = -\mathscr{E}_t^{(1)}[m_{t+1}] \tag{23}$$

and

$$\begin{aligned} i_t &= -\mathscr{E}_t^{(1)}[m_{t+1} - \pi_{t+1}] && (24) \\ &= r_t + \mathscr{E}_t^{(1)}[\pi_{t+1}] + \underbrace{\mathrm{cov}_t[m_{t+1}, \pi_{t+1}]}_{IRP_t^{(1)}} && (25) \end{aligned}$$

One-period real bonds carry no risk, and so the term premium on such bonds $RTP_t^{(1)}$ is always zero. The term $\mathrm{cov}_t[m_{t+1}, \pi_{t+1}]$ in (25) is the inflation-risk premium on the one-period nominal bond. 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.

The real two-period bond yield in the log-normal case is

$$\begin{aligned} y_t^{(2)} &= -\frac{1}{2}\log \mathrm{E}_t\left[\exp\left(m_{t+1} + m_{t+2}\right)\right] && (26) \\ &= \mathscr{E}_t^{(2)}\left[r_t\right] + \underbrace{\mathrm{cov}_t[m_{t+1}, r_{t+1}]}_{RTP_t^{(2)}} && (27) \end{aligned}$$

The first term is the expectation of short-term real rates. The covariance term constitutes the RTP. It 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 will generally increase the real term premium. In most macroeconomic models, including those we will describe below, monetary policy shocks have this effect because they cause real interest rates to rise in bad states of the world. Thus, in general, the real term premium should *rise* when monetary-policy risk increases.

Similarly, the two-period nominal yield can be written as

$$\begin{aligned} y_t^{\$(2)} &= \mathscr{E}_t^{(2)}[i_t] + \mathrm{cov}_t[m_{t+1} - \pi_{t+1}, i_{t+1}] && (28) \\ &= \mathscr{E}_t^{(2)}[r_t] + \mathscr{E}_t^{(2)}[\pi_t] + RTP_t^{(2)} \\ &\quad + \underbrace{IRP_t^{(1)} + \mathrm{cov}_t[\pi_{t+1}, m_{t+2}] + \mathrm{cov}_t[\pi_{t+2}, m_{t+1}] + \mathrm{cov}_t[\pi_{t+2}, m_{t+2}]}_{IRP_t^{(2)}} && (29) \end{aligned}$$

Again, the first term in (29) is the expectations component of interest rates (the sum of expected real rates and expected inflation). 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. 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. In economies with nominal rigidities, monetary-policy shocks move inflation and output in the same direction. When such shocks are more prevelant—i.e., MPR increases—the inflation risk of nominal bonds falls. This is true for both one- and two-period bonds, as long as tighter policy reduces both real activity and inflation in a persistent way. Thus, in contrast to the RTP, the IRP should *fall* when monetary-policy risk increases. Since the payoffs on two-period bonds are just the inverse of the corresponding interest rates in $t+1$, log nominal returns on nominal and real bonds are given by

$$\begin{aligned} r_{t+1}^{\$(2)} + \pi_{t+1} &= 2y_t^{\$(2)} - i_{t+1} && (30) \\ r_{t+1}^{(2)} + \pi_{t+1} &= 2y_t^{(2)} - r_{t+1} + \pi_{t+1} && (31) \end{aligned}$$

Finally consider the equity claim which pays a nominal dividend D_{t+1}. The stock price is

$$P_t^{eq} = \mathrm{E}_t \left[\frac{M_{t+1}}{\Pi_{t+1}} D_{t+1} \right] \qquad (32)$$

The log nominal return on this asset is given by

$$r_{t+1}^{eq} + \pi_{t+1} = \log D_{t+1} - \log P_t^{eq} \tag{33}$$

$$= \log D_{t+1} - \mathscr{E}_t^{(1)} \left[m_{t+1} - \pi_{t+1} + \log D_{t+1}\right] \tag{34}$$

and the expected return can be decomposed as:

$$\mathrm{E}_t \left[r_{t+1}^{eq} + \pi_{t+1}\right] = i_t \underbrace{-\mathrm{cov}_t[m_{t+1}, d_{t+1}]}_{ERP_t} + \underbrace{\mathrm{cov}_t[\pi_{t+1}, d_{t+1}] - \frac{1}{2}\mathrm{var}_t[d_{t+1}]}_{\text{Jensen}} \tag{35}$$

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. Monetary-policy shocks cause nominal dividends to fall in bad states of the world, so an increase in MPR pushes the equity premium higher. The last two terms 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.

The variance-covariance structure of asset returns also have simple expressions in this case. For example, the conditional one-period correlation between returns on equities and two-period nominal bonds is

$$\mathrm{corr}_t \left[r_{t+1}^{eq} + \pi_{t+1}, r_{t+1}^{\$(2)} + \pi_{t+1}\right] = \frac{-\mathrm{cov}_t \left[\log D_{t+1}, i_{t+1}\right]}{\sqrt{\mathrm{var}_t[\log D_{t+1}]\mathrm{var}_t[i_{t+1}]}} \tag{36}$$

Because monetary policy shocks move nominal interest rates in the opposite direction of nominal economic activity, an increase in the variance of such shocks (MPR) can be expected to increase this correlation.

3.2 Asset prices and risk in a reduced-form economy

To illustrate the effects of monetary-policy risk more explicitly and compare them to the effects of other structural sources of risk, we now work out the above asset-pricing equations in the case of a simple, linear macroeconomic model. We assume that equilibrium log dividends, inflation, and interest rates are linear functions of three exogenous shocks:

$$d_t = \epsilon_t^D + a_d^S \epsilon_t^S + a_d^M \epsilon_t^M \tag{37}$$

$$\pi_t = a_\pi^D \epsilon_t^D - \epsilon_t^S + a_\pi^M \epsilon_t^M \tag{38}$$

$$r_t = a_r^D \epsilon_t^D + a_r^S \epsilon_t^S + \epsilon_t^M \tag{39}$$

where ϵ_t^D ,ϵ_t^S , and ϵ_t^M are mean zero, independent, serially uncorrelated, and have variances σ_D^2, σ_S^2, and σ_M^2, respectively. We assume that the reduced-form coefficient values in (37) through (39) are such that these three shocks have the typical properties of aggregate-demand, aggregate-supply, and monetary-policy shocks in models with monetary non-neutrality. Specifically, we assume that the coefficients a_π^D, a_r^D, and a_d^S are positive and a_r^S, a_d^M, and a_π^M are negative. Since expected inflation is always zero in this model, from equation (25), the nominal short rate is

$$i_t = r_t + IRP^{(1)} + V_\pi \tag{40}$$

where $V_\pi = \mathrm{var}[\pi_t]/2$. Homoskedasticity and log-normality ensure that both V_π and $IRP^{(1)}$ are constant. Thus, real and nominal short-term interest rates respond identically to shocks. Table 1 summarizes these assumptions by showing the contemporaneous directional effect of a positive shock of each type on each of the endogenous variables.

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

$$m_t = -r_{t-1} + \mu^D \epsilon_t^d + \mu^S \epsilon_t^s + \mu^M \epsilon_t^M + V_m \tag{41}$$

for some parameters μ, and where $V_m = \mathrm{var}[m_t]/2$, a constant. (Note that this specification ensures consistency with equation (23).) We assume that positive monetary-policy shocks and negative supply shocks lower real activity, consistent with their directional affect on dividends, so that they raise the marginal utility of cash flows—i.e., $\mu^S < 0$ and $\mu^M > 0$. The effects of demand shocks on m_t are theoretically ambiguous but are not especially important for us. We assume that they are also positive, consistent with the structural model we present later.

Given these assumptions, it is straightforward to deduce the effects of an increase in the volatility of any of the three shocks on risk premia and on the first and second moments of asset returns. For example, consider the equity risk premium which is governed by the covariance between the SDF and dividends, as shown in equation (35):

$$\begin{aligned} ERP &= -\mathrm{cov}_t[m_{t+1}, d_{t+1}] && (42)\\ &= -\mu_d \sigma_D^2 - a_d^S \mu_S \sigma_s^2 - a_d^M \mu^M \sigma_M^2 && (43) \end{aligned}$$

One can read off of this equation the effects of changes in the variance of any of the shocks on the ERP. In particular, given our interest in monetary-policy risk, we have that $\partial ERP/\partial \sigma_M^2 = -a_d^M \mu^M > 0$. Thus, an increase in MPR raises the risk premium on stocks. The reason is straightforward. Positive monetary-policy shocks cause dividends to fall and, because they also reduce aggregate income and consumption, cause marginal utility to rise—stocks lose value in bad monetary-policy states. An increase in the variance of such shocks thus makes the covariance between dividends and marginal utility more negative, for which investors require compensation in the form of higher expected returns.

Working in this way for the other variables of interest, the following proposition summarizes the effects of MPR in this model:

Proposition 1. In the log-linear Gaussian model described above, an increase in monetary-policy risk σ_M 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;
- higher pairwise covariances between (a) returns on equities and nominal bonds and (b) returns on equities and inflation;
- for σ_M sufficiently large, higher pairwise correlations between (a) returns on equities and nominal bonds and (b) returns on equities and inflation.

The last item requires some explanation. Although the covariances noted in the proposition are both strictly increasing in σ_M, the corresponding correlation coefficients may be decreasing over some range of σ_M, particularly when it is small relative to σ_S.[6] Nonetheless, it is straightforward to verify that, in the limit, the correlations necessarily go to +1 as MPR gets large.

Tables 2 and 3 summarize these claims. 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 reported in the table 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 risk, like monetary 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 risk lead to higher real and nominal term premia, while demand risk reduces them. On the other hand, supply and monetary-policy

[6] To see this, set σ_D to zero. Then, the correlation between stock and bond returns is $\frac{-\left(a_d^S a_i^S \sigma_S^2 + a_d^M \sigma_M^2\right)}{\sqrt{\left((a_d^S)^2\sigma_S^2 + (a_d^M)^2\sigma_M^2\right)\left((a_i^S)^2\sigma_S^2 + \sigma_M^2\right)}}$. The derivative of this expression with respect to σ_M is positive iff $a_d^M \sigma_M^2 > a_i^S a_d^S \sigma_S^2$.

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 so far assume that shocks to MPR have no effects on the *levels* of dividends, inflation, or interest rates. However, in general, risk shocks of all types reduce economic activity. 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 (e.g., Leduc and Liu (2016)).

With this motivation, suppose that the increase in σ_M that we have examined above is accompanied by a negative realization of the demand shock ϵ_t^D. In this case, we can draw the following additional conclusions about asset prices:

Proposition 2. If, in the log-linear Gaussian model described above, an increase in monetary-policy risk σ^M induces a negative aggregate-demand shock, it has the following effects, in addition to those described in Proposition 1:

- A steeper nominal yield curve,
- Negative abnormal returns on equities,
- Positive abnormal returns on long-term nominal bonds

Again, these results are straightforward to show, but it is instructive to consider the case of the steepening yield curve. Suppose the increase in MPR causes a demand shock of magnitude -1. Then, the nominal short rate changes by $-a_r^D < 0$. Since the shock is assumed to be iid, the expectations component of the two period yield, $\mathscr{E}_t^{(2)}[i_t]$, shifts down by an amount $-a_r^D/2$. (If the shocks were persistent, the expectations component would shift down by a larger amount, but still by no more than the change in the short rate.) We have shown above that the nominal term premium rises in response to an increase in MPR, and demand shocks have no effect on term premia. Thus, long-term yields must shift down by less than short-term yields. The combination of a steeper yield curve and higher term premium is consistent with empirical evidence that a higher yield-curve slope predicts higher bond returns (e.g., Fama and Bliss (1987)).[7]

[7] As we have specified the coefficient a_r^D, the same conclusions hold for the real yield curve, but we do not emphasize this result because it is sensitive to assumptions about how quickly policy rates adjust to demand shocks. If nominal short rates are very inertial, real rates may rise in the short term in response to a change in MPR. Indeed, in the structural model below this turns out to be the case.

We also note that, while covariances, and therefore risk premia, can be affected by changes in the variance of shocks, they can also shift because of changes in the reduced form coefficients on the shocks. These changes could be either exogenous, due to secular shifts in the underlying structural environment, or endogenous, if the coefficients on shocks change in different states of the world (i.e. variables are nonlinear functions of the shocks). A particularly important endogenous change in reduced-form dynamics occurs at the ELB. The net effects of such changes on risk premia and asset-price dynamics depend on a variety of quantitative factors that require a structural model to parse.

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 asset prices. 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 suggestd 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

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:

$$\hat{i}_t = (1-\varphi_i)\left[\pi^* + r^* + \phi_y \widetilde{y}_t + \phi_\pi(\pi_t - \pi^*)\right] + \varphi_i \hat{i}_{t-1} + \epsilon_t^M \quad (44)$$

$$i_t = \max[\hat{i}_t, 0] \quad (45)$$

where π^* is the long-run inflation target, r^* is the steady-state real interest rate, $\widetilde{y}_t$ is the log deviation of output from its steady-state value, and ϵ_t^M is the monetary-policy shock.

We assume that $\epsilon_t^M \sim N(0, \sigma_{M,t})$. To introduce monetary-policy risk, we allow $\sigma_{M,t}$ to follow an AR(1) process, with truncation at zero to maintain positive variance:

$$\sigma_{M,t} = (1 - \varphi_\sigma)\sigma_M^* + \varphi_\sigma \sigma_{M,t-1} + \epsilon_t^\sigma \tag{46}$$

where ϵ_t^σ follows a distribution that is $N(0, \sigma_\sigma)$, subject to truncation below at $-\mathrm{E}_{t-1}[\sigma_{M,t}]$.[8] Because, under our calibration, $\sigma_{M,t-1}$ is typically far from zero, this process is nearly linear and homoskedastic over most of the relevant range, including near the steady state. 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 $\mathrm{var}_t\left[\sigma_{t+1}^M\right]$ is nearly constant. Under a log process, this variance increases with ϵ_t^σ. Since, as we argued above, MPR shocks act similarly to demand shocks, increases in their conditional variance have qualitatively different effects on asset prices than increases in the conditional variance of ϵ_t^M, and nonlinear processes for $\sigma_{M,t}$ confound these two effects. The linear process, in contrast, allows us to isolate the effects of $\mathrm{var}_t[\epsilon_{t+1}^M]$ that we are interested in.[9]

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.[10] Our approach is to assume that households' lifetime utility is given by preferences of the following form:

$$U_t = \left[(1-\beta)(C_t^\rho - \theta_t H_t^{\chi\rho}) + \beta\left(\mathscr{R}_t\left(V_{t+1}^C\right)^\rho + \mathscr{R}_t\left(V_{t+1}^H\right)^\rho\right)\right]^{1/\rho} \tag{47}$$

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

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

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

$$V_t^C = \left[(1-\beta)C_t^\rho + \beta\mathscr{R}_t\left(V_{t+1}^C\right)^\rho\right]^{1/\rho}$$

[8] Specifically, the PDF of ϵ_t^σ is $\frac{1}{\sqrt{2\pi}(\sigma_\sigma)^{3/2}} \exp\left[-\frac{\epsilon_t^\sigma}{2(\sigma_\sigma)^2}\right]\left(1 - \mathrm{erf}\left[\frac{(1-\varphi_\sigma)\sigma_i^* + \varphi_\sigma\sigma_{M,t-1}}{\sqrt{2}}\right]\right)$.

[9] The 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.

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

$$V_t^H = \left[(1-\beta)\theta_t H_t^{\chi\rho} + \beta \mathscr{R}_t \left(V_{t+1}^H\right)^\rho\right]^{1/\rho}$$

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.

Households hold one-period nominal bonds that have face value of one dollar in quantity B_t. We assume that they receive a time-varying nonpecuniary payoff from holding all types of government bonds (or, equivalently, that they purchase bonds at a time-varying premium relative to the arbitrage-free value). As noted in Section 2.1, this wedge, $\exp(\xi_t)$, captures time-varying preferences for safe and liquid assets. Since interest rates are pegged by the policy rule in this model, ξ_t functions similarly to an aggregate demand shifter with respect to consumption/savings behavior. We assume that it follows

$$\xi_t = \varphi_\xi \xi_{t-1} + \epsilon_t^D \tag{48}$$

where the "demand shock" ϵ_t^D is distributed normally with mean zero and variance σ_D^2.[11]

Letting P_t be the price of the consumption good and W_t be the nominal wage, the household maximizes lifetime utility subject to the flow budget constraint:

$$P_t C_t + P_t^{\$(1)} B_t \leq B_{t-1} + W_t H_t + \mathscr{P}_t \tag{49}$$

where $\mathscr{P}_t$ 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 rate a constant rate $-\rho g$, as in 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

$$y(i)_t = a_t + \psi \log H(i)_t$$

[11] We also solved a version of the model in which the demand factor was instead specified as a fluctuation in time preference, along the lines of Albuquerque et al. (2016). However, in the calibration, we found that that the model best matched the data when that type of shock was given a variance of zero. In contrast, as shown below, non-zero values of the bond-demand shock we use can help to match the asset-price data. Such shocks can be microfounded by adding bonds to the current utility kernel (but not to V_t^C) in equation (47).

where the technology process follows

$$a_t = (\phi_a - 1)a_t^* + \phi_a a_{t-1} + \epsilon_t^S$$

with $\epsilon_t^S \sim \mathrm{N}[0, \sigma_S]$, and 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 $\exp y(i)$.

4.2 Equilibrium and solution

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

$$\theta\chi H_t^{\chi\rho-1} C_t^{1-\rho} = \frac{W_t}{P_t} \tag{50}$$

$$\widetilde{C}_t = \exp[g]\left(\beta\exp[i_t + \xi_t]E_t\left[\frac{\widetilde{C}_{t+1}^{\rho-1}}{\Pi_{t+1}}\Big(\frac{\widetilde{V}_{t+1}^C}{\mathscr{R}_t(\widetilde{V}_{t+1}^C)}\Big)^{\alpha-\rho}\right]\right)^{\frac{1}{\rho-1}} \tag{51}$$

$$\widetilde{V}_t^C = \left[(1 - \beta\exp[g\rho])\widetilde{C}_t^\rho + \beta\exp[g\rho]\mathscr{R}_t\left(\widetilde{V}_{t+1}^C\right)^\rho\right]^{1/\rho} \tag{52}$$

where tildes denote percentage deviations from trend. After log-linearizing, the firm's problem, combined with the labor-supply curve (50), results in the New Keynesian Phillips Curve:

$$\pi_t = \pi^* + \beta\exp[g(\rho-1)]E_t[\pi_{t+1} - \pi^*] - \nu\widetilde{a}_t + \kappa\widetilde{y}_t \tag{53}$$

where ν and κ are reduced-form parameters that take non-negative values. Note that, in order to obtain meaningful risk premia, we do not linearize the Euler equation.

We note that although a "productivity" shock is the only shock we have allowed to enter the Phillips Curve, the random variable a_t can be thought of as standing in for a variety of supply-side factors for our purposes. For example, a hypothetical labor-supply shock or markup shock would enter the equation similarly. Although productivity, labor-supply, and markup shocks have different implications for the labor market, they have the same implications for asset prices (conditional on their dynamics). Thus, we refer to a_t as the "supply shock" to capture this broader meaning.

A solution to the model consists of values of C_t, H_t, W_t, P_t, and $P_t^{\$(1)}$ such that equations (50) through (53) are satisfied and the goods and bond markets clear ($C_t = \exp y_t$ and $B_t = 0$) for all t. Previous papers that have studied asset prices in DSGE models have often relied on higher-order perturbation methods to solve them.[12] Although this approach is relatively efficient and can give accurate approximations for macroeconomic dynamics near the steady state, it is not likely 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 policy functions are not sufficiently smooth, polynomial approximations far from this point can be quite poor.[13] These problems may be compounded if the model involves nondifferentiable functions, such as those induced by the ELB. To avoid these issues, we rely on discretization 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 B describes the numerical procedure in detail.

4.3 Asset Prices

The log SDF in this model is given by

$$m_{t+1} = \log \beta + (\rho - 1)(g + \widetilde{c}_{t+1} - \widetilde{c}_t) + (\alpha - \rho)(\log \widetilde{V}^C_{t+1} - \log \mathscr{R}_t(\widetilde{V}^C_{t+1})) \tag{54}$$

Real and nominal bond prices at maturity N are given as in equations (8) and (**??**), where we assume

$$\xi_t = \varphi_\xi \xi_{t-1} + \epsilon^\xi_t \tag{55}$$

with $\epsilon^\xi_t \tilde{N}[0, \sigma_\xi]$.

To price equities, first consider a hypothetical contract that pays a stream of cash flows equal to aggregate nominal output, $\{\exp(y_{t+1}), \exp(y_{t+1} + \pi_{t+1}), ...\}$. The price of such a contract can be found recursively as

$$P^y_t = \mathrm{E}_t[\exp(m_{t+1} - \pi_{t+1})(P^y_{t+1} + \exp(y_{t+1} + \pi_{t+1}))] \tag{56}$$

and its log returns are

$$r^y_t + \pi_t = \log\left[P^y_t + \exp(y_{t+1} + \pi_t)\right] - \log P^y_{t-1} \tag{57}$$

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

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

[13] See Aldrich and Kung (2019).

that the return on equities is given by

$$r_t^{eq} + \pi_t = \log\left[\frac{1}{\delta}\exp(r_t^y + \pi_t) - \frac{1-\delta}{\delta}\exp(i_{t-1})\right] \tag{58}$$

Finally, one-period implied volatilities on equities are given by equation (20). Risk premia and moments of returns can be found from these expressions, given the relationships noted in Section 2.

4.4 Calibration

Table 4 presents the parameter values we use in our baseline model. Most of the values governing the deterministic dynamics are standard. We set the quarterly rate of time preference β to 0.99, and ρ to 1/3, corresponding to an intertemporal elasticity of substitution of 1.5. The parameters governing labor supply and demand, θ and ψ, ultimately enter the equilibrium conditions through κ and ν in the Phillips Curve. We set ψ to .35, and θ is set such that κ (the parameter governing the Phillips Curve slope) is .05. We set the policy responses to output and inflation to standard Taylor Rule values and the inertial term to 0.75. We set π^* to 3.4% at an annual rate, matching the average level of PCE inflation since 1971. We assume that supply shocks have a quarterly persistence of $\varphi_a = 0.95$ and demand shocks of $\varphi_\xi = 0.9$.

Other values we calibrate to roughly match macroeconomic and asset-price moments. In particular, we search over a grid of model parameter values (δ, α, r^*) and state-dynamic parameters $(\sigma_a, \varphi_\sigma, \sigma_M^*, \sigma_\sigma, \sigma_\xi)$ for values that generate unconditional means, variances, and correlations of output, inflation, interest rates, and asset returns that are similar to those observed in the post-1971 data.[14] The model's success in matching these moments will be discussed below.

We set δ to 0.2, which helps to match the volatility of equity returns given the (relatively low) volatility of nominal output. The parameter α, governing risk aversion, we set to -160, which helps us to generate realistic risk premia. While this is a large value by microeconomic standards, several other papers working with recursive preferences in New Keynesian asset-pricing models have adopted values of this magnitude or higher (e.g., Rudebusch and Swanson (2012), Gourio and Ngo (2020)).[15] r^* is equal to 1.65% which helps to match the general level of interest rates.

The key parameters for us are those that determine the laws of motion for the demand and supply shocks and especially those associated with MPR. We set the standard deviation of supply and demand shocks to $\sigma_a = 0.005$ and $\sigma_\xi = 0.0007$, consistent with moderately

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

[15] Swanson (2021) obtains risk premia similar to those in the data in a recursive-preference NK model with α value of 60 by introducing permanent shocks to productivity.

sized productivity or markup shocks and relatively small fluctuations in bond demand. For MPR shocks we set the persistence φ_σ to .9, the unconditional mean σ_M^* to .0045, and the conditional standard deviation σ_σ to .0014. Because of the endogenous feedback, it is difficult to interpret these magnitudes directly. But, as we show next, they generate realistic values for interest-rate and macroeconomic volatility, as well as helping to replicate observed risk premia.

5 Baseline Results

5.1 Unconditional moments

Table 5 shows how the model matches the moments of macroeconomic variables and asset prices in the data. Given the simplicity of the model and the coarseness of our calibration, we should not expect a perfect match. Even so, the model is able to come reasonably close to generating the observed asset-price and macroeconomic values in most cases, including in the means of short- and long-term interest rates, the volatilities of interest rates, inflation, and output, and most of the pairwise correlations among the macro series and asset returns.

One place the model misses notably is in matching the volatility of long-term yields. This is likely because of insufficient persistence in the shocks. In particular, it is clear in the data that there have been low-frequency changes in interest rates over the last 50 years that have substantially contributed to the unconditional volatility of yields. Our model does not allow for what are effectively stochastic trends in the data. A second reason for the low yield volatility in the model is likely the absence of independent drivers of term premia that are sometimes present in reality (e.g., quantitive easing or the passthrough of short-term interest rates to term premia as in Hanson and Stein (2015)). The low volatility of yields in the model will have implications for the interpretation of several of our other results.

Although it performs better in matching historical risk premia than most other NK models, our model still falls somewhat short on the average values of both the nominal term premium (76 bp, rather than 162 bp) and the equity premium (3.5%, rather than 7.8%).[16] Most of the miss on the nominal term premium can be traced to the low model-implied volatility of long-term yields just noted. If we effectively control for this volatility by computing the bond's Sharpe ratio, the fit appears much better (0.30 versus 0.26). For similar reasons, we also do a better job of matching the term premium on shorter-term yields (not shown). Presumably, the model could do a better job on both bond and equity risk premia if we allowed for even higher levels of risk aversion.

[16]Because inflation-indexed (TIPS) yields are not available for most of our sample, we do not attempt to match them or report them in the table. However, it appears that the model does reasonably well in this dimension. The average TIPS yield in the available data exceeds the average ex-post (CPI-adjusted) real short rate by 37 bp; in the model the comparable value is 44 bp.

Although it is not of primary interest for us, the model also does not do a good job of matching the unconditional correlation between output and interest rates, implying a very negative value. This is a consequence of the relatively small and transitory values we calibrate for demand shocks, since those shocks push these two variables in the same direction. MPR shocks, as will be seen in a moment, also have this effect, but they are also relatively small.

Table 6 shows the unconditional and conditional correlations between the expectations and term-premium components of nominal yields in our model. Positive MPR shocks increase term premia. At the same time, they depress the economy, prompting a persistent policy-easing response. Thus they move the two components of bond yields in opposite directions, and their presence induces a negative correlation. This is broadly consistent with Cieslak and Povala (2016), who esitmate that the correlation between nominal term premia and rate expectations is indeed moderately negative most of the time in the U.S. data. Because expected short rates respond less aggressively to MPR shocks when the ELB is binding, this correlation becomes somewhat less negative at and near the ELB. This is also directionally consistent with Cieslak and Povala (2016), who find that the correlation rises during periods when policy rates are low, though they estimate a much more dramatic response in this direction than our model produces.

5.2 Sources of risk premia

Table 7 reports unconditional moments for the nominal ten-year term premiun, inflation risk premium, real term premium, equity premium, and the variance risk premia on stock returns and the short-term interest rate. To measure the risk premia contribution from MPR, we compute the levels of risk premia that would prevail in a model with only monetary-policy shocks (i.e., shutting down ϵ_t^S, ϵ_t^D, and ϵ_t^σ), leaving the rest of the model intact.[17] We then measure the contribution from MPR shocks as the difference between the model with strictly policy shocks and a model with policy shocks with stochastic variance. However, this decomposition of these risk premia into their structural sources is not uniquely defined due to the nonlinearity of the model. In Table 8 we illustrate how risk premia change when removing 1 or 2 shocks at at time.

The existence of monetary policy shocks contribute 39 basis points to the average value of the ten-year NTP, 49 basis points of which comes through the RTP and -10 basis points of which comes through the IRP. The average slope of the real yield curve is determined by real term premia. Thus, compensation for risk arsing 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. Additionally, MPR drives the bulk of the equity premium,

[17] σ_M^* is adjusted such that the policy shocks have exactly the same unconditional variance as the shocks in the baseline model.

contributing a total of about 320 basis points.

It may appear surprising that monetary-policy risk explains almost the entirety of the equity premium in our model. This result can be understood as follows. Our demand shocks are calibrated to be fairly small and transitory, and they thus have only short-run effects on the macro aggregates and negligible effects on risk premia. Supply shocks, on the other hand, are not trivial. Indeed, they explain the bulk of the nominal term premium and a significant portion of the variation in inflation (not shown). But they have relatively small effects on output and real equity returns and consequently cannot generate a large equity premium. Swanson (2021) also emphasizes that aggregate-supply shocks make a small equity-premium contribution in NK models unless those shocks are permanent (or risk aversion is extremely high). The absence of permanent shocks in our model is one possible source of missing risk and may thus explain why the risk premia we generate are, as noted above, somewhat too small overall.

It is also worth noting that risk shocks—that is, shocks to the conditional standard deviation of policy shocks—contribute -6 to nominal term premia. A negative contribution occurs because risk shocks behave like demand shocks in the sense that they send bond returns higher in bad states of the world and therefore increase the hedging property of bonds. These shocks also contribute to the existence of positive variance risk premia, though these effects are relatively modest.

5.3 Impulse-responses

Figure 2 shows impulse-response functions for the macroeconomic variables in our model in response to each of the four shocks. Figure **??** shows how asset prices respond, by focusing on 10-year real and nominal yields, along with cumulative returns on equities and real and nominal bonds. Because the ELB introduces large changes in how the endogenous variables respond to shocks, we show the IRFs initialized at two different values. In each panel, the red line shows the responses beginning from the steady state, while the blue line shows the responses beginning from a state vector in which the shadow rate is significantly below the ELB.[18] All of the shocks are scaled such that, when they occur in the steady state, they result in a .5% drop in output on impact. This makes the IRFs comparable across different shocks and between the steady-state and ELB cases.

The responses of the macroeconomic variables to supply, demand, and monetary-policy shocks at the steady state are intuitive and standard for New Keynesian models. The responses of stock and bond prices reflect a mix of changes in expected discounted cash flows and risk premia, discussed in more detail below. Negative supply and demand shocks and

[18]Specifically, in the latter case, we initialize the IRF with values of $i_{t-1} = -10\%$. This results in a short-term rate that would remain at the ELB for four quarters in the absence of shocks.

positive policy shocks all cause real dividends to fall, leading to a decrease in stock returns. The supply and policy shocks also cause real and nominal interest rates to rise, compounding the decline in equity prices and also causing negative returns on bonds. On the other hand, the demand shocks cause rates to fall, leading to positive excess bond returns. At the ELB, negative supply shocks become expansionary since the central bank no longer offsets the rise in inflation. Thus stocks experience large short-term gains in response to such shocks.[19] Meanwhile, the effects of monetary-policy shocks on both stock and bond returns become significantly weaker at the ELB since they do not pass through to current short-term interest rates, while the effects of demand shocks become stronger, as the countervailing policy response is temporarily unavailable.

Of more interest for our purposes are the responses to the MPR shock. Starting from the steady-state, in response to an increase in MPR that reduces output by 0.5% (2.8 standard deviations at the steady state), inflation falls by about 0.4%, and the policy rate eases by about 2%. These macroeconomic reactions are similar to the responses to MPR shocks studied in Mumtaz and Zanetti (2013). Meanwhile, equities experience significant negative abnormal returns in response to a positive MPR shock, reflecting a combination of lower real dividends, lower inflation, and an increase in the equity premium. In the periods after the initial shock, overall stock returns are slightly positive for several quarters, reflecting the increase in the nominal risk-free rate.

The responses of the macroeconomic variables and equity returns to MPR are very similar to their responses to aggregate demand, consistent with previous literature that has emphasized the demand-like nature of uncertainty shocks in general. However, when it comes to bond yields, the two types of shocks behave differently. Like the demand shock, the MPR shock causes a contemporaneous decrease in the ten-year nominal yield, but it causes an *increase* in the real yield. We will explore these changes in more detail below, but roughly speaking the decline in the nominal yield in response to MPR is driven mostly by the lower expected path of policy in response. The average expected real rate, on the other hand, is little changed in response to the shock, and most of the increase in real yields reflects a term-premium effect. In both cases, the yield responses are quantitatively small—on the order of a few basis points. However, because these are long-duration instruments, this translates into abnormal returns on the order of 0.5% (positive in the case of nominal bonds, negative in the case of real bonds).

All of the asset-price responses have the same sign at the steady state and at the ELB, but the magnitudes differ somewhat. Most notably, the impact of the MPR shock on contemporaneous stock returns doubles from about 2% at the steady state to 4% at the ELB, again reflecting the larger multipliers associated with the absence of a policy response.

[19]This result essentially mirrors that of Gourio and Ngo (2020)—stocks and inflation become positively correlated at the ELB in the presence of supply shocks.

MPR shocks also change the conditional second moments across macroeconomic and financial variables. Figure 3 shows the response of conditional standard deviations of output, inflation, short rates, stock returns, and bond returns to a shock of this type. Not surprisingly, all of these volatilities rise when the volatility of the policy shock rises, though at the ELB the effects on volatility are smaller. Figure 4 shows the IRFs for five particular conditional correlations that are of interest. Positive MPR shocks have modest positive effects on the correlation between output and inflation, between stock and nominal bond returns, and between stock returns and inflation. They have essentially no effects on the correlations between output and short-term interest rates (nominal or real) at the steady-state, but at the ELB they decrease these correlations somewhat in the short run.

The responses of conditional second moments map into the responses of risk premia. In particular, Figure 5 plots the responses of the equity premium, the real term premium, the nominal term premium, the inflation-risk premium, and the variance risk premia to MPR. All but the inflation-risk premium increase as a result of the MPR shock. Qualitatively, the logic behind these results is just as described in the Section 3. A higher conditional variance of the policy rate makes the covariance between returns and marginal utility more negative across all asset classes, and it also lowers the covariance between inflation and marginal utility. Since inflation eats into nominal bond returns, this last result causes the inflation-risk premium to fall. Quantitatively, the magnitudes of the term premia responses are modest—on the order of a few basis points—but these translate into increases of 0.4% - 0.6% in expected excess returns given the long duration of the bonds in question. Meanwhile, the equity premium rises by 2% (at the steady state) in response to the shock. These effects are somewhat damped at the ELB.

Looking at the responses of yields in a different way, panel (a) of Figure 7 shows how the entirety of the nominal yield curve and its components respond to a MPR shock on impact at the steady state. As seen above, nominal yields decline in response to the shock (black line). The decomposition shows that most of this response reflects a lower path of expected inflation (red squares), although expected real rates (blue diamonds) and inflation risk premia (yellow triangles) also move lower. Meanwhile, real term premia (green circles) increase across maturities, with the largest rise in the 2- to 4-year sector.[20] The shock increases the slope of the curve as the increase in real term premia effect partially offsets the fall in expected rates across maturities. This is consistent with the intuitive arguments we made in Section 3. As shown in panel (b), the effects of MPR shocks on the long end of the nominal yield curve are similar at the ELB 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

[20]The hump-shaped response of term premia across maturities reflects the tension between the greater impact a given-sized increase in volatility has on longer-duration bonds and the relatively short halflife of the risk shock itself.

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.

The bottom two panels of Figure 5 show the responses of the variance risk premia to MPR shocks. For both stocks and bonds, these responses are negative—implying that implied volatilities rise by more than expected realized volatilities in response to MPR—though they are quantitatively small. Interestingly, the ELB magnifies these responses substantially at horizons of about one year.

We note that, theoretically, it is ambiguous whether the effects of MPR on asset prices are larger or smaller at the ELB. On the one had, risk shocks behave like demand shocks in reducing output and inflation. Since the monetary authority is constrained by the ELB, it can not offset negative demand shocks by lowering rates, thereby amplifying the effect of a demand shock. As such, one might expect the effects of risk shocks to be amplified at ELB. However, we have seen that the effect of monetary-policy shocks (ϵ_t^M) are muted at the ELB. Therefore, a shock to monetary policy shocks' second moment (ϵ_t^σ) may not pass through as much to the conditional variance of interest rates at the ELB. Our quantitative results show that both sides of the story are at play in the model. Figures 2 and 3 show that the responses of output, inflation and stock returns to MPR are amplified at the ELB. On the other hand, the effect on bond returns is weaker as the short rate is stuck at the ELB. Figure 4 shows that the latter story also holds—the response of the conditional second moments is weaker at the ELB, which helps to explain why the equity and term premia effect shown in Figure 5 is also weaker at the ELB.

6 Monetary Policy Risk Shocks: Data vs. Model

Are the effects of MPR shocks found in our impulse responses consistent with the data? Empirical identification of a shock to risk about the exogenous component of the Taylor rule at low frequency is a challenge beyond the scope of the paper. However, Bauer, Lakdawala, and Mueller (2022) (BLM) use daily Eurodollar futures and options to construct a measure of monetary-policy risk at the one-year horizon and study its effects. Our model potentially provides a theoretical explanation for their results, and, in turn, their work provides external validity to our model.

BLM rely on an FOMC meeting-date event-study approach, regressing daily changes of various asset prices on changes in a measure of monetary-policy shocks and their measure of uncertainty about short-term interest rates. Under the assumption that changes in interest-rate uncertainty on FOMC days reflect only news about MPR, as we have defined it, these estimates capture the contemporaneous effects of MPR on asset prices.[21] Using the BLM

[21] In the data, changes in the conditional standard deviation of the short rate can also be the result of changes in the conditional variance of output and inflation, which we have referred to as the endogenous

data, we replicate their empirical findings in Table 9. According to these estimates, an increase in MPR, controlling for standard shocks to the fist moment of the monetary-policy stance, is associated with a fall in stock returns and an increase in the VIX, in long term real and nominal yields, and in term premia.

Additionally, motivated by the disparity between our IRFs at the ELB and away from it, we extend the BLM analysis by interacting their measures of monetary policy and short-rate uncertainty shocks with an indicator variable that takes a value of 1 at the ELB. Table 10 presents the results of the event-study regressions with these two interaction terms. We find that the ELB consistently amplifies the effect of MPR shocks, and in some cases this amplification is quite large. The effect of changes in MPR on stock returns at the ZLB is roughly 4 times larger than the estimated effect away from at the ZLB, while for the VIX, the magnification is nearly 6 times.

We compare these empirical results to analogous shocks generated from our model (again, under the assumption that the only second-moment shocks occurring on FOMC-announcement days are shocks to σ^M_{t+1}). Since the estimated marginal effect of MPR in Table 9 and Table 10 control for changes to the short rate—while, in the model, shocks to σ^M_{t+1} induce a simultaneous decline in the short rate—we neutralize the model-implied response of i_t by introducing a policy shock ϵ^M_t such that the 1-year interest rate remains constant at its initial value.[22] We then compute the ratio of the changes in our variables of interest relative to changes in the conditional standard deviation of the 1-year rates. That is, we look at the marginal effect on asset prices of an increase in risk surrounding the observable 1-year rate due to an increase in σ^M_{t+1}. We compute ratios starting from the steady state and at a state where $\hat{i}_{t-1}$ is relatively low to capture the effect of the ELB. Because the shadow rate is unobservable in the data, it is unclear where we should set it when doing this experiment. As such, we compute the ratio at three different negative values of $\hat{i}_{t-1}$ to study the relative effect of the ELB. (All of the exogenous state variables are set to their steady-state values in these experiments.)

The results are documented in Table 11. We find that the effects of shocks to MPR on stock returns and their implied volatility at the steady-state are strikingly similar to what is estimated in the data. Additionally, we find that the ELB magnifies these effects by roughly 5 and 3 times respectively, for moderately negative values of the shadow rate, again generally consistent with the empirical estimates. The direction of the effects at the steady-state and the magnification of these effects at the ELB are both consistent for the term premia and real 10 Year yields as well.

In contrast, we cannot match the empirical response of longer-term nominal yields to increases in MPR. In the data, yields rise in response to MPR primarily because of increases

component of interest-rate risk. Presumably, because they look at changes on FOMC meeting dates the entire variation in MPR is coming from changes in the exogenous component of risk.

[22]The BLM monetary-policy shocks are measured as unanticipated changes in one-year interest-rate futures.

in the term premium. We *do* generally match the estimates of such changes, as shown in the last four columns of the table. However, our model implies that the expectations component of yields falls in response to MPR shocks, while in the data this component rises. Although we neutralize the short-run effect of MPR on expected policy rates, because our policy shock has weaker persistence than our MPR shock, the expectation of short rates falls at the ten-year horizon. This discrepancy is then amplified at the ELB.

Why is a given increase in interest-rate uncertainty associated with so much larger changes in asset prices at the ELB? Figure 3 showed that the effects of an MPR shock MPR on stock returns and stock-market volatility do increase at the ELB, but only by a factor of 2, not a factor of 4 or so as we find in Table 11. And, as we saw in Figures 3 and 6, the effects of an MPR shock on stock-market volatility and term premia *decline* at the ELB, in contrast to the event-study results. The answer to this apparent contradiction is that the independent variable in Table 11 is interest-rate implied volatility (at the 1-year horizon), not the MPR shock itself. As we have show above, a shock to $\text{var}_t\left[\sigma_{t+1}^M\right]$ passes through less to the conditional standard deviation of the short rate $\text{var}_t\left[i_{t+1}\right]$ at the ELB. Thus, a larger MPR shock is required to generate a given amount of future interest-rate volatility. Thus, to a large degree, the "effects" of rate volatility appear bigger at the ELB because the size of the shock associated with a given change in rate volatility is itself bigger.

7 Varying the Risk Parameters

While our baseline model is set up to consider cyclical variation in monetary-policy risk, there are also reasons to suspect longer-term secular changes. As noted in the Introduction, there are strong reasons to suspect that the public has become more informed about the FOMC's behavior over time, especially following the introduction of the panoply of communications tools beginning in the early 1990s. It is thus likely that there has been a secular decline in monetary-policy risk over the last several decades. At the same time, it is also plausible that the increased frequency and volume of information has caused MPR to vary more at higher frequencies, even as it has fallen on average.[23] Our model allows us to study how secular changes in both the average level and the volatility of MPR would map into asset prices.

The two key parameters that jointly govern these unconditional moments in the model are σ_M^* and σ_σ. We solve the model across a grid of these two parameters and calculate unconditional moments. Figure 8 shows contour plots for several asset-pricing moments of interest across a range of unconditional standard deviations of the policy and risk shock. These plots

[23] For example, in his first post-FOMC press conference, Chairman Bernanke noted that providing more information about monetary policy could serve "to help the public understand what you're doing, to help the markets understand what you're doing," but "the counterargument has always been that, if there was a risk, that the Chariman speaking might create unnecessary volatility in financial markets...." (FOMC Press Conference Transcript, 27 April 2011, available at www.federalreserve.gov/monetarypolicy.)

have irregular shapes for two reasons. First, although we solve the model over a rectangular grid of $\{\sigma_M^*, \sigma_\sigma\}$ combinations, these values map nonlinearly into $\{\mathrm{E}[\sigma_M], \mathrm{var}[\sigma_M]\}$ space. (In particular, for any given level of $\mathrm{var}[\sigma_M]$, there is a minimum level that $\mathrm{E}[\sigma_M]$ can take, explaining why there are no points in the northwest corner of the plots.) Second, the presence of the ELB implies nonexistence of a solution some regions of the parameter space, particularly when the shock variances are very high. (See Richter and Throckmorton (2015).) The points corresponding to these parameter values are left blank.

A decrease in the average level of MPR is equivalent to a leftward movement within the contour plots. As shown in the first two panels, such a move is consistent with lower realized volatilities of both short- and long-term interest rates. (The same is true of implied volatilities, not shown.) As noted in the Introduction, interest-rate volatility has generally trended downward since the early 1980s, consistent with such a story. However, the model suggests that an increase in the variability of MPR—a vertical movement in the plots—would have increased the volatility of long-term yields, all else equal, offsetting some of this effect.

The bottom two panels show the effects of the parameter changes on two other variables that seem to exhibit longer-run trends: the nominal term premium and the correlation between stock and nominal bond returns. Section 5.3 showed that a transitory decrease in monetary-policy risk causes stock-bond correlations to fall and increases term premia in a conditional sense, and the same logic applies here unconditionally. Because risk shocks themselves act like demand shocks in driving bond returns and stock returns in the opposite direction, a decrease in their unconditional variance reduces the stock-bond correlation and lowers the likelihood of bonds paying off in bad states of the world, thus causing term premia to fall. Quantitatively, the effects of the reduction in MPR by itself is modest. For example, reducing the unconditional standard deviation of short-rate shocks from 80 bp per quarter to 40 bp per quarter reduces the stock-bond correlation by about 20 percentage points and the term premium by about 35 basis points when the variance of MPR is low.

While these possible effects of parameter changes are smaller than the observed changes in term premia and stock-bond correlations shown in Figure 1, two additional factors that are captured by our model may account for a significant portion of the discrepancy. First, in both cases increasing the variance of MPR (moving vertically in the triangles) works in the same direction as decreasing its average value—that is, it leads to a lower term premium and a lower stock-bond correlation. Thus, our story about more-frequent communication making MPR more volatile could partly explain the observed trends.

Second, proximity of the ELB, which has been a persistent feature of the data over much of the period when greater policy communication has prevailed, also contributes to a lower stock-bond correlation and a lower term premium, since, at the ELB, bonds are a hedge for supply shocks. To see this, in Figure 9 we plot the four variables in question over ranges of $\hat{i}_t$ in the model, holding all shocks at their steady-state levels. Not surprisingly, the ELB has a large dampening effect on short-rate volatility, though smaller effects on the volatility of long-term rates because of the limited passthrough of shocks to those rates in general. For similar reasons, it has a relatively small effect on the ten-year term premium, though this effect does go in the right direction to explain the trend. On the other hand, the ELB has a fairly large effect on the stock-bond correlation, reducing it by perhaps 40 percentage points over plausible ranges of the shadow rate. Thus, we conclude that a decrease in MPR, together with an increase in its volatility and a frequently-binding ELB can explain a significant portion of the downward trends displayed in Figure 1.

8 Conclusion

We have shown how risk surrounding the exogenous component of monetary policy contributes to risk premia across asset classes. MPR helps explain why the equity premium and real and nominal term premium are all positive on average. These qualitative results hold in a broad class of models where money is not neutral.

In our specific quantitative model—which broadly matches asset-price and macro moments as well as event-study evidence—MPR on average contributes 30 to 40 basis points to the nominal ten-year term premium, roughly half of its average value. Meanwhile, it explains about 45 bp of the real ten-year term premium and 320 bp of the equity premium. Fluctuations in MPR also help to explain negative variance-risk premia on both stocks and bonds. A secular reduction in MPR and an increase in its volatility, such as plausibly resulted from the explosion in Federal Reserve communication over the last three decades, helps explain the observed long-term declines in interest-rate volatility, the nominal term premium, and the correlation between bond and stock returns, as well as the increased correlation between output and policy rates. The proximity of the effective lower bound on nominal rates enhances these effects.

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 lower and steeper nominal yield curve; positive abnormal returns on nominal bonds and negative abnormal returns on real bonds; and a lower inflation-risk premium. Most of these effects are quantitatively significant. For example, an increase in MPR that temporarily raises the conditional volatility of the short rate by 50 bp near the steady state produces a a contemporaneous decline in stock returns of 2% and an increase in 10-year bond returns of 0.5%. The ELB alters the magnitudes of these responses—for example, it significantly increases the effect of MPR on stock returns—but it does not generally change their sign.

A Details of the New Keynesian Model

[TO BE COMPLETED]

B Solution Method

A perfect foresight "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 $S_t \in \mathbf{R}^{\mathbf{N_s}}$ denote the state of the economy. Then a solution consists of mappings $\{f_1, f_2, f_3\}$ where

$$\begin{aligned} f_1 : \mathbf{R}^{\mathbf{N_s}} \to \mathbf{R} \ s.t. \ \tilde{V^C}_t = f_1(S_t), \\ f_2 : \mathbf{R}^{\mathbf{N_s}} \to \mathbf{R} \ s.t. \ \tilde{C}_t = f_2(S_t), \\ f_3 : \mathbf{R}^{\mathbf{N_s}} \to \mathbf{R} \ s.t. \ \pi_t = f_3(S_t), \end{aligned} \tag{59}$$

and which satisfy the equilibrium conditions of equations 1,2 and 3.

To obtain our solution set we employ iterative methods. To employ these methods in practice we need to discretize the statespace S_t, over which we can approximate our set of functional solutions, and build an associated transition matrix Q. We begin with discretization of the statespace S_t. We partition our set of state variables into a set of N_1 variables which follow an ar(1) $\{\mathbf{z}_1, \mathbf{z}_2, ...\mathbf{z}_{N_1}\}$ and a set of N_2 variables $\{\mathbf{h}_1, \mathbf{h}_2, ...\mathbf{h}_{N_2}\}$ which follow an ar(1) but have stochastic variance governed by $\{\mathbf{v}_1, \mathbf{v}_2, ...\mathbf{v}_{N_2}\}$ respectively, where each follows an ar(1) itself.[24] For $\mathbf{z}_i \in \{\mathbf{z}_1, \mathbf{z}_2, ...\mathbf{z}_{N_1}\}$ we use the well-known Rouwenhorst to build a discrete grid $\hat{\mathbf{z}}_\mathbf{i}$ of length N_{z_i}. For $\mathbf{v}_i \in \{\mathbf{v}_1, \mathbf{v}_2, ...\mathbf{v}_{N_2}\}$, we start by using Rouwenhorst to build a discrete grid $\hat{\mathbf{v}}_\mathbf{i}$ of length N_{v_i}. However, since $\mathbf{v}_i$ governs the conditonal variance of $\mathbf{h}_i$, we must ensure that $\hat{\mathbf{v}}_\mathbf{i}$ remains strictly positive. Therefore, if $\hat{\mathbf{v}}_\mathbf{i}$ has only positive values then we use that grid. Otherwise, we keep the largest point of the chosen grid, and create a uniform grid of same length N_{v_i} but beginning at a number just above 0 rather than the negative number chosen by Rouwenhorst. For $\mathbf{h}_i \in \{\mathbf{h}_1, \mathbf{h}_2, ...\mathbf{h}_{N_2}\}$ we use a uniform grid of length N_{h_i} which runs from 3 conditional standard deviations below the unconditional mean to 3 conditional standard deviations above the unconditional mean - where the conditional standard deviation is conditional on the largest grid point in the associated state $\mathbf{v}_i$'s grid $\hat{\mathbf{v}}_\mathbf{i}$ governing $\mathbf{h}_i$'s variance. This is critical, because to accurately represent dynamics in the state variable governing $\mathbf{h}_i$'s stochastic variance, $\mathbf{v}_i$, there must be grid points $\hat{\mathbf{h}}_\mathbf{i}$ such that, for cases where $\mathbf{v}_i$ - and therefore the conditional standard deviation of $\mathbf{h}_i$ - is large, there is nonzero conditional likelihood of transitioning to those states. If grid points for $\hat{\mathbf{h}}_\mathbf{i}$ are truncated too low, then conditional on high values of $\hat{\mathbf{v}}_\mathbf{i}$, all of these points would lie to the left of where the mass of the distribution ought to be. As such, changes in $\hat{\mathbf{v}}_\mathbf{i}$ will not lead to equivalent changes in the numerical standard deviation of $\hat{\mathbf{h}}_\mathbf{i}$.

Given our set of grids $(\mathbf{Z} = \{\hat{\mathbf{z}}_\mathbf{1}, \hat{\mathbf{z}}_\mathbf{2}, ...\hat{\mathbf{z}}_{\mathbf{N_1}}\})$,$(\mathbf{H} = \{\hat{\mathbf{h}}_\mathbf{1}, \hat{\mathbf{h}}_\mathbf{2}, ...\hat{\mathbf{h}}_{\mathbf{N_1}}\})$, and $(\mathbf{V} = \{\hat{\mathbf{v}}_\mathbf{1}, \hat{\mathbf{v}}_\mathbf{2}, ...\hat{\mathbf{v}}_{\mathbf{N_1}}\})$,

[24]Note $N_1 + 2N_2 = N_s$

we build our discrete grid $\hat{\mathbf{S}}$ as the Cartesian product of all our individual grids: That is

$$\hat{\mathbf{S}} = \{(z_1, z_2, ...z_{N_1}, h_1, h_2, ...h_{N_2}, v_1, v_2, ...v_{N_2})|z_i \in \hat{\mathbf{z}}_i \in \mathbf{Z}, h_i \in \hat{\mathbf{h_i}} \in \mathbf{H}, v_i \in \hat{\mathbf{v}}_\mathbf{i} \in \mathbf{V}\},$$

with length:

$$N = (\prod_{j}^{N_1} N_{z_j})(\prod_{q}^{N_2} N_{h_q})(\prod_{i}^{N_2} N_{v_i})$$

We turn now to building transition matrix Q. Generic transition probabilities are given by:

$$\begin{aligned} q_{ij} &= Prob\left[\hat{\mathbf{S}}_{\mathbf{t+1}} = s_j|\hat{\mathbf{S}}_{\mathbf{t}} = s_i\right] \\ &= \prod_{g=1}^{N_1} Prob\left[\hat{\mathbf{z}}_{gt+1} = s_j(\hat{\mathbf{z}}_g)|\hat{\mathbf{S}}_{\mathbf{t}} = s_i\right] \cdot \\ &\prod_{g=1}^{N_2} Prob\left[\hat{\mathbf{v}}_{gt+1} = s_j(\hat{\mathbf{v}}_g)|\hat{\mathbf{S}}_{\mathbf{t}} = s_i\right] \cdot \\ &\prod_{g=1}^{N_2} Prob\left[\hat{\mathbf{h}}_{gt+1} = s_j(\hat{\mathbf{h}}_g)|\hat{\mathbf{S}}_{\mathbf{t}} = s_i\right], \end{aligned}$$

where i, j index elements of the state space ($s_j, s_i \in \hat{\mathbf{S}}$) and $s_j(\hat{\mathbf{z}}_g)$ refers to the element of s_j corresponding to state $\hat{\mathbf{z}}_g$. Note that since the transition probabilities of $\mathbf{Z}$ and $\mathbf{V}$ are only conditionally dependent on lagged values, we have,

$$Prob\left[\hat{\mathbf{z}}_{gt+1} = s_j(\hat{\mathbf{z}}_g)|\hat{\mathbf{S_t}} = s_i\right] = Prob\left[\hat{\mathbf{z}}_{gt+1} = s_j(\hat{\mathbf{z}}_g)|\hat{\mathbf{z}}_{gt} = s_i(\hat{\mathbf{z}}_g)\right] \ \forall \ g \in \{1, 2, ..., N_1\},$$

$$Prob\left[\hat{\mathbf{v}}_{gt+1} = s_j(\hat{\mathbf{v}}_g)|\hat{\mathbf{S_t}} = s_i\right] = Prob\left[\hat{\mathbf{v}}_{gt+1} = s_j(\hat{\mathbf{v}}_g)|\hat{\mathbf{z}}_{gt} = s_i(\hat{\mathbf{v}}_g)\right] \ \forall \ g \in \{1, 2, ..., N_1\},$$

Therefore we can rewrite q_{ij} as follows:

$$\begin{aligned} q_{ij} = \prod_{g=1}^{N_1} Prob\left[\hat{\mathbf{z}}_{gt+1} = s_j(\hat{\mathbf{z}}_g)|\hat{\mathbf{z}}_{gt} = s_i(\hat{\mathbf{z}}_g)\right] \cdot \prod_{g=1}^{N_2} Prob\left[\hat{\mathbf{v}}_{gt+1} = s_j(\hat{\mathbf{v}}_g)|\hat{\mathbf{v}}_{gt} = s_i(\hat{\mathbf{v}}_g)\right] \cdot \\ \prod_{g=1}^{N_2} Prob\left[\hat{\mathbf{h}}_{gt+1} = s_j(\hat{\mathbf{h}}_g)|\hat{\mathbf{h}}_{gt} = s_i(\hat{\mathbf{h}}_g), \hat{\mathbf{v}}_{gt} = s_i(\hat{\mathbf{v}}_g)\right], \end{aligned}$$

The transition probability $q_{ij}$is formed by taking the product of the individual conditional transition probabilities of the state variables. For the AR(1) variables the conditional likelihood of transition to state $s_j(\hat{\mathbf{z}}_g)$ only depends on its lagged value and none of the other states. However, for the set $\mathbf{H}$, the probability of transitioning from one state to the next

$\left[\hat{\mathbf{h}}_{gt+1} = s_j(\hat{\mathbf{h}}_g)|\hat{\mathbf{h}}_{gt} = s_i(\hat{\mathbf{h}}_g)\right]$ depends critically on the conditional variance at the time governed by $\hat{\mathbf{v}}_{gt} = s_i(\hat{\mathbf{v}}_g)$.

The Rouwenhorst method provides us with probabilities:

$$Prob\left[\hat{\mathbf{z}}_{gt+1} = s_j(\hat{\mathbf{z}}_g)|\hat{\mathbf{z}}_{gt} = s_i(\hat{\mathbf{z}}_g)\right], \ \forall\ g \in \{1,2,...N_1\}\ \forall\ (i,j) \in \{1,2,...N\}X\{1,2,...N\}$$

In principal, these come from the Rouwenhorst transition matrix for each $\hat{\mathbf{z}}_g \in \mathbf{Z}$. The same is true for all $\hat{\mathbf{v}}_g \in \mathbf{v}$ where $v_g > 0\ \forall\ v_g \in \hat{\mathbf{v}}_g$. In the case where we have to truncate the grid to be greater than zero, we compute the point on the conditional PDF of the transition state and re-normalize such that conditional probabilities sum to one. Specifically, consider the probobilty of transitioning from some $v_{gi} \in \hat{\mathbf{v}}_g$ to some $v_{gj} \in \hat{\mathbf{v}}_g$. We have:

$$Prob\left[\hat{\mathbf{v}}_{gt+1} = v_{gj}|\hat{\mathbf{v}}_{gt} = v_{gi}\right] = \frac{\Phi\left(\frac{v_{gj}-(1-\rho_{v_g})\delta_{v_g}+\rho_{v_g}v_{gi}}{\sigma_{v_g}}\right)}{\sum_{q=1}^{N_{v_g}}\Phi\left(\frac{v_{gq}-(1-\rho_{v_g})\delta_{v_g}+\rho_{v_g}v_{gi}}{\sigma_{v_g}}\right)},$$

where $\delta_{v_g}, \rho_{v_g}, \sigma_{v_g}$ govern the mean, persistence and conditional variation of $\mathbf{v}_g$. For,

$$Prob\left[\hat{\mathbf{h}}_{gt+1} = s_j(\hat{\mathbf{h}}_g)|\hat{\mathbf{h}}_{gt} = s_i(\hat{\mathbf{h}}_g), \hat{\mathbf{v}}_{gt} = s_i(\hat{\mathbf{v}}_g)\right],$$

we follow a similar procedure using the point on the standard normal pdf, conditional on the mean governed by $s_i(\hat{\mathbf{h}}_g)$ and variance governed by $s_i(\hat{\mathbf{v}}_g)$. Specifically, consider the probobilty of transitioning from some $h_{gi} \in \hat{\mathbf{h}}_g$ to some $h_{gj} \in \hat{\mathbf{h}}_g$ conditional on being in $s_i(\hat{\mathbf{v}}_g) \in \hat{\mathbf{v}}_g$. We have:

$$Prob\left[\hat{\mathbf{h}}_{gt+1} = h_{gj}|\hat{\mathbf{h}}_{gt} = h_{gi}, \hat{\mathbf{v}}_{gt} = s_i(\hat{\mathbf{v}}_g)\right] = \frac{\Phi\left(\frac{h_{gj}-(1-\rho_{h_g})\delta_{h_g}+\rho_{h_g}h_{gi}}{s_i(\hat{\mathbf{v}}_g)}\right)}{\sum_{q=1}^{N_{h_g}}\Phi\left(\frac{h_{gq}-(1-\rho_{h_g})\delta_{h_g}+\rho_{h_g}h_{gi}}{s_i(\hat{\mathbf{v}}_g)}\right)}.$$

We are now equipped with our discretization of the statespace S_t and our associated transition matrix Q. We then approximate our solution set of functions over the discretized grid by defining $\{\hat{\mathbf{f}}_1, \hat{\mathbf{f}}_2, \hat{\mathbf{f}}_3\}$ over our set of nodes $\hat{\mathbf{S}}$. Our goal is to employ time iteration as in Judd (1996). However, because we have inetia in the policy rule (equation ((44)) our "exogenous" state space $\hat{\mathbf{S}}$ does not fully summarize the state of the economy which depends criticality on i_{t-1}. As such, we consider the "box" of Cartesian Product of $\hat{\mathbf{S}}$ and a grid of nodes $\hat{\mathbf{I}}$ for i_{t-1}. This is equivalent to defining $\{\hat{\mathbf{f}}_1, \hat{\mathbf{f}}_2, \hat{\mathbf{f}}_3\}$ for each grid point $\hat{i} \in \hat{\mathbf{I}}$.

To employ time iteration we need to compute expectations. Without inertia, this is straightforward. The conditional expectation of the output gap for example would simply be:

$$E_t[\tilde{C}_{t+1}] = dot(Q, \hat{\mathbf{f}}_2)$$

That is, the conditional expectation of the output gap is the dot product of the transition matrix with the vector which approximates consumption over the states. However, with inertia, this is no longer the case. Denote $\hat{\mathbf{f}}_2^{\hat{i}}$ as the approximation for the output gap over the set of states conditional on $i_{t-1} = \hat{i}$. The dot product above then yields the expecation condtional on $i_t = \hat{i}$:

$$dot(Q, \hat{\mathbf{f}}_2^{\hat{i}}) = E_t[\tilde{C}_{t+1} | i_t = \hat{i}]$$

We compute the above expectation for all nodes $\hat{i} \in \hat{\mathbf{I}}$, and build an interpolating function $h(\hat{i}) = E_t[\tilde{C}_{t+1} | i_t = \hat{i}]$ which provides the conditional expectation in each node conditional on $i_t = \hat{i}$. To compute the expectation of any endogenous variable at any node we evaluate this interpolated function at the interest rate observed in that node, which is given at all states through the taylor rule.

This procedure allows us to compute expectations and therefore we can employ time iteration as in Judd (1998). We do this for all macro variables and then the same for the Price-Dividend ratio for equities. Computing the term structure of interest rates does not require iteration.

References

Adrian, Tobias, Richard K. Crump, and Emanuel Moench (2013). "Pricing the Term Structure with Linear Regressions". In: *Journal of Financial Economics* 110(1), pp. 110–138.

Albuquerque, Rui et al. (2016). "Valuation risk and asset pricing". In: *The Journal of Finance* 71(6), pp. 2861–2904.

Aldrich, Eric Mark and Howard Kung (2019). "Computational methods for production-based asset pricing models with recursive utility". In: *Studies in Nonlinear Dynamics & Econometrics* 25(1), pp. 1–26.

Backus, David K, Allan W Gregory, and Stanley E Zin (1989). "Risk premiums in the term structure: Evidence from artificial economies". In: *Journal of Monetary Economics* 24(3), pp. 371–399.

Basu, Susanto and Brent Bundick (2017). "Uncertainty Shocks in a Model of Effective Demand". In: *Econometrica* 85(3), pp. 937–958.

Bauer, Michael D., Aemit Lakdawala, and Phillipe Mueller (2022). "Market-Based Monetary Policy Uncertainty". In: *Economic Journal* 132, pp. 1290–1308.

Bekaert, Geert, Seonghoon Cho, and Antonio Moreno (2010). "New Keynesian macroeconomics and the term structure". In: *Journal of Money, Credit and Banking* 42(1), pp. 33–62.

Binsbergen, Jules H. van et al. (2012). "The Term Structure of Interest Rates in a DSGE Model with Recursive Preferences". In: *Journal of Monetary Economics* 59, pp. 634–648.

Bloom, Nicholas (2009). "The Impact of Uncertainty Shocks". In: *Econometrica* 77(3), pp. 623–685.

Bloom, Nicholas (2014). "Fluctuations in Uncertainty". In: *Journal of Economic Perspectives* 28(2), pp. 153–176.

Bloom, Nicholas et al. (2018). "Really uncertain business cycles". In: *Econometrica* 86(3), pp. 1031–1065.

Bundick, Brent, Trenton Herriford, and A. Lee Smith (2022). "The Term Structure of Monetary Policy Uncertainty". In: *FRB Kansas City working paper* 22-02.

Caggiano, Giovanni, Efrem Castelnuovo, and Giovanni Pellegrino (2017). "Estimating the real effects of uncertainty shocks at the zero lower bound". In: *European Economic Review* 100, pp. 257–272.

Campbell, John Y. (1986). "Bond and stock returns in a simple exchange model". In: *The Quarterly Journal of Economics* 101(4), pp. 785–803.

Campbell, John Y., Carolin Pflueger, and Luis M. Viceira (2020). "Macroeconomic Drivers of Bond and Equity Risks". In: *Journal of Political Economy* 128(8), pp. 3148–3185.

Christiano, Lawrence J., Martin Eichenbaum, and Charles Evans (1996). "The Effects of Monetary Policy Shocks: Evidence from the Flow of Funds". In: *Review of Economics and Statistics* 78(1), pp. 16–34.

Cieslak, Anna and Pavol Povala (2016). "Information in the term structure of yield curve volatility". In: *The Journal of Finance* 71(3), pp. 1393–1436.

Dew-Becker, Ian (2014). "Bond Pricing with a Time-Varying Price of Risk in and Estimated Medium-Scale Bayesian DSGE Model". In: *Journal of Money, Credit, and Banking* 46(5), pp. 837–888.

Epstein, Larry G and Stanley E Zin (1991). "Substitution, risk aversion, and the temporal behavior of consumption and asset returns: An empirical analysis". In: *Journal of political Economy* 99(2), pp. 263–286.

Fama, Eugene F and Robert R Bliss (1987). "The information in long-maturity forward rates". In: *The American Economic Review*, pp. 680–692.

Fernandez-Villaverde, Jesus et al. (2015). "Nonlinear Adventures at the Zero Lower Bound". In: *Journal of Economic Dynamics and Control* 57, pp. 182–204.

Gilchrist, Simon, Jae W. Sim, and Egon Zakrajsek (2014). "Uncertainty, Financial Frictions, and Investment Dynamics". In: *NBER working paper w20038*.

Gourio, Francois and Phuong Ngo (2020). "Risk Premia at the ZLB: A Macroeconomic Interpretation". In: *FRB Chicago working paper 2020-1*.

Hanson, Samuel G and Jeremy C Stein (2015). "Monetary policy and long-term real rates". In: *Journal of Financial Economics* 115(3), pp. 429–448.

Hordahl, Peter, Oreste Tristani, and David Vestin (2008). "The Yield Curve and Macroeconomic Dynamics". In: *Economic Journal* 118, pp. 1937–1970.

Hsu, Alex, Li, and Francisco Palomino (2021). "Real and nominal equilibrium yield curves". In: *Management Science* 67(2), pp. 1138–1158.

Husted, Lucas, John Rogers, and Bo Sun (2020). "Monetary Policy Uncertainty". In: *Journal of Monetary Economics* 115, pp. 20–36.

Judd, Kenneth L. (1996). *Numerical Methods in Economics*. MIT Press.

Jurado, Kyle, Sydney C Ludvigson, and Serena Ng (2015). "Measuring uncertainty". In: *American Economic Review* 105(3), pp. 1177–1216.

Justiniano, Alejandro and Giorgio E. Primiceri (2008). "The Time-Varying Volatility of Macroeconomic Fluctuations". In: *American Economic Review* 98(3), pp. 604–641.

Kim, Don H. and Jonathan H. Wright (2005). "An Arbitrage-Free Three-Factor Term Structure Model and the Recent Behavior of Long-Term Yields and Distant-Horizon Forward Rates". In: *FEDS working paper* 2005-33.

King, Thomas B. (2019). "Expectation and Duration at the Effective Lower Bound". In: *Journal of Financial Economics* 134(3), pp. 736–760.

Kung, Howard (2015). "Macroeconomic Linkages between Monetary Policy and the Term Structure of Interest Rates". In: *Journal of Financial Economics* 115(1), pp. 42–57.

Leduc, Sylvain and Zheng Liu (2016). "Uncertainty shocks are aggregate demand shocks". In: *Journal of Monetary Economics* 82, pp. 20–35.

Li and Francisco Palomino (2014). "Nominal Rigidities, Asset Returns, and Monetary Policy". In: *Journal of Monetary Economics* 66, pp. 210–225.

Mehra, Rajnish and Edward C Prescott (1985). "The equity premium: A puzzle". In: *Journal of monetary Economics* 15(2), pp. 145–161.

Mumtaz, Haroon and Francesco Zanetti (2013). "The Impact of the Volatility of Monetary Policy Shocks". In: *Journal of Money, Credit, and Banking* 45(4), pp. 535–558.

Nakata, Taisuke (2017). "Uncertainty at the zero lower bound". In: *American Economic Journal: Macroeconomics* 9(3), pp. 186–221.

Pflueger, Carolin (2023). "Back to the 1980s or Not? The Drivers of Inflation and Real Risks in Treasury Bonds". In: *Working paper (February)*.

Richter, Alexander W and Nathaniel A Throckmorton (2015). "The zero lower bound: frequency, duration, and numerical convergence". In: *The BE Journal of Macroeconomics* 15(1), pp. 157–182.

Rotemberg, Julio J. and Michael Woodford (1996). "An Optimization-Based Econometric Framework for the Evaluation of Monetary Policy". In: *NBER Macroeconomics Annual 1997* 12, pp. 297–346.

Rudebusch, Glenn D. and Eric T. Swanson (2012). "The Bond Premium in a DSGE Model with Long-Run Real and Nominal Risks". In: *American Economic Journal: Macroeconomics* 4(1), pp. 105–143.

Swanson, Eric T. (2006). "Have Increases in Federal Reserve Transparency Improved Private Sector Interest Rate Forecasts?" In: *Journal of Money, Credit, and Banking* 38(3), pp. 297–346.

Swanson, Eric T. (2021). "A Macroeconomic Model of Equities and, Real, Nominal, and Defaultable Debt". In: *Working paper*.

C Tables

Table 1: Assumptions about Effects of Shocks in the Linear Model

Shock	Dividends	Inflation	Real and nominal interest rates	SDF
Demand	↑	↑	↑	↓
Supply	↑	↓	↓	↓
Monetary	↓	↓	↑	↑

Table 2: Effects of Higher Risk - Risk Premia

Source of risk	Nominal Term Prem.	Inflation Risk Prem.	Real Risk Prem.	Equity Risk Prem.
Supply	↑	↑	↑	↑
Demand	↓	↓	↓	↑
Monetary	↑	↓	↑	↑

Table 3: Effects of Higher Uncertainty - Correlations

Source of risk	Inflation v Output	Stock v bond returns	Stock v inflation	Nominal Rates v output
Supply	↓	↑	↓	↓
Demand	↑	↓	↑	↑
Monetary	↑	↑	↑	↓

Table 4: Parameter Values in NK Model

Parameter	Value	Economic Interpretation
α	-160	Risk aversion
β	.99	Time preference
ρ	1/3	IES=1.5
δ	.2	Leverage
$\pi*$	.034	Inflation target (annualized)
$r*$	.016	Steady-state real rate (annualized)
ϕ_y	.5/4	Policy rule - output
ϕ_π	1.5	Policy rule - inflation
φ_i	.75	Policy rule - inertia
κ	.05	Phillips Curve slope
ν	.052	Phillips Curve supply shock multiplier
φ_a	.95	Supply shock persistence
σ_a	.005	Supply shock std. dev.
φ_σ	.9	MPR shock persistence
σ_M^*	.0045	MPR shock mean
σ_σ	.001438	MPR shock std. dev.
φ_ξ	.9	Demand shock persistence
σ_ξ	.00066	Demand shock std. dev.

[1] Notes: All values calibrated to a quarterly frequency, except where noted.

Table 5: Empirical and Model Moments

Unconditional Moment	Data	Model
$E[i_t]$	.044	.049
$Std[i_t]$	.034	.025
$E[y_t^{\$(40)}]$	.060	.056
$Std[y_t^{\$(40)}]$	.032	.006
$E[R^{eq}]$	.118	.084
$Std[R^{eq}]$	.173	.156
$Std[y_t]$	.021	.027
$Std[\pi_t]$	.028	.018
$Corr(y, \pi)$	.18	.3
$Corr(y, i)$	.16	-.82
$Corr(R^{eq}, R^{\$(40)})$	.16	.31
$Corr(R^{\$(1)}, R^{\$(40)})$	.03	.28
Sharpe ratio: 10y nom. bonds	.261	.302
Sharpe ratio: Equities	.425	.234

[1] Note: Data are the 3-month Treasury bill rate, the 10-year nominal Treasury yield, the change in the Wilshire 5000 total-return index, PCE inflation, and the CBO output gap. All data are quarterly, 1971 - 2023, reported as annualized values where appropriate, and obtained from the FRED database.

Table 6: Correlation of the Term Structure

Moment	Unconditional	Steady State	ELB
$Corr\left[NTP^{(10)}, \mathscr{E}^{(10)}[i]\right]$	-.12	-.25	-.14
$Corr\left[NTP^{(5)}, \mathscr{E}^{(5)}[i]\right]$	-.15	-.29	-.17

[1] Notes: The table shows the correlations between the term-premium and expectations components of nominal yields at 5- and 10-year horizons, measured unconditionally, at the steady-state, and at the ELB.

Table 7: Model-Implied Risk Premia

	$NTP^{(40)}$	$RTP^{(40)}$	$IRP^{(40)}$	ERP	VRP: equities	VRP: short rate
Total	76	44	32	347	-18	-5
Contribution from:						
Monetary policy risk ($\sigma_{M,t}$)	39	49	-10	324	0	0
Uncertainty about Uncertainty (σ_σ)	-6	0	-6	14	-18	-5

Notes

Table 8: Risk premia and their sources

Premium	$NTP^{(40)}$	$RTP^{(40)}$	$IRP^{(40)}$	ERP	VRP: equities	VRP: short rate	$\sigma[Supply]$	$\sigma[Demand]$	$\sigma[MP]$	$\sigma[MPR]$
Total	76	44	32	347	18	5	.0157	.0015	.0058	.0029
ϵ^S and ξ	47	-2	49	32	0	0	.0157	.0015	-	-
ϵ^S, ξ and ϵ^M	83	45	38	339	2	0	.0157	.0015	.0058	-
ϵ^M only	39	49	-10	324	0	0	-	-	.0058	-
ϵ^M and ϵ^σ	33	49	-16	338	18	5	-	-	.0058	.0029

Risk premia under various shock specifications. $\sigma[\cdot]$ refers to the unconditional standard deviation of the shocks in the model.

Table 9: Event Study Regressions

	SPX	VIX	5 year	10 Year	10 Year (Real)	5Y TP (ACM)	10Y TP (ACM)	5Y TP (KW)	10Y TP (KW)
SRU	-0.0866**	22.0590**	0.6010**	0.6849**	0.7235**	0.4338**	0.5179**	0.2735**	0.3581**
	(0.0307)	(4.4476)	(0.1487)	(0.1583)	(0.2001)	(0.1080)	(0.1539)	(0.0644)	(0.0767)
MPS	-0.0160	-0.2794	0.5272**	0.3207**	0.3287**	-0.0107	-0.1296*	0.1430**	0.1366**
	(0.0116)	(1.6773)	(0.0561)	(0.0597)	(0.0762)	(0.0407)	(0.0580)	(0.0243)	(0.0289)
Obs	197	197	197	197	157	197	197	197	197
R^2	0.0901	0.1442	0.5060	0.3376	0.2609	0.0967	0.0566	0.3583	0.3222
Adj. R^2	0.0807	0.1354	0.5009	0.3308	0.2513	0.0874	0.0469	0.3516	0.3153

Standard errors are in parentheses. Statistical significance indicators: $^+$: $p < 0.10$, *: $p < 0.05$, **: $p < 0.01$. Data is from Bauer, Lakdawala, and Mueller (2022), and regressions follow closely from their specification. The sample runs from February 1994 to September 2020 and removes the Financial Crisis period between July 2007 and August 2009. I(ZLB) is defined an indicator variable which takes a value of 1 in the OIS Rate is below 30 basis points at the time of a meeting.

Table 10: Event Study Regressions — With ZLB Effect

	SPX	VIX	5 year	10 Year	10 Year (Real)	5Y TP (ACM)	10Y TP (ACM)	5Y TP (KW)	10Y TP (KW)
SRU	-0.0527	11.4532*	0.3504*	0.4997**	0.2334	0.3247**	0.4641**	0.1706*	0.2239**
	(0.0333)	(4.4599)	(0.1579)	(0.1724)	(0.2368)	(0.1182)	(0.1698)	(0.0691)	(0.0818)
SRU * I(ELB)	-0.1766**	55.1580**	0.6024*	0.6093^{+}	0.8188*	0.3546	0.4030	0.3056*	0.4335**
	(0.0637)	(8.5283)	(0.3019)	(0.3298)	(0.3581)	(0.2260)	(0.3247)	(0.1320)	(0.1564)
MPS	-0.0211^{+}	1.2947	0.5280**	0.3298**	0.3165**	-0.0055	-0.1152^{+}	0.1463**	0.1428**
	(0.0116)	(1.5538)	(0.0550)	(0.0601)	(0.0756)	(0.0412)	(0.0592)	(0.0241)	(0.0285)
MPS * I(ELB)	0.0682	-21.1909**	0.4160^{+}	0.0927	0.3408	0.0600	-0.2686	0.0947	0.0777
	(0.0502)	(6.7146)	(0.2377)	(0.2596)	(0.2656)	(0.1780)	(0.2556)	(0.1040)	(0.1232)
SRU (Total ELB Effect)	-0.23	66.6	0.95	1.11	1.05	0.68	0.87	0.48	0.66
Obs	197	197	197	197	157	197	197	197	197
R^2	0.126	0.299	0.546	0.360	0.323	0.119	0.065	0.400	0.373
Adj. R^2	0.107	0.285	0.537	0.346	0.305	0.101	0.045	0.387	0.360

Standard errors are in parentheses. Statistical significance indicators: $^{+}$: $p < 0.10$, * : $p < 0.05$, ** : $p < 0.01$. Data is from Bauer, Lakdawala, and Mueller (2022), and regressions follow closely from their specification. The sample runs from February 1994 to September 2020 and removes the Financial Crisis period between July 2007 and August 2009. I(ELB) is defined an indicator variable which takes a value of 1 if the OIS Rate is below 30 basis points at the time of an FOMC announcement.

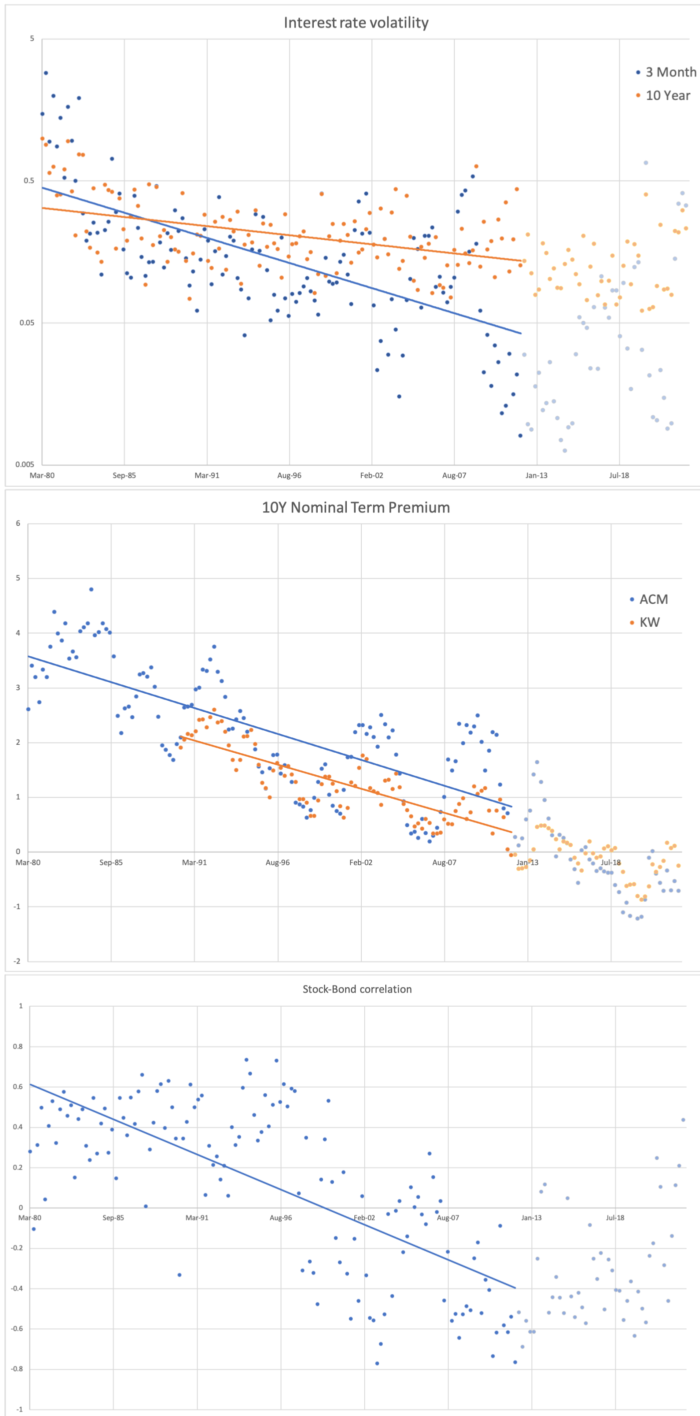
Table 11: Model Implied Effect of Interest Rate Uncertainty Shock

	SPX	VIX	5 year	10 Year	10 Year (Real)	5Y TP	10Y TP
Data: Non-ELB	-.053	11	.35	.50	.23	.17 - .32	.22 - .46
Model: Steady State	-.087	14	-.11	-.07	.12	.19	.10
Data: ELB	-.23	67	.95	1.11	1.05	.48 - .68	.66 - .87
Model: $i = -.025$	-.171	24	-.26	-.16	.24	.37	.20
Model: $i = -.050$	-.422	47	-.71	-.42	.58	.86	.47
Model: $i = -.075$	-2.48	224	-3.52	-2.11	3.33	4.33	2.38

Shown is the ratio of the change in the selected variable relative to the change in the conditional standard deviation of the 1 year yield induced by a shock to the conditional variance of the policy shock in the model ($\text{var}_t\left[\sigma^M_{t+1}\right]$). This is ratio is computed when the shock takes place at the steady state (row 2) and in three other locations where the $t-1$ shadow rate (i_{t-1}) is below zero to illustrate the effect of the ELB. The model-implied values are compared to the empirical estimates from different samples of the data, based on the point estimates reported in table 10.

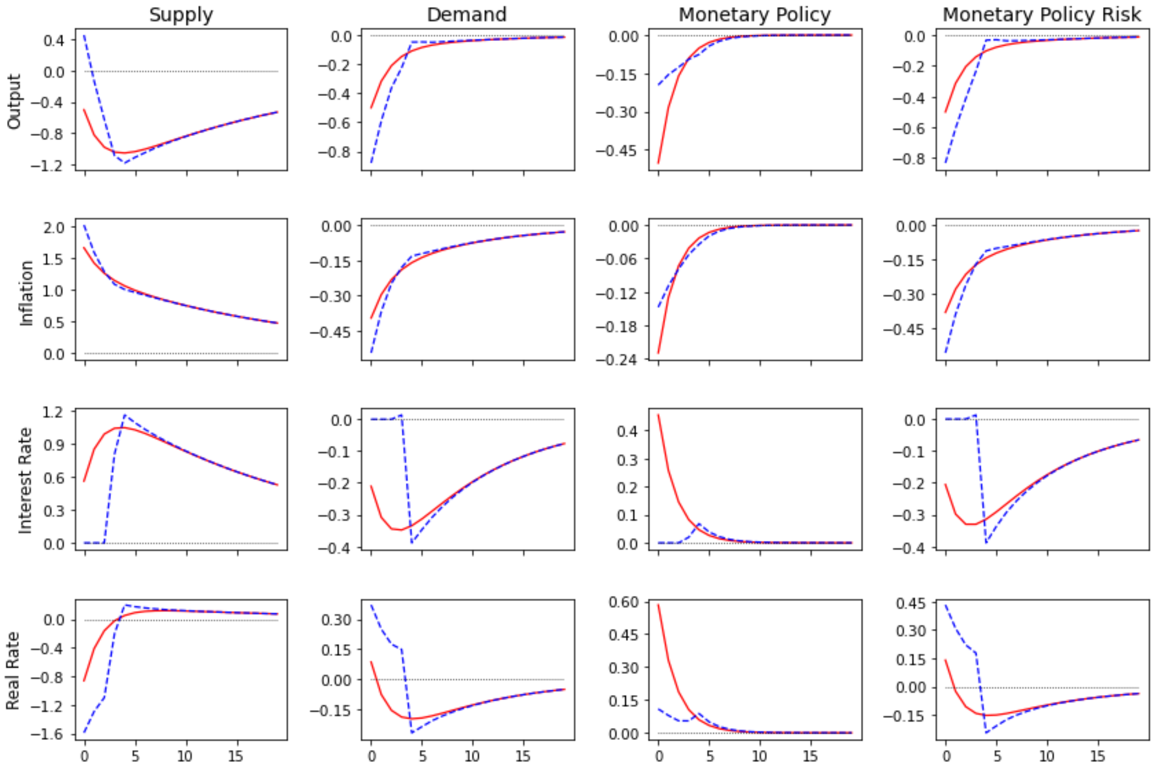
D Figures

Figure 1: Trends in Selected Asset-Price Moments



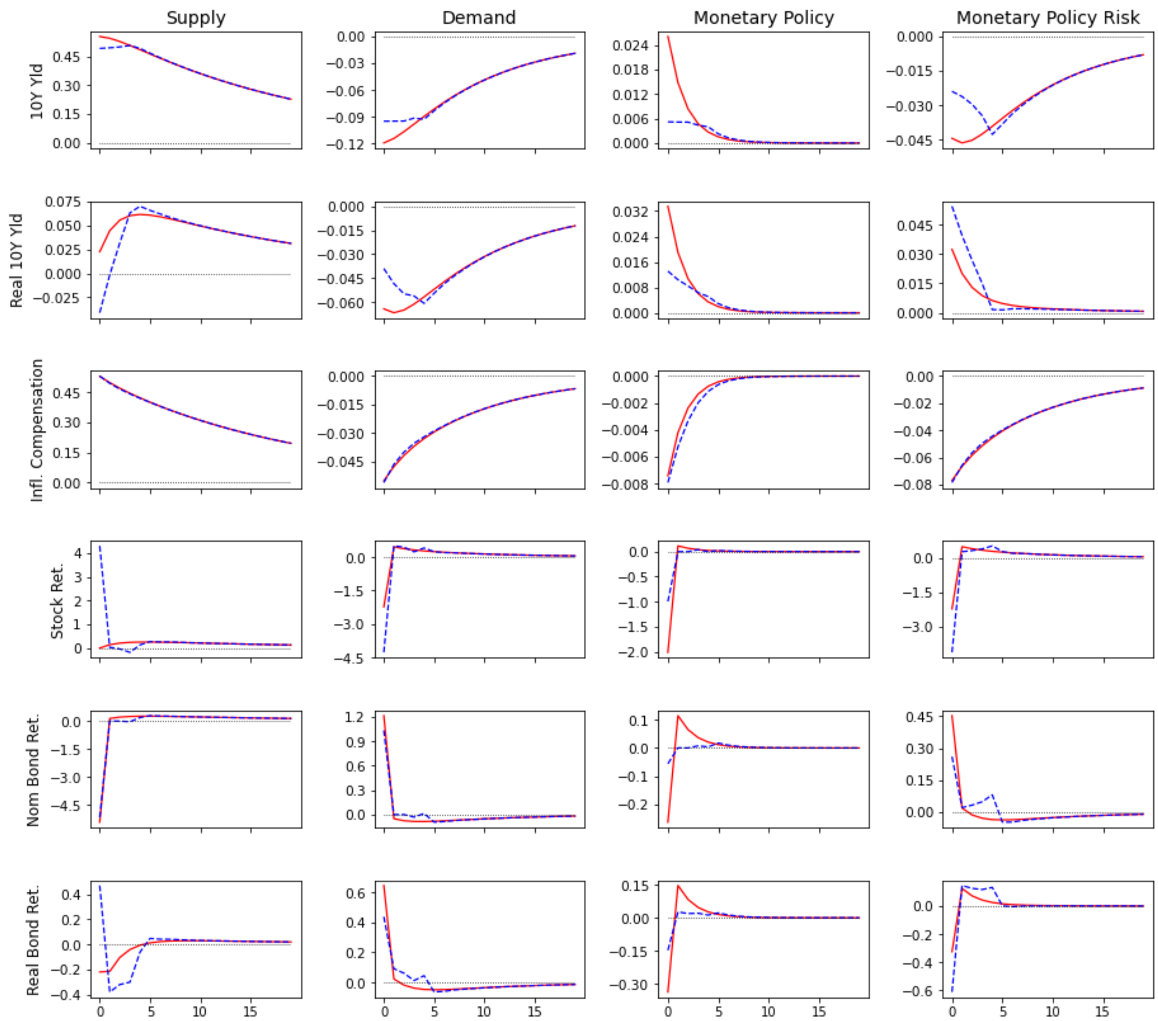
Dots are quarterly sample statistics computed from daily data. Solid lines are linear time trends over the period 1980-2012. For term premia, ACM denotes the model of Adrian, Crump, and Moench (2013) and KW denotes Kim and Wright (2005) (begins in 1990).

Figure 2: Impulse Responses to Shocks: Macro Variables



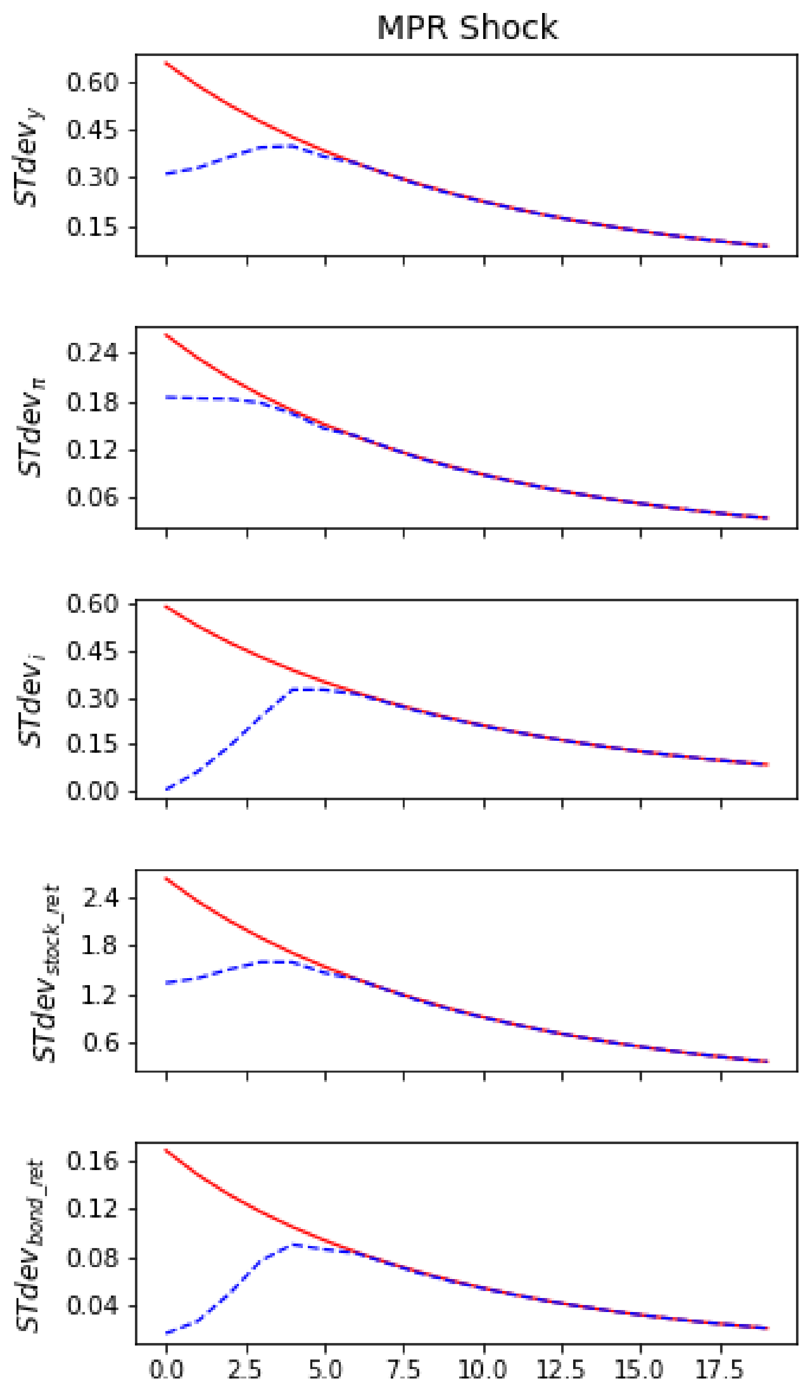
Units are in Percentage Points. Red is away from the ELB and blue-dashed is starting from a point which, with no other shocks, would remain at the ELB for 4 periods.

Figure 3: Impulse Responses to Shocks: Asset Prices



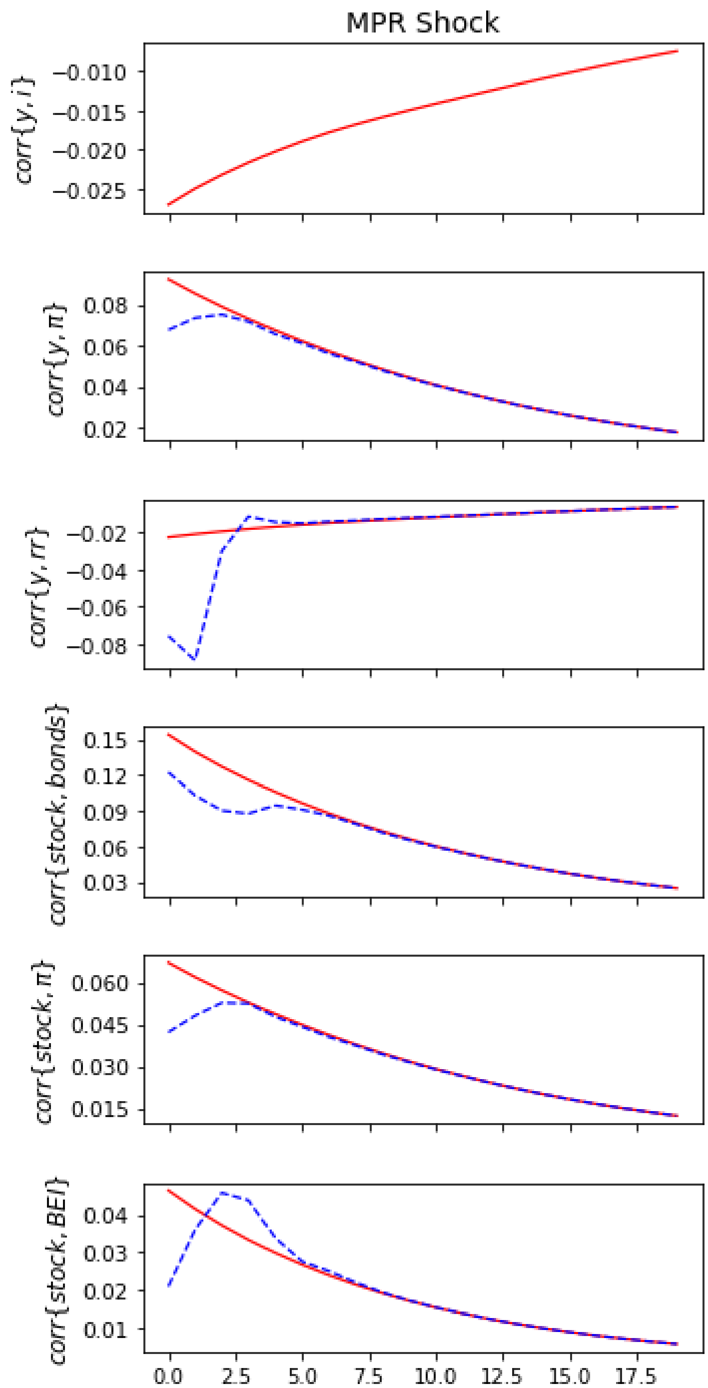
Units are in Percentage Points. Red is away from the ZLB and blue-dashed is starting from a point which, with no other shocks, would remain at the ZLB for 4 periods.

Figure 4: Impulse Responses to Shocks: Conditional Second Moments



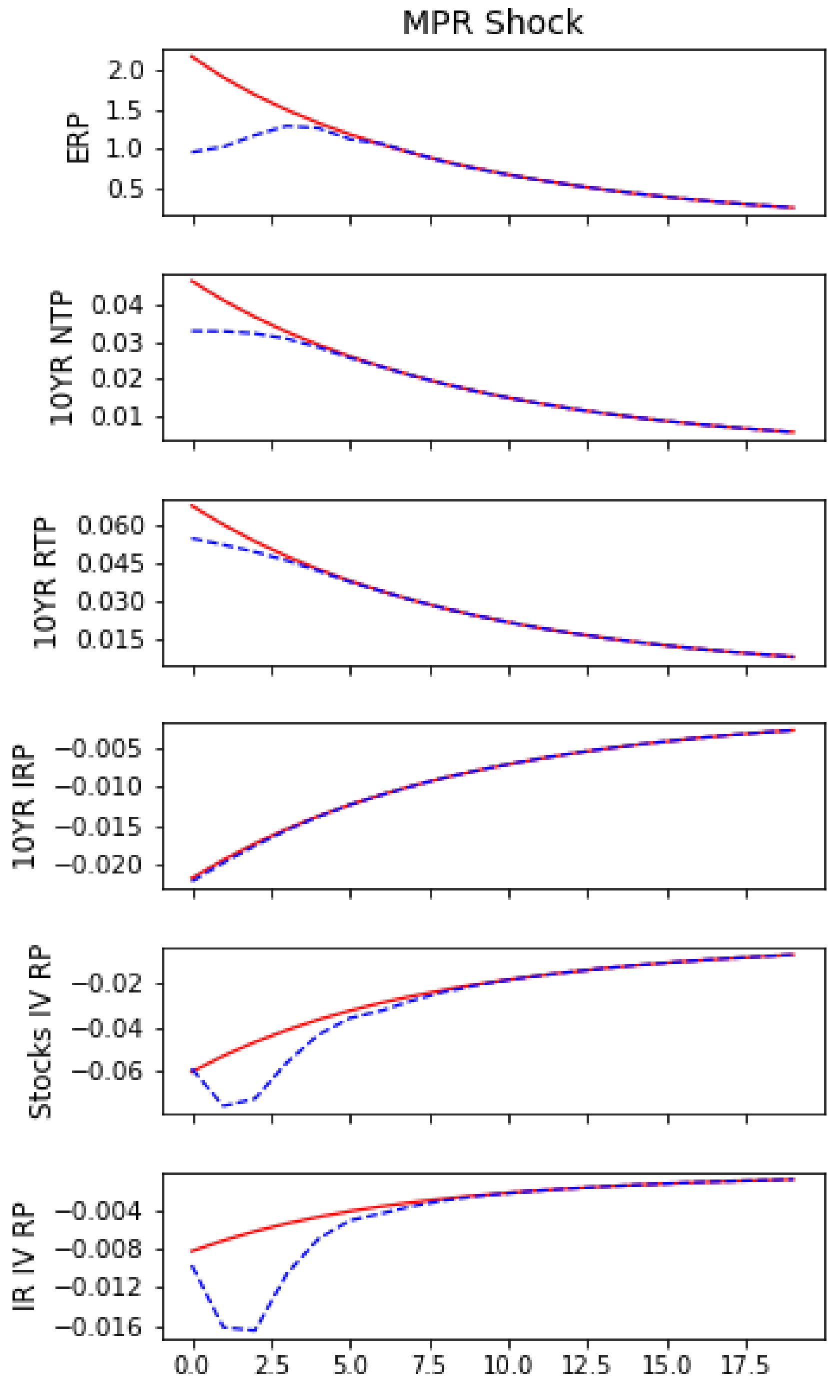
Units are in Percentage Points. Red is away from the ZLB and blue-dashed is starting from a point which, with no other shocks, would remain at the ZLB for 4 periods.

Figure 5: Impulse Responses to Shocks: Conditional Correlations



Red is away from the ZLB and blue-dashed is starting from a point which, with no other shocks, would remain at the ZLB for 4 periods.

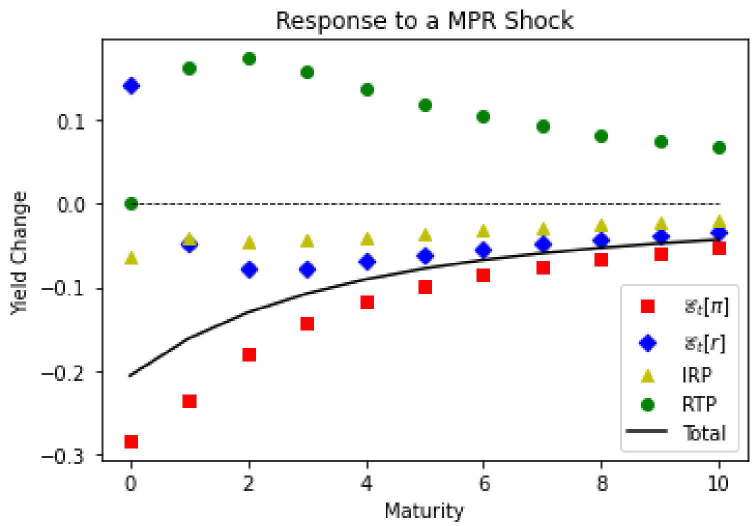
Figure 6: Impulse Responses to Shocks: Risk Premia



Units are in Percentage Points. Red is away from the ZLB and blue-dashed is starting from a point which, with no other shocks, would remain at the ZLB for 4 periods.

Figure 7: Impulse Responses to Shocks: Yield Curve

(a) Away From ELB



(b) At ELB

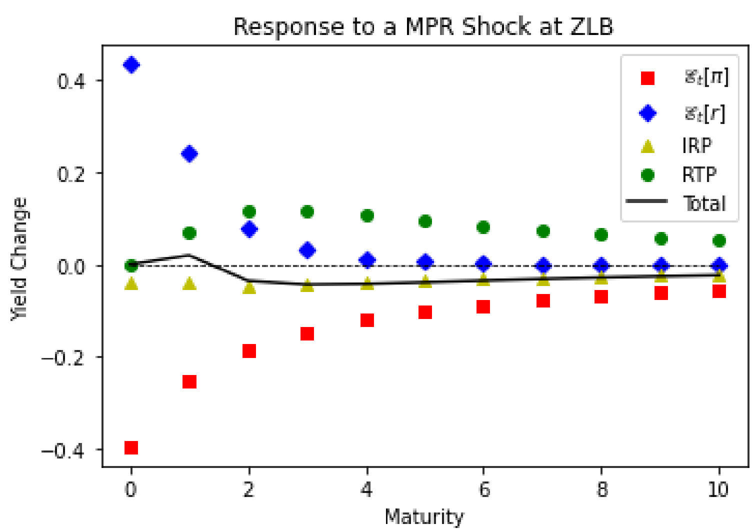
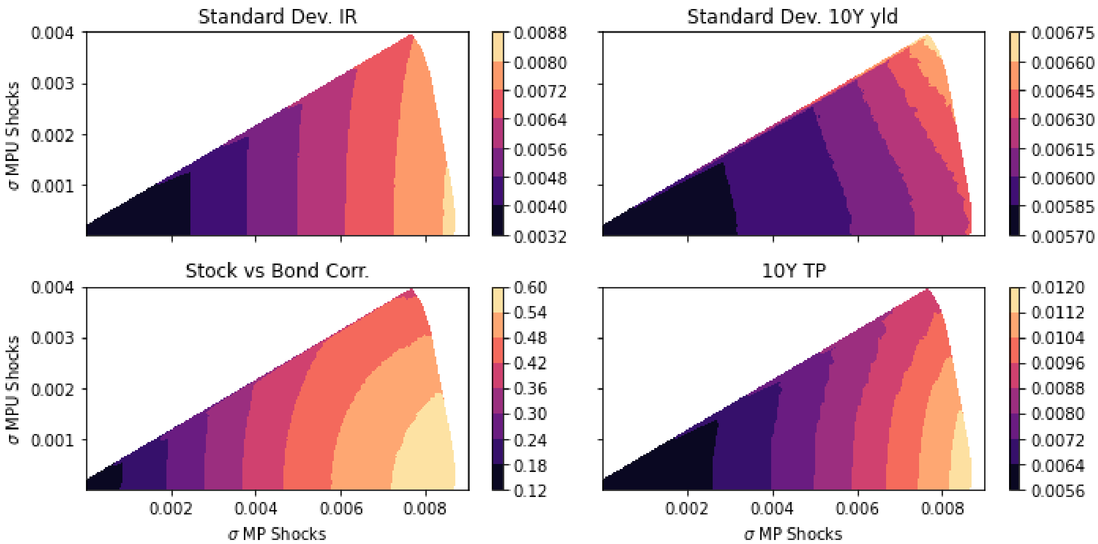
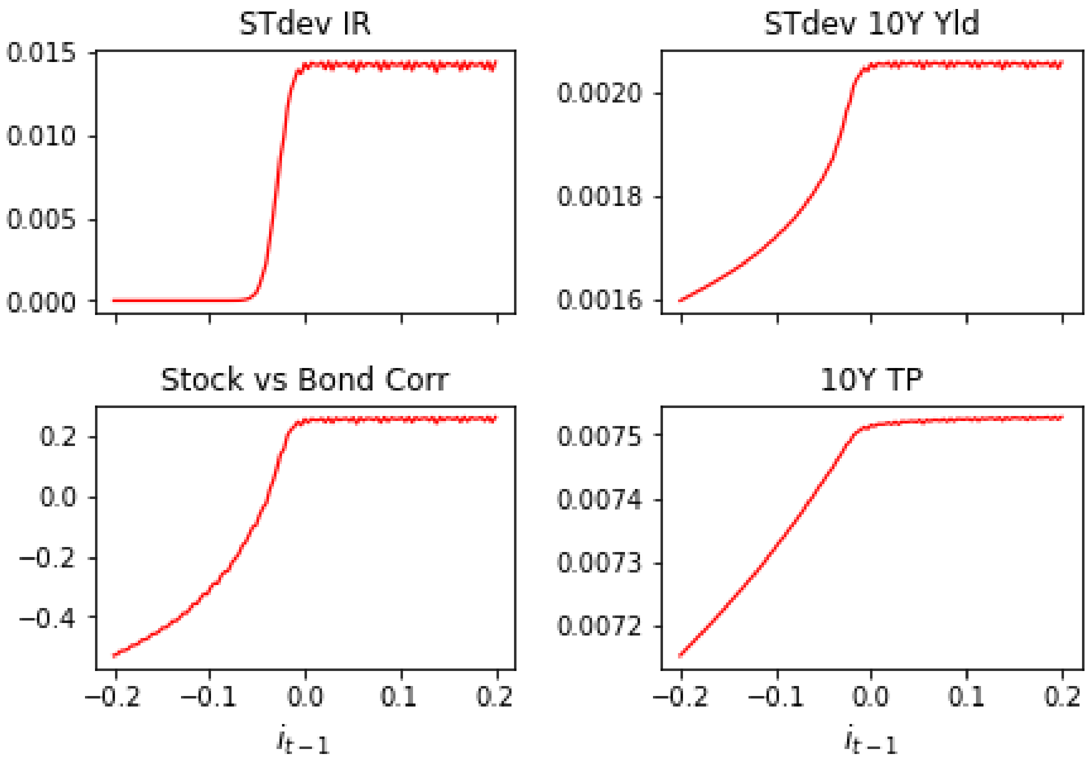


Figure 8: Unconditional Moment Contour Plots



Notes:

Figure 9: ZLB Effect



Notes: