(Un)expected Monetary Policy Shocks and Term Premia

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SUMMARY

The term structure of interest rates is crucial for the transmission of monetary policy to financial markets and the macroeconomy. Disentangling the impact of monetary policy on the components of interest rates, expected short rates and term premia, is essential to understanding this channel. To accomplish this, we provide a quantitative structural model with endogenous, time-varying term premia that are consistent with empirical findings. News about future policy, in contrast to unexpected policy shocks, has quantitatively significant effects on term premia along the entire term structure. This provides a plausible explanation for partly contradictory estimates in the empirical literature.

Keywords: DSGE model, Bayesian estimation, Time-varying risk premia, Monetary policy.

1. INTRODUCTION

Gauging how monetary policy tools affect the entire term structure of interest rates is important for understanding the monetary transmission mechanism. We provide an estimated medium-scale macro-finance DSGE model of the term structure with endogenous, time-varying term premia to do just that. As term structure comprises expected future short term interest rates and term premia, understanding the effect of monetary policy shocks on the term structure necessitates the disentanglement of the effects on both components. There is, however, no consensus so far on the impact of monetary policy shocks on these premia (for a discussion see Nakamura and Steinsson, 2018). The empirical literature faces the challenge of identifying monetary policy shocks when multiple instruments, say, unexpected changes in the policy rate and news about future policy, are used concurrently. This challenge has increased in relevance since the 1990s as the Federal Reserve has increasing relied on communication in transmitting monetary policy (see, for example, Gürkaynak et al., 2005a, Campbell et al., 2012). While a structural model in general can help to disentangle these instruments, the existing literature has trouble providing a tractable framework that enables an empirical analysis of time-varying term premia that price the equilibrium risk of the model's structural shocks.

This paper makes two important contributions to the literature. First, we provide an estimated medium scale macro-finance model using Bayesian likelihood methods and U.S. macroeconomic and Treasury bond time series. The estimated model implies historical time series of term premia that match those found in reduced form empirical estimates without sacrificing the macroeconomic fit. Second, we then use this model to study the impact of monetary policy shocks on the term structure. We specifically use our structural production-based[1] asset pricing framework to differentiate between unexpected changes of the policy rate and news about future monetary policy on

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[1] This is in contrast to the consumption-based frameworks of, for example, Wachter (2006), Bansal and Shaliastovich (2013), and Creal and Wu (2020).

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endogenous, time-varying term premia. Thus, our approach is similar in spirit to empirical studies such as Creal and Wu (2017) as news and uncertainty shocks are related conceptually. Our structural investigation however studies the effects of first moment shocks (e.g. announcement or news shocks) on the term structure, while Creal and Wu (2017) focus on the interaction of second moment shocks (uncertainty shocks) with the real economy and vice versa.

We find that unexpected changes in the policy rate have limited effects but news about future policy strong effects on term premia of all maturities. This highlights the importance of the nonlinear in risk approximation that we pursue here, with announcements of future policy operating through the risk channel alongside the expectations channel on many standard analyses. The differentiated impact of policy news is in line with the recent empirical literature. Furthermore, this difference can help understand partly contradictory estimates of the effects of monetary policy shocks on the term structure in the empirical literature.

Many studies have found that nominal term premia are sizable, volatile, and have been on the decline since the beginning of the 1980s (see Rudebusch et al. (2007) for a summary, as well as Bauer et al. (2012) and Bauer et al. (2014) for dissents).[2] The literature has also emphasized the role of real term premia relative to inflation risk premia, (see, e.g., Gürkaynak et al., 2010, Chernov and Mueller, 2012) manifesting itself in an upward sloping real yield curve. This sets the yardstick for our structural model. To this end, a joint model of the interaction between the macroeconomy and the term structure of interest rates that allows for time-varying risk premia is needed.

The structural macroeconomic literature has broadly taken two approaches to address this need. One is to incorporate financial market frictions or segmented markets (e.g., Gertler and Karadi, 2011, 2013, Carlstrom et al., 2017, Sims and Wu, 2021) that constrain financial market participants and generate premia as spreads between market rates of different maturities and the risk-free short rate. Swanson (2019) argues that such premia are risk-neutral expected losses from, say, default, motivating the second approach taken in the literature. Here, time-varying risk premia result through the pricing via the stochastic discount factor of endogenous conditional heteroskedasticities generated by the nonlinearities in the model. We contribute to this second approach and estimate a model in which risk premia arise in frictionless markets through their nonlinear consequences for risk.

Within this second approach, the macro-finance literature has offered punctual solutions for selective empirical facts,[3] but the investigation of a comprehensive model is hampered by the computationally burdensome solution and estimation methods. Consequentially, the literature has focused either on matching selected moments (Rudebusch and Swanson, 2012, Andreasen et al., 2018) or on highly stylized models (van Binsbergen et al., 2012). This tension is poignantly noted by Gürkaynak and Wright (2012, p. 354): "A general problem with a structural model ... is that it is challenging to maintain computational tractability and yet obtain time-variation in term premia."

This paper provides an estimated, fully structural joint model of the macroeconomy and the real and nominal term structure of interest rates with a frictionless production-based asset market approach. Thus, the term premia in our model price the risk in equilibrium generated by higher moments of structural shocks and do not rely on segmented markets or financial frictions. Our medium scale New Keynesian macro-finance DSGE model is fitted to U.S. macroeconomic and Treasury bond time series from 1983:Q1 to 2007:Q4 using Bayesian likelihood methods, covering the Great Moderation and stopping before the onset of the Great Recession.[4] To do so, we apply

[2]Bauer et al. (2012) and Bauer et al. (2014) argue for their bias-correction methods for the estimation of affine Gaussian dynamic term structure models and show that their approach, among others, points to a reduced downward trend in the term premia on forward rates. We have adapted their code to calculate the ten year term premia we examine here. While their bias-correction method predicts a less pronounced trend than other measures, it is still there, with a linear time trend implying a decline in the 10 year term premia by nearly two percentage points since 1984.

[3]E.g., Piazzesi and Schneider (2007) highlight the role of recursive preferences and supply shocks for an upward sloping nominal yield curve and Wachter (2006) points to the role of habit formation for real yields.

[4]The online appendix explores the consequences of extending the sample out to 2019:Q4. As discussed there, we present as our baseline sample the period here due especially to the constancy of monetary policy during this

a novel procedure that captures both constant and time-varying risk premia while maintaining linearity in states and shocks (Meyer-Gohde, 2016). The closest contributions to ours are Andreasen (2011) and Dew-Becker (2014). While the former is silent about model predictions of stylized macroeconomic facts and other financial facts besides the nominal term structure, the latter, in addition, predicts historical time-varying term premia which are at odds with the empirical literature.

In contrast, our estimated structural model implies a historical 10-year term premium comparable in level, pattern, and volatility with recent reduced-form empirical estimates, see figure 1.[5] The left side of figure 1 compares the 10-year nominal term premium implied by the model (black solid line) for the estimation sample between 1983 and 2008, i.e., the model parameters are estimated and the term premium is smoothed using the data from 1983 to 2008. The right side (red dashed line) compares the same term premium but is smoothed using data out to 2019Q4 without reestimating the parameters, i.e., the model parameters are held at the values estimated from the 1983-2008 baseline sample period.

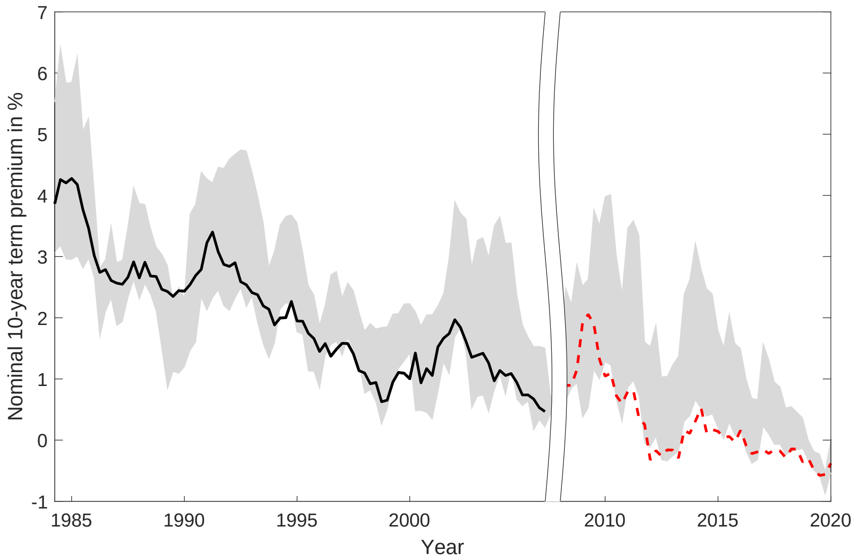


Figure 1: Model implied 10-year nominal term premium.

Note: The figure compares the model implied term premium (black line and red-dashed line) with a range of corresponding estimates in the literature (gray areas). In particular, the gray area presents the range (maximum and minimum) of the estimates from models developed by Kim and Wright (2005), Rudebusch and Wu (2008), Bernanke et al. (2004), Adrian et al. (2013), Bauer et al. (2012) and Bauer (2018). The first three measures were calculated by Rudebusch et al. (2007) and Rudebusch et al. (2006). A description of the estimates can be found there.

The model implies both an upward sloping nominal yield curve in line with the data and an upward sloping real yield curve in line with empirical estimates (see, for example, Gürkaynak et al., 2010, Chernov and Mueller, 2012). This is in contrast to work by van Binsbergen et al. (2012), Andreasen (2012), Swanson (2019) for example, which generally attribute a stronger insurance-like character to real bonds leading to flat or downward sloping real yield curves.[6]

This is accomplished through the combination of Rudebusch and Swanson (2012), who show

period.

[5]We are very grateful to Eric T. Swanson and Michael Bauer for sharing their estimates with us. For the measures associated with Bauer et al. (2012), we adapted their code to calculate our ten year term premium instead of the five-on-five year forward premium in the original analysis, as well as extending their sample to Q1 2020.

[6]The UK is often an example for such flat or downward sloped real yield curve, but this is maybe due to strong institutional (pensions legislation) impacts on real long-term bonds and therefore rather an anomaly (see Campbell et al., 2009, Bank of England, 1999, p. 8).

that supply shocks induce a negative correlation between inflation an consumption which, in turn, produces an upward sloping nominal yield curve; and the inducement of positive autocorrelation in consumption growth in the face of habits that leads households to desire to draw down precautionary savings at longer maturities, producing negative correlation between marginal utility and payouts on real bonds and, hence, an upward sloping real yield curve following Wachter (2006) and Hördahl et al. (2008). Additionally, our results suggest that 2/3 of the average slope of the nominal term structure is related to real rather than to inflation risk and an upward sloping inflation risk premium – both consistent with recent estimates in the literature (see, for example, Abrahams et al., 2016).

While the general impact of monetary policy on the term structure of interest rates is largely agreed upon (Piazzesi, 2005), there exists considerable disagreement on the effects of monetary policy shocks on term premia. Hanson and Stein (2015) argue for quantitatively strong effects of monetary policy shocks on real risk premia, while Nakamura and Steinsson (2018) find rather small effects despite sizable effects on nominal term premia overall. More strikingly, the literature disagrees even on the qualitative effects: Gertler and Karadi (2015) and Abrahams et al. (2016) find a positive, while Crump et al. (2016) and Nakamura and Steinsson (2018) find a negative correlation between the policy rate and nominal term premia. There are many potential reasons for this disagreement; we take our cue from Ramey (2016) and focus on disentangling the effects of structural shocks arising from different monetary policy tools on term premia. This disentangling follows naturally from the structural model and avoids ongoing challenge for empirical models,[7] especially given the unconventional monetary policy literature on, say, central bank communication (or news about future monetary policy) on the term structure and its components (Gürkaynak et al., 2005a, Nakamura and Steinsson, 2018).

We find that an unexpected tightening of monetary policy via a simple innovation to the Taylor rule reduces risk premia particularly at longer maturities, in line with the empirical work by Crump et al. (2016) but in contrast to, e.g., Gertler and Karadi (2015). Yet overall such a shock has quantitatively limited effects, in line with findings from other structural models (see, for example, Rudebusch and Swanson, 2012). In contrast, a shock to the inflation target or unconditional forward guidance[8] reveals news about future paths of macroeconomic variables, affecting households' precautionary motives and thereby their demand for risk premia. Specifically, a change in the inflation target can be interpreted as a change in the systematic component of monetary policy (see Cogley et al., 2010) as it affects agents' perception of the macroeconomy in the longer run. With, say, a reduced inflation target, the same absolute change of inflation is associated with a stronger automatic response of monetary policy. Similarly, forward guidance communicates the expected path of future short rates and is likewise informative as to the central bank's commitment to allow higher inflation in the future. The recent empirical literature (Gürkaynak et al., 2005a, Nakamura and Steinsson, 2018) argues that news revealed by monetary policy about its future path has strong effects on the term structure of interest rates. In line with this, we find that a shock to the inflation target has strong effects on the term structure of interest rates and term premia across all maturities. As laid out by Rudebusch and Swanson (2012), a change to the inflation target introduces long-run (nominal) risk that strongly affects households' expectation formation and precautionary savings motives. Unconditional forward guidance likewise affects term premia substantially, causing real term premia and inflation risk premia to rise as agents expect more volatile inflation and output in the future in line with the empirical findings of Akkaya et al. (2015).

Our analysis of the effects of monetary policy shocks on term premia suggests that the quantitatively large effects found in the empirical literature seem primarily driven by monetary policy news about its medium- or long-term stance rather than changes in the policy rate. Beyond identification, our Monte-Carlo analysis of small samples with common empirical models demonstrates

[7]See, for example, Campbell et al. (2016) for a discussion of potential shortcomings in isolating the effects of monetary policy in the recent literature.

[8]For a discussion of different forms of forward guidance see Campbell et al. (2012) and Akkaya et al. (2015). Particularly such a distinction is a significant challenge in many empirical approaches (see, for example, the discussion in Nakamura and Steinsson, 2018, Campbell et al., 2016).

that estimation uncertainty results in a wide range of quantitatively and qualitatively different point estimates. Thus, our structural model can help to rationalize empirical findings. Finally, given the model's tractability it can be applied and extended to serve as a building block for future research.

The remainder of the paper reads as follows: Section 2 presents the model. Then, section 3 describes the solution method, the data, and the Bayesian estimation approach in greater detail. Section 4 presents the estimation results and discusses the model fit. The effects of unexpected and expected monetary policy on the term structure and comparisons with empirical estimates are presented in section 5. Section 6 concludes the paper.

2. MODEL

In the following section, we present our dynamic stochastic general equilibrium (DSGE) model, a standard New Keynesian model but with recursive preferences (Epstein and Zin, 1989, 1991, Weil, 1989) and both real and nominal long-run risk (Bansal and Yaron, 2004, Gürkaynak et al., 2005b), highlighted in the literature for the explanation of many financial moments in consumption-based asset pricing.

2.1 Firms

A perfectly competitive representative firm produces the final good y_t, which is aggregated from a continuum of intermediate goods $y_{j,t}$ by $y_t = \left(\int_0^1 y_{j,t}^{(\theta_p-1)/\theta_p} dj\right)^{\theta_p/(1-\theta_p)}$ where $\theta_p > 1$ the elasticity of substitution. Cost-minimization yields the demand function for the intermediate good $y_{j,t} = \left(P_{j,t}/P_t\right)^{-\theta_p} y_t$ and the aggregate price level is then $P_t = \left(\int_0^1 P_{j,t}^{1-\theta_p} dj\right)^{\frac{1}{1-\theta_p}}$.

The intermediate good j is produced by a monopolistic competitive firm with

$$y_{j,t} = \exp\{a_t\} k_{j,t-1}^{\alpha} \left(z_t l_{j,t}\right)^{1-\alpha} - z_t^{+} \Omega_t \tag{1}$$

where $k_{j,t}$ and $l_{j,t}$ denote capital and labor inputs used for production by the jth intermediate good producer. The capital share is α and Ω_t the fixed costs of production. Short-run risk is present via the stationary technology shock a_t that follows

$$a_t = \rho_a a_{t-1} + \sigma_a \epsilon_{a,t}, \text{ with } \epsilon_{a,t} \overset{iid}{\sim} N(0,1) \tag{2}$$

and long-run risk via the stochastic aggregate productivity trend z_t

$$\mu_{z,t} \doteq \ln\{z_t/z_{t-1}\} = (1-\rho_z)\,\bar{\mu}_z + \rho_z \mu_{z,t-1} + \sigma_z \epsilon_{z,t}, \text{ with } \epsilon_{z,t} \overset{iid}{\sim} N(0,1). \tag{3}$$

Croce's (2014) specification for productivity, which mirrors Bansal and Yaron's (2004) for consumption, is captured as a special case.[9]

Alongside the stochastic trend z_t, we assume a deterministic trend in the relative price of investment Υ_t with $\exp\{\bar{\mu}_\Upsilon\} = \Upsilon_t/\Upsilon_{t-1}$. Following Altig et al. (2011) we define $z_t^+ = \Upsilon_t^{\frac{\alpha}{1-\alpha}} z_t$ as an overall measure of technology with associated trend $\mu_{z^+,t} = \frac{\alpha}{1-\alpha}\bar{\mu}_\Upsilon + \mu_{z,t}$.

Finally, we scale Ω_t by z_t^+ to ensure the existence of a balanced growth path and let production costs be time-varying as proposed by Andreasen (2011).

$$\ln\left(\Omega_t/\bar{\Omega}\right) = \rho_\Omega \ln\left(\Omega_{t-1}/\bar{\Omega}\right) + \sigma_\Omega \epsilon_{\Omega,t}, \text{ with } \epsilon_{\Omega,t} \overset{iid}{\sim} N(0,1). \tag{4}$$

[9] Short-run risk (SSR) is white noise as in Croce (2014) and Bansal and Yaron (2004) if $\rho_a = 1$

$$\varpi_t \doteq \exp\{a_t\} z_t^{1-\alpha} \Rightarrow \ln\{\varpi_t/\varpi_{t-1}\} = (1-\alpha)\ln\{z_t/z_{t-1}\} + a_t - a_{t-1} = \underbrace{(1-\alpha)\,\mu_{z,t}}_{\text{LRR}} + \underbrace{(\rho_a - 1)\,a_{t-1} + \sigma_a \epsilon_{a,t}}_{SRR}$$

Following Calvo (1983), intermediate goods firms face staggered price setting and adjust their prices only with probability $(1-\gamma_p)$ each period. Non-adjusted prices evolve according to the indexation rule: $P_{j,t} = P_{j,t-1}\pi_{t-1}^{\xi_p}$, where $\pi_t = P_t/P_{t-1}$ is gross inflation. Firms that are able to adjust their prices, choose the same price $\tilde{p}_t = P_{j,t}$ to maximize the present value of their expected future profits, accounting for demand, indexation, and the readjustment probability. Firms are owned by the households and discount with their real stochastic discount factor M_{t+1}. The optimality conditions, where mc_t are real marginal costs, are

$$\mathcal{K}_t = y_t\tilde{p}_t^{-\theta_p} + \gamma_p E_t\left[M_{t+1}\left(\pi_t^{\xi_p}/\pi_{t+1}\right)^{1-\theta_p}\left(\tilde{p}_t/\tilde{p}_{t-1}\right)^{-\theta_p}\mathcal{K}_{t+1}\right] \text{ and} \tag{5}$$

$$\mathcal{K}_t = \frac{\theta_p}{\theta_p-1}y_t mc_t\tilde{p}_t^{-\theta_p-1} + \gamma_p E_t\left[M_{t+1}\left(\pi_t^{\xi_p}/\pi_{t+1}\right)^{-\theta_p}\left(\tilde{p}_t/\tilde{p}_{t-1}\right)^{-\theta_p-1}\mathcal{K}_{t+1}\right]. \tag{6}$$

2.2 Households

We assume that the representative household has recursive preferences to disentangle risk aversion and the intertemporal elasticity of the substitution (IES). Following Rudebusch and Swanson (2012), the value function of the household is

$$V_t = \begin{cases} u_t + \beta\left(E_t\left[V_{t+1}^{1-\sigma_{EZ}}\right]\right)^{\frac{1}{1-\sigma_{EZ}}} & \text{if } u_t > 0 \text{ for all } t \\ u_t - \beta\left(E_t\left[\left(-V_{t+1}\right)^{1-\sigma_{EZ}}\right]\right)^{\frac{1}{1-\sigma_{EZ}}} & \text{if } u_t < 0 \text{ for all } t \end{cases} \tag{7}$$

where u_t is the period utility kernel and $\beta \in (0,1)$ the subjective discount factor.

Similarly to Andreasen et al. (2018), we assume the following utility kernel

$$u_t = \exp\{\varepsilon_{b,t}\}\left[\frac{1}{1-\gamma}\left(\left(\frac{c_t - bh_t}{z_t^+}\right)^{1-\gamma} - 1\right) + \frac{\psi_L}{1-\chi}\left(1-l_t\right)^{1-\chi}\right] \tag{8}$$

with consumption c_t, the habit h_t, hours l_t, and preference parameters γ, χ, and ψ_L. We assume an external habit in last period's aggregate consumption $h_t = C_{t-1}$, the degree of which is controlled by $b \in (0,1)$. The preference shock $\varepsilon_{b,t}$ is given by

$$\varepsilon_{b,t} = \rho_b\varepsilon_{b,t-1} + \sigma_b\epsilon_{b,t}, \text{ with } \epsilon_{b,t} \overset{iid}{\sim} N(0,1) \tag{9}$$

The household's period budget constraint equates real expenditures with income

$$c_t + I_t/\Upsilon_t + b_t + T_t = w_t l_t + r_t^k k_{t-1} + b_{t-1}\exp\left\{R_{t-1}^f\right\}/\pi_t + \int_0^1 \Pi_t(j)\,dj. \tag{10}$$

Expenditures comprise consumption, investment I_t, a lump-sum tax T_t, and a one-period bond b_t that accrues the risk-free nominal interest R_t^f in the following period; while income comprises labor income $w_t l_t$ with w_t the real wage, capital service income $r_t^k k_{t-1}$, the pay-off from last period's bonds b_{t-1}, and the dividends from the intermediate firms – indexed by j – owned by households $\Pi(j)$.

Households own the physical capital stock, which accumulates as

$$k_t = (1-\delta)\,k_{t-1} + \exp\{\varepsilon_{i,t}\}\left(1 - \frac{\nu}{2}\left(\frac{I_t}{I_{t-1}} - \exp\{\bar{\mu}_{z^+} + \bar{\mu}_\Upsilon\}\right)^2\right) I_t \tag{11}$$

δ is the depreciation rate and $\nu \geq 0$ controls the investment adjustment costs as in Christiano et al. (2005), which are zero along the balanced growth path via $\exp\{\bar{\mu}_{z^+} + \bar{\mu}_\Upsilon\}$. $\varepsilon_{i,t}$ represents an investment shock that evolves as

$$\varepsilon_{i,t} = \rho_i\varepsilon_{i,t-1} + \sigma_i\epsilon_{i,t}, \text{ with } \epsilon_{i,t} \overset{iid}{\sim} N(0,1) \tag{12}$$

2.3 Monetary Policy

We follow Rudebusch and Swanson (2008, 2012) and specify monetary policy via the following interest rate rule

$$R_t^f = \rho_R R_{t-1}^f + (1-\rho_R)\left(\bar{r} + \ln \pi_t + \frac{\eta_y}{4}\ln\left(\frac{y_t}{z_t^+ \bar{y}}\right) + \frac{\eta_\pi}{4}\ln\left(\frac{\pi_t^4}{\pi_t^*}\right)\right) + \frac{\sigma_m}{4}\epsilon_{m,t} \tag{13}$$

where $\bar{r}$ is the deterministic steady-state real interest rate. The policy parameters ρ_R, η_y, and η_π characterize the degree of monetary policy's aim to smooth the interest rate, stabilize deviations in output from its balanced growth path – $\ln\left(y_t/z_t^+\bar{y}\right)$ – and those in inflation from the central bank's inflation target π_t^* – $\ln\left(\pi_t^4/\pi_t^*\right)$. Departures from these aims are captured by $\epsilon_{m,t} \overset{iid}{\sim} N(0,1)$. Following Gürkaynak et al. (2005b), the inflation target is time-varying and is governed by

$$\log \pi_t^* - 4\log\bar{\pi} = \rho_\pi\left(\log \pi_{t-1}^* - 4\log\bar{\pi}\right) + 4\zeta_\pi\left(\log \pi_{t-1} - \log\bar{\pi},\right) + \sigma_\pi \epsilon_{\pi,t} \tag{14}$$

with $\epsilon_{\pi,t} \overset{iid}{\sim} N(0,1)$ a shock to the inflation target.

2.4 Aggregation and Market Clearing

The aggregate market clearing constraint in the goods market is given by

$$p_t^+ y_t = \exp\{a_t\} k_{t-1}^\alpha \left(z_t l_t\right)^{1-\alpha} - z_t^+ \Omega_t \tag{15}$$

where $l_t = \int_0^1 l(j,t)\ dj$ and $k_t = \int_0^1 k(j,t)\ dj$ are aggregated labor and capital. Price dispersion, $p_t^+ = \int_0^1 \left(\frac{P_{j,t}}{P_t}\right)^{-\theta_p} dj$, arises from staggered price setting and evolves as

$$p_t^+ = (1-\gamma_p)\left(\tilde{p}_t\right)^{-\theta_p} + \gamma_p\left(\pi_{t-1}^{\xi_p}/\pi_t\right)^{-\theta_p} p_{t-1}^+ \tag{16}$$

The economy's aggregate resource constraint implies that

$$y_t = c_t + I_t/\Upsilon_t + g_t \tag{17}$$

where government expenditures $g_t = \bar{g} z_t^+ \exp\{\varepsilon_{g,t}\}$ grow with the economy, are financed by lump-sum taxes $g_t = T_t$, and are subject to shocks via $\varepsilon_{g,t}$ given as

$$\varepsilon_{g,t} = \rho_g \varepsilon_{g,t-1} + \sigma_g \epsilon_{g,t}, \text{ with } \epsilon_{g,t} \overset{iid}{\sim} N(0,1) \tag{18}$$

Finally, the aggregate price index is $1 = \gamma_p\left(\pi_{t-1}^{\xi_p}/\pi_t\right)^{1-\theta_p} + (1-\gamma_p)\left(\tilde{p}_t\right)^{1-\theta_p}$.

2.5 The Nominal and Real Term Structures

The nominal and real term structures follow the procedures of, e.g., Rudebusch and Swanson (2008, 2012) and Andreasen (2012) identically: assets are priced following standard no-arbitrage arguments as the sum of their stochastically discounted state-contingent payoffs in period $t+1$. For example, the price of a default free n-period zero-coupon bond that pays one unit of cash at maturity satisfies

$$P_{n,t} = E_t\left[M_{t,t+n}^{\$} 1\right] = E_t\left[M_{t,t+1}^{\$} P_{n-1,t+1}\right] \tag{19}$$

where $M_{t,t+1}^{\$}$ is the household's nominal stochastic discount factor given by

$$M_{t,t+1}^{\$} = \beta\frac{\lambda_{t+1}}{\lambda_t \pi_{t+1}}\left(V_{t+1}\right)^{-\sigma_{EZ}} E_t\left[V_{t+1}^{1-\sigma_{EZ}}\right]^{\frac{\sigma_{EZ}}{1-\sigma_{EZ}}} \tag{20}$$

with λ_t the marginal utility of consumption. The continuously compounded yield to maturity on the n-period zero-coupon nominal bond is $\exp\left\{-nR^{\$}_{n,t}\right\} = P^{\$}_{n,t}$.

Following, e.g., Rudebusch and Swanson (2012), the term premium is the difference between a bond's yield and its unobserved risk-neutral equivalent yield. This risk-neutral bond, which also pays one unit of cash at maturity, is priced as

$$\hat{P}_{n,t} = \exp\left\{-R^f_t\right\} E_t\left[\hat{P}_{n-1,t+1}\right] \tag{21}$$

In contrast to eq. (19), discounting is performed using the risk-free rate and the nominal term premium on a bond with maturity n is given by

$$TP^{\$}_{n,t} = \frac{1}{n}\left(\log \hat{P}_{n,t} - \log P^{\$}_{n,t}\right) \tag{22}$$

Similarly, we can derive the yield to maturity of a real bond $R_{n,t}$ as well as its risk-neutral equivalent, leading analogously to the associated real term premium $TP_{n,t}$. Finally, we follow the literature and define inflation risk premia $TP^{\pi}_{n,t}$ as

$$TP^{\pi}_{n,t} = TP^{\$}_{n,t} - TP_{n,t} \tag{23}$$

3. MODEL SOLUTION AND ESTIMATION

3.1 Solution Method

We solve the model with the method of Meyer-Gohde (2016), delivering a nonlinear in risk, but linear in states approximation at the means of the endogenous variables.[10] Unlike standard higher order perturbations or affine approximation methods, this allows us to use the standard set of macroeconometric tools for estimation and analysis of linear models, without limiting the approximation to the certainty-equivalent approximation around the deterministic steady state. We adjust the points and slopes of the decision rules for risk out to the second moments of the exogenous processes to capture both constant and time-varying risk premium, as well as the effects of conditional heteroskedasticity (e.g. van Binsbergen et al., 2012). Our resulting linear in states approximation is

$$y_t \simeq \tilde{y}(\sigma) + y_y(\tilde{y}(\sigma), 0, \sigma)\left(y_{t-1} - \tilde{y}(\sigma)\right) + y_\varepsilon(\tilde{y}(\sigma), 0, \sigma)\varepsilon_t \tag{24}$$

where y_t are endogenous and ε_t exogenous variables, $\tilde{y}(\sigma)$ the means of the endogenous variables, and $y_y(\tilde{y}(\sigma), 0, \sigma)$ and $y_\varepsilon(\tilde{y}(\sigma), 0, \sigma)$ the first derivatives of the policy function evaluated at the means $\tilde{y}(\sigma)$. The hyperparameter σ scales the distribution of the exogenous variables with $\sigma = 1$ returning the stochastic model we analyze and $\sigma = 0$ its deterministic counterpart. All of the objects in the approximation can be recovered from Taylor series in σ using the derivative information available from standard higher order perturbations at the deterministic steady state. The online appendix provides a self-contained overview of the derivations involved in this approximation.

The tension between the nonlinearity need to capture the time-varying effects of risk underlying asset prices on the one hand and the difficulties of using nonlinear estimation routines on such models on the other is highlighted by van Binsbergen et al. (2012), who model inflation as exogenous in a New Keynesian model to make the particle filter tractable. The advantage of a linear in state approximation for estimation has also been noted by, e.g., Ang and Piazzesi (2003), Hamilton and Wu (2012), Dew-Becker (2014). Our approach compromises between nonlinearity in risk and the endogenous stochastic discount factor to price financial variables consistent with

[10] Meyer-Gohde (2016) provides derivations for adjustments around the deterministic and stochastic steady states, along with those around the mean that we derive in the online appendix and apply here, accuracy checks and formal justifications for the method.

the macroeconomy on the one hand, and the need for linearity in states to make the estimation of medium scale policy relevant models feasible on the other.[11] In contrast to other conceptually similar approximations, the method of Meyer-Gohde (2016) allows us to go beyond providing a linear in state approximation that depends somehow on risk and instead be able to define our approximation rigorously via perturbation as providing an approximation of the mean of the stochastic variables and their slopes at these means, accurate out to the second order moments of the stochastic driving forces. That is, it is explicitly transparent which components of risk we capture and which we do not. To further reduce the computational burden, we apply the PoP method of Andreasen and Zabczyk (2015) that solves the model in a two-step fashion.

3.2 Data

We estimate the model with quarterly U.S. data covering the Great Moderation (1983:Q1 to 2007:Q4) due to the constancy of monetary policy during this period – see online appendix for a discussion of extending the sample out to 2019:Q4. While the systematic behavior of monetary policy is an important driver of the yield curve, as pointed out, for example, by Campbell et al. (2014), we chose a time episode which is characterized by a relatively stable monetary policy regime.First, it is widely accepted in the literature that the U.S. faced a systematic change in monetary policy after Paul Volcker became chairman of the Federal Reserve. Second, the start of the Great Recession, the financial crisis of 2008, along with the zero interest policy rates that prevailed from December 2008 onward marks another structural change in U.S. monetary policy.

Our estimation uses four macroeconomic time series, six time series from the nominal yield curve, and two time series of survey data on interest rate forecasts.[12] The macroeconomic series comprise real GDP growth, real private investment growth, real private consumption growth, and annualized GDP deflator inflation. While the last is measured in levels, the remaining variables are expressed in per capita log-differences using the civilian noninstitutional population over 16 years (CNP16OV) series from the U.S. Department of Labor, Bureau of Labor Statistics.

The nominal yield curve is measured by the 1-quarter, 1-year, 3-year, 5-year, and 10-year annualized interest rates of U.S. Treasury bonds. The data are from Gürkaynak et al. (2007) with the exception of the 1-quarter rate, where we use the 3-month T-Bill rate from the Board of Governors of the Federal Reserve System. To have a consistent description of the yield curve, we use this interest rate as the policy rate ($R_t^f = R_{1,t}^{\$}$) in our model instead of the effective Fed funds rate.[13]

Among others, Kim and Orphanides (2012), Andreasen (2011) have shown that survey data on interest rate forecasts can improve the identification of term structure models. For this reason, we incorporate 1 and 4-quarter ahead expectations of the 3-month T-Bill from the Survey of Professional Forecasters into the estimation.

3.3 Bayesian Estimation

We now present our priors for the parameters we estimate and the calibration of those we do not. Given the choice of our observable variables and the characteristics of our model, for example, the highly stylized labor market, some of the model parameters can hardly be expected to be identified. These parameters are calibrated either following the literature or related to our observables and are summarized in table I. The remaining parameters of the model are estimated.

We calibrate the steady state growth rates, $\bar{z}^+$ and $\bar{\Psi}$ to 0.54/100 and 0.08/100 which implies quarterly growth rates of 0.54 and 0.62 percent for GDP and investment as in our sample.

[11] See the online appendix for a comparison with generalized impulse responses using a third order perturbation as well as an investigation into the role of certainty nonequivalence in our risk-adjusted linear approximation. The results contained there indicate that the risk-sensitive linear approximation is successful in capturing the dynamic effects of changes in risk on the model's endogenous variables.

[12] See online appendix for details on the source and a description of all data used in this paper.

[13] We have decided to focus on the nominal yield curve only. The reason is that valid data for real yields, TIPS, are only available from the beginning of the 2000s onwards and, secondly, those data are based on CPI inflation data while we have decided to use the GDP deflater.

Moreover, we calibrate the capital depreciation rate, δ, to 10% per year and the share of capital, α, in the production function to 1/3. We also assume that in the deterministic steady state, the labor supply $\bar{l}$ and government consumption to GDP ratio $\bar{g}/\bar{y}$ are 1/3 and 0.19, respectively. The discount rate β is set equal to 0.99 and the steady state elasticity of substitution between intermediate goods θ_p to 6, implying a markup of 20%. Following Andreasen et al. (2018), we set the price indexation $\xi_p = 0$ and calibrate the Frisch elasticity of labor supply FE to 0.5. Hence, we can solve recursively for $\chi = 1/FE \cdot \left(1/\bar{l} - 1\right)$.

Description	Symbol	Value
Technology trend in percent	$\bar{z}^{+}$	0.54/100
Investment trend in percent	$\bar{\Psi}$	0.08/100
Capital share	α	1/3
Depreciation rate	δ	0.025
Price markup	$\theta_p/(\theta_p - 1)$	1.2
Price indexation	ξ_p	0
Discount factor	β	0.99
Frisch elasticity of labor supply	FE	0.5
Labor supply	$\bar{l}$	1/3
Ratio of government consumption to output	$\bar{g}/\bar{y}$	0.19

Table I: Parameter calibration.

Since our focus is to jointly explain macroeconomic and asset pricing facts, we pay special attention to selected first and second moments. The practical problem boils down to having just one observation on the means, e.g., of the slope, curvature, and level of the yield curve, while there are many observations to identify parameters crucial for the model dynamics. To mitigate this imbalance, we apply an endogenous prior approach similar to Del Negro and Schorfheide (2008) and Christiano et al. (2011) and begin with a set of initial priors, $p(\theta)$, independent across parameters. Then, we use two sets of first and second moments from a pre-sample,[14] treating them in separate blocks to capture potentially different precisions of beliefs regarding these moments.

We focus on the first moments of inflation and as well as means of level, slope, and curvature factors of the yield curve. We include the mean of inflation because the non linearities in our model impose strong precautionary motives that push the predicted ergodic mean of inflation away from its deterministic steady state, $\bar{\pi}$, as is also discussed by Tallarini (2000) and Andreasen (2011). The second moments of interest are a set of variances of macroeconomic variables (GDP growth, consumption growth, investment growth, inflation, and the policy rate).[15]

4. ESTIMATION RESULTS

We now turn to the results of our estimation. We begin with the estimated parameters, turn then to the predicted first and second moments of endogenous variables, and conclude with a comparison of the estimated components of the ten-year yield predicted by our model with those from the literature.

4.1 Parameter Estimates

As discussed in section 3, our solution method, unlike standard perturbations (e.g. Andreasen et al., 2018), maintains linearity in states and shocks which allows us to estimate the model with

[14] We follow Christiano et al. (2011) and use the actual sample as our pre-sample because of the monetary regime changes.

[15] In the online appendix, we describe the method of endogenously formed priors regarding first and second moments as well as its practical application in the paper.

the standard Bayesian techniques familiar to linear DSGE analysis. We estimate the posterior mode of the distribution and employ a random walk Metropolis-Hasting algorithm to simulate the posterior distribution of the parameters, quantifying the uncertainty of our estimates. We run two chains, each with 100,000 parameter vector draws where the first 50% have been discarded. Table II provides the resulting posterior mode, posterior mean and the 90% posterior credible set of the estimated parameters. The results indicate that the posterior distributions of all structural parameters are well approximated and differ from the initial prior distribution.[16]

Name	Symbol	Mode	Mean	5%	95%
Relative risk aversion	RRA	89.860	91.427	75.581	108.489
Calvo parameter	γ_p	0.853	0.855	0.843	0.866
Investment adjustment	ν	1.417	1.440	1.204	1.667
Habit formation	b	0.685	0.679	0.614	0.741
Intertemporal elas. substitution	IES	0.089	0.089	0.077	0.101
Steady state inflation	$100\,(\bar{\pi}-1)$	1.038	1.034	0.981	1.091
Interest rate AR coefficient	ρ_R	0.754	0.752	0.718	0.786
Interest rate inflation coefficient	η_π	3.124	3.164	2.839	3.491
Interest rate output coefficient	η_y	0.156	0.159	0.114	0.204
Inflation target coefficient	$100\zeta_\pi$	0.210	0.242	0.109	0.366
AR coefficient technology	ρ_a	0.366	0.356	0.304	0.412
AR coefficient preference	ρ_b	0.820	0.817	0.793	0.843
AR coefficient investment	ρ_i	0.956	0.955	0.949	0.961
AR coefficient gov. spending	ρ_g	0.910	0.909	0.880	0.937
AR coefficient inflation target	ρ_π	0.934	0.925	0.901	0.950
AR coefficient long-run growth	ρ_z	0.630	0.611	0.500	0.729
AR coefficient fixed cost	ρ_Ω	0.928	0.928	0.922	0.933
S.d. technology	$100\sigma_a$	2.333	2.460	1.929	2.985
S.d. preference	$100\sigma_b$	4.878	4.880	4.180	5.570
S.d. investment	$100\sigma_i$	2.516	2.523	2.337	2.689
S.d. monetary policy shock	$100\sigma_m$	0.561	0.572	0.494	0.653
S.d. government spending	$100\sigma_g$	2.010	2.018	1.825	2.220
S.d. inflation target	$100\sigma_\pi$	0.167	0.180	0.130	0.226
S.d. long-run growth	$100\sigma_z$	0.345	0.353	0.253	0.446
S.d. fixed cost	$100\sigma_\Omega$	9.766	9.705	9.022	10.372
ME 1-year T-Bill	$400R^{\$}_{4,t}$	0.185	0.188	0.161	0.214
ME 2-year T-Bill	$400R^{\$}_{8,t}$	0.084	0.085	0.071	0.100
ME 3-year T-Bill	$400R^{\$}_{12,t}$	0.078	0.081	0.067	0.095
ME 5-year T-Bill	$400R^{\$}_{20,t}$	0.152	0.156	0.130	0.181
ME 10-year T-Bill	$400R^{\$}_{40,t}$	0.287	0.297	0.251	0.346
ME 1Q-expected policy rate	$400E_t\left[R^f_{t,t+1}\right]$	0.456	0.464	0.408	0.522
ME 4Q-expected policy rate	$400E_t\left[R^f_{t,t+4}\right]$	0.738	0.750	0.660	0.842

Table II: Posterior statistics. Posterior means and parameter distributions are based on a standard MCMC algorithm with two chains of 100,000 parameter vector draws each, 50% of the draws used for burn-in, and a draw acceptance rates about 1/3.

We find a low intertemporal elasticity of substitution ($IES = 0.089$) and a high relative risk aversion ($RRA \approx 90$).[17] Our estimated IES is in line with, e.g., Hall (1988) or Andreasen et al. (2018), but differs from other estimates in the literature especially those in the long-run and valuation risk literatures that argue for an IES above one which is in particular not investigated in this paper. With this in mind, it is not surprising that the model needs a high relative risk aversion to fit the data. Nevertheless, our estimate is still in line with much of the existing macro-finance literature (see, for example, van Binsbergen et al., 2012, Rudebusch and Swanson, 2012, Swanson, 2019). Though a direct comparison is difficult as all of these studies use different samples, ours

[16]See Figures in the online appendix for an illustration of the posterior distribution of each parameter relative to its initial prior distribution.

[17]See online appendix for the equations of IES and RRA and, hence, for a mapping of the structural parameters in those measures.

covers the Great Moderation, and their models differ in their specification of structural shocks. As pointed out by van Binsbergen et al. (2012), models that feature a higher volatility of shocks (higher risk) thereby increasing the volatility of the stochastic discount factor need, e.g., less risk aversion to match average bond yields. This notwithstanding, our estimate of risk aversion is higher than in endowment economy studies and in micro-studies (Barsky et al., 1997). Potential explanations are present in the literature, with Malloy et al. (2009) having shown that risk aversion estimated for stockholders in the U.S. is substantially lower than a representative agent using aggregate consumption (which they find increases to 81) and Barillas et al. (2009) argue that a small amount of model uncertainty can substitute for the large degree relative risk aversion often found in the literature.

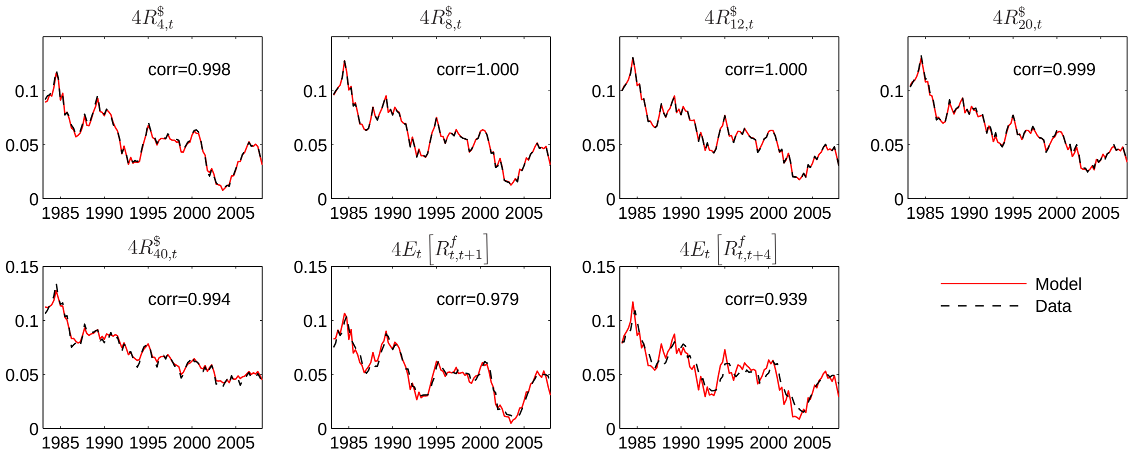
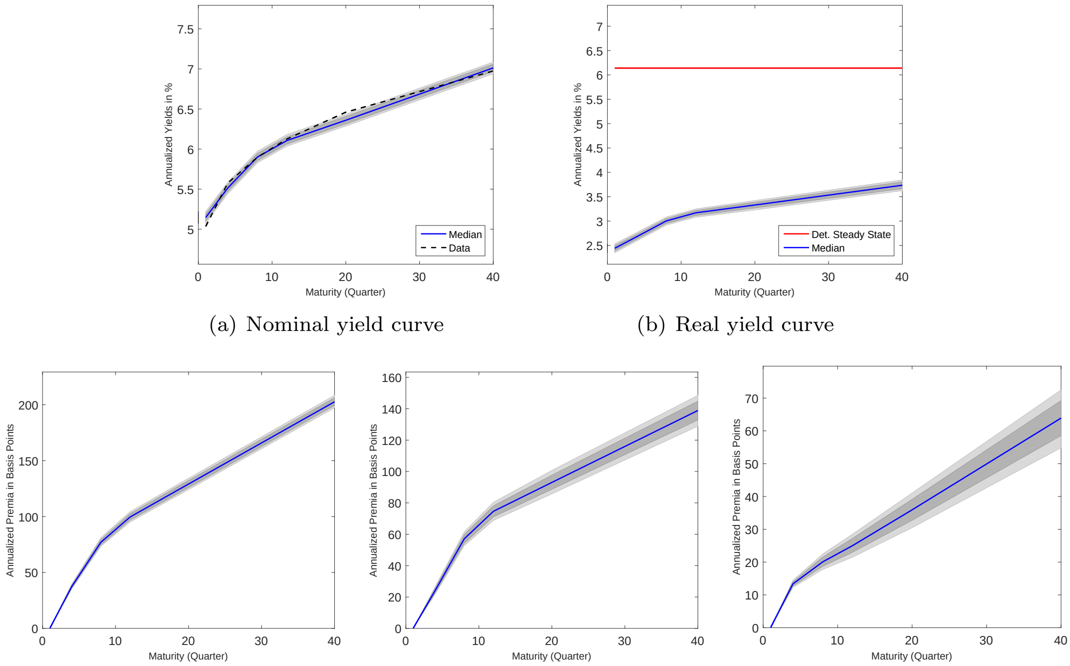


Figure 2: Observed and model implied nominal returns of treasury bills and returns of expected short rates at the posterior mode.

We estimate a quarterly deterministic steady state inflation of around 1.04% which is substantially higher than the average observed inflation rate (0.64%). Due to the non linearities in our model, the difference is related to households' precautionary motives, as also discussed by Tallarini (2000), but the approximated ergodic mean of inflation, see the subsequent subsection, is similar to the average U.S. inflation over our sample.

For the inflation target, we estimate $\rho_\pi = 0.93$ and $\zeta_\pi = 0.002$. The latter coefficient is similar to Rudebusch and Swanson (2012) while the former is smaller than their calibration, implying a less persistent effect of nominal risk in our model. This is an interesting result, because, first, a lower persistence puts the model a priori in a more difficult position to generate high and volatile nominal term premia, forcing other channels to become more important in this regard. Second, the reduction from, e.g., 0.99 to 0.93 reduces the half-life of the shock from around 69 to 10 quarters, so changes to the inflation target are expected to be significantly less long lasting. Moreover, we estimate a moderate size of investment adjustment costs ($\nu = 1.4$) and comparable estimates to the literature for price stickiness ($\gamma_p = 0.85$) and external habit formation ($b = 0.67$). Finally, we find that monetary policy puts more weight on stabilizing the inflation gap ($\eta_\pi = 3.13$) than on the output gap ($\eta_y = 0.16$) and smoothes changes in the policy rate ($\rho_R = 0.75$).

Figure 2 shows the historical time series (dash-dotted line) and the model implied smoothed time series (solid line) for the seven variables estimated with measurement error. Note that we estimate small measurement errors along the yield curve. In particular, the measurement errors range between 7 and 29 basis points, implying a correlation between the smoothed model implied yields and the data of 0.99 or higher. The measurement errors for the 1-quarter ahead and 1-year ahead expectations of the 3-month T-Bill are 45 and 74 basis points, respectively, delivering high correlations (0.94 and 0.98) of our model-based expectations with the data from the Survey of Professional Forecasters.



(a) Nominal yield curve (b) Real yield curve

(c) Nominal term premium (d) Real term premium (e) Inflation risk premium

Figure 3: Term structure of interest rates

4.2 Predicted Moments

In the following subsection, we begin our posterior analysis with respect to the predicted first and second moments. Figure 3(a) shows the predicted ergodic means of the nominal yields in relation to the means of the corresponding data. The figure illustrates the success of our estimation approach, with the a priori information about the level, slope, and curvature, based on only 3-month, 2-year, and 10-year nominal yields, sufficient to estimate first moments for all maturities.

Backus et al. (1989) and den Haan (1995) formulated the bond-pricing puzzle with the question as to why the yield curve is upward sloping. That is, long-term bonds should carry an insurance-like negative risk premium and therefore the yield curve should be downward sloping. However, the data for nominal yields as well as estimates for the nominal term premium suggest the opposite, as does our model (see figure 3(c)). The mechanism can be found in, e.g., Rudebusch and Swanson (2012): supply shocks move consumption and inflation in opposite directions, imposing a negative correlation between the two. Thus, inflation reduces the real value of nominal bonds precisely in states of low consumption when agents would particularly value higher payouts, thereby generating a positive term premium. To this end, Piazzesi and Schneider (2007) show that consumption and inflation were negatively correlated in the period 1952-2004 for the U.S., which suggests that supply shocks play a relatively important role in generating the upward sloping nominal term structure in the data and in our model.

The negative correlation between consumption growth and inflation can explain the positive slope in the nominal term structure by appealing to inflation risk, but absent another mechanism cannot account for the real term structure. If it is solely inflation risk driving the upward slope of the nominal term structure, then the real term structure should be downward sloping as spells of low consumption growth will be associated with low real rates (and hence high prices for real bonds). This gives agents a higher payout precisely when they would value it highly and implying that real bonds should carry negative, insurance-like risk premia. Nevertheless, as illustrated by figures 3(b) and 3(d), our model also predicts an upward-sloping real term structure which is in line with the literature (see, for example, Gürkaynak et al., 2010, Chernov and Mueller, 2012). The mechanism in our model follows that described in Wachter (2006) and Hördahl et al. (2008), as our households' habit formation introduces a hump-shaped response of consumption. This makes consumption growth positively autocorrelated while reducing agents' precautionary saving motive for longer maturities: households will seek to maintain their habit in the face of a slowdown in consumption, drawing down their precautionary savings and driving down real bond prices, implying that payouts on real bonds are negatively correlated with marginal utility and that real bonds demand a positive risk premium. The precautionary motive is illustrated in figure 3(b), where the red line shows the real yield curve in absence of risk, i.e., at the deterministic steady state. When confronted with risk, agents accumulate additional capital, driving down its return, consistent with Tallarini (2000) and Hansen et al. (1999). This reduction, however, is decreasing in the maturity due to the positive real risk premium, resulting in our estimated upward sloping real term structure.

Figure 3(e) shows that our model predicts an upward sloping inflation risk premium consistent with recent estimates in the literature (see, for example, Abrahams et al., 2016), with our ergodic mean term structure of inflation risk comfortably between the estimates of Buraschi and Jiltsov (2005) and Chen et al. (2010). The ergodic mean of inflation risk is approximately half the size of the real term premia for all maturities, consistent with Kim and Wright's (2005) estimates for the ten year inflation and real risk premia. Consequentially, our results suggest that most of the average slope of the nominal term structure is related to real rather than to inflation risk. Again, this finding is consistent with recent estimates for the U.S. (see, for example, Kim and Wright, 2005) and is also qualitatively comparable to the results by Hördahl and Tristani (2012) for the Euro area. So far, most of the DSGE models (see, for example, van Binsbergen et al., 2012, Swanson, 2019) generally attribute a stronger insurance-like character to real bonds that lead to flat or downward sloping real yield curves.

Name	Data		Model	
	Mean	S.d.	Mean	S.d.
GDP growth	0.540	0.573	0.540	0.790
			[0.368; 0.712]	[0.665; 0.934]
Consumption growth	.610	0.411	0.540	0.545
			[0.378; 0.702]	[0.453; 0.652]
Investment growth	0.620	2.049	0.620	2.253
			[0.297; 0.943]	[1.854; 2.720]
Annualized inflation	2.496	0.922	2.469	1.117
			[1.829; 3.108]	[0.930; 1.372]
Annualized policy rate	5.034	2.222	5.144	2.461
			[3.058; 7.235]	[1.763; 3.524]
1-year T-Bill	5.578	2.400	5.515	2.174
			[3.567; 7.464]	[1.517; 3.203]
2-year T-Bill	5.896	2.431	5.900	1.886
			[4.158; 7.642]	[1.281; 2.834]
3-year T-Bill	6.125	2.421	6.107	1.682
			[4.549; 7.665]	[1.135; 2.541]
5-year T-Bill	6.460	2.357	6.360	1.383
			[5.083; 7.641]	[0.932; 2.089]
10-year T-Bill	6.975	2.198	7.014	0.927
			[6.054; 7.974]	[0.624; 1.414]

Table III: Simulated and empirical moments of selected macro and financial variables.

Note: The simulated moments are based on 1200 parameter vector draws from the posterior. For each draw, we simulate 1000 time series for each variable of interest. After removing a burn-in of 5000 periods for each simulation the final simulated time series have the same length (T=100) as the vector of observables. The numbers in brackets indicate 5% and 95% probabilities.

Table III compares empirical and simulated first and second moments of selected variables.[18] The results illustrate that our estimation approach delivers an ergodic mean of inflation comparable to the mean of the data and, as a result, captures households' precautionary savings motives. Moreover, the simulated second moments regarding the macroeconomic variables are in line with the data, highlighting the ability of our New Keynesian DSGE model to match financial and macroeconomic moments jointly (see also Andreasen et al., 2018).[19] Regarding treasury bonds, our model misses the high volatility for longer maturities, but matches the monotonic decrease in volatility with the maturity. This result in general equilibrium models has been described in den Haan (1995) and is related to some missing source of persistence in the model (see Hördahl et al., 2008). We do not see this, however, as a fatal shortcoming of our analysis. Firstly, the uncertainty related to these moments is quite high and, secondly, it rather illustrates the tension in the competing goals the model faces: matching highly volatile nominal treasury bonds while predicting a very smooth inflation rate.

4.3 Model Implied Historical Fit

In the following subsection, we discuss our model implied historical time series for the nominal term premium, break-even inflation rate, real rate, and inflation risk premium. It is important to stress that these measures did not enter into our estimation and, instead, are produced as estimated latent variables in our analysis. To judge the quality of our estimated model, we contrast our estimates with various estimates from the literature. Following the majority of the empirical literature, we limit our discussion to 10-year maturities.

In figure 1, we compare our 10-year nominal term premium with several different prominent estimates from the literature. As Rudebusch et al. (2007) show, all of the estimated term premia, which they investigate, follow a similar pattern and are highly correlated. This is also true for our extented sample which includes two more recent estimates by Adrian et al. (2013) and Bauer (2018).[20] Table IV presents the correlations between these five measures of the term premium and the estimate of our model. Our estimate shows also a remarkably high correlation with all measures, but especially with those of Kim and Wright (2005) and Bauer (2018) (0.94 and 0.93, respectively). Given that our model is arguably closest in structure to the model used by Rudebusch and Wu (2008), we would have expected our model to display a much higher correlation with their measure than it actually does. Also, while the model by Rudebusch and Wu (2008) predicts a smooth term premium, all other models including the model presented in this paper predict a much more volatile measure.

The reason that our model produces a large and volatile term premium is similar to explanations postulated in the recent literature (see, for example, Andreasen et al., 2018). Beside the role of supply shocks in our model that generate a sizable term premium, the presence of long-run nominal risk is important in generating a volatile term premium (see Rudebusch and Swanson, 2012). Additionally, our model captures a channel recently postulated by Andreasen et al. (2018), namely the role of steady-state inflation for the mean and volatility of risk premia. In particular, steady-state inflation generates more heteroscedasticity in the stochastic discount factor which eventually produces more volatile risk premia. This channel is present despite the fact that the shocks in our model are all homoscedastic. More specifically, the endogenously generated heteroscedasticity in the pricing kernel is a byproduct of the heteroscedasticity in price dispersion due to positive steady-state inflation.

Figure 4(a) compares our 10-year real rate with the estimates provided by Gürkaynak et al. (2010) using TIPS data and those of Chernov and Mueller (2012) using survey-based forecasting data. Both measures are not fully identical with the real rate measured by our model, for example, while our real rates are based on GDP inflation the aforementioned measures are based on CPI

[18]Online appendix presents further statistics for the DSGE model.

[19]Additionally, the online appendix presents the autocorrelation of hp-filtered macroeconomic variables which also illustrates the good fit.

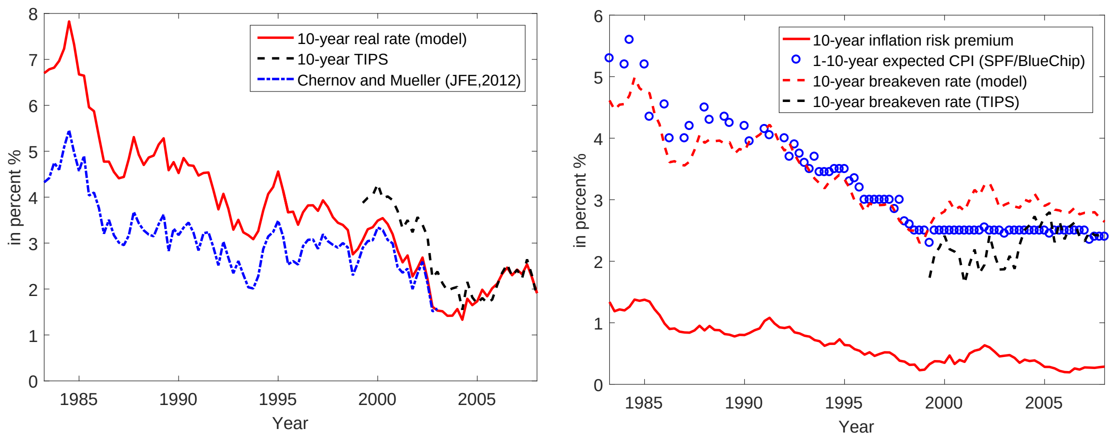
[20]The estimates by Bauer (2018) start in 1990, so all calculations using this estimate are restricted to a shorter sample.

	Bernanke et al.	Rude-busch and Wu	Kim and Wright	Adrian et al.	Bauer et al.	Bauer	S.d.
Bernanke et al. (2004)	1.000						1.294
Rudebusch and Wu (2008)	0.763	1.000					0.336
Kim and Wright (2005)	0.976	0.811	1.000				0.981
Adrian et al. (2013)	0.784	0.940	0.811	1.000			0.926
Bauer et al. (2012)	0.802	0.919	0.880	0.990	1.000		1.198
Bauer (2018)	0.853	0.734	0.936	0.859	0.865	1.000	1.182
Model	0.904	0.800	0.940	0.842	0.848	0.932	0.943

Table IV: Six measures of the 10-year term premium.

Note: Correlations among six measures of the 10-year term premium from 1984:Q1-2005:Q4. The last column presents the standard deviation over the sample. Statistics related to the estimates by Bauer (2018) are based on a shorter sample starting 1990 and those labeled Bauer et al. (2012) is their restricted OLS approach, following Hamilton and Wu (2012), adapted to our ten year term premia.

data. Also, our model has no role for a liquidity premium component that is arguably a non-negligible component of TIPS (see, for example, Abrahams et al., 2016). Nevertheless, our estimate captures the downward trend since the 1980s found likewise in Chernov and Mueller (2012). Additionally, our estimate demonstrates a high correlation with both (0.9 with Gürkaynak et al. (2010) and 0.94 with Chernov and Mueller (2012)) of these alternative measures, derived from empirical reduced-form models.



(a) 10-year real interest rate. (b) 10-year break-even rate and its components.

Figure 4: 10-year real interest rate and 10-year break-even rate.

Note: The left panel shows the model implied 10-year real rate (red solid), 10-year TIPS of Gürkaynak et al. (2010) (black dashed), and 10-year real rate of Chernov and Mueller (2012)(blue dash-dotted). The right panel shows the model implied 10-year inflation risk premium (red solid) and the 10-year break-even inflation rate (red-dashed), the 10-year break-even inflation rate of Gürkaynak et al. (2010) (black dashed), and 1 to 10-year average expected CPI inflation from SPF and BlueChip (blue circle).

Figure 4(b) shows the model implied 10-year break-even inflation rate. At the beginning of the sample, the breakeven inflation rate declines continuously until 1998. From 1999 onward we find a stable breakeven rate fluctuating around 3 percent. Over this period, our estimate is comparable in levels and pattern with those by Gürkaynak et al. (2010). Moreover, the continuous decline in the model's breakeven rate until 1998 is accompanied by a decreasing inflation risk premium. This pattern is commensurate with declining inflation expectations in this period.

In summary, our model-implied estimates of the components of 10-year bond yields demonstrate a considerable alignment with various empirical estimates in the literature. This alignment is all the more remarkable as these components of the yields were not used in our estimation procedure. Our results reiterates Swanson's (2019) conclusion that DSGE models with recursive preferences and nominal rigidities can jointly replicate observations on the macroeconomy and financial markets and extends this result to a richer and estimated model. This provides us with a high degree of confidence in our model's ability to replicate stylized term structure facts as we now turn to the structural analysis of the effects of monetary policy on the term structure of interest rates and its components.

5. MONETARY POLICY THROUGH THE LENS OF OUR MODEL

5.1 Comparison of Monetary Policy Tools

In this subsection, we analyze the effects of monetary policy shocks on term premia and distinguish between three different policy actions.[21] First, a surprise shock to the policy rate via the residual of the Taylor rule. Second, a shock to the inflation target that can be interpreted as a change in the systematic component of monetary policy (see Cogley et al., 2010) as it affects agents' perception of inflation in the long run. Third, we investigate the effects of a commitment by the monetary authority to a path for future short rates; i.e., forward guidance by means of a credible announcement to change the policy rate in the future while holding it constant until then.[22] We find that it is important to distinguish between term premia at shorter and longer maturities, with the difference driven by real risk. Monetary policy actions that convey information about the future path of the economy, as is the case in the last two actions we examine here, have substantially larger effects on term premia, consistent with Nakamura and Steinsson's (2018) Fed information effect.

The first two actions follow directly from the model, we implement the forward guidance scenario by altering the Taylor rule in eq. (13) following Laséen and Svensson (2011) and others by adding a sequence of anticipated shocks to the Taylor rule that allow the monetary authority to keep the policy rate upon announcement constant until the announced interest rate change (here a cut) is implemented as follows

$$r_t^f = R\left(r_{t-1}^f, \pi_t, y_t\right) + \sigma_m \left(\epsilon_{m,t} + \sum_{k=1}^{K} \epsilon_{m,t+k}\right), \ \epsilon_{m,t+k} \overset{iid}{\sim} N\left(0,1\right) \tag{25}$$

where $R(\cdot)$ characterizes the systematic response of monetary policy, $\epsilon_{m,t}$ is the usual contemporaneous policy shock, and $\sum_{k=1}^{K} \epsilon_{m,t+k}$ a sequence of policy shocks known to agents at time t but that affect the policy rule k periods later, i.e., at time $t+k$.

Figure 5 shows the impact responses of the nominal and real term structures.[23] The unexpected monetary policy shock (left column) shows that the response on impact of the term structure

[21] We again use the risk-adjusted linear method of Meyer-Gohde (2016). See the online appendix for a robustness exercise using generalized impulse responses in a third order perturbation as well as an investigation into the role of certainty nonequivalence in our risk-adjusted linear approximation.

[22] While this may seem a narrow aspect of recent experience with unconventional monetary policy, Woodford (2012), for example, argues that even quantitative easing itself can at least partially be interpreted as forward guidance through the signalling channel, building on results by e.g. Bauer and Rudebusch (2014). Furthermore, forward guidance has been a component of standard monetary policy at major central banks even before its explicit implementation since the financial crisis (see Gürkaynak et al., 2005a).

[23] The online appendix presents the dynamic responses of macroeconomic variables and of 1-year and 10-year maturities, respectively.

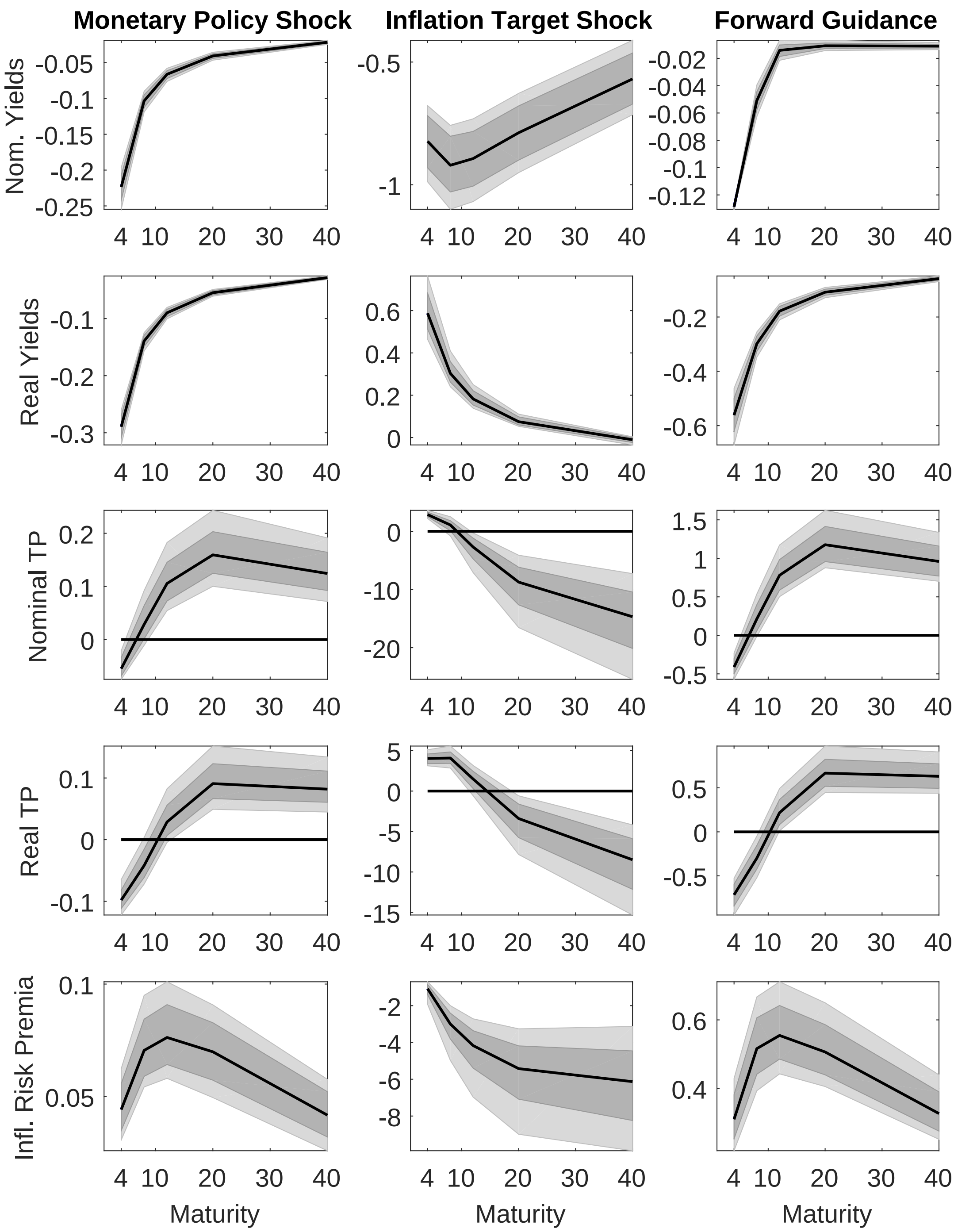


Figure 5: Impact responses of nominal and real term structures.

Note: The figure shows the impact response across all maturities to a surprise 50 basis point policy rate cut, a surprise cut in the inflation target leading to a 50 basis point policy rate cut, and forward guidance of a 50 basis point policy rate cut in 4 quarters. The deviations of yields are in percentage points while the deviations of risk premia are presented in basis points. Shaded areas represent the 90% and 68% posterior credible sets.

becomes more muted with the maturity, as would be expected in accordance with the expectations hypothesis and the path of the policy rate (assumed identical to the short rate). Similarly, the response on impact of the real yield curve, see the second row of Figure 5, is driven primarily by the expectations hypothesis and the Fisher equation with the response likewise becoming more muted with the maturity.

With the expectations hypothesis being the predominate driver of the impact on real and nominal rates, an unexpected monetary policy shock – a simple innovation to the Taylor rule – has limited, though nonzero, effects on the risk premia along all maturities. This finding is in line with those of other structural models (see, for example, Rudebusch and Swanson, 2012). On impact, see the third row of Figure 5, bond holders demand higher total premia for holding nominal bonds for longer maturities and lower total premia for shorter maturities. The effects on impact for the real term premia qualitatively mirror those of the nominal term premia, confirming that the primary driver of the nominal term premia is indeed the real economy and associated risks. On impact, the real term premia, see the fourth row of Figure 5, are shifted downward across all maturities relative to the impact response of the nominal term premia, reflecting the elevation in the inflation risk premia, see the bottom row of Figure 5, demanded by investors in response to the inflationary effects of the expansionary monetary policy. The negative initial response of real term premia associated with shorter maturities and positive response of those associated with longer maturities can be understood from the comovement of the real yields and the consumption relative to its habit in the pricing kernel.[24] Yields on real bonds at all maturities drop on impact whereas consumption relative to its habit initially rises but then falls. This generates a positive comovement between the pricing kernel and bond prices on shorter maturities that thus contain a negative, insurance-like premium. At longer maturities, this comovement becomes negative as consumption drops relative to its habit and thus real bonds of longer maturities bear a positive risk premium to induce households to hold these bonds that pay less when payoffs are more highly valued. The timing of when the ten year real term premium turns negative coincides with the onset of the contraction in the real economy. Finally, on impact, investors demand a higher premium across all maturities to compensate them for inflation risks associated with the surprise change in monetary policy.

In contrast, a surprise shock to the inflation target has a much stronger effect on the risk premia of interest rates across all maturities, see the second column of Figure 5, with the effects roughly two orders of magnitude larger. Rudebusch and Swanson (2012) found that the presence of long run nominal risk via a time-varying inflation target increases average nominal term premia by making bond holdings riskier, especially ta higher maturities. We find additionally that monetary policy's inflation target substantially affects real term premia. This is consistent with the interpretation of the shock to the inflation target as being a shift in the systematic monetary policy (see Cogley et al., 2010) by changing agents' longer run expectations of the macroeconomy. Both the inflation target and realized inflation are reduced by the more aggressive posture of monetary policy towards inflation, leading to a reduction in the inflation risk premia at all horizons on impact, see the bottom middle panel of Figure 5. While nominal term premia are still primarily driven by risks associated with the real economy in response to the inflation target shock, the effects of inflation risk premia are disproportionately increased in magnitude relative to the response to an unexpected monetary policy shock, consistent with the interpretation of this experiment being not only a change in the systematic response of monetary policy, but more specifically a more aggressive posture towards inflation.

Turning to forward guidance, the dynamic responses of interest rates are driven by the countervailing effects of the expectations hypotheses and risk premia. As in standard models under the expectations hypothesis, the dynamics of interest rates with longer maturities reflect the dynamic adjustment of the risk free short rate, determined by the monetary authority's Taylor rule. The large effects on inflation and output imply that the policy rate rises quickly above its ergodic mean only few quarters after its announced fall. This explains, at least in part, why we observe only a mild drop on impact in nominal bonds with a maturity longer than 2 years (see the upper

[24]The macroeconomic impulse responses are standard and can be found in the online appendix.

right panel of Figure 5). While the yield of a 1-quarter real bond falls by around 30 basis points on impact, the yield of a 10-year real bond falls by around 3 basis points (see the second row of the right column in Figure 5). Bondholders demand higher nominal premia on impact for all maturities from 2 years onward, in line with the empirical findings of Akkaya et al. (2015). While the real premium falls for two year real bonds on impact, this is outweighed by the larger increase in inflation risk, see the bottom two rows of the right column in Figure 5. This overall increase in risk premia prevents nominal and real long rates from falling as strongly as the expectations hypothesis would predict and therefore dampens the expansionary effects of the announced cut in the policy rate. Finally, the increase in inflation risk premia follows what theory would predict. While forward guidance does communicate the expected path of future short rate, it is just as informative about the central bank's commitment to allow higher inflation in the future. This commitment drives households' demand for higher inflation risk premia.

In sum, our findings show the importance to distinguish between different policy tools when assessing the effects of monetary policy shocks on term premia. Unexpected monetary policy shocks die out quite quickly, limiting their effects on business cycle frequencies and, consequentially, on risk premia. But news about monetary policy, which reveal information about macroeconomic variables in the future have quantitatively much stronger effects. To this end, our structural model confirms the empirical finding from Gürkaynak et al. (2005a) but highlights that monetary policy communication transmits on long-term bonds especially through risk premia. Most of the empirical literature, which investigates the effects of monetary policy shocks on term premia, focuses on samples starting in the early 1990s. At this time, the Federal Reserve had increasingly used communication as a policy tool. To this end, our findings suggest that the quantitative strong effects on term premia found by the empirical literature are primarily driven by monetary policy news about their mid- or long-term stance rather than changes in the policy rate.

5.2 Comparison with Empirical Findings

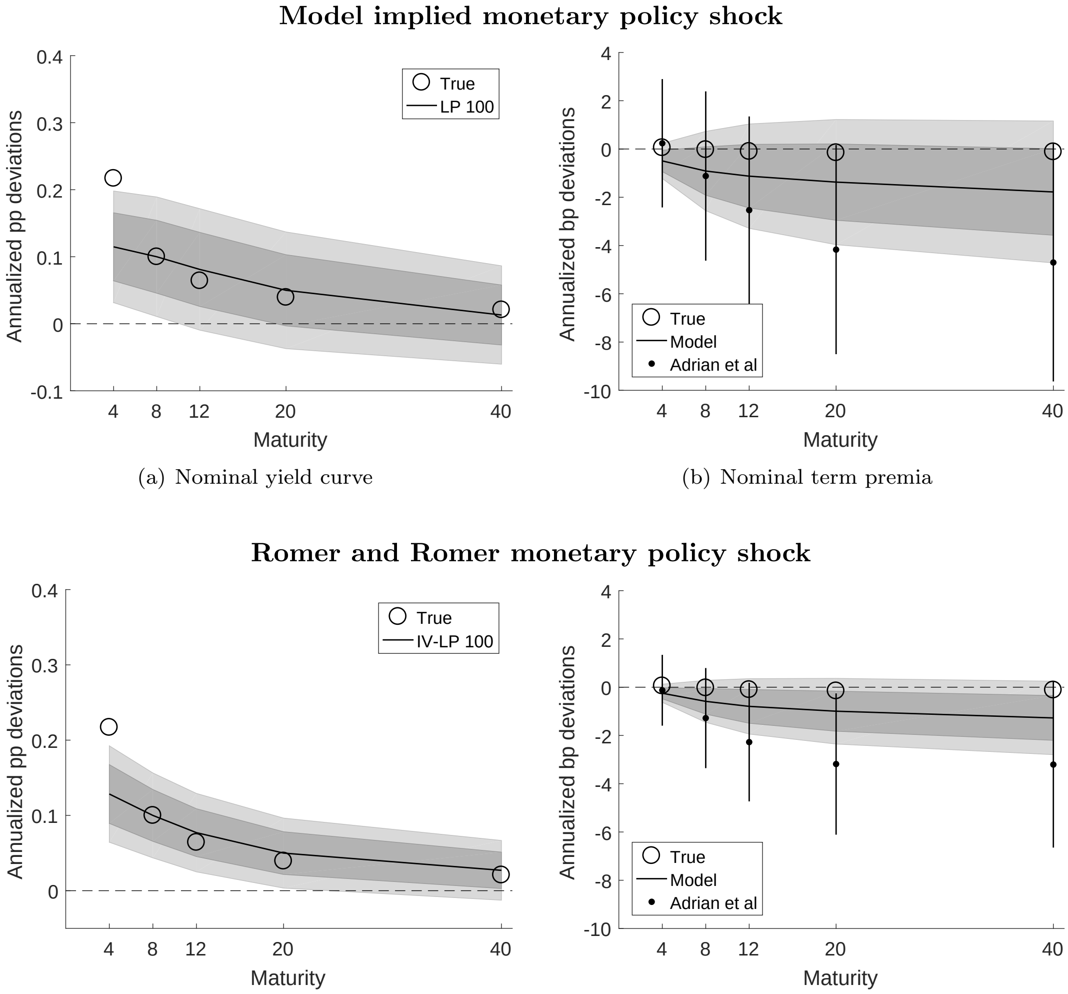
In the following subsection, we compare the findings from our structural model with those from the empirical literature, focusing on the effects of a surprise shock to the policy rate via the residual of the Taylor-rule. We run a local linear projection following Jordà (2005) by regressing the variables of interest like treasury yields and historical estimates for nominal term premia on the monetary policy shock identified by our model.[25]

We further validate our results by providing results based on an alternative measure for a monetary policy shock, the measure of Romer and Romer (2004) as updated by Wieland and Yang (2016). This measure is based on a structural interpretation of a monetary policy rule and, therefore, has a close relation to an innovation in the Taylor-rule as in our model. However, this measure is at its best a proxy for such a innovation. Accordingly, we use an instrumental variable local projection (IV-LP) as proposed by Stock and Watson (2018).

Figure 6 shows the impact effects of a surprise monetary policy shock across maturities. We scaled the median response of the 2-year treasury bond to be 0.1 annualized percentage points which ensures that the impact responses, especially at longer maturities, are comparable. Specifically, the first row of Figure 6 shows the effects of the model implied monetary policy shock while the second row shows the effects of Romer and Romer monetary policy shocks.

As can be seen in Figure 6(a), the local projection fails to recover the true impact effect at shorter maturities, underestimating the true monetary policy surprise and imposing too strong a persistence across maturities. Figure 6(b) shows the effects on nominal term premia. The impact responses from the empirical model contain the true theoretical responses, but are not significantly different from zero for all maturities (on 90% confidence level). At a 68% confidence level, the impact response is negative for longer maturities as are our theoretical responses. The IV-LP with Romer and Romer (2004) shocks as instruments gives qualitatively similar but quantitatively smaller results for the impact response of nominal term premia. This finding also holds when using different estimates for the nominal term premium (those of Adrian et al. (2013) are also depicted

[25]The online appendix provides additional details of the empirical model along with further results.



(a) Nominal yield curve (b) Nominal term premia

(c) Nominal yield curve (d) Nominal term premium

Figure 6: Impact effect of monetary policy shock on nominal yields and nominal term premia for different maturities.

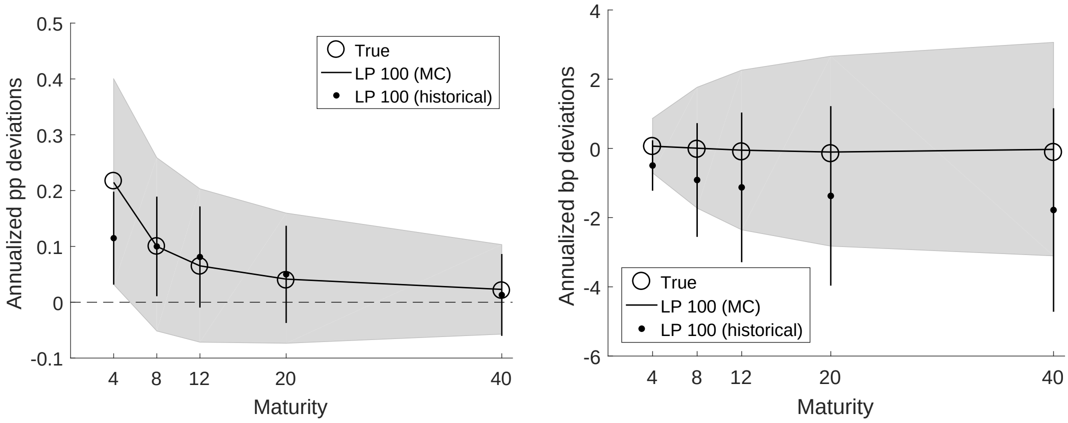
Note: The solid line and shaded areas show median response, the 68%, and 90% confidence bands from the local projection with the model implied historical term premia as dependent variable, respectively. The circles indicate the theoretical, true response. Additionally, the dots and vertical lines in the right panel show median response and 90% confidence bands from the local projection with term premia estimates from Adrian et al. (2013). We use the Newey-West correction for the standard errors.

in Figure 6(b) and 6(d)).[26]

All the local projection estimates deliver qualitatively and quantitatively similar results: A tightening of monetary policy rates reduces nominal term premia, especially for longer maturities, in line with the empirical work by Nakamura and Steinsson (2018) and Crump et al. (2016) but in contrast to, e.g., Gertler and Karadi (2015).

However, we also find that local projections using our model's smoothed series predict larger effects than the true, theoretical responses. For further investigation, we perform a Monte-Carlo exercise with simulated time series for nominal yields, monetary policy shocks, and nominal term

[26]In the online appendix, we show that this also holds for most of the empirical estimates of the 10-year nominal term premia in the literature.



(a) Nominal yield curve (b) Nominal term premium

Figure 7: Monte-Carlo exercise

Note: Impact effect of monetary policy shock on nominal yields and nominal term premia for different maturities. The solid line and shaded areas show median response and 90% confidence bands from the local projection with sample length 100 based on 1000 simulations, respectively. The circles indicate the theoretical, true response. Additionally, the dots and vertical lines show median response and 90% confidence bands from the local projection with historical model implied term premia.

premia from our model at the posterior mean.[27] Figure 7 shows that the empirical model on average replicates the true response of the DGP and, therefore, shows no systematic small sample bias (see discussion in Jordà, 2005). Nevertheless, the estimation uncertainty is especially pronounced in small samples, capturing a wide range of quantitatively and qualitatively different estimates in the empirical literature.

6. CONCLUSION

This paper provides a medium scale macro-finance model, estimated with full information on U.S. macroeconomic and Treasury bond data. We estimate historical time series of term premia that match those found in reduced form empirical estimates without sacrificing the macroeconomic fit or other financial variables. We therefore provide a structural framework for the analysis of endogenous, time-varying term premia.

Distinguishing between different monetary policy actions is important. While unexpected shocks to the policy rate have quantitatively small effects, shocks revealing information about the future of monetary policy (e.g. forward guidance) can have quantitatively much stronger effects. Hence, with this disentangling of shocks, an ongoing challenge for empirical models, our findings can provide insight on some of the seemingly contradictory findings in this literature (see, for example among others, Hanson and Stein, 2015, Nakamura and Steinsson, 2018, Gertler and Karadi, 2015).

We offer a first step toward understanding the transmission of monetary policy on the term structure of interest rates from a structural Bayesian perspective, but many salient questions need further investigation. For example, while our model features a frictionless asset trade, a model featuring market segmentation could affect the policy conclusions of our paper (see, for example, Fuerst, 2015). Additionally, a further extension would be the incorporation of the zero lower bound for interest rates, which remains a not fully resolved methodological challenge for nonlinear DSGE models as well affine term structure models. Moreover, investigating the impact of unconventional

[27]We run 1000 simulations with a sample length of 100 after having discarded the first 5000 observations. Afterwards, we run the same local projections as before.

monetary policy on risk premia or the impact of monetary policy on asset valuation more generally are natural questions of currently high interest. We acknowledge but leave these extensions for future work, providing an estimated macro-finance model in this paper able to provide a structural analysis of the impact of monetary policy on the term structure of interest rates.

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