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MONETARY POLICY DRIVERS OF BOND AND EQUITY RISKS

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ABSTRACT

How do monetary policy rules, monetary policy uncertainty, and macroeconomic shocks affect the risk properties of US Treasury bonds? The exposure of US Treasury bonds to the stock market has moved considerably over time. While it was slightly positive on average over the period 1960-2011, it was unusually high in the 1980s, and negative in the 2000s, a period during which Treasury bonds enabled investors to hedge macroeconomic risks. This paper develops a New Keynesian macroeconomic model with habit formation preferences that prices both bonds and stocks. The model attributes the increase in bond risks in the 1980s to a shift towards strongly anti-inflationary monetary policy, while the decrease in bond risks after 2000 is attributed to a renewed focus on output fluctuations, and a shift from transitory to persistent monetary policy shocks. Endogenous responses of bond risk premia amplify these effects of monetary policy on bond risks.

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1 Introduction

In different periods of history, long-term US Treasury bonds have played very different roles in investors' portfolios. During the Great Depression of the 1930s, and once again in the first decade of the 21st Century, Treasury bonds served to hedge other risks that investors were exposed to: the risk of a stock market decline, and more generally the risk of a weak macroeconomy, with low output and high unemployment. Treasuries performed well both in the Great Depression and in the two recessions of the early and late 2000s. During the 1970s and especially the 1980s, however, Treasury bonds added to investors' macroeconomic risk exposure by moving in the same direction as the stock market and the macroeconomy. A number of recent papers including Baele, Bekaert, and Inghelbrecht (2010), Campbell, Sunderam, and Viceira (2013), Christiansen and Ranaldo (2007), David and Veronesi (2013), Guidolin and Timmermann (2006), and Viceira (2012) have documented these developments. These stylized facts raise the question what macroeconomic forces determine the risk properties of US Treasury bonds, and particularly their changes over time.

The contribution of this paper is twofold. First, we develop a model combining a standard New Keynesian macroeconomy with habit formation preferences. Bonds and stocks in the model can be priced from assumptions about their payoffs. Second, we use the model to relate changes in bond risks to periodic regime changes in the parameters of the central bank's monetary policy rule and the volatilities of macroeconomic shocks, including a regime shift that we identify in the early 2000s. Since monetary policy in our model affects macroeconomic and risk premium dynamics, we capture both the direct and indirect effects of monetary policy changes.

Macroeconomic dynamics in our model follow a standard three-equation New Keynesian model. An investment-saving curve (IS) describes real equilibrium in the goods market based on the Euler equation of a representative consumer, a Phillips curve (PC) describes the effects of nominal frictions on inflation, and a monetary policy reaction function (MP) embodies a Taylor rule as in Clarida, Gali, and Gertler (1999), Taylor (1993), and Woodford (2001). A time-varying inflation target in the monetary policy rule captures investors' long-term perceived policy target and volatility in long-term bond yields.

While we model macroeconomic dynamics as loglinear, preferences in our model are nonlinear to capture time-varying risk premia in bonds and stocks. As in Campbell and Cochrane (1999), habit formation preferences generate highly volatile equity returns and address the "equity volatility puzzle" one of the leading puzzles in consumption-based asset pricing (Campbell, 2003). In contrast to Campbell and Cochrane (1999), our preferences are consistent with an exactly loglinear consumption Euler equation, time-varying conditionally homoskedastic real interest rates, and endogenous macroeconomic dynamics. The model is overall successful at matching the comovement of bond and stock returns, while generating time-varying bond and equity risk premia.

By using a New Keynesian macroeconomic framework, in which price stickiness allows monetary policy to have real effects, we overcome some of the limitations of affine term structure and real business cycle approaches. One common approach to studying macroeconomic bond risks is to use identities that link bond returns to movements in bond yields, and that link nominal bond yields to expectations of future short-term real interest rates, expectations of future inflation rates, and time-varying risk premia on longer-term bonds over short-term bonds. Barsky (1989), Shiller and Beltratti (1992), and Campbell and Ammer (1993) were early examples of this approach. A more recent literature has proceeded in a similar spirit, building on the no-arbitrage restrictions of affine term structure models

(Duffie and Kan 1996, Dai and Singleton 2000, 2002, Duffee 2002) to estimate multifactor term structure models with both macroeconomic and latent factors (Ang and Piazzesi 2003, Ang, Dong, and Piazzesi 2007, Rudebusch and Wu 2007). Although these exercises can be informative, they are based on a reduced-form econometric representation of the stochastic discount factor and the process driving inflation. This limits the insights they can deliver about the underlying macroeconomic determinants of bond risks.

A more ambitious approach is to build a general equilibrium model of bond pricing. Real business cycle models have an exogenous real economy, driven by shocks to either goods endowments or production, and an inflation process that is either exogenous or driven by monetary policy reactions to the real economy. Papers in the real business cycle tradition often assume a representative agent with Epstein-Zin preferences, and generate time-varying bond risk premia from stochastic volatility in the real economy and/or the inflation process (Song 2014, Bansal and Shaliastovich 2013, Buraschi and Jiltsov 2005, Burkhardt and Hasseltoft 2012, Gallmeyer et al 2007, Piazzesi and Schneider 2006). Some papers instead derive time-varying risk premia from habit formation in preferences, with or without stochastic macroeconomic volatility (Ermolov 2015, Bekaert, Engstrom, and Grenadier 2010, Bekaert, Engstrom, and Xing 2009, Buraschi and Jiltsov 2007, Dew-Becker 2013, Wachter 2006). Under either set of assumptions, this work allows only a more limited role for monetary policy, which determines inflation (at least in the long run) but has no influence on the real economy.[2]

Our model is most closely related to a recent literature exploring the asset pricing implications of New Keynesian models. Recent papers in this literature include Andreasen (2012), Bekaert, Cho, and Moreno (2010), Van Binsbergen et al (2012), Dew-Becker (2014), Kung

[2]A qualification to this statement is that in some models, such as Buraschi and Jiltsov (2005), a nominal tax system allows monetary policy to affect fiscal policy and, through this indirect channel, the real economy.

(2015), Li and Palomino (2014), Palomino (2012), Rudebusch and Wu (2008), and Rudebusch and Swanson (2012). While this literature has begun to focus on the term structure of interest rates, an integrated treatment of bonds and stocks, especially with risk premia not driven by counterfactually extreme heteroskedasticity in macroeconomic fundamentals, has so far proved elusive.[3]

We use our model to quantitatively investigate two candidate explanations for the empirical instability in bonds' risk properties: changes in monetary policy or changes in macroeconomic shocks. In this way we contribute to the literature on monetary policy regime shifts (Andreasen 2012, Ang, Boivin, Dong, and Kung 2011, Bikbov and Chernov 2013, Boivin and Giannoni 2006, Chib, Kang, and Ramamurthy 2010, Clarida, Gali, and Gertler 1999, Palomino 2012, Rudebusch and Wu 2007, Smith and Taylor 2009). While this literature has begun to focus on the implications of monetary regime shifts for the term structure of interest rates, previous papers have not looked at the implications for the comovements of bonds and equities as we do here. Our structural analysis takes account of various channels by which the monetary policy regime affects the sensitivities of bond and stock returns to macroeconomic shocks, including endogenous responses of risk premia.[4]

Both US monetary policy and the magnitude of macroeconomic shocks have changed substantially over our sample from 1960 to 2011. Testing for break dates in the relation between the Federal Funds rate, output, and inflation, we determine three distinct monetary policy regimes. The first regime comprises the period of rising inflation in the 1960s and

[3]In contrast, several papers have used reduced-form affine or real business cycle models to provide an integrated treatment of bonds and stocks (Ang and Ulrich 2012, Bansal and Shaliastovich 2013, Bekaert, Engstrom, and Grenadier 2010, Ermolov 2015, Koijen, Lustig, and Van Nieuwerburgh 2010, Campbell 1986, Campbell, Sunderam, and Viceira 2013, d'Addona and Kind 2006, Dew-Becker 2013, Eraker 2008, Hasseltoft 2009, Lettau and Wachter 2011, Wachter 2006).

[4]Song (2014), in a paper circulated after the first version of this paper, considers bond-stock comovements in a model with exogenous real dynamics.

1970s, while the second one covers the inflation-fighting period under Federal Reserve Board chairmen Paul Volcker and Alan Greenspan. The third regime is characterized by renewed attention to output stabilization and smaller, but more persistent, shocks to the monetary policy rule. If central bank policy affects the macroeconomy through nominal interest rates, it is natural to think that these significant changes in monetary policy should change the risks of bonds and stocks.

The nature of economic shocks has also changed over time, with potentially important implications for bond risks. While oil supply shocks were prominent during the 1970s and early 1980s, they became less important during the subsequent Great Moderation. In our model, Phillips curve shocks act as supply shocks, leading to high inflation recessions. Nominal bond prices fall with rising inflation expectations, while stock prices fall with recessions, so Phillips curve shocks move bonds and stocks in the same direction and give rise to positive nominal bond betas. In contrast, credible shocks to the central bank's long-term inflation target, or equivalently persistent shocks to the monetary policy rule, reduce the beta of nominal bonds. A downward drift in the perceived target drives down inflation and induces firms with nominal rigidities to reduce output. Consequently, a negative target shock raises the value of nominal bonds just as equity prices fall, decreasing the stock-market beta of nominal bonds.

Figure 1 shows a timeline of changing US bond risks, together with estimated monetary policy regimes, and oil price shocks from Hamilton (2009). We estimate monetary policy break dates using data on the Federal Funds rate, the output gap (the gap between real output and potential output under flexible prices), and inflation.[5] The statistically determined dates 1977Q2 and 2001Q1 line up remarkably closely with changes in bond betas and institu-

[5]For details on the econometric procedure, see Section 3.

tional and personal changes at the Federal Reserve, even if a reading of US Federal Reserve history might suggest a slightly later first break date. The 1977Q2 break date precedes by two years Paul Volcker's appointment as Federal Reserve chairman, a change that ushered in a new era of inflation fighting. The second break date corresponds to the end of the great economic expansion in the 1990s and the start of the Federal Reserve's accommodative response to the end of the technology boom and the attacks of 9/11.

The 5-year nominal bond CAPM betas and return volatilities plotted in Figure 1 illustrate important changes in the risks of nominal bonds over time.[6] Moreover, changes in nominal bond risks broadly line up with changes in monetary policy regimes. The nominal bond beta, shown in Panel A, was positive but close to zero before 1977, strongly positive thereafter, and turned negative after the year 2000. Bond return volatility, shown in Panel B, also increased during the middle subperiod, although there is higher-frequency variation as well, most notably a short-lived spike in the early 1980s. In contrast, oil price shocks are concentrated in the first two of our subperiods and there is no visually apparent relation between oil price shocks and nominal bond betas. The main empirical analysis in this paper systematically examines the role of time-varying shock volatilities and finds that they interact with monetary policy in important ways to jointly determine the risks of nominal bonds.

The organization of the paper is as follows. Section 2 describes the model for macroeconomic dynamics and preferences. This section also derives the New Keynesian IS curve from habit preferences.

Section 3 describes our data sources and presents summary statistics for our full sample period, 1954Q3 through 2011Q4, and for three subperiods, 1960Q2–1977Q1, 1977Q2–2000Q4, and 2001Q1–2011Q4. For each subperiod, this section also estimates a reduced-form

[6] We show filtered CAPM betas and standard deviations of daily returns on a benchmark 5-year nominal bond over a rolling 3-month window, together with 95% confidence intervals.

monetary policy rule and backs out naïve estimates of Taylor rule parameters. These naïve monetary policy parameter estimates, unlike estimates from our full model, do not account for regression bias caused by endogeneity of macroeconomic variables and time-variation in the central bank's inflation target.

Section 4 calibrates our model to fit both macroeconomic and asset pricing data over our three subperiods. The model fits bond-stock comovements and empirical Taylor-rule type regressions for each subperiod, while generating plausible consumption growth volatility and equity and bond risk premia. Section 5 presents counterfactual analysis, asking how bond risks would have evolved over time if the monetary policy rule, or the volatilities of macroeconomic shocks, had been stable instead of time-varying. Section 6 concludes, and an online appendix (Campbell, Pflueger, and Viceira 2015) presents additional details.

2 A New Keynesian Asset Pricing Model

Our model integrates a standard three-equation loglinear New Keynesian macroeconomic model with a habit-formation model of asset prices. While many variants of the basic New Keynesian model have been proposed, we use a small-scale New Keynesian framework to study the key implications of the broader class of New Keynesian models for asset prices.[7] We abstract from nonlinearities in macroeconomic dynamics in order to focus on nonlinearities in asset prices, where they are most salient. On the asset pricing side, we build on the habit model of Campbell and Cochrane (1999). The stochastic discount factor (SDF) links asset returns and macroeconomic and monetary variables in equilibrium.

The Euler equation is a standard New Keynesian building block and provides an equiv-

[7] Woodford (2003) argues that the three-equation log-linearized New Keynesian model captures many of the key dynamics of more complicated models, including models with investment.

alent of the investment-savings (IS) curve. We derive a habit-founded Euler equation in terms of the current, lagged, and expected output gaps and the short-term real interest rate. Euler equations with both backward-looking and forward-looking components are common in the dynamic stochastic general equilibrium (DSGE) literature (Christiano, Eichenbaum, and Evans 2005, Boivin and Giannoni 2006, Smets and Wouters 2007).[8] The backward-looking component is important for obtaining a unique equilibrium (Cochrane 2011) and for capturing the empirical output response to monetary policy shocks (Fuhrer 2000). The forward-looking component follows from standard household dynamic optimization.

The second building block of a New Keynesian model is the Phillips curve (PC) equation that links inflation and real output in equilibrium. We directly assume a PC with both forward- and backward-looking components, as may arise from different microfoundations for nominal rigidities. In the appendix, we derive a PC for firms that face Calvo (1983) price-setting frictions and partial indexing as in Smets and Wouters (2003), and maximize future expected profits discounted with our SDF. Alternative microfoundations, such as infrequent information updating, may yield variants of this benchmark PC (Mankiw and Reis 2002).

The third building block of the model is an equation describing the behavior of the central bank. We assume that the central bank's policy instrument is the short-term nominal interest rate. The central bank sets this interest rate according to a Taylor (1993) monetary policy (MP) rule, as a linear function of the "inflation gap" (the deviation of inflation from the central bank's target), the output gap, and the lagged nominal interest rate. Empiri-

[8]Christiano, Eichenbaum, and Evans (2005) and Boivin and Giannoni (2006) derive a backward- and forward-looking linearized Euler equation in a model where utility depends on the difference between consumption and an internal habit stock. A backward-looking component in the Euler equation can also be derived in a model with multiplicative external habit (Abel 1990, Fuhrer 2000). Our model differs from these previous works in that difference habit in our model gives rise to time-varying bond and equity risk premia. Rudebusch and Swanson (2008) allow for time-varying risk premia in a production-based model. However, their focus is on the endogenous labor response to habits and they use perturbation solution methods, which we have found not to work well for our model.

cally, the Fed appears to smooth interest rates over time, and we capture this by modeling the nominal short rate as adjusting gradually to the target rate. This approach is fairly standard in the New Keynesian literature, although there is some debate over the relative importance of partial adjustment and serially correlated unobserved fundamentals in the MP rule (Rudebusch 2002, Coibion and Gorodnichenko 2012).

While our preferences build on Campbell and Cochrane (1999) and Wachter (2006), they differ in that surplus consumption—or consumption relative to habit—can depend on the current and lagged output gaps. As a result, the consumption Euler equation takes the form of an exactly loglinear New Keynesian Euler equation depending on current, future, and lagged output gaps as in Clarida, Gali, and Gertler (1999, CGG). Risk premia increase when surplus consumption and the output gap are low, consistent with the empirical evidence on stock and bond return predictability (Chen 1991, Cochrane 2007, Cochrane and Piazzesi 2005, Fama 1990, Fama and French 1989, Lamont 1998, Lettau and Ludvigson 2001).

The New Keynesian macroeconomic model provides equilibrium dynamics for the output gap, inflation, and the policy rate, but preferences are over consumption. We bridge this gap between the macroeconomic and asset pricing sides of the model by assuming that consumption and the output gap are driven by the same shock. We model the output gap as the difference between current period consumption and an exponentially-weighted moving average of lagged consumption. This specification generates near random-walk dynamics for consumption, similar to the endowment consumption dynamics in Campbell and Cochrane (1999), while preserving stationarity for the output gap.

Figure 2, Panel A supports this description of the joint dynamics of consumption and the output gap. The figure plots the time series of stochastically detrended consumption—log real consumption of nondurables and services less an exponentially-weighted moving average with a half life of 2.6 years—and the log output gap. The two series move very closely together—almost surprisingly so given the measurement issues in both series—with a correlation of 90%.

We allow for shocks to the central bank's long-run policy rate. These shocks can temporarily boost demand in our model. We interpret them broadly as capturing persistent shocks to expected policy rates, or explicit and perceived changes in the long-run inflation target. As such, movements in the policy target capture changes in forward-looking public expectations of central bank behavior, that are accompanied by almost no movement in the Fed Funds rate. The central bank's ability to steer output and inflation through target expectations may vary with central bank credibility (Orphanides and Williams 2004).

We model the policy target rate as a unit root process, consistent with the extremely high persistence in US inflation data (Ball and Cecchetti 1990, Stock and Watson 2007). We choose a unit root specification rather than a highly persistent mean-reverting inflation target for several reasons. First, the inflation target reflects consumers' long-run target expectations, whose changes cannot be anticipated. A mean-reverting inflation target would imply counterintuitive predictability of target changes. Second, a highly persistent inflation target may lead to equilibrium existence and uniqueness issues, while we can factor out a unit root inflation target from equilibrium dynamics. Finally, while our unit root assumption means that the unconditional variance of nominal interest rates is undefined, we believe that this difficulty is driven by high empirical persistence in nominal bond yields rather than our specific modeling assumption. Even with a highly persistent inflation target this unconditional variance would be extremely sensitive to an imprecisely identified persistence parameter.

Following CGG, we assume that transitions from one regime to another are structural breaks, completely unanticipated by investors. We show in the appendix that model implications are unchanged if we include a small, constant regime-switching probability. The qualitative and quantitative implications from a model with unanticipated regime changes survive for two reasons. First, quarterly transition probabilities have to be small to match average empirical regime durations of ten to 25 years. Second, our regimes differ in an important dimension from the model of David and Veronesi (2013), where learning about regimes has important effects on bond and equity risks. David and Veronesi's regimes are characterized by exogenously given first moments for consumption growth and inflation. In contrast, our regimes have identical long-run consumption growth and inflation distributions, but differ endogenously in the co-movements of output, inflation, and interest rates. Thus our approach can be regarded as complementary to David and Veronesi.

2.1 Macroeconomic dynamics

We model business cycle and inflation dynamics using a standard log-linearized three equation New Keynesian model (CGG):

$$
\begin{aligned}
x_t &= \rho^{x-} x_{t-1} + \rho^{x+} \mathrm{E}_t x_{t+1} - \psi(\mathrm{E}_t i_t - \mathrm{E}_t \pi_{t+1}), && (1)\\
\pi_t &= \rho^{\pi} \pi_{t-1} + (1-\rho^{\pi}) \mathrm{E}_t \pi_{t+1} + \kappa x_t + u_t^{PC}, && (2)\\
i_t &= \rho^{i} i_{t-1} + (1-\rho^{i}) \left[\gamma^{x} x_t + \gamma^{\pi} \left(\pi_t - \pi_t^*\right) + \pi_t^*\right] + u_t^{MP}, && (3)\\
\pi_t^* &= \pi_{t-1}^* + u_t^*. && (4)
\end{aligned}
$$

We denote the log output gap—the deviation of real output from flexible price equilibrium—by x_t and log inflation by π_t. We write π_t^* for the inflation target. We write i_t to denote the log yield at time t—and return at time $t+1$—on a one-period nominal T-bill. Similarly, r_t

denotes the log yield on a one-period real Treasury bill. We use the subscript t for short-term nominal and real interest rates to emphasize that they are known at time t.

We do not include an IS shock in the Euler equation (1) because we require the IS equation to be consistent with the consumption Euler equation.[9] The New Keynesian PC (2) has parameters ρ^{π}, determining the relative weight on past inflation and expected future inflation, and κ governing the sensitivity of inflation to the output gap.

Equations (3) and (4) describe monetary policy. They determine the short-term nominal interest rate with parameters ρ^{i} controlling the influence of past interest rates on current interest rates, γ^{x} governing the reaction of the interest rate to the output gap, and γ^{π} governing the response of the interest rate to inflation relative to its target level π_t^*. Equation (4) specifies that the central bank's policy target follows a random walk.

Monetary policy in our model does not react directly to long-term nominal bond yields or stock prices, but only to macroeconomic determinants of these asset prices. However, a persistent policy target shifts the term structure similarly to a level factor. In that sense, our model is similar to models where the level factor of the nominal term structure directly enters the central bank's monetary policy function (Rudebusch and Wu 2007, 2008).

Finally, we assume that the vector of shocks is independently and conditionally normally distributed with mean zero and diagonal variance-covariance matrix:

$$u_t = [u_t^{PC}, u_t^{MP}, u_t^*]', \qquad \mathrm{E}_{t-1}[u_t u_t'] = \Sigma_u = \begin{bmatrix} (\sigma^{PC})^2 & 0 & 0 \\ 0 & (\sigma^{MP})^2 & 0 \\ 0 & 0 & (\sigma^*)^2 \end{bmatrix}. \tag{5}$$

[9] We found that modifying preferences to allow for an IS shock has little effect on the pricing of consumption claims.

Equation (5) has two important properties. First, the variances of all shocks in the model are conditionally homoskedastic. The previous version of this paper generated countercyclical risk premia from countercyclical volatility of shocks. While time-varying volatilities may provide a convenient tool for generating time-varying risk premia, strong countercyclical heteroskedasticity is not a feature of macroeconomic data.[10] The habit formation preferences in this model reconcile conditionally homoskedastic macroeconomic fundamentals with empirically plausible time-variation in bond and equity risk premia. Second, the assumption that monetary policy shocks u_t^{MP} and u_t^* are uncorrelated with PC shocks captures the notion that all systematic variation in the short-term nominal interest rate is reflected in the monetary policy rule.

2.2 Consumption and Preferences

Consider a habit formation model of the sort proposed by Campbell and Cochrane (1999), where utility is a power function of the difference between consumption C and habit H:

$$U_t = \frac{(C_t - H_t)^{1-\gamma} - 1}{1-\gamma} = \frac{(S_t C_t)^{1-\gamma} - 1}{1-\gamma}. \tag{6}$$

Here $S_t = (C_t - H_t)/C_t$ is the surplus consumption ratio and γ is a curvature parameter that controls risk aversion. Relative risk aversion varies over time as an inverse function of the surplus consumption ratio: $-U_{CC}C/U_C = \gamma/S_t$.

[10] We thank our discussants Jules van Binsbergen, Martin Lettau, and Monika Piazzesi for emphasizing this point. Ermolov (2015), in a paper circulated after the previous version of this paper, also uses changing fundamental volatility to generate changing risk premia.

Marginal utility in this model is

$$U_t' = (C_t - H_t)^{-\gamma} = (S_t C_t)^{-\gamma}, \tag{7}$$

and log marginal utility is given by $\ln U_t' = -\gamma(s_t + c_t)$.

2.2.1 Modeling Consumption

We model consumption in terms of the output gap, so inflation and the Federal Funds rate are relevant for consumption only to the extent that they are correlated with the output gap. In small-scale New Keynesian models, it is common to model consumption as equal to the output gap (CGG). However, the output gap is stationary while empirical consumption appears to have a unit root, with potentially important asset pricing implications. For constants τ and g, we model consumption as follows:

$$c_t = gt + \tau\left(x_t + (1-\phi)[x_{t-1} + x_{t-2} + ...]\right). \tag{8}$$

For any stationary output gap process, the relation (8) defines a consumption process with a unit root. As a leading example, if the output gap follows an AR(1) process with first-order autocorrelation parameter ϕ, (8) implies that consumption follows a random walk with drift. The parameter g regulates average consumption growth and τ regulates the relative volatility of consumption and the output gap.

Ignoring constants, we can equivalently rewrite the output gap in terms of consumption in excess of an exponentially decaying stochastic trend:

$$x_t \;\; = \;\; \tau^{-1}\left(c_t - (1-\phi)[c_{t-1} + \phi c_{t-2} + ...]\right). \tag{9}$$

While our model of consumption and output is reduced form, it captures a salient feature of the data. Both the left-hand and right-hand-sides in (9) are stationary, so it makes sense to consider the correlation in the data. Figure 2, Panel A shows that stochastically detrended real consumption and the log output gap (from the Congressional Budget Office) are 90% correlated. We use an annualized smoothing parameter of $\phi = 0.94$, corresponding to a half-life of 2.6 years. Importantly, Figure 2 suggests a stable consumption-output gap relation across monetary policy and shock regimes.

2.2.2 Modeling the Surplus Consumption Ratio

We specify the dynamics for log surplus consumption to satisfy the following two features. First, we require log surplus consumption to be stationary, as in Campbell and Cochrane (1999). Second, the loglinear Euler equation (1) is exact for our choice of preferences.

Denoting the demeaned log output gap by x_t, unexpected consumption innovations by $\varepsilon_{c,t+1}$, and the steady state log surplus consumption ratio by $\bar{s}$, we assume the following dynamics for the log surplus consumption ratio:

$$s_{t+1} = (1-\theta_0)\bar{s} + \theta_0 s_t + \theta_1 x_t + \theta_2 x_{t-1} + \lambda(s_t)\varepsilon_{c,t+1}, \tag{10}$$

$$\varepsilon_{c,t+1} = c_{t+1} - \mathrm{E}_t c_{t+1} = \tau\left(x_{t+1} - \mathrm{E}_t x_{t+1}\right). \tag{11}$$

We can use (9) to substitute out the output gap from (10) and re-write surplus consumption dynamics in terms of current and lagged consumption (ignoring constants):

$$\begin{aligned} s_{t+1} = {} & \theta_0 s_t + \theta_1 \tau^{-1}\left(c_t - (1-\phi)\left[c_{t-1} + \phi c_{t-2} + ...\right]\right) \\ & + \theta_2 \tau^{-1}\left(c_{t-1} - (1-\phi)\left[c_{t-2} + \phi c_{t-3} + ...\right]\right) + \lambda(s_t)\varepsilon_{c,t+1}. \end{aligned} \tag{12}$$

When $\theta_1 = \theta_2 = 0$, the surplus consumption dynamics (10) are the same as in Campbell and Cochrane (1999). In the appendix, we show that log habit can be approximated as a distributed lag of log consumption, with weights on distant lags converging to those of Campbell and Cochrane (1999). This distributed lag expression shows that θ_1 and θ_2 free up how strongly log habit loads onto the first two lags of consumption, relative to Campbell and Cochrane (1999). Increases in θ_2 and θ_2 reduce log habit loadings onto the most recent lags of consumption, while increasing the loadings on medium-term lags. In the calibrated model, we find that simulated log surplus consumption is closely related to the log output gap. A regression of the simulated log surplus consumption ratio onto the output gap yields a slope coefficient of around 12 and an R-squared of 63%, weighted across calibration periods.[11]

The specification (10) is similar in spirit to Wachter (2006). However, our approach has several advantages in a model that endogenously derives macroeconomic dynamics. While short-term real interest rates in Wachter (2006) and Menzly, Santos, and Veronesi (2004) depend on the surplus consumption ratio and are therefore heteroskedastic, the short-term real rate in our model depends on current, lagged, and future values of the output gap. This allows us to obtain equilibrium dynamics for the output gap and consumption that are conditionally homoskedastic—an assumption in Campbell and Cochrane (1999), Menzly, Santos, and Veronesi (2004), and Wachter (2006). In addition, the dynamics (10) allow us to derive exactly the loglinear New Keynesian Euler equation (1) from the consumption Euler equation.

[11]If θ_1 and θ_2 are different from zero, there is the theoretical possibility that the log surplus consumption ratio exceeds the maximal value s_{max}, where the sensitivity function is non-zero. However, the probability of this event is very small in our calibrated model (less than 0.1% per quarter). In this respect our model is similar to Campbell and Cochrane (1999), who also have an upper bound on surplus consumption that can never be crossed in continuous time.

2.2.3 Consumption Euler equation

Standard no-arbitrage conditions in asset pricing imply that the gross one-period real return $(1 + R_{t+1})$ on any asset satisfies

$$1 = \mathrm{E}_t \left[M_{t+1} \left(1 + R_{t+1} \right) \right], \tag{13}$$

where

$$M_{t+1} \quad = \quad \frac{\beta U'_{t+1}}{U'_t} \tag{14}$$

is the stochastic discount factor (SDF). The Euler equation for the return on a one-period real T-bill can be written in log form as:

$$\ln U'_t = r_t + \ln \beta + \ln \mathrm{E}_t U'_{t+1}. \tag{15}$$

For simplicity, we assume that short-term nominal interest rates contain no risk premia or that $i_t = r_t + E_t \pi_{t+1}$, where π_{t+1} is inflation from time t to time $t+1$. This approximation is justified if uncertainty about inflation is small at the quarterly horizon, as appears to be the case empirically. Substituting $r_t = i_t - E_t \pi_{t+1}$ into (15), and dropping constants to reduce the notational burden, we have:

$$\ln U'_t = (i_t - \mathrm{E}_t \pi_{t+1}) + \ln \mathrm{E}_t U'_{t+1}. \tag{16}$$

Substituting (10) into the Euler equation for the one-period real T-bill gives

$$\begin{aligned} r_t &= -\ln\beta + \gamma g + \gamma\theta_2 x_{t-1} + \gamma\left(\theta_1 - \tau\phi\right)x_t + \gamma\tau \mathrm{E}_t x_{t+1} \\ &\quad -\gamma(1-\theta_0)(s_t - \bar{s}) - \frac{\gamma^2\sigma_c^2}{2}\left(1+\lambda(s_t)\right)^2. \end{aligned} \tag{17}$$

Now, we use Campbell and Cochrane (1999)'s condition that the terms in (17) involving s_t cancel, which imposes restrictions on the sensitivity function $\lambda(s_t)$. Moreover, habit must be predetermined at and near the steady state. These conditions ensure that the steady-state surplus consumption ratio and the sensitivity function λ are given by

$$\bar{S} = \sigma_c\sqrt{\frac{\gamma}{1-\theta_0}}, \tag{18}$$

$$\bar{s} = log(\bar{S}), \tag{19}$$

$$s_{max} = \bar{s} + 0.5(1-\bar{S}^2), \tag{20}$$

$$\lambda(s_t, \bar{S}) = \begin{cases} \frac{1}{\bar{S}}\sqrt{1-2(s_t-\bar{s})} - 1 & , s_t \leq s_{max} \\ 0 & , s_t \geq s_{max} \end{cases}. \tag{21}$$

We then re-arrange the Euler equation in terms of the current, lagged, and future log output gaps and the short-term real interest rate r_t, ignoring constants for simplicity:

$$x_t = \underbrace{\frac{\tau}{\tau\phi - \theta_1}}_{\rho^{x+}}\mathrm{E}_t x_{t+1} + \underbrace{\frac{\theta_2}{\tau\phi-\theta_1}}_{\rho^{x-}} x_{t-1} - \underbrace{\frac{1}{\gamma(\tau\phi-\theta_1)}}_{\psi} r_t. \tag{22}$$

Several points are worth noting about the IS curve (22). First, the asset pricing Euler equation holds without shocks. Second, because $\theta_1 > 0$, $\theta_2 > 0$ and $\phi < 1$, the coefficients on the lagged output gap and the expected future output gap sum to more than one. Third, the slope of the IS curve ψ does not equal the elasticity of intertemporal substitution (EIS) of the representative consumer.

The lag coefficient ρ^{x-} in the IS curve (22) is non-zero whenever the lagged output gap enters into surplus consumption (i.e. $\theta_2 \neq 0$). Cochrane (2011) shows that solutions to purely forward-looking New Keynesian models are typically ill-behaved. We therefore need $\theta_2 \neq 0$ to obtain a partly backward-looking Euler equation and well-behaved macroeconomic dynamics.

2.3 Modeling bonds and stocks

We model stocks as a levered claim on consumption c_t. We assume that log dividend growth is given by:

$$\Delta d_t = \delta \Delta c_t. \tag{23}$$

We interpret δ as capturing a broad concept of leverage, including operational leverage. The interpretation of dividends as a levered claim on consumption is common in the asset pricing literature (Abel 1990, Campbell 1986, 2003). We maintain our previous simplifying approximation that risk premia on one-period nominal bonds equal zero, but risk premia on longer-term bonds are allowed to vary, so the expectations hypothesis of the term structure of interest rates does not hold.

2.4 Model solution and stability

We first describe the solution for macroeconomic dynamics. The state variable dynamics have a solution of the form

$$\hat{Y}_t = P\hat{Y}_{t-1} + Qu_t, \tag{24}$$

where

$$\hat{Y}_t = [x_t, \hat{\pi}_t, \hat{\imath}_t]', \tag{25}$$

$$\hat{\pi}_t = \pi_t - \pi_t^*, \tag{26}$$

$$\hat{\imath}_t = i_t - \pi_t^*. \tag{27}$$

We solve for $P \in \mathbb{R}^{3\times 3}$ and $Q \in \mathbb{R}^{3\times 4}$ using the method of generalized eigenvectors (see e.g. Uhlig 1999).

In principle, the model can have more than one solution. We only consider dynamically stable solutions with all eigenvalues of P less than one in absolute value, yielding non-explosive solutions for the output gap, inflation gap, and interest rate gap. Cochrane (2011) argues that there is no economic rationale for ruling out solutions solely on the basis of an explosive inflation path. In general, in our model an explosive solution for inflation is also explosive for the output gap and the real interest rate. We find it reasonable to rule out such solutions with explosive real dynamics.

The inclusion of backward-looking terms in the IS curve and PC implies that there exist at most a finite number of dynamically stable equilibria of the form (24). This is true even when the monetary policy reaction to inflation (γ^π) is smaller than one, which usually leads to an indeterminate equilibrium in highly stylized Keynesian models with only forward-looking components (Cochrane 2011).

Next, we require all our equilibria to satisfy a battery of equilibrium selection criteria to rule out unreasonable solutions and pick a unique solution. We require the solution to be real-valued and "expectationally stable" (Evans 1985, 1986, McCallum 2003). Expectational stability requires that for small deviations from rational expectations, the system returns to the equilibrium. We also impose the solution selection criterion of Uhlig (1999), which is closely related to the minimum state variable solution proposed by McCallum (2004).

While we formally model regimes as lasting an infinite period of time, one might think that agents understand that the regime will have to end eventually, potentially arbitrarily far in the future. We implement the Cho and Moreno (2011) criterion, which captures this limiting case. This criterion, also used by Bikbov and Chernov (2013), has two appealing interpretations. The first interpretation captures the notion that if monetary policy changes slowly over time and those changes are fully anticipated, even monetary policy regimes with weak inflation responses may have unique equilibria (Farmer, Waggoner, and Zha 2009).[12] The Cho and Moreno (2011) criterion is equivalent to assuming that the system returns to an equilibrium with all variables constant from period T^* onwards and then letting T^* go to infinity. An alternative interpretation of the Cho and Moreno (2011) criterion is closely related to expectational stability. If agents deviate from rational expectations and instead have constant expectations, the system returns to the Cho and Moreno (2011) equilibrium. The Appendix provides full details on the model solution and solution criteria.

2.5 Solutions for bond and stock returns

We solve numerically for bond prices and equity dividend-price ratios. On the numerical side, this paper contributes by extending the value function iteration methodology of Wachter (2005) to multiple state variables. We found that methods relying on analytic linear approximations to the sensitivity function λ (e.g. Lopez, López-Salido, and Vazquez-Grande 2014), numerical higher-order perturbation methods (Rudebusch and Swanson 2008), and

[12] We thank Mikhail Chernov for pointing out to us that when rational agents anticipate a return to a different equilibrium, even regimes with an inflation reaction coefficient less than one can have a determinate equilibrium.

numerical global projection methods led to substantial approximation error. For our baseline grid and simulation, it takes about 45 minutes to solve and simulate the model once for all three subperiod calibrations. The model solution is therefore too slow for estimation, but we can conduct grid searches over lower-dimensional parameter subsets.

Let P^d_{nt}/D_t denote the price-dividend ratio of a zero-coupon claim on the aggregate stock market dividend at time $t+n$. The price-dividend ratio on the aggregate stock market then equals the infinite sum of zero-coupon price-dividend ratios. The price of a zero-coupon claim for the dividend at time t is given by $P^d_{0t}/D_t = 1$. For $n \geq 1$, we solve for the n-period price-dividend ratio numerically using the iteration

$$\frac{P^d_{nt}}{D_t} = \mathrm{E}_t\left[M_{t+1}\frac{D_{t+1}}{D_t}\frac{P^d_{n-1,t+1}}{D_{t+1}}\right].$$

Denoting n-period real and nominal zero-coupon bond prices by $P_{n,t}$ and $P^\$_{n,t}$, one-period bond prices are given by

$$P^\$_{1,t} = \exp(-\hat{\imath}_t - \pi^*_t - r^f), \tag{28}$$

$$P_{1,t} = \exp(-\hat{\imath}_t + \mathrm{E}_t\hat{\pi}_{t+1} - r^f). \tag{29}$$

For $n > 1$, zero-coupon bond prices follow the recursions:

$$P_{n,t} = \mathrm{E}_t\left[M_{t+1}P_{n-1,t+1}\right], \tag{30}$$

$$P^\$_{n,t} = \mathrm{E}_t\left[M_{t+1}\exp(-\pi_{t+1})P^\$_{n-1,t+1}\right]. \tag{31}$$

Since n-period zero-coupon nominal bond prices are proportional to $\exp(-n\pi^*_t)$, we solve numerically for scaled nominal bond prices $\exp(n\pi^*_t)P^\$_{n,t}$.

We solve over a five-dimensional grid of the surplus consumption ratio s_t, the lagged output gap x_{t-1}, and the scaled vector $\tilde{Y}_t$. $\tilde{Y}_t$ scales and rotates the variables in $\hat{Y}_t$ such that shocks to $\tilde{Y}_t$ are independent standard normal and the first element of $\tilde{Y}_t$ is conditionally perfectly correlated with consumption innovations. The baseline solution grid uses 50 gridpoints for s_t spaced between -50 and s_{max}, and two gridpoints along every dimension of $\tilde{Y}_t$ at plus and minus two standard deviations from the unconditional mean. It also uses two gridpoints for x_{t-1} ranging over all output gap values covered by the grid for x_t. We therefore have a total of 800 gridpoints. We use Gauss-Legendre 40-point quadrature to integrate over consumption shocks, and 10-point quadrature for those dimensions of $\tilde{Y}_t$ that are contemporaneously uncorrelated with the SDF. We evaluate price-dividend ratios and bond prices between gridpoints using five-dimensional loglinear interpolation. Asset pricing properties are unchanged if we increase the grid size for any of these dimensions, indicating that the baseline grid is sufficient.

2.6 Properties of bond and stock returns

The solutions for bond and stock returns imply that returns are conditionally heteroskedastic (even though macroeconomic fundamentals are homoskedastic by assumption), and that conditional expected asset returns vary over time with the surplus consumption ratio. Time-varying risk premia generate a non-linear effect of fundamental shocks on bond betas which can amplify their linear effect. For example, consider a contractionary shock that simultaneously lowers output and inflation. The shock pushes bond valuations higher and stock valuations lower, generating a negative nominal bond beta. But the negative bond beta implies that nominal bonds are safe. The increase in risk aversion following the contractionary shock makes nominal bonds more valuable hedges, driving up their prices and making the bond beta even more negative. We show in our calibration that amplification through time-varying risk premia can be quantitatively important.

3 Preliminary Empirical Analysis

3.1 Monetary policy regimes

We explore monetary policy in three subperiods, which we determine using a Quandt Likelihood Ratio (QLR) test. The resulting break dates correspond closely to changes in bond betas in Figure 1, a potentially alternative criterion for determining breaks. Our first subperiod, 1960Q2–1977Q1, covers roughly the Fed chairmanships of William M. Martin and Arthur Burns. The second subperiod, 1977Q2–2000Q4, covers the Fed chairmanships of G. William Miller, Paul Volcker, and Alan Greenspan until the end of the long economic expansion of the 1990s. The third subperiod, 2001Q1-2011Q4, contains the later part of Greenspan's chairmanship and the earlier part of Ben Bernanke's chairmanship.[13]

Our identification of a third regime for monetary policy is supported by several observations. First, in the late 1990s and 2000s the Federal Reserve has placed increased emphasis on transparency and providing guidance to market expectations, in part through changing the policy rate gradually (Coibion and Gorodnichenko 2012, Stein and Sunderam 2015). As a result of greater transparency and credibility, the central bank may have been able to affect the real economy not only through immediate changes in the Fed Funds rate, as in the earlier periods, but also through expectations of future changes. In our model, inflation

[13] Sims and Zha (2006) argue for a break in the volatility of money demand shocks in 2000 and a break in the monetary policy rule in 1987. In the appendix, we show that the main findings in the paper are robust to assuming a break in the monetary policy rule in 1987 and a break in volatilities of shocks in 2000. With these alternative break dates, the model attributes changes in bond betas around 2000 to changes in monetary policy, and especially the volatilities of inflation target and monetary policy shocks.

target shocks can capture credible central bank announcements of future actions that are not accompanied by any immediate rate changes.

Second, the experience of moderate inflation and apparently well anchored inflation expectations from the mid-1980s through the mid-1990s seems to have encouraged the Federal Reserve to turn its attention back to output stabilization, after the single-minded focus on combating inflation under Fed chairman Paul Volcker. Rigobon and Sack's (2003) empirical evidence is also consistent with this interpretation.

Illustrating both investors' focus on the central bank's output response and the forward-looking nature of monetary policy in the 2000s, a typical *New York Times* bond market commentary argued in 2000: "Prices of Treasury securities were down (...), after a stronger-than-expected gain in industrial production raised investor concern about further interest rate increases by the Federal Reserve."[14]

3.2 Data and summary statistics

We use quarterly US data on output, inflation, interest rates, and aggregate bond and stock returns from 1954Q3 to 2011Q4. GDP in 2005 chained dollars and the GDP deflator are from the Bureau of Economic Analysis via the Fred database at the St.Louis Federal Reserve. We use the end-of-quarter Federal Funds rate from the Federal Reserve's H.15 publication and the availability of this data series determines the start date of our analysis. We use quarterly potential GDP in 2005 chained dollars from the Congressional Budget Office.[15]

[14]Hurtado, Robert, 2000, Treasury Prices Fall on Report of Higher Factory Output, *The New York Times*, February 16.

[15]Table 2-3 of the CBO's August 2012 report "An Update to the Budget and Economic Outlook: Fiscal Years 2012 to 2022" (http://www.cbo.gov/publication/43541). Averaging the Federal Funds rate over the last week of the quarter eliminates spikes in the Fed Funds rate due to banks' liquidity requirements on the last day of the quarter.

The end-of-quarter five year bond yield is from the CRSP monthly Treasury Fama-Bliss discount bond yields. The 5-year TIPS yield, 5-year breakeven, and daily 5-year nominal bond yields are from Gürkaynak, Sack and Wright (2010). We use the value-weighted combined NYSE/AMEX/Nasdaq stock return including dividends from CRSP, and measure the dividend-price ratio using data for real dividends and the S&P 500 real price.[16] Interest rates, and inflation are in annualized percent, while the log output gap is in natural percent units. All yields are continuously compounded. We consider log returns in excess of the log T-bill rate. The end-of-quarter three-month T-bill is from the CRSP monthly Treasury Fama risk-free rate files and is based on the average of bid and ask quotes.

Table 1 shows summary statistics for the log output gap, inflation, the Federal Funds rate, and the 5-year nominal bond yield for the US over the full sample period 1954Q3-2011Q4 and over the three subperiods. The log real output gap has a first-order quarterly autocorrelation of 0.96 over the full sample period, implying a half life of 5 years. Realized inflation, the Fed Funds rate and the 5-year nominal bond yield are also highly persistent in the full sample and across subperiods. The average log output gap was positive in the earliest subperiod, and negative afterwards. Inflation and interest rates were significantly lower in the latest subperiod compared to the early subperiods.

In our model, variation in the output gap drives consumption innovations and hence risk premia. We now verify empirically the relation between equity risk premia and the output gap, and examine the relation of the output gap with well known predictors of excess stock returns, such as the price-dividend ratio.

Figure 2, Panel B shows the log output gap and the log price-dividend ratio for the full sample period. The correlation between the two variables is 0.47, 0.54, and 0.62 for the first,

[16] The source is Robert Shiller's website at http://www.econ.yale.edu/ shiller/data.htm.

second, and third subperiod. The weighted average correlation across the three subperiods is 0.54. This average correlation is less than one, but nonetheless strongly positive, supporting the model's link between the output gap, consumption, and risk premia. While the log price-dividend ratio clearly varies cyclically with the output gap, Figure 2 also shows longer-term shifts in the price-dividend ratio across regimes, which our model does not capture and which drive up the volatility of the log price-dividend ratio in the data as compared to the model.

Table 2 uses the output gap to predict equity log excess returns:

$$r^e_{t+1} - i_t = a^0 + a^x x_t + \epsilon_{t+1}. \tag{32}$$

The point estimate of a^x is negative for each subperiod and significant in the first subperiod, consistent with our model specification. Subsample estimates vary around the full sample estimate of $a^x = -0.49$.

3.3 Estimating monetary policy rules

Following CGG, we estimate a monetary policy rule in terms of the output gap, inflation, and the Fed Funds rate:

$$i_t = c^0 + c^x x_t + c^\pi \pi_t + c^i i_{t-1} + \epsilon_t. \tag{33}$$

If monetary policy shocks had no contemporaneous effect on output and inflation, and if the inflation target were constant, we could use (33) to back out monetary policy rule parameters according to:

$$\hat{\rho}^i = \hat{c}^i, \tag{34}$$

$$\hat{\gamma}^x = \hat{c}^x/(1-\hat{c}^i), \tag{35}$$

$$\hat{\gamma}^\pi = \hat{c}^\pi/(1-\hat{c}^i). \tag{36}$$

In the full model, we do not use the potentially biased "naïve" estimates (34) through (36). Instead, as we explain in greater detail in Section 4.1, we use the model to account for regression bias, backing out monetary policy parameters γ^x, γ^π and ρ^i to match the empirical slope coefficients (33) for each subperiod. Alternatively, one could introduce additional modeling assumptions to ensure that (34) through (36) give unbiased estimates of the true monetary policy rule (Backus, Chernov, and Zin 2013). We consider the naïve monetary policy parameter estimates a useful sanity check for the model-implied monetary policy changes.

We start our preliminary analysis of monetary policy regimes by determining the start and end dates of subperiods. Even if naïve monetary policy parameter estimates are biased, a break in the relation (33) should indicate a change in monetary policy. CGG have argued forcefully that the monetary policy rule changed substantially in the early 1980s. It is therefore plausible that we should find one or more breaks in the monetary policy rule.

We estimate monetary policy break dates using a sequence of three QLR tests. First, we test for a break over our full sample period. We interact all coefficients in (33) with post-break date dummies for all potential break dates within the middle 50% of the sample. The estimated break date — the date with the highest F-statistic against the null hypothesis of no interaction terms — is 1977Q2 for our full sample. The test statistic for the null hypothesis of no break in 1977Q2 exceeds its 95% critical value if we treat the break date as known, but not if we treat the break date as unknown.[17] If there are two or more breaks in the sample and the post-2000 regime has similarities with the regime in the 1960s, this might make it harder to reject the null of no break in the full sample. Next, we test for breaks in the pre- and post-1977Q2 subsamples. The estimated break dates are 1960Q2 and 2001Q1. In both cases, we can reject the null of no break with or without known break date at the 95% confidence level.

Table 3 reports OLS Taylor rule regressions and naïve implied monetary policy parameters for the three subperiods.[18] Table 3 suggests that monetary policy has varied substantially over time. The output gap slope coefficient is positive and statistically significant for the first subperiod and is small for the two later subperiods. The inflation slope coefficient increases from about 0.2 in the pre-1977 period to 0.4 in the post-1977 period and comes back down to 0.2 during the post-2000 period. Finally, the monetary policy smoothing parameter ρ^i is stable during the first two periods and increases during the post-2000 period. During the most recent subperiod the regression explains 94% of the variation in the Federal Funds rate, consistent with a shift away from transitory monetary policy shocks towards expectational guidance in central bank policy.

While the slope coefficients in the upper panel of Table 3 estimate the short-run response of monetary policy to inflation and output fluctuations, the naïve monetary policy coefficients in the bottom panel give a better account of the long-run monetary policy response. The naïve implied parameters indicate that during the earliest subperiod, 1960Q2–1977Q1, the central bank raised nominal interest rates less than one-for-one with inflation. In contrast,

[17] We use the 5% critical value for 4 restrictions and 25% trimming tabulated in Andrews (2003).

[18] Standard errors for the naïve monetary policy parameters use the delta method. Asterisks indicate parameters that are significant at the 5% or 1% level based on a likelihood ratio test, which may differ from significance implied by the delta method if the relation between parameters is nonlinear.

the central bank raised nominal interest rates more than one-for-one with inflation during the both the later two subperiods (1977Q3–2011Q4). Hence, even though our statistically determined break date is slightly earlier than the break date in CGG, empirical results are similar.

The point estimates of $\hat{\gamma}^x$ in Table 3 also suggest that the central bank has put somewhat higher weight on output fluctuations in the earliest and latest subperiods than during the middle subperiod, although neither the estimates of $\hat{\gamma}^x$ nor $\hat{c}^x$ are statistically significant in the latest subperiod.

The estimated OLS monetary policy rule is similar for the last subperiod when we exclude the financial crisis. The Appendix estimates monetary policy rules for two parts of the third subperiod, before and after the start of the financial crisis, which we take to be the third quarter of 2008.

4 Model Calibration

We now calibrate our model to key empirical moments for the US over the three subsamples: 1960Q2-1977Q1, 1977Q2-2000Q4, and 2001Q1-2011Q4. Table 4 summarizes the calibration parameters, while Tables 5 and 6 compare key empirical and model moments.

4.1 Calibration procedure

We separate the parameters into two blocks. Parameters in the first block are held constant across subperiods, while parameters in the second block correspond to our main candidate explanations for changes in bond betas and change across subperiods. Time-invariant parameters include those governing the relation between the output gap, consumption, and dividends (ϕ, τ, g, δ), preference parameters (γ, θ_0, θ_1, θ_2, $\bar{r}$), and Phillips curve parameters (ρ^π, κ). Time-varying parameters include the monetary policy rule parameters γ^x, γ^π, and ρ^i and the shock volatilities σ^{PC}, σ^{MP}, and σ^*. Our selection of parameter blocks is consistent with Smets and Wouters (2007), who estimate a structural New Keynesian model separately for the periods 1966-1979 and 1984-2004. They find important changes in the shock volatilities and the monetary policy parameters across those two periods, whereas estimated preference parameters are largely stable across subperiods.

Within the first block of time-invariant parameters, we set the leverage parameter $\delta = 3.39$ to match the relative volatility of stochastically detrended dividends and consumption. This corresponds to firm leverage of 68%, where we interpret leverage broadly as incorporating operational leverage, leases, and fixed obligations to non-shareholders. We set the average riskfree rate $\bar{r} = 0.94\%$ and the average consumption growth rate $g = 1.89\%$. The persistence of the surplus consumption ratio $\theta_0 = 0.97$ per quarter, corresponding to an annual persistence of 0.87, is exactly as in Campbell and Cochrane (1999).

The parameter ϕ is an important determinant for the dynamic properties of consumption growth. If the output gap follows an AR(1) process and the consumption-output relation is governed by (8), consumption growth is serially uncorrelated if and only if ϕ equals the first-order autocorrelation of the output gap. Strong predictability in consumption growth generates excessive predictability of real rates and volatility of bond and stock returns in our model. In order to generate plausible volatilities of asset returns, we choose ϕ to generate a 12-quarter consumption variance ratio that averages one across subperiods. The resulting numerical value, $\phi = 0.94$, also generates the highest correlation between stochastically detrended consumption and the output gap, further corroborating our choice of ϕ. We choose the scaling parameter τ to match the ratio of full sample standard deviations for stochastically detrended consumption and the output gap.

We set the utility curvature parameter to $\gamma = 3$, somewhat higher than in Campbell and Cochrane (1999) and Wachter (2006). In our model, γ not only determines risk premia and the Sharpe ratio of risky assets, but it also enters into the Euler equation and equilibrium dynamics of the output gap, inflation, and Federal Funds rate. A higher value of γ flattens the relation between the output gap and the real short-term interest rate, and avoids explosive macroeconomic dynamics. At the same time, asset Sharpe ratios rise roughly with $\gamma/\bar{S} \propto \sqrt{\gamma}$. Therefore, setting $\gamma = 3$ leads to a small increase in asset Sharpe ratios as compared to Campbell and Cochrane (1999).

The new preference parameters θ_1 and θ_2 are both set to 0.02, so they are small but positive. The parameters θ_1 and especially θ_2 are also important for macroeconomic dynamics, since a unique equilibrium may not exist when they are set to zero. As we have seen, θ_1 and θ_2 enter into the Euler equation, so requiring a persistent output gap pins down their relative values.

We choose a Phillips curve slope of $\kappa = 0.04$. Rotemberg and Woodford (1997) and Woodford (2003) obtain a similarly small Phillips curve slope in a micro-founded New Keynesian model where prices on average remain constant for three quarters. The Phillips curve is strongly backward-looking with $\rho^{\pi} = 0.96$. A large backward-looking component helps generate unique and learnable equilibria and is consistent with empirical evidence by Fuhrer (1997). Gali and Gertler (1999) find some empirical evidence in favor of a forward-looking curve using the labor share of income instead of the output gap.

We next calibrate the subperiod-specific parameters in the second block. Due to computational constraints, we cannot jointly optimize over monetary policy rule coefficients and standard deviations of shocks. Therefore, we first optimize over monetary policy rule coefficients. Next, we optimize over the standard deviations of shocks while holding constant the monetary policy rule parameters.

We choose monetary policy parameters γ^x, γ^π and ρ^i to minimize the distance between the empirical OLS regressions reported earlier in Table 3 and identical regressions estimated in simulated data from the model. In this way we correct for potential regression bias caused by endogeneity of inflation and output and time-variation in the central bank's inflation target.

The calibrated monetary policy parameters mirror the broad changes in naïvely estimated monetary policy parameters in Table 3. The output gap weight γ^x decreased slightly from the pre-Volcker period to the Volcker period and then increased substantially in the post-2000 period. The inflation weight γ^π was below one during the pre-Volcker period, and greater than one thereafter. Finally, monetary policy persistence increased substantially after 2000.

For each subperiod, the three standard deviations of fundamental shocks are chosen to minimize the distance between model and empirical macroeconomic and asset second moments. For each subperiod, we run a grid search over the standard deviations of shocks. We minimize a weighted distance function in the residual standard deviations of a VAR(1) in the output gap, inflation, Federal Funds rate, and 5-year nominal bond yield, the standard deviations of bond and stock returns, and the betas of nominal and real bond returns.

As we explain in greater detail in Section 5.1, shock volatilities are identified because each volatility affects particular features of the data. More volatile PC shocks lead to more volatile inflation surprises and stock returns, and increase nominal bond betas. More volatile MP shocks lead to more volatile Fed Funds rate innovations and bond returns, and have a positive effect on nominal bond betas. More volatile shocks to the inflation target primarily increase the volatilities of output gap and nominal bond yield innovations, but have little effect on the volatilities of quarterly inflation and Fed Funds rate innovations, because target shocks act on inflation with a long lag. More volatile inflation target shocks also increase the volatilities of bond and stock returns, and decrease nominal bond betas.

The volatilities of fundamental shocks, σ^{IS}, σ^{MP} and σ^{*}, change substantially and plausibly across time periods. We estimate a substantially larger volatility of MP shocks for the period 1977-2000 than for the earliest subperiod and especially the latest subperiod. The estimated volatility of PC shocks is largest in the earliest subperiod, a period comprising major global oil price shocks, and smallest for the most recent subperiod. The calibrated volatility of inflation target shocks is small for all three subperiods, but increases in the third subperiod. As we will see, the model requires a higher volatility of inflation target shocks after 2000 to generate negative bond betas.

At first, it might seem counterintuitive that the Federal Reserve's inflation target was especially volatile during the most recent subperiod. However, it is important to keep in mind that inflation target shocks can be either positive or negative. The period 2001-2011 saw a steep decline in 5-year nominal bond yields from 4.6% to 0.9%, as would be the case if investors' perceived inflation target experienced a sequence of negative shocks that moved it towards or even below the Federal Reserve's officially stated target. Survey evidence is also consistent with economically meaningful uncertainty about the Federal Reserve's inflation target. In a 2012 special question by the Survey of Professional Forecasters, 21% of forecasters reported that their long-run inflation forecasts differed in an economically meaningful way from the officially stated inflation target of 2%. Individual forecasters' long-run annual-average personal consumption expenditures (PCE) inflation forecasts varied between 1.14% and 3.40%, suggesting substantial room for shocks to investors' inflation targets. Stein and Sunderam (2015) present a theoretical analysis of contemporary monetary policy in which inflation target uncertainty plays a key role.

Panel B of Table 4 shows implied calibration parameters. The calibrated Euler equation has a large forward-looking and a small backward-looking component. The implied slope of the IS curve with respect to the real interest rate equals $\psi = 0.41$ for each subperiod, which is within the range of empirical estimates by Yogo (2004) and earlier work by Hall (1988).[19]

4.2 Evaluating the fit of the model

Table 5 evaluates the model fit for all three subperiods. Panel A of Table 5 shows that the model matches the empirical OLS monetary policy rules estimated in Table 3, validating our choice of monetary policy parameters.[20] The top half of Table 5, Panel B reports volatilities of equity, nominal bond, and real bond returns, and the betas of nominal bonds and real bonds with respect to equities. The bottom half of Panel B shows model-implied and empirical volatilities of VAR(1) residuals in the output gap, inflation, the Federal Funds rate, and the 5-year nominal bond yield.

The middle panel shows that our model fits well the overall changes in nominal bond betas across subperiods, which are the primary object of interest in our analysis. Similarly to the data, the model generates a small but positive nominal bond beta in the pre-Volcker period, a strongly positive bond beta during the Volcker-Greenspan period, and a negative bond beta in the post-2000 period.

The last row in the middle panel shows that in the third subperiod, when US inflation-indexed bonds were available, the model implies a small but negative real bond beta that is

[19]The long-run risk literature, following Bansal and Yaron (2004), prefers a value greater than one for the elasticity of intertemporal substitution. We need an IS curve with a real rate slope less than one, because otherwise the effect of monetary policy on the output gap grows disproportionately, leading to a non-persistent output gap and loss of a stable equilibrium.

[20]All model moments are calculated from 2 simulations of length 50000. We choose a long simulation period to capture the steady-state distribution of the state variables.

comparable to that in the data.[21] If inflation-indexed bonds had been available during earlier periods, the model indicates that real bonds would have been especially valuable hedges with negative betas during the 1960Q2-1977Q1 subperiod, when PC shocks were dominant. However, real bonds would have had positive betas in the 1977Q2-2000Q4 subperiod.

Comparing the first rows of the middle and bottom panels shows that the time-varying risk premia in the model are sufficient to reconcile a low volatility of the output gap with much higher volatility of equity returns, addressing the "equity volatility puzzle". The model fits overall stock return volatilities quite well, although it slightly overstates volatility in the second period and understates it in the third period. Bond return volatilities are however lower in the model than in the data.

The bottom panel of Table 5 shows mixed results for the overall ability of our model to fit the empirical volatilities of VAR(1) residuals. The model matches well the level and time-variation in the volatility of Fed Funds rate innovations. However, it somewhat overstates the volatility of the output gap and understates the volatilities of inflation and the log nominal yield. The former understatement is more pronounced in the last two subperiods, while the latter is relatively constant across the subperiods.

Table 6 shows that the model generates empirically plausible implications for consumption dynamics (top panel), equity returns (middle panel), and bond returns (bottom panel). The top panel shows that the average annualized volatility of consumption innovations in the model is 1.94%, as compared to a standard deviation of 1.53% for quarterly consumption growth in the data. The twelve-quarter consumption variance ratio averages very close

[21] In the third subperiod, the model generates reasonable volatilities for 5-year breakeven, or the difference between nominal and inflation-indexed bond yields, supporting the calibrated volatility of long-term inflation expectations. The model standard deviation of quarterly changes in 5-year breakeven is 29 bps. The analogous empirical standard deviation is 38 bps, after adjusting for TIPS liquidity as in Pflueger and Viceira (2015) and TIPS indexation lags as in D'Amico, Kim, and Wei (2008).

to one across subperiods. Even though the model necessarily generates some predictability in consumption growth over the short run, medium-term consumption dynamics therefore closely resemble a random walk benchmark. Table 6 also shows that the output gap is similarly persistent in the model and in the data.

The middle panel of Table 6 shows that the model generates a high equity premium of 8%, which even exceeds that in the data. At the same time, the model obtains an average price-dividend ratio of 34.85, which is somewhat higher than the empirical price-dividend ratio over our sample. The combination of a high equity premium and a high equity price-dividend ratio can be reconciled by the fact that the model overstates the average dividend growth rate at $\delta \times g = 3.39 \times 1.89\% = 6.40\%$. The log price-dividend ratio has similar persistence but slightly less volatility than in the data. The somewhat lower volatility of the price-dividend ratio is not surprising in light of the longer-term non-cyclical shifts in the price-dividend ratio we observed in Figure 2, Panel B. The calibration generates a positive and empirically plausible correlation between the output gap and the log price-dividend ratio. Model stock returns are predictable from the log dividend-price ratio and the output gap with empirically plausible slope coefficients, indicating that the model has reasonable variation in equity risk premia. Moreover, the model generates heteroskedasticity in stock returns. Rather than regressing squared quarterly stock returns onto the lagged output gap, we use quarterly absolute stock returns in annualized percent standard deviation units, which gives more easily interpretable magnitudes. In the model, stock returns are conditionally more volatile when the output gap is low, similarly to the data.

The bottom panel of Table 6 evaluates the implications of the model for bond returns. The model implies a term structure that is on average slightly upward-sloping. A more strongly upward-sloping term structure in the data might reflect extreme liquidity of short-term Treasury debt (Greenwood and Vayanos 2008, Krishnamurthy and Vissing-Jorgensen 2012, Nagel 2014). The average standard deviation of the real interest rate in the model is a reasonable 1.83% per annum.

The average slope of the term structure reflects upward-sloping term structures in the first two subperiods (46 bps and 83 bps) and a downward-sloping term structure of -55 bps in the third subperiod, consistent with negative bond betas during this subperiod. Table 6 reports a cross-subperiod Campbell and Shiller (1991) regression. Model bond risk premia vary two times more strongly across subperiods than the slope of the term structure, in line with a full-sample Campbell and Shiller (1991) regression of quarterly bond excess returns onto the lagged term spread. While the model does not generate substantial comovement between the slope of the term structure and bond risk premia within each subperiod, these results are consistent with empirical evidence of stronger bond return predictability from the slope of the term structure in longer samples (Pflueger and Viceira 2011). While it is beyond the scope of this paper, it would be natural to build on our framework to explore gradually changing monetary policy coefficients and macroeconomic uncertainty as drivers of higher-frequency changes in bond risk premia. This might also help increase the somewhat low volatility of bond returns, as Campbell and Ammer (1993) report that time-varying risk premia contribute meaningfully to bond volatility.

A comparison with Wachter's (2006) results is in order. Wachter's (2006) framework differs from ours in two key respects. First, Wachter (2006) models bonds as having a strongly positive beta, making bonds risky and generating an upward-sloping term structure. Our calibration exercise shuts down this channel, because we constrain ourselves to matching bond betas in different subperiods. Second, Wachter (2006)'s real short-term interest rates inherit the heteroskedasticity of surplus consumption. This heteroskedasticity generates highly time-varying risk premia and bond return predictability, but also a positive relation between the nominal short rate and the slope of the term structure. The more moderately cyclical risk premia in our model are consistent with a negative empirical correlation between the Fed Funds rate and the slope of the term structure.

5 Counterfactual Analysis of Changing Bond Risks

We are now in a position to investigate the role of changing monetary policy and macroeconomic shocks for bond betas. Our calibrated model replicates the broad shift in nominal bond betas over time. The calibration procedure explicitly incorporates changes in monetary policy and changes in the volatilities of shocks across subperiods and both may contribute to time-varying bond risks to varying degrees.

5.1 Impulse responses

Impulse response functions clarify the mechanism by which individual shocks act on stocks and bonds. Figure 3 shows responses of the output gap, inflation, the nominal short-term interest rate, the real short-term interest rate, the 5-year nominal bond yield, and the equity dividend yield to one-standard-deviation fundamental shocks. Figure 4 further decomposes bond and equity market responses into risk neutral and risk premium components.[22] Each panel shows three lines, each corresponding to one subperiod calibration. Vertical bars

[22]Impulse responses for the nominal bond yield and dividend yield are averaged over 10000 simulations, starting from the stochastic steady-state at time 0. The responses for the output gap, inflation, the nominal short rate and the real short rate can be computed analytically, since these variables follow a VAR(1) in equilibrium. Risk neutral bond yields are the average of expected future nominal short-term interest rates. Risk neutral equity dividend yields are from a Campbell and Shiller (1988) approximate loglinear decomposition and correspond to the sum of expected future real short-term interest rates less expected future dividend growth. Risk premium components reflect future expected excess returns.

indicate the size of the initial shock reaction. The size of each shock is identical across subperiods, and equal to the sample-size weighted average of the shock standard deviation.

Figure 3 shows that a typical PC shock acts as a persistent inflationary and contractionary supply shock, decreasing the output gap and raising inflation. For subperiods 1 and 2, a PC shock moves the log dividend-price ratio and bond yields, and stock and bond prices, in the same direction and hence contributes to a positive bond beta. Short term nominal bond yields increase following a PC shock, reflecting the central bank's positive weight on inflation fluctuations. In contrast, the real rate drops as the central bank accommodates the decline in the output gap.

The three subperiod calibrations show differential responses to PC shocks and those differences are related to changes in monetary policy. The second subperiod shows especially positive responses for nominal bond yields and equity dividend yields. Intuitively, the Federal Reserve raises the policy rate aggressively and immediately in response to an inflationary shock, which reduces both bond prices and stock prices at the same time and drives up the bond beta.

In contrast, in subperiod 3 nominal bond yields decline following a PC shock. A positive PC shock increases expected nominal short rates and the risk neutral component of 5-year nominal bond yields. However, the risk premium component in nominal bond yields decreases sufficiently to dominate the overall response, as shown in the first column in Figure 4.

A monetary policy (MP) shock acts as a strongly positive impulse to nominal and real short-term interest rates. It induces consumers to postpone consumption, leading to a decrease in the output gap and a slow decrease in inflation. Both bond yields and equity dividend yields increase following a MP shock, so MP shocks tend to raise bond betas.

Finally, a shock to the central bank's policy target has a delayed but permanent effect on inflation and generates negative bond betas. Monetary policy in our model cannot affect real interest rates in the long-run, so expectations of permanently higher policy rates lead to a gradual increase in inflation towards the new target. Increasing inflation induces firms with nominal rigidities to produce more, as captured by the Phillips curve relation in the model. The increase in the output gap leads to a decline in risk premia and an increase in stock returns. At the same time, increased inflation expectations and short-term nominal interest rates raise nominal bond yields and lower nominal bond prices. Inflation target shocks therefore act to reduce the beta of nominal bonds. In our model, an announced and fully credible disinflation, as proxied by a negative inflation target shock, moves inflation and output in the same direction similarly to Mankiw and Reis (2002).

Bond and stock responses following monetary policy shocks are in line with empirical results in the literature. Bernanke and Kuttner (2005) estimate that a one percentage point surprise increase in the policy rate leads to a decrease in equity prices of between 2.5% and 4.6%. In our model, a one percentage point MP shock leads to a drop in stock prices ranging from 1.9% to 4.1%, depending on the subperiod. Bond yields in our model increase by 12 to 15 bps following a one percentage point MP shock and by 48 to 91 bps following a one percentage point inflation target shock. These results are in line with Cochrane and Piazzesi (2002), who estimate an increase in medium-term bond yields of 20 to 80 bps following a one percentage point surprise in the Federal Funds rate. Cochrane and Piazzesi's proxies of Fed Funds rate surprises have large effects on long-term bond yields and appear to forecast increases in inflation and economic activity, consistent with these proxies loading at least partially onto policy target shocks in our model.

Figure 4 decomposes bond and equity dividend yield responses into risk neutral and risk premium components. Rows one and two show that dividend yield responses to all shocks are largely driven by risk premia. The risk neutral nominal bond yield in row three reflects expected future short-term interest rates, while the risk premium component in row four reflects future expected excess returns on nominal bonds.

We first describe bond risk premia for subperiod 3. Because nominal bonds have a negative beta in the third subperiod, their safety value is especially large during times of high risk aversion. Risk premia increase during recessions, leading to a flight to quality and lowering bond yields. As a result, bond risk premia decrease following PC and MP shocks, but increase following inflation target shocks. This is an illustration of the important risk premium channel operative in our model. Bonds in periods 1 and 2 are risky, so bond risk premium responses have the opposite sign to that in period 3.

Time-varying risk premia are quantitatively significant and amplify the comovement of bonds and stocks. Table 7 reports correlations between risk neutral and risk premium components of bond and stock returns for all three subperiods. In the pre-Volcker period and the Volcker-Greenspan period, risk premium components of bonds and stocks are highly positively correlated, driving up the overall co-movement of bond and stock returns, as in the empirical analysis of Campbell and Ammer (1993). In the post-2000 period, time-varying risk premia in bonds and stocks are highly negatively correlated, making the overall bond-stock correlation substantially more negative than the corresponding risk neutral correlation.

Impulse responses illustrate the qualitative effects of shocks on bond betas, holding constant the risk properties of bonds and stocks. The next section compares the full effects of shock volatilities and monetary policy changes on bond betas.

5.2 Counterfactual bond risks

We now answer our initial question, namely how changes in monetary policy and shock volatilities contribute to shifts in bond betas across subperiods. The model has implications for how bond betas would have changed if only monetary policy or the volatility of PC shocks had changed. Figure 5 shows model-implied nominal (Panel A) and real (Panel B) bond betas as we vary individual parameters while holding the remaining parameters constant at their pre-Volcker values. The left panels decompose model-implied betas into changes due to monetary policy and changes due to PC shock volatilities. The right panels further decompose monetary policy changes into the monetary policy rule, the volatility of transitory MP shocks, and the volatility of persistent policy target shocks.

As we have already discussed, PC shocks tend to create stagflationary recessions, thereby increasing the beta of nominal bonds. The red dash-dot line in the top left panel of Figure 5 shows that as PC shocks became less dominant over time, this acted to generate a smaller, but not a negative, nominal bond beta after the late 1970s. In order to understand why we saw an increased nominal bond beta during the Volcker period and subsequently a negative bond beta after 2000, we turn to monetary policy changes.

The green dashed line in the top left panel of Figure 5 shows that if only monetary policy had changed in the late 1970s, but the volatility of PC shocks had remained constant, we should have seen a strong increase in nominal bond betas in the late 1970s and a negative bond beta after 2000, consistent with the data. The top right panel shows that both the change towards a more anti-inflationary monetary policy rule and the increase in MP shocks contributed to the increased bond beta in the late 1970s. During the Volcker-Greenspan period, the central bank reacts to PC shocks by raising the short-term nominal policy rate strongly and immediately, amplifying the negative effect on both nominal bond and stock prices and increasing the nominal bond beta.

Two changes in monetary policy drive the decrease in nominal bond betas after 2000. First, as monetary policy moved away from surprise movements in the policy rate towards providing more forward guidance to markets, the composition of perceived monetary policy innovations shifted towards a larger persistent component, driving down nominal bond betas. Second, the renewed focus on output gap fluctuations implies a more accommodative monetary policy response to PC shocks, decreasing the effect of PC shocks on nominal bond betas.

Monetary policy may affect nominal bond betas because it affects real bond betas, inflation and inflation risk premia, or both. Panel B of Figure 5 shows that while changes in the monetary policy rule and the volatility of transitory MP shocks imply broadly similar changes in nominal and real bond betas, changing PC shock volatilities and inflation target volatilities have dramatically different effects on nominal and real bond betas. The bottom left panel indicates that as PC shocks became less dominant, this pushed up real bond betas both after the late 1970s and after 2000. In contrast to nominal bond betas, real bond betas are unaffected by the volatility of policy target shocks.

We conclude that changes in the monetary policy rule explain the increase in the empirical nominal bond beta in the late 1970s. On the other hand, the model attributes the negative nominal bond beta after 2000 to a combination of a renewed monetary policy focus on output gap fluctuations, more volatile persistent monetary policy shocks, and smaller supply shocks.

6 Conclusion

Given the importance of nominal US Treasury bonds in investment portfolios, and in the design and execution of fiscal and monetary policy, financial economists and macroeconomists need to understand the determinants of Treasury bond risks. This is particularly challenging because the risk characteristics of nominal Treasury bonds are not stable over time.

This paper provides a new framework for modeling the influence of monetary policy on the macroeconomy and asset prices. The model not only allows us to study the co-movement of bonds and stocks but also how macroeconomic supply shocks, central bank responses to those shocks, and monetary policy uncertainty affect asset prices. We propose a model that integrates the building blocks of a standard New Keynesian model into a habit formation asset pricing framework, where risk premia can vary in response to macroeconomic conditions. We calibrate our model to US data between 1960 and 2011, a period in which macroeconomic conditions, monetary policy, and bond risks have experienced significant changes. We allow for discrete regime changes just before the second quarter of 1977 and the first quarter of 2001.

The model generates empirically plausible volatilities for stock returns, low consumption and output gap volatilities, and time-variation in bond and equity risk premia. At the same time, it matches the pattern of changing nominal bond betas across subperiods.

Our model is sufficiently rich to allow for a detailed exploration of the monetary policy drivers of bond and equity risks. We find that several elements of monetary policy have been especially important drivers of bond risks during the last half century. First, a strong reaction of monetary policy to inflation shocks increases both the beta of nominal bonds and the volatility of nominal bond returns. Large increases in short-term nominal interest rates in response to inflation shocks tend to lower real output and stock prices, while causing bond prices to fall. Our model attributes the large positive beta and high volatility of nominal bonds after 1977 to a change in monetary policy towards a more anti-inflationary stance. Evidence of such a change has been reported by Clarida, Gali, and Gertler (1999) and other papers studying monetary policy regimes, but our model clarifies how this alters the behavior of the bond market.

Second, our model implies that changes in the volatility of supply shocks, or shocks to the Phillips curve, can also affect bond risks. Supply shocks, such as oil price shocks, move inflation and output in opposite directions, making bond returns procyclical. We estimate that the volatility of these shocks decreased after 1977, consistent with a Great Moderation in macroeconomic shocks. If monetary policy had remained constant, decreases in the volatilities of supply shocks would have implied decreases in nominal bond betas after the late 1970s. Monetary policy changes in the late 1970s counteracted the effect of smaller PC shocks, instead leading to higher nominal bond betas. However, after 2000 the smaller volatility of PC shocks was reinforced by monetary policy and helped generate negative nominal bond betas.

Fourth, we find that the size and composition of shocks to the monetary policy rule affect bond betas. A smaller volatility of transitory monetary policy shocks and a higher volatility of persistent shocks contributes to negative nominal bond betas after 2000. These persistent inflation target shocks may be interpreted literally, as the result of shifting central bank preferences, or more broadly as changes in investors' perceived long-term policy target (Orphanides and Williams 2004).

An important lesson of our model is that changing fundamental risks can be amplified by time-variation in risk premia. Because risk premia are countercyclical in our model, assets with positive betas have risk premia that increase in recessions, driving down their prices and further increasing their betas. Assets with negative betas, on the other hand, become even more desirable hedges during recessions; this increases their prices and makes their betas even more negative. Thus the dynamic responses of risk premia amplify sign changes in betas that originate in changes in monetary policy, and underline the importance of nonlinear effects in understanding the impact of changes in monetary policy and macroeconomic shocks on asset prices.

We combine a loglinear New Keynesian model and habit formation preferences to match empirical properties of bond and equity risks. At the same time, there are additional channels that could fruitfully be explored by building on this benchmark model. First, we could model dividends not just as levered consumption claims but as claims also on the output gap to capture short-run cyclical variation in dividends; we could introduce demand shocks to the IS curve (a change that has little effect on the pricing of consumption claims but may be more important for the pricing of output gap claims); and we could allow the log surplus consumption ratio to enter the IS curve by adapting the preference specification of Wachter (2006). Second, our use of a habit-formation model shuts down the pricing of long-run risks that is the focus of a large literature following Bansal and Yaron (2004). Third, it could be instructive to explicitly model linear and nonlinear effects of monetary policy on corporate investment decisions, labor, and asset prices. Production and labor can be particularly challenging in habit formation models, even though some of these questions are starting to be addressed (Jermann 1998, Lettau and Uhlig 2000, Boldrin, Christiano, and Fisher 2001, Uhlig 2007, Rudebusch and Swanson 2008, Lopez, López-Salido, and Vazquez-Grande 2014). Fourth, while the regime shifts we consider are unanticipated, our model could provide a useful platform for understanding how expectations of future monetary policy and forward guidance affect asset prices today. Finally, we calibrate our model to US historical data but it will be valuable to extend this analysis to comparative international data on monetary policy in relation to bond and stock returns. Countries such as the UK, where inflation-indexed bonds have been issued for several decades, will provide particularly useful evidence on the comparative risks of real and nominal bonds, and their changes over time.

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Tables and Figures

Table 1: Summary Statistics

1954.Q3-2011.Q4	Output Gap	Inflation	Fed Funds	Nom. Bond Yield
Mean	-0.63	3.38	5.39	5.95
Std	(2.91)	(2.34)	(3.54)	(2.77)
AR(1) Coefficient	0.96	0.85	0.93	0.97
	(0.02)	(0.03)	(0.03)	(0.02)
AR(4) Coefficient	0.69	0.77	0.81	0.91
	(0.05)	(0.04)	(0.04)	(0.03)

1960.Q2-1977.Q1	Output Gap	Inflation	Fed Funds	Nom. Bond Yield
Mean	0.72	4.06	5.16	5.54
Std	(2.95)	(2.65)	(2.34)	(1.49)
AR(1) Coefficient	0.95	0.88	0.90	0.94
	(0.04)	(0.06)	(0.05)	(0.04)
AR(4) Coefficient	0.64	0.73	0.58	0.86
	(0.10)	(0.08)	(0.10)	(0.07)

1977.Q2-2000.Q4	Output Gap	Inflation	Fed Funds	Nom. Bond Yield
Mean	-1.20	3.72	7.80	8.14
Std	(2.30)	(2.41)	(3.44)	(2.49)
AR(1) Coefficient	0.95	0.89	0.85	0.95
	(0.04)	(0.05)	(0.05)	(0.03)
AR(4) Coefficient	0.65	0.85	0.73	0.80
	(0.09)	(0.05)	(0.07)	(0.06)

2001.Q1-2011.Q4	Output Gap	Inflation	Fed Funds	Nom. Bond Yield
Mean	-2.20	2.20	2.09	3.26
Std	(3.11)	(1.08)	(1.85)	(1.18)
AR(1) Coefficient	0.97	0.50	0.92	0.88
	(0.03)	(0.14)	(0.04)	(0.08)
AR(4) Coefficient	0.80	0.23	0.54	0.64
	(0.10)	(0.15)	(0.10)	(0.11)

Full sample and subperiod summary statistics. US quarterly log output gap (%), GDP deflator inflation (%, Annualized), Fed Funds rate (%, Annualized) averaged over last week of the quarter, and 5-year nominal yield (%, Annualized). Yields and inflation continuously compounded. Standard errors in parentheses. AR(4) coefficient refers to slope coefficient of variable onto its own 4-quarter lag.

Table 2: Predicting Stock Returns with Output Gap

Log Exc. Stock Ret. xr^e_{t+1}	60.Q2-11.Q4	60.Q2-77.Q1	77.Q2-00.Q4	01.Q1-11.Q4
Output Gap x_t	-0.49*	-0.66*	-0.27	-0.71
	(0.20)	(0.31)	(0.39)	(0.43)
Constant	0.63	0.98	1.42	-1.27
	(0.61)	(1.07)	(0.82)	(1.74)
R^2	0.02	0.02	0.04	0.01

Quarterly realized log excess stock returns (%, Quarterly) from quarter t to quarter $t+1$ onto the output gap (%) in quarter t. Newey-West standard errors with 2 lags in parentheses. * and ** denote significance at the 1% and 5% levels. The table reports averages across subperiods weighted by empirical sample length.

Table 3: Estimating the Monetary Policy Function

Fed Funds i_t	60.Q2-77.Q1	77.Q2-00.Q4	01.Q1-11.Q4
Output Gap x_t	0.17**	0.03	0.04
	(0.05)	(0.07)	(0.03)
Inflation π_t	0.21**	0.41*	0.21**
	(0.07)	(0.17)	(0.07)
Lagged Fed Funds i_{t-1}	0.69**	0.66**	0.83**
	(0.05)	(0.14)	(0.08)
Constant	0.67**	1.17*	-0.12
	(0.22)	(0.56)	(0.22)
R^2	0.86	0.77	0.94
Naïve Implied $\hat{\gamma}^x$	0.54**	0.09	0.22
	(0.17)	(0.23)	(0.12)
Naïve Implied $\hat{\gamma}^\pi$	0.66**	1.21**	1.19**
	(0.17)	(0.21)	(0.68)
Naïve Implied $\hat{\rho}^i$	0.69**	0.66**	0.83**
	(0.05)	(0.14)	(0.08)

We estimate $i_t = c^0 + c^x x_t + c^\pi \pi_t + c^i i_{t-1} + \epsilon_t$. All variables are described in Table 1. Naïve implied monetary policy parameters are calculated according to $\hat{\rho}^i = \hat{c}^i$, $\hat{\gamma}^x = \hat{c}^x/(1-\hat{c}^i)$, and $\hat{\gamma}^\pi = \hat{c}^\pi/(1-\hat{c}^i)$. Newey-West standard errors with 6 lags in parentheses. Standard errors for $\hat{\gamma}^x$ and $\hat{\gamma}^\pi$ are calculated by the delta method. * and ** denote significance at the 5% and 1% levels. Significance levels for implied parameters are based on an ordinary least squares likelihood ratio test.

Table 4: Parameter Choices

Panel A: Calibration Parameters

Time-Invariant Parameters				
Smoothing Parameter Consumption	ϕ		0.94	
Scaling Parameter Consumption	τ		0.89	
Consumption Growth Rate	g		1.89	
Leverage	δ		3.39	
Persistence Surplus Cons.	θ_0		0.97	
Dependence Output Gap	θ_1		0.02	
Dependence Lagged Output Gap	θ_2		0.02	
Utility Curvature	γ		3.00	
Steady-State Riskfree Rate	$\bar{r}$		0.94	
PC Lag Coefficient	ρ^{π}		0.96	
Phillips Curve Slope	κ		0.04	
Monetary Policy Rule		60.Q2-77.Q1	77.Q2-00.Q4	01.Q1-11.Q4
Output Weight	γ^{x}	0.33	0.28	0.84
Inflation Weight	γ^{π}	0.60	1.61	1.60
Persistence MP	ρ^{i}	0.60	0.64	0.82
Std. Shocks				
Std. PC		0.80	0.35	0.27
Std. MP		0.77	1.56	0.61
Std. Infl. Target		0.10	0.11	0.40

Panel B: Implied Parameters

		60.Q2-77.Q1	77.Q2-00.Q4	01.Q1-11.Q4
Discount Rate	β	0.85	0.85	0.85
IS Curve lag Coefficient	ρ^{x-}	0.02	0.02	0.02
IS Curve Forward Coefficient	ρ^{x+}	1.10	1.10	1.10
IS Curve Real Rate Slope	ψ	0.41	0.41	0.41
Steady-State Surplus Cons. Ratio	$\bar{S}$	0.10	0.08	0.08
Log Max. Surplus Cons. Ratio	s_{max}	-1.80	-1.98	-2.04

Table 5: Model and Empirical Moments

Panel A: Estimated MP Rule – Fed Funds onto Output Gap, Infl. and Lag. Fed Funds

	60.Q2-77.Q1		77.Q2-00.Q4		01.Q1-11.Q4	
	Empirical	Model	Empirical	Model	Empirical	Model
Output Gap	0.17**	0.15	0.03	-0.10	0.04	0.06
Inflation	0.21**	0.29	0.41*	0.36	0.21**	0.16
Lagged Fed Funds	0.69**	0.71	0.66*	0.63	0.83**	0.84

Panel B: Subperiod Second Moments

	60.Q2-77.Q1		77.Q2-00.Q4		01.Q1-11.Q4	
Std. Asset Returns	Empirical	Model	Empirical	Model	Empirical	Model
Std. Eq. Ret.	18.35	16.19	15.68	16.49	20.34	15.48
Std. Nom. Bond Ret.	4.92	1.38	8.11	2.65	5.92	3.58
Nominal Bond Beta	0.07**	0.05	0.12	0.12	-0.18**	-0.16
Std. Real Bond Ret.		5.06		2.13	4.27	2.20
Real Bond Beta		-0.29		0.05	-0.08	-0.07
Std. VAR(1) Residuals						
Output Gap	0.83	1.23	0.76	1.05	0.67	0.96
Inflation	1.05	0.78	1.04	0.32	0.86	0.26
Fed Funds Rate	0.90	0.75	1.55	1.51	0.47	0.58
Log Nominal Yield	0.47	0.14	0.77	0.27	0.56	0.38

This table reports average model moments from 2 simulations of length 50000. * and ** denote significance at the 5% and 1% levels. We use Newey-West standard errors with 2 lags for the nominal bond beta and Newey-West standard errors with 6 lags for the empirical Taylor rule estimation.

Table 6: Consumption and Risk Premia

Consumption and Output Gap	Empirical	Model
Std. Cons. Innovation	1.53	1.94
AR(1) Coefficient Output Gap	0.96	0.92
Twelve Quarter Cons. Variance Ratio		0.95
Equities		
Equity Premium	4.50	8.37
Price-Dividend Ratio (exp(mean(p-d))	30.05	34.85
Std(d-p)	0.40	0.21
AR(4) Coefficient d-p	0.92	0.89
Correlation(x,p-d)	0.54	0.80
Slope 1 YR Exc. Stock ret. wrt d-p	0.08	0.19
Slope 5 YR Exc. Stock ret. wrt d-p	0.25	0.73
Slope quarterly stock ret. wrt x	-0.49	-0.32
Slope quarterly annualized std. stock ret. wrt x	-1.03	-0.61
Nominal 5-Year Zero Coupon Bond		
Nominal Bond Excess Return	1.60	0.23
Slope Term Structure	1.06	0.28
Subperiod Average Bond Excess Return onto Term Spread	2.88	2.14
Correlation(Fed Funds, Slope Term Structure)	-0.44	-0.59
Std(real rate) (Percent, Ann.)		1.83

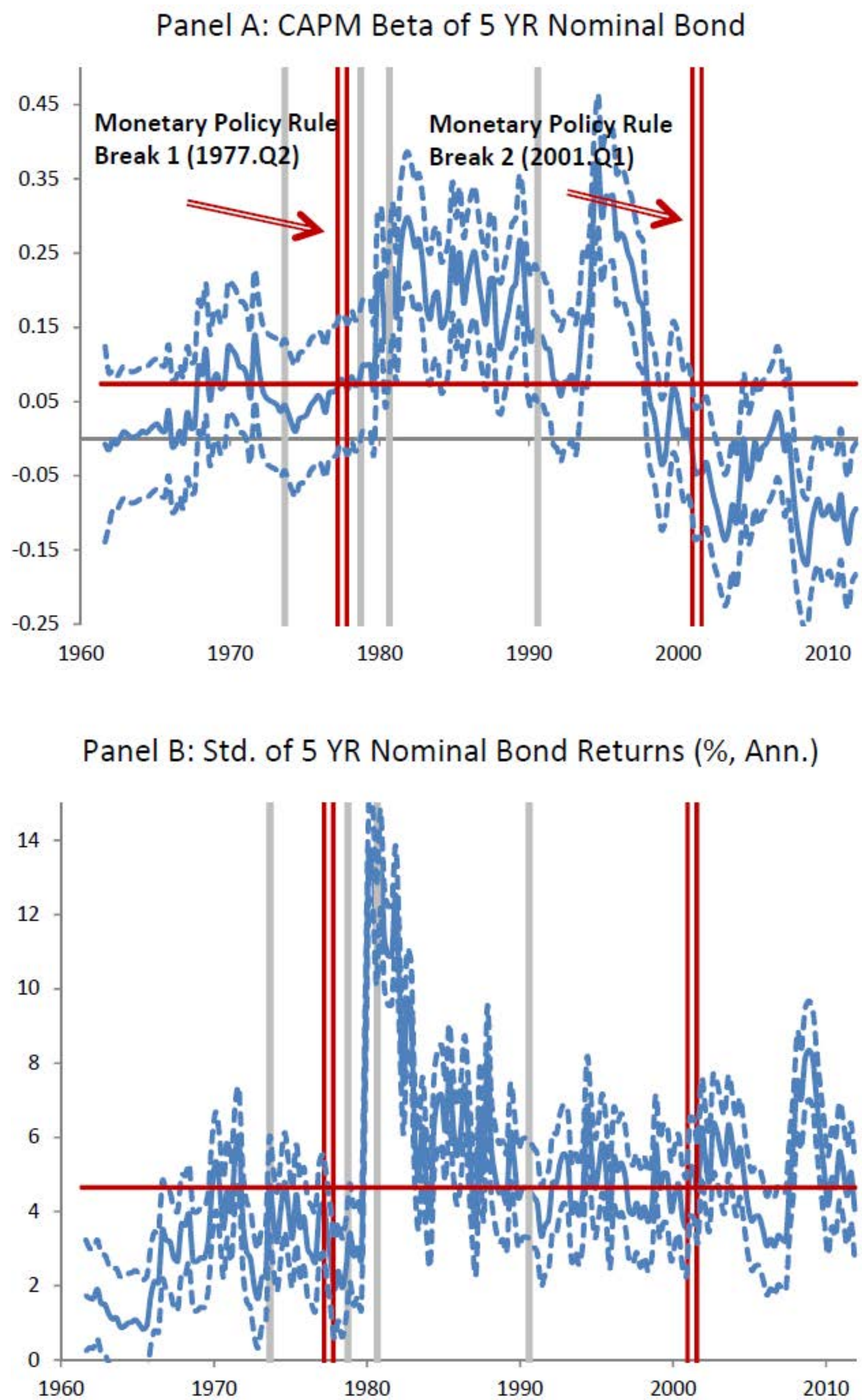
The model consumption and output gap dynamics are computed analytically. All other model moments are simulated as described in Table 5. Model moments report weighted averages across subperiods, where weights are proportional to the length of the empirical period. Empirical moments are for the full sample. We proxy for the quarterly annualized standard deviation of stock returns with two times the absolute one-quarter stock return in percent. The subperiod average bond excess return onto the term spread is compared to a quarterly Campbell and Shiller (1991) regression for the full sample.

Table 7: Risk Premium Decomposition

Nominal 5-Year Bond-Stock Correlation	60.Q2-77.Q1	77.Q2-00.Q4	01.Q1-11.Q4
Risk Neutral	0.41	0.70	-0.57
Risk Premium	0.83	0.79	-0.77
Combined	0.54	0.75	-0.71
Real 5-Year Bond-Stock Correlation			
Risk Neutral	-0.84	0.41	-0.28
Risk Premium	-0.77	0.43	-0.30
Combined	-0.93	0.41	-0.50

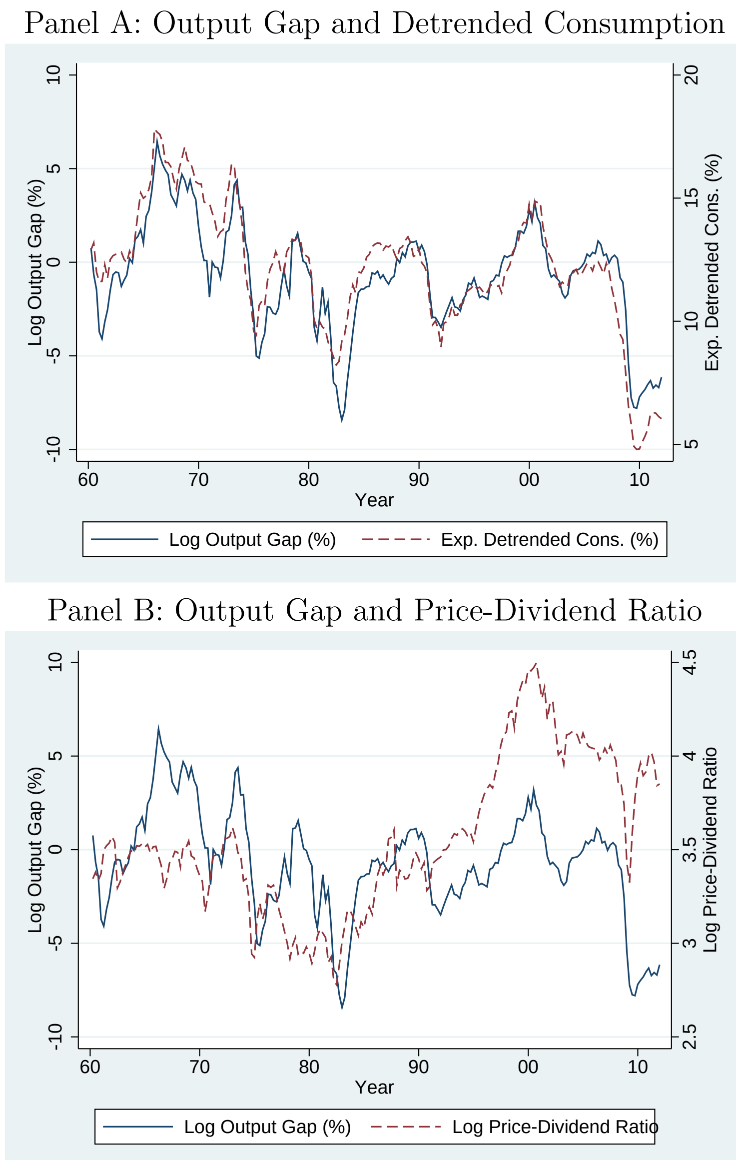
Combined returns are the sum of risk-neutral and risk premium components. We compute risk neutral nominal and real bond returns as innovations to expected interest rates (Campbell and Ammer 1993). The risk premium component of nominal and real bond returns corresponds to innovations in expected excess returns. Risk neutral stock returns are the sum of innovations to dividends and real interest rates in a Campbell and Shiller (1988) approximate loglinear decomposition. The risk premium component corresponds to innovations in expected excess returns. The loglinearization constant is chosen to match simulated subperiod average equity price-dividend ratios for combined equity returns for each subperiod.

Figure 1: US Nominal Bond Beta, Bond Volatility, and Monetary Policy Break Dates



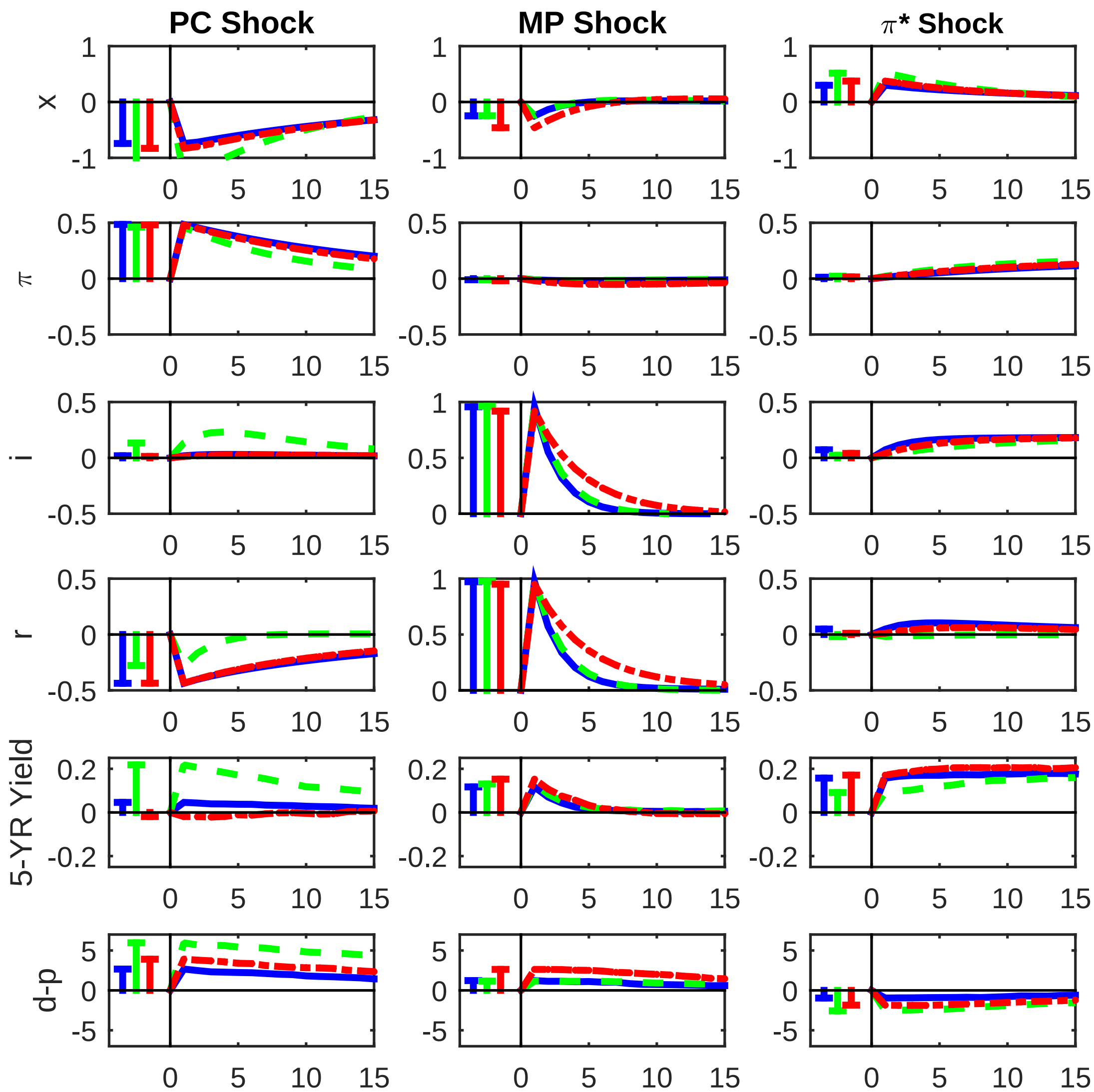
Nominal bond beta and standard deviation of nominal bond returns from daily bond and stock returns over past three months. The unconditional mean is shown with a red horizontal line. We model time-varying second moments as an unobserved trend AR(1) component plus white measurement noise. We show trend second moments estimated using the Kalman filter. 95% confidence intervals, which do not take into account parameter uncertainty, are shown in dashed. Gray vertical lines depict Hamilton (2009) oil price shocks.

Figure 2: US Output Gap, Detrended Consumption, and Equity Price-Dividend Ratio



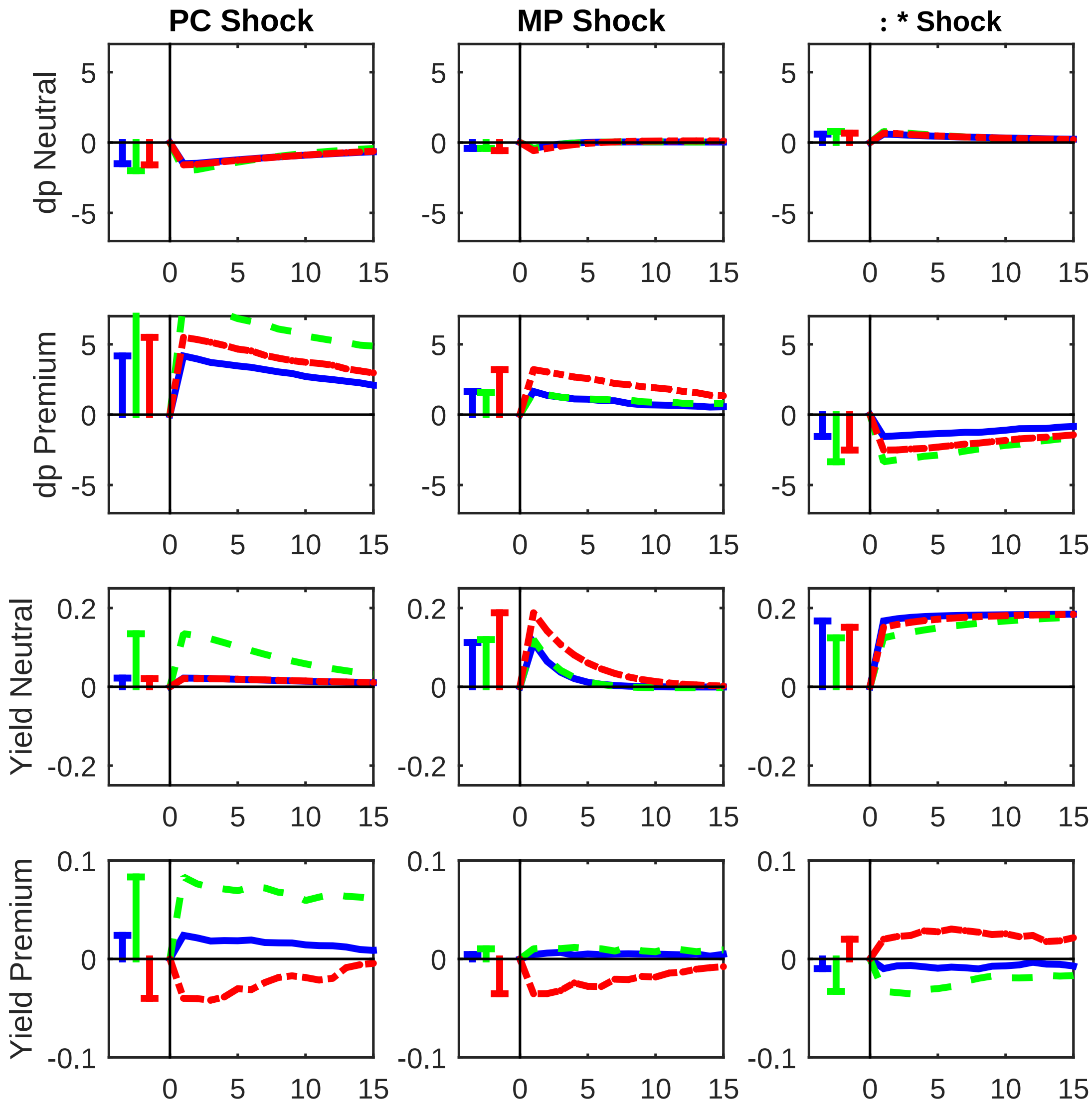
Panel A plots the time series of the US log real output gap together with log real consumption in excess of its exponential moving average. An annualized smoothing parameter of $\phi = 0.94$ corresponds to a half-life of 2.6 years. We use real consumption expenditures data for nondurables and services from the Bureau of Economic Analysis National Income and Product Accounts Tables. The US log output gap (%) is described in Table 1. The end-of-quarter price dividend ratio is computed as the S&P 500 real price divided by real dividends averaged over the past 10 years.

Figure 3: Impulse Response Functions



This figure shows average simulated impulses for the output gap, inflation, the nominal and real short rates, the 5-year nominal yield, and the log dividend price ratio. We show impulse responses for the subperiods 1960.Q2-1977.Q1 (blue solid line), 1977.Q2-2000.Q4 (green dashed line), and 2001.Q1-2011.Q4 (red dash-dot line). Vertical bars indicate the magnitude of the initial response for each variable for period 1 (blue, left), period 2 (middle, green), and period 3 (red, right). This figure shows impulse responses to the same size shocks for all three subperiods. The shock size equals the weighted average subsample standard deviation. Impulse responses for the nominal bond yield and dividend yield are averaged over 2000 simulations, starting from the unconditional steady-state at time 0. The output gap and the dividend price ratios are in percent deviations from the steady state. All other variables are in annualized percent.

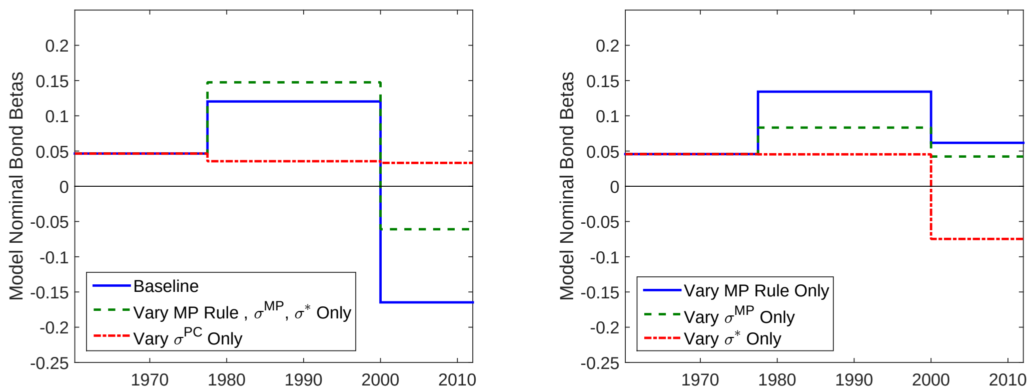
Figure 4: Risk Neutral and Risk Premium Impulse Responses



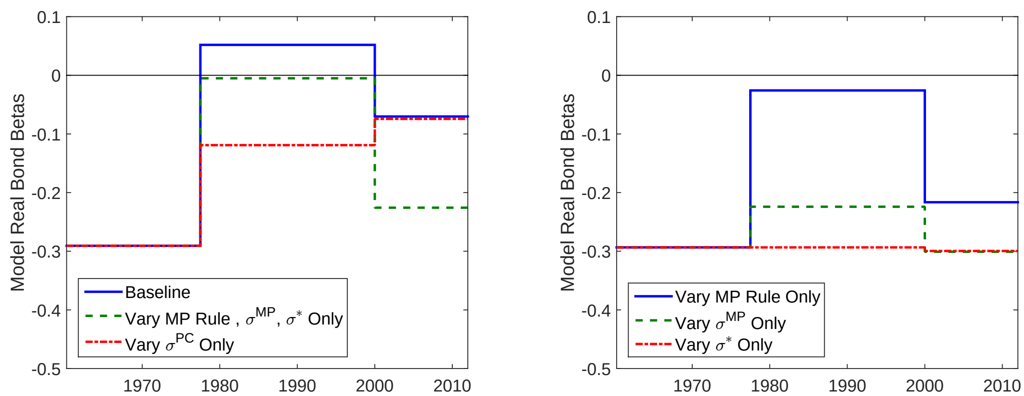
This figure shows average simulated impulses for risk neutral and risk premium components of 5-year nominal bond yields ('Yield') and the log equity dividend-price ratio ('dp'). Risk neutral and risk premium components add up to bond yield and dividend yield responses shown in Figure 3. Risk neutral bond yields are the average of expected future nominal short-term interest rates. Risk neutral equity dividend yields are from a Campbell and Shiller (1988) approximate loglinear decomposition and correspond to the sum of expected future real short-term interest rates less expected future dividend growth. We show impulse responses for the subperiods 1960.Q2-1977.Q1 (blue solid line), 1977.Q2-2000.Q4 (green dashed line), and 2001.Q1-2011.Q4 (red dash-dot line). Impulse responses are simulated exactly as in Figure 3.

Figure 5: Counterfactual Changes in Nominal and Real Bond Betas

Panel A: Nominal 5-Year Bonds



Panel A: Real 5-Year Bonds



This figure shows model-implied bond betas for three subperiods. The left panels decompose the total model-implied bond beta (baseline) into changes in monetary policy and changes in PC shock volatilities. The right panels decompose changes in bond betas due to monetary policy into different monetary policy components. All parameters not listed in the corresponding legend are held constant at their period 1 values.