Monetary Policy Drivers of Bond and Equity Risks

John Y. Campbell, Carolin Pflueger, and Luis M. Viceira

First draft: March 2012
This draft: April 2015

1Campbell: Department of Economics, Littauer Center, Harvard University, Cambridge MA 02138, USA, and NBER. Email john_campbell@harvard.edu. Pflueger: University of British Columbia, Vancouver BC V6T 1Z2, Canada. Email carolin.pflueger@sauder.ubc.ca. Viceira: Harvard Business School, Boston MA 02163 and NBER. Email lviceira@hbs.edu. We are grateful to Alberto Alesina, Yakov Amihud, Robert Barro, Philip Bond, Mikhail Chernov, Paul Beaudry, Ian Dew-Becker, Alexander David, Adlai Fisher, Ben Friedman, Lorenzo Garlappi, Joao Gomes, Gita Gopinath, Robin Greenwood, Joshua Gottlieb, Howard Kung, Leonid Kogan, Deborah Lucas, Greg Mankiw, Harald Uhlig, Michael Woodford, conference and seminar participants at the University of British Columbia, the Harvard Monetary Economics Seminar, the 2013 HBS Finance Research Retreat, the University of Calgary, the University of Miami, MIT Sloan, the Vienna Graduate School of Finance, ECWFC 2013, PNWCF 2014, the Jackson Hole Finance Conference 2014, the ASU Sonoran Winter Finance Conference 2014, the Duke/UNC Asset Pricing Workshop, the Monetary Policy and Financial Markets Conference at the Federal Reserve Bank of San Francisco, the Adam Smith Asset Pricing Workshop, NYU Stern School, the Federal Reserve Bank of New York, the Bank of Canada, the SFS Cavalcade 2014, and especially our discussants Jules van Binsbergen, Olivier Coibion, Gregory Duffee, Martin Lettau, Francisco Palomino, Monika Piazzesi, Rossen Valkanov, and Stanley Zin for helpful comments and suggestions. We thank Jiri Knesl for able research assistance. This material is based upon work supported by Harvard Business School Research Funding and the PH&N Centre for Financial Research at UBC.
Abstract

The exposure of US Treasury bonds to the stock market has moved considerably over time. While it was slightly positive on average in 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 explores the effects of monetary policy rules, monetary policy uncertainty, and macroeconomic shocks on nominal bond risks, using a New Keynesian model with habit formation and discrete regime shifts in 1977 and 2000. The increase in bond risks after 1977 is attributed primarily to a shift in monetary policy towards a more anti-inflationary stance, while the more recent decrease in bond risks after 2000 is attributed to a renewed focus on output stabilization combined with decreased volatility of supply shocks and increased volatility of the Fed’s long-run inflation target. Endogenous responses of bond risk premia amplify these effects of monetary policy on bond risks.
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 particularly 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.

In this paper we ask what macroeconomic forces determine the risk properties of US Treasury bonds, and particularly their changes over time. One common approach to this question uses 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 economic 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 limited role for monetary policy, which determines inflation (at least in the long run) but has no influence on the real economy.\(^2\)

Accordingly a recent literature has explored the asset pricing implications of New Keynesian models, in which price stickiness allows monetary policy to have real effects. Recent papers in this literature include Andreasen (2012), Bekaert, Cho, and Moreno (2010), Van Binsbergen et al (2012), Dew-Becker (2014), Kung (2015), Li and Palomino (2014), Palomino (2012), Rudebusch and Wu (2008), and Rudebusch and Swanson (2012).

\(^2\)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.
We follow this second approach and quantitatively investigate two candidate explanations for the empirical instability in bonds’ risk properties: changes in monetary policy or changes in macroeconomic shocks. Importantly, in order to account for general equilibrium effects of monetary policy on bond and equity risks, we allow monetary policy to alter how consumption, output and inflation respond to shocks.

Both US monetary policy and the magnitude of macroeconomic shocks have changed substantially over our sample 1960-2011. 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 1970s, while the second one covers the inflation-fighting period under Federal Reserve Board chairmen Paul Volcker and Alan Greenspan. The third, newly identified, regime is characterized by renewed attention to output stabilization and increased central bank gradualism after the turn of the millennium. If central bank policy acts on 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 and 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 and therefore Phillips curve shocks give rise to positive nominal bond betas. In contrast, shocks to the perceived central bank long-term inflation target reduce the beta of nominal bonds. A downward drift in the perceived inflation target induces firms with nominal rigidities to reduce output. Consequently,
a negative inflation 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, but no information from asset markets. The statistically determined dates 1977Q2 and 2001Q1 line up remarkably closely with important institutional 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. Next, we turn to changes in bond risks over the same sample.

The 10-year nominal bond CAPM betas and return volatilities plotted in Figure 1 illustrate important changes in the risks of nominal bonds over time. 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 also higher-frequency variation, 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.

Footnotes:
3 For details on the econometric procedure, see Section 3.
4 We show filtered CAPM betas and standard deviations of daily returns on a benchmark 10-year nominal bond over a rolling 3-month window, together with 95% confidence intervals.
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 risks of nominal bonds.

This paper builds on the New Keynesian asset pricing literature and makes two contributions. First, we formulate a New Keynesian model in which bonds and stocks can both be priced from assumptions about their payoffs, and in which time-varying risk premia, driven by habit formation, generate realistic variances and covariances for these asset classes. Most previous New Keynesian asset pricing papers have concentrated on the term structure of interest rates, and have paid little attention to the implied pricing of equities. This contrasts with the integrated treatment of the bond and stock markets in several papers that use reduced-form affine or real business cycle models (Ang and Ulrich, 2012, Bansal and Shaliastovich 2013, Bekker, 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).

Second, we use our 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 the regime shift that we newly identify in the early 2000s. 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.

The organization of the paper is as follows. Section 2 lays out a standard New Keynesian model that explains interest rates, inflation, and medium-term deviations of output from trend (the “output gap”) using three structural equations: an investment-saving curve (IS) that describes real equilibrium in the goods market based on the Euler equation of a representative consumer, a Phillips curve (PC) that describes the effects of nominal frictions on inflation, and a monetary policy reaction function (MP) embodying a Taylor rule as in Clarida, Gali, and Gertler (1999), Taylor (1993), and Woodford (2001). This section also derives the New Keynesian IS curve from preferences with slowly moving habit, building on Campbell and Cochrane (1999).

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 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 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. 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
2 A New Keynesian Asset Pricing Model

Our model integrates a standard three-equation New Keynesian macroeconomic model with a habit-formation model of asset prices. On the macroeconomic side, we use a standard New Keynesian specification, taking the Phillips curve as already log-linearized. On the asset pricing side, we extend the habit formation preferences of Campbell and Cochrane (1999), which we do not log-linearize. 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 equivalent of the Investment and 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, Canova and Sala 2009).\(^5\) 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

\(^5\)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, where risk premia vary over time. 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.
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, real rate dynamics are described by a standard 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 defined 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. However, consumption is typically thought to have a unit root, while the output gap has stationary dynamics. We therefore model the output gap as the difference between current period consumption and an exponentially-weighted moving average of lagged consumption.

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 half life 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%.

The second building block of a New Keynesian model is the Phillips curve (PC) equation that links inflation and real output in equilibrium. We do not take a stand on the precise
source of nominal rigidities in the economy and instead directly assume a PC with both forward- and backward-looking components. While a Calvo (1983) model of monopolistically competitive firms and staggered price setting implies a forward-looking Phillips curve, a backward-looking Phillips curve can arise when price setters update their information infrequently (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. Empirically, 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).

We allow for shocks to the central bank’s long-run inflation target. These shocks can temporarily boost demand in our model. We interpret them broadly as capturing explicit and perceived changes in the long-run inflation target. As such, movements in the inflation target may capture changing public expectations of central bank behavior, or changing central bank credibility, even if the central bank’s true target is relatively stable (Orphanides and Williams 2004).

Inflation target shocks have permanent effects because we model the target as a unit root process. The unit root assumption is consistent with the extremely high persistence in
U.S. inflation data (Ball and Cecchetti 1990, Stock and Watson 2007). We choose the unit root specification rather than a highly persistent mean-reverting inflation target for several reasons. First, the inflation target reflects consumers’ long-run inflation expectations, whose changes cannot be anticipated. A mean-reverting inflation target implies counterintuitive predictability of changes in the target. Second, a highly persistent inflation target may lead to equilibrium existence and uniqueness issues. Third, when the inflation target is mean-reverting and an equilibrium exists, we may obtain counterintuitive impulse responses for inflation. Inflation can respond to a positive inflation target shock as it would to a positive monetary policy shock by declining, rather than converging to the new, higher, target inflation level. Finally, while our unit root assumption means that the unconditional variance of nominal interest rates in the model is not defined, even with a highly persistent inflation target this unconditional variance would be extremely sensitive to the imprecisely identified persistence parameter.

To close the model we need to make identification assumptions. Dynamic stochastic general equilibrium (DSGE) models are often under-identified or only very weakly identified (Canova and Sala 2009, An and Schorfheide 2007) because the mapping between underlying parameters and model moments can be highly nonlinear. Restrictions on the form of the monetary policy shock may be necessary to identify monetary policy parameters (Backus, Chernov, and Zin 2013). We adopt identification assumptions commonly used in the structural vector autoregression literature to help identify the central bank’s monetary policy rule, using exclusion restrictions that allow us to estimate the monetary policy rule by Ordinary Least Squares (OLS).

Following CGG, we assume that transitions from one regime to another are structural
breaks, completely unanticipated by investors. While we recognize the importance of allowing agents to anticipate potential future changes in policy and to optimize according to such expectations, we think that a parsimonious model like ours still brings substantive insights which are likely to survive in a more sophisticated but less analytically tractable model. We focus on unanticipated regime changes as an interesting approximation for two reasons. First, quarterly transition probabilities would 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 (2013) 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 (2013).

2.1 Euler equation with habit formation

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

$$1 = E_t [M_{t+1} (1 + R_{t+1})],$$

(1)

where $M_{t+1}$ is the stochastic discount factor (SDF). Household optimization implies a SDF of the form

$$M_{t+1} = \frac{\beta U'_{t+1}}{U'_t},$$

(2)
where $U'_t$ is the marginal utility of consumption at time $t$ and $\beta$ is a time discount factor. Substitution of (2) into (1) produces the standard Euler equation.

The Euler equation for the return on a one-period real T-bill can be written in log form as:

$$\ln U'_t = rt + \ln \beta + \ln E_t U'_{t+1},$$

where we write $r_t$ for the log yield at time $t$—and return at time $t + 1$—on a one-period real Treasury bill. Similarly, we write $i_t$ to denote the log yield on a one-period nominal T-bill. We use the subscript $t$ for short-term nominal and real interest rates to emphasize that they are known at time $t$.

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 $\pi_{t+1}$ 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 (3), and dropping constants to reduce the notational burden, we have:

$$\ln U'_t = (i_t - E_t \pi_{t+1}) + \ln E_t U'_{t+1}.$$  

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 $H$:

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

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

Marginal utility in this model is

$$U_t' = (C_t - H_t)^{-\gamma} = (S_tC_t)^{-\gamma},$$

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

### 2.1.1 Modeling Consumption

We assume that consumption is driven by the same shock as the output gap. However, in contrast to the output gap, consumption has a unit root. The other two New Keynesian state variables, inflation and the Federal Funds rate, are relevant for consumption only to the extent that they are correlated with the output gap.

As a leading example, consider the case where consumption is a random walk and the output gap is an AR(1), but they share the same innovation. Then

$$c_t = c_{t-1} + \varepsilon_t,$$  
$$x_t = \phi x_{t-1} + \varepsilon_t.$$  

It follows that the output gap is the difference between current consumption and an exponentially-weighted stochastic moving average. Equivalently, consumption can be written as a weighted
sum of the current output gap and all lagged output gaps.

\[
x_t = c_t - (1 - \phi)[c_{t-1} + \phi c_{t-2} + ...],
\]

(9)

\[
c_t = x_t + (1 - \phi)[x_{t-1} + x_{t-2} + ...].
\]

(10)

We find new and striking support in U.S. data for a relation of the form (9). We choose a quarterly smoothing parameter of \( \phi = 0.94 \), corresponding to a half-life of 2.6 years. For our choice of \( \phi \), the log output gap from the Congressional Budget Office and stochastically detrended consumption are 90% correlated over our sample period. Panel A of Figure 2 shows the time series of stochastically detrended real consumption and the log output gap. Importantly, this relation does not appear to be driven by any particular subsample, suggesting that the relation between consumption and the output gap is stable across different monetary policy and shock regimes.

Relation (10) is useful for illustrative purposes. We use a slightly more general form for the full model, allowing for different volatilities of consumption and output gap shocks. For a constant scaling parameter \( \tau \) and growth rate of consumption \( g \), we model consumption as follows:

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

(11)

In the full model, the output gap does not follow a simple AR(1) process. However, any stationary stochastic process for the output gap defines a consumption process with a unit root through (11).
2.1.2 Modeling the Surplus Consumption Ratio

We choose a functional form for marginal utility to satisfy the following two features. First, we require log surplus consumption to be stationary, as in Campbell and Cochrane (1999). Second, the real short-term interest rate is determined by a standard New Keynesian Euler equation with terms in the contemporaneous, lagged, and expected output gaps.

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

\begin{align}
    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}, \\
    \varepsilon_{c,t+1} &= c_{t+1} - E_t c_{t+1} = \tau (x_{t+1} - E_t x_{t+1}).
\end{align}

The parameter $\theta_0$ determines the persistence of the log surplus consumption ratio. When $\theta_1 = \theta_2 = 0$, equation (12) nests the model of Campbell and Cochrane (1999). When habit reacts slowly to the overall macroeconomic environment, it is plausible that both current and lagged output gaps should enter into utility and raise surplus consumption. We calibrate the parameters $\theta_1$ and $\theta_2$ to small but positive numbers. Since $x_t$ has an unconditional mean of zero, $\bar{s}$ gives the steady state surplus consumption ratio.\(^6\)

Including current and lagged output gaps in the surplus consumption dynamics allows us to reverse engineer an Euler equation with both forward- and backward-looking terms.

\(^6\)If $\theta_1$ and $\theta_2$ are different from zero, there is the theoretical possibility that the log surplus consumption ratio exceeds the maximal value $s_{\text{max}}$, 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.
Substituting (12) into the Euler equation for the one-period real T-bill gives

\[ r_t = -\ln \beta + \gamma g + \gamma \theta_2 x_{t-1} + \gamma (\theta_1 - \tau \phi) x_t + \gamma \tau E_t x_{t+1} \]
\[ -\gamma (1 - \theta_0)(s_t - \bar{s}) - \frac{\gamma^2 \sigma^2_c}{2} (1 + \lambda(s_t))^2. \]  

(14)

Now, we use Campbell and Cochrane (1999)’s condition that the terms in (14) 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 \( \lambda \) are given by

\[ \bar{S} = \sigma_c \sqrt{\frac{\gamma}{1 - \theta_0}}, \]  
\[ \bar{s} = \log(\bar{S}), \]  
\[ s_{max} = \bar{s} + 0.5(1 - \bar{S}^2), \]  
\[ \lambda(s_t, \bar{S}) = \begin{cases} \frac{1}{\bar{s}} \sqrt{1 - 2(s_t - \bar{s})} - 1, & s_t \leq s_{max} \\ 0, & s_t \geq s_{max} \end{cases}. \]  

(15)  
(16)  
(17)  
(18)

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:

\[ x_t = \frac{\tau}{\tau \phi - \theta_1} E_t x_{t+1} + \frac{\theta_2}{\tau \phi - \theta_1} x_{t-1} - \frac{1}{\gamma (\tau \phi - \theta_1)} r_t. \]  

(19)

Several points are worth noting about the IS curve (19). 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 $\psi$ does not equal the elasticity of intertemporal substitution (EIS) of the representative consumer.

The lag coefficient $\rho$ in the IS curve (19) is non-zero whenever the lagged output gap enters into surplus consumption (i.e. $\theta_2 \neq 0$). Moreover, 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.

### 2.2 Macroeconomic dynamics

We complement the consumers’ Euler equation with standard building blocks of New Keynesian macroeconomic models. We assume that consumers and price-setting firms do not incorporate contemporaneous monetary policy shocks into their time $t$ decisions, similarly to Christiano, Eichenbaum and Evans (2005). Instead, they form their time $t$ expectations based on monetary policy and inflation target shocks up to time $t-1$ and PC and inflation target shocks up to time $t$. We denote the expectation with respect to this information set by:

$$
E_t(\cdot) = E(\cdot | u^P_{t}, u^P_{t-1}, u^P_{t-2}, ..., u^P_{t-1}, u^P_{t-2}, u^P_{t-1}, \ldots).
$$

(20)

The assumption that consumers and firms make decisions based on $E_t(\cdot)$ expectations implies that monetary policy shocks do not affect macroeconomic aggregates contemporaneously, but only with a lag. This identification assumption is common in the structural VAR literature (Christiano, Eichenbaum, and Evans, 1999) and it is helpful for our empirical strategy in that in the absence of inflation target shocks we could estimate the monetary policy Taylor rule by OLS. In the full model with inflation target shocks, the link between
OLS Taylor rule estimates and monetary policy parameters is no longer one-to-one, but we are still able to choose monetary policy parameters to match the empirical Taylor rule OLS estimates. Note that we do not assume the same lags for financial markets; we allow asset prices to react to monetary policy shocks contemporaneously.

The dynamics of the output gap, inflation, and Fed Funds rate are summarized by the following linearized system of equations:

\[
x_t = \rho^x x_{t-1} + \rho^x E_{t-1} x_{t+1} - \psi (E_{t-1} i_t - E_{t-1} \pi_{t+1}), \\
\pi_t = \rho^\pi \pi_{t-1} + (1 - \rho^\pi) E_{t-1} \pi_{t+1} + \lambda x_t + u_t^{PC}, \\
i_t = \rho^i (i_{t-1} - \pi^*_t) + (1 - \rho^i) [\gamma^x x_t + \gamma^\pi(\pi_t - \pi^*_t)] + \pi^*_t + u_t^{MP}, \\
\pi^*_t = \pi^*_{t-1} + u^*_t.
\]  

Equation (21) is the IS curve (19) with the expectational timing assumption (20).

Equation (22) is a standard New Keynesian equation that determines inflation from the price-setting behavior of firms. It has parameters \(\rho^\pi\), determining the relative weight on past inflation and expected future inflation, and \(\lambda\), governing the sensitivity of inflation to the output gap.

Equations (23) and (24) describe monetary policy. Equation (23) is a central bank reaction function along the lines of Clarida, Gali, and Gertler (1999), Taylor (1993), and Woodford (2001). It determines the short-term nominal interest rate with parameters \(\rho^i\), controlling the influence of past interest rates on current interest rates, \(\gamma^x\), governing the reaction of the interest rate to the output gap, and \(\gamma^\pi\), governing the response of the interest rate to inflation relative to its target level \(\pi^*_t\). Equation (24) specifies that the central bank’s
inflation 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 inflation 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

$$u_t = [u_t^{PC}, u_t^{MP}, u_t^∗]'$$

is independently and conditionally normally distributed with mean zero and variance-covariance matrix:

$$E_{t-1} [u_t u_t'] = \Sigma_u = \begin{bmatrix} (\sigma^{PC})^2 & 0 & 0 \\ 0 & (\sigma^{MP})^2 & 0 \\ 0 & 0 & (\sigma^*)^2 \end{bmatrix}.$$  \hspace{1cm} (26)

Equation (26) has two important properties. First, for parsimony we assume that all the shocks in the model are uncorrelated with each other. 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. Second, the variances of all shocks in the model, not just the shock to the Euler equation, are conditionally homoskedastic. The previous version of this paper generated countercyclical risk premia from countercyclical volatility of shocks. However, strong countercyclical heteroskedasticity is not a feature of macroeconomic data.\footnote{We thank our discussants Jules van Binsbergen, Martin Lettau, and Monika Piazzesi for emphasizing...}
in this model generate countercyclical risk premia and heteroskedasticity in asset returns even without heteroskedasticity of macroeconomic fundamentals.

2.3 Modeling bonds and stocks

In our model, risk premia vary over time and the expectations hypothesis of the term structure of interest rates does not hold. 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. We also assume that asset prices react contemporaneously to all shocks in the model, including monetary policy and inflation target shocks.

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

\[
\Delta d_t = \delta \Delta c_t. \tag{27}
\]

We interpret \( \delta \) 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).

Ermolov (2015), in a paper circulated after the previous version of this paper, also uses changing fundamental volatility to generate changing risk premia.
2.4 Model solution and stability

We define the inflation and nominal interest rate gaps as:

\[ \hat{\pi}_t = \pi_t - \pi_t^*, \quad \hat{\mathring{i}}_t = \mathring{i}_t - \pi_t^*. \] (28) (29)

We solve for the dynamics of the vector of state variables

\[ \hat{Y}_t = [x_t, \hat{\pi}_t, \hat{\mathring{i}}_t]' . \] (30)

The state variable dynamics have a solution of the form

\[ \hat{Y}_t = P\hat{Y}_{t-1} + Qu_t . \] (31)

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 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 Phillips curve means that
there exist at most a finite number of dynamically stable equilibria of the form (31). This is true even when the monetary policy reaction to inflation ($\gamma^\pi$) 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 apply multiple equilibrium selection criteria proposed in the literature 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 into 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 is closely linked to a possibility that the regime will switch back to a stable equilibrium in the distant future. 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.$^8$ 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.

$^8$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.
2.5 Solutions for bond and stock returns

We solve for bond and stock prices numerically, extending the value function iteration methodology of Wachter (2005) to multiple state variables. Let \( \frac{P^{d}_t}{D_t} \) denote the price-dividend ratio of a zero coupon claim on the aggregate stock market dividend at time \( t + n \). The price of a zero coupon claim for the dividend at time \( t \) is given by \( \frac{P^{d}_t}{D_t} = 1 \). For \( n \geq 1 \), we solve for the \( n \)-period price-dividend ratio numerically using the iteration

\[
\frac{P^{d}_{nt}}{D_t} = E_t \left[ M_{t+1} \frac{D_{t+1}}{D_t} \frac{P^{d}_{n-1,t+1}}{D_{t+1}} \right].
\]

The price-dividend ratio on the aggregate stock market is then given as the infinite sum of zero-coupon price-dividend ratios.

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

\[
P^{s}_{1,t} = \exp(-\hat{i}_t - \pi^*_t - r^f), \quad (32)
\]
\[
P_{1,t} = \exp(-\hat{i}_t + E_t \hat{\pi}_{t+1} - r^f). \quad (33)
\]

For \( n > 1 \), the prices of real and nominal zero coupon bonds with maturity \( n \) follow the recursions:

\[
P_{n,t} = E_t \left[ M_{t+1} P_{n-1,t+1} \right], \quad (34)
\]
\[
P^{s}_{n,t} = E_t \left[ M_{t+1} \exp(-\pi_{t+1}) P^{s}_{n-1,t+1} \right]. \quad (35)
\]
In contrast to Campbell and Cochrane (1999) and Wachter (2005), prices for bonds and dividend claims in our model depend on more than one state variable. To simplify the numerical analysis, we define the scaled vector of state variables $\tilde{Y}_t$. $\tilde{Y}_t$ scales and rotates the variables in $\hat{Y}_t$ such that shocks to $\tilde{Y}_t$ are independent standard normal. Moreover, the first element of $\tilde{Y}_t$ is conditionally perfectly correlated with innovations to consumption, and hence the SDF. Since the inflation target follows a random walk, $n$-period zero coupon nominal bond prices are proportional to $\exp(-n\pi^*_t)$. We therefore solve numerically for scaled nominal bond prices $\exp(n\pi^*_t)P^s_{n,t}$.

We solve for equity price-dividend ratios and bond prices as functions of the surplus consumption ratio $s_t$, the lagged output gap $x_{t-1}$, and the vector $\hat{Y}_t$. The baseline solution has 50 gridpoints for $s_t$ spaced between $\log(-50)$ 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 four gridpoints for $x_{t-1}$ ranging over all output gap values covered by the grid for $\hat{Y}_t$. We therefore have a grid with a total of 1600 gridpoints. Solutions are almost unchanged if we increase the grid size for any of these dimensions, indicating that the baseline grid is sufficient.

Following Wachter (2005), we use Gauss-Legendre 40-point quadrature to integrate over consumption shocks. We find that Gauss-Legendre 10-point quadrature is sufficient for the other dimensions of $\hat{Y}_t$, which are contemporaneously uncorrelated with the SDF. We use five-dimensional loglinear interpolation to evaluate price-dividend ratios and bond prices between gridpoints.
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 raises inflation. The shock pushes both stock and nominal bond valuations lower, generating a positive nominal bond beta. But the positive bond beta implies that nominal bonds are risky. The increase in risk aversion at the beginning of a recession increases nominal bond risk premia, amplifying the effect of the shock on bond prices. We show in our calibration of the model that amplification mechanisms of this kind, operating 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 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
Our identification of a third regime for monetary policy is supported by several observations. First, the Federal Reserve appears to have followed a more gradual policy after 2000, as documented by Hamilton, Pruitt, and Borger (2011) using Fed Funds futures. Increased gradualism may have been the result of greater central bank transparency and credibility, which has allowed the Federal Reserve to adopt more cautious policies, and implement them effectively over longer time horizons.

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. The empirical evidence in Rigobon and Sack (2003) is also consistent with this interpretation.

Finally, Figure 1 in the introduction shows that substantial changes in the sign and magnitude of nominal bond betas and bond return volatilities line up reasonably well with our proposed monetary policy regimes. Our estimates of the monetary policy rule shown below also provide robust empirical support for the existence of a third monetary policy regime.

### 3.2 Data and summary statistics

Our empirical analysis uses quarterly US data on output, inflation, interest rates, and aggregate bond and stock returns from 1954Q3 to 2011Q4. GDP in 2005 chained dollars and

---

9See also Bernanke (2004) for anecdotal evidence for increased gradualism in the 2000s.
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 during the last week of the quarter 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.\(^\text{10}\) The end-of-quarter five year bond yield is from the CRSP monthly Treasury Fama-Bliss discount bond yields. 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.\(^\text{11}\) 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 rates 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.

\(^\text{10}\) 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.

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, 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 necessarily drives up the volatility of the log price-dividend ratio in the data as compared to the model.

Table 2 estimates the relation between the output gap and equity risk premia in US data and shows that this relation is stable across subperiods. Table 2 estimates the following predictive relation between quarterly equity excess returns and the output gap:

\[ r_{t+1}^e - i_t = a_0 + a^x x_t + \epsilon_{t+1}. \]  \hspace{1cm} (36)

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 weighted subperiod average of \( a^x = -0.54 \).
3.3 Estimating monetary policy rules

The monetary policy parameters are important inputs into the calibrated model for each subperiod. The model’s inclusion restrictions are such that if we knew the output gap, the inflation gap, and the interest rate gap, we could estimate the monetary policy function by OLS. Unfortunately, the inflation gap and interest rate gap are not directly observable. We therefore follow CGG in estimating a monetary policy rule in terms of the output gap, inflation, and the Fed Funds rate:

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

(37)

If the inflation target were constant over time, we could use the estimates \( \hat{c}^x \), \( \hat{c}^\pi \), and \( \hat{c}^i \) to back out the monetary policy rule parameters according to:

\[ \hat{\rho}^i \quad = \quad \hat{c}^i, \]  

(38)

\[ \hat{\gamma}^x \quad = \quad \hat{c}^x / (1 - \hat{c}^i), \]  

(39)

\[ \hat{\gamma}^\pi \quad = \quad \hat{c}^\pi / (1 - \hat{c}^i). \]  

(40)

In the full model, we do not use the potentially biased “naïve” estimates (38) through (40). 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 \( \gamma^x \), \( \gamma^\pi \) and \( \rho^i \) to match the empirical slope coefficients (37) for each subperiod. Nonetheless, the naïve monetary policy parameter estimates (38) through (40) are 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 (37) should indicate a change in monetary policy regime. CGG have argued forcefully that the monetary policy rule changed substantially in the early 1980s. It is therefore plausible that we should find at least 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 (37) 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 of 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. 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.

Second, we test for a break in the post-1977Q2 subsample. The estimated break date is 2001Q1 and we can reject the null of no break with or without known break date at the 95% confidence level. Finally, we test for a break in the pre-1977Q2 subsample. The estimated break date is 1960Q2 and we can reject the null of no break with or without known break date at the 95% confidence level. Based on the evidence from these break date tests, we therefore calibrate the model to the following three subperiods: 1960Q2-1977Q1, 1977Q2-2000Q4, and 2001Q1-2011Q4.

---

12 We use the 5% critical value for 4 restrictions and 25% trimming tabulated in Andrews (2003).
Table 3 reports estimates of the OLS Taylor rule regression (37) for the three subperiods together with naïve implied monetary policy parameters. Table 3 reports standard errors for the naïve monetary policy parameters using the delta method, and indicates with stars the parameters that are significant at the 5% or 1% level based on a likelihood ratio test.

The estimates in Table 3 suggest 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 $\rho^{i}$ is stable during the first two periods and increases during the post-2000 period.

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, 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, we find similar empirical results.

The point estimates of $\hat{\gamma}^{\pi}$ in Table 4 also suggest that the central bank has put somewhat higher weight on output fluctuations in the earliest and latest subperiods than during the

---

13 $\hat{\gamma}^{\pi}$ appears to be less precisely estimated in the latest subperiod, while $\hat{c}^{\pi}$ is precisely estimated. This results from the nonlinear relation between the parameters. The coefficient $\hat{\gamma}^{\pi}$ is $\hat{c}^{\pi}$ divided by $(1 - \hat{c}^{i})$. Because $\hat{c}^{i}$ is very close to 1 in the latest subperiod, standard errors for $\hat{\gamma}^{\pi}$ based on the delta method tend to be very large.
middle subperiod, although the estimates of neither $\gamma x$ nor $\delta x$ are statistically significant in the latest subperiod.

Consistent with our informal observations about increased gradualism in monetary policy, during the most recent subperiod the regression explains 94% of the variation in the Federal Funds rate, implying only small deviations from the monetary policy rule. The coefficient on the lagged Fed Funds rate in the monetary policy function is large at 0.83.

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. Not surprisingly, none of the estimated coefficients of the monetary policy rule are statistically different from zero during the post-crisis period.

4 Model Calibration

We now calibrate our model to key empirical moments for the US over the three subsamples: 1960Q2-1977Q2, 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 are allowed to change across subperiods.
Time-invariant parameters include those governing the relation between the output gap, consumption, and dividends ($\phi, \tau, g, \delta$), preference parameters ($\gamma, \theta_0, \theta_1, \theta_2, \bar{r}$), and Phillips curve parameters ($\rho^\pi, \lambda$). Time-varying parameters include the monetary policy rule parameters $\gamma^x, \gamma^\pi$, and $\rho^i$ and the shock volatilities $\sigma^{PC}, \sigma^{MP}$, and $\sigma^*$. 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 $\bar{r}$ and the average growth rate of consumption $g$ exactly as in Campbell and Cochrane (1999).

The parameter $\phi$ 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 (11), consumption growth is serially uncorrelated if and only if $\phi$ equals the first-order autocorrelation of the output gap. Strong predictability in consumption growth generates predictability of real rates and large bond and stock return volatilities in our model. In order to generate plausible volatilities of asset returns, we choose $\phi$ to generate a 12-quarter consumption variance ratio that averages to 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 $\phi$. For a
given value of $\phi$, we choose the scaling parameter $\tau$ to match the ratio of full sample standard deviations for stochastically detrended consumption and the output gap.

We calibrate the persistence of the surplus consumption ratio as in Wachter (2006) to $\theta_0 = 0.89$. We set the utility curvature parameter to $\gamma = 3$. By comparison, Campbell and Cochrane (1999) and Wachter (2006) use a curvature parameter of $\gamma = 2$. In our model, $\gamma$ 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 $\gamma$ flattens the relation between the output gap and the real short-term interest rate, and helps us generate non-explosive macroeconomic dynamics. At the same time, asset Sharpe ratios rise roughly with $\gamma/S$, which is proportional to $\sqrt{\gamma}$. Therefore, setting $\gamma = 3$ leads to a small increase in asset Sharpe ratios as compared to Campbell and Cochrane (1999).

The new preference parameters $\theta_1$ and $\theta_2$ are both set to 0.02, so they are small but positive. Positive $\theta_1$ and $\theta_2$ imply that a positive output gap raises surplus consumption this period and next, as might be the case with slowly moving external habit. If we set $\theta_1$ and $\theta_2$ too high, we obtain implausibly high real bond return volatilities. The parameters $\theta_1$ and especially $\theta_2$ are also important for macroeconomic dynamics. We need $\theta_2 > 0$ to ensure that the Euler equation has a backward-looking component, which helps generate a unique equilibrium. Since $\theta_1$ and $\theta_2$ enter into the Euler equation coefficients, requiring a persistent output gap pins down their relative values.

We choose a Phillips curve slope of $\lambda = 0.04$. 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). Backward-looking Phillips curves can arise in a model with infrequent information updating (Mankiw and Reis, 2002). 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. We choose monetary policy parameters $\gamma^x$, $\gamma^\pi$ and $\rho^i$ 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 time-variation in the central bank’s inflation target.

The calibrated monetary policy parameters mirror the broad changes in naively estimated monetary policy parameters in Table 3. The output gap weight $\gamma^x$ decreased from the pre-Volcker period to the Volcker period, but increased again in the post-2000 period. The inflation weight $\gamma^\pi$ was below 1 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 slightly increase 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, $\sigma^{IS}$, $\sigma^{MP}$ and $\sigma^*$, 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 the Federal Reserve’s officially stated target.

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 forward- and
backward-looking Euler equation components sum to more than one as a result of time-varying risk premia. 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). The steady-state surplus consumption ratio is smallest and hence risk premia are most volatile in the most recent subperiod.

### 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 provides a good fit for the empirical OLS monetary policy rules estimated in Table 3. The fact that the model matches regressions of the Federal Funds rate onto the output gap, inflation, and lagged Fed Funds rate validates our choice of monetary policy parameters. The top half of Table 5, 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 bottom half of Panel B reports volatilities of returns on equities, nominal bonds, and real bonds, and the betas of nominal bonds and real bonds with stocks. 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. Thus our model addresses the “equity volatility” puzzle, one of the leading puzzles in consumption-based asset pricing (Campbell, 2003).

---

14 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.

15 All model moments are calculated from 5 simulations of length 10000. We choose a long simulation period to capture the steady-state distribution of the state variables.
The middle 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 particularly extreme in the third subperiod, while the latter is relatively constant across the subperiods.

The bottom panel shows that our model fits overall stock return volatilities quite well, although it slightly overstates volatility in the second period and understates it in the third period. Mirroring the behavior of 5-year nominal bond yields, bond return volatilities are lower in the model than in the data. The calibrated model fits very well the time variation in nominal bond betas across subperiods. Both empirical and model nominal bond betas are positive in the first subperiod, increase in the second subperiod, and turn negative in the third subperiod. These changes in bond risks are the primary object of interest in our analysis.

The last two rows in Table 5 show that in the third period, when inflation-indexed bonds were available in the U.S., the implied model real bond beta is very close to zero, but not actually negative as in the data. If inflation-indexed bonds had been available during earlier periods, the model indicates that real bond returns would have been more volatile in the second and especially in the first subperiod. Moreover, real bonds would have been especially valuable hedges with negative real bond betas during the 1960Q2-1977Q1 subperiod, when PC shocks were dominant.

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 2.03%, as compared to a standard deviation of 1.54% for quarterly consumption growth in the data. The twelve-quarter consumption variance ratio averages to exactly 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 37.73, 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 the model has reasonable variation in equity risk premia.

The bottom panel of Table 6 evaluates the implications of the model for bond returns. The model implies an average upward-sloping term structure. This average is driven by a more strongly upward-sloping term structure during the first two subperiods and a
downward-sloping nominal term structure during the third subperiod, consistent with negative bond betas during this subperiod. A model regression of one-quarter nominal bond excess returns onto the lagged bond yield spread generates a predictability coefficient very close to that in the data (Campbell and Shiller, 1991).\textsuperscript{16} Table 6 also reports the average standard deviation of the real interest rate in the model, which is a reasonable 2.24\% per annum.

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 nominal 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.\textsuperscript{17} Each panel shows three

\textsuperscript{16} Wachter (2006) reports a similar result using a different modification of Campbell-Cochrane preferences in a model with an exogenous consumption process.

\textsuperscript{17} Impulse response functions for the nominal bond yield and dividend yield are averaged over 2000 simulations, starting from the unconditional steady-state of the system at time 0. The responses for the output gap, inflation, the nominal short rate and the real short rate can be computed analytically, since
lines, each corresponding to one subperiod calibration. Vertical bars 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 across subsamples.

Figure 3 shows that a typical PC shock acts as a persistent inflationary and contractionary supply shock. It leads to a decline in the output gap and an increase in inflation. For all three subperiods, but especially for the second subperiod, short-term and 5-year nominal bond yields rise in the shock period and nominal bond prices fall. In contrast, the real short-term rate drops in response to a PC shock. This is the case even for subperiods 2 and 3, where the monetary policy rule has an inflation weight $\gamma^\pi$ greater than one. The reason is that the central bank dampens the increase in the nominal policy rate in response to the steep drop in the output gap following a PC shock. The persistent decrease in the output gap raises equity risk premia, lowers stock prices, and raises the dividend-price ratio. A PC shock therefore 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.

The three subperiod calibrations show differential responses to PC shocks and those differences are related to changes in monetary policy. The especially accommodative monetary policy in the first subperiod implies that the output gap response is smallest and the inflation response is largest for this subperiod. On the other hand, the positive impact of a PC shock on the long-term nominal yield and the nominal bond beta is strongest in the second subperiod. In this period, the Federal Reserve reacts aggressively and immediately to the increase in inflation, driving up nominal bond yields.

The nominal bond yield response to a PC shock is smallest in the third subperiod for these variables follow a VAR(1) in equilibrium.
two reasons. First, the central bank’s renewed focus on output gap fluctuations leads to a more accommodative policy rate response than in the second subperiod. Second, risk premia further dampen the effect of PC shocks on bond betas during this period. Because nominal bonds have a negative beta in the third subperiod, nominal bonds benefit from a flight to safety effect when risk aversion increases in recessions. This is an illustration of the important risk premium channel operative in our model.

A monetary policy (MP) shock acts as a strongly positive impulse to nominal and real short-term interest rates. The effect on the nominal short-term interest rate is especially long-lasting for subperiod 3, where monetary policy is characterized by a high persistence coefficient. The MP shock induces consumers to postpone consumption, generating a strongly negative response in the output gap and a slow decrease in inflation. The output gap and inflation react with a lag due to our identification assumption that real variables are determined using an information set that excludes contemporaneous monetary policy shocks.

In subperiod 3, nominal bonds are hedges and a MP shock has counteracting effects on nominal bond yields. Risk aversion increases at the beginning of recessions, leading to a drop in nominal bond yields. Subsequently, the effect from higher expected short-term nominal rates dominates, leading to an increase in nominal bond yields. The decrease in equity dividend yields reflects a strong decrease in dividends, dominating a smaller decrease in equity prices.

Finally, a shock to the central bank’s inflation target has a delayed but permanent effect on inflation and generates negative bond betas. As inflation gradually increases towards the new target, the output gap experiences a temporary boom. This increase in the output gap occurs because an increase in inflation induces firms with nominal rigidities to produce
more, as captured by the Phillips curve relation in the model. The increase in output 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.

Impulse responses illustrate the qualitative effects of shocks on bond betas, holding constant the risk properties of bonds and stocks. In general, volatilities of shocks affect the riskiness of bonds and stocks priced into impulse responses. A full understanding of the changes in betas requires that we compare the full effects of volatilities of shocks and monetary policy changes. We turn to this next.

5.2 Counterfactual bond risks

We now answer our initial question, namely how changes in monetary policy and volatilities of shocks contribute to shifts in bond betas across our different periods. Figure 4 plots the model’s implied nominal bond betas across subperiods. We show baseline model nominal bond betas (also reported in Table 5) in blue. Baseline bond betas reflect changes in both monetary policy and macroeconomic shocks. However, the model has implications for how nominal bond betas would have changed if only monetary policy or only the volatility of macroeconomic shocks had changed.

Figure 4 decomposes changes in model bond betas into changes due to monetary policy and macroeconomic shocks. The green line depicts model-implied nominal bond betas in a scenario where PC shocks are equally volatile across periods, but monetary policy rule coef-
coefficients \((\gamma^x, \gamma^\pi, \rho^i)\) and the volatilities of central bank-driven shocks \((\sigma^{MP}, \sigma^*)\) change across subperiods. In contrast, the red line shows model-implied bond betas in a scenario where only the volatility of PC shocks changes across subperiods, but monetary policy remains unchanged from the pre-Volcker period.

Figure 4 shows that in the late 1970s, changes in monetary policy acted to increase the nominal bond beta, while changes in the volatility of Phillips curve shocks acted to decrease the nominal bond beta. In contrast, monetary policy and the volatility of PC shocks acted in the same direction in 2000, both decreasing the beta of nominal bonds.

As we have already discussed, PC shocks tend to create stagflationary recessions, thereby increasing the beta of nominal bonds. As PC shocks became less dominant over time, this acted to generate a smaller bond beta in the 1977Q2-2000Q4 period and even a negative bond beta during the post-2000 period. While we did see a negative bond beta during the post-2000 period, the empirical beta of nominal bonds increased during the Volcker-period. In order to understand this increase in bond betas, we turn to monetary policy changes.

The green line in Figure 4 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, consistent with the data. Monetary policy during the second subperiod is characterized by a strong and immediate response to inflation fluctuations and relatively smaller concern for output fluctuations. During the Volcker-Greenspan period, the central bank reacts to PC shocks by raising the short-term nominal bond yield strongly and immediately, amplifying the negative effect on both bond and stock returns and increasing the bond beta.

If only monetary policy had changed after 2000, the model implies a decrease in the
nominal bond beta. Two changes in monetary policy drive this decrease in 2000. First, the renewed focus on output gap fluctuations implies a more accommodative monetary policy response to PC shocks, decreasing the effect of PC shocks on bond betas. Second, a higher volatility of inflation target shocks also drives the nominal bond beta downward. However, with unchanged volatility of PC shocks the nominal bond beta would have remained positive (albeit smaller) in the third subperiod.

Next, we further decompose the effects of monetary policy changes, focusing on changes in the monetary policy rule coefficients. Figure 5 shows counterfactual changes in bond betas when only the monetary policy rule coefficients (\(\gamma^x\), \(\gamma^\pi\), \(\rho^i\)) change across subperiods, but the volatilities of all shocks (\(\sigma^{PC}\), \(\sigma^{MP}\), \(\sigma^*\)) are constant at their period 1 values. We compare these changes to a scenario where the monetary policy rule is constant but the volatilities of all shocks change. The green lines in Figures 4 and 5 show similar changes in the late 1970s and after 2000. The most notable difference is that the green line shows a smaller post-2000 decline in Figure 5 than in Figure 4.

We conclude that changes in the monetary policy rule in the late 1970s are most important for understanding the increase in the empirical nominal bond beta that occurred at that time. On the other hand, the model attributes the negative nominal bond beta after 2000 to a combination of smaller supply shocks, more volatile inflation target shocks, and an increased weight on output fluctuations in the monetary policy rule.
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 argues that understanding bond risks requires modeling the influence of monetary policy on the macroeconomy, particularly the relation between output and inflation, and understanding 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 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, the model 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, the response of monetary policy to output is also important. When policy is more responsive to output, as it has been after 2000, interest rates tend to fall in recessions, driving up bond prices. This pattern tends to create a negative beta for nominal bonds.

Third, 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 both after 1977 and after 2001, consistent with a continued 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 both in the late 1970s and after 2000. 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 generated negative nominal bond betas.

Fourth, we find that the volatility of shocks to the central bank’s inflation target contributes to lower nominal bond betas after 2000. These inflation target shocks may be interpreted literally, as the result of shifting central bank preferences, or more broadly as the result of changes in the credibility in monetary policy (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.

Our analysis has several limitations that provide opportunities for future research. First, we have made a number of specific assumptions that could be modified in variants of our approach. For example, 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, since we use a New Keynesian model, the micro-foundations of our model are not as clear and detailed as is standard in the dynamic stochastic general equilibrium literature. We have little to say about the production side of the economy or the labor market. 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). Third, 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). Fourth, the regime shifts we consider are unanticipated, once-and-for-all events rather than stochastically recurring events whose probabilities are understood by market participants. 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.
References


51


Tables and Figures

Table 1: Summary Statistics

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

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

<table>
<thead>
<tr>
<th>Period</th>
<th>Output Gap</th>
<th>Inflation</th>
<th>Fed Funds</th>
<th>Nom. Bond Yield</th>
</tr>
</thead>
<tbody>
<tr>
<td>Mean</td>
<td>(2.30)</td>
<td>(2.41)</td>
<td>(3.44)</td>
<td>(2.49)</td>
</tr>
<tr>
<td>Std</td>
<td>0.95</td>
<td>0.89</td>
<td>0.85</td>
<td>0.95</td>
</tr>
<tr>
<td>AR(1) Coefficient</td>
<td>0.04</td>
<td>0.05</td>
<td>0.05</td>
<td>(0.03)</td>
</tr>
<tr>
<td>AR(4) Coefficient</td>
<td>0.65</td>
<td>0.85</td>
<td>0.73</td>
<td>0.80</td>
</tr>
<tr>
<td></td>
<td>(0.09)</td>
<td>(0.05)</td>
<td>(0.07)</td>
<td>(0.06)</td>
</tr>
</tbody>
</table>

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

Table 2: Predicting Stock Returns with Output Gap

<table>
<thead>
<tr>
<th>Log Exc. Stock Ret. $x_{t+1}$</th>
<th>Average</th>
<th>60.Q2-77.Q1</th>
<th>77.Q2-00.Q4</th>
<th>01.Q1-11.Q4</th>
</tr>
</thead>
<tbody>
<tr>
<td>Output Gap $x_t$</td>
<td>-0.54</td>
<td>-0.66*</td>
<td>-0.27</td>
<td>-0.71</td>
</tr>
<tr>
<td></td>
<td>(0.31)</td>
<td>(0.39)</td>
<td>(0.43)</td>
<td></td>
</tr>
<tr>
<td>Constant</td>
<td>0.48</td>
<td>0.98</td>
<td>1.42</td>
<td>-1.27</td>
</tr>
<tr>
<td></td>
<td>(1.07)</td>
<td>(0.82)</td>
<td>(1.74)</td>
<td></td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.03</td>
<td>0.02</td>
<td>0.04</td>
<td>0.01</td>
</tr>
</tbody>
</table>

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

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

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. Since the inflation target is not directly observable it is omitted. 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

<table>
<thead>
<tr>
<th>Parameter</th>
<th>Value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Smoothing Parameter Consumption</td>
<td>0.94</td>
</tr>
<tr>
<td>Scaling Parameter Consumption</td>
<td>0.89</td>
</tr>
<tr>
<td>Consumption Growth Rate</td>
<td>1.89</td>
</tr>
<tr>
<td>Leverage</td>
<td>3.39</td>
</tr>
<tr>
<td>Persistence Surplus Cons.</td>
<td>0.89</td>
</tr>
<tr>
<td>Dependence Output Gap</td>
<td>0.02</td>
</tr>
<tr>
<td>Dependence Lagged Output Gap</td>
<td>0.02</td>
</tr>
<tr>
<td>Utility Curvature</td>
<td>3.00</td>
</tr>
<tr>
<td>Steady-State Riskfree Rate</td>
<td>0.94</td>
</tr>
<tr>
<td>PC Lag Coefficient</td>
<td>0.96</td>
</tr>
<tr>
<td>Phillips Curve Slope</td>
<td>0.04</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Monetary Policy Rule</th>
<th>60.Q2-77.Q1</th>
<th>77.Q2-00.Q4</th>
<th>01.Q1-11.Q4</th>
</tr>
</thead>
<tbody>
<tr>
<td>Output Weight</td>
<td>0.37</td>
<td>0.31</td>
<td>0.40</td>
</tr>
<tr>
<td>Inflation Weight</td>
<td>0.63</td>
<td>1.50</td>
<td>1.85</td>
</tr>
<tr>
<td>Persistence MP</td>
<td>0.78</td>
<td>0.62</td>
<td>0.92</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Std. Shocks</th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Std. PC</td>
<td>0.88</td>
<td>0.51</td>
<td>0.05</td>
</tr>
<tr>
<td>Std. MP</td>
<td>0.87</td>
<td>1.69</td>
<td>0.70</td>
</tr>
<tr>
<td>Std. Infl. Target</td>
<td>0.03</td>
<td>0.12</td>
<td>0.32</td>
</tr>
</tbody>
</table>

Panel B: Implied Parameters

<table>
<thead>
<tr>
<th>Parameter</th>
<th>60.Q2-77.Q1</th>
<th>77.Q2-00.Q4</th>
<th>01.Q1-11.Q4</th>
</tr>
</thead>
<tbody>
<tr>
<td>Discount Rate</td>
<td>0.88</td>
<td>0.88</td>
<td>0.88</td>
</tr>
<tr>
<td>IS Curve lag Coefficient</td>
<td>0.02</td>
<td>0.02</td>
<td>0.02</td>
</tr>
<tr>
<td>IS Curve Forward Coefficient</td>
<td>1.10</td>
<td>1.10</td>
<td>1.10</td>
</tr>
<tr>
<td>IS Curve Real Rate Slope</td>
<td>0.41</td>
<td>0.41</td>
<td>0.41</td>
</tr>
<tr>
<td>Steady-State Surplus Cons. Ratio</td>
<td>0.11</td>
<td>0.12</td>
<td>0.07</td>
</tr>
<tr>
<td>Log Max. Surplus Cons. Ratio</td>
<td>-1.70</td>
<td>-1.61</td>
<td>-2.14</td>
</tr>
<tr>
<td>Max Surplus Cons. Ratio</td>
<td>0.18</td>
<td>0.20</td>
<td>0.12</td>
</tr>
</tbody>
</table>
Table 5: Model and Empirical Moments


<table>
<thead>
<tr>
<th></th>
<th>60.Q2-77.Q1</th>
<th>77.Q2-00.Q4</th>
<th>01.Q1-11.Q4</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Empirical</td>
<td>Model</td>
<td>Empirical</td>
</tr>
<tr>
<td>Output Gap</td>
<td>0.17**</td>
<td>0.16</td>
<td>0.03</td>
</tr>
<tr>
<td>Inflation</td>
<td>0.21**</td>
<td>0.26</td>
<td>0.41*</td>
</tr>
<tr>
<td>Lagged Fed Funds</td>
<td>0.69**</td>
<td>0.67</td>
<td>0.66*</td>
</tr>
</tbody>
</table>

Panel B: Subperiod Second Moments

<table>
<thead>
<tr>
<th></th>
<th>60.Q2-77.Q1</th>
<th>77.Q2-00.Q4</th>
<th>01.Q1-11.Q4</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Empirical</td>
<td>Model</td>
<td>Empirical</td>
</tr>
<tr>
<td>Std. VAR(1) Residuals</td>
<td>0.83</td>
<td>1.16</td>
<td>0.76</td>
</tr>
<tr>
<td>Output Gap</td>
<td>1.05</td>
<td>0.81</td>
<td>1.04</td>
</tr>
<tr>
<td>Inflation</td>
<td>0.90</td>
<td>0.82</td>
<td>1.55</td>
</tr>
<tr>
<td>Fed Funds Rate</td>
<td>0.47</td>
<td>0.27</td>
<td>0.77</td>
</tr>
<tr>
<td>Log Nominal Yield</td>
<td>18.35</td>
<td>18.27</td>
<td>15.68</td>
</tr>
<tr>
<td>Std. Asset Returns</td>
<td>Std. Eq. Ret.</td>
<td>4.92</td>
<td>2.74</td>
</tr>
<tr>
<td>Std. Nom. Bond Ret.</td>
<td>0.07**</td>
<td>0.14</td>
<td>0.12</td>
</tr>
<tr>
<td>Nominal Bond Beta</td>
<td>3.52</td>
<td>1.00</td>
<td>4.27</td>
</tr>
</tbody>
</table>

This table reports average model moments from 5 simulations of length 10000. * 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 in the bottom panel.
Table 6: Consumption and Risk Premia

<table>
<thead>
<tr>
<th>Consumption and Output Gap</th>
<th>Empirical</th>
<th>Model</th>
</tr>
</thead>
<tbody>
<tr>
<td>Std. Cons. Innovation</td>
<td>1.53</td>
<td>2.03</td>
</tr>
<tr>
<td>AR(1) Coefficient Output Gap</td>
<td>0.96</td>
<td>0.91</td>
</tr>
<tr>
<td>Twelve Quarter Cons. Variance Ratio</td>
<td></td>
<td>1.00</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Equities</th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Equity Premium</td>
<td>4.50</td>
<td>8.21</td>
</tr>
<tr>
<td>Price-Dividend Ratio (exp(mean(p-d)))</td>
<td>30.05</td>
<td>37.73</td>
</tr>
<tr>
<td>Std(d-p)</td>
<td>0.40</td>
<td>0.18</td>
</tr>
<tr>
<td>AR(4) Coefficient d-p</td>
<td>0.92</td>
<td>0.88</td>
</tr>
<tr>
<td>Correlation(x, p-d)</td>
<td>0.54</td>
<td>0.51</td>
</tr>
<tr>
<td>Slope 1 YR Exc. Stock ret. wrt d-p</td>
<td>0.08</td>
<td>0.19</td>
</tr>
<tr>
<td>Slope 5 YR Exc. Stock ret. wrt d-p</td>
<td>0.25</td>
<td>0.61</td>
</tr>
<tr>
<td>Slope quarterly stock ret. wrt x</td>
<td>-0.49</td>
<td>-0.63</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Nominal 5-Year Zero Coupon Bond</th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Nominal Bond Excess Return</td>
<td>1.60</td>
<td>0.84</td>
</tr>
<tr>
<td>Slope Term Structure</td>
<td>1.06</td>
<td>0.49</td>
</tr>
<tr>
<td>1-Quarter Bond Excess Return onto Lagged Term Spread</td>
<td>2.76</td>
<td>2.71</td>
</tr>
<tr>
<td>Std(real rate) (Percent, Ann.)</td>
<td>2.24</td>
<td></td>
</tr>
</tbody>
</table>

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.
Nominal bond beta and standard deviation of nominal bond returns from daily bond and stock returns over past three months as in Campbell, Sunderam, and Viceira (2013). 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.
Panel A: Output Gap and Detrended Consumption

Panel B: Output Gap and 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. The smoothing parameter \( \phi = 0.94 \) 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.
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 following one standard deviation shocks. 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 average subsample standard deviation, where the average is weighted by sample length. Impulse responses for the nominal bond yield and dividend yield are averaged over 2000 simulations, starting from the unconditional steady-state of the system 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 units.
This figure shows model-implied bond betas for three subperiods. The blue solid line replicates the baseline model bond betas from Table 5. The green dashed line holds the standard deviations of PC shocks constant at their period 1 values and varies the monetary policy rule coefficients, MP shock volatilities and inflation target shock volatilities across subperiods. The red dash-dot line changes only the volatilities of PC shocks across subperiods and holds constant the MP rule coefficients, and volatilities of MP and inflation target shocks.
This figure shows model-implied bond betas for three subperiods. The blue solid line replicates the baseline model bond betas from Table 5. This figure differs from Figure 5 in that the green dashed line varies only the MP coefficients ($\gamma^x, \gamma^\pi, \rho^i$) across subperiods and holds all volatilities of shocks, including $\sigma^{MP}$ and $\sigma^*$, constant. The red dash-dot line holds the monetary policy coefficients constant and changes only the volatilities of shocks ($\sigma^{PC}, \sigma^{MP}, \sigma^*$) across subperiods.