

Monetary Policy Drivers of Bond and Equity Risks

John Y. Campbell, Carolin Pflueger, and Luis M. Viceira¹

First draft: March 2012

This draft: August 2013

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. This material is based upon work supported by Harvard Business School Research Funding and the Bureau of Asset Management 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 parameters and macroeconomic shocks on nominal bond risks, using a New Keynesian model with habit formation and discrete regime shifts in 1979 and 1997. The increase in bond risks after 1979 is attributed primarily to a shift in monetary policy towards a more anti-inflationary stance, while the more recent decrease in bond risks after 1997 is attributed primarily to an increase in the persistence of monetary policy interacting with continued shocks to the central bank's 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 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 (2010) have documented these developments. In this paper we ask how monetary policy has contributed to these changes in bond risks.

Given the importance of Treasury bonds as an asset class, it is natural to ask what economic factors determine their risk properties. One way to do this is to use identities that link bond returns to movements in bond yields, and that link nominal bond yields to expectations of future short-term real interest rates, expectations of future inflation rates, and time-varying risk premia on longer-term bonds over short-term bonds. Barsky (1989), Shiller and Beltratti (1992), and Campbell and Ammer (1993) were early examples of this approach. A more recent literature has proceeded in a similar spirit, building on the no-arbitrage restrictions of affine term structure models (Duffie and Kan 1996, Dai and

Singleton 2000, 2002, Duffee 2002) to estimate multifactor term structure models with both macroeconomic and latent factors (Ang and Piazzesi 2003, Ang, Dong, and Piazzesi 2007, Rudebusch and Wu 2007). Although these exercises can be informative, they are based on a reduced-form econometric representation of the stochastic discount factor and the process driving inflation. This limits the insights they can deliver about the 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 (Bansal and Shaliastovich 2010, 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 (Bekaert, Engstrom, and Grenadier 2010, Bekaert, Engstrom, and Xing 2009, Buraschi and Jiltsov 2007, 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

²A 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.

(2012), Kung (2013), Li and Palomino (2013), Palomino (2012), Rudebusch and Wu (2008), and Rudebusch and Swanson (2012).

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 and stochastic volatility, 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 (Bansal and Shaliastovich 2010, Bekaert, Engstrom, and Grenadier 2010, Campbell 1986, Campbell, Sunderam, and Viceira 2013, d’Addona and Kind 2006, Eraker 2008, Hasseltoft 2008, 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 the shocks hitting the economy. 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 basic 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 solves for the stochastic discount factor (SDF) implied by the New Keynesian IS curve, and uses it to price bonds and stocks.

Section 3 describes our data sources and presents summary statistics for our full sample period, 1960Q1 through 2011Q4, and for three subperiods, 1960Q1–1979Q2, 1979Q3–1996Q4, and 1997Q1–2011Q4. These subperiods are chosen to match both shifts in monetary policy and changes in measured bond risks. This section also estimates the parameters of the monetary policy reaction function, over the full sample and the three subperiods, using reduced-form regression methodology.

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 through time if the parameters of the monetary policy rule, or the volatilities of shocks hitting the economy, had been stable instead of time-varying. Section 6 concludes, and an online appendix (Campbell, Pflueger, and Viceira 2013) presents additional details.

2 A New Keynesian Asset Pricing Model

We model the dynamics of macroeconomic time series with a standard New Keynesian framework consisting of a log-linearized Euler equation, a Phillips curve, and a monetary policy function. We integrate asset pricing into the framework by deriving the Euler equation from a stochastic discount factor (SDF) that also prices stocks and bonds in the model. The SDF links asset returns and macroeconomic and monetary variables in equilibrium in a standard no-arbitrage setup.

The Euler equation is a standard New Keynesian building block and provides an equivalent of the Investment and Savings (IS) curve. We provide a micro-founded log-linearized Euler equation relating current output to the lagged output gap, the expected future output gap, and the 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).³ Fuhrer (2000) argues that allowing for a backward-looking component is important for capturing the empirical hump-shaped output response to a monetary policy shock. The forward-looking component follows from standard household dynamic optimization.

We derive an Euler equation with both backward-looking and forward-looking components from a consumption-based SDF in which the marginal utility of consumption depends

³Christiano, 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).

on the current and lagged values of the output gap, and its conditional volatility varies inversely with the output gap. This assumption about the volatility of marginal utility implies that real risk premia increase during recessions, consistent with the empirical evidence on stock and bond return predictability (Fama and French 1989). A parametric model that exhibits these properties and produces analytically tractable expressions for asset prices and expected returns after suitable log-linearization of the SDF is the habit-formation model of Campbell and Cochrane (1999), in which utility is a power function of the difference between consumption and habit. We therefore adopt this specification of utility for analytical convenience. Finally, shocks to marginal utility, or demand shocks, introduce shocks to the Euler equation.

The second building block of a New Keynesian model is the Phillips curve (PC) equation that links inflation and real output in equilibrium. We assume a PC with both forward- and backward-looking components to capture the price setting behavior of firms. 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 a time-varying central bank inflation target. Historical US inflation appears highly persistent (Ball and Cecchetti 1990, Stock and Watson 2007). We capture this empirical regularity by modeling the inflation target as a unit root process. Movements in our estimated inflation target may capture episodes where public expectations of central bank behavior are not well anchored, because the central bank lacks credibility, even if the central bank's true target is relatively stable (Orphanides and Williams 2004).

To close the model we need to make identification assumptions. 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. 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 central bank's monetary policy rule by Ordinary Least Squares (OLS).

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

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

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

$$M_{t+1} = \frac{\beta U'_{t+1}}{U'_t}, \quad (2)$$

where U'_t is the marginal utility of consumption at time t and β 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:

$$\ln U'_t = r_t + \ln \beta + \ln E_t U'_{t+1}, \quad (3)$$

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 π_{t+1} is inflation from time t to time $t + 1$. This approximation is justified if uncertainty about inflation is small at the quarterly horizon, as appears to be the case empirically.

Equation (3) also describes the nominal interest rate given a model for expected inflation. Substituting $r_t = i_t - E_t\pi_{t+1}$ into (3), and dropping constants to reduce the notational burden, we have:

$$\ln U'_t = (i_t - E_t\pi_{t+1}) + \ln E_t U'_{t+1}. \quad (4)$$

We assume that $\ln U'_t$ is a linear function of the current and lagged log output gap x_t and that its conditional volatility is also an exponential affine function of x_t with a negative slope so that the volatility of marginal utility is higher when the output gap is low. These assumptions imply an Euler equation for the real riskfree rate that relates the real interest rate to the current output gap, its first-order lag, and its expected value.

To see this, 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 :

$$U_t = \frac{(C_t - H_t)^{1-\alpha} - 1}{1-\alpha} = \frac{(S_t C_t)^{1-\alpha} - 1}{1-\alpha}. \quad (5)$$

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

Marginal utility in this model is

$$U'_t = (C_t - H_t)^{-\alpha} = (S_t C_t)^{-\alpha}, \quad (6)$$

and log marginal utility is given by $\ln U'_t = -\alpha(s_t + c_t)$. Assuming lognormality, or taking a

second-order Taylor approximation, the Euler equation (4) becomes

$$-\alpha(s_t + c_t) = (i_t - E_t\pi_{t+1}) - \alpha E_t(s_{t+1} + c_{t+1}) + \frac{\alpha^2}{2}\sigma_t^2, \quad (7)$$

where $\sigma_t^2 = \text{Var}_t(s_{t+1} + c_{t+1})$.

Now suppose that

$$s_t + c_t = x_t - \theta x_{t-1} - v_t, \quad (8)$$

where the error term v_t is white noise uncorrelated with current or lagged x_t , or any other information variables known in advance. Furthermore, assume that the volatility of marginal utility is higher when the output gap is low. For some $0 < b < 1$:

$$\sigma_t^2 = \bar{\sigma}^2 \exp(-bx_t) \approx \bar{\sigma}^2(1 - bx_t). \quad (9)$$

Here $\bar{\sigma}$ is the conditional volatility of the SDF when the output gap is zero.

Substituting (8) and (9) into (7) yields the Euler equation:

$$x_t = \rho^{x^-} x_{t-1} + \rho^{x^+} E_t x_{t+1} - \psi (i_t - E_t\pi_{t+1}) + u_t^{IS}, \quad (10)$$

where $\rho^{x^-} = \theta/(1 + \theta^*)$, $\rho^{x^+} = 1/(1 + \theta^*)$, $\psi = 1/\alpha(1 + \theta^*)$, $u_t^{IS} = v_t/(1 + \theta^*)$, and $\theta^* = \theta - \alpha b \bar{\sigma}^2/2 < \theta$.

Several points are worth noting about the IS curve (10). First, because $\theta^* < \theta$, the coefficients on the lagged output gap and the expected future output gap sum to more than

one. Second, the slope of the IS curve ψ does not equal the elasticity of intertemporal substitution (EIS) of the representative consumer. Third, shocks to the IS curve result from marginal utility or demand shocks in equation (8).

2.2 Macroeconomic dynamics

We complement the consumers' Euler equation with standard building blocks of New Keynesian macroeconomic models.

We assume that consumers and firms do not incorporate contemporaneous monetary policy shocks into their time t decisions, similarly to Christiano, Eichenbaum and Evans (2005). Consumers and price-setting firms form their time t expectations based on monetary policy shocks up to time $t-1$ and IS, PC and inflation target shocks up to time t . We denote the expectation with respect to this information set by:

$$E_{t-}(\cdot) = E(\cdot | u_t^{IS}, u_{t-1}^{IS}, u_{t-2}^{IS}, \dots, u_t^{PC}, u_{t-1}^{PC}, u_{t-2}^{PC}, \dots, u_{t-1}^{MP}, u_{t-2}^{MP}, u_t^*, u_{t-1}^*, \dots). \quad (11)$$

The assumption that consumers and firms make decisions based on E_{t-} 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 we can estimate the monetary policy Taylor rule by OLS.

The dynamics of the output gap, inflation, and Fed Funds rate can then be summarized

by the linearized system of equations:

$$x_t = \rho^{x^-} x_{t-1} + \rho^{x^+} E_{t-} x_{t+1} - \psi(E_{t-} i_t - E_{t-} \pi_{t+1}) + u_t^{IS}, \quad (12)$$

$$\pi_t = \rho^\pi \pi_{t-1} + (1 - \rho^\pi) E_{t-} \pi_{t+1} + \lambda x_t + u_t^{PC}, \quad (13)$$

$$i_t = \rho^i (i_{t-1} - \pi_{t-1}^*) + (1 - \rho^i) [\gamma^x x_t + \gamma^\pi (\pi_t - \pi_t^*)] + \pi_t^* + u_t^{MP}, \quad (14)$$

$$\pi_t^* = \pi_{t-1}^* + u_t^*. \quad (15)$$

Equation (12) is the IS curve (10) with the expectational timing assumption (11). Equation (13) is a standard New Keynesian equation that determines inflation from the price-setting behavior of firms. It has parameters ρ^π , determining the relative weight on past inflation and expected future inflation, and λ , governing the sensitivity of inflation to the output gap.

Equation (14) 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 ρ^i , controlling the influence of past interest rates on current interest rates, γ^x , governing the reaction of the interest rate to the output gap, and γ^π , governing the response of the interest rate to inflation relative to its target level π_t^* . Equation (15) specifies that the central bank's inflation target follows a random walk.

Our monetary policy specification does not explicitly depend on nominal long-term yields. 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). Similarly, monetary policy does not respond directly to stock prices in our model, contrary to

the suggestion of Rigobon and Sack (2003), although it does respond to output and inflation which in turn move stock prices.

The online Appendix, Campbell, Pflueger, and Viceira (2013), shows that it is straightforward to rewrite the model in terms of inflation and nominal interest rate gaps defined as

$$\hat{\pi}_t = \pi_t - \pi_t^*, \quad (16)$$

$$\hat{i}_t = i_t - \pi_t^*. \quad (17)$$

The alternative formulation that results from this substitution is convenient for solving and interpreting the model.

Finally, we assume that the vector of shocks

$$u_t = [u_t^{IS}, u_t^{PC}, u_t^{MP}, u_t^*]' \quad (18)$$

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

$$E_{t-1} [u_t u_t'] = \Sigma_u \times (1 - bx_{t-1}) = \begin{bmatrix} (\bar{\sigma}^{IS})^2 & 0 & 0 & 0 \\ 0 & (\bar{\sigma}^{PC})^2 & 0 & 0 \\ 0 & 0 & (\bar{\sigma}^{MP})^2 & 0 \\ 0 & 0 & 0 & (\bar{\sigma}^*)^2 \end{bmatrix} \times (1 - bx_{t-1}). \quad (19)$$

Equation (19) has two important properties. First, the variances of all shocks in the model, not just the shock to the Euler equation, are proportional to $(1 - bx_{t-1})$, and thus linear in the output gap. This proportionality assumption makes the model relatively tractable and helps us fit the volatilities of bond and stock returns. Second, 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 the *IS* and *PC* shocks captures the notion that all systematic variation in the short-term nominal interest rate is reflected in the monetary policy rule.

2.3 Modeling bonds and stocks

We use the exact loglinear framework of Campbell and Ammer (1993) to express excess log returns on nominal and real bonds as a function of changes in expectations of future short-term interest rates, inflation, and risk premia. 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 write $r_{n-1,t+1}$ for the real one-period log return on a real n-period bond from time t to time $t+1$ and $xr_{n-1,t+1}$ for the corresponding return in excess of r_t . $r_{n-1,t+1}^{\$}$ denotes the nominal one-period return on a similar nominal bond and $xr_{n-1,t+1}^{\$}$ the corresponding excess

return over i_t . We use the identities:

$$\begin{aligned}
r_{n-1,t+1}^{\$} - E_t r_{n-1,t+1}^{\$} &= - (E_{t+1} - E_t) \sum_{j=1}^{n-1} (\hat{i}_{t+j} + \pi_{t+j}^*) \\
&\quad - (E_{t+1} - E_t) \sum_{j=1}^{n-1} x r_{n-j-1,t+1+j}^{\$}
\end{aligned} \tag{20}$$

$$\begin{aligned}
r_{n-1,t+1} - E_t r_{n-1,t+1} &= - (E_{t+1} - E_t) \sum_{j=1}^{n-1} r_{t+j} \\
&\quad - (E_{t+1} - E_t) \sum_{j=1}^{n-1} x r_{n-j-1,t+1+j},
\end{aligned} \tag{21}$$

Nominal bond excess returns reflect shocks to the long-term inflation target, news about the nominal interest rate gap, and news about future nominal bond excess returns. Real long-term excess bond returns reflect news about the real interest rate gap and news about future real bond excess returns.

We model stocks as a levered claim on the log output gap x_t . We assume that log dividends are given by:

$$d_t = \delta x_t. \tag{22}$$

We interpret δ as capturing a broad concept of leverage, including operational leverage. The interpretation of dividends as a levered claim on the underlying fundamental process is common in the asset pricing literature (Abel 1990, Campbell 1986, 2003).

We write r_{t+1}^e for the log stock return and $x r_{t+1}^e$ for the log stock return in excess of r_t . Following Campbell (1991) we use a loglinear approximation to decompose stock returns into

dividend news, news about real interest rates, and news about future excess stock returns ignoring constants:

$$\begin{aligned}
r_{t+1}^e - E_t r_{t+1}^e &= \delta (E_{t+1} - E_t) \sum_{j=0}^{\infty} \rho^j \Delta x_{t+1+j} - (E_{t+1} - E_t) \sum_{j=1}^{\infty} \rho^j r_{t+j} \\
&\quad - (E_{t+1} - E_t) \sum_{j=1}^{\infty} \rho^j x r_{t+1+j}^e.
\end{aligned} \tag{23}$$

Here ρ is a loglinearization constant close to 1.

2.4 Model solution

Denote the vector of state variables by:

$$\hat{Y}_t = [x_t, \hat{\pi}_t, \hat{l}_t]'. \tag{24}$$

The state variable dynamics have a solution of the form

$$\hat{Y}_t = P \hat{Y}_{t-1} + Q u_t. \tag{25}$$

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). We only consider solutions that are real-valued, stable and have finite entries for Q . In principle, the model can have more than one stable solution when there are more than three stable generalized eigenvalues, but in practice there is often a unique solution. We require the model solution be ‘expectationally stable’ (Evans 1985, 1986, McCallum

2003) both in the local sense and globally starting from constant expectations. Intuitively, expectational stability requires that for a deviation from rational expectations, the system returns to the equilibrium. This criterion helps us select a unique solution. For more details on the model solution and solution criteria see the Appendix.

2.4.1 Stochastic discount factor

We can express innovations to log consumption plus habit as:

$$s_{t+1} + c_{t+1} - E_t(s_{t+1} + c_{t+1}) = Q^M u_{t+1}, \quad (26)$$

$$Q^M = e_1 Q - (1 + \theta^*) e_1. \quad (27)$$

In (9) we assumed that the conditional variance of the log SDF is linear in the output gap. Using equation (19) for the changing variance of the shock vector u_{t+1} , we can now verify this assumption with:

$$\sigma_t^2 = Q^M \Sigma_u Q^{M'} (1 - bx_t). \quad (28)$$

The variance of the SDF conditional on a zero output gap is $\bar{\sigma}^2 = Q^M \Sigma_u Q^{M'}$.

2.4.2 Solutions for bond and stock returns

We obtain solutions for unexpected nominal and real bond returns of the form:

$$r_{n-1,t+1}^{\$} - E_t r_{n-1,t+1}^{\$} = A^{\$,n} u_{t+1}, \quad (29)$$

$$r_{n-1,t+1} - E_t r_{n-1,t+1} = A^n u_{t+1}. \quad (30)$$

Up to constant terms, log yields of nominal and real zero coupon bonds equal:

$$y_{n,t}^{\$} = \pi_t^* + B^{\$,n} \hat{Y}_t, \quad (31)$$

$$y_{n,t} = B^n \hat{Y}_t. \quad (32)$$

The loglinear decompositions (20) and (21) for nominal and real bonds are exact, so the solutions for model bond returns and yields (29), (30), (31), and (32) are also exact conditional on our loglinearization of the stochastic discount factor. The vectors $A^{\$,n} \in \mathbb{R}^{1 \times 4}$, $A^n \in \mathbb{R}^{1 \times 4}$, $B^{\$,n} \in \mathbb{R}^{1 \times 3}$, and $B^n \in \mathbb{R}^{1 \times 3}$ are defined recursively.

We obtain that unexpected and expected log excess stock returns take the following loglinear approximate forms:

$$r_{t+1}^e - E_t r_{t+1}^e = A^e u_{t+1}, \quad (33)$$

$$E_t x r_{t+1}^e = (1 - b x_t) b^e, \quad (34)$$

for some $A^e \in \mathbb{R}^{1 \times 4}$ and some $b^e \in \mathbb{R}$.

The online Appendix presents solution details, including bond betas, log dividend price ratios, and multi-period expected log equity returns in excess of short-term T-bills.

2.4.3 An estimable VAR

While standard empirical measures are available for the output gap, we do not observe the interest rate and inflation gaps. We therefore cannot directly estimate the recursive law of motion (25). However, for a long-term bond maturity n , we can estimate a VAR(1) in the vector:

$$Y_t = [x_t, \pi_t, i_t, y_{n,t}^{\$}]'. \quad (35)$$

The model implies that:

$$Y_{t+1} = P^Y Y_t + Q^Y u_{t+1}^Y. \quad (36)$$

Here, $u_t^Y = u_t$ and P^Y and Q^Y are determined by P , Q and the loadings for long-term nominal bond yields $B^{\$,n}$.

3 Data Sources and Summary Statistics

Our empirical analysis uses quarterly US data on output, inflation, interest rates, and aggregate stock returns for the period between the first quarter of 1960 and the fourth quarter

of 2011. Our calibration distinguishes between three subperiods corresponding to major shifts in monetary policy. The first subperiod, 1960.Q1–1979.Q2, covers the Fed chairmanships of William M. Martin, Arthur Burns, and G. William Miller. The second subperiod, 1979.Q3–1996.Q4, covers the Fed chairmanship terms of Paul Volcker and Alan Greenspan. The third subperiod, 1997.Q1–2011.Q4, covers the period after Alan Greenspan’s “Irrational Exuberance” speech at the end of 1996.Q4. Observers of the Fed have often interpreted this widely noted speech as signalling that the Fed became more aware of and responsive to US and international asset market conditions in the late 1990s.⁴ Our first two subperiods are identical to those considered by Clarida, Gali, and Gertler (CGG, 1999), while our third covers data that has become available since their analysis.

Figure 1 shows rolling window nominal bond betas and bond return volatilities, calculated using daily data within each quarter, together with dates of major shifts in monetary policy. The volatility of supply shocks may also have varied over time and contributed to variation in bond risks. We therefore also show dates of major global oil shocks from Hamilton (2009). We can see that the monetary policy breaks line up fairly well with the changes in bond betas. Moreover, two of the oil shocks were very close to the first break in monetary policy, suggesting that the volatility of supply shocks may have been greater during the earlier part of our second subperiod and the later part of the first subperiod. The volatility of bond returns was highest at the beginning of the middle subperiod and lines up roughly with the monetary policy regimes.

⁴“The Challenge of Central Banking in a Democratic Society”, remarks by Chairman Alan Greenspan at the Annual Dinner and Francis Boyer Lecture of The American Enterprise Institute for Public Policy Research, Washington, D.C., December 5, 1996. Available online at <http://www.federalreserve.gov/boarddocs/speeches/1996/19961205.htm>.

We collect data on output, inflation, interest rates, and stock returns from the following sources. GDP in 2005 chained dollars and the GDP deflator are from the Bureau of Economic Analysis via the Fred database at the St. Louis Federal Reserve. The end-of-quarter Federal Funds rate is from the Federal Reserve’s H.15 publication. We use quarterly potential GDP in 2005 chained dollars from the Bureau of Economic Analysis.⁵ The end-of-quarter three-month T-bill is from the CRSP monthly Treasury Fama risk free rates files. We use log yields based on average of bid and ask quotes. 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.⁶ 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.

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 1960-2011 and over each subperiod. 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 have been significantly lower in the latest subperiod compared to the early subperiods.

⁵Table 2-3 of the BEA’s August 2012 report “An Update to the Budget and Economic Outlook: Fiscal Years 2012 to 2022” (<http://www.cbo.gov/publication/43541>)

⁶The source is Robert Shiller’s website at <http://www.econ.yale.edu/shiller/data.htm>.

In our model, expected excess log stock returns vary negatively with the output gap. It is therefore important to verify empirically if this relation exists, and to examine the relation of the output gap with well known predictors of excess stock returns in the data such as the price-dividend ratio.

Figure 2 shows the log output gap and the log price-dividend ratio for the full sample period. The correlation between the two variables in Figure 2 is 0.18 for the full sample period, but it is 0.39 for 1960.Q1–1979.Q3, 0.41 for 1979.Q3–1996.Q4, and 0.74 for 1997.Q1–2011.Q4. The procyclicality of the price-dividend ratio is evident in Figure 2. Occasional long-lasting shifts in the relative levels of the two variables, particularly the secular increase in the price-dividend ratio during the bull market at the end of the 20th Century, decrease the full sample correlation relative to the subsample correlations.

Time-varying volatilities of shocks imply that the equity premium in our model varies inversely with the output gap. We wish to compare time-variation in empirical and model risk premia. Table 2 estimates the predictive relation between quarterly equity excess returns and the output gap:

$$r_{t+1}^e - i_t = a^0 + a^x x_t + \epsilon_{t+1}. \quad (37)$$

Table 2 shows that the full sample estimate of a^x is negative and significant, consistent with our model specification. Subsample estimates vary around the full sample estimate of $a^x = -0.49$.

We also consider the log dividend-price ratio (the negative of the log price-dividend ratio) as a predictor of equity excess returns. Table 3 reports regressions of one through five year

log equity excess returns onto the lagged log dividend-price ratio. We report regressions over our full sample period 1960-2011 and an extended sample period 1947-2011. Table 3 shows that the log dividend-price ratio predicts equity excess returns with positive coefficients. The coefficients in Table 3 are comparable to those reported in Campbell and Cochrane (1999), but smaller in magnitude due to our inclusion of more recent data.

3.1 Estimating monetary policy rules

The central bank's Taylor rule parameters are key inputs for calibrating the model for each subperiod. The model incorporates exclusion restrictions, 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 the monetary policy rule in terms of the output gap, inflation, and the Fed Funds rate:

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

We use the estimated values \hat{c}^x , \hat{c}^π , and \hat{c}^i to pin down the calibrated values of the monetary policy parameters according to: $\hat{\rho}^i = \hat{c}^i$, $\hat{\gamma}^x = \hat{c}^x / (1 - \hat{c}^i)$, and $\hat{\gamma}^\pi = \hat{c}^\pi / (1 - \hat{c}^i)$.

The estimated monetary policy functions in Table 4 yield consistent estimates of the monetary policy parameters only if the inflation target is constant or contemporaneously uncorrelated with the output gap and inflation gaps. We will therefore need to verify that

this bias is quantitatively small in the calibrated model.

The estimates in Table 4 suggest that monetary policy has varied substantially over time. During the earliest subperiod, 1960.Q1–1979.Q2, 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 Volcker-Greenspan-Bernanke periods (1979.Q3–2011.Q4).⁷ This finding is consistent with the empirical evidence reported by CGG and updates it to the most recent period.

The point estimates of $\hat{\gamma}^x$ 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 middle subperiod, although the estimates of neither $\hat{\gamma}^x$ nor \hat{c}^x are statistically significant in the latest subperiod. This empirical finding is similar to Hamilton, Pruitt, and Borger (2010) who use the reaction of Fed Funds futures to macroeconomic announcements to estimate monetary policy rules before and after 2000.

During the most recent period, the monetary policy rule explains 91% of the variation in the Federal Funds rate implying that deviations from the monetary policy rule have been extremely small. Moreover, the loading of the lagged Fed Funds rate into the monetary policy function during this subperiod is very large at 0.89, almost twice as large as in the earlier subperiods. We will see that this increase in estimated policy inertia is important for understanding changing bond risks.

⁷ $\hat{\gamma}^\pi$ appears to be less precisely estimated in the latest subperiod, while \hat{c}^π is precisely estimated. This results from the nonlinear relation between the parameters. The coefficient $\hat{\gamma}^\pi$ is \hat{c}^π 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.

4 Calibration

We now present a calibration matching key empirical moments for the US over the three sub samples: 1960.Q1-1979.Q2, 1979.Q3-1996.Q4, and 1997.Q1-2011.Q4. Table 5 summarizes the calibration parameters, while Tables 6 and 7 compare key empirical and model moments.

Panel A in Table 5 shows the calibration parameters. We separate the parameters into two blocks. We treat the first block of parameter values ρ , δ , α , ρ^π , ρ^{x+} , ρ^{x-} , and λ as deep structural parameters which are time-invariant. In contrast, we allow the second block of parameters to vary across time periods. The second block of parameters comprises the monetary policy rule parameters γ^x , γ^π , and ρ^i and the shock volatilities $\bar{\sigma}^{IS}$, $\bar{\sigma}^{PC}$, $\bar{\sigma}^{MP}$, and $\bar{\sigma}^*$. We use the empirical evidence in Table 4 to pin down the monetary policy parameters γ^x , γ^π , and ρ^i for each subperiod.

We set the Phillips curve slope to $\lambda = 0.30$ following CGG. We set the backward-looking component of the Phillips curve to $\rho^\pi = 0.80$ following the empirical estimates of Fuhrer (1997). Over the full sample period the standard deviations of four quarter real dividend growth and four quarter output gap growth are 5.35% and 2.20%, respectively. We use the ratio of these empirical standard deviations to pin down the model leverage parameter at $\delta = 2.43$, corresponding to a leverage ratio of 59%. We interpret δ as capturing a broad concept of leverage, including operational leverage. We choose the loglinearization parameter ρ as in Campbell and Ammer (1993), scaled to quarterly frequency for a value of 0.99. We set the preference parameter to $\alpha = 30$ to generate plausible equity return volatility.

Let superscript p denote subperiod parameters. The remaining parameters, ρ^{x+} , ρ^{x-} ,

$\sigma^{IS,p}$, $\sigma^{PC,p}$, $\sigma^{MP,p}$, and $\sigma^{*,p}$ for $p = 1, 2, 3$, are chosen to minimize the distance between model and empirical moments. We choose parameters to fit the slope coefficients of a VAR(1) in the log output gap, log inflation, log Fed Funds, and five year nominal log bond yield, the standard deviations of the VAR(1) residuals in annualized percent, equity return volatility and bond return volatility in annualized percent and the nominal bond beta.⁸

The standard deviations of some of the shocks, evaluated at a zero value of the output gap, change considerably across time periods. In particular, we estimate the volatilities of shocks to the MP function and to the inflation target to be significantly larger in the period 1979–1996 than in the earliest and latest subperiods. We also estimate the volatility of shocks to the PC curve to be largest in the earliest subperiod, and the volatility of the shocks to the IS curve to be smallest in the most recent subperiod.

Panel B in Table 5 shows the values of the remaining calibration parameters. These values are implied by the parameter choices shown in Panel A of the table. The calibrated Euler equation has economically significant forward-looking and backward-looking components with a backward-looking component of 0.45 and a forward-looking component of 0.62. The forward- and backward-looking Euler equation components sum to more than one as a result of time-varying risk premia. The parameters imply $\theta = \rho^{x^-} / \rho^{x^+} = 0.73$, so investors' habit moves less than one for one with the lagged output gap.

The slope of the IS curve ψ with respect to the real interest rate is linked to the curvature

⁸The objective function is the sum of squared differences between model and empirical moments summed over all three sub-periods. The equity and bond volatilities are scaled by 0.1 and the nominal bond beta is scaled by a factor of 10 to ensure that moments have roughly equal magnitudes. The online Appendix describes details of the optimization procedure.

parameter that controls risk aversion α and ρ^{x+} by $\psi = \rho^{x+}/\alpha$, as in Section 2.1. The implied slope of the IS curve is 0.02. This value is close to zero, but is in line with the empirical findings in Yogo (2004) and earlier work by Hall (1988).⁹ The parameters that control the volatility of the SDF, b and $\bar{\sigma}$, change little across subperiods.

Table 6 shows calibrated and empirical volatilities of VAR(1) residuals, volatilities of stock and bond returns, nominal bond betas, and Taylor rule regressions of the same form as in Table 4. Model moments are calculated from 2000 simulations of length 250, corresponding closely to our empirical sample size of 61 years of quarterly observations.

The calibrated model provides a close fit for the standard deviations of VAR(1) residuals and stock and bond return volatilities for each of the subperiods. Moreover, the calibrated model fits well the time variation in the nominal bond beta. Both empirical and model nominal bond betas were small but positive in the first subperiod, larger and positive in the second subperiod, and negative during the third subperiod. These changes in bond risks are the primary object of interest in our analysis.

The model's simulated Taylor rules also correspond closely to their empirical counterparts for each subperiod. This finding is reassuring in that it suggests that we can indeed identify time-varying monetary policy parameters from the regressions reported in Table 4 even without estimating the unobservable inflation target.

Our calibration generates additional model moments not used in the fitting procedure. Table 7 shows that many of these moments are reasonably comparable to their empirical

⁹However the long-run risk literature, following Bansal and Yaron (2004), presents an opposing view.

counterparts. The calibration generates reasonable correlations between the output gap and the log price-dividend ratio, especially given that empirical subperiod correlations between these two variables tend to exceed full sample correlations. The log dividend-price ratio is less persistent and less volatile than in the data. The model dividend-price ratio predicts stock excess returns with coefficients that are somewhat larger than the empirical counterparts for the period 1960.Q1–2011.Q4 but closer to the empirical counterparts for a longer sample 1947.Q1–2011.Q4. The model implies that the output gap predicts quarterly stock excess returns with a coefficient which is about six times larger than in the data. The model real short-term rate has a low average standard deviation of 1.90%. The coefficient from regressing $s_t + c_t$ onto the output gap x_t is positive, implying that marginal utility is low when the output gap is high.

5 Counterfactual Analysis of Changing Bond Risks

We next investigate the role of changing monetary policy and macroeconomic shocks for nominal bond betas. Our calibrated model replicates the shift in nominal bond betas over time. Since we allow both monetary policy and the volatilities of shocks to vary across subperiods, both may contribute to time-varying bond risks to varying degrees.

Figure 3 plots nominal bond betas against the monetary policy reaction coefficients γ^x and γ^π . γ^π captures the long-run response of monetary policy to an increase in the inflation gap, while γ^x captures the response to an increase in the output gap. Each panel corresponds to one subperiod. The following parameters vary across panels and equal the calibration values

for the respective subperiod: the persistence of monetary policy ρ^i , and the volatilities of shocks $\bar{\sigma}^{IS}$, $\bar{\sigma}^{PC}$, $\bar{\sigma}^{MP}$, and $\bar{\sigma}^*$. Red and orange colors indicate positive bond betas, while blue and green colors indicate negative bond betas. White areas indicate that no stable solution exists. We show the estimated combinations of monetary policy parameters for each subperiod.

The contour lines in Figure 3 show combinations of reaction coefficients γ^x and γ^π that keep bond betas fixed. These contours are upward-sloping in all three panels, showing that γ^π increases the bond beta while γ^x reduces it. However, the contour lines are relatively flat, which shows that the effect of γ^π on the bond beta is much larger than the effect of γ^x . In addition, the contours shift across panels of the Figure, which tells us that other parameters of the model are also relevant.

Turning to specific results for our three subsamples, panel A of Figure 3 shows that nominal bond betas are positive for a wide range of monetary policy reaction coefficients γ^x and γ^π in the presence of period 1 shock volatilities and period 1 monetary policy persistence. A monetary policy rule with a higher weight on stabilizing inflation γ^π , such as those estimated for our second and third subsamples, would have produced even more positive bond betas than were actually observed during 1960.Q1–1979.Q2.

Panel B shows that in the presence of period 2 shocks and monetary policy persistence, the sign of the nominal bond beta is highly sensitive to the central bank’s weight on inflation stabilization γ^π . The central bank’s strong emphasis on inflation stabilization during 1979.Q3–1996.Q4 is reflected in the strongly positive nominal bond beta during this period. This emphasis continued in period 3, so in the presence of period 2 shock volatilities and

period 2 monetary policy persistence, period 3 monetary policy reaction coefficients would have produced a positive nominal beta, and not the negative beta actually observed.

Panel C indicates that the negative nominal bond beta since the late 1990s cannot be attributed to a change in the monetary policy reaction coefficients γ^π or γ^x . In the presence of period 3 shock volatilities and period 3 monetary policy persistence, the nominal bond beta would have been negative even if the weights on output and inflation stabilization, γ^x and γ^π , had remained constant between periods 2 and 3. It would have been even more negative than actually observed if monetary reaction coefficients in period 3 had been those measured for period 1.

Given the secondary role for the output reaction coefficient γ^x in Figure 3, we replace this parameter with the monetary persistence parameter in Figure 4, which plots nominal bond betas against the monetary policy parameters γ^π and ρ^i . The contour lines in this figure are upward-sloping and convex, becoming extremely steep as monetary persistence approaches its maximum value of one. This shows that γ^π increases the bond beta while ρ^i reduces it, and ρ^i has a nonlinear effect that is stronger at high levels of persistence. This is important given the increase in persistence we have estimated for our final subperiod.

Figure 4 indicates that the strong monetary policy persistence during the most recent subperiod has helped generate a negative nominal bond beta. Moreover, a move to period 3 monetary policy persistence would have acted to decrease the nominal bond beta for all subperiod calibrations. However, a move to period 3 monetary policy persistence and weight on inflation stabilization would not have been sufficient to make the bond beta negative given the other parameters estimated for the first subperiod.

5.1 Impulse responses

To develop further insight into the mechanisms of the model that produce the patterns illustrated in Figures 3 and 4, we now report impulse response functions implied by the model in each subperiod. Figure 5 shows dynamic responses for macroeconomic variables and asset prices following one-standard-deviation increases in each of the fundamental model shocks in period 1 starting from zero. Each panel shows three lines, each corresponding to one subperiod calibration. We consider the responses of the output gap, the inflation gap, the nominal and real interest rates, the dividend-price ratio, and the nominal long-term yield to shocks.

An IS shock has qualitatively similar effects across all three calibrations. In particular, an expansionary IS shock increases both the nominal long-term yield and the dividend-price ratio, because higher expected inflation raises the long-term nominal yield, while slow real growth following an initial increase in the output gap pushes up the dividend-price ratio. This implies that IS shocks tend to generate a positive nominal bond beta for all three estimated monetary policy rules. The magnitudes of the responses of the nominal and real Federal Funds rates i_t and r_t do vary across subperiods. The immediate increase in i_t is strongest for the period 2 calibration, but the period 3 calibration generates the most persistent increases in i_t and r_t .

An inflationary PC shock lowers the output gap, moving output and inflation in opposite directions, in all three subperiods. However, the responses of the short-term interest rate, the dividend-price ratio, and the nominal long-term yield vary across calibrations. The dividend-

price ratio and the nominal long-term yield move in the same direction for the period 1 and 2 calibrations contributing to positive nominal bond betas during those periods. In contrast, for the period 3 calibration a PC shock moves the dividend-price ratio and the long-term nominal yield in opposite directions, contributing to a negative nominal bond beta during that period. For all three monetary policy rules, the nominal short-term interest rate and the long-term yield overshoot their steady state values before eventually returning to steady state. The greater persistence of period 3 monetary policy decreases the immediate response in the Federal Funds rate, so the real short-term interest rate and the long-term nominal yield both fall after an inflationary PC shock.

The negative nominal bond beta in period 3 implies that the nominal-bond risk premium is procyclical. This means that the recession following an inflationary PC shock decreases the risk premium, further lowering the nominal bond yield and amplifying the negative bond beta. This is an example of an important amplification mechanism in our model.

A monetary policy (MP) shock has very small effects on both the dividend price ratio and the nominal yield. Intuitively, our calibration incorporates a small slope of the IS curve, so monetary policy shocks have little effect on the output gap and the stochastic discount factor. In period 3, the standard deviation of the monetary policy shock is small, so the effect of a monetary policy shock is especially small.

A shock to the central bank's inflation target has much more pronounced effects on inflation and nominal interest rates. An inflation target shock leads to permanent increases in inflation and nominal short-term and long-term yields. In the transition period, inflation is below the new target and this negative inflation gap boosts output through the New Keyne-

sian Phillips curve. A positive inflation target shock acts similarly to a supply shock inducing firms to increase supply. Inflation target shocks therefore induce opposite movements in the dividend-price ratio and the long-term nominal yield, implying that inflation target shocks tend to generate a negative nominal bond beta.

Figure 5 helps illustrate why the persistence parameter ρ^i affects the nominal bond beta. ρ^i determines how quickly the monetary policy rule (14) incorporates shocks to the inflation target and the inflation gap. A higher value of ρ^i means that the initial central-bank reaction to PC shocks is smaller, but the initial reaction to inflation target shocks is larger. As a result, even in the absence of risk premia a higher value of ρ^i generates a flat response of long-term yields to PC shocks and a steep response of long-term yields to inflation target shocks. PC shocks tend to increase nominal bond betas but inflation target shocks tend to decrease nominal bond betas. An increase in ρ^i increases the relative importance of inflation target shocks and tends to decrease the nominal bond beta.

The effect of monetary policy persistence ρ^i on the nominal bond beta is compounded by the dynamic behavior of risk premia. If ρ^i is large, then bond returns are largely driven by inflation target shocks and are therefore countercyclical. Since nominal bond returns are then positively correlated with marginal utility, nominal bond yields incorporate low or even negative risk premia during recessions when volatility is high. This procyclicality of bond risk premia amplifies the countercyclicality of bond returns.

5.2 Marginal analysis

Table 8 analyzes the marginal effect of each parameter on the nominal bond beta, the volatility of nominal bond returns, and the volatility of equity returns across subperiods. We provide derivatives with respect to the MP parameters γ^x , γ^π and ρ^i , and also with respect to the log standard deviations of shocks—thus we can interpret the derivatives with respect to the standard deviations of shocks as semi-elasticities. All other parameters are held constant at their 1960.Q1-1979.Q2, 1979.Q3-1996.Q4, or 1997.Q1-2011.Q4 values.

We also decompose the same three asset pricing moments using partial semi-elasticities while holding constant the loadings of bond and stock returns onto the fundamental model shocks A^e and $A^{s,n}$. These partial effects take account of bond and stock responses to each of the fundamental model shocks, similarly to the impulse responses shown in Figure 5, without taking account of alterations in responses caused by changing second moments (for example, the risk premium amplification mechanism discussed in the previous subsection).¹⁰

The top panel in Table 8 shows that the MP inflation coefficient γ^π tends to increase the nominal bond beta, while the MP output gap coefficient γ^x and MP persistence ρ^i tend to decrease the nominal bond beta. Monetary policy persistence has a strongly nonlinear effect on the nominal bond beta, and the magnitude of this effect is particularly large for the

¹⁰The nominal bond beta partial semi-elasticity also holds constant the standard deviation of equity returns. Equivalently, this partial semi-elasticity captures the effect of shock volatilities on the covariance of bond and stock returns scaled by the inverse of a constant equity volatility. The nominal bond beta partial semi-elasticities sum to two times the calibrated nominal bond beta for each subperiod. The partial semi-elasticities for the standard deviations of asset returns sum to the calibrated standard deviation of asset returns for each subperiod. The online Appendix presents detailed formulas for semi-elasticities and partial semi-elasticities.

1997.Q1–2011.Q4 calibration. The nominal bond beta increases in the PC shock volatility and decreases in the inflation target shock volatility. These signs are consistent with the partial-semi elasticities but are substantially amplified by endogenous responses of risk premia. IS and MP shock volatilities have small and negative effects on the nominal bond beta.

The middle panel of Table 8 shows that the effect of parameters on bond return volatility is especially nonlinear, with total derivatives switching sign exactly when the nominal bond beta switches sign. But partial semi-elasticities with respect to shock volatilities, which ignore the endogenous responses of risk premia, are all positive. The partial semi-elasticity with respect to PC shock volatility is especially important in the middle subperiod, while the partial semi-elasticity with respect to inflation target shocks is especially important in the most recent subperiod.

The third panel of Table 8 looks at equity volatility. The volatility of the PC shock is clearly the single most important driver of equity volatility in the model, followed by the inflation target shock volatility.

Table 9 uses the semi-elasticities of Table 8 to decompose the changes in bond and equity risks across the three subperiods. We report linear approximations of how much the change in each monetary policy parameter and in each shock volatility has contributed to changes in the nominal bond beta, nominal bond return volatility, and equity return volatility. Table 9 weights total derivatives as reported in Table 8 by the corresponding parameter change from one period to the next. Total derivatives can vary across subperiods and we therefore average total derivatives across lagged and led periods.

Table 9 also reports the total linearized change in bond and equity risks due to the combined change in monetary policy parameters γ^x , γ^π , ρ^i , $\bar{\sigma}^{MP}$, and $\bar{\sigma}^*$ and due to the combined changes in the volatilities of supply and demand shocks $\bar{\sigma}^{IS}$, and $\bar{\sigma}^{MP}$. Comparing the sum of linearized changes in bond and equity risks and the model-implied changes in bond and equity risks gives the change due to model nonlinearity. The linear effects of the individual parameters and the nonlinearity effect sum to the total model change reported at the top of Table 9.

Table 9 shows that the reduction in the central bank's response to output, γ^x , and the increase in the central bank's response to inflation, γ^π , were important contributors to the increases in the nominal bond beta and the volatility of bond returns that occurred in 1979. On the other hand, changes in the volatility of supply (PC) shocks and inflation target shocks acted to decrease the nominal bond beta and the volatility of bond returns at the 1979 regime change. Thus monetary policy changes and shock volatility changes offset each other to some degree in 1979.

In 1997, the increase in monetary policy persistence is most important for understanding the decline in the nominal bond beta, but the the increase in the central bank's reaction to output also plays a role. Once again changes in the volatilities of macroeconomic shocks worked against these changes and offset them to some degree. The bottom of the table shows that nonlinear interaction effects, for example between the persistence of monetary policy and the volatility of inflation target shocks, are important at both regime changes but particularly so in 1997.

6 Conclusion

Given the importance of nominal 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 nominal bond risks. This is particularly challenging because the risk characteristics of nominal 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 and demand shocks and central bank responses to those shocks affect asset prices. We propose a model that integrates the building blocks of a New Keynesian model into an asset pricing framework in which risk and consequently 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 discrete regime changes before the third quarter of 1979 and the first quarter of 1997.

Our model is sufficiently rich to allow for a detailed exploration of the monetary policy drivers of bond and equity risks. We find that two 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 1979 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, a monetary policy that smooths nominal interest rates over time implies that positive shocks to long-term target inflation cause real interest rates to fall, driving up output and equity prices, while increasing nominal long-term interest rates. This makes nominal bond returns countercyclical, implying a negative risk premium because nominal bonds hedge against deflationary recessions. Our model attributes the negative beta of nominal bonds since 1997 to a significant increase in the persistence of monetary policy, together with continuing shocks to the central bank's inflation target. These inflation target shocks may be interpreted literally, as the result of shifting central bank preferences, or more broadly as the result of imperfect credibility in monetary policy (Orphanides and Williams 2004).

Our model also indicates that changes in the volatility of supply shocks (that is, shocks to the Phillips curve) can affect bond risks. Supply shocks move inflation and output in opposite directions, making bond returns procyclical. They also have a strong effect on the volatility of equity returns. However, we do not estimate large changes over time in the volatility of supply shocks and so our model does not attribute the historical changes in bond betas to this source.

Our results imply that it is particularly important to take account of changing risk premia. Because macroeconomic volatility is 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 are originated by changes in monetary policy.

Our analysis has several limitations that can be addressed in future research. 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. 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). The regime shifts we consider are unanticipated, once-and-for-all events rather than stochastically recurring events whose probabilities are understood by market participants. Our model makes predictions about real bonds, but Treasury inflation-protected securities (TIPS) were issued only during our last subsample (Campbell, Shiller, and Viceira 2009) so we do not use evidence on TIPS in our empirical work. 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

- Abel, Andrew B., 1990, "Asset Prices under Habit Formation and Catching up with the Joneses", *American Economic Review Papers and Proceedings* 80, 38–42.
- An, Sungbae, and Frank Schorfheide, 2007, "Bayesian Analysis of DSGE Models", *Econometric Reviews* 26, 113–172.
- Andreasen, Martin M., 2012, "An Estimated DSGE Model: Explaining Variation in Nominal Term Premia, Real Term Premia, and Inflation Risk Premia", *European Economic Review* 56, 1656–1674.
- Ang, Andrew, Jean Boivin, Sen Dong, and Rudy Loo-Kung, 2011, "Monetary Policy Shifts and the Term Structure", *Review of Economic Studies* 78, 429–457.
- Ang, Andrew and Monika Piazzesi, 2003, "A No-Arbitrage Vector Autoregression of Term Structure Dynamics with Macroeconomic and Latent Variables", *Journal of Monetary Economics* 50, 745–787.
- Ang, Andrew, Sen Dong, and Monika Piazzesi, 2007, "No-Arbitrage Taylor Rules", unpublished paper, Columbia University and University of Chicago.
- Baele, Lieven, Geert Bekaert, and Koen Inghelbrecht, 2010, "The Determinants of Stock and Bond Return Comovements", *Review of Financial Studies* 23, 2374–2428.
- Ball, Laurence and Stephen Cecchetti, 1990, "Inflation and Uncertainty at Long and Short Horizons", *Brookings Papers on Economic Activity* 1, 215–245.
- Bansal, Ravi and Ivan Shaliastovich, 2010, "A Long-Run Risks Explanation of Predictability Puzzles in Bond and Currency Markets", unpublished paper, Duke University.
- Bansal, Ravi, and Amir Yaron, 2004, "Risks for the Long Run: A Potential Resolution of Asset Pricing Puzzles," *Journal of Finance* 59, 1481–1509.
- Barsky, Robert B., 1989, "Why Don't the Prices of Stocks and Bonds Move Together?", *American Economic Review* 79, 1132–1145.
- Bekaert, Geert, Seonghoon Cho, and Antonio Moreno, 2010, "New Keynesian Macroeconomics and the Term Structure", *Journal of Money, Credit, and Banking* 42, 33–62.
- Bekaert, Geert, Eric Engstrom, and Steve Grenadier, 2010, "Stock and Bond Returns with Moody Investors", *Journal of Empirical Finance* 17, 867–894.

- Bekaert, Geert, Eric Engstrom, and Yuhang Xing, 2009, “Risk, Uncertainty, and Asset Prices”, *Journal of Financial Economics* 91, 59–82.
- Bikbov, Ruslan and Mikhail Chernov, 2013, “Monetary Policy Regimes and the Term Structure of Interest Rates”, *Journal of Econometrics* 174, 27–43.
- van Binsbergen, Jules H., Jesús Fernández-Villaverde, Ralph Koijen, and Juan Rubio-Ramírez, 2012, “The Term Structure of Interest Rates in a DSGE Model with Recursive Preferences”, *Journal of Monetary Economics* 59, 634–648.
- Blanchard, Olivier and Jordi Gali, 2007, “The Macroeconomic Effects of Oil Shocks: Why Are the 2000s So Different From the 1970s?”, NBER Working Paper No. 13368.
- Burkhardt, Dominic and Henrik Hasseltoft, 2012, “Understanding Asset Correlations”, unpublished paper, University of Zurich and Swiss Finance Institute.
- Boivin, Jean, and Marc P. Giannoni, 2006, “Has Monetary Policy Become More Effective?”, *Review of Economics and Statistics* 88, 445–462.
- Buraschi, Andrea and Alexei Jiltsov, 2005, “Inflation Risk Premia and the Expectations Hypothesis”, *Journal of Financial Economics* 75, 429–490.
- Buraschi, Andrea and Alexei Jiltsov, 2007, “Habit Formation and Macroeconomic Models of the Term Structure of Interest Rates”, *Journal of Finance* 62, 3009–3063.
- Calvo, Guillermo, 1983, “Staggered Prices in a Utility Maximizing Framework”, *Journal of Monetary Economics* 12, 383–98.
- Campbell, John Y., 1986, “Bond and Stock Returns in a Simple Exchange Model”, *Quarterly Journal of Economics* 101, 785–803.
- Campbell, John Y., 1991, “A Variance Decomposition for Stock Returns”, *Economic Journal* 101:157–179.
- Campbell, John Y., 2003, “Consumption-Based Asset Pricing”, Chapter 13 in George Constantinides, Milton Harris, and Rene Stulz eds. *Handbook of the Economics of Finance Vol. IB*, North-Holland, Amsterdam, 803–887.
- Campbell, John Y. and John Ammer, 1993, “What Moves the Stock and Bond Markets? A Variance Decomposition for Long-Term Asset Returns”, *Journal of Finance* 48, 3–37.

- Campbell, John Y. and John H. Cochrane, 1999, “By Force of Habit: A Consumption-Based Explanation of Aggregate Stock Market Behavior”, *Journal of Political Economy* 107, 205–251.
- Campbell, John Y., Carolin Pflueger, and Luis Viceira, 2013, “Appendix to Monetary Policy Drivers of Bond and Equity Risks”, online at <http://scholar.harvard.edu/campbell/publications>.
- Campbell, John Y., Robert J. Shiller, and Luis M. Viceira, 2009, “Understanding Inflation-Indexed Bond Markets”, *Brookings Papers on Economic Activity* 79-120, Spring 2009.
- Campbell, John Y., Adi Sunderam, and Luis M. Viceira, 2013, “Inflation Bets or Deflation Hedges? The Changing Risks of Nominal Bonds”, unpublished paper, Harvard University.
- Canova, Fabio, and Luca Canova, 2009, “Back to Square One: Identification Issues in DSGE Models” *Journal of Monetary Economics*, 56, 431–449.
- Chib, Siddhartha, Kyu Ho Kang, and Srikanth Ramamurthy, 2010, “Term Structure of Interest Rates in a DSGE Model with Regime Changes”, unpublished paper, Washington University in St. Louis.
- Christiano, Lawrence J., Martin Eichenbaum, and Charles L. Evans, 1999, “Monetary Policy Shocks: What Have We Learned and to What End?”, in John B. Taylor and Michael Woodford (eds.) *Handbook of Macroeconomics 1999*, 65–148, Elsevier.
- Christiano, Lawrence J., Martin Eichenbaum, and Charles L. Evans, 2005, “Nominal Rigidities and the Dynamic Effects of a Shock to Monetary Policy”, *Journal of Political Economy* 113, 1–45.
- Christiansen, Charlotte and Angelo Ranaldo, 2007, “Realized Bond-Stock Correlation: Macroeconomic Announcement Effects”, *Journal of Futures Markets* 27, 439–469.
- Clarida, Richard, Jordi Gali, and Mark Gertler, 1999, “The Science of Monetary Policy: A New Keynesian Perspective”, *Journal of Economic Literature* 37, 1661–1707.
- Coibion, Olivier and Yuriy Gorodnichenko, 2012, “Why Are Target Interest Rate Changes So Persistent?”, *American Economic Journal: Macroeconomics* 4, 126–162.
- d’Addona, Stefano and Axel H. Kind, 2006, “International Stock-Bond Correlations in a Simple Affine Asset Pricing Model”, *Journal of Banking and Finance* 30, 2747–2765.

- Dai, Qiang and Kenneth Singleton, 2000, “Specification Analysis of Affine Term Structure Models”, *Journal of Finance* 55, 1943–1978.
- Dai, Qiang and Kenneth Singleton, 2002, “Expectations Puzzles, Time-Varying Risk Premia, and Affine Models of the Term Structure”, *Journal of Financial Economics* 63, 415–442.
- David, Alexander and Pietro Veronesi, 2013, “What Ties Return Volatilities to Fundamentals and Price Valuations?”, forthcoming *Journal of Political Economy*.
- Duffee, Greg, 2002, “Term Premia and Interest Rate Forecasts in Affine Models”, *Journal of Finance* 57, 405–443.
- Duffie, Darrell and Rui Kan, 1996, “A Yield-Factor Model of Interest Rates”, *Mathematical Finance* 6, 379–406.
- Eraker, Bjørn, 2008, “Affine General Equilibrium Models”, *Management Science* 54, 2068–2080.
- Evans, George, 1985, “Expectational Stability and the Multiple Equilibria Problem in Linear Rational Expectations”, *The Quarterly Journal of Economics* 100, 1217–1233.
- Evans, George, 1986, “Selection Criteria for Models with Non-Uniqueness”, *Journal of Monetary Economics* 18, 147–157.
- Fama, Eugene F. and Kenneth R. French, 1989, “Business Conditions and Expected Returns on Stocks and Bonds”, *Journal of Financial Economics* 25, 23–50.
- Fuhrer, Jeffrey C., 1997, “The (Un)Importance of Forward-Looking Behavior in Price Specifications”, *Journal of Money, Credit, and Banking* 29, 338–350.
- Fuhrer, Jeffrey C., 2000, “Habit Formation in Consumption and Its Implications for Monetary Policy Models”, *American Economic Review* 90, 367–390.
- Gallmeyer, Michael, Burton Hollifield, Francisco Palomino, and Stanley Zin, 2008, “Term Premium Dynamics and the Taylor Rule”, unpublished paper, Texas A&M University, Carnegie Mellon University, and University of Michigan.
- Guidolin, Massimo, and Allan Timmermann, 2007, “Asset Allocation under Multivariate Regime Switching”, *Journal of Economic Dynamics and Control* 31, 3503–3544.

- Hall, Robert E., 1988, “Intertemporal Substitution in Consumption”, *Journal of Political Economy* 96, 221–273.
- Hamilton, James, 2009, “Causes and Consequences of the Oil Shock of 2007-08”, *Brookings Papers on Economic Activity* 215–261, Spring 2009.
- Hamilton, James, Seth Pruitt, and Scott Borger, 2010, “Estimating the Market-Perceived Monetary Policy Rule”, NBER Working Paper 16412.
- Hasseltoft, Henrik, 2008, “The ‘Fed-Model’ and the Changing Correlation of Stock and Bond Returns: An Equilibrium Approach”, unpublished paper, Stockholm School of Economics.
- Kung, Howard, 2013, “A Macroeconomic Foundation for the Equilibrium Term Structure of Interest Rates”, unpublished paper, University of British Columbia.
- Lettau, Martin and Jessica Wachter, 2011, “The Term Structures of Equity and Interest Rates”, *Journal of Financial Economics* 101, 90–113.
- Li, Erica X.N. and Francisco Palomino, 2013, “Nominal Rigidities, Asset Returns, and Monetary Policy”, unpublished paper, Cheung Kong Graduate School of Business and University of Michigan.
- Mankiw, N. Gregory, and Ricardo Reis, 2002, “Sticky Information Versus Sticky Prices: A Proposal to Replace the New Keynesian Phillips Curve”, *Quarterly Journal of Economics* 177, 1295–1328.
- McCallum, Bennett T., 2003, “Multiple-Solution Indeterminacies in Monetary Policy Analysis”, *Journal of Monetary Economics* 50, 1153–1175.
- Orphanides, Athanasios and John Williams, 2004, “Imperfect Knowledge, Inflation Expectations, and Monetary Policy”, Chapter 5 in Ben S. Bernanke and Michael Woodford eds. *The Inflation-Targeting Debate*, University of Chicago Press.
- Palomino, Francisco, 2012, “Bond Risk Premiums and Optimal Monetary Policy”, forthcoming *Review of Economic Dynamics*.
- Piazzesi, Monika and Martin Schneider, 2006, “Equilibrium Yield Curves”, in D. Acemoglu, K. Rogoff, and M. Woodford (eds.) *NBER Macroeconomics Annual 2006*, 317-363, MIT Press, Cambridge, MA.

- Rigobon, Roberto and Brian Sack, 2003, “Measuring the Reaction of Monetary Policy to the Stock Market”, *Quarterly Journal of Economics* 639–669.
- Rudebusch, Glenn D., 2002, “Term Structure Evidence on Interest Rate Smoothing and Monetary Policy Inertia”, *Journal of Monetary Economics* 49, 1161–1187.
- Rudebusch, Glenn D. and Eric T. Swanson, 2012, “The Bond Premium in a DSGE Model with Long-Run Real and Nominal Risks”, *American Economic Journal: Macroeconomics* 4, 105–143.
- Rudebusch, Glenn D. and Tao Wu, 2007, “Accounting for a Shift in Term Structure Behavior with No-Arbitrage and Macro-Finance Models”, *Journal of Money, Credit, and Banking* 39, 395–422.
- Rudebusch, Glenn D. and Tao Wu, 2008, “A Macro-Finance Model of the Term Structure, Monetary Policy, and the Economy”, *Economic Journal* 118, 906–926.
- Shiller, Robert J. and Andrea Beltratti, 1992, “Stock Prices and Bond Yields: Can Their Comovements Be Explained in Terms of Present Value Models?”, *Journal of Monetary Economics* 30, 25–46.
- Smets, Frank, and Rafael Wouters, 2010, “Shocks and Frictions in US Business Cycles: A Bayesian DSGE Approach”, *American Economic Review* 97, 586–606.
- Smith, Josephine M. and John B. Taylor, 2009, “The Term Structure of Policy Rules”, *Journal of Monetary Economics* 56, 907–917.
- Stock, James H. and Mark Watson, 2007, “Why Has US Inflation Become Harder to Forecast?”, *Journal of Money Credit and Banking* 39, 3–33.
- Taylor, John B., 1993, “Discretion versus Policy Rules in Practice”, *Carnegie-Rochester Conference Series on Public Policy* 39, 195–214.
- Uhlig, Harald, 1999, “A Toolkit for Analyzing Nonlinear Dynamic Stochastic Models Easily”, in Ramon Marimon and Andrew Scott (eds.) *Computational Methods for the Study of Dynamic Economies*, 30–61, Oxford University Press.
- Viceira, Luis M., 2010, “Bond Risk, Bond Return Volatility, and the Term Structure of Interest Rates”, forthcoming *International Journal of Forecasting*.
- Wachter, Jessica A., 2006, “A Consumption-Based Model of the Term Structure of Interest Rates”, *Journal of Financial Economics* 79, 365–399.

- Woodford, Michael, 2001, “The Taylor Rule and Optimal Monetary Policy”, *American Economic Review Papers and Proceedings* 91, 232–237.
- Yogo, Motohiro, 2004, “Estimating the Elasticity of Intertemporal Substitution When Instruments Are Weak”, *Review of Economics and Statistics* 86, 797–810.

Tables and Figures

Table 1: **Summary Statistics**

1960.Q1-2011.Q4	Output Gap	Inflation	Fed Funds	Nom. Bond Yield
Mean	-0.77	3.50	5.78	6.24
Std.	2.91	2.38	3.77	2.75
AR(1) Coefficient	0.96	0.88	0.89	0.98
	(0.02)	(0.03)	(0.03)	(0.02)
AR(4) Coefficient	0.73	0.79	0.81	0.91
	(0.05)	(0.04)	(0.04)	(0.03)
1960.Q1-1979.Q2	Output Gap	Inflation	Fed Funds	Nom. Bond Yield
Mean	0.64	4.38	5.27	5.80
Std.	2.79	2.72	2.45	1.63
AR(1) Coefficient	0.93	0.88	0.86	0.96
	(0.04)	(0.06)	(0.06)	(0.35)
AR(4) Coefficient	0.61	0.78	0.60	0.92
	(0.09)	(0.07)	(0.10)	(0.06)
1979.Q3-1996.Q4	Output Gap	Inflation	Fed Funds	Nom. Bond Yield
Mean	-1.92	3.76	8.67	8.73
Std.	2.14	2.24	4.04	2.55
AR(1) Coefficient	0.93	0.87	0.78	0.94
	(0.04)	(0.05)	(0.08)	(0.04)
AR(4) Coefficient	0.50	0.78	0.74	0.77
	(0.10)	(0.06)	(0.08)	(0.08)
1997.Q1-2011.Q4	Output Gap	Inflation	Fed Funds	Nom. Bond Yield
Mean	-1.27	2.06	3.08	3.90
Std.	3.13	1.01	2.32	1.51
AR(1) Coefficient	1.00	0.51	0.95	0.94
	(0.03)	(0.12)	(0.04)	(0.05)
AR(4) Coefficient	0.89	0.28	0.78	0.80
	(0.09)	(0.13)	(0.09)	(0.08)

Full sample and sub period summary statistics. US quarterly log output gap (%), GDP deflator inflation (% Annualized), Fed Funds rate (% Annualized), and 5 year nominal yield (% Annualized). Yields and inflation continuously compounded. Standard errors in parentheses.

Table 2: **Predicting Stock Returns with Output Gap**

Log Exc. Stock Ret. xr_{t+1}^e	60.Q1-11.Q4	60.Q1-79.Q2	79.Q2-96.Q4	97.Q1-11.Q4
Output Gap x_t	-0.49*	-0.61*	-0.32	-0.47
	(0.20)	(0.30)	(0.48)	(0.40)
Constant	0.59	0.81	1.12	0.18
	(0.61)	(0.95)	(1.14)	(1.36)
R^2	0.03	0.04	0.01	0.02

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.

Table 3: **Predicting Stock Returns with Dividend Price Ratio**

Log Exc. Stock Ret. $xr_{t \rightarrow t+k}^e$	k=4	k=8	k=12	k=16	k=20
Log Dividend Price Ratio $d_t - p_t$	0.08	0.15	0.19*	0.21**	0.25**
	(0.05)	(0.09)	(0.09)	(0.08)	(0.06)
Constant	0.34	0.62*	0.79*	0.91**	1.10**
	(0.18)	(0.29)	(0.31)	(0.27)	(0.23)
R^2	0.04	0.07	0.09	0.10	0.12
Sample	1960.Q1-2011.Q1				
$xr_{t \rightarrow t+k}^e$	k=4	k=8	k=12	k=16	k=20
Log Dividend Price Ratio $d_t - p_t$	0.11*	0.21**	0.28**	0.34**	0.43**
	(0.04)	(0.07)	(0.08)	(0.09)	(0.10)
Constant	0.44**	0.82**	1.13**	1.40**	1.76**
	(0.15)	(0.24)	(0.27)	(0.29)	(0.33)
R^2	0.08	0.15	0.22	0.27	0.33
Sample	1947.Q1-2011.Q4				

k quarter log excess stock returns onto lagged log dividend price ratio, both in natural units. Newey-West standard errors with $k + 4$ lags in parentheses. * and ** denote significance at the 1% and 5% levels.

Table 4: **Estimating the Monetary Policy Function**

Fed Funds i_t	60.Q1-11.Q4	60.Q1-79.Q2	79.Q3-96.Q4	97.Q1-11.Q4
Output Gap x_t	0.06 (0.04)	0.18** (0.06)	-0.04 (0.13)	0.05 (0.04)
Inflation π_t	0.21 (0.11)	0.30** (0.07)	0.83** (0.21)	0.21** (0.07)
Lagged Fed Funds i_{t-1}	0.81** (0.05)	0.56** (0.10)	0.43* (0.17)	0.89** (0.06)
Constant	0.42 (0.26)	0.91* (0.38)	1.75 (0.92)	-0.12 (0.29)
R^2	0.79	0.75	0.69	0.91
Implied $\hat{\gamma}^x$	0.32 (0.21)	0.42** (0.13)	-0.07 (0.22)	0.44 (0.21)
Implied $\hat{\gamma}^\pi$	1.08** (0.43)	0.69** (0.16)	1.44** (0.19)	1.92* (1.26)
Implied $\hat{\rho}^i$	0.81** (0.05)	0.56** (0.10)	0.43** (0.17)	0.89** (0.06)

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. Implied 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 5: Parameter Choices

Panel A: Calibration Parameters

Time-Invariant Parameters				
Log-Linearization Constant	ρ			0.99
Leverage	δ			2.43
Preference Parameter	α			30
Backward-Looking Comp. PC	ρ^π			0.80
Slope PC	λ			0.30
Forward-Looking Comp. IS	ρ^{x+}			0.62
Backward-Looking Comp. IS	ρ^{x-}			0.45
Monetary Policy Rule				
		60.Q1-79.Q2	79.Q3-96.Q4	97.Q1-11.Q4
MP Coefficient Output	γ^x	0.42	-0.07	0.44
MP Coefficient Infl.	γ^π	0.69	1.44	1.92
Backward-Looking Comp. MP	ρ^i	0.56	0.43	0.89
Std. Shocks				
Std. IS	$\bar{\sigma}^{IS}$	0.45	0.43	0.26
Std. PC shock	$\bar{\sigma}^{PC}$	1.08	0.80	0.93
Std. MP shock	$\bar{\sigma}^{MP}$	1.04	2.03	0.26
Std. infl. target shock	$\bar{\sigma}^*$	0.37	0.70	0.53
Panel B: Implied Parameters				
Time-Invariant Implied Parameters				
SDF Lag Output Gap	θ			0.72
SDF Lag with Varying Risk Premia	θ^*			0.61
Slope IS	ψ			0.02
Time-Varying Implied Parameters				
Heteroskedasticity Parameter	b	7.18	8.27	7.89
Volatility SDF	σ	0.32	0.30	0.31

Table 6: Model and Empirical Moments

Std. VAR(1) Residuals	60.Q1-79.Q2		79.Q3-96.Q4		97.Q1-11.Q4	
	Empirical	Model	Empirical	Model	Empirical	Model
Output Gap	0.92	0.81	0.75	0.77	0.65	0.51
Inflation	1.12	1.22	0.89	0.92	0.80	1.02
Fed Funds Rate	1.22	1.15	2.07	2.17	0.66	0.55
Log Nominal Yield	0.48	0.41	0.85	0.72	0.55	0.70
Std. Asset Returns						
Std. Eq. Ret.	17.62	18.82	15.34	17.50	20.08	17.95
Std. Nom. Bond Ret.	4.85	3.91	9.11	7.09	5.55	5.83
Nominal Bond Beta	0.06	0.07	0.20	0.22	-0.17	-0.15
Taylor Rule: Fed Funds onto Output, Infl. and Lag. Fed Funds						
Output	0.18	0.21	-0.04	-0.12	0.05	0.01
Inflation	0.30	0.33	0.83	0.65	0.21	0.19
Lagged Fed Funds	0.56	0.61	0.43	0.39	0.89	0.80

Table 7: **Additional Empirical and Model Moments**

	Empirical	Model
AR(1) Coefficient Output Gap	0.96	0.83
AR(4) Coefficient Output Gap	0.73	0.18
Correlation(x, p-d)	0.18	0.47
Std(d-p)	0.40	0.15
AR(4) Coefficient d-p	0.92	0.11
Slope 1 year exc. Stock ret. wrt d-p	0.08	0.84
Slope 5 year exc. Stock ret. wrt d-p	0.25	1.03
Slope quarterly stock ret. wrt x	-0.49	-3.30
Std(real rate) (% Ann.)		1.90
Regression s+c onto x		0.23

Additional model and empirical moments are not explicitly fitted by the calibration procedure. Model moments show averages across three sub sample calibrations weighted by sub sample length.

Table 8: Marginal Effects of Parameters

		Total Derivative			Partial Derivative		
		60.Q1-79.Q2	79.Q3-96.Q4	97.Q1-11.Q4	60.Q1-79.Q2	79.Q3-96.Q4	97.Q1-11.Q4
Nominal Bond Beta							
MP Coefficient Output	γ^x	-3.92	-1.50	-1.37			
MP Coefficient Inflation	γ^π	5.01	1.80	1.86			
MP Persistence	ρ^i	-1.85	-1.91	-20.90			
IS Shock Std.	$\bar{\sigma}^{IS}$	-0.56	-0.11	-0.09	-0.01	-0.01	-0.01
PC Shock Std.	$\bar{\sigma}^{PC}$	3.43	3.87	5.25	0.22	0.58	-0.05
MP Shock Std.	$\bar{\sigma}^{MP}$	-0.28	-0.33	-0.06	0.01	0.02	0.00
Infl. Target Shock Std.	$\bar{\sigma}^*$	-2.59	-3.42	-5.09	-0.08	-0.15	-0.24
Std. Nominal Bond Returns							
MP Coefficient Output	γ^x	-16.06	-11.72	14.20			
MP Coefficient Inflation	γ^π	20.94	14.35	-19.04			
MP Persistence	ρ^i	-6.98	-14.48	215.46			
IS Shock Std.	$\bar{\sigma}^{IS}$	-1.99	-0.28	1.81	0.22	0.42	0.89
PC Shock Std.	$\bar{\sigma}^{PC}$	14.49	32.35	-52.90	1.24	5.55	0.05
MP Shock Std.	$\bar{\sigma}^{MP}$	-0.83	-2.30	0.78	0.35	0.42	0.10
Infl. Target Shock Std.	$\bar{\sigma}^*$	-7.76	-22.68	56.14	2.11	0.71	4.80
Std. Equity Returns							
MP Coefficient Output	γ^x	-1.59	-1.29	-1.12			
MP Coefficient Inflation	γ^π	0.77	0.75	0.66			
MP Persistence	ρ^i	-0.20	-0.32	-1.70			
IS Shock Std.	$\bar{\sigma}^{IS}$	0.23	0.05	0.00	0.24	0.07	0.01
PC Shock Std.	$\bar{\sigma}^{PC}$	17.41	11.87	14.89	17.24	11.63	14.78
MP Shock Std.	$\bar{\sigma}^{MP}$	-0.08	-0.16	-0.06	0.08	0.18	0.06
Infl. Target Shock Std.	$\bar{\sigma}^*$	1.27	5.74	3.11	1.26	5.63	3.09

Derivatives with respect to monetary policy rule parameters and log standard deviations of shocks (semi-elasticities). Partial derivatives hold constant the loadings of bond and stock returns. Nominal bond beta partial derivatives also hold equity volatility constant.

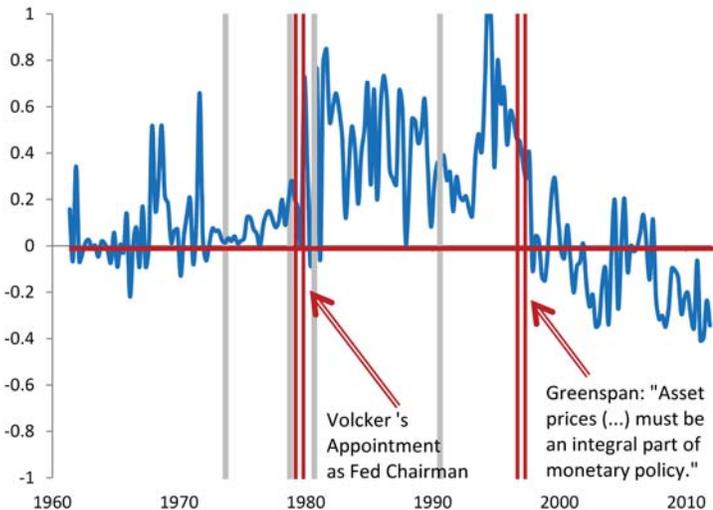
Table 9: **Decomposing Changes in Bond and Equity Risks**

Change Date	Nominal Bond Beta		Std. Nom. Bond Returns		Std. Equity Returns		
	79.Q3	97.Q1	79.Q3	97.Q1	79.Q3	97.Q1	
Empirical Change	0.13	-0.36	4.26	-3.56	-2.28	4.74	
Model Change	0.15	-0.37	3.18	-1.27	-1.32	0.45	
Total Derivative x Parameter Change							
MP Coefficient Output	γ^x	1.35	-0.74	6.83	0.12	0.71	-0.62
MP Coefficient Inflation	γ^π	2.63	0.86	13.43	-0.50	0.57	0.34
MP Persistence	ρ^i	0.26	-4.97	1.44	42.64	0.04	-0.44
MP Shock Std.	$\bar{\sigma}^{MP}$	-0.20	0.42	-1.01	1.80	-0.08	0.23
Infl. Target Shock Std.	$\bar{\sigma}^*$	-1.92	1.14	-9.55	-3.72	2.18	-1.23
IS Shock Std.	$\bar{\sigma}^{IS}$	0.01	0.05	0.04	-0.36	0.00	-0.01
PC Shock Std.	$\bar{\sigma}^{PC}$	-1.10	0.67	-6.93	-1.04	-4.47	1.97
Combined Effects							
MP Sub-Total		2.11	-3.28	11.14	40.34	3.42	-1.73
IS&PC Shocks Sub-Total		-1.09	0.72	-6.90	-1.40	-4.47	1.95
Total Linear Changes		1.03	-2.56	4.24	38.94	-1.06	0.22
Nonlinearity Effect		-0.87	2.19	-1.06	-40.21	-0.26	0.23

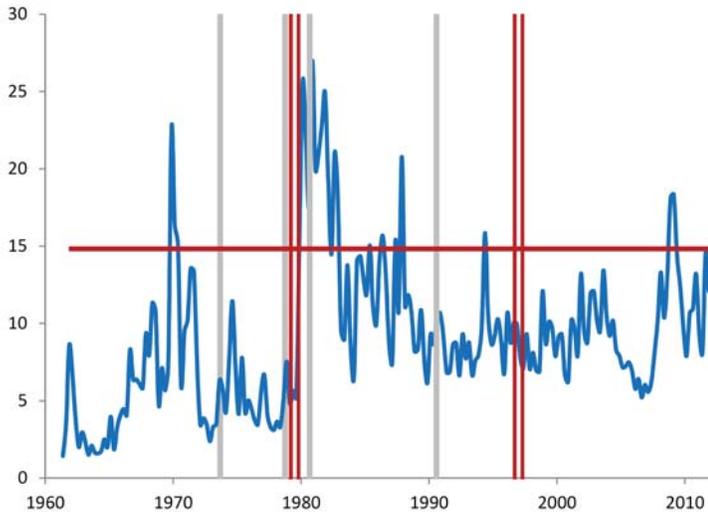
We average total derivatives across subsequent periods weighting sub periods according to their sample size. We multiply average weighted total derivatives by the parameter change between subsequent periods. Total derivatives for each sub period are reported in Table 8 and parameter values for each sub period are reported in Table 5. The “MP Sub-Total” row sums the linear effects of γ^x , γ^π , ρ^i , $\bar{\sigma}^{MP}$, and $\bar{\sigma}^*$. The “IS&PC Shocks Sub-Total” row sums the linear effects of $\bar{\sigma}^{IS}$ and $\bar{\sigma}^{PC}$. The row “Total Linear Changes” reports the sum of all linear effects reported in the panel above. The row “Nonlinearity Effect” shows the difference between the model change and the total of linear changes.

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

Panel A: CAPM Beta of 10 YR Nominal Bond



Panel B: Std. of 10 YR Nominal Bond Returns (% Ann.)



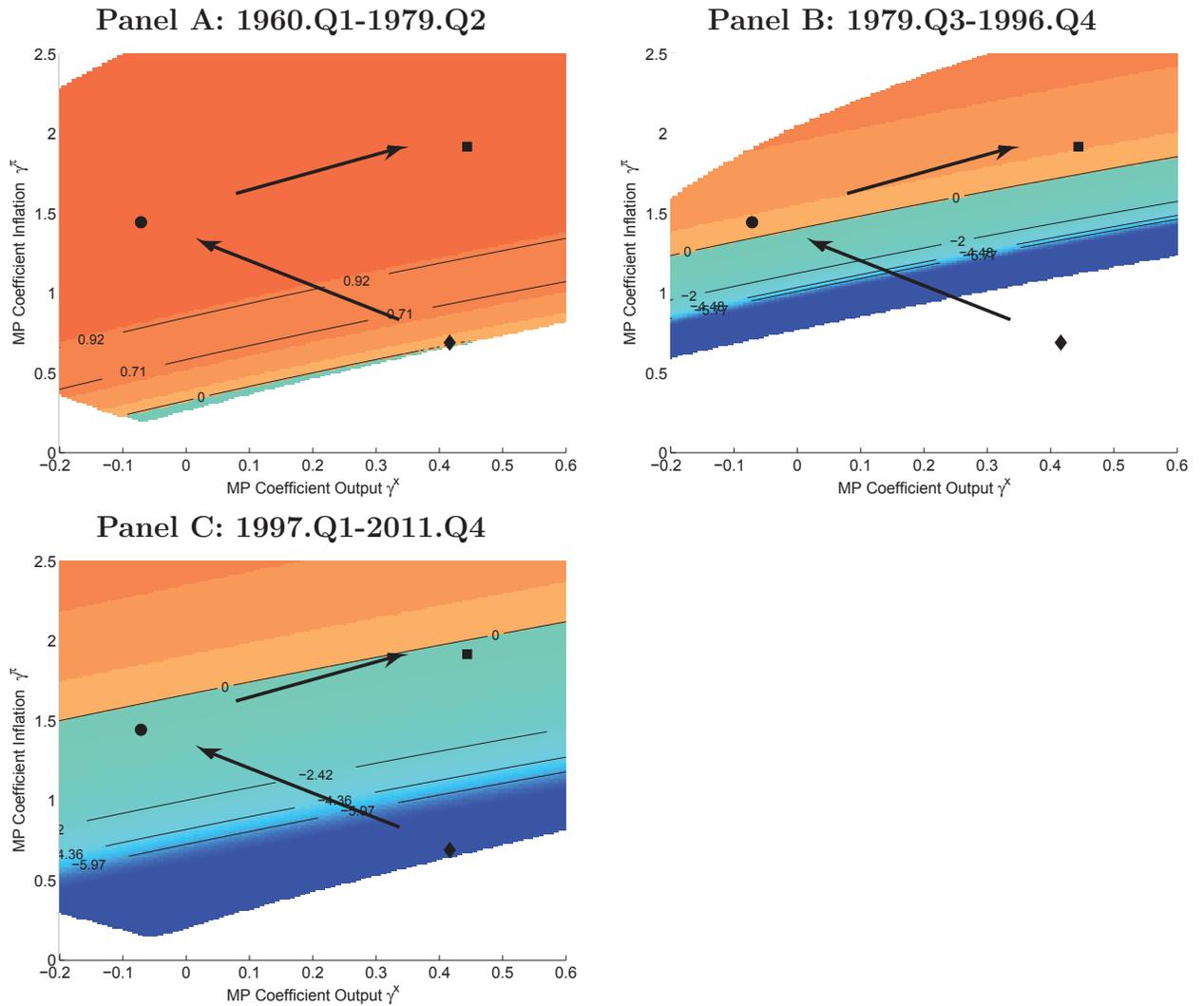
Nominal bond beta and standard deviation of nominal bond returns from daily bond and stock returns over past three months as in Campbell, Sunderam, and Viceira (2013). Gray vertical lines depict Hamilton (2009) oil price shocks.

Figure 2: US Log Output Gap and Log Price-Dividend Ratio



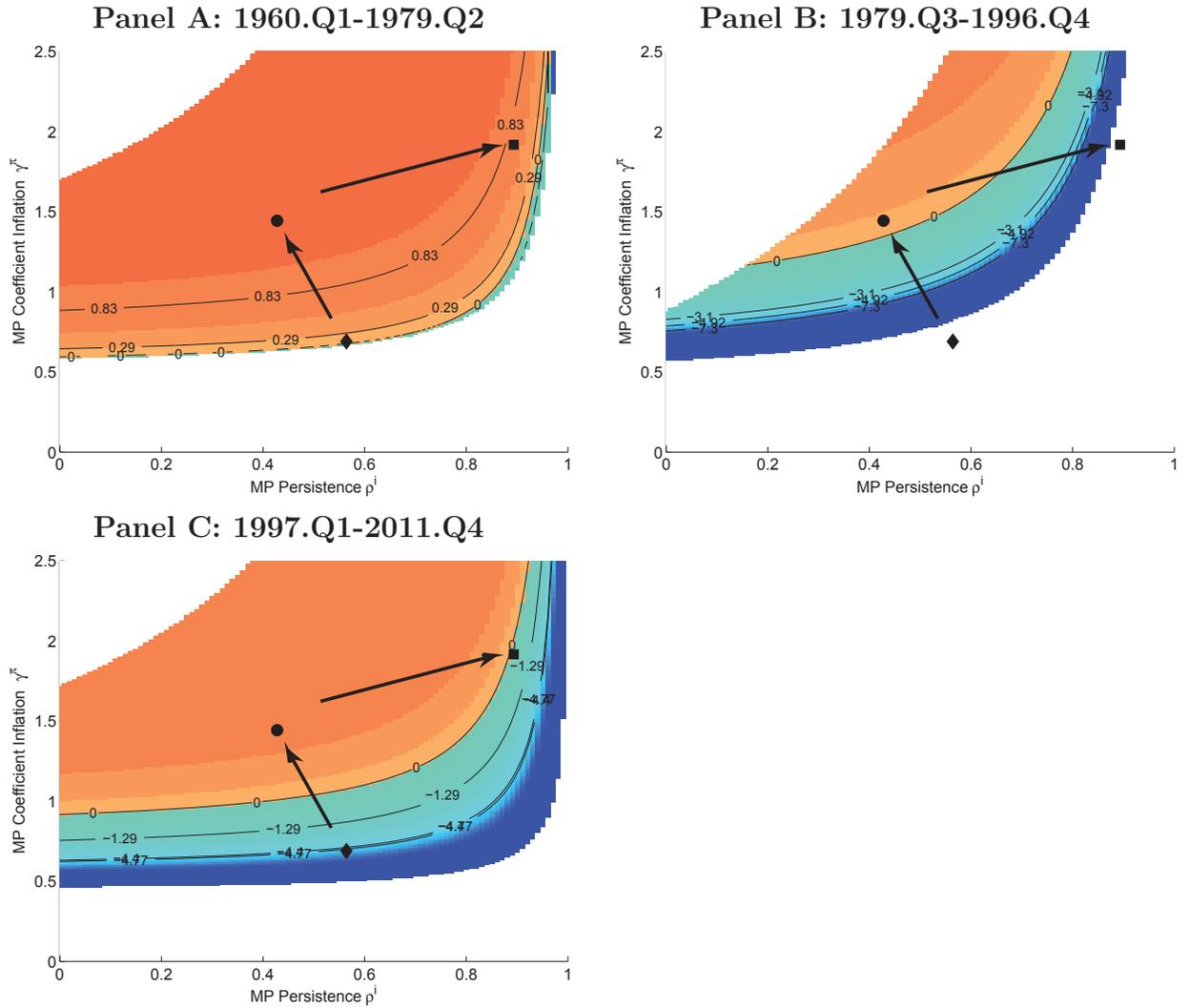
US log output gap (%) as described in Table 1. The end-of-quarter price dividend ratio is computed as the S&P 500 real price divided by real dividends averaged over the past 10 years.

Figure 3: Nominal Bond Betas Against Monetary Policy Parameters γ^π and γ^x



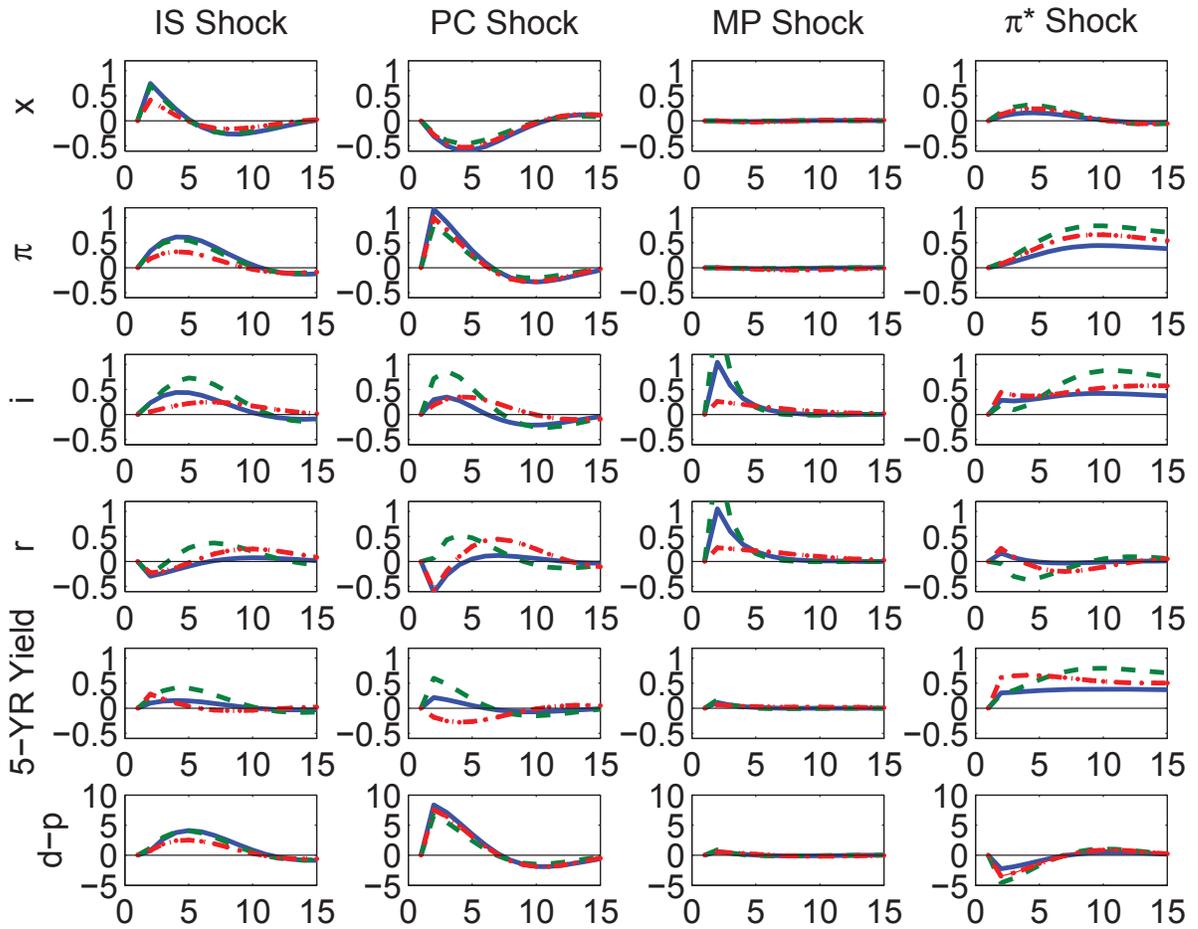
The volatilities of shocks and the monetary policy persistence parameter ρ^i are fixed at their calibrated values for 1960.Q1-1979.Q2 (Panel A), 1979.Q3-1996.Q4 (Panel B) or 1997.Q1-2011.Q4 (Panel C). We indicate calibrated monetary policy parameters as follows: 1960.Q1-1979.Q2 = diamond, 1979.Q3-1996.Q4 = circle, 1997.Q1-2011.Q4 = square. Arrows connect the sub period monetary policy rules in chronological order.

Figure 4: Nominal Bond Betas Against Monetary Policy Parameters γ^π and ρ^i



The volatilities of shocks and the monetary policy coefficient γ^x are fixed at their calibrated values for 1960.Q1-1979.Q2 (Panel A), 1979.Q3-1996.Q4 (Panel B) or 1997.Q1-2011.Q4 (Panel C). We indicate calibrated monetary policy parameters as follows: 1960.Q1-1979.Q2 = diamond, 1979.Q3-1996.Q4 = circle, 1997.Q1-2011.Q4 = square. Arrows connect the sub period monetary policy rules in chronological order.

Figure 5: Impulse Response Functions



We show impulses for the output gap, inflation, the nominal and real Federal Funds rates, the 5 year nominal yield, and the log dividend price ratio following one standard deviation shocks in period 1. We show impulse responses for the sub periods 1960.Q1-1979.Q2 (blue solid line), 1979.Q3-1996.Q4 (green dashed line), and 1997.Q1-2011.Q4 (red dash-dot line). Note that standard deviations of shocks vary across sub periods. The output gap and the dividend price ratios are in percent deviations from the steady state. All other variables are in annualized percent units.