# Monetary Policy and the Term Structure of Interest Rates*

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**Abstract**

We study how well a New Keynesian business cycle model can explain the observed behavior of nominal interest rates. We focus on two puzzles raised in previous literature. First, Donaldson, Johnsen, and Mehra (1990) show that while in the U.S. nominal term structure the interest rates are pro-cyclical and term spreads counter-cyclical the stochastic growth model predicts that the interest rates are counter-cyclical and term spreads pro-cyclical. Second, according to Backus, Gregory, and Zin (1989) the standard general equilibrium asset pricing model can account for neither the sign nor the magnitude of average risk premiums in forward prices. Hence, the standard model is unable to explain rejections of the expectations hypothesis. We show that a New Keynesian model with habit-persistent preferences and a monetary policy feedback rule produces pro-cyclical interest rates, counter-cyclical term spreads, and creates enough volatility in the risk premium to account for the rejections of expectations hypothesis. Moreover, unlike Buraschi and Jiltsov (2005), we identify the systematic monetary policy, not monetary policy shocks, as the key factor behind rejections of expectations hypothesis.

**Keywords:** Term Structure of Interest Rates, Monetary Policy, Sticky Prices, Habit Formation, Expectations Hypothesis.

**JEL classification:** E43, E44, E5, G12.

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# 1. Introduction

One of the oldest problems in economic theory is the interpretation of the term structure of interest rates. It has been long recognized that the term structure of interest rates conveys information about economic agents' expectations about future interest rates, inflation rates, and exchange rates. In fact, it is widely agreed that the term structure is *the best* source of information about economic agents' inflation expectations for one to four years ahead.[1]

Since it is generally recognized that monetary policy can only have an impact on real and nominal variables with "long and variable lags" as Friedman (1968) put it, the term structure is an invaluable source of information for monetary authorities.[2] Moreover, empirical studies indicate that the slope of the term structure predicts consumption growth better than vector autoregressions or leading commercial econometric models.[3]

While the empirical performance of the term structure as a predictor of future economic conditions has been amply documented, currently available macroeconomic models do not seem to capture neither the basic quantitative nor qualitative features of the term structure. First, Donaldson, Johnsen, and Mehra (1990) show that while in the U.S. nominal term structure the interest rates are pro-cyclical and term spreads counter-cyclical, the standard stochastic growth model predicts that interest rates are counter-cyclical and term spreads pro-cyclical.[4] King and Watson (1996) compare a real business cycle model, a sticky price model, and a liquidity effect model. They emphasize that none of the models captures the empirical fact that high real or nominal interest rates predict low level of economic activity two to four quarters in the future.

Second, if agents are risk-averse the term structure will depend on both the private sector's expectations and on the term premium. In order for the policy makers to extract information about market expectations from the term structure they need to have knowledge about the sign and the magnitude of the term premia. But as Söderlind and Svensson (1997) note in their review:

> "We have no direct measurement of this (potentially) time-varying covariance [term premium], and even ex post data is of limited use since the stochastic discount factor is not observable. It has unfortunately proved to be very hard to explain (U.S. ex post) term premia by either utility based asset pricing models or various proxies for risk."

We develop a Dynamic Stochastic General Equilibrium model with habit-persistent preferences and nominal rigidities to explain the nominal interest rate term structure. It turns out that the model produces pro-cyclical interest rates, counter-cyclical term spreads, and creates enough volatility in the risk premium to account for the rejections of expectations hypothesis. In addition, we show that the conduct of monetary policy is a key factor behind these results.

[1] See, e.g., Fama (1975, 1990) and Mishkin (1981, 1990a, 1992) for studies on inflation expectations and the term structure of interest rates using U.S. data. Mishkin (1991) and Jorion and Mishkin (1991) use international data. Abken (1993) and Blough (1994) provide surveys of the literature.

[2] Svensson (1994ab) and Söderlind and Svensson (1997) discuss monetary policy and the role of the term structure of interest rates as a source of information. Evans and Marshall (1998), Piazzesi (2005), Cochrane and Piazzesi (2002, 2005), and Buraschi and Jiltsov (2005) are recent contributions to this literature.

[3] See, e.g., Harvey (1988), Chen (1991), and Estrella and Hardouvelis (1991).

[4] Donaldson, Johnsen, and Mehra derived their theoretical results under the assumption that the capital depreciates fully. With less than full depreciation, the results differ, see Vigneron (1999).

The rest of the paper is organized as follows. The rest of this Section discusses in more detail the two problems in previous literature that we identified and goes through the related literature in more detail. Section 2 explains the New Keynesian model we use, Section 3 discusses the techiniques we use to solve the model numerically, and Section 4 explains the parameterization of the model. Section 5 reports the results related the term structure—particularly in relation to the term spread and term premium puzzles. Section 6 discusses the relationship between monetary policy and the term structure. Section 7 investigates the role of the business cycle shocks in explaining the term structure behaviour. Finally, Section 8 concludes. Appendix A derives the inflation rate dynamics in our model. Appendix B presents results from seven different experiments to illustrate features of the model.

*Cyclical Behavior of the Interest Rates*

Donaldson, Johnsen, and Mehra (1990) show that in a stochastic growth model with full depreciation the term structure of (ex-ante) real interest rates is rising at the peak of the business cycle and falling at the trough of the cycle. In addition, at the peak of the cycle the term structure lies uniformly below the term structure at the trough.

The economic intuition for the behavior of the interest rates in economic models is straightforward. At the cycle peak, the aggregate and individual consumption are expected to be, on average, lower in the future and thereby the agents will want to save more thereby driving the interest rates down. At the cycle trough the aggregate and individual consumption are expected to be higher in the future and the agents, consequently, have less need to save and push the interest rates up.

Fama (1990) summarizes the empirical evidence on the nominal interest rate term structure:

> "A stylized fact about the term structure is that interest rates are pro-cyclical. (...) [In] every business cycle of the 1952–1988 period the one-year spot rate is lower at the business trough than at the preceding or following peak. (...) Another stylized fact is that long rates rise less than short rates during business expansions and fall less during contractions. Thus spreads of long-term over short-term yields are countercyclical. (...) [In] every business cycle of the 1952–1988 period the five-year yield spread (the five-year yield less the one-year spot rate) is higher at the business trough than at the preceding or following peak."

It should be emphasized that Donaldson, Johnsen, and Mehra compare theoretical results concerning *real* interest rates with empirical data on *nominal* interest rates. They provide two arguments to justify this. First, in the empirical literature (e.g., Mishkin 1981, 1990a, 1992) the real and the nominal term structure move in tandem. Second, the results for the nominal and the real term structure in the theoretical model developed by Backus, Gregory, and Zin (1989) were qualitatively very similar.

Both of these arguments have a flaw. With few exceptions, the empirical literature had used ex-post real term structure derived from the Fisher hypothesis. In addition, Labadie (1994) shows that in a monetary endowment economy the results concerning the shape of the term structure depend crucially on whether the real GDP is assumed to be trend-stationary or difference-stationary. Vigneron (1999) shows that the same is true in a real production economy, and, moreover, the degree of depreciation in capital affects the results.

King and Watson (1996) compare a real business cycle model, a sticky price model, and a liquidity effect model. They emphasize that none of the models captures the empirical fact that high real or nominal interest rates predict low level of economic activity two to four quarters in the future. Compared to their work, we include habit-persistent preferences and a monetary feedback rule, and study the whole term structure.

*Rejections of the Expectations Hypothesis*

The empirical research on the term structure of interest rates has for the most part focused on testing the (pure) expectations hypothesis. In this strand of the literature the hypothesis examined is whether forward rates are unbiased predictors of future spot rates. The most popular way to test the hypothesis has been to run a linear regression:

$$r_{t+1} - r_t = a + b(f_t - r_t) + \varepsilon_t,$$

where $r_t$ is the one-period spot rate at time $t$ and $f_t$ is the one-period-ahead forward rate at time $t$. The *pure expectations hypothesis* implies that $a = 0$ and $b = 1$. Rejection of the first restriction, $a = 0$, gives the *expectations hypothesis*: the term premium is nonzero but constant.

By and large the literature rejects both restrictions.[5] Rejection of the second restriction, $b = 1$, requires, under the alternative, a risk premium that varies through time and is correlated with the forward premium, $f_t - r_t$. Most studies—e.g., Fama and Bliss (1987) and Fama and French (1989)—interpret this result as evidence of the existence of a time-varying risk premium. What models are capable of generating risk premia variability similar to the ones observed in the time series?

To address this question, Backus, Gregory, and Zin (1989) a standard dynamic general equilibrium asset pricing model as developed by LeRoy (1973), Rubinstein (1976), Lucas (1978), and Breeden (1979). The important result in Backus, Gregory, and Zinis that the model can account for neither the sign nor the magnitude of average risk premiums in forward prices and holding-period returns. The model is unable to explain rejections of the expectations hypothesis. A similar puzzle has been shown to exist for equity premia by Mehra and Prescott (1985).

*Related Literature*

In previous literature, Evans and Marshall (1998) show that a limited participation model is broadly consistent with the impulse response functions of the real and nominal yields to a monetary policy shock. However, their real yields are *ex-post* real yields. In addition, Piazzesi (2005) criticizes their methodology on the grounds that it doesn't impose the no arbitrage condition on the yield movements. Seppälä and Xie (2004) study the cyclical behavior of nominal and (ex-ante) real term structures of interest rates in the UK data, and in a real business cycle, a limited participation, and a New Keynesian model. Their result is that only the New Keynesian model gets closest to matching the cyclical behavior for both the nominal and the real term structure. Unlike in this paper, they use exogenously defined money supply processes, assume that the Fisher hypothesis holds, and employ standard households preferences.

[5] The literature is huge. Useful surveys are provided by Melino (1988), Shiller (1990), Mishkin (1990b), and Campbell, Lo, and MacKinlay (1997). The most important individual studies are probably Shiller (1979), Shiller, Campbell, and Schoenholtz (1983), Fama (1984, 1990), Fama and Bliss (1987), Froot (1989), Campbell and Shiller (1991), and Campbell (1995).

Piazzesi (2005) shows that the Federal Reserve policy can be better approximated by assuming that it responds to the information contained *only* in the term structure rather than in other macroeconomic variables. Cochrane and Piazzesi (2005) show that the monetary policy shocks can explain 45% of excess nominal bond returns, and Cochrane and Piazzesi (2002) show that the term structure explains 64% of the changes in the federal funds target rate.

Buraschi and Jiltsov (2005) study the inflation risk premium in a continuous-time general equilibrium model in which the monetary authority sets the money supply based on targets on the long-term growth of the nominal money supply, inflation, and economic growth. They identify the time-variation of the inflation risk premium as an important explanatory variable of deviations from the expectations hypothesis. In contrast, in our model the monetary policy authority follows an interest rate rule which is closer to the actual conduct of monetary policy in most countries. Moreover, we find that the in our model monetary policy shocks and inflation risk premium are not the explation behind the rejections of the expectations hypothesis. Instead the *systematic* monetary policy drives our results. This result is closely related to earlier studies by Mankiw and Summers (1984), Mankiw and Miron (1986), and Dotsey and Otrok (1995) in reduced form models.

Seppälä (2004) studies the asset pricing implications of an endowment economy when agents can default on contracts. The results show that this limited commitment model is one potential solution of the term premium puzzle. Dai (2002) shows that a model with limited participation is another. However, both of these models study only the real term structure.[6] One of our objectives is to study whether a habit-formation model that in previous work has been successful in accounting for asset pricing puzzles can also explain the term premium puzzle.[7] Buraschi and Jiltsov (2003) and Wachter (2004) show that an external habit model in the style of Campbell and Cochrane (1999) can explain this puzzle.

In reduced form macro models, Bekaert, Cho, and Moreno (2003) estimate a log-linear three equation New Keynesian model using a Maximum Likelihood estimator, and examine the term structure generated by the model. The dynamics of the endogenous variables is driven by three exogenous shocks and two unobserved state variables. While their specifications is justified on empirical grounds—and helps to generate persistent dynamics—it allows many degrees of freedom, so that the role played in the term structure behaviour by the endogenous shock transmission mechanism of the optimizing model is difficult to ascertain.

The research on joint macro-finance model poses similar problems (Hördahl, Tristani, and Vestin, 2002, Rudebusch and Wu, 2004). This literature aims at integrating small scale optimizing models of output, inflation and interest rates with affine no-arbitrage specifications for bond prices. In this way, it is possible to identify the affine model latent factor with the macroeconomic aggregates. The bond pricing portion of the model does not have a structural interpretation. Therefore, it is difficult to evaluate the importance of the exogenous persistence introduced in this family of models for the term structure results.

[6] In addition, there are a few recent papers such as Duffee (2002) and Dai and Singleton (2002) that study the term premium puzzle in the nominal yields using reduced form no arbitrage models. We use a structural general equilibrium model (which naturally implies no arbitrage).

[7] Previous studies include Jermann (1998) and Boldrin, Christiano, and Fisher (2001).

# 2. The Model

The theoretical interest rate term structure is derived from a dynamic stochastic general equilibrium model of the business cycle. An important objective of the paper is to evaluate the role of monetary policy in generating an empirically plausible term structure. Hence, we adopt a money-in-utility-function model where nominal rigidities allow monetary policy to affect the dynamics of real variables. We follow Calvo (1983) and the New Keynesian literature on the business cycle by assuming that prices cannot be updated to the profit-maximizing level in each period. Firms face an exogenous, constant probability of being able to reset the price in any period $t$. This setup can also be derived from a menu cost model, where firms face a randomly distributed fixed cost $k_t$ of updating the price charged, and the support of $k_t$ is $[0;\overline{k}]$, $\overline{k} \to \infty$ (see Klenow and Kryvtsov, 2004).

While more sophisticated pricing mechanism can be introduced—such as state-dependent pricing (Dotsey, King and Wolman, 1999), partial indexation to past prices (Christiano, Eichenbaum and Evans, 2001), a mix of rule-of-thumb and forward-looking pricing (Gali and Gertler, 1999)—we limit the model to the more essential ingredients of the New Keynesian framework. This allows us to investigate the impact on the term structure of four key features: (i) systematic monetary policy modeled as an interest rate rule; (ii) nominal price rigidity; (iii) habit-persistent preferences; (iv) positive steady state money growth rate. Woodford (2003) and Walsh (2003) offer a comprehensive treatment of the New Keynesian framework, and describe in detail the microfoundations of the model.

Each consumer owns shares of all firms, and households are rebated any profit from the monopolistically competitive output sector. Savings can be accumulated in money balances, or in a range of riskless nominal and real bonds spanning several maturities. The government runs a balanced budget in every period, and rebates to consumers any seigniorage revenue from issuing the monetary asset. Output is produced with undifferentiated labor, supplied by the household-consumers, via a linear production function.

*Households*

There is a continuum of infinitely lived households, indexed by $j \in [0, 1]$. Consumers demand differentiated consumption goods, choosing from a continuum of goods, indexed by $z \in [0, 1]$. In the notation used throughout the paper, $C_t^j(z)$ indicates consumption by household $j$ at time $t$ of the good produced by firm $z$.

Households' preferences over the basket of differentiated goods are defined by the CES aggregator:

$$C_t^j = \left[ \int_0^1 C_t^j(z)^{\frac{\theta-1}{\theta}} dz \right]^{\frac{\theta}{\theta-1}}, \ \theta > 1 \tag{1}$$

The representative household chooses $\left\{ C_{t+i}^j, C_{t+i}^j(z), N_{t+i}^j, \frac{X_{t+i}^j}{P_{t+i}}, \frac{B_{t+i}^j}{P_{t+i}} \right\}_{i=0}^{\infty}$ where $N_t$ denotes labor supply, $X_t$ nominal money balances, $P_t$ the aggregate price level, and $B_t$ bond holdings, to maximize:

$$E_t \sum_{i=0}^{\infty} \beta^i \left\{ \frac{(C_{t+i}^j - bC_{t+i-1}^j)}{1-\gamma}^{1-\gamma} D_{t+i} - \frac{\ell N_{t+i}^{1+\eta}}{1+\eta} + \frac{\xi}{1-\gamma_m} \left( \frac{X_{t+i}^j}{P_{t+i}} \right)^{1-\gamma_m} \right\} \tag{2}$$

subject to

$$\int_0^1 C_t^j(z) P_t(z) dz = W_t N_t^j + \Pi_t^j - (X_t^j - X_{t-1}^j) - (\overrightarrow{p}_t \overrightarrow{B}_t^j - B_{t-1}^j) - \tau_t^j, \tag{3}$$

and (1). When $b > 0$ the preferences are characterized by habit persistence (Boldrin, Christiano, and Fisher, 2001). $D_t$ is an aggregate stochastic preference shock. Each element of the row vector $\overrightarrow{p}_t$ represents the price of an asset with maturity $k$ that will pay one unit of currency in period $t+k$. The corresponding element of $\overrightarrow{B}_t$ represents the quantity of such claims purchased by the household. $B_{t-1}^j$ indicates the value of the household portfolio of claims maturing at time $t$.

The solution to the intratemporal expenditure allocation problem between the varieties of differentiated goods gives the individual good $z$ demand function:

$$C_t^j(z) = \left[ \frac{P_t(z)}{P_t} \right]^{-\theta} C_t^j. \tag{4}$$

Equation (4) is the demand of good $z$ from household $j$, where $\theta$ is the price elasticity of demand. The associated price index $P_t$ measures the least expenditure for differentiated goods that buys a unit of the consumption index:

$$P_t = \left[ \int_1^0 P_t(z)^{1-\theta} dz \right]^{\frac{1}{1-\theta}}. \tag{5}$$

Since all household solve an identical optimization problem and face the same aggregate variables, in the following we omit the index $j$. Using equations (4) and (5), we can write the budget constraint as:

$$C_t = \frac{W_t}{P_t} N_t + \frac{\Pi_t}{P_t} - \frac{X_t - X_{t-1}}{P_t} - \frac{\overrightarrow{p}_t \overrightarrow{B}_t - B_{t-1}}{P_t} - \frac{\tau_t}{P_t},$$

The first order conditions with respect to labor and real money balances are:

$$MUC_t = E_t \left[ \frac{D_t}{(C_t - bC_{t-1})^\gamma} - \beta b \frac{D_{t+1}}{(C_{t+1} - bC_t)^\gamma} \right]$$

$$0 = MUC_t \frac{W_t}{P_t} - \ell N_t^\eta \tag{6}$$

$$0 = \xi \left( \frac{X_t}{P_t} \right)^{-\gamma_m} - MUC_t + E_t \left[ \beta MUC_{t+1} \frac{P_t}{P_{t+1}} \right] \tag{7}$$

where $MUC$ is the marginal utility of consumption. The Euler equation for the time $t$ price $p_{t,k}^b$ of a bond paying a unit of consumption aggregate at time $t+k$ is:

$$p_{t,k}^b = E_t \left[ \beta^k \frac{MUC_{t+k}}{MUC_t} \right]$$

We define the real gross interest rate for maturity $k$ as $(1 + r_{t,k}) = 1/p^b_{t,k}$. In a similar fashion, the Euler equation associated withe the nominal bond yielding one unit of money with certainty in period $t + k$ is:

$$p^B_{t,k} = E_t \left[ \beta^k \frac{P_t}{P_{t+k}} \frac{MUC_{t+k}}{MUC_t} \right] = \frac{1}{(1 + R_{t,k})} \tag{8}$$

where $(1 + R_{t,k})$ is the gross nominal interest rate for maturity $k$.

*Firms and price setting*

The firm producing good $z$ employs a linear technology:

$$Y_t(z) = A_t N_t(z)$$

where $A_t$ is an aggregate productivity shock. Minimizing the nominal cost of producing a given amount of output $\overline{Y}$:

$$Cost = W_t N_t(z)$$

yields the labor demand schedule:

$$MC^N_t(z) MPL_t(z) = W_t \tag{9}$$

where $MC^N$ is the nominal marginal cost, $MPL$ is the marginal product of labor $(Y_t(z)/N_t(z))$. Equation (9) implies that the real marginal cost $MC_t$ of producing one unit of output is:

$$MC_t(z) MPL_t(z) = W_t/P_t$$

Firms adjust their prices infrequently. In each period there is a constant probability $(1 - \theta_p)$ that the firm will be able to adjust its price, independently of past history. This implies that the fraction of firms setting prices at $t$ is $(1 - \theta_p)$ and the expected waiting time for the next price adjustment is $\frac{1}{1-\theta_p}$. The problem of the firm setting the price at time $t$ consists of choosing $P_t(z)$ to maximize the expected discounted stream of profits:

$$E_t \sum_{i=0}^{\infty} (\theta_p \beta)^i \frac{MUC_{t+i}}{MUC_t} \left[ \frac{P_t(z)}{P_{t+i}} Y_{t,t+i}(z) - \frac{MC^N_{t+i}}{P_{t+i}} Y_{t,t+i}(z) \right] \tag{10}$$

subject to

$$Y_{t,t+i}(z) = \left[ \frac{P_t(z)}{P_{t+i}} \right]^{-\theta} Y_{t+i}, \tag{11}$$

In (11), $Y_{t,t+i}(z)$ is the firm's demand function for its output at time $t+i$, conditional on the price set at time $t$, $P_t(z)$. Market clearing insures that $Y_{t,t+i}(z) = C_{t,t+i}(z)$ and $Y_{t+i} = C_{t+i}$. Substituting (11) into (10), the objective function can be written as:

$$E_t \sum_{i=0}^{\infty} (\theta_p \beta)^i \frac{MUC_{t+i}}{MUC_t} \left\{ \left[ \frac{P_t(z)}{P_{t+i}} \right]^{1-\theta} Y_{t+i} - \frac{MC^N_{t+i}}{P_{t+i}} \left[ \frac{P_t(z)}{P_{t+i}} \right]^{-\theta} Y_{t+i} \right\}. \tag{12}$$

Since $P_t(z)$ does not depend on $i$, the optimality condition is:

$$P_t(z)E_t\sum_{i=0}^{\infty}(\theta_p\beta)^i MUC_{t+i}\left[\frac{P_t(z)}{P_{t+i}}\right]^{1-\theta}Y_{t+i}=\mu E_t\sum_{i=0}^{\infty}(\theta_p\beta)^i MUC_{t+i}MC^N_{t+i}\left[\frac{P_t(z)}{P_{t+i}}\right]^{1-\theta}Y_{t+i}. \quad (13)$$

where

$$\mu=\frac{\theta}{\theta-1}$$

is the flexible-price level of the markup, and also the markup that would be observed in a zero-inflation (zero money growth rate) steady state. To use rational expectations solution algorithms when the steady state money growth rate is non-zero, we need to express the first order condition as a difference equation (see Ascari, 2004, and King and Wolman, 1996). This can be accomplished expressing $P_t(z)$ as the ratio of two variables:

$$P_t(z)=\frac{G_t}{H_t},$$

and

$$G_t=\frac{(G_t/H_t)^{1-\theta}}{MUC_t}\hat{G}_t \quad (14)$$

$$H_t=\frac{(G_t/H_t)^{1-\theta}}{MUC_t}\hat{H}_t, \quad (15)$$

where

$$\hat{G}_t=\mu MUC_t MC_t P_t^{\theta-1}Y_t+\theta_p\beta\hat{G}_{t+1} \quad (16)$$

$$\hat{H}_t=MUC_t P_t^{\theta-1}Y_t+\theta_p\beta\hat{H}_{t+1}. \quad (17)$$

*Market Clearing*

Since the measure of the economy is unitary, in the symmetric equilibrium it holds that:

$$M_t^j=M_t\ ;\ C_t^j=C_t$$

and the consumption shadow price is symmetric across households: $MUC_t^j=MUC_t$. Given that all firms are able to purchase the same labor service bundle, and so are charged the same aggregate wage, they all face the same marginal cost. The linear production technology insures that $MC$ is equal across all firms—whether they are updating or not their price—regardless of the level of production, which will indeed be different. Firms are heterogeneous in that a fraction $(1-\theta_p)$ of firms in the interval $[0,1]$ can optimally choose the price charged at time $t$. In equilibrium each producer that chooses a new price $P_t(z)$ in period $t$ will choose the same new price $P_t(z)$ and the same level of output. Then the dynamics of the consumption-based price index will obey

$$P_t=\left[\theta_p P_{t-1}^{1-\theta}+(1-\theta_p)P_t(z)^{1-\theta}\right]^{\frac{1}{1-\theta}}. \quad (18)$$

The Appendix A shows that the inflation rate dynamics is given by:

$$[(1+\pi_t)]^{1-\theta} = \theta_p + (1-\theta_p)\left[\frac{\tilde{G}_t}{\tilde{H}_t}(1+\pi_t)\right]^{1-\theta} \tag{19}$$

$$\tilde{G}_t \equiv \frac{\hat{G}_t}{P_t^{\theta}} \quad ; \quad \tilde{H}_t \equiv \frac{\hat{H}_t}{P_t^{\theta-1}}$$

In a steady state with gross money growth rate equal to $\Upsilon$, and gross inflation equal to $\Pi = \Upsilon$,

$$\frac{G}{HP_t} = \frac{\tilde{G}}{\tilde{H}} = \frac{P_t(z)}{P_t}$$

$$\frac{P_t(z)}{P_t} = \mu * MC * \frac{(1-\theta_p\beta\Pi^{\theta-1})}{(1-\theta_p\beta\Pi^{\theta})}$$

Since $P_t(z)$ is the optimal price chosen by the fraction of firms that can re-optimize at time $t$, it is the inverse of what King and Wolman (1996) define as the *price wedge.* With zero steady state inflation the steady state average markup is equal to $1/MC$, therefore there is no price wedge. But when steady state inflation is positive, the price wedge is less than one: the average price is always smaller than the optimal price, since some firms would like to increase the price, but are constrained not to do so. Combining this equation with equation (19) gives the steady state marginal cost and price wedge as a function of $\Pi$:

$$\frac{\tilde{G}}{\tilde{H}} = \left[\frac{(1-\theta_p)}{(1-\theta_p\Pi^{\theta-1})}\right]^{\frac{1}{\theta-1}}$$

$$MC = \frac{1}{\mu}\left[\frac{\Pi^{1-\theta}-\theta_p}{1-\theta_p}\right]^{\frac{1}{1-\theta}} \frac{1}{\Pi}\frac{(1-\theta_p\beta\Pi^{\theta})}{(1-\theta_p\beta\Pi^{\theta-1})}$$

*Asset markets*

The government rebates the seigniorage revenues to the household in the form of lump-sum transfers, so that in any time $t$ the government budget is balanced. Since we defined in equation (3) $\tau^j$ as the amount of the tax levied by the government on household $j$, assuming $\tau_t^j = \tau_t^i \;\; \forall\, j, i \;\in [0,1]$ , at every date $t$ the transfer will be equal to:

$$-\int_0^1 \tau_t^j dj = -\tau_t \int_0^1 dj =$$

$$= -\tau_t = M_t^s - M_{t-1}^s$$

Equilibrium in the money market requires:

$$M_t^s = M_t^{dj} = M_t^d$$

We assume the monetary policy instrument is the short term nominal interest rate $(1+R_{t,1})$. The money supply is set by the monetary authority to satisfy whatever money demand is consistent with the target rate.

Domestic bonds are in zero-net supply, since the government does not issue bonds. Therefore in equilibrium it must hold that:

$$B_{t,i} = 0$$

for any component of the vector $\overrightarrow{B}_t$. However, because we have complete markets, we can still price both nominal and real bonds.

*Monetary Policy*

The economy's dynamics is driven by business cycle shocks temporarily away from the non-stochastic steady state. In this instances, the domestic monetary authority follows a forward-looking, instrument feedback rule:

$$\frac{\left(1+\overline{R}_{t,t+1}\right)}{(1+R^{ss})} = E_t \left(\frac{1+\pi_{t+1}}{1+\pi_{SS}}\right)^{\omega_\pi} \left(\frac{Y_t}{Y_{SS}}\right)^{\omega_y} \tag{20}$$

where $\omega_\pi$, $\omega_y \geq 0$ are the feedback coefficients to CPI inflation and output. The monetary authority adjusts the interest rate in response to deviations of the target variables from the steady state. In the steady state, a constant money growth rate rule is followed. The choice of the parameters $\omega_\pi$, $\omega_y$ allows us to specify alternative monetary policies. When the central bank responds to current rather than expected inflation equation (20) returns the rule suggested by Taylor (1993) as a description of U.S. monetary policy.

We assume the central bank assigns positive weight to an interest rate smoothing objective, so that the domestic short-term interest rate at time $t$ is set according to

$$(1+R_{t,1}) = \left[\left(1+\overline{R}_{t,t+1}\right)\right]^{(1-\chi)} \left[(1+R_{t-1,1})\right]^{\chi} \varepsilon_t^{mp} \tag{21}$$

where $\chi \in [0,1)$ is the degree of smoothing and $\varepsilon_t^{mp}$ is an unanticipated exogenous shock to monetary policy.

## 3. *Algorithm*

We solve the model using a second-order approximation around the non-stochastic steady state. The numerical solution is done using the approach and the routines in Schmitt-Grohe and Uribe (2004). It is well known that taking a first-order approximation to the bond prices will give no risk premia and that a second-order approximation will give only constant premia. The reason is simple: the second-order approximation involves only squared error terms that have constant expectation.

For this reason, in the first step, we solve our model for six state variables and seven control variables in 13 equations using the Schmitt-Grohe and Uribe methods. In the second step, we generate 200,000 observations of state and control variables. In the final step, we regress the future marginal rates of substitution—see equations (24) and (25) below—on the third-order complete polynomials of the state variables to generate bond prices.

Our approach is very similar to the parameterized expectations algorithm employed by Evans and Marshall (1998). While they didn't study the premia, notice that the algorithm amounts to taking a third-order approximation to bond prices. With third-order approximation, the current state variables multiply squared future error terms, and hence risk premia are time-varying.

## 4. *Parametrization*

Preference, technology and policy parameters are parameterized consistently with the New Keynesian monetary business cycle literature. Estimated and calibrated staggered-price adjustment models are discussed in Bernanke and Gertler (2000), Christiano Eichenbaum, and Evans(2001), Ireland (2001), Ravenna (2002), Rabanal and Rubio-Ramirez (2003), Walsh (2003), Woodford (2003).

Households' preferences are modeled following the internal habit-persistence framework of Boldrin, Christiano, and Fisher (2001). The persistence parameter $b$ is set to 0.6, a value that Fuhrer (2000) finds optimizes the match between sticky-price models and consumption data. The value of $\gamma$ is set to 2.1, and is chosen to provide adequate curvature in the utility function so as to generate enough risk-premia volatility. The parametrization of habit-persistent preferences plays a very important role in the model's term-structure properties. Its impact on the results is discussed in detail in Section 5. Labor supply elasticity $(1/\eta)$ is equal to 2, and the parameter $\ell$ is chosen to set steady state labor hours at about 20% of the available time. This is a value consistent with many OECD countries postwar data, although on the low side for the US. The quarterly discount factor $\beta$ is parametrized so that the steady state real interest rate is equal to 1%. The demand elasticity $\theta$ is set to obtain a flexible-price equilibrium producers' markup $\mu = \vartheta/(\vartheta - 1) = 1.1$. While Bernanke and Gertler (2000) use a higher value of 1.2, in our model positive steady state inflation implies the steady state markup is larger than in the flexible-price equilibrium.

The production technology is linear in labor hours. Given that the model is parameterized at business-cycle frequencies, this a fair approximation widely used in the literature. To parametrize the Calvo (1983) pricing adjustment mechanism, the probability $\theta_p$ faced by firms of not adjusting the price in any given period is set to 0.75 , implying that the average time between price adjustments for a producer is 1 year. This value is in line with estimates for the US reported by Gali and Gertler (1999) and Rabanal and Rubio-Ramirez (2003).

Variants of the instrument rule (21) have been estimated both in single-equation and in simultaneous-equation contexts. We set the inflation feedback coefficients $\omega_\pi$ to 2.15, which is a value close to the one found by Clarida, Gali and Gertler (2000) for the Greenspan tenure in the US. The choice of a value for $\omega_y$ is more controversial, depending on the operational definition of output gap used by the central bank at any given point in time. We choose a value of $\omega_y = 0$. The smoothing parameter $\chi$ is equal to 0.8, as estimated by Rabanal (2004). Section 6 and Appendix B discuss different monetary policy specifications. Quarterly steady state inflation is set equal to the average U.S. value over the period $1994 - 2004$, about 0.75% on a quarter basis. This implies an annualized steady state nominal interest rate of 7%.

The preference and technology exogenous shocks follow an AR(1) process:

$$\log Z_t = (1 - \rho_Z) \log \overline{Z} + \rho_Z \log Z_{t-1} + \varepsilon_t^Z \qquad \varepsilon_t^Z \sim i.i.d.\ N(0, \sigma_Z^2)$$

Table 1: Selected variables volatilities and correlations. Sample: 1984–2004.

| | *Standard Deviation* | | *Correlation with output* | |
|---|---|---|---|---|
| Variable | Model | U.S. data | Model | U.S. Data |
| $Y_t$ | 1.8 | 0.91 | 1 | 1 |
| $\pi_t$ | 0.8 | 0.42 | 0.69 | 0.3 |
| $R_t$ | 1.27 | 0.53 | 0.47 | 0.42 |
| $r_t$ | 0.78 | 0.56 | 0.1 | 0.12 |

where $\overline{Z}$ is the steady state value of the variable. The policy shock $\varepsilon_t^{mp}$ is a Gaussian i.i.d. stochastic process. The autocorrelation parameters are equal to $\rho_a = \rho_d = 0.9$. The standard deviation of the innovations $\varepsilon$ is set to $\sigma_a = 0.35$, $\sigma_d = 8$, $\sigma_{mp} = 0.1$ (percent values). The low value of the policy shock implies the largest part of the short term nominal interest rate dynamics is driven by the systematic monetary policy reaction to the state of the economy. The preference shock volatility is large, but very close to the one estimated by Rabanal and Rubio-Ramirez (2003) on U.S. data. Compared to their estimates, the technology shock volatility is low. But as the volatility of this shock increases, the correlation between nominal interest rate and GDP becomes smaller and smaller, since a technology shock generates a negative correlation. Note that the authors cited above adopt a model which includes a cost-push shock. This shock generates a strong positive correlation between $R_t$ and $y_t$, and they estimate its volatility to be equal to 41.

An important concern in the parametrization of the shocks has been to match the correlations between output and nominal and real rates with U.S. data, to be able to evaluate whether the term structure generated by the model can predict output variation, as many empirical studies have found in the US. Table 1 summarizes the main statistics for the model, and compares them to U.S. data.[8] The match with empirical correlations is satisfactory. To obtain this result though the model simulated volatilities for output, nominal interest rate and inflation turn out to be substantially larger than in the data.

In Table 1 the data sample spans the last two decades, in the hope of summarizing the business cycle properties of the U.S. economy under a homogenous monetary policy regime. Table 2 compares the model's second moments to the whole U.S. post-war data sample.[9] This sample is heterogenous with respect to the U.S. monetary policy goals and the US Federal Reserve operating procedures, and includes the 1970s inflationary episode. On the other hand, the sample can be considered more

[8]Note: Standard deviation measured in percent. The output series is logged and Hodrick-Prescott filtered. U.S. data: $Y_t$ is real GDP, $\pi_t$ is CPI inflation, $R_t$ is 3-months T-bill rate, $r_t$ is ex-post short term real interest rate. All rates are on a quarter basis. Quarterly data sample is 2:1984–1:2004. Data are taken from the St. Louis Federal Reserve Bank FRED II database.

[9]Standard deviation measured in percent. The output series is logged and Hodrick-Prescott filtered. U.S. data: $Y_t$ is real GDP, $\pi_t$ is CPI inflation, $R_t$ is 3-months T-bill rate, $r_t$ is ex-post short term real interest rate. All rates are on a quarter basis. Quarterly data sample is 1:1947–1:2004. Data are taken from the St. Louis Federal Reserve Bank FRED II database.

Table 2: Selected variables volatilities and correlations. Sample: 1947–2004.

| | *Standard Deviation* | | *Correlation with output* | |
|---|---|---|---|---|
| Variable | Model | U.S. data | *Model* | *U.S. Data* |
| $Y_t$ | 1.8 | 1.69 | 1 | 1 |
| $\pi_t$ | 0.8 | 0.87 | 0.69 | 0.27 |
| $R_t$ | 1.27 | 0.74 | 0.47 | 0.17 |
| $r_t$ | 0.78 | 0.8 | 0.1 | −0.16 |

representative of the variety of shocks that drove the U.S. business cycle. As expected, the standard deviations of all U.S. variables increases. The correlations of output with inflation and the ex-post real interest rate drop significantly, and become much smaller compared to the 1984-2004 sample values and to the model theoretical prediction.

# 5. *The Term Structure of Interest Rates*

*The Real and Nominal Term Structures*

Let $m_{t+1}$ denote the real stochastic discount factor

$$m_{t+1} \equiv \beta \frac{MUC_{t+1}}{MUC_t}, \tag{22}$$

and let $M_{t+1}$ denote the nominal stochastic discount factor

$$M_{t+1} \equiv \beta \frac{MUC_{t+1}}{MUC_t} \frac{P_t}{P_{t+1}}, \tag{23}$$

The price of an $n$-period zero-coupon real bond is given by

$$\begin{aligned} p^b_{t,n} &= E_t \left[ \prod_{j=1}^{n} m_{t+j} \right] \\ &= E_t[m_{t+1} p^b_{t+1,n-1}], \end{aligned} \tag{24}$$

and similarly the price of an $n$-period zero-coupon nominal bond is given by

$$\begin{aligned} p^B_{t,n} &= E_t \left[ \prod_{j=1}^{n} M_{t+j} \right] \\ &= E_t[M_{t+1} p^B_{t+1,n-1}]. \end{aligned} \tag{25}$$

The bond prices are invariant with respect to time, and hence equations (24) and (25) give a recursive formula for pricing zero-coupon real and nominal bonds of any maturity. For simplicity, we next express rates for only real rates. Nominal rates are obtained in a similar manner.

Forward prices are defined by

$$p^f_{t,n} = \frac{p^b_{t,n+1}}{p^b_{t,n}},$$

and the above prices are related to interest rates (or yields) by

$$f_{t,n} = -\log(p^f_{t,n}) \quad \text{and} \quad r_{t,n} = -(1/n)\log(p^b_{t,n}). \tag{26}$$

To define the risk premium as in Sargent (1987), write (24) for a two-period bond using the conditional expectation operator and its properties:

$$\begin{aligned} p^b_{t,2} &= E_t[m_{t+1}p^b_{t+1,1}] \\ &= E_t[m_{t+1}]E_t[p^b_{t+1,1}] + \text{cov}_t[m_{t+1}, p^b_{t+1,1}] \\ &= p^b_{t,1}E_t[p^b_{t+1,1}] + \text{cov}_t[m_{t+1}, p^b_{t+1,1}], \end{aligned}$$

which implies that

$$p^f_{t,1} = \frac{p^b_{t,2}}{p^b_{t,1}} = E_t[p^b_{t+1,1}] + \text{cov}_t\left[m_{t+1}, \frac{p^b_{t+1,1}}{p^b_{t,1}}\right]. \tag{27}$$

Since the conditional covariance term is zero for risk-neutral investors, we call it the *risk premium* for the one-period forward contract, $rp_{t,1}$, given by

$$rp_{t,1} \equiv \text{cov}_t\left[m_{t+1}, \frac{p^b_{t+1,1}}{p^b_{t,1}}\right] = p^f_{t,1} - E_t[p^b_{t+1,1}],$$

and similarly $rp_{t,n}$ is the risk premium for the $n$-period forward contract:

$$rp_{t,n} \equiv \text{cov}_t\left[\prod_{j=1}^{n} m_{t+j}, \frac{p^b_{t+n,1}}{p^b_{t,1}}\right] = p^f_{t,n} - E_t[p^b_{t+n,1}].$$

Table 3 presents the means, standard deviations, and correlations with for the nominal and real term structure in the model, and for the U.S. nominal data as estimated by McCulloch and Kwon (1993) from the first quarter of 1947 until the fourth quarter of 1990 and by Duffee (2001) from the first quarter of 1991 until the fourth quarter of 1998. Output is filtered using the Hodrick-Prescott (1980) filter with a smoothing parameter of 1600 both in the model data in data.

The Table shows that both real and nominal term structures are procyclical. In data, short maturities are procyclical and long maturities countercyclical. In contrast, the nominal term spreads are clearly countercyclical both in data and in model. Both the nominal and real term structures generated by the model are mostly upward-sloping, but the long end of nominal term structure slightly slopes downward. In data, the term structure is clearly upward-sloping. Means are matched quite well; in model the nominal yields from three months until 20 years vary from 5% to 6.73% and in data from 5% to 6.5%. Similarly, the average term spreads are produced by the model are quite close to the average term spreads in data. The term structure of volatilities in model is strongly downward-sloping while in data it is essentially flat.[10] However, the average standard deviations across maturities are roughly the same in data and in model.

In addition, the model produces strong positive correlation between yields and (the cyclical component of) output while in data the correlation is low and positive for short maturities and essentially zero for long maturities. The strong positive correlation in the model is a product of the large shocks autocorrelation. Persistent shocks are needed to obtain sufficient volatility of rates at the long end. The downside is that correlations with output will be very high as shocks die out slowly. Possible remedies are introducing hybrid inflation and/or time-varying inflation target. Both would give a larger volatility of interest rate at long maturities, with smaller shocks variance, probably lowering the correlation.

*The Expectations Hypothesis*

The oldest and simplest theory about the information content of the term structure is so called (pure) expectations hypothesis. According to the pure expectations theory forward rates are unbiased predictors of future spot rates. It is also common to modify the theory so that constant risk-premium is allowed—this is usually called the expectations hypothesis. However, it should be noted that both versions of the expectations hypothesis are always incorrect. To see this, let us assume, for a sake of an argument, that the agents are risk-neutral: $\gamma = 0$. Equation (27) reduces then into

$$p^f_{t,1} = E_t[p^b_{t+1,1}]$$

and from (26) we obtain

$$\exp^{-f_{t,1}} = E_t\left[\exp^{-r_{t+1,1}}\right].$$

From the Jensen's inequality it follows that

$$f_{1,t} < E_t[r_{1,t+1}] \tag{28}$$

and the difference between the left and right hand side of (28) varies with $E_t[r_{1,t+1}]$ and $\mathrm{var}_t[r_{1,t+1}]$. This effect is known as *convexity premium* or *bias*.

Backus, Gregory, and Zin (1989), on the other hand, tested the expectations hypothesis in the complete markets endowment economy (Lucas model) by starting with (27), assuming that the risk premium was constant

$$E_t[p^b_{1,t+1}] - p^f_{1,t} = a,$$

and then regressed

$$p^b_{1,t+1} - p^f_{1,t} = a + b(p^f_{1,t} - p^b_{1,t}) \tag{29}$$

to see if $b = 0$. They generated 200 observations 1000 times and used Wald test with White (1980) standard errors to check if $b = 0$ with 5% significance level. They could reject the hypothesis only roughly 50 times out of 1000 regressions which is what one would expect from chance alone. On

[10] If one restricts the attention to 1980:1 to 1998:4 data sample, the the term structure of volatility is clearly downward-sloping. The standard deviation of the three-month yield is 3.05 and the standard deviation of the 20-year yield is 1.77. In addition, in the UK nominal and real data the term structure of volatilities is downward-sloping, see Seppälä (2000).

Table 3: Main term structure statistics. (N/A missing due to shortage of data.)

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| 3-month real yield (model) | 2.06381 | 3.23813 | 0.08726 |
| 1-year real yield (model) | 3.32266 | 2.91786 | 0.35683 |
| 10-year real yield (model) | 3.93816 | 1.00484 | 0.58524 |
| 20-year real yield (model) | 3.97521 | 0.55389 | 0.58615 |
| 30-year real yield (model) | 3.99029 | 0.37829 | 0.58336 |
| 10-year minus 3-month real (model) | 1.87435 | 2.71671 | 0.11245 |
| 20-year minus 3-month real (model) | 1.91140 | 2.92054 | 0.01442 |
| 30-year minus 3-month real (model) | 1.92648 | 3.01215 | −0.02054 |
| 10-year minus 1-year real (model) | 0.61550 | 2.08035 | −0.21780 |
| 20-year minus 1-year real (model) | 0.65255 | 2.44158 | −0.29346 |
| 30-year minus 1-year real (model) | 0.66764 | 2.58823 | −0.31701 |
| 3-month nominal yield (model) | 5.00436 | 5.15380 | 0.47255 |
| 1-year nominal yield (model) | 6.33660 | 4.97113 | 0.56127 |
| 10-year nominal yield (model) | 6.75250 | 1.53730 | 0.61543 |
| 20-year nominal yield (model) | 6.71831 | 0.87280 | 0.60586 |
| 30-year nominal yield (model) | 6.69836 | 0.59335 | 0.60582 |
| 10-year minus 3-month nominal (model) | 1.74813 | 3.78465 | −0.39352 |
| 20-year minus 3-month nominal (model) | 1.71394 | 4.35994 | −0.43731 |
| 30-year minus 3-month nominal (model) | 1.69399 | 4.61190 | −0.45013 |
| 10-year minus 1-year nominal (model) | 0.41589 | 3.47529 | −0.53061 |
| 20-year minus 1-year nominal (model) | 0.38170 | 4.11731 | −0.54923 |
| 30-year minus 1-year nominal (model) | 0.36175 | 4.39051 | −0.55362 |
| 3-month nominal yield (data) | 5.06450 | 3.05130 | 0.14735 |
| 1-year nominal yield (data) | 5.47900 | 3.12905 | 0.12193 |
| 10-year nominal yield (data) | 6.22462 | 2.96448 | −0.00705 |
| 20-year nominal yield (data) | 6.54877 | 3.15684 | −0.05252 |
| 30-year nominal yield (data) | N/A | N/A | N/A |
| 10-year minus 3-month nominal (data) | 1.16012 | 1.15935 | −0.40586 |
| 20-year minus 3-month nominal (data) | 1.04854 | 1.32304 | −0.41947 |
| 30-year minus 3-month nominal (data) | N/A | N/A | N/A |
| 10-year minus 1-year nominal (data) | 0.74561 | 0.96807 | −0.41570 |
| 20-year minus 1-year nominal (data) | 0.62854 | 1.13526 | −0.39922 |
| 30-year minus 1-year nominal (data) | N/A | N/A | N/A |

Table 4: The number of rejects in each regressions in the benchmark model for nominal term structure.

| $y_{t+1}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ |
|---|---|---|---|---|
| $x_t$ | $p^f_{1,t} - p^b_{1,t}$ | $p^f_{1,t} - p^b_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ |
| $Wald(a = b = 0)$ | 1000 | 52 | 1000 | 66 |
| $Wald(b = 0)$ | 948 | 61 | 92 | 63 |
| $Wald(b = -1)$ | 1000 | 1000 | 1000 | 1000 |

Table 5: The number of rejects in each regressions in the benchmark model for nominal term structure when $b = 0$.

| $y_{t+1}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ |
|---|---|---|---|---|
| $x_t$ | $p^f_{1,t} - p^b_{1,t}$ | $p^f_{1,t} - p^b_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ |
| $Wald(a = b = 0)$ | 72 | 56 | 54 | 42 |
| $Wald(b = 0)$ | 67 | 56 | 66 | 56 |
| $Wald(b = -1)$ | 1000 | 1000 | 1000 | 1000 |

the other hand, for all values of $b$ except $-1$, the forward premium is still useful in forecasting the changes in spot prices. The hypothesis $b = -1$ was rejected every time.

Table 4 presents the number of rejections of different Wald tests in the regressions

$$y_{t+1} = a + bx_t$$

in our benchmark model for nominal term structure. Table 5 presents the same tests when the habit-formation parameter $b = 0$. Table 6 displays the same tests for real term structure, and table 7 displays the test for real term structure with $b = 0$. Only our benchmark model is roughly consistent with empirical evidence on the expectations hypothesis. The model can generate enough variation in the risk premia to account for the rejections of the expectations hypothesis 95% of the time. On the hand, when the risk premium is substracted from $p^b_{1,t+1} - p^f_{1,t}$ $b$ is equal to zero with 5% significance level. Comparing the tables, is is clear that habit-formation is a necessary condition for the rejection of expectations hypothesis. However, since the hypothesis is rejected for real term structure only about 75% of the time, it seems to be the case the monetary policy, which mostly affects nominal rates, plays also an important role. This issue is studied in more detail in Section 6.

In Table 8 the results of the regression (29) are presented for one realization of 200 real and nominal observations and for the data. The data are quarterly observations from 1960:1 to 1998:4 of three and six-month U.S. Treasury bills. In Table 8, *Wald* rows refer to the marginal significance level of the corresponding Wald test. The expectations hypothesis can be rejected at 5% critical level for simulated nominal data but not for simulated real data. We return to this question in Section 6.

Table 6: The number of rejects in each regressions in the benchmark model for real term structure.

| $y_{t+1}$ <br> $x_t$ | $p^b_{1,t+1} - p^f_{1,t}$ <br> $p^f_{1,t} - p^b_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ <br> $p^f_{1,t} - p^b_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t}$ <br> $p^b_{1,t} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ <br> $p^b_{1,t} - p^f_{1,t}$ |
|---|---|---|---|---|
| $Wald(a = b = 0)$ | 1000 | 58 | 1000 | 55 |
| $Wald(b = 0)$ | 775 | 47 | 102 | 61 |
| $Wald(b = -1)$ | 1000 | 1000 | 1000 | 1000 |

Table 7: The number of rejects in each regressions in the benchmark model for real term structure when $b = 0$.

| $y_{t+1}$ <br> $x_t$ | $p^b_{1,t+1} - p^f_{1,t}$ <br> $p^f_{1,t} - p^b_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ <br> $p^f_{1,t} - p^b_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t}$ <br> $p^b_{1,t} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ <br> $p^b_{1,t} - p^f_{1,t}$ |
|---|---|---|---|---|
| $Wald(a = b = 0)$ | 68 | 68 | 61 | 60 |
| $Wald(b = 0)$ | 71 | 68 | 65 | 52 |
| $Wald(b = -1)$ | 1000 | 1000 | 1000 | 1000 |

Table 8: The tests of the expectations hypothesis in a single regression.

| Variable/Test | Benchmark Real | Benchmark Nominal | Data |
|---|---|---|---|
| $a$ | 0.0035 | 0.0035 | 0.0008 |
| se($a$) | 0.0004 | 0.0003 | 0.0005 |
| $b$ | $-0.1198$ | $-0.2189$ | $-0.4866$ |
| se($b$) | 0.0795 | 0.0601 | 0.1458 |
| $R^2$ | 0.0148 | 0.0915 | 0.1505 |
| $Wald(a = b = 0)$ | 0 | 0 | 0.0003 |
| $Wald(b = 0)$ | 0.13206 | 0.000027 | 0.0008 |
| $Wald(b = -1)$ | 0 | 0 | 0.0004 |

Table 9: Expectations hypothesis regressions in rates.

| Regression | $a$ | se($a$) | $b$ | se($b$) | $R^2$ |
|---|---|---|---|---|---|
| Benchmark ($n = 2$) | −0.1028 | 0.2181 | 0.3581 | 0.2164 | 0.0114 |
| Benchmark ($n = 3$) | 0.1122 | 0.4317 | 0.2046 | 0.2307 | 0.0032 |
| Benchmark ($n = 4$) | 0.2858 | 0.5953 | 0.1245 | 0.2318 | 0.0013 |
| Benchmark ($n = 5$) | 0.3477 | 0.7080 | 0.1214 | 0.2315 | 0.0014 |
| Benchmark ($n = 6$) | 0.4433 | 0.7838 | 0.1686 | 0.2216 | 0.0029 |
| Benchmark ($n = 11$) | 0.4736 | 0.9207 | 0.2916 | 0.1871 | 0.0129 |
| Data ($n = 2$) | 0.1259 | 0.1720 | −1.3078 | 0.4016 | 0.0682 |
| Data ($n = 3$) | 0.4494 | 0.3349 | −1.6873 | 0.5588 | 0.0758 |
| Data ($n = 4$) | 0.9220 | 0.4552 | −1.9917 | 0.6277 | 0.0851 |
| Data ($n = 5$) | 1.2799 | 0.5718 | −2.2363 | 0.7023 | 0.0861 |
| Data ($n = 6$) | 1.5000 | 0.6808 | −2.5348 | 0.7715 | 0.0922 |
| Data ($n = 11$) | 3.0204 | 1.2524 | −3.3127 | 1.2728 | 0.0722 |

Recent empirical literature has concentrated on the Log Pure Expectations Hypothesis. According to the hypothesis, the $n$-period forward rate should equal the expected one-period interest rate $n$ periods ahead:

$$f_{n,t} = E_t[r_{1,t+n}].$$

To test the hypothesis, one can run the regression

$$(n-1) * (r_{n-1,t+1} - r_{n,t}) = a + b(r_{n,t} - r_{1,t}) \quad \text{for } n = 2, 3, 4, 5, 6, 11 \text{ years.} \tag{30}$$

According to the Log Pure Expectations Hypothesis, one should find that $b = 1$.[11] Table 9 summarizes the results from this regression for the models and data from 1960:1 to 1998:4. The expectations hypothesis is again clearly rejected both in the model and in data.[12]

*The Term Structure Predictions of Future Economic Activity*

Despite the fact that the expectations hypothesis has been rejected over and over again in the empirical literature, it has also been found that the term and forward spreads forecast changes in the interest rates, consumption growth, and other economic activity. In this section, we will compare the predictions of our benchmark model to two famous empirical papers on the term structure and the future economic activity.

The first paper is by Fama and Bliss (1987) who use forward spread to predict the future changes in one-year interest rates one to four years ahead. Table 10 presents the regression results of equation

$$r_{1,t+n} - r_{1,t} = a + b(f_{n,t} - r_{1,t}) \quad \text{for } n = 1, 2, 3, 4 \text{ years}$$

[11] See, e.g., Campbell, Lo, and McKinley (1997).

[12] The coefficients of $b$, however, have different signs in the model and in data. The model doesn't seem to capture the full magnitude of how clearly the expectations hypothesis is rejected in data. It is interesting to note that the model coefficients are quite close to the data on UK nominal yields as presented in Seppälä (2000).

Table 10: Forward spread forecasts of future interest rate changes $n$ years ahead.

| Regression | $a$ | se($a$) | $b$ | se($b$) | $R^2$ |
|---|---|---|---|---|---|
| Benchmark ($n = 1$) | 0.2468 | 0.1296 | 0.0811 | 0.0267 | 0.0424 |
| Benchmark ($n = 2$) | 1.5229 | 0.4965 | 0.4651 | 0.0758 | 0.1384 |
| Benchmark ($n = 3$) | 1.8843 | 0.5520 | 0.5627 | 0.0731 | 0.2175 |
| Benchmark ($n = 4$) | 1.9653 | 0.5424 | 0.6389 | 0.0689 | 0.2995 |
| Data ($n = 1$) | 0.6546 | 0.1255 | −1.2593 | 0.1828 | 0.2719 |
| Data ($n = 2$) | 0.0019 | 0.2788 | 0.1575 | 0.2533 | 0.0034 |
| Data ($n = 3$) | −0.1298 | 0.3074 | 0.5794 | 0.2744 | 0.0431 |
| Data ($n = 4$) | −0.3871 | 0.2911 | 0.8665 | 0.2163 | 0.0950 |

for the data from 1960:1 to 1998:4 and the benchmark model with 200 observations. The standard errors are White (1980) heteroskedasticity consistent standard errors. In both cases, $b$ increases with the forecast horizon. With longer maturities, the match is quite good. On the other hand, $R^2$ increases in the model and decreases in data with the forecast horizon. It should be noted that in the original Fama and Bliss paper, the $R^2$ also increased with forecast horizon as in our model. The main difference between our data and the data used by Fama and Bliss is that the latter used monthly data from January 1965 to December 1984. In their data sample, the interest rates have strong mean reverting property that increases the forecast power in longer horizons. On the other hand, in our sample the downward trend in data since the early 1980's dominates the data and decreases the forecasting power in longer horizons.

The second paper is Estrella and Hardouvelis (1991) who use the term spread to predict the future changes in the log consumption growth one to four years ahead. The data are quarterly observations from 1960:1 to 1998:4 of U.S. consumption non-durables plus services regressed on 10-year government bonds less three-month Treasury bill rates. Table 11 presents the regression results of equation

$$(100/n) * (\log(c_{t+n}) - \log(c_t)) = a + b(r_{10,t} - r_{1,t}) \quad \text{for } n = 1, 2, 3, 4 \text{ years}$$

for the data and the benchmark model. The standard errors are White (1980) heteroskedasticity consistent standard errors. Upward-sloping term structure clearly predicts expansions both in our model and in data, and downward-sloping term structure clearly predicts recessions, again, both in the model and in data. Again, $R^2$ increases in the model and decreases in data with the forecast horizon. This feature of the model is largely a result the high shocks autocorrelation (see the discussion page 16).

Table 11: Term spread forecasts of future consumption growth $n$ years ahead.

| Regression | $a$ | se($a$) | $b$ | se($b$) | $R^2$ |
|---|---|---|---|---|---|
| Benchmark ($n=1$) | 0.0385 | 0.1939 | 0.1413 | 0.0384 | 0.0656 |
| Benchmark ($n=2$) | 0.0697 | 0.1462 | 0.1418 | 0.0270 | 0.1194 |
| Benchmark ($n=3$) | 0.0403 | 0.1155 | 0.1255 | 0.0227 | 0.1583 |
| Benchmark ($n=4$) | 0.0090 | 0.0899 | 0.1176 | 0.0178 | 0.2141 |
| Data ($n=1$) | 3.0990 | 0.1353 | 0.3164 | 0.1828 | 0.0758 |
| Data ($n=2$) | 3.1785 | 0.1156 | 0.2070 | 0.0772 | 0.0504 |
| Data ($n=3$) | 3.2580 | 0.1027 | 0.0564 | 0.0657 | 0.0052 |
| Data ($n=4$) | 3.3688 | 0.0882 | −0.0427 | 0.0523 | 0.0038 |

## *6. Monetary Policy and Inflation Risk Premium*

Recall the definitions of one period zero-coupon nominal bond (25) and the nominal stochastic discount factor (23)

$$p_t^B = E_t[M_{t+1}] = E_t\left[\beta\frac{MUC_{t+1}P_t}{MUC_tP_{t+1}}\right]. \tag{31}$$

To define the inflation risk premium, write (31) using the definition conditional covariance and the definition of real bond price (24):

$$\begin{aligned} p_t^B &= E_t\left[\beta\frac{MUC_{t+1}P_t}{MUC_tP_{t+1}}\right] \\ &= E_t\left[\beta\frac{MUC_{t+1}}{MUC_t}\right]E_t\left[\frac{P_t}{P_{t+1}}\right] + \text{cov}_t\left[\beta\frac{MUC_{t+1}}{MUC_t},\frac{P_t}{P_{t+1}}\right] \\ &= p_t^bE_t\left[\frac{P_t}{P_{t+1}}\right] + \text{cov}_t\left[m_{t+1},\frac{P_t}{P_{t+1}}\right]. \end{aligned}$$

Since the conditional covariance term is zero for risk-neutral investors and when inflation process is deterministic, we call it the *inflation risk premium*, $irp_{t,1}$, given by

$$irp_{t,1} \equiv \text{cov}_t\left[m_{t+1},\frac{P_t}{P_{t+1}}\right] = p_t^B - p_t^bE_t\left[\frac{P_t}{P_{t+1}}\right],$$

and similarly $irp_{t,n}$ is the $n$-period inflation risk premium:

$$irp_{t,n} \equiv \text{cov}_t\left[\prod_{j=1}^n m_{t+j},\frac{P_t}{P_{t+n}}\right] = p_{t,n}^B - p_{t,n}^bE_t\left[\frac{P_t}{P_{t+n}}\right].$$

Assuming that the inflation risk premium is zero, we get the Fisher hypothesis:

$$p_{t,n}^B = p_{t,n}^bE_t\left[\frac{P_t}{P_{t+n}}\right]$$

Table 12: Main inflation risk premia statistics.

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| Inflation risk premium ($n = 1$) | $-0.02152$ | 0.00587 | $-0.41487$ |
| Inflation risk premium ($n = 2$) | $-0.05897$ | 0.01254 | $-0.40707$ |
| Inflation risk premium ($n = 4$) | $-0.16307$ | 0.02295 | $-0.36519$ |
| Inflation risk premium ($n = 8$) | $-0.38759$ | 0.03534 | $-0.20202$ |
| Inflation risk premium ($n = 12$) | $-0.55246$ | 0.04527 | 0.18787 |
| Inflation risk premium ($n = 16$) | $-0.64817$ | 0.06306 | 0.42333 |
| Inflation risk premium ($n = 20$) | $-0.69205$ | 0.08089 | 0.47417 |
| Inflation risk premium ($n = 24$) | $-0.70403$ | 0.09668 | 0.47655 |
| Inflation risk premium ($n = 28$) | $-0.69623$ | 0.10807 | 0.48398 |
| Inflation risk premium ($n = 32$) | $-0.67399$ | 0.11668 | 0.48397 |
| Inflation risk premium ($n = 36$) | $-0.64554$ | 0.12241 | 0.47302 |
| Inflation risk premium ($n = 40$) | $-0.61329$ | 0.12529 | 0.49829 |
| Inflation risk premium ($n = 60$) | $-0.45001$ | 0.09238 | 0.48880 |
| Inflation risk premium ($n = 80$) | $-0.33307$ | 0.06904 | 0.44080 |

or by taking logs and multiplying by $-(1/n)$:

$$R_{t,n} \approx r_{t,n} + \frac{1}{n} E_t \left[ \log \left( \frac{P_{t+n}}{P_t} \right) \right].$$

That is, nominal interest rate equals the sum of the (ex-ante) real interest rate and the average expected inflation.

Table 12 presents the main statistics for the inflation risk premia in our benchmark case, and Figures 1–6 show the inflation risk premium (and other variables) respond to one standard deviation shock in preferences, productivity, and monetary policy, respectively.

As Table 12 shows, the inflation risk premium is always negative, and hence positive preference shock that increases inflation decreases inflation risk premium. That is, as inflation increases inflation risk premium becomes more negative or larger. Similarly, positive productivity shock decreases both inflation and the size of inflation risk premium (moves it closer to zero). Finally, positive monetary policy shock decreases both inflation and the size of inflation risk premium.

Because the inflation risk premium is unobservable in data, it is hard to access the mean and standard deviation of premia in Table 12. The best we can do is to compare how different parameters affect the size and volatility of premia. Buraschi and Jiltsov (2005) argue that the time-variation of the inflation risk premium is an important explanatory variable of deviations from the expectations hypothesis. We address this question by varying the "aggressiveness" of the monetary with respect to inflation, i.e., the parameter $\omega_\pi$. Table 13 presents the inflation risk premia statistics when $\omega_\pi = 3$. Notice that premia are considerably smaller and less volatile than in our benchmark with more aggressive monetary policy. This is not surprising. The more aggressive is the policy the lower and less volatile is inflation, and hence investors do not ask as much compensation for inflation volatility.

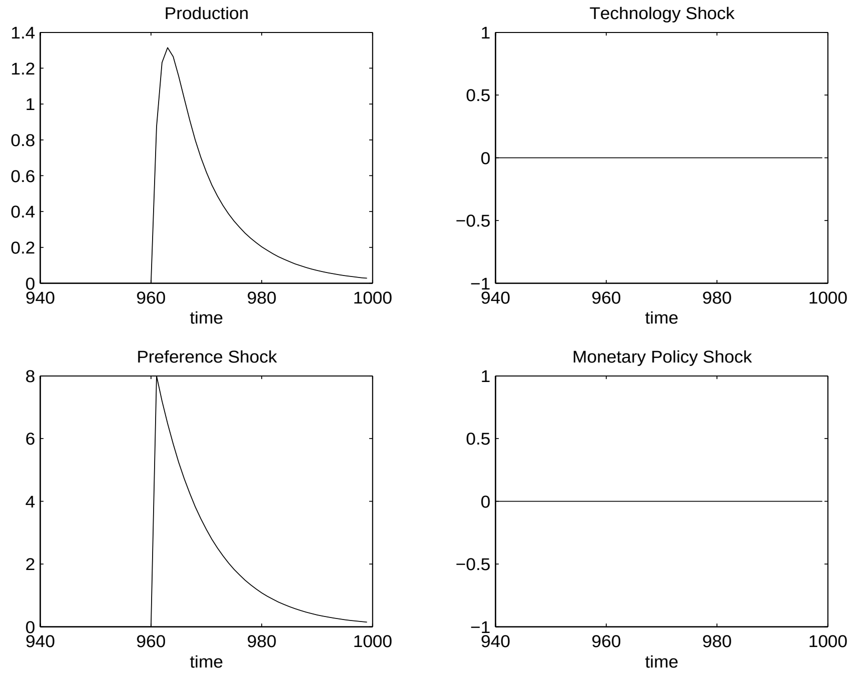


Figure 1: Impulse response functions to one standard deviation preference shock in the benchmark parameterization.

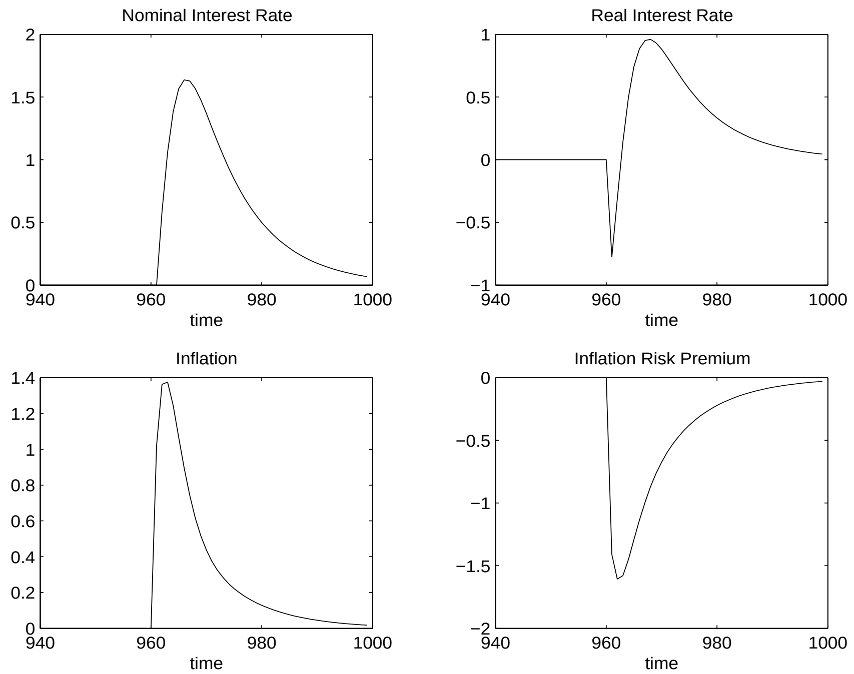


Figure 2: Impulse response functions to one standard deviation preference shock in the benchmark parameterization.

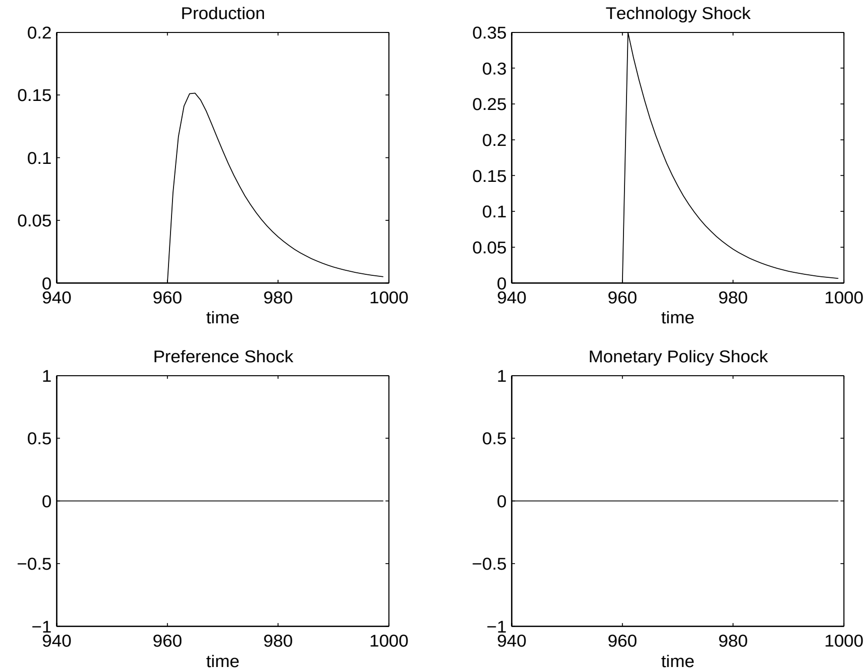


Figure 3: Impulse response functions to one standard deviation productivity shock in the benchmark parameterization.

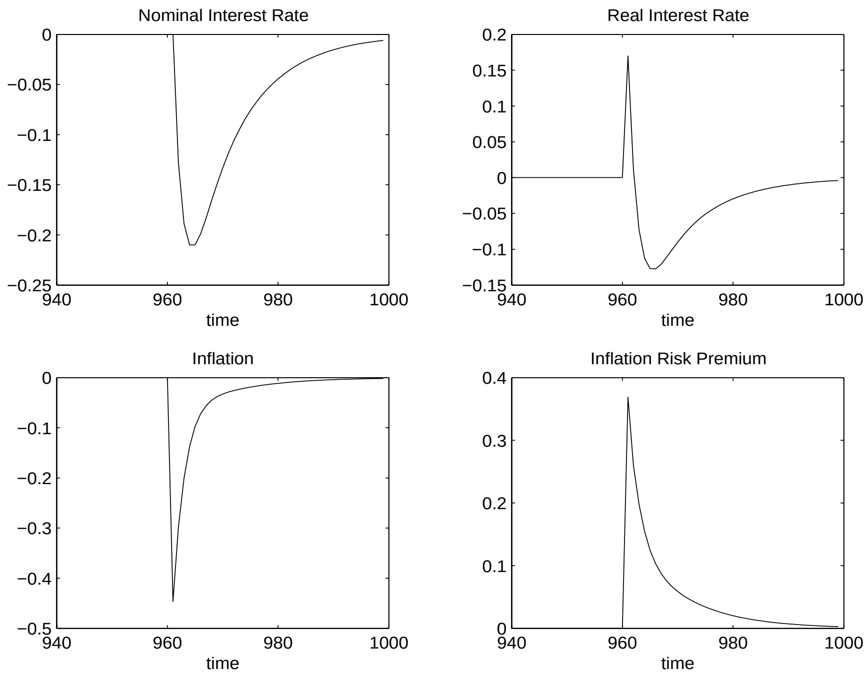


Figure 4: Impulse response functions to one standard deviation productivity shock in the benchmark parameterization.

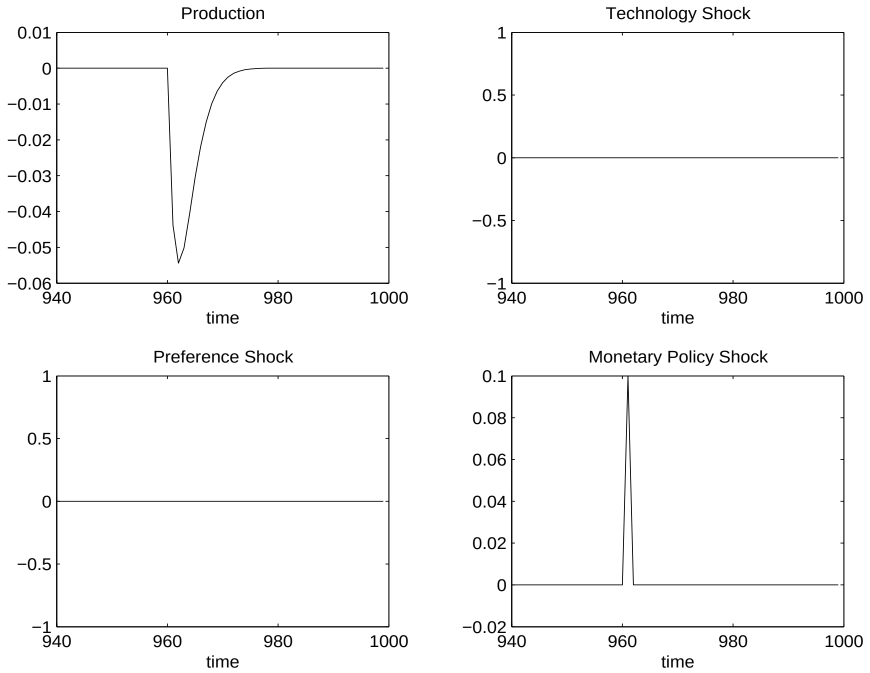


Figure 5: Impulse response functions to one standard deviation monetary policy shock in the benchmark parameterization.

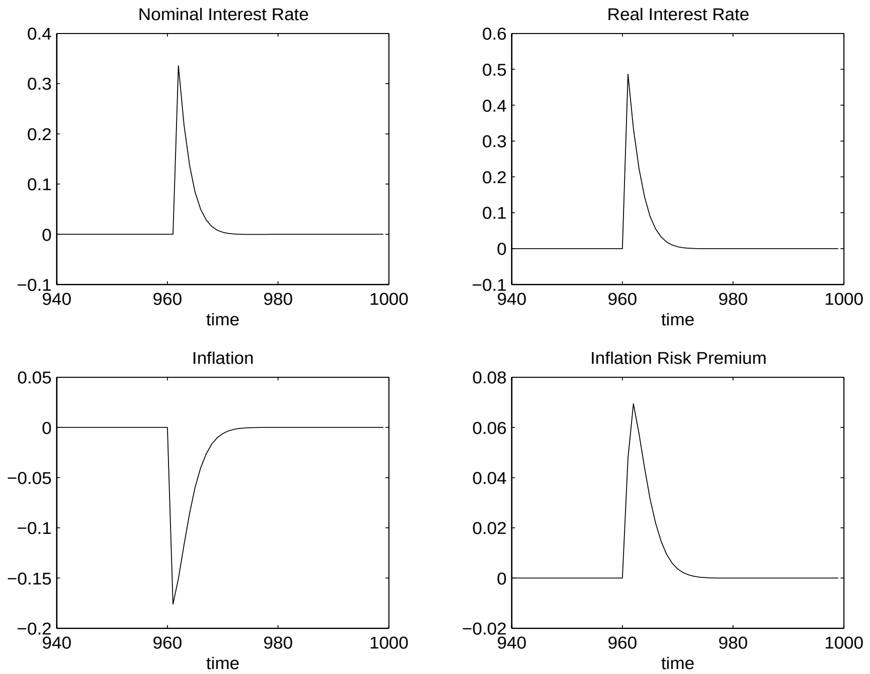


Figure 6: Impulse response functions to one standard deviation monetary policy shock in the benchmark parameterization.

Table 13: Main inflation risk premia statistics when $\omega_\pi = 3$.

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| Inflation risk premium ($n = 1$) | −0.00862 | 0.00417 | −0.54447 |
| Inflation risk premium ($n = 2$) | −0.02999 | 0.00873 | −0.53235 |
| Inflation risk premium ($n = 4$) | −0.09462 | 0.01505 | −0.51060 |
| Inflation risk premium ($n = 8$) | −0.23582 | 0.02018 | −0.37049 |
| Inflation risk premium ($n = 12$) | −0.33871 | 0.02286 | 0.07227 |
| Inflation risk premium ($n = 16$) | −0.39823 | 0.03216 | 0.38099 |
| Inflation risk premium ($n = 20$) | −0.42545 | 0.04301 | 0.44461 |
| Inflation risk premium ($n = 24$) | −0.43273 | 0.05210 | 0.45677 |
| Inflation risk premium ($n = 28$) | −0.42779 | 0.05892 | 0.47216 |
| Inflation risk premium ($n = 32$) | −0.41364 | 0.06449 | 0.46797 |
| Inflation risk premium ($n = 36$) | −0.39583 | 0.06882 | 0.45334 |
| Inflation risk premium ($n = 40$) | −0.37563 | 0.07068 | 0.48285 |
| Inflation risk premium ($n = 60$) | −0.27331 | 0.05142 | 0.47720 |
| Inflation risk premium ($n = 80$) | −0.20098 | 0.03938 | 0.41296 |

Table 14: The number of rejects in each regressions in the benchmark model for nominal term structure when $\omega_\pi = 3$.

| $y_{t+1}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ |
|---|---|---|---|---|
| $x_t$ | $p^f_{1,t} - p^b_{1,t}$ | $p^f_{1,t} - p^b_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ |
| $Wald(a = b = 0)$ | 1000 | 50 | 1000 | 71 |
| $Wald(b = 0)$ | 983 | 56 | 97 | 69 |
| $Wald(b = -1)$ | 1000 | 1000 | 1000 | 1000 |

However, Table 14 shows that the expectations hypothesis is actually rejected *more* often when the policy is more aggressive. Moreover, Table 15 presents the inflation risk premia statistics when $\omega_\pi = 1.2$. The premia are much larger and more volatile. But Table 16 shows that the expectations hypothesis is actually rejected *less* often when the policy is less aggressive. That is, unlike in Buraschi and Jiltsov (2005), inflation risk premium is *not* the explanation for the rejections of expectations hypothesis.

How can that be? Mankiw and Summers (1984) and Mankiw and Miron (1986) show that the estimate of $b$ in the expectations hypothesis regression in rates (30) converges to

$$\text{plim}\,\hat{b} = \frac{\text{var}(E_t\Delta r_{t+1}) + 2\,\text{corr}(\psi_t, E_t\Delta r_{t+1})\,\text{std}(E_t\Delta r_{t+1})\,\text{std}(\psi_t)}{\text{var}(E_t\Delta r_{t+1}) + 4\,\text{var}(\psi_t) + 4\,\text{corr}(\psi_t, E_t\Delta r_{t+1})\,\text{std}(E_t\Delta r_{t+1})\,\text{std}(\psi_t)}, \tag{32}$$

where $\psi_t$ is the term premium. Notice that if $\psi_t$ is deterministic, i.e., $\text{std}(\psi_t) = 0$, $\text{plim}\,\hat{b} = 1$. Also, observe that as $\text{std}(\psi_t)$ increases, $\text{plim}\,\hat{b}$ decreases. Finally, (32) is a complicated function

Table 15: Main inflation risk premia statistics when $\omega_\pi = 1.2$.

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| Inflation risk premium ($n = 1$) | $-0.04423$ | 0.01387 | 0.42590 |
| Inflation risk premium ($n = 2$) | $-0.11610$ | 0.03258 | 0.41053 |
| Inflation risk premium ($n = 4$) | $-0.33723$ | 0.07986 | 0.40715 |
| Inflation risk premium ($n = 8$) | $-0.88928$ | 0.17094 | 0.38337 |
| Inflation risk premium ($n = 12$) | $-1.33503$ | 0.25524 | 0.45048 |
| Inflation risk premium ($n = 16$) | $-1.61328$ | 0.33652 | 0.50424 |
| Inflation risk premium ($n = 20$) | $-1.75971$ | 0.39659 | 0.51349 |
| Inflation risk premium ($n = 24$) | $-1.82118$ | 0.45157 | 0.49452 |
| Inflation risk premium ($n = 28$) | $-1.82818$ | 0.49612 | 0.48720 |
| Inflation risk premium ($n = 32$) | $-1.79790$ | 0.52048 | 0.49139 |
| Inflation risk premium ($n = 36$) | $-1.74663$ | 0.52422 | 0.48779 |
| Inflation risk premium ($n = 40$) | $-1.68178$ | 0.52357 | 0.50495 |
| Inflation risk premium ($n = 60$) | $-1.32811$ | 0.42376 | 0.47136 |
| Inflation risk premium ($n = 80$) | $-1.04778$ | 0.34960 | 0.40516 |

Table 16: The number of rejects in each regressions in the benchmark model for nominal term structure when $\omega_\pi = 1.2$.

| $y_{t+1}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ |
|---|---|---|---|---|
| $x_t$ | $p^f_{1,t} - p^b_{1,t}$ | $p^f_{1,t} - p^b_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ |
| $Wald(a = b = 0)$ | 1000 | 59 | 1000 | 77 |
| $Wald(b = 0)$ | 286 | 56 | 77 | 64 |
| $Wald(b = -1)$ | 1000 | 1000 | 1000 | 1000 |

Table 17: Main inflation risk premia statistics when $\chi = 0$.

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| Inflation risk premium ($n = 1$) | $-0.04142$ | 0.02792 | $-0.58377$ |
| Inflation risk premium ($n = 2$) | $-0.10549$ | 0.05441 | $-0.58067$ |
| Inflation risk premium ($n = 4$) | $-0.27086$ | 0.09143 | $-0.56336$ |
| Inflation risk premium ($n = 8$) | $-0.61949$ | 0.11888 | $-0.42942$ |
| Inflation risk premium ($n = 12$) | $-0.89103$ | 0.12351 | $-0.12946$ |
| Inflation risk premium ($n = 16$) | $-1.06158$ | 0.14742 | 0.18314 |
| Inflation risk premium ($n = 20$) | $-1.14887$ | 0.17832 | 0.33769 |
| Inflation risk premium ($n = 24$) | $-1.18473$ | 0.21687 | 0.41650 |
| Inflation risk premium ($n = 28$) | $-1.18562$ | 0.24489 | 0.46531 |
| Inflation risk premium ($n = 32$) | $-1.15817$ | 0.26308 | 0.49031 |
| Inflation risk premium ($n = 36$) | $-1.12056$ | 0.28413 | 0.49048 |
| Inflation risk premium ($n = 40$) | $-1.07382$ | 0.29335 | 0.49224 |
| Inflation risk premium ($n = 60$) | $-0.81736$ | 0.22491 | 0.44980 |
| Inflation risk premium ($n = 80$) | $-0.62867$ | 0.17684 | 0.44566 |

Table 18: The number of rejects in each regressions in the benchmark model for nominal term structure when $\chi = 0$.

| $y_{t+1}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ |
|---|---|---|---|---|
| $x_t$ | $p^f_{1,t} - p^b_{1,t}$ | $p^f_{1,t} - p^b_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ |
| $Wald(a = b = 0)$ | 1000 | 63 | 1000 | 75 |
| $Wald(b = 0)$ | 764 | 66 | 86 | 74 |
| $Wald(b = -1)$ | 1000 | 1000 | 1000 | 1000 |

of $\text{std}(E_t\Delta r_{t+1})$, but if $\text{std}(E_t\Delta r_{t+1}) = 0$, then $\text{plim}\,\hat{b} = 0$. Moreover, as $\text{std}(E_t\Delta r_{t+1}) \to \infty$, $\text{plim}\,\hat{b} \to 1$. Hence, the lower is $\text{std}(E_t\Delta r_{t+1})$ the easier it is to reject the expectations hypothesis.

Empirical evidence supports this interpretation. Mankiw and Miron (1986) show that it is much more difficult to reject the expectations hypothesis using data prior to the founding of the Fed.[13] They suggest that the explanation is the Federal Reserve's commitment to stabilizing interest rates.

We next study the issue by setting the interest rate smoothing parameter $\chi = 0$. As shown in Tables 17 and 18, similarly as with $\omega_\pi = 1.2$, we get *larger* inflation risk premium and *fewer* rejections of expectations hypothesis.

[13] Choi and Wohar (1991) cannot reject the expectations hypothesis over the sample period of 1910–1914.

Table 19: Main term structure statistics when $\sigma_{mp} = 0$.

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| 3-month nominal yield | 5.02277 | 5.13179 | 0.47615 |
| 1-year nominal yield | 6.34453 | 4.96382 | 0.56358 |
| 10-year nominal yield | 6.75754 | 1.53661 | 0.61642 |
| 20-year nominal yield | 6.72312 | 0.87224 | 0.60621 |
| 30-year nominal yield | 6.70322 | 0.59275 | 0.60425 |

Table 20: Main term structure statistics when $\sigma_a = 0$.

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| 3-month nominal yield | 5.03020 | 5.10918 | 0.48620 |
| 1-year nominal yield | 6.34650 | 4.93666 | 0.57515 |
| 10-year nominal yield | 6.76003 | 1.52846 | 0.63147 |
| 20-year nominal yield | 6.72685 | 0.86730 | 0.62396 |
| 30-year nominal yield | 6.70767 | 0.58829 | 0.62600 |

## 7. *Which Shocks Matter?*

In our baseline model, we have three different shocks: a technology shock, a preference shock, and a monetary policy shock. In this section we investigate the importance of these different shocks.

*No Monetary Policy Shocks* ($\sigma_{mp} = 0$)

With no monetary policy shocks, the results are almost exactly the same as in our benchmark case, see Table 19.

*No Technology Shocks* ($\sigma_a = 0$)

With no technology shocks, the results are almost exactly the same as in our benchmark case, see Table 20. Since neither monetary policy nor technology shocks matter for our results, it must be the case that the demand shock plays the most crucial role in the model.

*No Demand Shocks* ($\sigma_d = 0$)

Without demand shocks, the interest rates are much less volatile and flatter as presented in Table 21. In addition, interest rates and inflation are countercyclical. However, the expectations hypothesis is still rejected roughly 50% of the time as shown in Table 22.

Table 21: Main term structure statistics when $\sigma_d = 0$.

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| 3-month nominal yield | 6.90371 | 1.29485 | −0.63981 |
| 1-year nominal yield | 6.97045 | 1.13743 | −0.63810 |
| 10-year nominal yield | 6.98185 | 0.30401 | −0.61539 |
| 20-year nominal yield | 6.97930 | 0.17013 | −0.61407 |
| 30-year nominal yield | 6.97820 | 0.11473 | −0.61255 |

Table 22: The number of rejects in each regressions in the benchmark model for nominal term structure when $\sigma_d = 0$.

| $y_{t+1}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ |
|---|---|---|---|---|
| $x_t$ | $p^f_{1,t} - p^b_{1,t}$ | $p^f_{1,t} - p^b_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ |
| $Wald(a = b = 0)$ | 879 | 52 | 668 | 58 |
| $Wald(b = 0)$ | 549 | 64 | 118 | 56 |
| $Wald(b = -1)$ | 999 | 1000 | 1000 | 1000 |

*Summary*

Summarizing, monetary policy and technology shocks are not crucial for our results. Preference shocks are crucial for matching the cyclical behavior of the interest rates but not for the rejections of the expectations hypothesis. These results are somewhat similar to Nakajima (2003) who shows that the standard RBC model driven by the Solow residual cannot explain the "preference residual" (the difference between real wage and the marginal rate of substitution between consumption and leisure), but the model driven by the preference residual can account for the Solow residual.

# 8. Conclusions

We show that a New Keynesian model with habit-persistent preferences and a monetary policy feedback rule with interest rate smoothing produces pro-cyclical interest rates, counter-cyclical term spreads, and creates enough volatility in the risk premium to account for the rejections of expectations hypothesis. Our results are related to conclusion reached by Dotsey and Otrok (1995)

> "[R]egression results [for the expectations hypothesis] that are in accord with those obtained in practise can be generated by the combination of (i) Fed behavior that both smooths the movements in interest rates... and (ii) time-varying term premia that are calibrated to match data moments."

In our model, habit formation delivers (ii) and interest rate smoothing delivers (i). Without habit formation, we reject the expectations hypothesis only 5% of the time. With habit formation but without interest rate smoothing, we reject 75% of the time. With habit formation and interest rate smoothing, we reject the expectation hypothesis 95% of the time. It is important to note that in our model, it is the *systematic* monetary policy that brings our results, and not the monetary policy shocks as in Buraschi and Jiltsov (2005). We also find that monetary policy and the technology shocks are not crucial for our results. Preference shocks are crucial for matching the cyclical behavior of the interest rates but not for the rejections of the expectations hypothesis.

## A. The Inflation Rate Dynamics

By iterating (16) and (17), we obtain:

$$\begin{aligned}
\hat{G}_{t+1} &= \mu MUC_{t+1} MC_{t+1} P_{t+1}^{\theta-1} Y_{t+1} + \theta_p \beta \hat{G}_{t+2} \\
\hat{G}_{t+2} &= \mu MUC_{t+2} MC_{t+2} P_{t+2}^{\theta-1} Y_{t+2} + \theta_p \beta \hat{G}_{t+3} \\
\hat{H}_{t+1} &= MUC_{t+1} P_{t+1}^{\theta-1} Y_{t+1} + \theta_p \beta \hat{H}_{t+2} \\
\hat{H}_{t+2} &= MUC_{t+2} P_{t+2}^{\theta-1} Y_{t+2} + \theta_p \beta \hat{H}_{t+3}
\end{aligned}$$

or

$$\begin{aligned}
\hat{G}_t &= \mu MUC_t MC_t P_t^{\theta-1} Y_t + \theta_p \beta \left[ \mu MUC_{t+1} MC_{t+1} P_{t+1}^{\theta-1} Y_{t+1} + \theta_p \beta \hat{G}_{t+2} \right] \\
&= \mu MUC_t MC_t P_t^{\theta-1} Y_t + \theta_p \beta \Big[ \mu MUC_{t+1} MC_{t+1} P_{t+1}^{\theta-1} Y_{t+1} \\
&+ \theta_p \beta \left( \mu MUC_{t+2} MC_{t+2} P_{t+2}^{\theta-1} Y_{t+2} + \theta_p \beta \hat{G}_{t+3} \right) \Big] \\
&= \mu MUC_t MC_t P_t^{\theta-1} Y_t + \theta_p \beta \left[ \mu MUC_{t+1} MC_{t+1} P_{t+1}^{\theta-1} Y_{t+1} \right] \\
&+ (\theta_p \beta)^2 \left[ \mu MUC_{t+2} MC_{t+2} P_{t+2}^{\theta-1} Y_{t+2} \right] + (\theta_p \beta)^3 \hat{G}_{t+3},
\end{aligned}$$

and

$$\begin{aligned}
\hat{H}_t &= MUC_t P_t^{\theta-1} Y_t + \theta_p \beta \left[ MUC_{t+1} P_{t+1}^{\theta-1} Y_{t+1} + \theta_p \beta \hat{H}_{t+2} \right] \\
&= MUC_t P_t^{\theta-1} Y_t + \theta_p \beta \left[ MUC_{t+1} P_{t+1}^{\theta-1} Y_{t+1} + \theta_p \beta \left( MUC_{t+2} P_{t+2}^{\theta-1} Y_{t+2} + \theta_p \beta \hat{H}_{t+3} \right) \right] \\
&= MUC_t P_t^{\theta-1} Y_t + \theta_p \beta \left[ MUC_{t+1} P_{t+1}^{\theta-1} Y_{t+1} \right] + (\theta_p \beta)^2 \left[ MUC_{t+2} P_{t+2}^{\theta-1} Y_{t+2} \right] + (\theta_p \beta)^3 \hat{H}_{t+3}.
\end{aligned}$$

Plugging these expressions into (14) and (15):

$$\begin{aligned}
G_t &= \frac{(G_t/H_t)^{1-\theta}}{MUC_t} \Big\{ \mu MUC_t MC_t P_t^{\theta-1} Y_t + \theta_p \beta \left[ \mu MUC_{t+1} MC_{t+1} P_{t+1}^{\theta-1} Y_{t+1} \right] \\
&+ (\theta_p \beta)^2 \left[ \mu MUC_{t+2} MC_{t+2} P_{t+2}^{\theta-1} Y_{t+2} \right] + (\theta_p \beta)^3 \hat{G}_{t+3} \Big\} \\
&= \mu MC_t (P_t(z)/P_t)^{1-\theta} Y_t + \mu \theta_p \beta \left[ (MUC_{t+1}/MUC_t) MC_{t+1} (P_t(z)/P_{t+1})^{1-\theta} Y_{t+1} \right] \\
&+ \mu (\theta_p \beta)^2 \left[ (MUC_{t+2}/MUC_t) MC_{t+2} (P_t(z)/P_{t+2})^{1-\theta} Y_{t+2} \right] + \frac{P_t(z)^{1-\theta}}{MUC_t} (\theta_p \beta)^3 \hat{G}_{t+3}. \qquad (33)
\end{aligned}$$

and

$$
\begin{aligned}
H_t &= \frac{(G_t/H_t)^{1-\theta}}{MUC_t}\Big\{MUC_t P_t^{\theta-1} Y_t + \theta_p\beta\left[MUC_{t+1}P_{t+1}^{\theta-1}Y_{t+1}\right] \\
&+ (\theta_p\beta)^2\left[MUC_{t+2}P_{t+2}^{\theta-1}Y_{t+2}\right] + (\theta_p\beta)^3\hat{H}_{t+3}\Big\} \\
&= MC_t(P_t(z)/P_t)^{1-\theta}Y_t + \theta_p\beta\left[(MUC_{t+1}/MUC_t)(P_t(z)/P_{t+1})^{1-\theta}Y_{t+1}\right] \\
&+ (\theta_p\beta)^2\left[(MUC_{t+2}/MUC_t)(P_t(z)/P_{t+2})^{1-\theta}Y_{t+2}\right] + \frac{P_t(z)^{1-\theta}}{MUC_t}(\theta_p\beta)^3\hat{H}_{t+3}. \qquad (34)
\end{aligned}
$$

Dividing (33) by (34), we get (13). Note also that:

$$
P_t(z) = \frac{G_t}{H_t} = \frac{\frac{(G_t/H_t)^{1-\theta}}{MUC_t}\hat{G}_t}{\frac{(G_t/H_t)^{1-\theta}}{MUC_t}\hat{H}_t} = \frac{\hat{G}_t}{\hat{H}_t},
$$

To obtain stationary variables under a positive money growth rate steady state regime start from equations (16)–(17) and divide $\hat{G}_t$ by $P_t^{\theta}$ and $\hat{H}_t$ by $P_t^{\theta-1}$ to get

$$
\begin{aligned}
\tilde{G}_t &\equiv \frac{\hat{G}_t}{P_t^{\theta}} = \mu MUC_t\frac{MC_t}{P_t}Y_t + \theta_p\beta\frac{\hat{G}_{t+1}}{P_t^{\theta}} \qquad (35)\\
&= \mu MUC_t\frac{MC_t}{P_t}Y_t + \theta_p\beta\frac{\hat{G}_{t+1}}{P_{t+1}^{\theta}}\frac{P_{t+1}^{\theta}}{P_t^{\theta}} = \mu MUC_t mc_t Y_t + \theta_p\beta\tilde{G}_{t+1}(1+\pi_{t+1})^{\theta} \qquad (36)\\
\tilde{H}_t &\equiv \frac{\hat{H}_t}{P_t^{\theta-1}} = MUC_tY_t + \theta_p\beta\frac{\hat{H}_{t+1}}{P_t^{\theta-1}} \qquad (37)\\
&= MUC_tY_t + \theta_p\beta\frac{\hat{H}_{t+1}}{P_{t+1}^{\theta-1}}\frac{P_{t+1}^{\theta-1}}{P_t^{\theta-1}} = MUC_tY_t + \theta_p\beta\tilde{H}_{t+1}(1+\pi_{t+1})^{\theta-1} \qquad (38)
\end{aligned}
$$

where $mc_t \equiv MC_t/P_t$ is the real marginal cost. Since

$$
\tilde{H}_t = \frac{\hat{H}_t}{P_t^{\theta-1}} = \frac{\hat{H}_tP_t}{P_t^{\theta}} \quad \Longrightarrow \quad \frac{\hat{H}_t}{P_t^{\theta}} = \frac{\tilde{H}_t}{P_t}
$$

and

$$
P_t(z) = \frac{G_t}{H_t} = \frac{\hat{G}_t}{\hat{H}_t} = \frac{\hat{G}_t/P_t^{\theta}}{\hat{H}_t/P_t^{\theta}} = \frac{\tilde{G}_tP_t}{\tilde{H}_t}.
$$

the law of motion for the price index

$$
P_t^{1-\theta} = \theta_pP_{t-1}^{1-\theta} + (1-\theta_p)P_t(z)^{1-\theta} = \theta_pP_{t-1}^{1-\theta} + (1-\theta_p)\left[\frac{\hat{G}_t}{\hat{H}_t}\right]^{1-\theta}
$$

can be divided by $P_t^{1-\theta}$ to obtain

$$
[(1+\pi_t)]^{1-\theta} = \theta_p + (1-\theta_p)\left[\frac{P_t(z)}{P_{t-1}}\right]^{1-\theta} = \theta_p + (1-\theta_p)\left[\frac{\hat{G}_t}{\hat{H}_tP_{t-1}}\right]^{1-\theta} = \theta_p + (1-\theta_p)\left[\frac{\tilde{G}_t}{\tilde{H}_t}(1+\pi_t)\right]^{1-\theta}.
$$

Table 23: Main term structure statistics when $b = 0$.

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| 3-month nominal yield | 6.98623 | 1.44943 | 0.30692 |
| 1-year nominal yield | 6.99047 | 1.34408 | 0.39468 |
| 10-year nominal yield | 6.99022 | 0.40513 | 0.46418 |
| 20-year nominal yield | 6.98662 | 0.23048 | 0.45700 |
| 30-year nominal yield | 6.98467 | 0.15837 | 0.45925 |

Table 24: Main term structure statistics when $\gamma = 1$.

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| 3-month nominal yield | 6.44966 | 2.83037 | 0.54434 |
| 1-year nominal yield | 6.82969 | 2.67349 | 0.60100 |
| 10-year nominal yield | 6.94103 | 0.81430 | 0.62670 |
| 20-year nominal yield | 6.93018 | 0.46307 | 0.61998 |
| 30-year nominal yield | 6.92388 | 0.31558 | 0.62013 |

## *B. Sensitivity Analysis*

In this Section, we briefly describe seven different experiments to illustrate the features of our model. We only concentrate on most dramatic differences with the benchmark model. Details are available on request.

*Test #1: No Habit Formation ($b = 0$)*

Table 23 shows the average yield curve when there is no habit formation present in the model. The term structure is flat, and hence it is no surprise that—as shown earlier in Table 5—habit formation is a necessary condition for the rejection of expectations hypothesis.

*Test #2: Logarithmic Preferences ($\gamma = 1$)*

With less curvature in the utility function, the the term structure is much flatter than in our benchmark case, as shown in Table 24.

*Test #3: No Interest Rate Smoothing ($\chi = 0$)*

When the monetary policy authority conducts no interest rate smoothing, the interest rates, not surprisingly, are much more volatile and steeper as presented in Table 25. Moreover, it is more difficult to reject expectations hypothesis as discussed in Section 6.

Table 25: Main term structure statistics when $\chi = 0$.

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| 3-month nominal yield | 1.75643 | 9.31439 | 0.54641 |
| 1-year nominal yield | 5.30754 | 8.63865 | 0.57931 |
| 10-year nominal yield | 6.20336 | 2.34314 | 0.59335 |
| 20-year nominal yield | 6.11159 | 1.33083 | 0.58849 |
| 30-year nominal yield | 6.05910 | 0.90156 | 0.58630 |

Table 26: Main term structure statistics when the steady inflation is zero.

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| 3-month nominal yield | 2.45515 | 4.44680 | 0.45301 |
| 1-year nominal yield | 3.47719 | 4.32154 | 0.55143 |
| 10-year nominal yield | 3.82458 | 1.37027 | 0.61767 |
| 20-year nominal yield | 3.80241 | 0.77660 | 0.60937 |
| 30-year nominal yield | 3.78915 | 0.52855 | 0.60917 |

*Test #4: Zero Steady State Inflation*

When the steady state inflation is zero, the nominal interest rates are lower, as presented in Table 26. However, this doesn't affect the model's ability to reject the expectations hypothesis as shown in Table 27.

*Test #5: Higher Steady State Inflation*

When the steady state inflation is twice as large, 6% per year, the nominal interest rates are much higher as presented in Table 28. However, this doesn't affect the model's ability to reject the expectations hypothesis as shown in Table 29.

*Test #6: Monetary Policy Responds to Output ($\omega_y = 0.2$)*

When the monetary policy authority responds, in addition to expected inflation, to output, inflation becomes countercyclical, its correlation with Hodrick-Prescott filtered output is now $-0.49965$. In addition, the inflation risk premium is now positive as shown in Table 30 and the responses of inflation and inflation risk premium to productivity and policy shocks are opposite to the benchmark as shown in Figures 7–10.

*Test #7: Less Aggressive Policy ($\omega_\pi = 1.2$)*

In addition to discussion in Section 6, Table 31 presents the main term structure statistics when $\omega_\pi = 1.2$. The term structure is now downward-sloping.

Table 27: The number of rejects in each regressions in the benchmark model for nominal term structure when the steady state inflation is zero.

| | | | | |
|---|---|---|---|---|
| $y_{t+1}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ |
| $x_t$ | $p^f_{1,t} - p^b_{1,t}$ | $p^f_{1,t} - p^b_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ |
| $Wald(a = b = 0)$ | 1000 | 52 | 1000 | 69 |
| $Wald(b = 0)$ | 982 | 68 | 111 | 75 |
| $Wald(b = -1)$ | 1000 | 1000 | 1000 | 1000 |

Table 28: Main term structure statistics when the steady inflation is zero.

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| 3-month nominal yield | 6.95375 | 7.11755 | 0.43617 |
| 1-year nominal yield | 9.00367 | 6.39357 | 0.50942 |
| 10-year nominal yield | 9.53720 | 1.89879 | 0.56305 |
| 20-year nominal yield | 9.46099 | 1.08623 | 0.54786 |
| 30-year nominal yield | 9.41742 | 0.73584 | 0.54986 |

Table 29: The number of rejects in each regressions in the benchmark model for nominal term structure when the steady state inflation is zero.

| | | | | |
|---|---|---|---|---|
| $y_{t+1}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t}$ | $p^b_{1,t+1} - p^f_{1,t} - rp_{1,t}$ |
| $x_t$ | $p^f_{1,t} - p^b_{1,t}$ | $p^f_{1,t} - p^b_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ | $p^b_{1,t} - p^f_{1,t}$ |
| $Wald(a = b = 0)$ | 1000 | 68 | 1000 | 86 |
| $Wald(b = 0)$ | 945 | 61 | 554 | 66 |
| $Wald(b = -1)$ | 1000 | 1000 | 1000 | 1000 |

Table 30: Main inflation risk premia statistics when $\omega_y = 0.2$.

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| Inflation risk premium ($n = 1$) | 0.01011 | 0.00264 | 0.35059 |
| Inflation risk premium ($n = 2$) | 0.02017 | 0.00543 | 0.34348 |
| Inflation risk premium ($n = 4$) | 0.03963 | 0.01039 | 0.32541 |
| Inflation risk premium ($n = 8$) | 0.07367 | 0.01727 | 0.32784 |
| Inflation risk premium ($n = 12$) | 0.09690 | 0.01878 | 0.28703 |
| Inflation risk premium ($n = 16$) | 0.10847 | 0.01702 | 0.23165 |
| Inflation risk premium ($n = 20$) | 0.11100 | 0.01517 | 0.20666 |
| Inflation risk premium ($n = 24$) | 0.10814 | 0.01295 | 0.15303 |
| Inflation risk premium ($n = 28$) | 0.10227 | 0.01136 | 0.07226 |
| Inflation risk premium ($n = 32$) | 0.09455 | 0.01063 | −0.02135 |
| Inflation risk premium ($n = 36$) | 0.08635 | 0.01094 | −0.08923 |
| Inflation risk premium ($n = 40$) | 0.07819 | 0.01108 | −0.15924 |
| Inflation risk premium ($n = 60$) | 0.04539 | 0.00676 | −0.28747 |
| Inflation risk premium ($n = 80$) | 0.02646 | 0.00455 | −0.20864 |

Table 31: Main term structure statistics when $\omega_\pi = 1.2$.

| | Mean | Standard Deviation | Correlation with Output |
|---|---|---|---|
| 3-month nominal yield | 6.44017 | 7.28137 | 0.46761 |
| 1-year nominal yield | 6.83520 | 7.01192 | 0.54884 |
| 10-year nominal yield | 6.08160 | 2.19782 | 0.60213 |
| 20-year nominal yield | 5.80191 | 1.29816 | 0.57477 |
| 30-year nominal yield | 5.64862 | 0.86674 | 0.58781 |

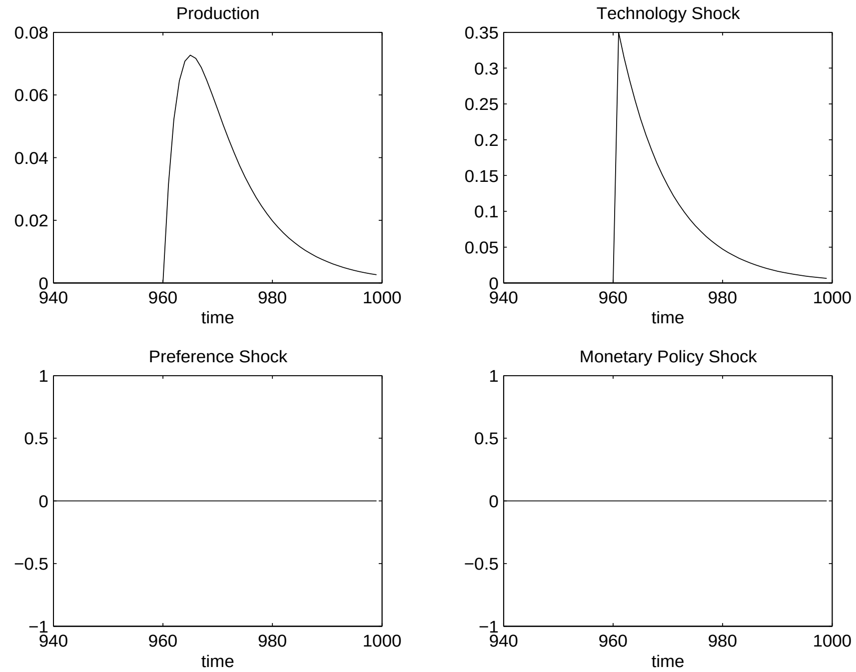


Figure 7: Impulse response functions to one standard deviation productivity shock in the benchmark parameterization when $\omega_y = 0.2$.

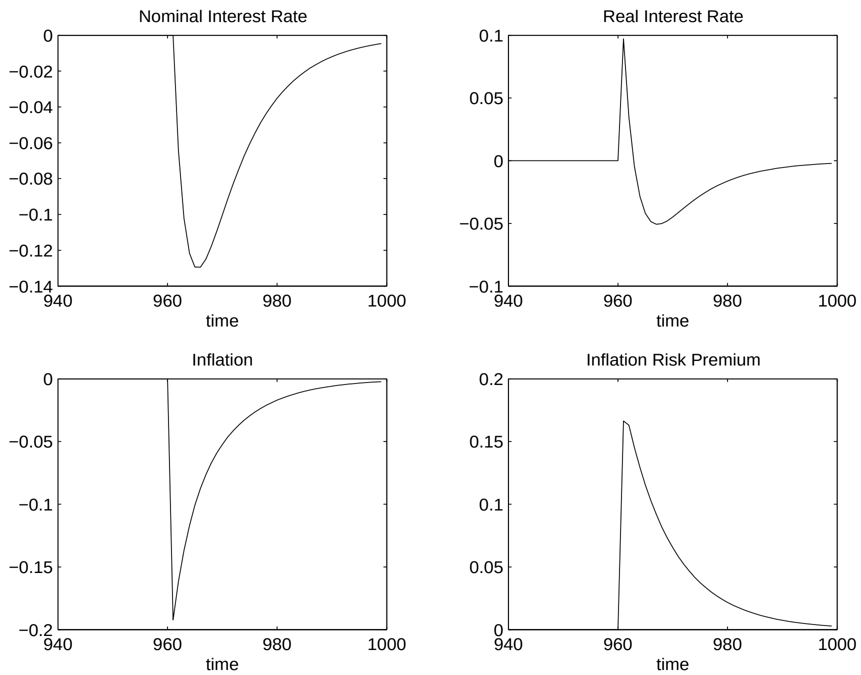


Figure 8: Impulse response functions to one standard deviation productivity shock in the benchmark parameterization when $\omega_y = 0.2$.

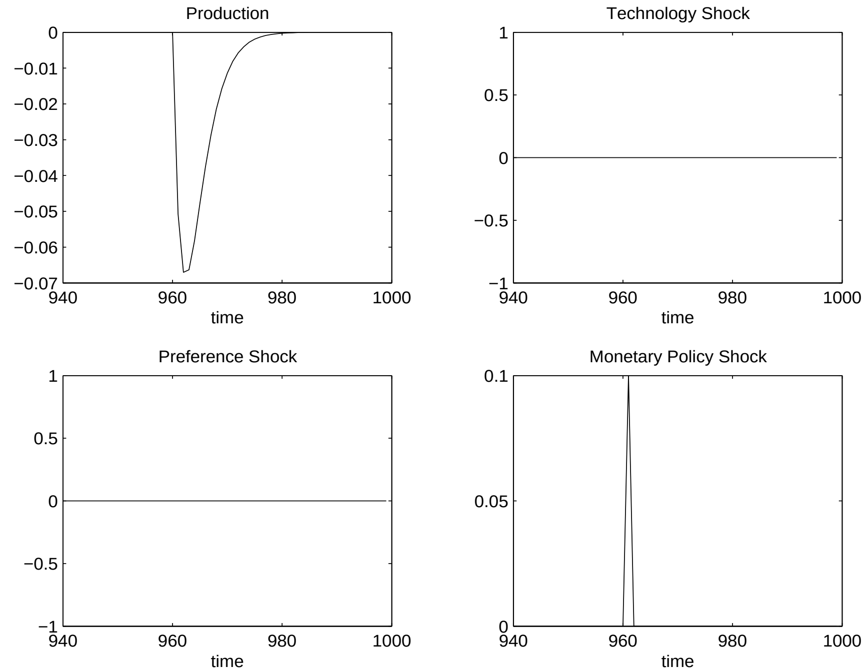


Figure 9: Impulse response functions to one standard deviation monetary policy shock in the benchmark parameterization when $\omega_y = 0.2$.

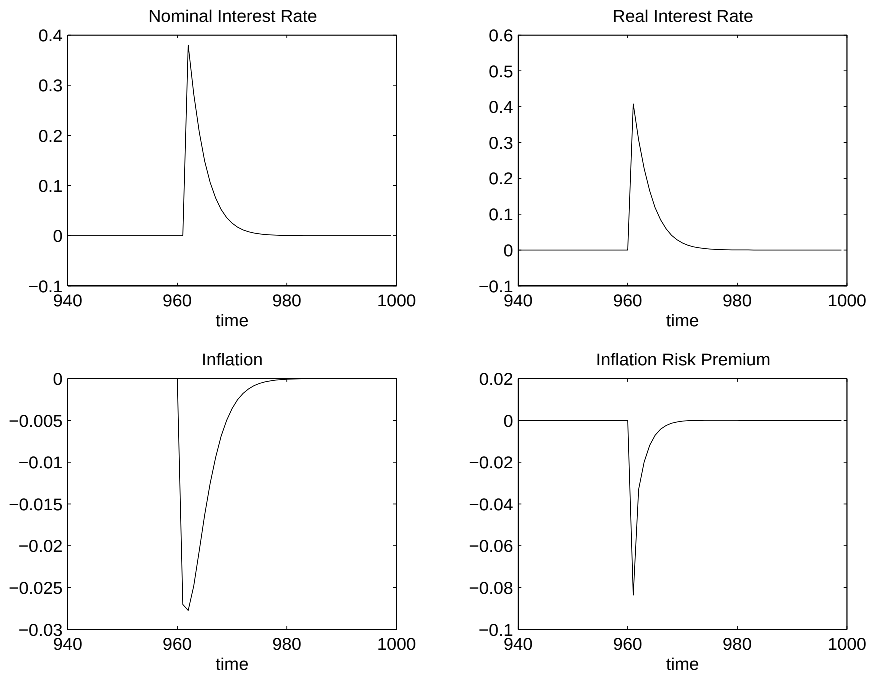


Figure 10: Impulse response functions to one standard deviation monetary policy shock in the benchmark parameterization when $\omega_y = 0.2$.

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