TERM PREMIA: ENDOGENOUS CONSTRAINTS ON MONETARY POLICY

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Abstract

Monetary policy evaluation using structural macro models suggests that historical monetary policy responds less aggressively to inflation and the output gap than would an optimal policy rule. However, these results are obtained using models with constant term premia. This paper shows how term premia may depend on the policy rule specification and policy rate uncertainty. A more aggressive policy rule involves an economically important increase in term premia. Consequently, conclusions about the specification of optimal monetary policy rules based on counterfactual simulations of models that exclude term premia effects may not be valid.

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Key words: optimal policy; term structure of interest rates; monetary policy transmission

1 Introduction

Interest in the analysis of monetary policy increased considerably in the 1990s with the development of a new variety of optimizing macro models. These models assume intertemporal optimizing behavior on the part of agents, sticky prices, no-arbitrage models of financial asset valuation, and interest rate feedback characterizations of monetary policy. Typically, alternative policy rules are evaluated in these models on the basis of what they imply for output and inflation variability.

A standard result in such counterfactual policy simulations is that historical monetary policy was not optimal and an outcome with lower inflation variability and lower output gap variability could have been achieved. Compared to empirical estimates of historical policy, a policy that responds more aggressively to output gaps or deviations of inflation from the policy goal would have been preferred. However, the benefits of more aggressive policy do not come without costs. In particular, although the variance of the output gap and inflation may be lower, the variance of the policy interest rate is higher. This paper suggests that increased policy rate variability may negatively impact economic activity through elevated term premia. Consequently, conclusions about the specification of optimal monetary policy rules based on counterfactual simulations of models that exclude term premia effects may not be valid.

Several theories have been proposed to explain why policymakers have historically set monetary policy in a way that imparts a less variable path of interest rates than theory suggests is optimal.[1] Brainard (1969) shows how uncertainty about model parameters can reduce the response of policy to economic disturbances. Data uncertainty should also reduce the responsiveness of policy to measures of economic activity (Orphanides (1998)). Rotemberg and Woodford (1997) argue that dependence of the policy rate setting on the lagged policy rate may increase the influence of a given policy action on longer term interest rates, and hence economic activity. While their discussion shows how such a policy rule may lessen excessive

[1] Clarida, Gali, and Gertler (1999) and Sack and Wieland (2000) review arguments for policy conservatism.

policy rate volatility, they don't theoretically justify a preference for limited policy rate variability. Goodfriend (1991) suggests that policymakers may prefer a smoother policy rate path because they fear that sharp changes in policy rates may disrupt financial markets.

Some researchers have modified their economic models to recognize the apparent real-world constraints on policy rate variability. In studies with parameter uncertainty, smoother policy is typically captured by reducing the response of the policy rate to output gaps and/or deviations of inflation from the policy target. In other studies, policy rate variability is assumed to be an exogenous constraint on monetary policy. One approach assumes the central bank minimizes output and inflation variability subject to an exogenously set bound on acceptable policy rate variability. Williams (1999) constrains policy by setting an upper bound on the variance of the policy rate, while Levin, Wieland, and Williams (1999) set an upper bound on the variance of the change in the policy rate. A second approach penalizes policy rate variability by specifying a central bank loss function that depends on policy rate variability in addition to output and inflation variability. Batini and Haldane (1999), for example, specify policy loss functions that assign a weight to policy rate variability in addition to squared deviations of inflation from the inflation target and of output from trend. Rudebusch and Svensson (1999), include the variance of the change in the policy rate in their specification of a policy loss function. In models where the variance of the change in the policy rate is included in the policy loss function, the tendency of policymakers to smooth policy rate changes is captured by including a lagged policy rate in the policy rate equation. This lagged interest rate is meant to capture that the policy rate is only partially adjusted to what a rule without lags would recommend.

This paper suggests an alternative motivation for constraints on policy rate variability. Term premia may depend on policy rate variability. The typical modeling approach specifies the the model's equations as log-linear approximations of the equilibrium conditions and assumes that term premia are constant (often zero). With this modeling approach, potential effects of increased policy rate variability on term premia will not be captured.

This paper examines how term premia may be related to monetary policy. A theoretical derivation shows how term premia depend on policy rate variability, economic uncertainty, and the treatment of apparent sluggishness in policy rate changes. Increased policy rate variability associated with more aggressive policy may raise term premia and longer-maturity market interest rates, and, thus, constrain economic activity. This potential effect of policy is excluded in standard simulations of alternative policy rules. Results suggest that the effect on term premia of more aggressive policy is nontrivial.

The next section discusses two alternative models of persistence in the policy rate, a partial adjustment model of policy and an AR error model of policy. Two models are considered since the data does not clearly support one model over the other, yet implications of changes in policy rate and economic uncertainty for term premia may depend on the specification chosen to represent policy. Section 3 reviews a model of the term structure, including an explicit expression for term premia. Section 4 calculates the 10-year term premium for estimates of model coefficients based on different historical policy episodes. Empirical results suggest that estimated term premia depend on how policy is modeled and that time variation in term premia has likely been sizable historically. In addition, by using results from published studies, the effects on term premia of more aggressive policy are shown to be economically important. The analysis raises serious questions about the applicability of optimal policy rule specifications based on models that exclude term-premium effects.

2 Modeling policy and economic activity

This section describes the macroeconomic structure that will be used in the model of the term structure of interest rates and term premia presented in section 3. The basic Taylor (1993a) rule representation of monetary policy is reviewed in the first subsection. The next two subsections present partial adjustment and AR-error generalizations of the Taylor rule that provide better fits of historical policy decisions. In the final subsection, a small reduced-form VAR model of macroeconomic activity is introduced. The specification nests both partial adjustment and AR-error models of policy.

2.1 The Taylor rule

Taylor (1993a) suggested that the target of the policy rate be set according to the deviation of inflation from the policy target for inflation and the deviation of output from potential. Thus, Taylor recommended a rule of the form:

$$r^* = \bar{\rho} + \pi + (\delta_\pi - 1)(\pi - \bar{\pi}) + \delta_y y \qquad (1)$$

where r^* is the Taylor-rule recommendation for the policy rate target, $\bar{\rho}$ is the equilibrium real policy rate, π is inflation, $\bar{\pi}$ is the policy target for inflation, and y is the output gap. Taylor used $\delta_y = 0.5$ as the weight on the output gap and $\delta_\pi = 1.5$ as the total weight on inflation.

Taylor rule specifications that are forward-looking with respect to inflation have been found to be empirically supported by Kozicki (1999) and Clarida, Gali, and Gertler (2000), among others. Table 1 summarizes estimation results over a collection of samples for regressions of the federal funds rate on a constant (assumed to include the equilibrium policy rate and the policy target), expected inflation, and the output gap:

$$r_t = c + \delta_\pi E_t \pi_{t,t+4} + \delta_y y_t + w_t. \qquad (2)$$

In these regressions, r_t is the federal funds rate, $E_t \pi_{t,t+4}$ is expected inflation over the four quarters from t to $t+4$ as measured using the median of expected inflation from the Survey of Professional Forecasters, y_t is the percent GDP gap based on estimates of potential output published by the Congressional Budget Office in August 2001, w_t is a possibly serially correlated regression residual, and $c = \bar{\rho} - 0.5\bar{\pi}$ is assumed to be constant.[2] Instrumental variables estimation uses four lags of each variable in the regression equation as well as contemporaneous expected inflation and four lags of

[2]The analysis uses CBO estimates of potential output that were published in August 2001. As these estimates did not incorporate the revisions to GDP data that were published in July 2001, GDP data, as available in the second quarter of 2001 was used to construct the output gap.

actual inflation.[3]

Empirical analysis supports weights close to those suggested by Taylor. As shown in the top panel of the table, labeled Estimated Taylor Rule, unmodeled serial correlation in residuals, estimates of the weight on inflation range from 1.27 to 2.02, but are insignificantly different from Taylor's suggested weight of 1.5. Estimates of the weight on the output gap range from 0.08 to 0.66 and are insignificantly different from Taylor's suggested weight of 0.5. The estimated rules fit historical policy better in the 1980s and 1990s than in the 1970s. The standard error of regression residuals is 1.59 percent in the 1970s, but 0.99 percent in the 1982-2000 sample and 0.78 percent in the 1987-2000 sample.

2.2 A Partial Adjustment Model of Policy

The Taylor (1993a) specification and versions with estimated coefficients yield policy recommendations that are at times similar to the target of the policy rate, but, as noted by Kozicki (1999), deviations tend to be persistent. The most common approach to improving the fit of policy rule recommendations to actual policy decisions, is to assume that policy only partially adjusts to the target in any one period:

$$r_t = \lambda r_{t-1} + (1-\lambda) r_t^* + \epsilon_{1,t} \tag{3}$$

where r_t is the one-period interest rate, and $\epsilon_{1,t}$ is a transitory, white noise policy shock. Clarida, Gali, and Gertler (2000) estimate similar rules.

Empirical results from estimation of a partial adjustment model are provided in the second panel of Table 1, labeled Partial Adjustment Model. A single equation is estimated:

$$r_t = \lambda r_{t-1} + (1-\lambda)(c + \delta_\pi E_t \pi_{t,t+4} + \delta_y \tilde{y}_t) + \epsilon_{1,t}, \tag{4}$$

corresponding to (3) with $r_t^* = c + \delta_\pi E_t \pi_{t,t+4} + \delta_y \tilde{y}_t$. Estimation uses nonlinear instrumental variables with the same set of instruments as was used for the first panel. Estimates of the responsiveness of policy to output gaps and inflation are

[3] Inflation and expected inflation are measured using the GDP deflator.

generally larger than in the first panel, and in several samples are significantly larger than Taylor's weights. Estimates of the persistence parameter λ range from 0.71 to 0.84, suggesting considerable persistence in the policy rate.

Woodford (1999), Levin, Wieland, and Williams (1999), and Sack (1998) are recent examples of papers that interpret estimates of λ close to one as implying that the monetary authority only gradually adjusts to its target policy rate, so that central banks smooth interest rate movements. Under this interpretation, estimates of λ on the order of 0.8 for quarterly data imply considerable monetary policy inertia, as in any quarter, the central bank only adjusts the policy rate by 20 percent of the change recommended by a Taylor-type policy rule.

2.3 An AR Error Model of Policy

Rudebusch (2002) questions the partial adjustment interpretation of empirical estimates of equations such as (4). He argues that such quarterly policy inertia would imply considerably more forecastable variation in interest rates at horizons of more than three months than is suggested from term structure evidence. Estimates of inertia may reflect misspecification of the Taylor rule if policymakers set policy according to variables not included in the Taylor rule (see the discussion in Kozicki (1999) and Rudebusch (2002)), or if variables in the Taylor rule are measured with error (Lansing (2002)). If omitted variables or measurement errors are serially correlated, then econometric estimates of λ in expressions such as (3) would overstate the degree of smoothing.[4]

These arguments suggest considering the following alternative description of policy decisions:

$$\begin{aligned} r_t &= r_t^* + \eta_t \\ \eta_t &= \rho\eta_{t-1} + \epsilon_{1,t} \end{aligned} \tag{5}$$

in which the deviation of the policy rate from the Taylor-rule recommendation is assumed to follow an AR(1) error process. This description of policy can be rewritten

[4]Real-time data concerns, for example, are likely to result in serially correlated differences between real-time and latest available estimates of the output gap.

as

$$r_t = \rho r_{t-1} + r_t^* - \rho r_{t-1}^* + \epsilon_{1,t} \tag{6}$$

where $\epsilon_{1,t}$ is assumed to serially uncorrelated.

Empirical results from estimation of:

$$r_t = \rho r_{t-1} + (1-\rho)c + \delta_\pi E_t \pi_{t,t+4} + \delta_y \tilde{y}_t - \rho \delta_\pi E_{t-1} \pi_{t-1,t+3} - \rho \delta_y \tilde{y}_{t-1} + \epsilon_{1,t} \tag{7}$$

are provided in the third panel of Table 1, labeled AR Error Model. Estimation is by nonlinear instrumental variables using the same instrument set as was used for the first two panels. Estimates of the responsiveness of policy to the output gap are insignificantly different from Taylor's weight of 0.5, however, estimates of the coefficient on expected inflation are generally smaller than 1.5, and significantly so in some cases.

The two policy specifications are nested by the following reduced form specification

$$r_t = a_1 r_{t-1} + a_2 r_t^* + a_3 r_{t-1}^* + \epsilon_{1,t}. \tag{8}$$

In the partial adjustment model, $a_1 + a_2 = 1$ and $a_3 = 0$, and in the AR error model, $a_1 - a_3 = 0$ and $a_2 = 1$. Rudebusch (2001) tested the restrictions of these two specifications to assess whether the data rejected one model but not the other. However, he found that results were fragile, depending on the sample used during estimation.

Another approach to assessing the interpretation of the partial adjustment model compares properties of estimated Taylor rule deviations to deviations of policy from estimated partial adjustment model recommendations after substituting out the lagged policy rate. After backward substitution, the partial adjustment model can be rewritten as:

$$\begin{aligned} r_t &= (1-\rho)\sum_{i=0}^{\infty} \rho^i r_{t-i}^* + \sum_{i=0}^{\infty} \rho^i \epsilon_{1,t} \\ &= \frac{(1-\rho)}{(1-\rho^n)} \sum_{i=0}^{n-1} \rho^i r_{t-i}^* + \upsilon_t + \sum_{i=0}^{\infty} \rho^i \epsilon_{1,t} \end{aligned} \tag{9}$$

where υ_t is an approximation error equal to the difference between the infinite sum in the first line above and the finite sum in the second. As the number of terms in the finite sum, n, increases, the approximation error converges to zero. If the partial adjustment interpretation is correct, then it seems reasonable to conclude that most of the variation in the policy rate should be due to the sum of lagged r^*. In addition, the deviations of the policy rate from the sum of lagged r^*s,

$$r_t - \frac{(1-\rho)}{(1-\rho^n)} \sum_{i=0}^{n-1} \rho^i r^*_{t-i}, \tag{10}$$

should be smaller than deviations of the policy rate from the estimated Taylor rule, $r_t - r^*_t$.

Table 2 contains root mean squared (RMS) deviations of the policy rule from the estimated Taylor rule (with no lags) and from the weighted sum of lagged r^*s. For the partial adjustment model, RMS deviations are reported for three choices of the number of terms in the approximating sum, one, eight, and sixteen. For most samples, RMS deviations are smallest for the estimated Taylor rule. The results question the validity of the partial adjustment model as capturing purposeful gradual adjustment of policy on behalf of policymakers.

2.4 A reduced form VAR model

To distinguish between the implications of the partial adjustment policy rule (3) from those of the AR-error policy rule (5) for the term structure of interest rates, it is sufficient to consider a two-variable AR model of the economy which follows a simple autoregression defined in terms of the policy rate, r_t, and the Taylor-rule recommendation, r^*_t. However, to do so requires a specification for the time-evolution of r^*_t. Assume r^*_t evolves according to:

$$r^*_t = \gamma_r r_{t-1} + \gamma_* r^*_{t-1} + (1 - \gamma_r - \gamma_*)(\bar{\rho} - \delta_\pi \bar{\pi}) + \epsilon_{2,t} \tag{11}$$

with $E_{t-1}[\epsilon_{1,t}\epsilon_{2,t}] = 0$. The Taylor rule recommendation is a function of the output gap and inflation, $r^*_t = (\bar{\rho} - \delta_\pi \bar{\pi}) + \delta_\pi E_t \pi_{t,t+4} + \delta_y \tilde{y}_t$ with $\delta_\pi \geq 0$ and $\delta_y \geq 0$, so it provides a summary of the state of the economy. Because estimates of the output gap and inflation are strongly persistent, γ_* is expected to be close to one. By contrast, γ_r is likely to be negative, reflecting that increases in the policy rate are expected to reduce both output and inflation.

Estimates of γ_r and γ_* are provided in Table 3. Four panels of results are included, with each corresponding to different models of policy, and, consequently, different estimates of the policy rule coefficients δ_π, δ_y, $\bar{\rho}$, and $\bar{\pi}$. Results in the top panel are based on r^* constructions that use Taylor weights of $\delta_y = 0.5$ and $\delta_\pi = 1.5$. Results in the second, third, and fourth panels use r^* constructions based on coefficient estimates as reported in the first, second, and third panels of Table 1, respectively. Estimates of γ_r and γ_* are as expected. Estimates of γ_r are generally negative and range from -0.01 to -0.24 for the 1982-2000 and 1987-2000 samples. Estimates of γ_* are positive and insignificantly different from one. Standard errors show sizable variation across the different samples suggesting considerable differences in economic uncertainty across time.

Both partial adjustment and AR-error policy specifications imply a vector AR(1) representation

$$z_t = H_1 z_{t-1} + H_2 \epsilon_t \tag{12}$$

where $z_t = [r_t \quad r_t^* \quad 1]'$, $\epsilon_t = [\epsilon_{1,t} \quad \epsilon_{2,t} \quad 0]'$, $E[\epsilon_{1,t}\epsilon_{2,s}] = 0$, and $E_{t-1}\epsilon_t\epsilon_t' = \Sigma = diag(\sigma_i^2)$ is a diagonal matrix with ith diagonal entry σ_i^2 and $\sigma_3^2 = 0$. For the partial adjustment specification

$$H_1 = \begin{bmatrix} \lambda + \gamma_r(1-\lambda) & \gamma_*(1-\lambda) & (1-\gamma_r-\gamma_*)(1-\lambda) \\ \gamma_r & \gamma_* & (1-\gamma_r-\gamma_*) \\ 0 & 0 & (\bar{\rho}-\delta_\pi\bar{\pi}) \end{bmatrix} \qquad H_2 = \begin{bmatrix} 1 & 1-\lambda & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 0 \end{bmatrix} \tag{13}$$

and for the AR-error specification

$$H_1 = \begin{bmatrix} \gamma_r + \rho & \gamma_* - \rho & (1-\gamma_r-\gamma_*) \\ \gamma_r & \gamma_* & (1-\gamma_r-\gamma_*) \\ 0 & 0 & (\bar{\rho}-\delta_\pi\bar{\pi}) \end{bmatrix} \qquad H_2 = \begin{bmatrix} 1 & 1 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 0 \end{bmatrix}. \tag{14}$$

The general structure in (12) will be used to represent the evolution of the state variables that are relevant for pricing nominal assets. The next section reviews the model of the term structure. The expression for the term premium will be shown to depend on the policy rule specification through the VAR coefficient matrices, H_1 and H_2, and on policy rate variability through the variance covariance matrix of economic shocks, Σ.

3 A model of the government term structure

The model of the nominal term structure is the same as in Kozicki and Tinsley (2001a and 2001b). The nominal price in period t of an n-period nominal bond which pays \$1 at $t+n$ is:

$$P_{n,t} = E_t[P_{n-1,t+1}M_{t+1}] \tag{15}$$

where M_{t+1} is a nominal stochastic discount factor. Recursive substitution yields the following alternative expression for $P_{n,t}$:

$$P_{n,t} = E_t(\Pi_{i=1}^{n} M_{t+i}) \tag{16}$$

Assuming joint log normality,

$$p_{n,t} = E_t(\sum_{i=1}^{n} m_{t+i}) + \frac{1}{2}Var_t(\sum_{i=1}^{n} m_{t+i}) \tag{17}$$

where $p_{n,t} = \log(P_{n,t})$ and $m_{t+i} = \log(M_{t+i})$. Since the continuous-time yield to maturity on the bond is defined as $r_{n,t} \equiv (-1/n)p_{n,t}$, it follows that

$$r_{n,t} = \frac{-1}{n}E_t(\sum_{i=1}^{n} m_{t+i}) - \frac{1}{2n}Var_t(\sum_{i=1}^{n} m_{t+i}) \tag{18}$$

so that the bond rate is determined by moments of the finite sum of the stochastic discount factor. The stochastic discount factor provides the link between the valuation of asset payoffs and states of the economy.

In conventional finance models, the stochastic discount factor is often specified to be a function of a short-term interest rate that evolves according to an AR process. For the purposes of macroeconomic analaysis, it would be more natural to assume the policy rate is generated by a macro VAR which includes the policy feedback rule, but the algebra for this can get tedious, vid. Kozicki and Tinsley (2001a). Here, we use the two-variable VAR outlined in section 2.4 to model evolution of the states, z_t.

The log stochastic discount factor is assumed to be a linear function of the states,

$$-m_{t+1} = a'z_t + \beta'\epsilon_{t+1} + v_{t+1} \tag{19}$$

where β, often referred to as the "price of risk," determines the covariance between innovations to the stochastic discount factor and innovations to the variables in z, v_{t+1}

is a shock with $E_t v_{t+1} = 0$, $E_t(v_{t+1}^2) = \sigma_v^2$, and $E_t(v_{t+i}\epsilon_{t+j}) = 0$ for $i, j > 0$. Although the derivation allows for the pricing of economic uncertainty, $E_t \epsilon_{2t}^2$, we leave analysis of this to future research as the focus of this article is on the relationship between policy uncertainty and term premia. Consequently, for the remainder of the paper, the assumption $\beta = [\beta_1,\ 0,\ 0]'$ is employed.

Conditional on the information in period t, the VAR specification in (12) implies that the future evolution of the state variable will take the form

$$z_{t+i} = H_1^i z_t + \sum_{j=1}^{i} H_1^{i-j} H_2 \epsilon_{t+j}. \tag{20}$$

Using (19) and (20), the evolution of the stochastic discount factor is

$$-m_{t+i} = a' H_1^{i-1} z_t + a' \sum_{j=1}^{i-1} H_1^{i-1-j} H_2 \epsilon_{t+j} + \beta' \epsilon_{t+i} + v_{t+i} \tag{21}$$

and an expression for the finite sum of the stochastic discount factor obtained by summing (21) over i is

$$\begin{aligned} -\sum_{i=1}^{n} m_{t+i} &= a'(I - H_1)^{-1}(I - H_1^n) z_t + \\ & \sum_{i=1}^{n} (\beta' + a'(I - H_1)^{-1}(I - H_1^{n-i}) H_2) \epsilon_{t+i} + \sum_{i=1}^{n} v_{t+i}. \end{aligned} \tag{22}$$

The coefficients in (21) can be determined from the one-period version of (18) which defines an expression for the short rate:

$$\begin{aligned} r_t &= -p_{1,t} \\ &= -[E_t(m_{t+1}) + \frac{1}{2} var_t(m_{t+1})] \\ &= a' z_t - \frac{1}{2}(\beta' \Sigma \beta + \sigma_v^2). \end{aligned} \tag{23}$$

Since z_t contains r_t, this expression is satisfied for $a = [1 \quad 0 \quad \frac{1}{2}(\beta' \Sigma \beta + \sigma_v^2)]'$.

An expression that relates a multi-period yield to expected short rates and a term premium can be derived by substituting for the mean and variance of the stochastic discount factor sum in (18). Using the solution for a, the mean of the stochastic discount factor sum (21) is

$$\begin{aligned}\frac{1}{n}E_t[\sum_{i=1}^{n}(-m_{t+i})] &= \frac{1}{n}[a'(I-H_1)^{-1}(I-H^n)z_t] \\ &= \frac{1}{2}(\beta'\Sigma\beta+\sigma_v^2)+\frac{1}{n}\sum_{i=0}^{n-1}E_t r_{t+1}, \qquad (24)\end{aligned}$$

and the variance of the stochastic discount factor sum is

$$\begin{aligned}-\frac{1}{2n}var_t(\sum_{i=1}^{n}m_{t+i}) &= -\frac{1}{2n}var[\sum_{i=1}^{n}(\beta'+a'(I-H_1)^{-1}(I-H_1^{n-i})H_2)\epsilon_{t+i}+\sum_{i=1}^{n}v_{t+i}] \\ &= -\frac{1}{2n}\sum_{i=1}^{n}(\beta'+a'(I-H_1)^{-1}(I-H_1^{n-i})H_2)\Sigma(\beta'+a'(I-H_1)^{-1}(I-H_1^{n-i})H_2)' \\ &\quad -\frac{1}{2}\sigma_v^2. \qquad (25)\end{aligned}$$

Substituting into (18), the n-period bond yield is the sum of expected short rates and a term premium, ϕ_n:

$$r_{n,t}=\frac{1}{n}\sum_{i=0}^{n-1}E_t r_{t+i}+\phi_n \qquad (26)$$

where, the term premium is equal to:

$$\begin{aligned}\phi_n &= \frac{-1}{2n}\sum_{i=1}^{n}[a'(I-H_1)^{-1}(I-H_1^{n-i})H_2\Sigma(a'(I-H_1)^{-1}(I-H_1^{n-i})H_2)'] \\ &\quad -\frac{1}{n}\sum_{i=1}^{n}[a'(I-H_1)^{-1}(I-H_1^{n-i})H_2\Sigma\beta]. \qquad (27)\end{aligned}$$

The term premium will be positive if β_1 is sufficiently negative.

The dependence of the term premium on monetary policy shows up in H_1 and H_2, the matrices which summarize the evolution of the economy including the policy rule specification, and in Σ, which summarizes economic and policy rate uncertainty. For both policy specifications, H_1 depends on the responsiveness of economic activity to the policy rate (captured by γ_r) and on the degree of persistence in economic activity (captured by γ_*). In addition, for the partial adjustment model, H_1 and H_2 depend on the persistence parameter λ that influences policy smoothness. More aggressive monetary policy may involve less policy persistence (λ closer to zero), or may reduce the persistence of shocks to economic activity (reduce γ_*) and thereby affect the term premium. In addition, increased policy rate variability that accompanies more aggressive monetary policy may affect the term premium through Σ.

4 Estimates of the 10-year Term Premium

This section starts with an examination of the effects on the size of the estimated 10-year term premium of shifts in estimated coefficients. Since results are based on empirical estimates using U.S. data, this analysis can only provide evidence on the size of model-based explanations of likely shifts in term premia historically. To provide a sense of the likely effects on term premia of increases in interest rate variability associated with a move to more aggressive "optimal" monetary policy rules, calculations must be based on results from counterfactual simulations. Statistics from published studies of such counterfactual simulations are used to assess the likely impact on term premia of more aggressive policy.

Key to these calculations is the link between policy rate uncertainty and Σ. The first subsection uses standard errors from estimated policy rules to estimate Σ. The second subsection explores the consequences of assumptions that relate Σ directly to policy rate variability.[5]

4.1 Theory-based estimates of historical variation in the 10-year term premium

The 10-year term premium ($n = 40$ quarters) is calculated using the expression in (27) for each of the five samples examined in the previous two sections: 1971:Q1-2000:Q4, 1971:Q1-1979:Q3, 1971:Q1-1982:Q3, 1982:Q4-2000:Q4, and 1987:Q4-2000:Q4. Parameters in H, and Σ are taken from the estimates provided in Tables 1 and 3. The standard errors of the estimated policy rule, provided in the final column of Table 1, are used to approximate policy-rate uncertainty—σ_1, the standard deviation of ϵ_{1t}.[6]

Estimates of the term premium also depend on β_1. In studies where the stochastic

[5] Tinsley (1999) discusses the consequences for term premia of reduced policy rate uncertainty and policies aimed at credible enforcement of upper or lower boundaries on segments of the term structure.

[6] The standard errors of the reduced form r^* equation, provided in the final column of Table 3, could be used to approximate economic uncertainty—σ_2, the standard deviation of ϵ_{2t}. However, as the price associated with this source of risk is restricted to equal zero, the value chosen for σ_2 is irrelevant for cacluations of the term premium.

discount factor is restricted to be a function of a single variable equal to the one-period rate, estimates of β_1 have been obtained to fit the average yield spread between long and short maturity yields, given estimates of coeffiecients in H_1, H_2, and Σ. Campbell, Lo, and MacKinlay (1997) estimate β_1 to be -122. Kozicki and Tinsley (2001b) provide several estimates of β_1, that differ across alternative time-series specifications for the short rate. Their estimates range from -148 to -383. Since it is difficult to identify β_1 without a more complete model, a large range of values of β_1 are considered.

Term premium estimates are provided in Table 4 for $n = 40$ quarters. Estimates of the term premium for the partial adjustment model of policy are provided in the top panel with those for the AR error model of policy in the bottom panel. A separate column is provided for each of the five samples.

Several results are evident in every sample for constant β_1. First, estimates of the term premium differ considerably for the two models of policy. Estimates are much larger for the AR error model than for the partial adjustment model. For instance, for $\beta_1 = -150$, over 1971-1979, the AR error model implies a 40-quarter term premium of 1.18 percent, compared to only 0.03 percent for the other model. Over 1982-2000, the difference is smaller, but still sizable, with term premium estimates of 0.42 for the AR error model and 0.17 for the partial adjustment model.

Second, term premia are unlikely to be constant over time. For instance, for $\beta_1 = -150$, estimates of the term premium vary between 0.04 in 1987-2000, to 1.18 in 1971-1979, and to 2.61 in 1971-1982 for the AR error model of policy. In the same periods, the partial adjustment model implies term premium estimates of -0.08, 0.11, and 0.51 for $\beta_1 = -350$. These differences are economically important. Of course, it is possible that the term premium is actually constant over time, but that β_1 varies over time in such a way as to negate the implications of shifts in Σ or other parameters. However, while time variation in β_1 is possible, variation constrained to exactly counteract the implications of shifts in other parameters is unlikely.

Third, to obtain the same magnitude estimate of the term premium, the partial adjustment model requires a much more negative value of β_1. For instance, over the 1971 through 2000 sample, estimates of the term premium for the AR error model are roughly twice as large as those for the partial adjustment model for any give value of β_1. In other words, for this sample, the partial adjustment model would require a price of risk that is twice as large in magnitude to generate the same size term premium.

Overall, these results suggest two important conclusions: estimates of term premia depend on how policy is modeled and historical variation in term premia is sufficiently large to be economically significant. An implication of the conclusions is that log-linear models that assume constant term premia are missing a feature of the economy that is relevant for policy analysis. Consequently, conclusions about the specification of optimal policy rules based on counterfactual simulations of these models are subject to the Lucas Critique.

The next section explores the implications for term premia of increases in policy rate variability associated with "optimal" policy that is more aggressive than empirical estimates of historical policy.

4.2 Estimates of effects on term premia of more aggressive policy

Several recent studies have suggested that historical U.S. monetary policy has not been optimal. In particular, these studies argue that a policy that responded more aggressively to output gaps and/or deviations of inflation from target would have achieved a better macroeconomic outcome—better in the sense of lower inflation or output gap variability. However, more aggressive policy also tends to lead to higher policy rate variability. This section provides estimates of the effect on term premia of higher policy rate variability.

Table 5 provides examples from three recent studies of the effect of more aggressive monetary policy on the standard devation of the policy interest rate and the standard deviation of the change in the policy rate. Using their own model (RS), Rudebusch and Svensson (1999) found that an increase in the weight on the output gap in a Taylor rule from 0.5 to 1.0 increased the standard deviation of the change in the policy rate from 0.71 to 1.03, a factor of 1.45. The optimal policy for their specified policy loss function was much more aggressive than either of these policy rule specifications and involved weights of 2.72 on inflation and 1.57 on the output gap. This policy resulted in more than a doubling of the standard deviation of the change in the policy rate compared to the original Taylor rule. Effects on the standard deviation of the policy rate were much smaller.

Levin, Wieland, and Williams (1999) also calculated the standard deviation of the policy rate and the standard deviation of the change in the policy rate for the original Taylor rule and for a rule with a unit weight on the output gap. They provided results for four different structural macroeconomic models: the Fuhrer-Moore (1995) model (FM), the Monetary Studies Research model of Orphanides and Wieland (1998) (MSR), the Federal Reserve Board staff model (cf. Brayton, Levin, et. al. 1997) (FRB), and Taylor's (1993b) multi-country model (TMCM). For all models, an increase in the responsiveness of monetary policy to the output gap increased the standard deviation of the policy rate and the standard deviation of the change in the policy rate. The effect on the standard deviation of the change in the policy rate was larger, with factors ranging from 0.90/0.75=1.20 for the Fuhrer-Moore model, to 0.50/0.30=1.67 for the MSR model. Effects on the standard deviation of the policy rate were somewhat smaller and ranged from 3.83/3.57=1.07 to 2.51/3.16=1.26, respectively, for these two models.

Simulation results from Williams (1999) find large effects on the standard deviation on the change in the interest rate when both weights in a Taylor rule are increased. Williams used two versions of the Federal Reserve Board staff model, FRB-RE which assumes rational expectations and FRB-VAR which uses VAR approximations to expectations. While effects on the standard deviation of the change in the interest rate were similar to those obtained by Rudebusch and Svensson (1999), effects on the standard deviation of the policy rate were larger.

Increases in policy rate variability will have implications for term premia if the conditional variance of the policy rate, σ_1^2, is related to unconditional policy rate variability, $SD(r)$, or to unconditional variability of the change in the policy rate, $SD(\Delta r)$. Either one of these is likely to be a reasonable assumption. Chan, Karolyi, Longstaff, and Sanders (1992) find that most models that successfully capture dynamics of the short term interest rate allow the volatility of interest rate changes to be a function of the level of the interest rate.[7] And, the change in the policy rate and the policy shock $\epsilon_{1,t}$ are highly correlated over 1971-2000. The correlation coefficient is 0.86 for the partial adjustment model of policy and 0.83 for the AR error model of policy. Table 6 provides additional evidence that suggests that the variance of $\epsilon_{1,t}$ is likely related to the level of the policy rate or to policy variability (measured as either the variance of the policy rate or the variance of the change in the policy rate). The table contains correlations between approximations to the conditional variance of the policy shock and properties of policy. The table provides two approximations to the conditional variance of the policy shock. One measure is the square of the policy shock (squared residuals). The second is an 8-quarter moving average of the square of the policy shock, i.e., the policy shock variance calculated over the current and prior seven quarters (MA(8) variance). These variance approximations are provided for policy shocks estimated as residuals from the Taylor rule, residuals from the estimated Taylor rule, residuals from the estimated partial adjustment model of policy, and innovations from the AR-error model of policy. Four policy summary statistics are included. The first is the level of the policy rate (Level), the second is the variance of the policy rate calculated over the current and prior seven quarters (MA(8) variance of level), the third is the squared change in the policy rate (Squared change $(\Delta r)^2$), and the fourth is the 8-quarter moving average of the square of the policy rate change (MA(8) variance of change).

Correlations between both approximations to the conditional variance and the four policy variables are high.[8] MA(8) approximations to the conditional variances

[7]It is commonplace in the finance literature to assume that the conditional variance of a short-term interest rate is a function of the level of the interest rate. See, for example, the square-root model of Cox, Ingersoll, and Ross (1985), the geometric Brownian motion model of Black and Scholes (1973) and related specifications by Dothan (1978) and Brennan and Schwartz (1980), the variable rate model of Cox, Ingersoll, and Ross (1980), and the constant elasticity of variance process of Cox and Ross (1976).

[8]Note: Results are robust to the following alternative definitions: replace squared residuals with absolute residuals, MA(8) variance with MA(8) standard deviation, MA(8) variance of the funds rate level with MA(8) standard deviation of the funds rate level, and MA(8) variance of the funds rate change with MA(8) standard deviation of the funds rate change.

of the policy shocks from the partial adjustment model and AR error model are on the order of 0.86 for the MA(8) variance of the level of the policy rate and 0.99 for the MA(8) variance of the change in the policy rate. Correlations between variance approximations for the Taylor rule and estimated Taylor rule and the MA(8) variances of the level and change of the policy rate are somewhat smaller, but still exceed 0.55. These results suggest that it is reasonable to map changes in policy variability that accompany more aggressive policy into changes in the conditional variance of policy, Σ, that appears in term premia expressions.

The results of Rudebusch and Svensson, suggest that the standard deviation of the change in the policy rate increases by a factor of between 1.69 and 2.45 when a rule that resembles historical policy is replaced in counterfactual simulations with an "optimal rule." This implies that the variance increases by a factor of between 2.8 and 6.0. Given the apparent proportional relationship between the variance of the change in the policy rate and Σ, this implies that, all else equal, term premia will be higher by a factor of betwen 2.8 and 6.0. This is a huge effect. Over the 1971-2000 period, the average spread between the 10-year Treasury yield and the federal funds rate (an approximation to the average term premium on the 10-year Treasury yield) was 0.82 percent. Multiplying this by the higher variance factors implies an increase in the term premium of between 1.48 percentage points and 4.1 percentage points.[9]

While the results in this section are only meant to be suggestive, they do provide strong evidence to suggest that term premia may respond to changes in monetary policy regimes. Consequently, findings that optimal policy is more aggressive than estimates of historical policy, may not hold up once the models are generalized to include a link between increased policy rate variability and increased term premia.

[9]Because these estimates hold everything else constant, they don't take account of other effects of more aggressive policy that may reduce the term premium. For example, in the VAR of section 2.4, more aggressive policy may influence the responsiveness persistence of non-policy shocks (by affecting γ_* in the VAR) or the responsiveness of economic activity to policy (by affecting γ_r in the VAR). Thus, the estimates of the effects on term premia likely provide an upper bound to the true effect.

5 Concluding Comments

Structural macro models have been used to evaluate alternative monetary policies and identify the "optimal" policy that would minimize a specified loss function. By using structural specifications, researchers hope their results are not subject to the Lucas critique. Unfortunately, the real world applicability of results from such counterfactual simulations is questionable. By not modeling term premia, the structural specifications used for policy evaluation are missing an important link between policy and economic activity.

This paper shows that estimates of term premia depend on how policy is modeled and that historical variation in term premia has likely been considerable. Furthermore, moving from a representative estimated historical policy rule to a typical optimal rule with more aggressive policy responses involves an economically important increase in term premia. Consequently, until structural models used to evaluate monetary policy are generalized to include potential links between policy and term premia, it would be premature to assume more aggressive monetary policy would be preferable when making real world policy decisions.

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Table 1: Estimates of the Policy Rule

Equation Number	Estimation Sample	Coeff on $E_t\pi_{t,t+4}$ δ_1	Coeff on $\tilde{y}_t$ δ_2	Coeff on r_{t-1} λ	Coeff on η_{t-1} ρ	Standard Error
Estimated Taylor Rule, unmodeled serial correlation in residuals						
2	1971:Q1 - 2000:Q4	1.32	0.13			2.13
		0.23	*0.19*			
	1971:Q1 - 1979:Q3	1.27	0.62			1.59
		0.19	*0.12*			
	1971:Q1 - 1982:Q3	1.79	0.08			2.88
		0.25	*0.36*			
	1982:Q4 - 2000:Q4	2.02	0.43			0.99
		0.26	*0.15*			
	1987:Q4 - 2000:Q4	1.64	0.66			0.78
		0.16	*0.12*			
Partial Adjustment Model						
4	1971:Q1 - 2000:Q4	1.87	1.22	0.84		1.01
		0.34	*0.51*	*0.05*		
	1971:Q1 - 1979:Q3	1.15	1.64	0.78		0.88
		0.35	*0.40*	*0.06*		
	1971:Q1 - 1982:Q3	2.58	1.66	0.84		1.52
		0.52	*1.01*	*0.07*		
	1982:Q4 - 2000:Q4	2.21	0.73	0.74		0.43
		0.33	*0.19*	*0.06*		
	1987:Q4 - 2000:Q4	1.72	0.92	0.71		0.30
		0.17	*0.13*	*0.05*		
AR Error Model						
7	1971:Q1 - 2000:Q4	0.90	0.55		0.91	1.04
		0.24	*0.22*		*0.05*	
	1971:Q1 - 1979:Q3	0.67	0.41		0.80	1.10
		0.47	*0.26*		*0.18*	
	1971:Q1 - 1982:Q3	0.75	0.32		0.90	1.61
		0.50	*0.36*		*0.10*	
	1982:Q4 - 2000:Q4	0.75	0.69		0.89	0.48
		0.21	*0.17*		*0.03*	
	1987:Q4 - 2000:Q4	1.15	0.49		0.90	0.39
		0.26	*0.17*		*0.07*	

Heteroskedasticity and serial correlation robust standard errors, calculated according to Newey and West (1987) with 6 lags, are italicized under the coefficient estimates.

Table 2: Fit of Backward Substituted Representations of the Partial Adjustment Model

Estimation Sample	Estimated Taylor Rule (no lags)	Partial Adjustment: Terms in Back-Substitution		
		1	8	16
1971:Q1 - 2000:Q4	2.13	3.17	2.39	2.27
1971:Q1 - 1979:Q3	1.59	3.02	1.49	1.66
1971:Q1 - 1982:Q3	2.88	4.72	2.44	2.41
1982:Q4 - 2000:Q4	0.99	1.15	1.58	1.53
1987:Q4 - 2000:Q4	0.78	0.90	1.32	1.36

Entries are square root of mean squared deviations of back-substituted approximations to the partial adjustment model from the funds rate.

Table 3: Estimates of coefficients in reduced form r^* equation

Estimation Sample	Coeff on r_{t-1} γ_r	Coeff on r^*_{t-1} γ_*	Standard Error
r^* coefficients from Taylor (1993)			
1971:Q1 - 2000:Q4	-0.09	1.05	0.67
	0.04	*0.04*	
1971:Q1 - 1979:Q3	0.09	0.85	0.87
	0.05	*0.11*	
1971:Q1 - 1982:Q3	-0.11	1.06	0.88
	0.05	*0.08*	
1982:Q4 - 2000:Q4	-0.01	0.95	0.51
	0.06	*0.09*	
1987:Q4 - 2000:Q4	-0.11	1.07	0.41
	0.06	*0.09*	
r^* coefficients from estimated Taylor Rule			
1971:Q1 - 2000:Q4	-0.04	1.02	0.50
	0.04	*0.05*	
1971:Q1 - 1979:Q3	-0.01	0.94	0.89
	0.05	*0.08*	
1971:Q1 - 1982:Q3	0.00	0.93	0.88
	0.09	*0.11*	
1982:Q4 - 2000:Q4	-0.02	0.95	0.61
	0.07	*0.08*	
1987:Q4 - 2000:Q4	-0.17	1.10	0.48
	0.08	*0.09*	
r^* coefficients from Partial Adjustment Model			
1971:Q1 - 2000:Q4	-0.16	1.03	1.17
	0.05	*0.05*	
1971:Q1 - 1979:Q3	-0.37	1.04	1.66
	0.11	*0.05*	
1971:Q1 - 1982:Q3	-0.30	1.07	2.06
	0.05	*0.04*	
1982:Q4 - 2000:Q4	-0.02	0.95	0.74
	0.09	*0.09*	
1987:Q4 - 2000:Q4	-0.24	1.13	0.58
	0.10	*0.09*	
r^* coefficients from AR Error Model			
1971:Q1 - 2000:Q4	-0.07	1.03	0.54
	0.02	*0.05*	
1971:Q1 - 1979:Q3	-0.05	0.99	0.53
	0.03	*0.07*	
1971:Q1 - 1982:Q3	-0.07	1.08	0.48
	0.02	*0.06*	
1982:Q4 - 2000:Q4	-0.01	0.92	0.44
	0.04	*0.05*	
1987:Q4 - 2000:Q4	-0.13	1.11	0.35
	0.06	*0.09*	

Heteroskedasticity and serial correlation robust standard errors, calculated according to Newey and West (1987) using 6 lags, are italicized under the coefficient estimates.

Table 4: Estimates of the 10-year Term Premium

β_1	1971:Q1 to 2000:Q4	1971:Q1 to 1979:Q3	1971:Q1 to 1982:Q3	1982:Q4 to 2000:Q4	1987:Q4 to 2000:Q4
Partial Adjustment Model					
-10	-0.01	-0.03	-0.10	-0.02	-0.02
-50	0.19	-0.01	-0.03	0.04	-0.03
-100	0.44	0.01	0.06	0.10	-0.04
-150	0.69	0.03	0.15	0.17	-0.05
-200	0.94	0.05	0.24	0.23	-0.05
-250	1.19	0.07	0.33	0.30	-0.06
-300	1.43	0.09	0.42	0.37	-0.07
-350	1.68	0.11	0.51	0.43	-0.08
AR Error Model					
-10	0.05	0.05	0.11	0.01	-0.01
-50	0.43	0.37	0.82	0.13	0.00
-100	0.90	0.78	1.71	0.27	0.02
-150	1.38	1.18	2.61	0.42	0.04
-200	1.85	1.59	3.50	0.57	0.06
-250	2.33	1.99	4.39	0.72	0.08
-300	2.80	2.40	5.28	0.86	0.10
-350	3.28	2.80	6.17	1.01	0.12

Table 5: Variance Tradeoffs

Study (Model)	δ_π	δ_y	SD(r)	SD(Δr)
Rudebusch and Svensson (1999)				
(RS)	1.50	0.50	4.94	0.71
(RS)	1.50	1.00	4.97	1.03
(RS)	2.72	1.57	5.11	1.74
Levin, Wieland, and Williams (1999)				
(FM)	1.50	0.50	3.57	0.75
	1.50	1.00	3.83	0.90
(MSR)	1.50	0.50	1.01	0.30
	1.50	1.00	1.19	0.50
(FRB)	1.50	0.50	2.51	0.90
	1.50	1.00	3.16	1.20
(TMCM)	1.50	0.50	4.00	1.58
	1.50	1.00	4.35	2.41
Williams (1999)				
(FRB-RE)	1.50	0.50	2.51	0.90
	2.00	2.00	4.32	2.00
(FRB-VAR)	1.50	0.50	2.07	0.94
	2.00	2.00	2.52	2.07

Table 6: Explaining policy rate uncertainty

Variance Approximation	Policy rate measure Level	MA(8) variance of level	Squared Change $(\Delta r)^2$	MA(8) variance of change
Taylor rule				
Squared residuals	0.62	0.82	0.14	0.77
MA(8) variance	0.41	0.65	0.09	0.58
Estimated Taylor rule				
Squared residuals	0.37	0.51	0.08	0.61
MA(8) variance	0.29	0.59	0.11	0.55
Partial Adjustment Model				
Squared residuals	0.44	0.32	0.94	0.52
MA(8) variance	0.67	0.86	0.47	0.99
AR Error Model				
Squared residuals	0.46	0.34	0.91	0.53
MA(8) variance	0.67	0.87	0.47	0.99