Term Structure Evidence on Interest Rate Smoothing and Monetary Policy Inertia

Glenn D. Rudebusch*
Federal Reserve Bank of San Francisco

September 2000

Abstract

Numerous studies have used quarterly data to estimate monetary policy rules or reaction functions that appear to exhibit a very slow partial adjustment of the policy interest rate. The conventional wisdom is that this gradual adjustment reflects a policy inertia or interest rate smoothing behavior by central banks. However, such quarterly monetary policy inertia would imply a large amount of forecastable variation in interest rates at horizons of more than three months, which is contradicted by evidence from the term structure of interest rates. The illusion of monetary policy inertia may reflect the episodic, unforecastably persistent shocks that central banks face.

*Homepage: www.sf.frb.org/econrsrch/economists/grudebusch.html; email: Glenn.Rudebusch@sf.frb.org; address: 101 Market Street, San Francisco, CA, 94105. For helpful comments, I thank Kevin Lansing, Tim Cogley, and participants in the 2000 CEPR Summer Institute in Amsterdam, especially my discussant, Peter van Els. Niloofar Badie, Kirran Bari, Casey Cornwell, and Carolina Marquez provided excellent research assistance. The views expressed in this paper do not necessarily reflect those of anyone else in the Federal Reserve System.

1. Introduction

How quickly do central banks adjust monetary policy in response to developments in the economy? A common view among economists is that the short-term policy interest rate in many countries is changed at a very sluggish pace over several quarters. The evidence supporting this view is found in the many monetary policy rules or reaction functions estimated in the literature with quarterly data. These policy rules take the general partial adjustment form $i_t = (1-\rho)i_t^* + \rho i_{t-1}$, where i_t is the average level of the policy interest rate in quarter t, which is set as a weighted average of the current desired level, i_t^*, and last quarter's actual value, i_{t-1}. Based on historical data, estimates of ρ are often in the range of 0.8, so these empirical rules appear to imply a very slow speed of adjustment of the policy rate to its fundamental determinants. This gradual adjustment of the policy rate over several quarters to its desired level is widely interpreted as evidence of an "interest rate smoothing" or "monetary policy inertia" behavior by central banks. For example, Clarida, Gali, and Gertler (2000, pp. 157-158) describe their U.S. estimates of various partial adjustment policy rules as ". . . the estimate of the smoothing parameter ρ is high in all cases, suggesting considerable interest rate inertia: only between 10 and 30 percent of a change in the [desired interest rate] is reflected in the Funds rate within the quarter of the change. Thus, our estimates confirm the conventional wisdom that the Federal Reserve smooths adjustments in the interest rate." A few of the many other recent papers with a similar inertial interpretation of monetary policy rules include Woodford (1999), Goodhart (1999), Levin, Wieland, and Williams (1999), Amato and Laubach (1999), Orphanides (1998), and Sack (1998).

Furthermore, some researchers have also argued recently that the monetary policy inertia apparently present in the real world may be an optimal behavioral response on the part of central banks. For example, one popular such normative argument contends that the quarterly policy inertia and interest rate smoothing behavior helps the central bank focus the expectations of agents in the economy on its stabilization goals and thereby achieve a better outcome (e.g., Levin, Wieland, and Williams, 1999, Woodford, 1999, and Sack and Wieland, 2000).

There is another quite separate literature on "interest rate smoothing," which, at least superficially, may appear to be consistent with the quarterly interest rate smoothing described above. This earlier literature analyzes changes in policy interest rates on a day-by-day basis. Both in the U.S. (e.g., Goodfriend, 1991, and Rudebusch, 1995) and in Europe, Japan, and Australia (e.g., Goodhart, 1997, and Lowe and Ellis, 1997), central banks appear to follow a pattern of behavior in which changes in the policy rate are undertaken at discrete intervals and in discrete amounts.[1] For example, Rudebusch (1995, p. 264) defines short-term (or weekly) interest rate smoothing as the Fed adjusting interest rates ". . . in limited amounts . . . over the course of several weeks with gradual increases or decreases (but not both) . . . ."[2]

Many have assumed—including the monetary policy rule papers cited above—that such short-term interest rate smoothing implies the quarterly interest rate smoothing found in the empirical policy rules. However, the earlier short-term interest rate smoothing refers to a partial adjustment over the course of several weeks, while quarterly interest rate smoothing refers to a partial adjustment over the course of several quarters. With such disparate time frames, the two types of partial adjustment are in fact largely independent, so a central bank could conduct either type of smoothing without much of the other. Indeed, a crucial point of the short-term interest rate smoothing literature is that although central banks smooth interest rates on a week-to-week or month-to-month basis, there is essentially no quarterly interest rate smoothing. This description follows Mankiw and Miron (1986, p. 225), who note that the postwar term structure of interest rates suggests that at a quarterly frequency ". . .the Fed set the short rate at a level that it expected to maintain. Under this characterization of policy, while the Fed might change the short rate in response to new information, it always (rationally) expected to maintain the short rate at its current level." Goodfriend (1991, p. 10) provides an identical random-walk assessment of the policy rate and argues that changes in the rate set by the Fed ". . . are essentially unpredictable at forecast horizons longer than a month or two." Similarly, Rudebusch (1995, p. 264) characterizes the Fed's behavior as, ". . . beyond a horizon of about a month, there are no planned movements to react to information already known." Thus, the earlier short-term interest rate smoothing literature rejects any partial adjustment or policy inertia at a quarterly frequency.[3]

This paper argues that quarterly interest rate smoothing and monetary policy inertia is a

[1] Also, see Balduzzi, Bertola, and Foresi (1997), Dotsey and Otrok (1995), and Eijffinger, Schaling, and Verhagen (1999).

[2] The short-term interest rate smoothing literature distinguishes three interest rates: the market rate at which funds are actually traded, i_t^m; the "target" rate that the central bank enforces in the market on a week-by-week basis, i_t; and the desired rate, i_t^*, that the central bank would set as its target if unconstrained by a desire to adjust rates slowly. Note that the target rate is not the desired rate. Furthermore, although the market and target rates, which are the ones reported in the popular press, can differ substantially on any given day, they are largely indistinguishable on a monthly average basis as the central bank hits its target, so both are denoted as i_t in this paper (which considers quarterly average data). As examples, Rudebusch (1995) explicitly models i_t^m and i_t on a daily basis (with i_t^* implicit), while Dotsey and Otrok (1995) model i_t^* and i_t on a monthly average basis (so $i_t^m = i_t$).

[3] This can be shown formally by simulating the models of short-term interest rate smoothing and fitting a partial adjustment model to the quarterly averaged simulated data. For example, data on the desired and actual funds rate (i_t^* and i_t) can be generated according to equations (5) and (6) in Dotsey and Otrok (1995), and after quarterly averaging, the $\hat{\rho}$ in an estimated partial adjustment policy rule is only about 0.1.

very modest phenomenon in practice, which is in accord with the earlier characterization of monetary policy partial adjustment as involving only a very short-term smoothing of rates. This argument, however, must account for the many estimated policy rules that appear to indicate that a high degree of quarterly interest rate smoothing is present in the real world. This seemingly straightforward descriptive evidence of slow adjustment from the inertial empirical policy rules is summarized in the next section, while Section 3 outlines the related normative arguments for the optimality of inertial behavior in a New Keynesian model of output and inflation.

Evidence against the existence of an inertial policy rule is obtained from the behavior of market interest rates at the short-term end of the yield curve. As documented in Section 4, there appears to be very little information generally available in financial markets regarding future interest rate movements beyond the next one or two months, which is consistent with the results of Mankiw and Miron (1986) and many others. In contrast, Section 5 derives the term structure implications of monetary policy inertia in a New Keynesian model and shows that the large ρ in an inertial rule implies that typically there *are* predictable future changes in the policy rate, which under rational expectations should be embodied in the term structure. Thus, there is an inconsistency between the term structure implications of quarterly interest rate smoothing and the historical term structure evidence. Furthermore, this inconsistency is robust to a variety of different assumptions about the specification of the model and the policy rule.

Assuming financial markets process information efficiently, the term structure evidence implies that the empirical policy rules displaying substantial partial adjustment are misspecified. Section 6 argues that such partial adjustment could be spuriously attributed to a non-inertial central bank, that is, one that displays no quarterly interest rate smoothing. The first part of this argument is based on the econometric near-observational equivalence of the partial adjustment rule and a non-inertial rule with serially correlated shocks. The second part argues that policymakers—along with financial markets—are uncertain as to the size and the duration of these shocks in real time. This scenario of significant shocks of ex ante unknown persistence may explain the illusion of monetary policy inertia.

2. The Policy Inertia in Estimated Rules

Many recent studies have estimated models of central bank behavior. A sizable fraction of these empirical policy rules or reaction functions follow Taylor (1993) and set the quarterly average level of the short-term policy interest rate (i_t) in response to (four-quarter) inflation ($\bar{\pi}_t$) and the output gap (y_t); however, a lagged policy rate is also included. Accordingly, a typical equation has the generic partial adjustment form (ignoring constants) of

$$i_t = (1 - \rho_1)(g_\pi \bar{\pi}_t + g_y y_t) + \rho_1 i_{t-1} + \xi_t, \tag{2.1}$$

where ρ_1, g_π, and g_y are the coefficients of what is denoted here as Rule 1.

For example, a least squares regression of Rule 1 on U.S. data from 1987:Q4 to 1999:Q4 yields (ignoring constants)

$$\begin{array}{ccccccccc} i_t & = & .27\,(\,1.53\,\bar{\pi}_t & + & .93\,y_t\,) & + & .73\,i_{t-1} & + & \xi_t, \\ & & (.30) & & (.14) & & (.07) & & \end{array} \tag{2.2}$$

$$\sigma_\xi = .36, \quad \bar{R}^2 = .96\ ,$$

where the interest rate is the quarterly average federal funds rate, inflation is defined using the GDP chain-weighted price index (denoted P_t so $\pi_t = 400(\ln P_t - \ln P_{t-1})$ and $\bar{\pi}_t = \frac{1}{4}\Sigma_{j=0}^{3}\pi_{t-j}$), and the output gap is defined as the percent difference between actual real GDP (Q_t) and potential output (Q_t^*) estimated by the Congressional Budget Office (i.e., $y_t = 100(Q_t - Q_t^*)/Q_t^*$). In this regression, the estimated values of the response coefficients—namely, $g_\pi = 1.53$ for the inflation response and $g_y = 0.93$ for the output response—are just above the 1.5 and 0.5 that Taylor (1993) originally proposed. Similar estimates are obtained in other empirical studies.[4]

However, what is particularly notable about such partial adjustment Taylor rule regressions are the large and highly significant estimates of the coefficient on the lagged policy rate, which in the above regression is $\hat{\rho}_1 = 0.73$. Indeed, such significant lagged dependence in the empirical estimation of Rule 1 appears to be an extremely robust result in the literature. For example, across six different quarterly U.S. data samples (differing in output gap definitions), Kozicki (1999) reports a range of $\hat{\rho}_1$ from 0.75 to 0.82, while across 16 different quarterly samples of U.S. data (differing in output, inflation, and sample period definitions), Amato and Laubach (1999) report a range of $\hat{\rho}_1$ from 0.78 to 0.92.

[4] See, for example, Kozicki (1999), Amato and Laubach (1999), Sack (1998), Orphanides (1998), and Judd and Rudebusch (1998).

In contrast to (2.2), if the constraint that $\rho_1 = 0$ is imposed (in the original spirit of Taylor, 1993), then the regression of this non-inertial form of Rule 1 from 1987:Q4 to 1999:Q4 yields

$$\begin{aligned} i_t \;=\; \underset{(.13)}{1.59}\, \bar{\pi}_t \;+\; \underset{(.09)}{.68}\, y_t \;+\; \xi_t, \end{aligned} \tag{2.3}$$

$$\sigma_\xi = .73, \quad \bar{R}^2 = .84, \quad DW = .33\ ,$$

which has a significantly worse fit and severely serially correlated errors, although the estimates of g_π and g_y are not very different. (In both equations, robust standard errors for the coefficients are reported in parentheses.[5])

The evidence for significant lagged dependence is also robust across different variations of the Taylor rule. In particular, Clarida, Gali, and Gertler (2000) recommend a forecast-based specification of the Taylor rule, which I denote as Rule 2,

$$i_t = (1-\rho_2)(g_\pi E_{t-1}\bar{\pi}_{t+4} + g_y E_{t-1} y_t) + \rho_2 i_{t-1} + \xi_t, \tag{2.4}$$

where $E_{t-1}\bar{\pi}_{t+4}$ is the forecast of annual inflation five quarters ahead based on the $t-1$ information set and $E_{t-1}y_t$ is the forecast of the time t output gap based on the $t-1$ information set. An instrumental variables estimate of Rule 2 over the 1987-99 sample is[6]

$$i_t \;=\; .21\,(\, \underset{(.50)}{1.40}\, E_{t-1}\bar{\pi}_{t+4} \;+\; \underset{(.28)}{.90}\, E_{t-1} y_t \,) \;+\; \underset{(.06)}{.79}\, i_{t-1} \;+\; \xi_t, \tag{2.5}$$

$$\sigma_\xi = .41, \quad \bar{R}^2 = .95\ .$$

These parameter estimates are broadly similar to ones for this specification given in Clarida, Gali, and Gertler (2000, table 5), although they report even slower partial adjustment with a $\hat{\rho}_2 = 0.91$. As above, there is a significant contrast in fit with the estimated non-inertial Rule 2, which has the restriction that $\rho_2 = 0$, although again the sizes of the $\hat{g}_\pi$ and $\hat{g}_y$ are similar,

$$i_t \;=\; \underset{(.32)}{1.33}\, E_{t-1}\bar{\pi}_{t+4} \;+\; \underset{(.18)}{.59}\, E_{t-1} y_t \;+\; \xi_t, \tag{2.6}$$

$$\sigma_\xi = 1.09, \quad \bar{R}^2 = .65, \quad DW = .35\ .$$

In short, as many have noted, the partial adjustment forms of Rules 1 and 2 appear to fit the data significantly better than those without partial adjustment. This significant lagged dependence in empirical Taylor-type rules also appears to be a quite general feature found in

[5] There is some residual serial correlation in regression (2.2) as well. Several papers use further lags of the interest rate to capture this correlation, but for simplicity this paper just considers first order terms.

[6] Four lags of inflation, the output gap, and the interest rate are used as instruments.

a variety of countries in Europe and elsewhere. For example, Clarida, Gali, and Gertler (1998) estimate Rule 2 on quarterly European data and obtain estimates of ρ_2 above 0.90 in Germany, France, Italy, and the UK, and Nelson (2000) provides estimates of Rule 1 for the UK that also display significant lagged dependence.

Besides a uniform set of estimates of the size and significance of the lagged policy interest rate in the empirical rules, there is also a standard partial adjustment interpretation of this term. When ρ_1 or ρ_2 equals zero, the current policy rate is based solely on *current* macroeconomic performance (actual or expected). When these lag coefficients are positive (but less than one), then the current policy rate is set equal to some fraction of this current desired interest rate and some fraction of last quarter's rate. This conventional wisdom of quarterly monetary policy partial adjustment has been advanced by numerous authors, including Goodhart (1999), Levin, Wieland, and Williams (1999), Woodford (1999), Amato and Laubach (1999), Clarida, Gali, and Gertler (2000), and Sack and Wieland (2000). Such partial adjustment behavior is often termed "interest rate smoothing" because the resulting interest rate series will be less volatile than would be suggested by the determinants of policy. Indeed, the degree of quarterly interest rate smoothing or inertia is often measured by the size of the speed of adjustment coefficient (ρ_1 or ρ_2) because as it increases for a given policy rule, the standard deviation of Δi_t falls.[7]

Given the simple form of Rules 1 and 2, it may be that the significant estimated partial adjustment reflects misspecification. One such misspecification might involve structural shifts in the parameters of the policy rule for different policy regimes; however, Clarida, Gali, and Gertler (2000) and others provide rule estimates over numerous subsamples, and all display a large partial adjustment lag coefficient. Alternatively, Rules 1 and 2 may be misspecified because of the omission of a persistent, serially correlated variable that influences monetary policy. Such an omitted variable could also produce the spurious appearance of partial adjustment in the estimated rule. However, in a wide variety of less parsimonious specifications of i_t^*, a significant estimate of ρ is still obtained. For example, McNees (1992), McCallum and Nelson (1999), and Fair (2000) estimate more complicated structural monetary policy rules and obtain significant evidence of policy inertia with partial adjustment coefficients on the order of 0.8 or higher. Similarly, numerous monetary VAR estimates, which contain as an interest rate equation

[7] This is true for a single stochastic equation but is, of course, not necessarily true in the context of a complete model. For example, in the Rudebusch and Svensson (1999) model, increasing ρ_1 can increase the variance of Δi_t and even lead to dynamic instability. Also, a rule with a larger autoregressive coefficient than another rule does not necessarily produce smoother interest rates, because the policy rate volatility also depends on the volatility of the other arguments of the rules.

the most popular recent implementation of an empirical policy reaction function, also show significant inertia. These VAR estimated interest rate equations contain large and significant lagged interest rate coefficients despite including a wide variety of other regressors. For example, Rudebusch (1998, table 2) reports the sum of the lagged funds rate coefficients in the reduced form of a well-known quarterly VAR interest rate equation (which has 24 non-interest-rate regressors) as 0.95.

Still, this paper takes issue with the "conventional wisdom" that quarterly monetary policy inertia exists and argues that the common empirical monetary policy rules are indeed misspecified. However, as described below, this misspecification is difficult to detect directly; thus, this paper focuses on indirect term structure evidence of the misspecification. As a first step, the next section introduces a model of the economy and considers the optimality of policy inertia.

3. Optimal Monetary Policy Inertia

The above empirical policy rules imply a very slow speed of adjustment. A $\hat{\rho}_1$ or $\hat{\rho}_2$ of 0.8 implies a 20 percent adjustment *each quarter*, so in a year, a central bank would complete only 60 basis points of a desired one percentage point change. Still, such sluggish behavior may be optimal for a central bank. An obvious explanation is that in a realistic model, i_{t-1} is likely a state variable, so the fully optimal instrument rule would include a response to its value (e.g., Rudebusch and Svensson, 1999).[8] An important example of this occurs in an explicitly forward-looking model, where partial adjustment can be optimal if the private sector is forward-looking and the monetary policymaker is credibly committed to a gradual policy rule (see Woodford, 1999, Rotemberg and Woodford, 1999, Levin, Wieland, and Williams, 1999, Sack and Wieland, 2000, and Amato and Laubach, 1999). In such a situation, the small inertial changes in the policy interest rate that are expected in the future can have a large effect on current supply and demand and can help the central bank control macroeconomic fluctuations.[9]

This argument can be elucidated within an empirical New Keynesian model. The key aggregate relationships of the simple theoretical version of this model are

$$\pi_t = \mu_\pi E_t \pi_{t+1} + (1 - \mu_\pi)\pi_{t-1} + \alpha_y y_t + \varepsilon_t, \tag{3.1}$$

[8] As noted by Levin, Wieland, and Williams (1999), this may be especially true for restricted rules, such as Rules 1 and 2, because the lagged policy rate may proxy for excluded lags of other variables.

[9] As a second reason why partial adjustment may be optimal, Sack (2000), Sack and Wieland (2000), and Söderström (2000) cite multiplicative parameter uncertainty; however, the results of Rudebusch (1999) and Peersman and Smets (1999) indicate that the effect of such uncertainty is quite modest empirically.

$$y_t = \mu_y E_t y_{t+1} + (1-\mu_y) y_{t-1} - \beta_r (i_t - E_t \pi_{t+1} - r^*) + \eta_t, \tag{3.2}$$

where $E_t\pi_{t+1}$ and $E_t y_{t+1}$ are the expectations of period $t+1$ inflation and output conditional on a time t information set and r^* is the equilibrium real rate. Much of the appeal of this model lies in its foundations in a dynamic general equilibrium model with temporary nominal price rigidities. An empirical version of this model suitable for quarterly data, where longer leads and lags appear appropriate given the institutional length of contracts and delays in information flows and processing, reformulates (3.1) and (3.2) as

$$\pi_t = \mu_\pi E_{t-1}\bar{\pi}_{t+3} + (1-\mu_\pi)\Sigma_{j=1}^{4}\alpha_{\pi j}\pi_{t-j} + \alpha_y y_{t-1} + \varepsilon_t, \tag{3.3}$$

$$y_t = \mu_y E_{t-1} y_{t+1} + (1-\mu_y)\Sigma_{j=1}^{2}\beta_{yj} y_{t-j} - \beta_r r_{t-1} + \eta_t, \tag{3.4}$$

where $E_{t-1}\bar{\pi}_{t+3}$ represents the expectation of average inflation over the next year and r_{t-1} is the real rate relevant for output.[10] In particular, r_{t-1} is defined as a weighted combination of an ex ante 1-year rate and an ex post 1-year rate:

$$r_{t-1} = \mu_r(E_{t-1}\bar{\imath}_{t+3} - E_{t-1}\bar{\pi}_{t+4}) + (1-\mu_r)(\bar{\imath}_{t-1} - \bar{\pi}_{t-1}), \tag{3.5}$$

where $\bar{\imath}_t$ is a four-quarter average of past interest rates, i.e., $\bar{\imath}_t = \frac{1}{4}\Sigma_{j=0}^{3} i_{t-j}$.

This model allows the analysis below to consider a wide range of explicit forward-looking behavior, which is important given the uncertainty about the quantitative importance of expectations. As a theoretical matter, the values of μ_π, μ_y, and μ_r are not clearly determined.[11] Furthermore, the empirical evidence on the appropriate values of these parameters is not decisive. At one extreme, the model with μ_π, μ_y, and μ_r set equal to zero matches the completely adaptive expectations model of Rudebusch and Svensson (1999) and Rudebusch (1999), which has had some success in approximating the time series data in the manner of a small estimated VAR (see Fuhrer, 1997, and Estrella and Fuhrer, 1998). In this extreme model, inflation and output are not based on explicit expectations but are based completely on lags (which may implicitly represent adaptive expectations), and the real rate is an average of the past four quarters of real rates (which may represent planning and production lags from interest rates to output or an adaptive expectations version of the term structure as in Modigliani and Schiller, 1973).

[10] For explicit derivations or discussion, see Woodford (1996), Goodfriend and King (1997), Walsh (1998), Clarida, Gali, and Gertler (1999), Svensson (1999a, b), McCallum and Nelson (1999), and Rudebusch (2000).

[11] From well-known contracting models of price-setting behavior it is possible to derive an inflation equation with $\mu_\pi \approx 1$ (e.g., Roberts, 1995). However, many authors assume that with realistic costs of adjustment and overlapping price and wage contracts there will be some inertia in inflation, so μ_π will be less than one, and with even higher costs for adjusting output, μ_y is likely much less than one as well. See Svensson (1999a, b) and Fuhrer and Moore (1995) and Fuhrer (1997).

However, estimated forward-looking models have also had some success in fitting the data, as in Rotemberg and Woodford (1999), Fuhrer (2000), McCallum and Nelson (1999), and Fuhrer and Moore (1995). The analysis below takes a very eclectic view and conditions on a wide range of possible values for μ_π, μ_y, and μ_r. In contrast, there is less contention regarding the values of the other parameters in the model, and these are set equal to the values given in Table 1, which are obtained from the data in Rudebusch (2000) for a very similar model.[12]

Table 2 summarizes the optimal amount of monetary policy inertia for various models, rules, and loss functions. The table displays the lag coefficients ρ_1 and ρ_2 from the *optimal* versions of Rules 1 and 2, across models with a range of forward-looking behavior. In particular, for inflation, μ_π is set equal to 0.1, 0.3, or 0.5 because the many available empirical estimates described in Rudebusch (2000) suggest that a very broad plausible range for μ_π is between 0 and 0.6.[13] Similarly for output, μ_y is set equal to either 0 or 0.3. Almost all empirical estimates have assumed that $\mu_y = 0$ (e.g., Fuhrer and Moore 1995); however, Fuhrer (2000) estimates a habit persistence model that implies that μ_y is approximately equal to 0.3 (see Rudebusch, 2000). Finally for interest rates, μ_r is varied over essentially the entire range, so $\mu_r = 0.1$, 0.5, or 0.9, because the colinearity of many interest rates makes it hard to obtain decisive empirical evidence on its value (e.g., Fuhrer and Moore, 1995).

Rules 1 and 2 are optimized according to a standard loss function in which the central bank minimizes the variation in inflation around its target π^*, the output gap, and changes in the interest rate (see Rudebusch and Svensson, 1999, and Clarida, Gali, and Gertler, 1999):

$$\mathrm{E}\left[L_t\right] = \mathrm{Var}\left[\bar{\pi}_t - \pi^*\right] + \lambda \mathrm{Var}[y_t] + \nu \mathrm{Var}\left[\Delta i_t\right], \tag{3.6}$$

where $\Delta i_t = i_t - i_{t-1}$, and the parameters $\lambda \geq 0$ and $\nu \geq 0$ are the relative weights on output and interest rate stabilization, respectively, with respect to inflation stabilization. Two different parameterizations of this loss function are considered in Table 2. Columns 4 and 5 provide the optimal ρ_1 and ρ_2 with $\lambda = 1$ and $\nu = 0.5$, the baseline case in Rudebusch and Svensson (1999), while columns 6 and 7 provide the optimal ρ_1 and ρ_2 with $\lambda = 1$ and $\nu = 0.1$, which involves a *very* minimal motive to reduce interest rate volatility.[14] With ν equal to 0.5 or 0.1, respectively,

[12] These are also little different from the values given in Rudebusch and Svensson (1999); in any case, the qualitative results below are robust to their variation.

[13] For example, Fuhrer (1997) estimates μ_π of about zero, while Fuhrer and Moore (1995) assumes μ_π is 0.5.

[14] The results in Table 2 are obtained by numerically maximizing the loss function over the parameters g_π, g_y, and ρ_j in the model of (3.3), (3.4), and Rule j. The policy rule is subject to an i.i.d. error with $\sigma_\xi = 0.4$, which is in the range of the empirical estimates in Section 2. As usual, the policy rule is assumed to be perfectly credible, so agents know the rule and assume (correctly) that it will be followed. The results are obtained using the "AIM" algorithm (Anderson and Moore, 1985) available at http://www.bog.frb.fed.us/pubs/oss/oss4/aimindex.html).

these loss functions equally penalize a 1 percent output gap, a 1 percentage point inflation gap, and a 1.41 or a 3.16 percentage point quarterly change in the funds rate. This appears to be a plausible range of penalty on interest rate volatility given the various reasons to reduce such volatility that have been proposed in the literature.[15]

As is evident in Table 2, a large range of optimal lag coefficients can be rationalized given the uncertainty about the specification of the model and the loss function. For some combination of model and loss function any optimal ρ_1 or ρ_2 between 0 (or just below) and 0.9 can be obtained. Surprisingly, there is little dependence of the optimal ρ_1 or ρ_2 on the values of μ_π and μ_y, the weights on explicit inflation or output expectations in the model, or on the value of ν, the weight on the "interest rate smoothing motive" in the loss function (except when $\mu_r = .1$). Instead, the degree of optimal quarterly interest rate smoothing is crucially dependent on the value of μ_r, which determines the degree to which interest rate expectations are forward-looking. This is consistent with the interpretation of Woodford (1999) and others that policy inertia is optimal when it alters expectations of future interest rates, which are also important determinants of current demand.

4. Term Structure Evidence on Interest Rate Predictability

The preceding two sections documented the large and significant estimated coefficient on the lagged interest rate in quarterly central bank reaction functions as well as the optimality of such monetary policy partial adjustment or inertia when economic agents are forward-looking with respect to future interest rate movements. This section focuses on measuring how much financial market participants actually know about future interest rate movements. This evidence will provide some crucial benchmarks for the next section, which quantifies the term structure implications of monetary policy inertia.

The partial adjustment of monetary policy by a central bank suggests that there are forecastable future movements in the policy interest rate. The amount of such forecastable variation can be measured with a standard term structure regression such as:

$$i_{t+j} - i_{t+j-1} = \delta + \gamma(E_t i_{t+j} - E_t i_{t+j-1}) + \psi^j_{t+j}, \tag{4.1}$$

[15] There are three broad such motives (e.g., Lowe and Ellis, 1997). First, interest rate volatility may induce instability in financial markets (e.g., Goodfriend, 1991, Rudebusch, 1995, Cukierman, 1996). Second, large interest rate changes may be difficult to achieve politically because of the decision-making process (e.g., Goodhart, 1997) or because such changes may be taken as an adverse signal of inconsistency and incompetence (e.g., Goodhart, 1999). Finally, smaller interest rate changes seem to make it less likely that the zero bound on nominal interest rates would be reached (though Woodford, 1999, disagrees).

(for $j \geq 1$). This equation regresses the realized change in the policy rate between two adjacent quarters on the expected such change.[16] Under rational expectations, $i_{t+j} = E_t i_{t+j} + e_{t+j}$, where e_{t+j}, the expectational error, has a mean of zero and is uncorrelated with time t information. In this case, the interest rate forecasting regression (4.1) would yield in the limit an estimate of $\hat{\delta} = 0$ and $\hat{\gamma} = 1$. Furthermore, for assessing monetary policy inertia, a statistic of particular interest is the R^2 of this regression, which provides the proportion of the variance of the future change in interest rates that can be forecasted at time t.

Many papers have estimated term structure regressions such as (4.1) for the postwar period using 3- and 6-month or 6- and 12-month Treasury bill spreads as proxies for expectations (see, for example, Mankiw and Miron, 1986, Mishkin, 1988, Cook and Hahn, 1990, and Rudebusch, 1995). These studies typically have obtained R^2's very close to zero. For example, in a 1959-1979 sample, Mankiw and Miron (1986, Table 1) obtain an R^2 of 0.02 in a regression of the change in the 3-month rate on the 3- and 6-month spread. However, these results might be too pessimistic because they typically cover a long sample that may not encompass a consistent monetary policy regime; this is important because the term structure implications derived in the next section assume that agents know the policy rule that the central bank is committed to (see Fuhrer, 1996). As an alternative, I estimate (4.1) with rates on 3-month eurodollar and eurodollar futures, which have been the trading vehicle of choice for hedging short-run future interest rate movements since the mid-1980s.[17] The eurodollar regressions use a short sample from 1988:01 to 2000:01, covering what is arguably a single consistent policy regime.

Denote $ED(t+j)_t$ as the interest rate on eurodollar deposits during quarter $t+j$ that is expected as of the end of quarter t. Thus, $ED(t+1)_t$ is the spot 3-month eurodollar rate at the end of quarter t, and $ED(t+2)_t$ is the rate on a eurodollar futures contract that settles 3 months ahead.[18] Then assume that $ED(t+j)_t = E_t i_{t+j} + \phi_t^j$, where ϕ_t^j is the term premium associated with the jth contract. Under the expectations hypothesis of the term structure, the term premia are assumed to be constant over time, but in practice it is widely recognized that there is some time variation. The consequences of time-varying term premia are discussed below.

Using eurodollar data to predict the one-quarter-ahead change in the quarterly average funds

[16] These "marginal" regressions are common in the literature (e.g., Mishkin, 1988); however, I obtained similar results with other forms as well.

[17] Eurodollar futures contracts are based on the 90-day London Interbank Offered Rate (LIBOR). For further details, see Jegadeesh and Pennacchi (1996).

[18] For the three regressions using eurodollar rates, quarters are defined to start at the eurodollar futures contract settlement dates (which occur about two weeks before the start dates of the usual quarters) in order to capture true two- and three-quarter-ahead expectations.

rate, equation (4.1) with $j = 1$ is estimated from 1988:Q1 to 2000:Q1 as

$$\underset{}{i_{t+1} - i_t} = \underset{(.05)}{-.25} + \underset{(.10)}{.81}\,(ED(t+1)_t - i_t) + \psi^1_{t+1}, \tag{4.2}$$

$$\sigma_{\psi 1} = .30, \quad R^2 = .56 \; .$$

This equation indicates that the 3-month eurodollar rate forecasts Δi_{t+1} quite well (with an average term premium of about 25 basis points). The R^2 indicates that just over 50 percent of the one-quarter-ahead change in the quarterly average funds rate is known at the end of a typical quarter. This is consistent with the evidence and interpretation in Rudebusch (1995) of interest rate smoothing at a weekly and monthly frequency. That is, at the end of quarter t, financial markets have some information about changes during the first several weeks of the following quarter.[19] In addition, in this regression, changes in the funds rate *during* quarter t (which are of course known at the end of quarter t) will also help predict the quarterly average change Δi_{t+1}. However, after replacing i_t with the end of quarter t funds rate, there still is substantial predictive power with $R^2 \approx .3$.

However, of particular interest in assessing quarterly monetary policy inertia will be the predictive ability at slightly longer horizons. Predicting the one-quarter change in the funds rate two quarters ahead (equation (4.1) with $j = 2$) yields[20]

$$i_{t+2} - i_{t+1} = \underset{(.07)}{-.06} + \underset{(.18)}{.47}\,(ED(t+2)_t - ED(t+1)_t) + \psi^2_{t+2}, \tag{4.3}$$

$$\sigma_{\psi 2} = .42, \quad R^2 = .13 \; .$$

Predicting Δi_{t+3} at quarter t yields

$$i_{t+3} - i_{t+2} = \underset{(.08)}{-.09} + \underset{(.30)}{.40}\,(ED(t+3)_t - ED(t+2)_t) + \psi^3_{t+3}, \tag{4.4}$$

$$\sigma_{\psi 3} = .44, \quad R^2 = .04 \; .$$

These regressions indicate that there is little if any information usually available in financial markets for predicting the level of the funds rate three to six months out ($R^2 = .13$) and no information for predicting it six to nine months out ($R^2 = .04$).[21] These R^2's will be used as benchmarks for assessing the plausibility of monetary policy inertia in the next section. These

[19] In particular, this significant predictive ability for Δi_{t+1} is consistent with the documented ability of a two-month and one-month interest rate spread to predict the one-month-ahead change.

[20] For $j > 1$, the forecast errors will have an $MA(j-1)$ moving average correlation, so robust standard errors are reported in parentheses.

[21] For predicting Δi_{t+2}, the p-value for the hypothesis that $R^2 = 0$ is 0.01, and for predicting Δi_{t+3}, the p-value is 0.15.

results turn out to be only marginally better than the standard ones described above with Treasury bills. They are also consistent with the effect of day-to-day monetary policy surprises on the slope of the term structure as in Kuttner (2000). A surprise change in the policy rate target on a particular day shifts the level of the term structure by a similar amount across all horizons, but carries little information about future changes in rates. This lack of information is also reflected in the general absence of trading in federal funds futures beyond a horizon of four months (see Rudebusch 1998).

Finally, the presence of time-varying term premia should be noted, which, as stressed by Mankiw and Miron (1986), can have severe consequences for empirical regressions like (4.3). The sample estimates of the γ and R^2 of this regression will depend positively on the covariance between the independent and dependent variables, Δi_{t+2} and $ED(t+2)_t - ED(t+1)_t$, and inversely on their variances. Accordingly, as the time variation in the term premia becomes more significant (boosting the independent, noisy variation in the eurodollar spread), the estimates γ and R^2 can be driven away from 1 even in the limit. The standard deviation of the residual to the term structure regression provides a rough upper bound on the size of the term premium. For example, in (4.3), $\psi_t^2 = \phi_t^1 - \phi_t^2 + e_{t+2} - e_{t+1}$, which is a combination of term premia and the expectational errors. The expectational errors are orthogonal to the term premia; thus, the standard deviation of the term premium associated with the $t+2$ and $t+1$ eurodollar spread (i.e., $\phi_t^1 - \phi_t^2$) is smaller than 0.42, the standard deviation of the regression ($\sigma_{\psi 2}$).

5. Term Structure Implications of Policy Inertia

The previous section provided evidence that beyond a horizon of three months there is little predictive information in financial markets about the future path of short-term interest rates. This section explores whether that evidence can be reconciled with a significant degree of quarterly monetary policy inertia. Intuitively, such a reconciliation seems unlikely, for if the funds rate is typically adjusted by only 20 percent toward its desired target in a given quarter, then the remaining 80 percent adjustment should be expected to occur in future quarters. The partial adjustment of the short-term policy interest rate embodied in Rule 1 or 2 with high ρ_1 or ρ_2 implies that there typically is a large amount of predictable future variation in the policy rate. Indeed, this is the essence of *optimal* policy inertia: Because private agents know that the policy rate is likely to be adjusted by a certain amount in the future, they change their behavior today.

The relationship between the forecastable variation in the interest rate, as measured by the

R^2 of the Δi_{t+2} prediction equation, and quarterly policy inertia, as measured by the ρ_1 and ρ_2 in Rules 1 and 2, is illustrated in Figure 1. This figure graphs the (analytical population) value of the R^2 of the regression (4.1), with $j = 2$, as a function of the value of ρ_1 or ρ_2 for a representative case of the model described above, namely, with $\mu_\pi = .3$, $\mu_r = .5$, and $\mu_y = 0$ (and the other parameters given in Table 1).[22] Also, for both policy Rules 1 (2.1) and 2 (2.4), g_π and g_y are set equal to 1.5 and 0.8, respectively, and the rule error is i.i.d. with $\sigma_\xi = 0.4$. (This calibration is in the range of the empirical rule estimates given in Section 2.) Note that even for the non-inertial policy rules there is some predictable future movement in interest rates (with $R^2 = .11$ when $\rho_1 = 0$ and $R^2 = .03$ when $\rho_2 = 0$). For example, the forecasting power with Rule 1 when $\rho_1 = 0$ reflects the fact that there are predictable movements two quarters ahead in the output gap and in the four-quarter inflation rate, which partly determine interest rates. Still, as ρ_1 and ρ_2 increase, the amount of predictable variation in Δi_{t+2} also increases, with R^2 values of .47 at $\rho_1 = 0.8$ and .45 at $\rho_2 = 0.8$.

This basic relationship between predictable interest rate variation and monetary inertia is robust across a wide variety of models and rules. Table 3 examines the same 18 different parameterizations of the model considered in Section 3 and non-inertial ($\rho_1 = \rho_2 = 0$) and inertial ($\rho_1 = \rho_2 = 0.8$) versions of both Rules 1 and 2. (Again, $g_\pi = 1.5$, $g_y = 0.8$, and $\sigma_\xi = 0.4$.) In addition, a time-varying term premium, $\phi_t^1 - \phi_t^2$, is included, which is assumed to be i.i.d. with a standard deviation of 0.15.[23] As noted above, such a term premium reduces the R^2 values. Each model and rule combination reports R_L^2 and R_U^2, which are the 5 percent lower and upper critical values, respectively, for the small-sample distribution of the R^2 (which are appropriate for 95 percent one-sided or 90 percent two-sided tests). These critical values are calculated from 10,000 simulated samples of the model and the given rule (with 100 observations each), and they allow a probabilistic assessment of the historical term structure regression results given in Section 4. The bottom line in the table gives the median R_L^2 and R_U^2 values across all models. Although there is interesting variation across models, the value of ρ_1 or ρ_2 is the key determinant of interest rate predictability, so I focus on the median values. Based on the historical results

[22] As above for Table 2, the unique stationary rational expectations solution for each specified policy rule and model is solved via AIM (see Levin, Wieland, and Williams, 1999, and Anderson and Moore, 1985). The reduced-form representation of the saddle-point solution is computed, the unconditional variance-covariance matrix of the model variables and the term spreads is obtained analytically, and the term structure regression asymptotic R^2 is calculated using the appropriate variances and covariances.

[23] Mankiw and Miron (1986, Table 3) estimate the standard error of this term premium to be 0.16, while in a more complicated time series specification with monthly data, Dotsey and Otrok (1995) estimate it to be 0.13. Also this standard deviation is about one-third the size of the regression standard error of (4.3), which includes the eurodollar term premia and the orthogonal expectational error.

with eurodollar data, the benchmark R^2 value for the Δi_{t+2} prediction regression is 0.13. This value is always included in the confidence intervals for the non-inertial $\rho_1 = 0$ and $\rho_2 = 0$ cases; indeed, it is quite close to the small-sample means (which are not shown).[24] In contrast, for the inertial policy rules, the R^2 confidence intervals with $\rho_1 = .8$ and with $\rho_2 = .8$ both lie above the historical R^2 value.

As shown in Table 4, very similar results are obtained for the Δi_{t+3} prediction regression. Again, a R_L^2 and R_U^2 pair is calculated for each of the model and rule combinations used in Table 3. Recall from the previous section that the benchmark value of R^2 values from the historical data is 0.04. As before, the historical value is contained in the confidence intervals of the non-inertial policy rules, but it is not contained in the median inertial policy rule interval.

In sum, quarterly partial adjustment and interest rate smoothing or inertia do not appear to be consistent with the lack of information in the term structure of interest rates.

6. The Illusion of Monetary Policy Inertia

The large estimated lag coefficients in the empirical partial adjustment policy rules appear to be strong evidence of monetary policy inertia. However, such quarterly inertia is inconsistent with the very low interest rate forecastability in the term structure of interest rates. This section shows how the partial adjustment evidence in the empirical rules may be explained by a rationale other than policy inertia.

As a first step, note that there is a large literature that argues that partial adjustment models such as Rules 1 and 2 are very difficult to identify and estimate empirically in the presence of serially correlated shocks (e.g., Griliches, 1967, Blinder, 1986, Hall and Rosanna, 1991, and McManus, Nankervis, and Savin, 1994). In particular, a standard policy rule with slow partial adjustment and no serial correlation in the errors will be very difficult to distinguish empirically from a policy rule that has immediate policy adjustment but highly serially correlated shocks. Using the 1987-1999 data from Section 2, this latter form of Rule 1 is estimated as[25]

$$i_t = \underset{(.24)}{1.24}\,\bar{\pi}_t + \underset{(.10)}{.33}\,y_t + \xi_t, \qquad \xi_t = \underset{(.06)}{.92}\xi_{t-1} + \omega_t, \tag{6.1}$$

$$\sigma_\omega = .36, \quad \bar{R}^2 = .96 \ .$$

[24] The term premia also reduce the slope estimates in the term structure regression to close to the historical values.

[25] Rule 1 with an AR(1) error is estimated via maximum likelihood, while rule 2 with an AR(1) error is estimated with an instrumental variables version of the Hildreth-Lu procedure.

This rule assumes that $\rho_1 = 0$ but allows for first order serial correlation where $\rho_e = 0.92$ is defined as the AR(1) error coefficient. For Rule 2, the corresponding AR(1) estimate is

$$i_t = \underset{(.66)}{2.00}\, E_{t-1}\bar{\pi}_{t+4} + \underset{(.24)}{.39}\, E_{t-1}y_t + \xi_t, \qquad \xi_t = \underset{(.11)}{.77}\xi_{t-1} + \omega_t, \tag{6.2}$$

$$\sigma_\omega = .36, \quad \bar{R}^2 = .86\ .$$

These two estimated autocorrelated error versions of Rules 1 and 2 display a fit to the data as well as $\hat{g}_\pi$ and $\hat{g}_y$ that are broadly comparable to the partial adjustment forms (2.2) and (2.5). In theory, it is possible to estimate a rule that allowed for both partial adjustment and serially correlated shocks; however, as argued in the research cited above, it will be very hard to distinguish these two models given the other arguments of the rules. As in Blinder (1986), distinguishing the two cases depends crucially on the other regressors in the rule and will be especially difficult in this setting where output and inflation depend on interest rates. This near-observational-equivalence provides the motivation for examining the indirect term structure evidence as above.

The estimated partial adjustment policy rules failed the term structure test in Section 5 by implying too much interest rate forecastability. However, the near-observationally-equivalent estimated rules (6.1) and (6.2) with serially correlated AR(1) shocks may not pass this test either. If agents know the persistent nature of the shocks, then even if the central bank adjusts policy immediately to output and inflation, the projection of the lingering effects of the known autocorrelated shocks should imply significant interest rate predictability. Indeed, such predictability is displayed in panel A of Table 5 (which, for brevity, only considers the model with $\mu_\pi = .3$, $\mu_r = .5$, and $\mu_y = 0$). Rules 1 and 2 take a form similar to the ones above, with $g_\pi = 1.5$, $g_y = .8$, $\rho_1 = \rho_2 = 0$, and an AR(1) shock calibrated with $\rho_e = 0.8$ and $\sigma_\omega = 0.4$. The model is solved in the usual full commitment fashion so all agents in the economy know the serial correlation pattern of the policy rule shocks.[26] The resulting R^2_L and R^2_U pairs for the Δi_{t+2} and Δi_{t+3} prediction regressions, given as the middle four entries in each line, show that these rules with known serially correlated shocks display even more interest rate forecastability than the partial adjustment rules. Also included as the final two entries in each line are the mean and standard error of the small-sample distributions of $\hat{\rho}_1$ and $\hat{\rho}_2$ from the estimated partial adjustment Rules 1 and 2 under the incorrect assumption that the policy rate shocks are serially

[26] This is done by solving the model with the quasi-differenced forms of the rules in AIM.

uncorrelated.[27] Note the mean $\hat{\rho}_1$ and $\hat{\rho}_2$ are about 0.8, closely matching the empirical partial adjustment rule estimates; thus, serially correlated shocks could account for the conventional policy rule estimates (2.2) and (2.5).

The interest rate forecastability in panel A reflects the economy-wide knowledge of the time series persistence properties of the policy rule shocks. The alternative is shown in panel B, which provides results for the same specifications as in panel A except the serial correlation of the shocks is unknown to everyone in the economy. In particular, agents assume that the shocks are serially uncorrelated rather than following the true AR(1) process. This scenario is able to match the term structure evidence of a very low forecastability of interest rates. The confidence intervals contain the benchmark 0.13 and 0.04 R^2 values for the Δi_{t+2} and Δi_{t+3} predictability regressions. In addition, this scenario is able to account for some of the spurious evidence of sluggish policy partial adjustment, supplying mean $\hat{\rho}_1$ and $\hat{\rho}_2$ in estimates of the (incorrect) partial adjustment forms for Rules 1 and 2 that are significantly greater than zero.

The key remaining issue is plausibility: Are there shocks to the policy rule that appear serially uncorrelated ex ante but are serially correlated ex post? The rest of this section argues that panel B does provide a plausible reconciliation of the empirical term structure and policy rule results.

Indeed, one such shock has already been discussed in the policy rules literature. Several recent papers (e.g., Smets, 1999, Rudebusch, 1999, 2000, and Orphanides, et al., 1999) have argued that setting monetary policy according to Taylor-type rules requires relying on a contemporaneous or real-time estimate of the output gap, denoted $y_{t|t}$, which is difficult to measure accurately. Real-time output gap estimates are often modeled as noisy estimates of the final series, i.e., $y_{t|t} = y_t + n_t$, where the stochastic error is the measurement error plaguing the policymaker (i.e., the revision from the standpoint of the end of the sample). Furthermore, based on the historical data, the estimated revisions, n_t, are large and persistent and often modeled as an AR(1) process with lag coefficient of 0.85. In this case, if the central bank follows Rule 1 with no partial adjustment or error in real time,

$$i_t = g_\pi \bar{\pi}_t + g_y y_{t|t}, \tag{6.3}$$

[27] That is, in this scenario, regressions (2.1) and (2.4) are estimated assuming $\rho_e = 0$ and neither ρ_1 nor ρ_2 is equal to zero, when in fact $\rho_e = 0.8$ and $\rho_1 = \rho_2 = 0$. The regressions are estimated on 10,000 samples of simulated data with 100 observations each.

the econometrician working with the final data will estimate

$$i_t = \hat{g}_\pi \bar{\pi}_t + \hat{g}_y y_t + k_t, \tag{6.4}$$

where the error $k_t = \hat{g}_y n_t$ is the highly serially correlated real-time data noise. From the standpoint of the econometrician, these are serially correlated ex post errors; however, in real time, the errors are ex ante uncorrelated as required in panel B. Even worse, the econometrician may estimate the partial adjustment form of Rule 1 and incorrectly interpret the serially correlated ex post revisions as evidence of policy inertia and interest rate smoothing. Indeed, Lansing (2000) provides a careful simulation study that demonstrates the potential effectiveness of real-time output gap errors to account for the spurious evidence of policy inertia in exactly this fashion.

Still, ex post serially correlated output gap revisions are probably not a general enough explanation to reconcile the policy inertia and term structure evidence completely. First, although such an explanation applies to Taylor-type rules, there are many other estimated historical reaction functions that do not include an output gap, but still exhibit significant persistence, for example, McNees (1992), McCallum and Nelson (1999), Fair (2000), and the VAR interest rate equations described above. Second, based on a very tentative reconstruction of real-time output gap data, Orphanides (1998) reports significant policy inertia even when estimating partial adjustment rules using the real-time data (also see Evans, 1998). Thus, although output gap revisions are unlikely to be the main source of ex post serially correlated shocks, they are at the very least illustrative.

To find a more convincing candidate, it is useful to re-examine the original analysis of Taylor (1993), which put forward a description of policy without interest rate smoothing or partial adjustment. Taylor argued that recent historical monetary policy had followed a rule only as a guide, so deviations from the rule during various episodes were an appropriate response to special circumstances, not evidence of partial adjustment. This view is illustrated in Figure 2, which displays the historical values of the funds rate (solid line) and the fitted values from the estimated Rule 1 in (6.1), which allows for serially correlated shocks. These large persistent shocks, the deviations between the two lines, appear to correspond to several special episodes (Kozicki, 1999). Most notably, the deviations in 1992 and 1993 are commonly interpreted as responses to a disruption in the flow of credit. As Fed Chairman Alan Greenspan testified to Congress on June 22, 1994: "Households and businesses became much more reluctant to borrow and spend and lenders to extend credit—a phenomenon often referred to as the 'credit crunch.' In an endeavor to defuse these financial strains, we moved short-term rates lower in a long series of steps that ended in the late summer of 1992, and we held them at unusually low levels through the end of 1993—both absolutely and, importantly, relative to inflation." Thus, this episode is better described as a persistent "credit crunch" shock than as a sluggish partial adjustment to a known desired rate.[28] In terms of the Taylor rule, the disruption of credit supply can be treated as a temporary lowering of the equilibrium real rate, which the Fed responds to by lowering the funds rate (relative to readings on output and inflation).[29] Furthermore, assuming that financial markets and the Fed did not know when the credit crunch would end, the persistence of the shock was unknown ex ante, so it would not have been incorporated into the term structure.[30] Similarly, a worldwide financial crisis appeared to play a large role in lowering rates in 1998 and 1999, and commodity price scares pushed rates up in 1988-89 and 1994-95.[31] The duration of these episodic factors also appeared to be difficult to predict ex ante, and their termination would not be incorporated into the term structure with any great certainty.

As this description of credit crunches and financial crises should make clear, these special circumstances or shocks are essentially unique episodes and are not being generated according to a single given time series process, as suggested by (6.1). The modeling of the shocks to the Taylor rule as a single, well-specified AR(1) process is simply a convenient econometric approximation. This approximation is misleading because it seems to imply that the persistence of the shocks or omitted factors can be known ex ante, when in fact they are unique episodes. In precisely this spirit, for example, Gerlach and Schnabel (2000) provide an alternative modeling strategy for their estimates of the Taylor rule for Europe in the 1990s—essentially providing dummy variable intercept shifts for a large persistent rule deviation. They find (p. 167) that a European Taylor rule fits quite well without any partial adjustment but with ". . . dummies for the period 1992:3-1993:3 to control for policy responses to intra-European exchange market pressures in this period" Assuming such dummies are of unknown duration ex ante, as is likely, this rule

[28] Kozicki (1999, p. 24) also makes the point that ". . . information and events outside the scope of Taylor-type rule specifications . . ." often appear to influence policy actions. This would be troubling if the non-inertial forms of Rules 1 and 2 were too parsimonious and excluded a serially correlated explanatory variable such as a commodity price index, so that the persistent shocks reflected a serially correlated omitted variable. However, as noted above, the empirical reaction function literature, including monetary VARs, has placed the proverbial kitchen sink on the right-hand side in attempts to explain the policy rate, yet serially correlated errors remain, which are modeled through lagged interest rates and partial adjustment. In addition, a serially correlated omitted explanatory variable would provide information about future movements that would be incorporated into the term structure.

[29] In contrast to the "shocks" in the VAR literature, these deviations from the rule are unanticipated at a quarterly frequency but are not exogenous because they respond to developments in the econmy.

[30] Rudebusch (1998, figure 7) provides an analysis of financial market expectations for this particular episode.

[31] Federal Reserve Governor Laurence Meyer (1999, p. 7) had this explanation for the easing of policy during late 1998: "There are three developments, each of which, I believe, contributed to this decline in the funds rate relative to Taylor Rule prescription. The first event was the dramatic financial market turbulence, following the Russian default and devaluation. The decline in the federal funds rate was, in my view, appropriate to offset the sharp deterioration in financial market conditions, including wider private risk spreads, evidence of tighter underwriting and loan terms at banks, and sharply reduced liquidity in financial markets."

formulation will also match the term structure results.

7. Conclusion

Empirical monetary policy rules with large estimated coefficients on the lagged policy interest rate, which are very prevalent in the literature, are widely interpreted as indicating a sluggish adjustment of the policy rate to its determinants—on the order of only about 20 percent per quarter. This partial adjustment implies predictable future changes in the policy rate over horizons of a few quarters, which does not accord with the lack of information about such changes in financial markets. This paper proposes a resolution of this empirical inconsistency by providing an alternative interpretation of the large lag coefficients in the estimated policy rules. These coefficients reflect serially correlated or persistent special factors or shocks that cause the central bank to deviate from the policy rule in unpredictable ways and for unpredictable durations.

This argument uses indirect term structure evidence to dismiss the partial adjustment rule. As noted above, it appears difficult to develop direct evidence against the partial adjustment rule (in the form of non-rejection of the $\rho = 0$ hypothesis). In particular, the uncertainty in modeling the desired policy rate, given the endogeneity of its determinants, the real-time nature of the information set, as well as small samples available, makes any direct evidence from estimated rules difficult to obtain. For example, the rule with partial adjustment and the rule with serially correlated shocks both appear to fit the data as empirical reaction functions. However, they have very different economic interpretations. In the former rule, persistent deviations from an output and inflation response occur because policymakers are slow to react. In the latter rule, these deviations reflect the policymaker's response to other influences of unknown duration. The two types of rules can be distinguished, however, by their very different implications for the term structure. Only the serially correlated shocks rule (with episodic or unique shocks) is consistent with the historical evidence showing that the term structure is largely uninformative about the future course of the policy rate.

There may be other possible reconciliations of the policy rule and term structure empirical results. For example, it may be that the rational expectations hypothesis of the term structure cannot be applied and the associated term structure interpretations above are spurious. One way in which this hypothesis may fail is that expectations are not rational, but this would undermine many aspects of any explicitly forward-looking macroeconomic modeling exercise such as the one above. Or term premia for short-term interest rates may be even more volatile than assumed above; however, if rates are driven by volatile term premia, then it seems unlikely that they can communicate the subtle expectations of future monetary policy as required in the literature on optimal monetary policy inertia.

It is also possible that there is some intermediate case of partial adjustment, a ρ_1 or ρ_2 of 0.4, say, along with some serially correlated shocks, that is not strictly rejected by the term structure evidence. (Note, for example, the nonlinearity in Figure 1.) However, it should be noted that while real-world discussions of monetary policy sometimes mention the "incrementalism" and "gradualism" of smoothing the policy rate over the next several weeks, there is no acknowledgment of *quarterly* interest rate smoothing.[32] As the *New York Times* (July 26, 2000) summarized of recent Congressional testimony: "Alan Greenspan, the Federal Reserve chairman, said today that the central bank's decision about whether to raise interest rates again at its meeting next month would hinge in large part on economic data released in coming weeks." That is, there was little if any pent-up pressure from the past for further adjustment.

In future research, the empirical rules given in Section 6 can be improved as further effort is made in estimating rules without the crutch of partial adjustment. Given the similar estimates above of g_π and g_y across rules, it may be that past conclusions about these coefficients, as in Clarida, Gali, and Gertler (2000), are robust to the exact formulation of serial correlation in the rule. However, the lagged policy rate, though useful in mopping up residual serial correlation, should not be given a structural partial adjustment interpretation with regard to central bank behavior. In particular, using the partial adjustment rule in a model as a representation of historical policy (as in Levin, Wieland, and Williams, 1999, and many other studies) may give very misleading results, especially about the nature of optimal policy inertia.[33]

With regard to optimality, the maintained hypothesis of economics for central banks, as for other agents in the economy, is that the non-inertial policy rule apparently used in practice is optimal, and certainly, the rule can be rationalized as such in particular models as in Table 2. However, it should be stressed that there are many aspects of the monetary policy process still to be modeled, especially imperfect credibility and uncertainty (see Rudebusch, 1999).

[32] And as noted in the introduction, models of monthly interest rate smoothing imply very little if any quarterly interest rate smoothing.

[33] This is the view of Donald Kohn, a Federal Reserve research director, who dismisses optimal policy inertia (in Taylor, 1999, p. 317): ". . . a sluggish reaction is working better here, in a forward-looking model, because the private sector knows what is going on and therefore stabilizes the economy. In the real world, this is not what happens. Both the private sector and the policymaker learn gradually about the size of a shock, which could perhaps lead to instability with a slow policy reaction."

Also, the absence of partial adjustment does not mean that central banks are not trying to influence long-term interest rates. However, in order to influence the long rate, central banks only must present a clear path for the policy rate that can shape expected future rates. The partial adjustment rule provides one such path, but it is not the only one. As noted by Goodfriend (1991) and Rudebusch (1995), an ex ante constant path, which is approximately what the non-inertial rules deliver, is another obvious choice.

Finally, further careful analysis of the empirical policy rule is required in modeling and identifying the shocks. Section 6 provides a simple formulation for adding shocks to a policy rule. A better specification may link persistent shocks in both the rule and the rest of the model. A bout of credit frictions or impediments may lower the equilibrium real rate and provide a persistent negative shock to the policy rule and to the output equation as well (see Rudebusch, 1999). Alternatively, an idiosyncratic inflation scare may provide a shock to the rule and to inflation expectations more broadly.

Table 1
Model Parameter Values

Coefficient	Value
$\alpha_{\pi 1}$	.67
$\alpha_{\pi 2}$	-.14
$\alpha_{\pi 3}$	.40
$\alpha_{\pi 4}$	.07
α_y	.13
β_{y1}	1.15
β_{y2}	-.27
β_r	.09
σ_ε	1.012
σ_η	.833

Table 2
Optimal Lag Coefficients for Policy Rules 1 and 2

Model			Loss: $\lambda = 1; \nu = .5$		Loss: $\lambda = 1; \nu = .1$	
μ_r	μ_π	μ_y	ρ_1	ρ_2	ρ_1	ρ_2
.1	.1	0	.29	.46	.00	.22
.1	.3	0	.26	.47	-.01	.24
.1	.5	0	.26	.54	.00	.33
.1	.1	.3	.19	.40	-.06	.15
.1	.3	.3	.19	.43	-.06	.19
.1	.5	.3	.21	.52	-.05	.27
.5	.1	0	.57	.64	.48	.56
.5	.3	0	.56	.65	.47	.56
.5	.5	0	.57	.71	.47	.63
.5	.1	.3	.56	.67	.50	.64
.5	.3	.3	.58	.70	.49	.66
.5	.5	.3	.58	.77	.50	.71
.9	.1	0	.72	.76	.71	.76
.9	.3	0	.72	.77	.71	.78
.9	.5	0	.74	.83	.71	.82
.9	.1	.3	.74	.82	.73	.86
.9	.3	.3	.74	.85	.73	.89
.9	.5	.3	.76	.91	.74	.89

Table 3
Predicting Δi_{t+2} with Various Models and Rules

Model			Rule 1				Rule 2			
			$\rho_1 = 0$		$\rho_1 = .8$		$\rho_2 = 0$		$\rho_2 = .8$	
μ_r	μ_π	μ_y	R_L^2	R_U^2	R_L^2	R_U^2	R_L^2	R_U^2	R_L^2	R_U^2
.1	.1	0	.03	.21	.15	.57	.01	.12	.24	.58
.1	.3	0	.04	.23	.27	.69	.01	.12	.30	.65
.1	.5	0	.10	.37	.72	.98	.02	.13	.39	.74
.1	.1	.3	.05	.24	.11	.40	.03	.17	.13	.42
.1	.3	.3	.05	.24	.14	.45	.03	.17	.15	.44
.1	.5	.3	.05	.28	.20	.54	.05	.19	.16	.45
.5	.1	0	.02	.23	.16	.51	.01	.12	.22	.55
.5	.3	0	.04	.23	.25	.66	.01	.11	.27	.61
.5	.5	0	.09	.35	.52	.89	.02	.12	.35	.70
.5	.1	.3	.05	.24	.10	.39	.03	.17	.12	.40
.5	.3	.3	.05	.24	.13	.43	.03	.17	.13	.41
.5	.5	.3	.06	.25	.15	.45	.05	.18	.13	.40
.9	.1	0	.03	.20	.15	.49	.00	.12	.20	.52
.9	.3	0	.03	.22	.23	.63	.01	.10	.24	.58
.9	.5	0	.09	.33	.42	.81	.02	.11	.32	.66
.9	.1	.3	.04	.23	.10	.38	.03	.17	.11	.38
.9	.3	.3	.04	.23	.11	.40	.03	.17	.12	.39
.9	.5	.3	.05	.24	.12	.38	.04	.18	.11	.35
Median			.05	.24	.15	.50	.03	.15	.18	.48

Table 4

Predicting Δi_{t+3} with Various Models and Rules

Model			Rule 1				Rule 2			
			$\rho_1 = 0$		$\rho_1 = .8$		$\rho_2 = 0$		$\rho_2 = .8$	
μ_r	μ_π	μ_y	R^2_L	R^2_U	R^2_L	R^2_U	R^2_L	R^2_U	R^2_L	R^2_U
.1	.1	0	.02	.19	.10	.45	.01	.12	.10	.45
.1	.3	0	.03	.21	.17	.64	.01	.12	.14	.52
.1	.5	0	.09	.35	.63	.98	.02	.13	.22	.63
.1	.1	.3	.02	.19	.06	.33	.01	.11	.05	.30
.1	.3	.3	.03	.19	.08	.38	.01	.12	.05	.32
.1	.5	.3	.05	.22	.13	.48	.01	.11	.06	.32
.5	.1	0	.02	.19	.09	.43	.01	.10	.09	.41
.5	.3	0	.03	.20	.16	.60	.01	.10	.12	.48
.5	.5	0	.09	.33	.44	.87	.02	.12	.18	.57
.5	.1	.3	.02	.19	.05	.31	.00	.11	.04	.29
.5	.3	.3	.03	.21	.07	.35	.01	.10	.05	.29
.5	.5	.3	.08	.31	.09	.38	.01	.10	.05	.28
.9	.1	0	.02	.19	.08	.40	.00	.09	.07	.38
.9	.3	0	.03	.21	.15	.56	.01	.10	.10	.44
.9	.5	0	.08	.31	.33	.77	.02	.10	.15	.52
.9	.1	.3	.02	.19	.05	.30	.00	.10	.04	.27
.9	.3	.3	.02	.19	.06	.32	.00	.10	.04	.27
.9	.5	.3	.04	.21	.07	.30	.01	.10	.03	.24
Median			.03	.20	.09	.42	.01	.10	.07	.35

Table 5

Serially Correlated Shocks and Interest Rate Forecastability

Rule	Δi_{t+2} forecastability		Δi_{t+3} forecastability		$\hat{\rho}_1$ or $\hat{\rho}_2$	
	R^2_L	R^2_U	R^2_L	R^2_U	mean	s.e.
	Panel A. Shocks of Known Persistence					
1	.60	.89	.48	.84	.82	.01
2	.53	.80	.29	.66	.81	.03
	Panel B. Shocks of Unknown Persistence					
1	.05	.24	.04	.22	.37	.06
2	.01	.12	.01	.11	.39	.07

References

[1] Amato, Jeffery, and Thomas Laubach (1999), "The Value of Interest Rate Smoothing: How the Private Sector Helps the Federal Reserve," *Economic Review*, Federal Reserve Bank of Kansas, Third Quarter, 47-64.

[2] Anderson, Gary S., and George Moore (1985), "A Linear Algebraic Procedure for Solving Linear Perfect Foresight Models," *Economics Letters* 17, 247-52.

[3] Balduzzi, Pieluigi, Guiseppe Bertola, and Silverio Foresi (1997), "A Model of Target Changes and the Term Structure of Interest Rates," *Journal of Monetary Economics* 39, 223-249.

[4] Blinder, Alan S. (1986), "More on the Speed of Adjustment in Inventory Models," *Journal of Money, Credit and Banking* 18, 355-365.

[5] Clarida, Richard, Jordi Gali, and Mark Gertler (1998), "Monetary Policy Rules in Practice: Some International Evidence," *European Economic Review* 42, 1033-1067.

[6] Clarida, Richard, Jordi Gali, and Mark Gertler (1999), "The Science of Monetary Policy: A New Keynesian Perspective," *Journal of Economic Literature* 37, 1661-1707.

[7] Clarida, Richard, Jordi Gali, and Mark Gertler(2000), "Monetary Policy Rules and Macroeconomic Stability: Evidence and Some Theory," *Quarterly Journal of Economics* 115, 147-180.

[8] Cook, Timothy and Thomas Hahn (1990), "Interest Rate Expectations and the Slope of the Money Market Yield Curve," *Economic Review*, Federal Reserve Bank of Richmond, Sept/Oct, 3-15

[9] Cukierman, Alex (1996), "Why Does the Fed Smooth Interest Rates?" in *Monetary Policy on the 75th Anniversary of the Federal Reserve System*, edited by M. Belognia, Boston: Kluwer Academic Publishers, 111-147.

[10] Dotsey, Michael, and Christopher Otrok (1995), "The Rational Expectations Hypothesis of the Term Structure, Monetary Policy and Time-Varying Term Premia," *Economic Quarterly,* (Federal Reserve Bank of Richmond), 81(1), 65-81.

[11] Eijffinger, Sylvester, Eric Schaling, and Willem Verhagen (1999), "A Theory of Interest Rate Stepping: Inflation Targeting in a Dynamic Menu Cost Model," working paper, Tilburg University, 1-48.

[12] Estrella, Arturo, and Jeffrey C. Fuhrer (1998), "Dynamic Inconsistencies: Counterfactual Implications of a Class of Rational Expectations Models," Working Paper No. 98-5, Federal Reserve Bank of Boston.

[13] Evans, Charles L. (1998), "Real-Time Taylor Rules and the Federal Funds Futures Market," *Economic Perspectives*, Federal Reserve Bank of Chicago, Third Quarter, 44-55.

[14] Fair, Ray (2000), "Estimated, Calibrated and Optimal Interest Rate Rules," manuscript, Yale University.

[15] Fuhrer, Jefferey C. (1996), "Monetary Policy Shifts and Long-Term Interest Rates," *The Quarterly Journal of Economics*, November, 1183-1209.

[16] Fuhrer, Jeffrey C. (1997), "The (Un)Importance of Forward-Looking Behavior in Price Specifications," *Journal of Money, Credit, and Banking* (August), 338-350.

[17] Fuhrer, Jeffrey C. (2000), "Habit Formation in Consumption and Its Implications for Monetary-Policy Models," *The American Economic Review* 90, 367-90.

[18] Fuhrer, Jeffery C., and George R. Moore (1995), "Monetary Policy Trade-Offs and the Correlation Between Nominal Interest Rates and Real Output," *American Economic Review* 85, 219-239.

[19] Gerlach, Stefan, and Gert Schnabel (2000), "The Taylor Rule and Interest Rates in the EMU Area," *Economics Letters* 67, 165-171.

[20] Goodfriend, M. (1991), "Interest Rates and the Conduct of Monetary Policy," *Carnegie-Rochester Series on Public Policy* 34, 7-30.

[21] Goodfriend, Marvin, and Robert G. King (1997), "The New Neoclassical Synthesis and the Role of Monetary Policy," in *NBER Macroeconomics Annual*, 231-283.

[22] Goodhart, Charles (1997), "Why Do the Monetary Authorities Smooth Interest Rates?" in *European Monetary Policy,* edited by Stefan Collignon, Washington, DC: Pinter, 119-178.

[23] Goodhart, Charles (1999), "Central Bankers and Uncertainty," *Bank of England Quarterly Bulletin* 39(1), 102-115.

[24] Griliches, Zvi (1967). "Distributed Lags: A Survey," *Econometrica* 35, 16-49.

[25] Hall, Alastair, and Robert J. Rosanna (1991), "Estimating the Speed of Adjustment in Partial Adjustment Models," *Journal of Business and Economic Statistics* 9, 441-53.

[26] Jegadeesh, Narasimhan, and George G. Pennacchi (1996), "The Behavior of Interest Rates Implied by the Term Structure of Eurodollar Futures," *Journal of Money, Credit and Banking* 28, 420-446.

[27] Judd, John, and Glenn Rudebusch (1998), "Taylor's Rule and the Fed: 1970-1997," *Economic Review*, Federal Reserve Bank of San Francisco, no. 3, 3-16.

[28] Kozicki, Sharon (1999), "How Useful Are Taylor Rules for Monetary Policy?" *Economic Review*, Federal Reserve Bank of Kansas City, Second Quarter, 5-33.

[29] Kuttner, Kenneth (2000), "Monetary Policy Surprises and Interest Rates: Evidence from the Fed Funds Futures Market," manuscript, Federal Reserve Bank of New York.

[30] Lansing, Kevin (2000), "Learning About a Shift in Trend Output: Implications for Monetary Policy and Inflation," manuscript, Federal Reserve Bank of San Francisco.

[31] Levin, Andrew, Volker Wieland, and John C. Williams (1999), "Robustness of Simple Monetary Policy Rules under Model Uncertainty," in *Monetary Policy Rules*, edited by John B. Taylor, Chicago: Chicago University Press, 263-299.

[32] Lowe, Philip, and Lucy Ellis (1997), "The Smoothing of Official Interest Rates," in *Monetary Policy and Inflation Targeting: Proceedings of a Conference*, Sydney, Australia: Reserve Bank of Australia, 286-312.

[33] Mankiw, N. Gregory, and Jeffrey A. Miron (1986), "The Changing Behavior of the Term Structure of Interest Rates," *The Quarterly Journal of Economics* 101, 211-28.

[34] McCallum, Bennett T., and Edward Nelson (1999), "Nominal Income Targeting in an Open-Economy Optimizing Model," *Journal of Monetary Economics* 43, 553-578.

[35] McManus, Douglas, A., John C. Nankervis, and N.E. Savin (1994), "Multiple Optima and Asymptotic Approximations in the Partial Adjustment Model," *Journal of Econometrics* 62, 91-128.

[36] McNees, Stephen K. (1992), "A Forward-Looking Monetary Policy Reaction Function: Continuity and Change," *New England Economic Review,* November/December, 3-14.

[37] Meyer, Laurence (1999), "The Global Economic Outlook and Challenges Facing Monetary Policy around the World," Remarks at the Annual Dinner of the Society of Business Economists, February 25, available at http://www.federalreserve.gov/boarddocs/speeches.

[38] Mishkin, Frederic S. (1988), "The Information in the Term Structure: Some Further Results," *Journal of Applied Econometrics* 3, 307-314.

[39] Modigliani, Franco, and Robert Schiller (1973), "Inflation, Rational Expectations and the Term Structure of Interest Rates," *Economica* 40, 12-43.

[40] Nelson, Edward (2000), "UK Monetary Policy 1972-1997: A Guide Using Taylor Rules," Bank of England Working paper 120.

[41] Orphanides, Athanasios (1998), "Monetary Policy Rules Based on Real-Time Data," working paper, Federal Reserve Board.

[42] Orphanides, Athanasios, Richard Porter, David Reifschneider, Robert Tetlow, and Frederico Finan (1999), "Errors in the Measurement of the Output Gap and the Design of Monetary Policy," *Journal of Economics and Business* 52, 117-42.

[43] Peersman, Gert, and Frank Smets (1999), "The Taylor Rule: A Useful Monetary Policy Benchmark for the Euro Area?" *International Finance* 2 , 85-116.

[44] Roberts, John M. (1995), "New Keynesian Economics and the Phillips Curve," *Journal of Money and Banking* (November), 975-984.

[45] Rotemberg, Julio, and Michael Woodford (1999), "Interest Rate Rules in an Estimated Sticky Price Model," in *Monetary Policy Rules*, edited by John B. Taylor, Chicago: Chicago University Press, 57-119.

[46] Rudebusch, Glenn (1995), "Federal Reserve Interest Rate Targeting, Rational Expectations, and the Term Structure," *Journal of Monetary Economics* 35, 245-274.

[47] Rudebusch, Glenn D. (1998), "Do Measures of Monetary Policy in a VAR Make Sense?" *International Economic Review* 39, 907-931.

[48] Rudebusch, Glenn D. (1999), "Is the Fed Too Timid? Monetary Policy in an Uncertain World," Working Paper 99-05, Federal Reserve Bank of San Francisco, forthcoming in the *Review of Economics and Statistics.*

[49] Rudebusch, Glenn D. (2000), "Assessing Nominal Income Rules for Monetary Policy with Model and Data Uncertainty," Working Paper 2000-03, Federal Reserve Bank of San Francisco.

[50] Rudebusch, Glenn D., and Lars E.O. Svensson (1999), "Policy Rules for Inflation Targeting," in *Monetary Policy Rules*, edited by John B. Taylor, Chicago: Chicago University Press, 203-246.

[51] Sack, Brian (1998) "Uncertainty, Learning and Gradual Monetary Policy," FEDS working paper 34, Federal Reserve Board.

[52] Sack, Brian (2000), "Does the Fed Act Gradually? A VAR Analysis," *Journal of Monetary Economics* 46, 229-256.

[53] Sack, Brian, and Volker Wieland (2000), "Interest Rate Smoothing and Optimal Monetary Policy: A Review of Recent Empirical Evidence," *Journal of Economics and Business* 52, 205-228.

[54] Smets, Frank (1999), "Output Gap Uncertainty: Does It Matter for the Taylor Rule?" in *Monetary Policy Under Uncertainty*, edited by Benjamin Hunt and Adrian Orr, Wellington, New Zealand: Reserve Bank of New Zealand, 10-29.

[55] Söderström, Ulf (2000), "Should Central Banks Be More Aggressive?" manuscript, Central Bank of Sweden.

[56] Svensson, Lars E.O. (1999a), "Inflation Targeting: Some Extensions," *Scandinavian Journal of Economics* 101, 337-361.

[57] Svensson, Lars E.O. (1999b), "Inflation Targeting as a Monetary Policy Rule," *Journal of Monetary Economics* 43, 607-654.

[58] Taylor, John B. (1993), "Discretion versus Policy Rules in Practice," *Carnegie-Rochester Conference Series on Public Policy* 39, 195-214.

[59] Taylor, John B. (1999), (ed.), *Monetary Policy Rules*, Chicago, IL: Chicago University Press.

[60] Walsh, Carl E. (1998), *Monetary Theory and Policy*, Cambridge MA: MIT Press.

[61] Woodford, Michael (1996), "Control of the Public Debt: A Requirement for Price Stability?" NBER Working Paper No. 5984.

[62] Woodford, Michael (1999), "Optimal Monetary Policy Inertia," *The Manchester School Supplement*, 1-35.

Figure 1
Interest Rate Forecastability and Policy Inertia

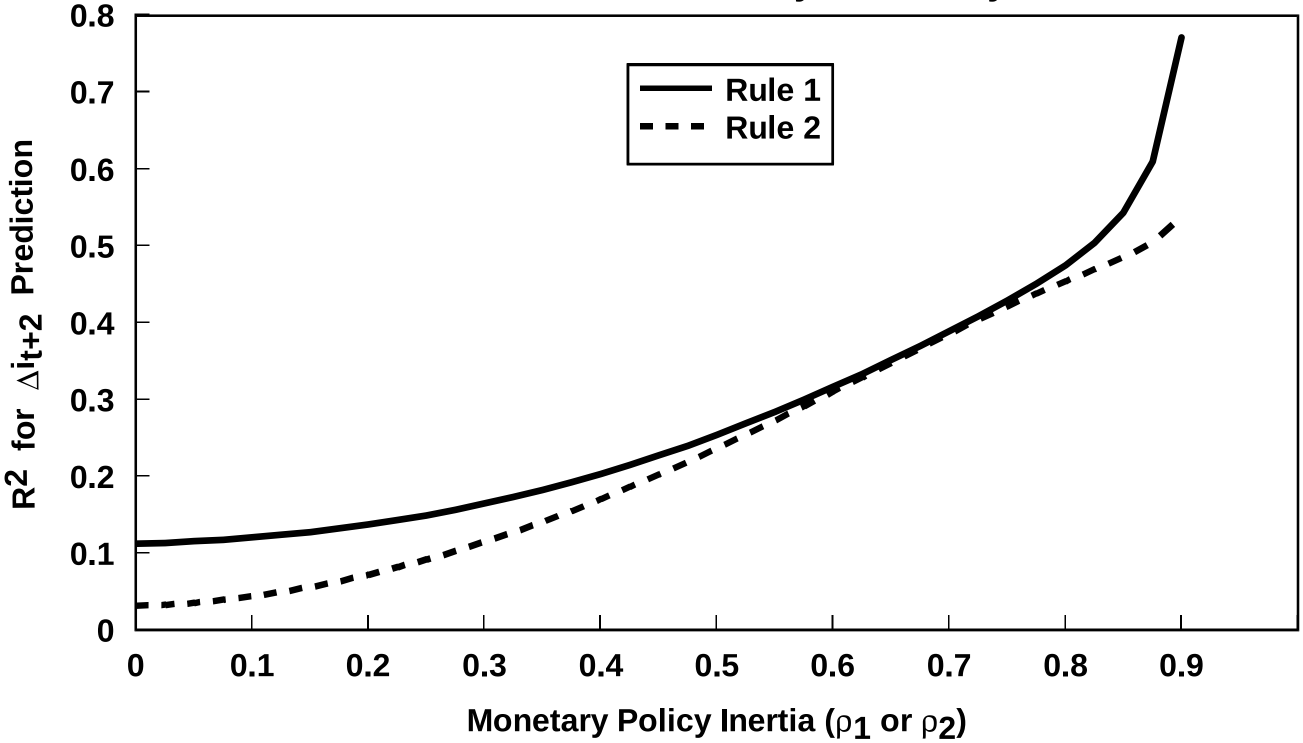

Figure 2
Historical Funds Rate and Fitted Rule 1