Monetary Policy Inertia: Fact or Fiction?

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Abstract

This paper examines interest rate inertia in empirical and optimal monetary policy rules. Estimated policy rules are often interpreted as showing extremely slow policy partial adjustment by central banks. Alternatively, the illusion of such inertia may reflect spuriously omitted persistent influences on policy. Simple policy rule regressions and monetary theory do not appear able to provide a compelling case for or against real-world inertia. However, the yield curve is very informative about expectations of future short rates and the monetary policy rule, and a variety of term structure evidence indicates that the amount of actual policy inertia is very low.

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1. Introduction

In recent years, there has been a clear shift in the focus of monetary policy research. While a decade or so ago, monetary aggregates were often used to model monetary policy, now the most common representation uses a short-term interest rate as the monetary policy instrument, and the literature on how central banks manipulate policy interest rates has grown very rapidly.[1] Especially since the introduction of the now ubiquitous Taylor (1993) rule, many researchers have examined monetary policy rules or reaction functions that relate the policy interest rate to a small set of observables.[2] There has been voluminous research on the theoretical optimal design of monetary policy rules and on the empirical estimation of such rules with historical data. A variety of important issues regarding the normative and positive forms of these rules have been considered—notably, the choice of the relevant argument variables in the rules and the nature of the dynamic adjustment embodied in the rules. This paper will examine the latter issue and broadly characterize the amount of monetary policy inertia or partial adjustment contained in empirical and optimal interest rate rules.

The dynamic adjustment process of monetary policy is a particularly interesting topic because of the lively debate about its fundamental nature. Both sides of the debate acknowledge the existence of smooth cyclical fluctuations in central bank policy rates—as illustrated in Figure 1 for the U.S.[3] The dispute instead involves a disagreement about the underlying nature of this slow adjustment of policy rates. One school of thought, the partial adjustment view, asserts that the persistence of policy rates reflects an inertia that is *intrinsic* or *endogenous* to the central bank. Under this view, there is a long, intentionally drawn-out adjustment of the policy rate in response to economic news. Such partial adjustment implies that the central bank knowingly distributes desired changes in the policy interest rate over an extended period of time; therefore, the smoothness of a policy rate reflects deliberate "interest rate smoothing" or "partial adjustment" or "gradualism" or "inertia" on the part of the central bank. For example, given typical empirical estimates, if a central bank knew it wanted to increase the policy rate by a percentage point, it would only raise it by about 20 basis points in the first three months and by about 60 basis points after one year. That is, there is a very slow convergence of the policy

[1] For some of the intellectual arguments underlying this shift, see Rudebusch and Svensson (2002).

[2] See Svensson (2003) for a discussion of targeting rules as an alternative representation.

[3] Figure 1 shows the quarterly average federal funds rate as the U.S. monetary policy instrument. As has been widely noted (e.g., Rudebusch, 1995), central banks tend to adjust their policy interest rates in sequences of relatively small steps over several years with only rare reversals of direction.

rate to its desired level.

The opposing view to partial adjustment is that the persistence of the policy rate simply reflects the response of the central bank to the slow cyclical fluctuations in the key macroeconomic driving variables of monetary policy, notably inflation and output, which are also illustrated in Figure 1 for the U.S.[4] In this case, the persistence of the policy rate reflects an inertia that is *extrinsic* or *exogenous* to the central bank. Therefore, from this second perspective, the slow adjustment of the policy rate simply reflects the slow accretion of information relevant to the setting of the policy interest rate by policymakers, who then completely adjust the policy rate fairly promptly—within, say, a few months—when confronted with new information.

This disagreement is not just an academic debate but is quite relevant to the conduct of monetary policy.[5] Indeed, policymakers themselves appear unclear about the source of the slow adjustment of policy interest rates. For example, Bernanke (2004) is a proponent of the intrinsic view in which the slow adjustment of the policy rate reflects "partial adjustment and monetary policy inertia." In contrast, Poole (2003) argues that there has been no partial adjustment: "In my view of the world, future policy actions are almost entirely contingent on the arrival of new information... Given information at the time of a meeting, I believe that the standing assumption should be that the policy action at the meeting is expected to position the stance of policy appropriately."

I will review the basic evidence for and against monetary policy inertia in Section 2. The inertial view appears widely supported by estimated monetary policy rules. When such rules are estimated without policy inertia, the residuals indicate significant, persistent deviations of the rule recommendation from the historical policy rate. With the addition of partial adjustment (in the form of a lagged dependent variable), these deviations are greatly reduced. The alternative view, as noted above, is that the deviations represent persistent influences on central-bank behavior that are not captured in a simple Taylor-type rule. These persistent influences may include, for example, responses to financial crises, judgemental adjustments, or differences between real-time and final revised data. Unfortunately, as is well-known in econometrics, at least since Griliches (1967), the two dynamic representations of partial adjustment and persistent

[4] Figure 1 displays the quarterly average federal funds rate, the four-quarter percent change in the price index for personal consumption expenditures excluding food and energy, and the output gap as estimated by the Congressional Budget Office.

[5] This same debate also occurs for other macroeconomic time series. For example, as many have noted, the infrequent adjustment of prices could reflect the inertial nature of price determination—menu costs, etc.—or it could indicate the sluggish economic determinants of completely flexible prices.

omitted variables can be very hard to distinguish in simple single-equation regressions. Indeed, this appears to be the case for the monetary policy rule regressions, especially since there is so much uncertainty about the appropriate arguments of the policy rules.

Therefore, Section 3 turns to theory and examines whether a central bank would want to engage in sluggish partial adjustment from the perspective of optimal monetary policy prescriptions. There are three key rationales for inertial behavior, namely, to reduce interest rate volatility, to exploit the expectational channel for monetary policy, and to respond optimally to data and model uncertainty. While there appears to be some validity to each of these rationales, they do not appear to be able to justify the extremely slow monetary policy inertia suggested by the estimated monetary policy rules.

In contrast to the weak and inconclusive single-equation evidence and theoretical rationales in Sections 2 and 3, a very powerful set of evidence on monetary policy inertia is introduced in Section 4. This evidence is contained in the term structure of interest rates, which can bring a vast amount of information to bear on the appropriate monetary policy rule. Assuming financial market participants understand the policy rule that links short-term interest rates to the realizations of macroeconomic variables, they then will also use that rule in pricing forward interest rates; accordingly, any deviations between expected future short-term rates and expected rule recommendations based on future macroeconomic conditions will be arbitraged away. Therefore, at any point in time, multi-period interest rates, which embody expectations of future short rates, will contain much information about the properties of the monetary policy reaction function. Section 4 presents three different ways to use such yield curve information—predictability regressions, system macro-finance estimates, and event studies based on macroeconomic data surprises. These procedures differ in amount of economic structure imposed and also operate at three different frequencies—quarterly, monthly, and intraday. The resulting consistent set of results from these diverse methodologies appears to provide decisive evidence against the presence of significant monetary policy inertia.

Finally, Section 5 concludes with some suggestions for future research.

2. Gradualism and Inertia in Estimated Policy Rules

An inertial view of monetary policy dynamic adjustment implies that the short-term policy rate is changed at a very sluggish pace, so a central bank would distribute the monetary policy reaction to new economic data over several quarters. The evidence underlying the belief in sluggish adjustment is based on estimated policy rules. These rules follow the standard partial adjustment form: $i_t = (1-\rho)\hat{\imath}_t + \rho i_{t-1}$, where i_t is the level of the policy interest rate set in quarter t, which is a weighted average of the current desired level, $\hat{\imath}_t$, 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—about 20 percent per quarter—of the policy rate to its fundamental determinants. The large coefficient on the lagged dependent variable is widely interpreted as evidence for a "monetary policy inertia" behavior by central banks.[6]

To examine this argument, it is useful to consider a specific case. The most commonly estimated inertial policy rules have been dynamic forms of the Taylor rule. In such rules, the actual interest rate partially adjusts to a desired interest rate that depends on inflation and the output gap; specifically,

$$i_t = (1-\rho)\hat{\imath}_t + \rho i_{t-1} + \xi_t \tag{1}$$

$$\hat{\imath}_t = k + g_\pi \bar{\pi}_t + g_y y_t, \tag{2}$$

where k is a constant incorporating an equilibrium real rate, r^*, and an inflation target, π^*, and g_π and g_y are the central bank response coefficients to (four-quarter) inflation ($\bar{\pi}_t$) and the output gap (y_t).[7]

To provide a benchmark for comparison, first consider an estimated *non-inertial* Taylor rule that assumes $\rho = 0$, as in Taylor (1999) and Yellen (2004). A least squares regression on U.S. data from 1987:Q4 to 2004:Q4 yields

$$\begin{array}{ccccccccccccc} i_t & = & \underset{(.28)}{2.04} & + & \underset{(.09)}{1.39\,\bar{\pi}_t} & + & \underset{(.06)}{.92\,y_t} & + & \xi_t & \equiv & \hat{\imath}_t^{NI} & + & \xi_t, \end{array} \tag{3}$$

$$\sigma_\xi = .97, \quad \bar{R}^2 = .82, \quad DW = .34\ .$$

The monetary policy response coefficients—namely, $g_\pi = 1.39$ for inflation response and $g_y = 0.92$ for output response—are not too far from the 1.5 and 0.5 that Taylor (1993) originally

[6] For example, Clarida, Gali, and Gertler (2000, pp. 157-158) describe their U.S. estimates of various partial adjustment policy rules as follows: ". . . 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." For European estimates, see Gerdesmeier and Roffia (2004) and Sauer and Sturm (2003).

[7] The federal funds rate is a quarterly average rate. Inflation is defined using the price index for personal consumption expenditures excluding food and energy (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., $y_t = 100(Q_t - Q_t^*)/Q_t^*$).

used. The fitted values from this non-inertial Taylor rule regression, which will be denoted $\hat{\imath}_t^{NI}$, are shown as the dotted line in Figure 2 and show a fairly good fit to the actual funds rate—the dark solid line. However, there are some large persistent deviations between the non-inertial rule and the historical funds rate, especially during 1992, 1993, 1999, and 2004 when the actual rate was held below the rule, and during 1991, 1995, and 1996 when the rate was pushed above the rule. Understanding the source of these deviations, as discussed below, is key to interpreting the evidence and arguments for and against policy inertia.

A partial adjustment mechanism is a standard econometric response to such persistent deviations, and a least squares regression for an inertial policy rule on U.S. data from 1987:Q4 to 2004:Q4 yields

$$i_t = .22\, \hat{\imath}_t^I + \underset{(.04)}{.78}\, i_{t-1} + \xi_t, \tag{4}$$

$$\hat{\imath}_t^I = \underset{(.18)}{2.13} + \underset{(.18)}{1.33}\, \bar{\pi}_t + \underset{(.13)}{1.29}\, y_t \tag{5}$$

$$\sigma_\xi = .38, \quad \bar{R}^2 = .97\ .$$

In this regression, the estimated values of the response coefficients are not so different from the non-inertial rule; however, the estimate of the partial adjustment coefficient ($\hat{\rho} = 0.78$) is economically and statistically significant. Such lagged dependence is an extremely robust empirical result in the literature.[8] Indeed, after taking into account the dynamic adjustment in equation (4), the fitted values in the inertial rule—which are shown as the light solid line in Figure 2—match the historical path of the funds rate much more closely than the non-inertial rule. This difference in fit is also apparent in Figure 3, which charts the residuals (ξ_t) from the inertial and non-inertial rules. The mean absolute error for the non-inertial rule of .82 percentage point is almost three times larger than the .29 percentage point error for the inertial rule.

The significance of ρ and the dramatic improvement in R^2 have been widely taken to be convincing evidence of monetary policy inertia. However, Rudebusch (2002a) argues that the monetary policy rule estimates are misleading and provide only the illusion of monetary policy inertia. In particular, if the desired policy interest rate depends on persistent factors other than the current output and inflation in the Taylor rule, then such a misspecification could result in a spurious finding of partial adjustment. Accordingly, based only on these types of policy rule

[8] Similar estimates are discussed by Kozicki (1999) and Rudebusch (2002a) for the U.S., and by Gerdesmeier and Roffia (2004) and Sauer and Sturm (2003) for the euro area.

estimates, it would be very difficult to distinguish whether the Fed's adjustment was sluggish or whether the Fed generally followed the Taylor rule with no policy inertia, but sometimes deviated from the rule for several quarters at a time in response to other factors.

The intuition for this argument is illustrated in Figures 4 and 5. Figure 4 displays the actual funds rate (dark solid line) and the "desired" funds rates from the two rules. The non-inertial rule desired rates are the fitted values $\hat{\imath}_t^{NI}$ from equation (3) (shown as the dotted line), and the inertial rule desired values are the $\hat{\imath}_t^{I}$ from equation (5) (shown as the light solid line). The two desired levels generally move together, so deviations of these desired rates from the actual funds rate are similar across the two rules. Understanding the persistent deviations of the historical interest rate from the two Taylor rule recommendations is key to interpreting the empirical evidence. Under the monetary policy inertia interpretation, these persistent deviations are the result of sluggish central bank responses to output and inflation gaps; that is, the central bank only gradually adjusts the policy rate to the level it would like to set in the absence of some partial adjustment constraint. However, there are several episodes evident in Figure 4 that appear to contradict such an interpretation. For example, at the beginning of 1995, the actual funds rate matched the desired funds rate (as recommended by either rule), but over the rest of that year, the desired funds rate dropped almost 200 basis points, while the actual funds rate jumped 100 basis points. Conversely, after the third quarter of 1998, when the actual rate equaled the desired values, desired rates rose sharply for the next year, while the actual funds rate dropped. Adding a lagged funds rate to the equation will certainly improve the regression fit, but it appears misleading to characterize these episodes as central bank partial adjustment when the actual and desired funds rates moved so dramatically in opposite directions.

The deviations of the two desired funds rates series from the actual funds rate are shown in Figure 5 (namely $i_t - \hat{\imath}_t^{NI}$ as the solid line and $i_t - \hat{\imath}_t^{I}$ as the dotted line). Instead of a partial adjustment explanation for these deviations, an alternative explanation is that the deviations in Figure 5 reflect the incomplete description of monetary policy provided by the Taylor rule. Indeed, it is easy to provide a narrative history of a variety of macroeconomic developments, in addition to estimates of the contemporaneous output gap and inflation, that the Fed appeared to respond to, some of which are indicated in Figure 5.[9] For example, relative to what the Taylor rule would have recommended, a response to the stock market crash may have lowered

[9] The original analysis of Taylor (1993) put forward a description of monetary policy that did not involve interest rate smoothing or partial adjustment. Taylor argued that deviations from the rule during various episodes were an appropriate response to special circumstances. Kozicki (1999) also makes this point.

rates in 1988, and inflation worries—at least, as discussed below, when judged using real-time data—appear to have led to a greater-than-Taylor-rule tightening during 1989. The deviations toward looser monetary policy in 1992 and 1993 have been interpreted as the Fed's response to disruptions in the flow of credit or severe financial headwinds.[10] An inflation scare at the end of 1994—evidenced by a rapid rise in long-term interest rates—preceded a sustained period of tight policy. Another factor that emerged during this period was the remarkable increase in the growth rate of productivity and potential output. At the time, most economists didn't recognize these changes and hence overestimated the degree of utilization in labor and product markets, which likely was reflected in tighter policy. In 1998 and 1999, a worldwide financial crisis following the Russian default and devaluation appears to have played a role in lowering rates.[11] Similarly, there was a rapid easing in response to events of September 11, 2001. Finally, 2003 and 2004 were dominated by fears of deflation, which would likely be reflected in lower rates than a simple Taylor rule would recommend, given potential concerns at the zero lower bound for the policy rate (as discussed in McGough, Rudebusch, and Williams, 2005).

This narrative suggests that some Taylor rule residuals reflect differences between policy judgements made with real-time data and Taylor rule estimations conducted with final revised data—a topic that deserves special attention (see Rudebusch, 1998, 2001, 2002a, b, and Orphanides, 2001, 2003). Figure 6 provides some evidence on the importance of these effects in the U.S. by showing the difference between real-time and current estimates of the output gap, which is denoted $y_{t|t} - y_{t|T}$, and the difference between real-time and current estimates of inflation, which is denoted $\bar{\pi}_{t|t} - \bar{\pi}_{t|T}$.[12] (The output gap revisions end in 1998 because of data confidentiality.) For example, Figure 6 shows that in real time the output gap from 1996 through 1998 was estimated to be about a percentage point higher than it is now (because the estimated level of potential output was lower in real time). This underestimation of the degree

[10] 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."

[11] 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."

[12] The output gap series is Federal Reserve Board staff's real-time estimate—kindly supplied by David Small from the FOMC Secretariat—minus the current (as of 2005) CBO output gap estimate. The inflation series is the real-time 4-quarter GDP deflator inflation rate—obtained from the Federal Reserve Bank of Philadelphia real-time data website—minus the current release.

of macroeconomic slack would be reflected in higher interest rates in real time than a Taylor rule estimated with current data would recommend. Similarly, during 1989, inflation was thought to be running about a half of a percentage point faster than current estimates would indicate, so the actual policy rate would likely be higher than the final-data rule would recommend.

It is possible to provide a rough indication of the importance of the data revisions in accounting for the Taylor rule residuals. The predicted Taylor rule residuals based on the real-time to final data revisions can be constructed under the assumption that the Fed followed the estimated non-inertial rule (3) in real time. Specifically, if the Fed used the estimated non-inertial Taylor rule coefficients to conduct policy, so $i_t = 2.04 + 1.39\bar{\pi}_{t|t} + .92y_{t|t}$, then the predicted residuals in equation (3) would equal $1.39(\bar{\pi}_{t|t} - \bar{\pi}_{t|T}) + .92(y_{t|t} - y_{t|T})$. These constructed residuals predicted by the data revisions are shown in Figure 7 along with the non-inertial rule residuals from equation (3). The fairly close correlation between the predicted residuals from real-time data revisions and the actual residuals from the estimated non-inertial rule suggests that a substantial amount of the deviations of the actual rate from the rule estimated with the current vintage of data can be accounted for by the reactions to real-time data and not to central bank partial adjustment.[13]

However, while real-time data revisions are undoubtedly part of the story, it is unlikely, as described in Figure 5, that the Fed follows a Taylor rule in real-time. Instead, like other central banks, it reacts in a less simplistic fashion to a wide variety of macroeconomic developments; that is, the alternative to partial adjustment is the misspecification of the Taylor rule. This omitted variables view of the non-inertial Taylor rule residuals is supported by much contemporaneous press coverage and the narrative policy record. Still, it would be more satisfying to be able to provide econometric evidence distinguishing between the partial adjustment and omitted variables interpretations of the policy rule estimates.

Unfortunately, Rudebusch (2002a) argues that conclusive evidence from simple policy rule estimates on the extent of inertia is inherently difficult to obtain. For example, suppose the non-inertial rule deviations were simply modeled as a first-order autoregressive process. Then, instead of the inertial model of central bank behavior in equations (1) and (2), we would obtain

[13] Lansing (2002) provides a careful simulation study that demonstrates the potential effectiveness of such real-time output gap errors to account for spurious evidence of policy inertia. Also, see Mehra (2001) and Apel and Jansson (2005) in the U.S., and Sauer and Sturm (2003) for the euro area. In addition, given the large inflation policy rule response coefficient, inflation data revisions should not be ignored.

the serially correlated shock model:

$$i_t = k + g_\pi \bar{\pi}_t + g_y y_t + \xi_t \tag{6}$$

$$\xi_t = \rho^e \xi_{t-1} + \omega_t. \tag{7}$$

These shocks would represent the persistent factors—credit crunches, financial crises, etc.—that central banks respond to. The salient question is whether it is possible to distinguish between the model (1) and (2) and model (6) and (7). Rudebusch (2002a) estimates a single equation that nests the inertial and serially correlated shocks rules and finds that the evidence distinguishing these two rules appears fragile to even modest changes in the sample period. His argument draws on a large literature in econometrics showing that estimates of partial adjustment models commonly indicate an unrealistically slow adjustment—whether applied to inventory behavior (Blinder, 1986) or money demand (Goodfriend, 1985).[14] In particular, a standard policy rule with slow partial adjustment and no serial correlation in the errors will be difficult to distinguish empirically from a policy rule that has immediate policy adjustment but highly serially correlated shocks. The choice between these two dynamic structures, which depends crucially on separating the influences of contemporaneous and lagged regressors, is especially difficult to untangle for empirical monetary policy rules for several reasons (also see Carare and Tchaidze, 2005). First, the arguments of the rules—four-quarter inflation and the output gap—are highly serially correlated, so distinguishing the effect of, say, $\bar{\pi}_t$ from $\bar{\pi}_{t-1}$ is not easy. Second, the arguments of the rules are not exogenous. Third, only short data samples of plausibly consistent rule behavior are available with a limited amount of business cycle variation in output and inflation. Fourth, there is some uncertainty about the appropriate arguments of the historical policy rule. Finally, as noted above, the actual interest rates are set on the basis of real-time data on output and inflation, which can also make it difficult to determine the correct dynamics.

There been have several interesting extensions of the analysis in Rudebusch (2002a). English, Nelson, and Sack (2003) and Gerlach-Kristen (2004) provide two slightly different tests of the inertial and serially correlated shock interpretations that, unlike in Rudebusch (2002a), allow for both partial adjustment and serially correlated shocks to be jointly present. These authors find evidence that both features are significant elements in the data; therefore, the standard

[14] There is a large literature that argues that partial adjustment models are difficult to identify and estimate empirically in the presence of serially correlated shocks (e.g., Griliches, 1967, Hall and Rosanna, 1991, and McManus, Nankervis, and Savin, 1994).

policy rule estimates in (4) and (5) are omitting important persistent factors (similar results for the euro area are provided by Castelnuovo, 2005). However, considerable uncertainty remains, as illustrated by the insightful small-sample calculations conducted by English, Nelson, and Sack (2003). They investigate how much of the deviations of the actual rate from the desired rate (the $i_t - \hat{i}_t^I$ in Figure 5) can be accounted for by partial adjustment. They find that a 95 percent confidence interval stretches from 8 percent to 88 percent; therefore, it is difficult to ascertain how economically important partial adjustment is. Furthermore, this wide range of uncertainty is only for a particular rule specification and estimation sample, so it ignores the broader uncertainty noted above.

Furthermore, the assumption that the persistent omitted variable is an AR(1) appears to be a gross simplification that may bias the results and boost the evidence for partial adjustment. Indeed, the narrative history summarized in Figure 5 suggests a more subtle reaction function than can be captured by equations (6) and (7). A few have tried to augment the estimated Taylor rule with other variables in order to capture directly the omitted persistent influences on policy that spuriously induce the appearance of policy inertia. For example, Gerlach-Kristen (2004) and Driffill et al. (2005) find evidence that proxies for financial stability concerns, such as a private-public credit spread, have explanatory power in the Taylor rule. Also, expectations appear to play an important role in tempering the policy response to current readings on output and inflation, and Mehra (2001) suggests that expectations of future inflation—and, in particular, inflation scares in the bond market—are an important consideration for policy, which when omitted will appear as policy inertia. Finally, as shown by Trehan and Wu (2003), ignoring a true time-varying equilibrium real rate (r_t^*) can lead to finding policy inertia when there is really none.[15]

Overall, however, the literature on augmenting the Taylor rule with the important determinants of policy other than current output and inflation appears to be incomplete at best. Svensson (2003) argues that such an endeavor is doomed to failure because the missing elements are so completely judgemental in nature that they cannot be captured by an instrument rule such as the Taylor rule:

> Monetary policy by the world's more advanced central banks these days is at least as

[15] In Europe, Gerlach and Schnabel (2000) find that a Taylor rule fits well without partial adjustment but with dummies for the period 1992:3-1993:3 to control for intra-European exchange market pressures. Gerlach-Kristen (2003) finds that the long rate is significant in a euro-area Taylor rule, while Gerdesmeier and Roffia (2004) recommend inclusion of a money growth gap.

> optimizing and forward-looking as the behavior of the most rational private agents. I find it strange that a large part of the literature on monetary policy still prefers to represent central bank behavior with the help of mechanical instrument rules. [p. 429] . . . Whereas simple instrument rules, like variants of the Taylor rule, may to some extent serve as very rough benchmarks for good monetary policy, they are very *incomplete* rules, because they don't specify when the central bank should or should not deviate from the simple instrument rule. Such deviations, by discretion and judgement, have been and will be frequent . . . [p. 467]

3. Rationales for Sluggish Adjustment by Central Banks

The discussion above indicates that, given the distinct possibility of omitted persistent variables from the monetary policy rules, the usual single-equation evidence from estimated policy reaction functions is inconclusive regarding the empirical importance of policy inertia. In this section, we take a different tack and examine the normative case for interest rate smoothing. Presumably, if theory can provide a fairly compelling rationale for the existence of inertia as a feature of optimal monetary policy, then the case for real-world partial adjustment would be strengthened. Therefore, in this section, I consider the empirical relevance the three most important explanations for why central banks might find partial adjustment attractive.

3.1. Gradualism and volatility reduction

One consequence of policy inertia is to produce interest rates that are less volatile than would be suggested by the determinants of policy. As the speed of adjustment coefficient ρ increases, the variances of the level and changes in the policy instrument decline. Therefore, an obvious rationale for policy gradualism would be some desire on the part of the central bank to reduce the volatility in interest rates and, more generally, in asset prices. Such a desire can be modeled directly in the central bank's loss function, and then, together with a model of the economy, the optimal ρ coefficient can be calculated for an optimal simple Taylor rule (as in, for example, Rudebusch and Svensson, 1999). If the optimal monetary policy partial adjustment coefficient matched the high empirical estimates of ρ, then those estimates would have some greater credence.

The most common way to model a desire for smooth interest rates is to specify a loss function in which the central bank minimizes a weighted sum of the squared inflation gap, the squared output gap, and changes in the policy rate (see Rudebusch and Svensson, 1999, and Clarida, Gali, and Gertler, 1999):

$$L_t = 1/2[(\bar{\pi}_t - \pi^*)^2 + \lambda y_t^2 + \nu_{\Delta i}(\Delta i_t)^2], \tag{8}$$

where $\Delta i_t = i_t - i_{t-1.}$ The parameters $\lambda \geq 0$ and $\nu_{\Delta i} \geq 0$ are the relative weights on output and interest rate stabilization with respect to inflation stabilization. The intertemporal loss function in quarter t is the discounted sum of the expected future per quarter losses,

$$\mathrm{E}_t \sum_{\tau=0}^{\infty} \delta^\tau L_{t+\tau,} \tag{9}$$

with a discount factor δ $(0 < \delta < 1)$. For $\delta = 1$, this loss function can be represented by the unconditional mean of the period loss function (Rudebusch and Svensson, 1999)

$$\mathrm{E}\left[L_t\right] = \mathrm{Var}\left[\bar{\pi}_t - \pi^*\right] + \lambda \mathrm{Var}[y_t] + \nu_{\Delta i}\mathrm{Var}\left[\Delta i_t\right], \tag{10}$$

which equals the weighted sum of the unconditional variances of the three goal variables and is the standard loss function in the literature.[16]

The presence of an interest rate smoothing motive in the loss function has some superficial plausibility, especially in light of the literature that analyzes changes in policy interest rates on a day-by-day basis. In the U.S. (e.g., Goodfriend, 1991, and Rudebusch, 1995) and many other countries (e.g., Goodhart, 1997, and Lowe and Ellis, 1997), central banks generally make changes in the policy rate at discrete intervals and in discrete amounts. Rudebusch (1995, p. 264), for example, describes a short-term interest rate smoothing in which the Fed adjusts interest rates ". . . in limited amounts . . . over the course of several weeks with gradual increases or decreases (but not both) . . . ." Such short-term partial adjustment of the funds rate involves cutting the policy rate by two 25-basis-point moves in fairly quick succession, rather than reducing the rate just once by 50 basis points.[17] This smoothing likely reflects various institutional rigidities, such as a fixed monthly meeting schedule and perhaps certain sociological and political factors.[18] However, short-term partial adjustment within a quarter is essentially independent of whether there is monetary policy inertia over the course of several *quarters*, and this latter issue is the one that is relevant to the empirical monetary policy rules. Indeed, as noted in Rudebusch

[16] However, the choice of δ is not innocuous. As shown in Dennis (2005), greater discounting leads to less concern about the future and less interest rate smoothing.

[17] This behavior has become less prevalent in the U.S. as the frequency of inter-meeting moves has declined.

[18] At a single meeting, 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).

(2002a), given their disparate time frames, the two types of partial adjustment are in fact largely independent, so a central bank could conduct either one without the other. Indeed, if the underlying rationale for reducing interest rate volatility is to reduce instability in financial markets (as described by, for example, Goodfriend, 1991, Rudebusch, 1995, Cukierman, 1996, Lowe and Ellis, 1997), then not wanting to move the policy rate by 50 basis points on a particular day is very different from not wanting to move it by 50 basis points on a quarterly average basis.

This issue is highlighted in trying to specify the weight $\nu_{\Delta i}$ on quarterly interest rate volatility relative to the variability of the output and inflation gaps. If λ and $\nu_{\Delta i}$ are both set equal to 1, then the loss function equally penalizes a 1 percent output gap, a 1 percentage point inflation gap, and a 1 percentage point quarterly change in the funds rate. This penalty on interest rate volatility appears to be implausibly high, given the overwhelming emphasis among central banks on the first two objectives relative to the third (e.g., the "dual mandate" in the United States). Also, in practice, central banks have at times implemented large changes in policy rates, which contradicts the notion of a significant penalty. Perhaps the most extreme example occurred in September 1992, when the Swedish central bank raised its policy rate from 20 percent to 500 percent in one week in an attempt to maintain a fixed exchange rate. However, during the 1979-82 monetary experiment, the U.S. also had much greater interest rate volatility which did not appear to carry, on its own, large costs. In the academic literature, $\nu_{\Delta i}$ is often set equal to equal to 0.5 or 0.1. These loss functions equally penalize a 1 percent output gap, a 1 percentage point inflation gap, and a 1.41 or 3.16 percentage point quarterly change in the funds rate. Such weights still seem at the high end of the plausible range of penalties to reduce volatility, especially in a world with a wide variety of financial market instruments that allow for hedging against interest rate volatility.

Finally, I should note that even the specification of the interest smoothing objective in the loss function is unclear. Svensson (2003) notes that if the motive in reducing interest rate volatility is to avoid financial instability, then the loss function should be specified to minimize the *surprise* in the policy rate:

$$\mathrm{E}\left[L_t\right] = \mathrm{Var}\left[\bar{\pi}_t - \pi^*\right] + \lambda \mathrm{Var}[y_t] + \nu_{\mathrm{E}i} \mathrm{Var}\left[\mathrm{E}_{t-1}\left[i_t\right] - i_t\right], \tag{11}$$

where $\nu_{\mathrm{E}i} \geq 0$ is the relative weight on policy rate surprises. A third specification, advocated by Woodford (1999), penalizes the variability in the *level* of the policy rate:

$$\mathrm{E}\left[L_t\right] = \mathrm{Var}\left[\bar{\pi}_t - \pi^*\right] + \lambda \mathrm{Var}[y_t] + \nu_i \mathrm{Var}\left[i_t - r^* - \pi^*\right], \tag{12}$$

where $\nu_i \geq 0$ weight deviations of the nominal rate from a neutral level.[19]

On its own, motivating a large partial adjustment coefficient through a central bank loss function desire for interest rate smoothing appears unrealistic (e.g., Svensson 2003). This is particularly true in a model with no explicit forward-looking expectational terms, as in Rudebusch and Svensson (1999), where an optimal ρ in a dynamic Taylor rule of greater than .2 or .3 is difficult to obtain. However, results can be very different in forward-looking models, which are considered in the next subsection.

3.2. Central Bank Inertia as a Lever on Expectations

The most passionate advocates for optimal monetary policy partial adjustment base their case on the ability of such inertia to allow the central bank to influence the current state of the economy by promising future actions; that is, sluggish adjustment can be a lever to help move and manage expectations. In particular, 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, 2003, Rotemberg and Woodford, 1999, Levin, Wieland, and Williams, 1999, and Sack and Wieland, 2000). 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.[20]

This argument can be elucidated and assessed within a simple expectational model. Rudebusch (2002a, 2005) describes an empirical version of the New Keynesian model[21] suitable for quarterly data, where inflation and output are determined by future expectations and lags on the past:

$$\pi_t = \mu_\pi E_{t-1}\bar{\pi}_{t+3} + (1-\mu_\pi)\Sigma_{j=1}^{4}\alpha_{\pi j}\pi_{t-j} + \alpha_y y_{t-1} + \varepsilon_t, \tag{13}$$

$$y_t = \mu_y E_{t-1}y_{t+1} + (1-\mu_y)\Sigma_{j=1}^{2}\beta_{yj}y_{t-j} - \beta_r(r_{t-1} - r^*) + \eta_t, \tag{14}$$

where $E_{t-1}\bar{\pi}_{t+3}$ represents the expectation of average inflation over the next year and $E_{t-1}y_{t+1}$ represents the expectation of period $t+1$ output conditional on a time $t-1$ information set.

[19] Woodford (1999, 2003) argues that smaller interest rate fluctuations reduce the likelihood of reaching the zero bound on nominal interest rates and the associated adverse effects on macroeconomic stability; however, with a properly specified model, such concerns should be captured in the output and inflation stabilization concerns in the loss function. Woodford (2003) also tries to motivate this specification of the loss function by appealing to the transactions frictions underlying money demand (so-called "shoe-leather" costs).

[20] This argument can be thought of as a special case of the more general rational that i_{t-1} is likely an important state variable given the dynamic structure of the economy, so the optimal instrument rule would include a response to it (e.g., Rudebusch and Svensson, 1999).

[21] Much of the appeal of the new Keynesian model lies in its foundations in a dynamic general equilibrium model with nominal price rigidities; see Woodford (2003) and Walsh (2003).

The real rate relevant for output, r_{t-1}, is defined as a weighted combination of an ex ante 1-year rate and an ex post 1-year rate:

$$r_{t-1} = \mu_r(E_{t-1}\bar{\imath}_{t+3} - E_{t-1}\bar{\pi}_{t+4}) + (1-\mu_r)(\bar{\imath}_{t-1} - \bar{\pi}_{t-1}), \tag{15}$$

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 consideration of a wide range of explicit forward-looking behavior. 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 (2001), which has had some success in approximating the time series data in the manner of a small estimated VAR (see Estrella and Fuhrer, 2002, and Fuhrer and Rudebusch, 2004). However, estimated forward-looking models also have had some success in fitting the data, as in Rotemberg and Woodford (1999) and Fuhrer (2000). The analysis below takes an eclectic view and conditions on a wide range of possible values for μ_π, μ_y, and μ_r.[22]

Table 1 summarizes the optimal amount of monetary policy inertia for various models and loss functions. The table displays the lag coefficients ρ from the *optimal* versions of the inertial Taylor rule in equations (1) and (2) across models with a range of forward-looking behavior and using the three different loss functions in equations (10), (11), and (12). For each loss function, the weight on the interest rate smoothing ($\nu_{\Delta i}$, $\nu_{\mathrm{E}i}$, or ν_i)is set equal to .5 or .1, while $\lambda = 1$.[23] Clearly in Table 1, a large range of optimal lag coefficients—between -.6 and 1.0—can be rationalized for some combination of model and loss function. Most interesting, however, is how the expectational channel can magnify even a small cost of interest rate fluctuations in the central bank loss function to produce a sizable partial adjustment coefficient in the policy rule. Also, note that the degree of optimal monetary policy inertia varies most strongly with the value of μ_r, which determines the degree to which interest rate expectations are forward-looking. Such variation is consistent with the interpretation of Woodford (1999, 2003) and Levin, Wieland, and Williams (1999) that policy inertia is optimal when it alters expectations of future interest rates that are also important determinants of current demand.

While an expectational channel for optimal monetary policy inertia is valid in principle, it seems unlikely that such a channel is responsible for empirical monetary policy inertia because

[22] 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 of Rudebusch (2002a).

[23] The results in Table 1 are obtained by numerically minimizing the loss function over the parameters g_π, g_y, and ρ in the model. The results are obtained using the "AIM" algorithm (Anderson and Moore, 1985) available at http://www.federalreserve.gov/pubs/oss/oss4/aimindex.html.

its underlying assumption of a perfectly credible policy rule seems so unlikely historically. That is, even if economic agents were sufficiently forward-looking (which is a separate unresolved issue), the monetary policy rule must also be assumed to be perfectly credible, so agents know the rule and assume (correctly) that it will be followed. This seems an unlikely description even for the relatively homogeneous 1987 through 2004 U.S. sample period underlying the above inertial policy rule estimates. In practice, there is no commitment technology, and the Fed has not been able to exercise the requisite control over expectations that would be used to make the expectational channel relevant (as evidenced most recently by Fed Chairman Greenspan's declaration in February 2005 that the low level of bond yields was an inexplicable "conundrum").

3.3. Uncertainty and Partial Adjustment

Uncertainty is the third general rationale often used to motivate optimal monetary policy inertia. The intuition appears clear: Uncertainty breeds caution, and caution suggests a gradual adjustment of the policy rate. As noted by Bernanke (2004), "Because policymakers cannot be sure about the underlying structure of the economy or the effects that their actions will have on economic outcomes, and because new information about the economic situation arrives continually, the case for policymakers to move slowly and cautiously when changing rates seems intuitive." However, the implication that greater uncertainty produces greater inertia is not a general theoretical result, and the empirical evidence for this proposition appears weak as well.

Because economic data can be quite noisy, policymakers inevitably operate with imperfect knowledge about the current state of the economy. In addition, it may be the case that the noisier the economic data are, the less aggressive policymakers should be in responding to it (Rudebusch, 2001, and Orphanides, 2003).[24] However, in practice, as noted by Rudebusch (2001) any such inducement toward timidity (that is, a low g_π and g_y) appears fairly modest and does not necessarily translate into greater sluggishness (that is, a high ρ). It should also be noted that a delayed policy response in the presence of uncertainty does not necessarily represent policy "partial adjustment" (in the sense of a policymaker's gradual adjustment of the policy rate to a desired or planned level). For example, the lagged Taylor rule response to prior quarterly inflation rates π_{t-1}, π_{t-2}, and π_{t-3} (through the annual inflation term $\bar{\pi}_t$)

[24] The general certainty-equivalence guideline is that optimal policy requires the same response under both partial and full information about the state of the economy. However, as discussed in Rudebusch (2001), the use of simple rules and inefficient output gap estimates are two relevant exceptions for this analysis.

represents not partial adjustment but an information filtering mechanism by the central bank in recognition of the uncertainty and noise in quarterly inflation rates. More generally, a simple information smoothing model lets the central bank react immediately to developments in an underlying measure of inflation

$$i_t = \beta\pi_t^u, \tag{16}$$

where underlying inflation is obtained via a linear updating filter on past actual inflation

$$\pi_t^u = \varphi\pi_{t-1}^u + (1-\varphi)\pi_t. \tag{17}$$

The resulting representation is observationally equivalent to the partial adjustment model,

$$i_t = (1-\varphi)\beta\pi_t + \varphi i_{t-1}, \tag{18}$$

but the information smoothing and partial adjustment models are conceptually distinct (as noted early on by Waud, 1968). Of course, in the presence of noisy data some filtering may be optimal; however, the estimated lag coefficient in equation (4) would imply $\varphi = .22$, which would be an extreme discounting of the current observation of annual inflation and the output gap in the Taylor rule.

Uncertainty about the model provides another rationale for caution. Indeed, ever since the classic Brainard (1967) analysis, uncertainty about the quantitative impact of policy and the dynamics of the economy has been widely cited as a rationale for damped policy action. However, in the general case, as Chow (1975, chapter 10) makes clear, *almost nothing can be said even qualitatively* about how the optimal rule under model uncertainty changes relative to the optimal rule under certainty. For example, the optimal policy response parameters are not necessarily reduced in the presence of uncertainty about several parameters. Thus, quantitative answers are required. Rudebusch (2001) provides some simple but instructive evidence that suggests that parameter uncertainty is not responsible for policy inertia. The policymaker is assumed to face an economy like (13), (14) and (15) *on average* (with μ_π, μ_y, and μ_r set equal to zero), but in any given quarter, the coefficients may take on a random value. These parameter shifts occur every quarter or every few years. The policymaker has to choose the g_π, g_y, and ρ parameters of the inertial Taylor rule (1) and (2), so that the loss function (15) is minimized. After allowing for uncertainty about all of the coefficients of the model, the optimal partial adjustment coefficient actually falls a bit.[25]

[25] This conclusion accords with much research on parameter uncertainty. Notably, in Estrella and Mishkin

Overall, although perhaps intuitive, the argument that uncertainty could account for the very gradual persistence in the data remains unproven.

4. Term Structure Evidence on Inertial Policy Rules

To summarize the discussion so far, the single-equation estimation of policy rules has yielded inconclusive results regarding the existence of policy inertia, and the theoretical case for substantial interest-rate smoothing appears unconvincing as well. To make some progress, this section turns to a vast and rich set of information about central bank reaction functions: the yield curve of interest rates. The yield curve contains such information because if financial market participants understand the policy rule that links short-term interest rates to the realizations of macroeconomic variables, then they will also use that rule in forming expectations of future short-term interest rates, which will be priced into long-term bonds.[26] In particular, any deviations from the policy rule embedded in expected future short-term rates and expected macroeconomic conditions would be arbitraged away. Therefore, at any point in time, multiperiod interest rates, which embody expectations of future short rates, contain much information about the properties of the reaction function. In this section, I outline three different methods by which this information can be extracted to inform the debate on policy inertia. These methods differ primarily by the amount of economic structure imposed and by the frequency of data employed.

4.1. Interest rate predictability at a quarterly frequency

Policy inertia has important implications for interest rate forecastability: in brief, the greater the delayed adjustment of the policy rate in reaction to current information, the greater the amount of forecastable future variation. Intuitively, if the funds rate typically is adjusted 20 percent toward its desired target in a given quarter, then the remaining 80 percent of the adjustment should be expected to occur in future quarters. Furthermore, assuming financial markets understand the inertial nature of monetary policy, they should anticipate the future partial adjustment of the funds rate and incorporate it into the pricing of longer-term maturities.

(1999), Peersman and Smets (1999), Shuetrim and Thompson (1999), and Tetlow and von zur Muehlen (2001), there is no significant attenuation of the rule parameters. Some attenuation is found in Sack (2000), Salmon and Martin (1999), and Söderström (1999).

[26] Note that the assumption is not one of credibility and commitment as in subsection 3.2, but one of transparency and learnability.

Rudebusch (2002a) shows that this general intuition is true in a wide variety of macroeconomic models. The amount of such forecastable variation in interest rates can be measured via a standard term structure regression at a quarterly frequency such as

$$\Delta i_{t+2} = \delta + \gamma E_t(\Delta i_{t+2}) + \psi_t. \tag{19}$$

This equation regresses the realized change in the policy rate in quarter $t+2$ (i.e., $\Delta i_{t+2} = i_{t+2} - i_{t+1}$) on the change that was expected in quarter t. Under rational expectations, this interest rate forecasting regression would yield in the limit an estimate of $\hat{\delta} = 0$ and $\hat{\gamma} = 1$. However, for assessing the forecastable variation in the interest rate and hence the degree of monetary policy inertia, the statistic of particular interest is the R^2 of this regression, which provides a natural measure of forecastability.

The theoretical relationship between the forecastable variation in the interest rate, as measured by the R^2 of the above prediction equation, and quarterly policy inertia, as measured by the ρ in the Taylor rule (1) and (2), is illustrated in Figure 8. This figure graphs the implied (population) value of the R^2 of the regression (19) as a function of ρ for a representative case of the model described in Section 3, namely, with $\mu_\pi = .3$, $\mu_r = .5$, and $\mu_y = 0$.[27] Note that even for the non-inertial policy rules there is some predictable future movement in interest rates (with $R^2 = .10$ when $\rho = 0$), since there are predictable changes two quarters ahead in the output gap and in the four-quarter inflation rate, which partly determine future changes in interest rates. Even though the output gap and inflation are highly persistent in levels, the associated slow mean reversion implies only a modest predictability of future quarterly *changes* in these series and the desired funds rate. However, as ρ increases, the amount of predictable variation in Δi_{t+2} also increases, with an R^2 value of .45 at $\rho = 0.8$.

Rudebusch (2002a) shows that this theoretical relationship between partial adjustment and predictability is robust across a vide variety of models (and for forecast-baed policy rules as well). This relationship can be empirically assessed by examining the extent of forecastable future movements in the policy interest rate in the data. Specifically, if policy is highly inertial, as the single-equation reaction functions suggest, then financial markets should anticipate the future partial adjustment of the funds rate. In that case, a regression of actual changes in the

[27] Also, g_π and g_y are set equal to 1.5 and 0.8, respectively. As in Table 1, 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.

funds rate on predicted changes embedded in the yield curve should provide a good explanatory fit and a fairly high R^2. In fact, researchers have found the opposite. They have estimated a variety of interest rate forecasting regressions and, using financial market expectations, have found little predictive information at quarterly frequencies with R^2's very close to zero.[28] For example, Rudebusch (2002a) shows that eurodollar futures from 1988 to 2000 have very little ability to predict the quarterly change in the funds rate two quarters ahead. The R^2 of such a regression is .11, which from Figure 8 suggests that ρ is probably close to zero.

This lack of predictive ability is well illustrated by the most recent episode of monetary policy easing. Figure 9 gives the actual funds rate target and various expected funds rate paths as of the middle of each quarter based on fed funds futures. Under quarterly policy inertia, the long sequence of target changes in the same direction in 2001 would be viewed as a set of gradual partial adjustments to a low desired rate. However, although the funds rate gradually fell in 2001, market participants actually anticipated few of these declines at a 6- to 9-month horizon, as they would have if policy inertia were in place. Instead, markets assumed at each point in time that the Fed had adjusted the funds rate down to just about where it wanted the funds rate to remain based on current information available. Under this interpretation, the long sequence of declines is the result of a series of fairly prompt responses to new information that turned progressively more pessimistic. That is, the presence of quarterly partial adjustment or policy inertia is contradicted by the lack of forecastability of changes in the funds rate.

Rudebusch (2002a) used a variety of structural models to show that the large estimated lag coefficients in the empirical inertial policy rules provided were inconsistent with the very low interest rate forecastability in the term structure of interest rates.[29] Söderlind, Söderström, and Vredin (2005) examine whether this is true using a vector auto-regressive (VAR) model and using survey data or macroeconomic forecasts. They also find that both sets of evidence are inconsistent with the inertial Taylor rule.

4.2. Term structure model estimation at a monthly frequency

While the evidence in section 4.1 on the predictability of interest rates is quite intuitive, it is somewhat indirect; that is, the absence of policy inertia is inferred from the lack of interest

[28] See, for example, Mankiw and Miron (1986), and Rudebusch (1995).

[29] Rudebusch (2002a) also shows that the rule with highly serially correlated errors, equations (6) and (7), does not imply such forecastability.

rate predictability evident in financial markets. More direct estimates of the degree of interest rate smoothing would perhaps be more compelling, and this section considers direct estimates of ρ. However, these estimates of ρ differ from the single-equation ones given in Section 2 because they are obtained in a complete system that combines key macroeconomic equations and information from the yield curve. The particular structure employed is from Rudebusch and Wu (2004). Their analysis uses monthly data to formally estimate a model that combines a fairly standard macroeconomic model with an off-the-shelf no-arbitrage finance representation from the empirical bond pricing literature. Again, it is the incorporation of yield curve information that allows precise inference about the absence of monetary policy inertia.

Almost all movements in the yield curve can be captured in a no-arbitrage framework in which yields are linear functions of a few unobservable or latent factors (e.g., Duffie and Kan, 1996, and Dai and Singleton, 2000). The Rudebusch-Wu macro-finance model employs such a framework: specifically, it features a constant factor volatility with state-dependent risk pricing of volatility, which implies conditionally heteroskedastic risk premiums. The one-month short rate is the sum of a constant and two unobserved term structure factors

$$i_t = \delta_0 + L_t^m + S_t^m, \tag{20}$$

where L_t^m and S_t^m are termed level and slope factors. The dynamics of these latent factors are given by

$$\begin{aligned} L_t^m &= \rho_L L_{t-1}^m + (1-\rho_L)\pi_t + \varepsilon_{L,t} && (21)\\ S_t^m &= \rho_S S_{t-1}^m + (1-\rho_S)[g_y y_t + g_\pi(\pi_t - L_t^m)] + u_{S,t} && (22)\\ u_{S,t} &= \rho_u u_{S,t-1} + \varepsilon_{S,t}, && (23) \end{aligned}$$

where π_t and y_t are inflation and the output gap.[30] These equations can be given a Taylor rule interpretation, with the factor L_t^m interpreted as the inflation rate targeted by the central bank, as perceived by private agents. Private agents slowly modify their views about L_t^m as actual inflation changes, so L_t^m is associated with an interim or medium-term inflation target (as in Bomfim and Rudebusch, 2000) with associated underlying inflation expectations over the next 2 to 5 years. The slope factor S_t^m captures the central bank's dual mandate to stabilize the real economy and keep inflation close to its target level. In addition, the dynamics of S_t^m allow for both partial adjustment and serially correlated shocks. If $\rho_u = 0$, the dynamics of S_t^m arise

[30] In this substitution with monthly data, π_t is the 12-month inflation rate and y_t is capacity utilization.

from monetary policy partial adjustment, as in an inertial Taylor rule. Conversely, if $\rho_S = 0$, the dynamics reflect the Fed's reaction to autocorrelated information or events not captured by output and inflation, as in the Taylor rule with AR(1) shocks.

Appended to the above equations is a small macroeconomic model of inflation and output suitable for estimation with monthly data, which also has some appeal as a New Keynesian specification:

$$\pi_t = \mu_\pi L_t^m + (1-\mu_\pi)[\alpha_{\pi_1}\pi_{t-1} + \alpha_{\pi_2}\pi_{t-2}] + \alpha_y y_{t-1} + \varepsilon_{\pi,t}. \tag{24}$$

$$y_t = \mu_y E_t y_{t+1} + (1-\mu_y)[\beta_{y1} y_{t-1} + \beta_{y2} y_{t-2}] - \beta_r(i_{t-1} - L_{t-1}^m) + \varepsilon_{y,t}\ . \tag{25}$$

That is, inflation in the current month is set as a weighted average of the public's expectation of the medium-term inflation target, identified as L_t^m, and two lags of inflation. Also, there is a one-month lag on the output gap to reflect adjustment costs and recognition lags. Current output is determined by expected future output, $E_t y_{t+1}$, lagged output, and the ex ante real interest rate, which is proxied by $i_{t-1} - L_{t-1}^m$ (that is, agents judge nominal rates against their view of underlying future inflation rate, not just next month's inflation). Finally, the specification of longer-term yields follows the standard no-arbitrage formulation. For pricing longer-term bonds, the risk price associated with the structural shocks is assumed to be a linear function of L_t^m and S_t^m.

The above macro-finance model was estimated by maximum likelihood for the sample period from January 1988 to December 2000. The complete parameter estimates and details are in Rudebusch and Wu (2004); however, of particular interest for policy inertia are the estimates of ρ_S, which is a minuscule .026, and of ρ_u, which is .975. These estimates decisively dismiss the interest rate smoothing or monetary policy inertia interpretation of the Taylor rule. The persistent rule deviations occur not because the Fed was slow to react to output and inflation, but because the Fed responds to a variety of persistent determinants beyond current output and inflation. Some intuition for this result is given in Figure 10, which displays the initial response of the entire yield curve to inflation and output shocks from the estimated macro-finance model. Positive shocks to inflation and output in this model are followed by immediate increases in short-term interest rates, and, for the inflation shock, these increases are more than one-for-one. These responses—shown as the solid lines—reflect the absence of monetary policy partial adjustment or inertia. In contrast, the dashed lines in Figure 10 display the yield curve responses from a model that is identical to the estimated macro-finance model except that ρ_S is set equal to .9 and ρ_u equals 0. This hypothetical alternative model has substantial monetary policy inertia, and it displays markedly weaker responses to inflation and output shocks of yields that have maturities less than two years. The two quite different responses of the yield curve in these models illustrate the potential importance of the information contained in the term structure for discriminating between the two models. Given the system ML estimates, it is clear that the data prefer the macro-finance model without policy inertia.

4.3. Intraday interest rate reactions to macroeconomic data

As a third illustration of the power of the term structure to illuminate the nature of the monetary policy reaction function, I provide some new evidence on interest rate smoothing based on intraday movements of the yield curve. The underlying insight exploited here is similar to the one above: an inertial policy rule has important implications for the evolution of the entire term structure through time. Again, this approach is extremely powerful because financial markets will enforce their understanding of the monetary policy reaction function at each point in time and across interest rates at all maturities. However, while the above results implement this idea with models estimated at a monthly or quarterly frequency and substantial economic structure, the results in this section are based on the intraday response of the yield curve to macroeconomic data releases and impose minimal structure. The resulting event study provides further compelling evidence against the existence of monetary policy inertia using very different data and information.

Intuitively, changes in the path of expected future interest rates following the release of news about the state of the economy should reveal the degree of interest rate smoothing because financial markets will expect an inertial central bank to distribute the policy rate changes over several periods. To illustrate this mechanism in a simple formal structure, consider the policy inertia framework:

$$i_t = (1-\rho)\beta\bar{\pi}_t + \rho i_{t-1}, \tag{26}$$

where i_t is the average short-term (daily) policy rate during quarter t, which is set by the central bank to respond gradually over time to the annual inflation rate, $\bar{\pi}_t$ (the 4-quarter percent change). Also, annual inflation is assumed to be a simple AR(1) process

$$\bar{\pi}_t = \delta\bar{\pi}_{t-1} + \varepsilon_{1,t} + \varepsilon_{2,t}, \tag{27}$$

with two sources of independent random variation. These two shocks are distinguished by the timing of their release dates during the quarter. News about inflation in $\varepsilon_{1,t}$ is revealed at the very beginning of quarter t, while the news in $\varepsilon_{2,t}$ is revealed sometime later in quarter t. We will explore the effects of news in $\varepsilon_{1,t}$, while $\varepsilon_{2,t}$ is just included in the model to emphasize that knowledge of $\varepsilon_{1,t}$ does not determine $\bar{\pi}_t$. In addition, to be consistent with the large literature on the time series properties of various macroeconomic series, annual inflation is assumed to be a highly persistent process, with δ close to 1.[31]

To calculate the immediate response of interest rates to the revelation of $\varepsilon_{1,t}$, note that at the end of period $t-1$, the expected value of the average interest rate over the next quarter is

$$\mathrm{E}[i_t|\mathrm{e}(t-1)] = \rho i_{t-1} + (1-\rho)\beta \mathrm{E}[\pi_t|\mathrm{e}(t-1)] \tag{28}$$

$$= \rho i_{t-1} + (1-\rho)\beta\delta\pi_{t-1}, \tag{29}$$

where $\mathrm{E}[\cdot|\mathrm{e}(t-1)]$ is the expectation conditional on the information set at the end of quarter $t-1$. Similarly, just after the revelation of $\varepsilon_{1,t}$ at the beginning of quarter t, the expected value of the quarter-t interest rate is

$$\mathrm{E}[i_t|\mathrm{b}(t)] = \rho i_{t-1} + (1-\rho)\beta \mathrm{E}[\pi_t|\mathrm{b}(t)] \tag{30}$$

$$= \rho i_{t-1} + (1-\rho)\beta(\delta\pi_{t-1} + \varepsilon_{1,t}), \tag{31}$$

where $\mathrm{E}[\cdot|\mathrm{b}(t)]$ is the expectation conditional on the information set at the beginning of quarter t. Therefore, the size of the revision to the expectation of i_t in response to $\varepsilon_{1,t}$ news about inflation equals

$$\Delta\mathrm{E}[i_t|\Delta] \equiv \mathrm{E}[i_t|\mathrm{b}(t)] - \mathrm{E}[i_t|\mathrm{e}(t-1)] = (1-\rho)\beta\varepsilon_{1,t}. \tag{32}$$

That is, the change in the expectation of i_t equals the amount of inflation news multiplied by the policy response coefficient and reduced by a fraction for interest rate smoothing. Still, even with data on the change in interest rate expectations, it is difficult to determine the size of ρ from this equation, on its own, because $\beta\varepsilon_{1,t}$ must be measured in some way.[32]

However, combining the revisions in expectations of i_t with revisions of other expected future interest rates does allow the partial adjustment coefficient to be determined. Specifically, note

[31] For evidence on this point and references to the voluminous literature, see Rudebusch (1992), Rudebusch and Svensson (1999), and Pivetta and Reis (2004).

[32] Macroeconomic data surprises relative to surveys of market participants may help but are clouded by revisions to earlier data.

that at the end of quarter $t-1$, the expected value of i_{t+1} is

$$\mathrm{E}[i_{t+1}|\mathrm{e}(t-1)] = \rho\mathrm{E}[i_t|\mathrm{e}(t-1)] + (1-\rho)\beta\mathrm{E}[\pi_{t+1}|\mathrm{e}(t-1)] \tag{33}$$

$$= \rho^2 i_{t-1} + (\delta+\rho)(1-\rho)\beta\delta\pi_{t-1}. \tag{34}$$

At the beginning of quarter t, the expected value of i_{t+1} is

$$\mathrm{E}[i_{t+1}|\mathrm{b}(t)] = \rho\mathrm{E}[i_t|\mathrm{b}(t)] + (1-\rho)\beta\mathrm{E}[\pi_{t+1}|\mathrm{b}(t)] \tag{35}$$

$$= \rho^2 i_{t-1} + (\delta+\rho)(1-\rho)\beta(\delta\pi_{t-1} + \varepsilon_{1,t}), \tag{36}$$

so the revision to expectations of i_{t+1} in response to $\varepsilon_{1,t}$ is equal to

$$\Delta\mathrm{E}[i_{t+1}|\Delta] \equiv \mathrm{E}[i_{t+1}|\mathrm{b}(t)] - \mathrm{E}[i_{t+1}|\mathrm{e}(t-1)] = (\delta+\rho)(1-\rho)\beta\varepsilon_{1,t}. \tag{37}$$

Finally, the ratio of the two revisions provides a straightforward expression

$$\Delta\mathrm{E}[i_{t+1}|\Delta] \ / \ \Delta\mathrm{E}[i_t|\Delta] = \delta + \rho. \tag{38}$$

If, as noted above, the value of δ is pinned down by the well-known macroeconomic dynamics of inflation, then this ratio of revisions in expected future rates will identify ρ.

To estimate the ratio above, I use intraday data on yields of 3- and 6-month U.S. Treasury securities.[33] The revisions in these two yields are calculated over the half-hour period from 5 minutes before a release of macroeconomic data to 25 minutes after that release.[34] Changes in the 3-month yield during this window provide a reading on $\Delta\mathrm{E}[i_t|\Delta]$, while changes in a combination of the two yields give $\Delta\mathrm{E}[i_{t+1}|\Delta]$ via the expectations hypothesis—namely, twice the 6-month yield minus the 3-month yield.[35] For example, if the 3-month rate increases by 5 basis points in response to a release of higher-than-expected consumer price inflation, and the 3-month rate expected three months ahead increases by 9 basis points, then their ratio provides an estimate of $\delta + \rho$ equal to 1.8. Assuming inflation follows a unit root process, so $\delta = 1$, then

[33] I also obtained similar results using interest rate expectations from daily federal funds futures and eurodollar futures. However, an advantage to using the Treasury yields is that they enforce a consistent timing so that the macroeconomic news always occurs at the beginning of the monetary policy adjustment. In addition, Treasury markets are the most liquid ones.

[34] This 30-minute window eliminates noise from extraneous sources, such as other data releases or monetary policy actions or communications. The data are discussed in Gürkaynak, Sack, and Swanson (2004, 2005) and were kindly supplied by the authors. They obtained tick-by-tick on-the-run Treasury yield data back to 1991 from a consortium of interdealer brokers. They also show that a 30-minute window is sufficiently wide to capture the full response of financial markets to news.

[35] This calculation ignores the time-varying term premium modeled in Rudebusch-Wu (2004) and discussed above; however, changes in the ratio of these premiums at these very short maturities are likely insignificant.

the monetary policy partial adjustment coefficient is 0.8. That is, in response to news about persistently higher inflation, financial markets assume that an inertial central bank will boost the policy rate higher over the next few months but will also gradually raise it even higher in subsequent months. Alternatively, at the opposite end of the spectrum, if 3- and 6-month yields change by an identical amount in response to a persistent shock (so $\Delta \mathrm{E}[i_t|\Delta] = \Delta \mathrm{E}[i_{t+1}|\Delta]$), then $\delta + \rho = 1$ and financial markets assume that there is no monetary policy partial adjustment by central banks.

In fact, the data indicate quite clearly that the case of little or no inertia is the relevant one. I consider 315 macroeconomic data releases from July 1991 to September 2004 for the unemployment and CPI series, which are two of the most important and closely watched data releases. Of course, the formal structure outlined above applies to any persistent macroeconomic determinant of monetary policy, so the unemployment and CPI releases are pooled to increase the precision of the estimates. The median value of $\Delta \mathrm{E}[i_{t+1}|\Delta]/\Delta \mathrm{E}[i_t|\Delta]$ is 1.00; the mean value is 1.06 with a standard error of 0.15.[36] Again, with the uncontroversial assumption that macroeconomic time series are highly persistent, these results imply a central tendency for ρ of around zero to .1 and a 95 percent confidence interval that lies entirely below .4.[37]

5. Conclusion

Does the persistence of the short-term policy interest rate reflect deliberate "partial adjustment" or "inertia" on the part of the central bank? As in many other areas of economics, understanding the nature of dynamic adjustment is a hard problem and one that simple regression estimates often cannot solve. However, in contrast to many other macrodynamic puzzles, interest rates have a rich set of term structure information that can help provide answers. One of the key insights above is that although the short rate is a policy instrument, it is also a fundamental driver of long yields, so a joint macro-finance perspective can sharpen inference about the policy reaction function. The yield curve results above—for quarterly predictability, monthly system estimation, and intraday responses to news—all point to fairly rapid central bank reactions to news and information and little real-world policy inertia. In essence, quarterly monetary policy partial adjustment does not appear to be consistent with the financial market's understanding

[36] The median expectational revision ratios for inflation and unemployment releases separately are also both equal to 1.0.

[37] These results also appear robust to consideration of longer maturities, as in $\Delta \mathrm{E}[i_{t+k}|\Delta]/\Delta \mathrm{E}[i_t|\Delta]$.

of the monetary policy rule. This result appears in accord with the views of central bankers, who often note that future policy actions will be largely contingent on incoming data and future changes in the economic outlook.[38]

In terms of future research, much work can still be done to exploit yield curve information about the monetary policy reaction function. Of course, so far, this research has largely been limited to the U.S., but the above methods should be applied more widely to other countries as well. Also, other procedures can be used to extract information from the yield curve. In addition, further analysis of empirical monetary policy rules is recommended. The lagged policy rate in empirical policy rules, though useful in mopping up residual serial correlation, should not be given a structural partial adjustment interpretation with regard to central bank behavior. A better strategy is to identify and model the underlying persistent factors that influence central bank actions.

[38] The absence of partial adjustment does not mean that central banks are not trying to influence long-term interest rates. In order to influence the long rate, central banks only must present a 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.

Table 1
Partial adjustment coefficients for optimal inertial Taylor rules

Model			Optimal ρ for different loss functions					
μ_r	μ_π	μ_y	$\nu_{\Delta i} = .1$	$\nu_{\Delta i} = .5$	$\nu_{\mathrm{E}i} = .1$	$\nu_{\mathrm{E}i} = .5$	$\nu_i = .1$	$\nu_i = .5$
.0	.0	.0	-.12	.19	-.04	.34	-.57	-.51
.3	.3	.3	.18	.37	.27	.48	-.27	-.12
.5	.5	.5	.64	.70	.63	.68	.70	.80
.8	.8	.8	.90	.94	.90	.92	.93	.96
.0	.0	.5	.03	.17	.05	.26	-.34	-.23
.0	.5	.0	-.12	.16	-.04	.31	-.54	-.44
.5	.0	.0	.49	.61	.49	.67	.28	.30

Notes: The optimal lag coefficients for an inertial Taylor rule are reported for each of 7 parameterizations of the model, which have varying μ_π, μ_y, and μ_r weights on expectational terms, and for 6 variations of the loss function. The loss functions have equal weight on output and inflation volatility ($\lambda = 1$) but a stronger or weaker interest rate smoothing motive—which may take the form of minimizing $\nu_{\Delta i}\mathrm{Var}\left[\Delta i_t\right]$, $\nu_{\mathrm{Ei}}\mathrm{Var}\left[\mathrm{E}_{t-1}\left[i_t\right] - i_t\right]$, or $\nu_i\mathrm{Var}\left[i_t - r^* - \pi^*\right]$. The associated optimal g_π and g_y are not reported.

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Figure 1
U.S. economic data

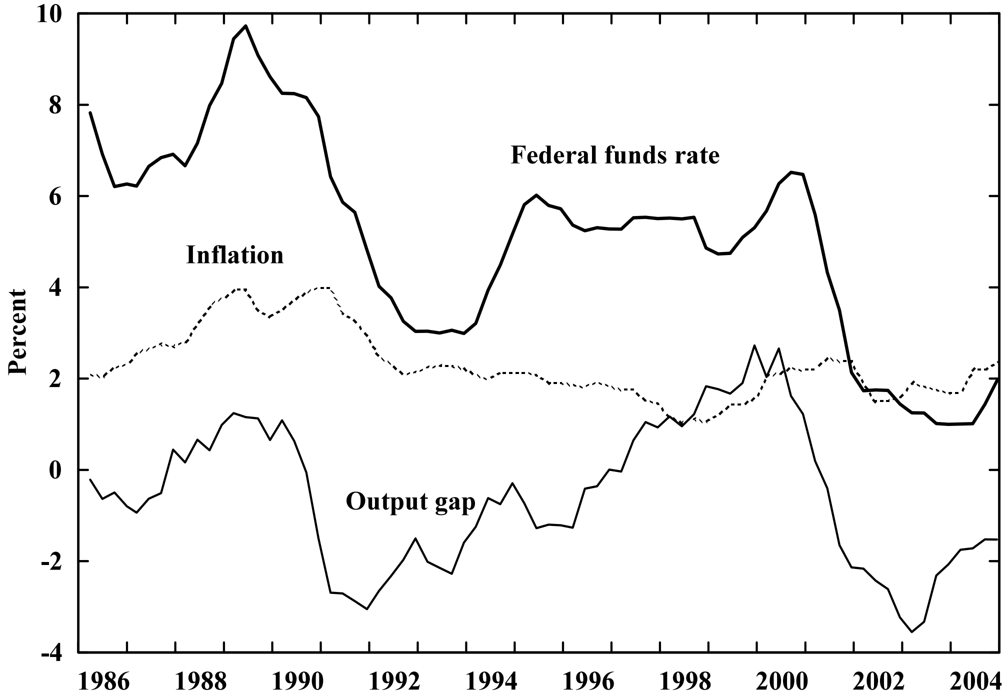


Figure 2
Actual and fitted federal funds rate

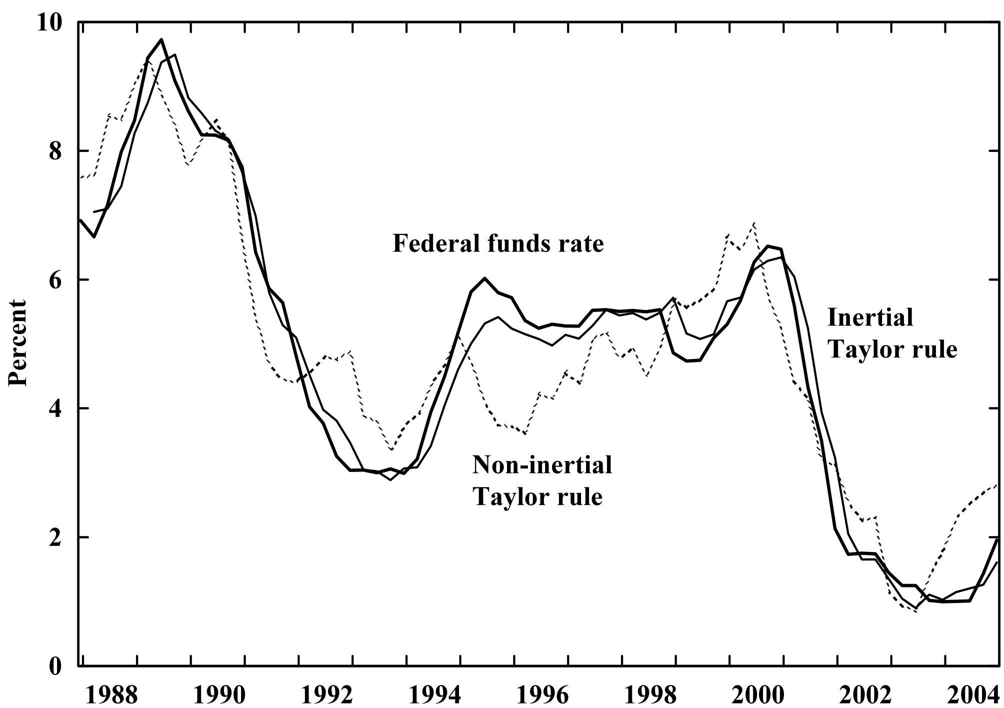


Figure 3
Residuals from estimated inertial and non-inertial Taylor rules

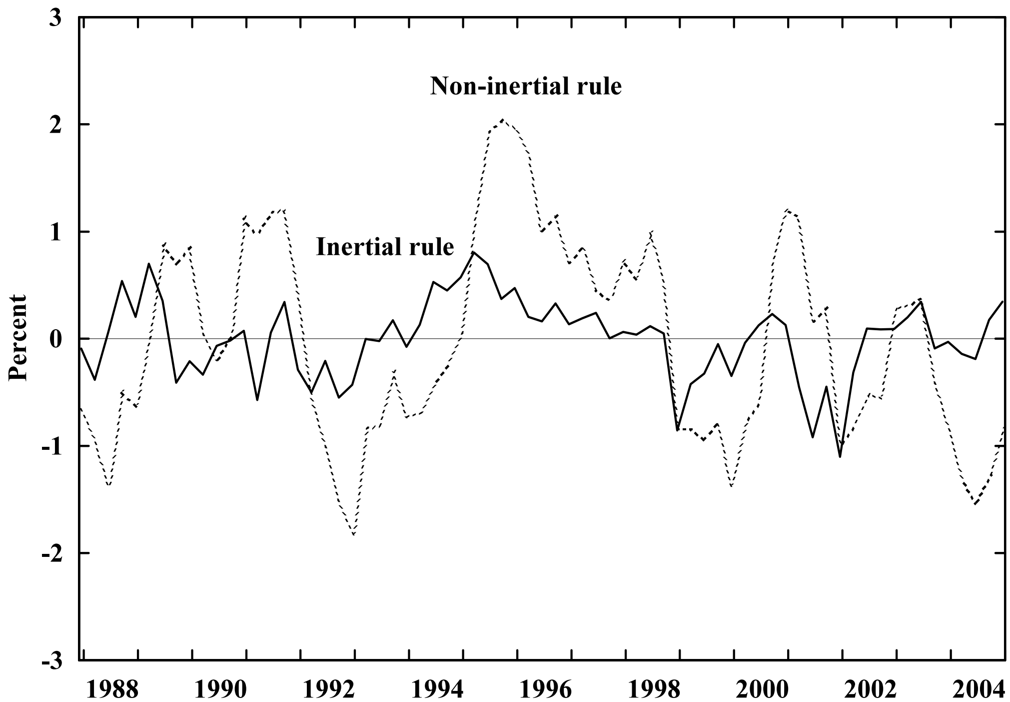


Figure 4
Actual and desired federal funds rate

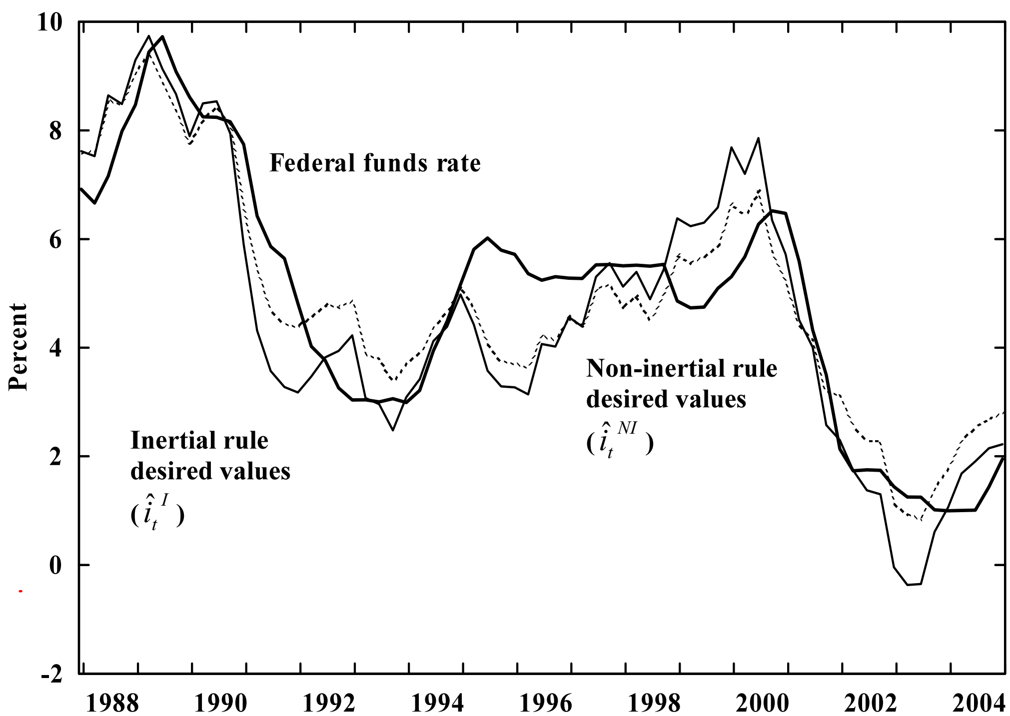


Figure 5
Deviations of actual funds rate from desired rule value

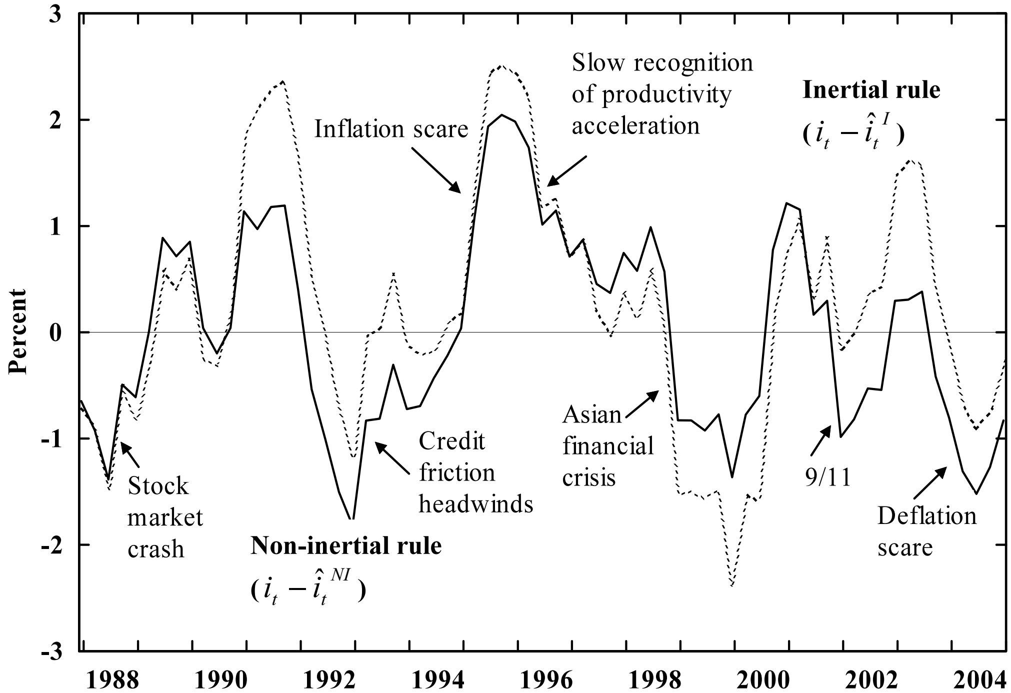


Figure 6
Differences between real-time and 2005 data vintages

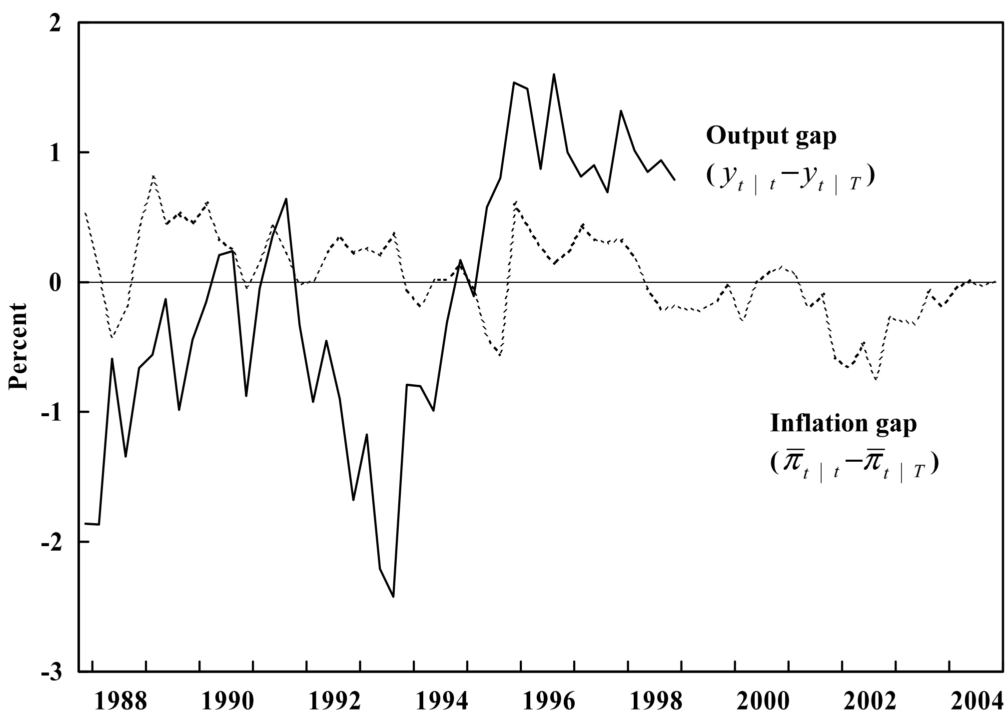


Figure 7
Matching non-inertial rule residuals with real-time data revisions

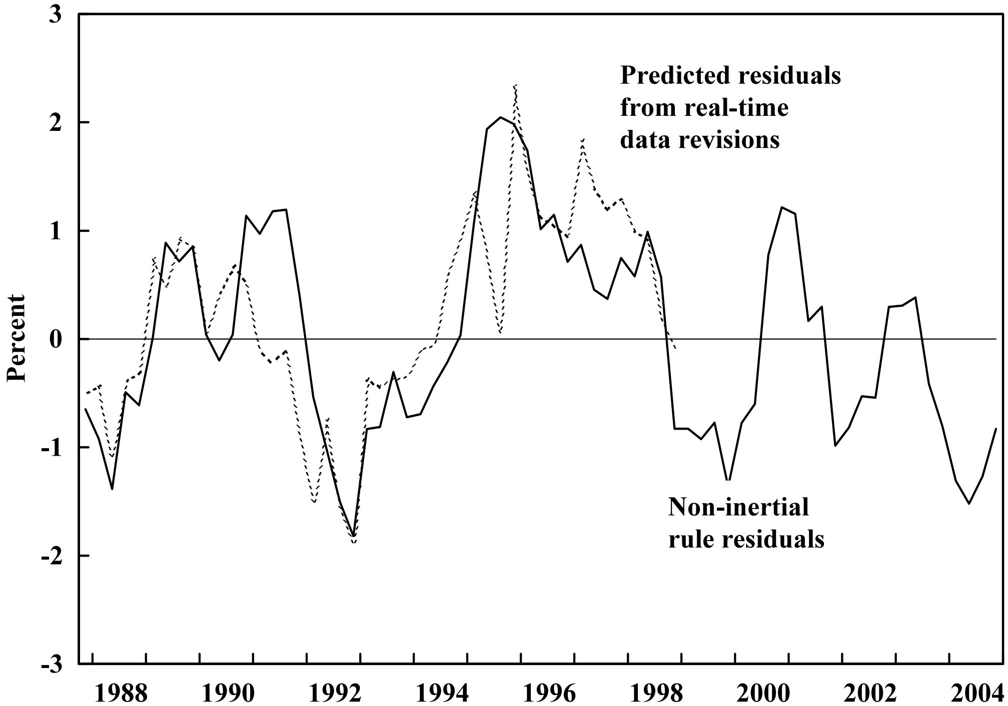


Figure 8
Implications for interest rate forecastability from policy inertia

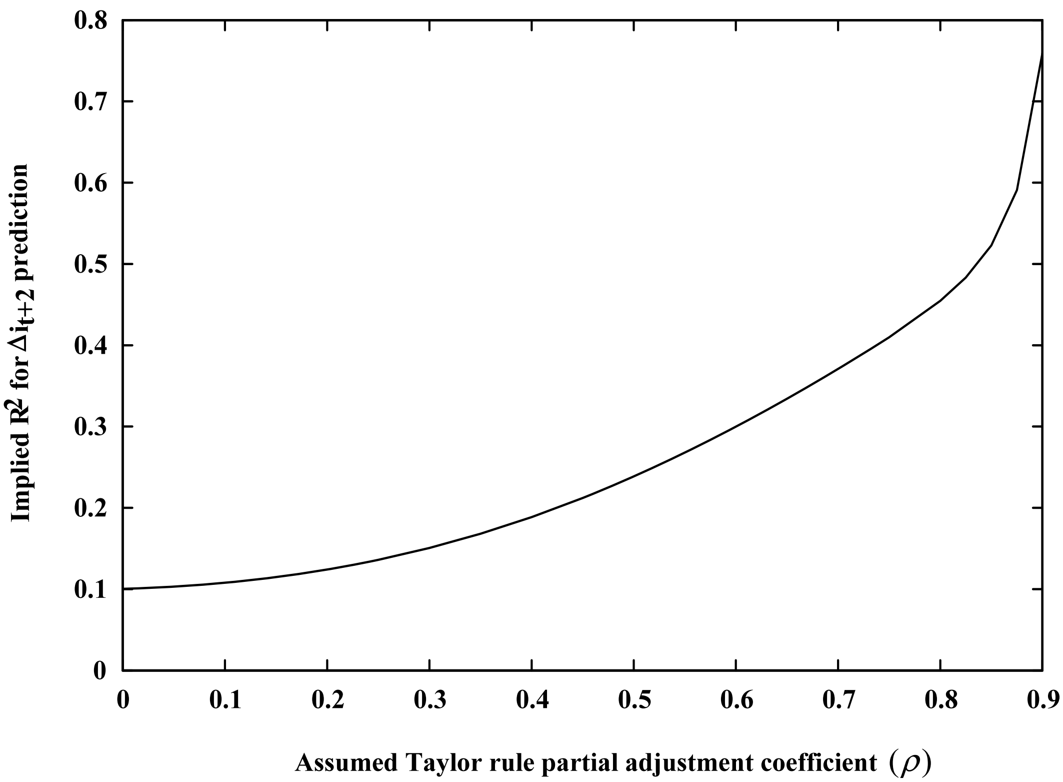


Figure 9
Actual and expected federal funds rate

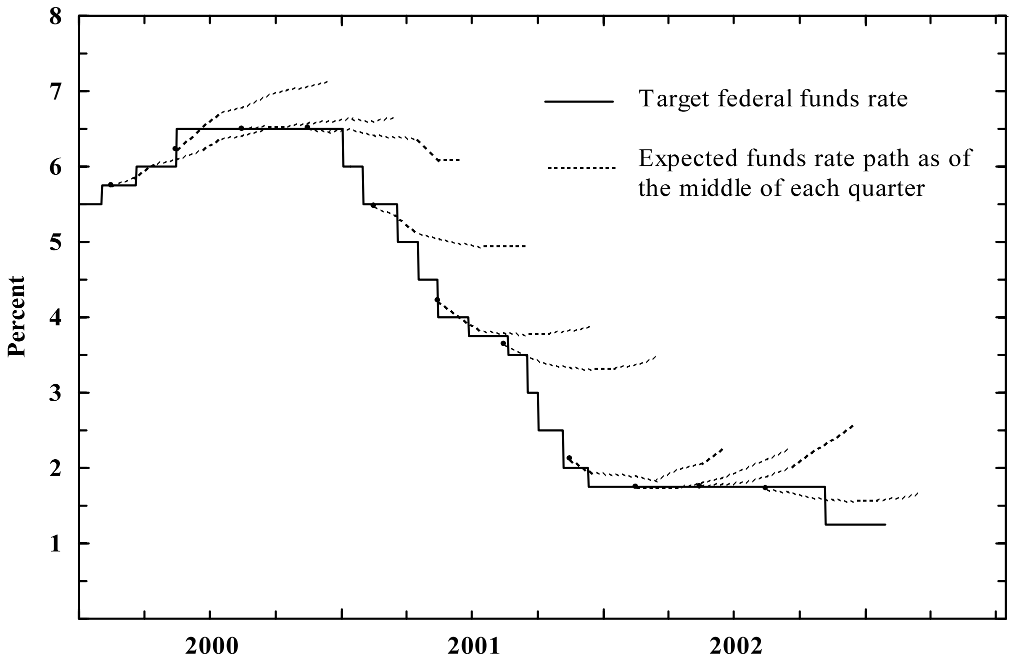
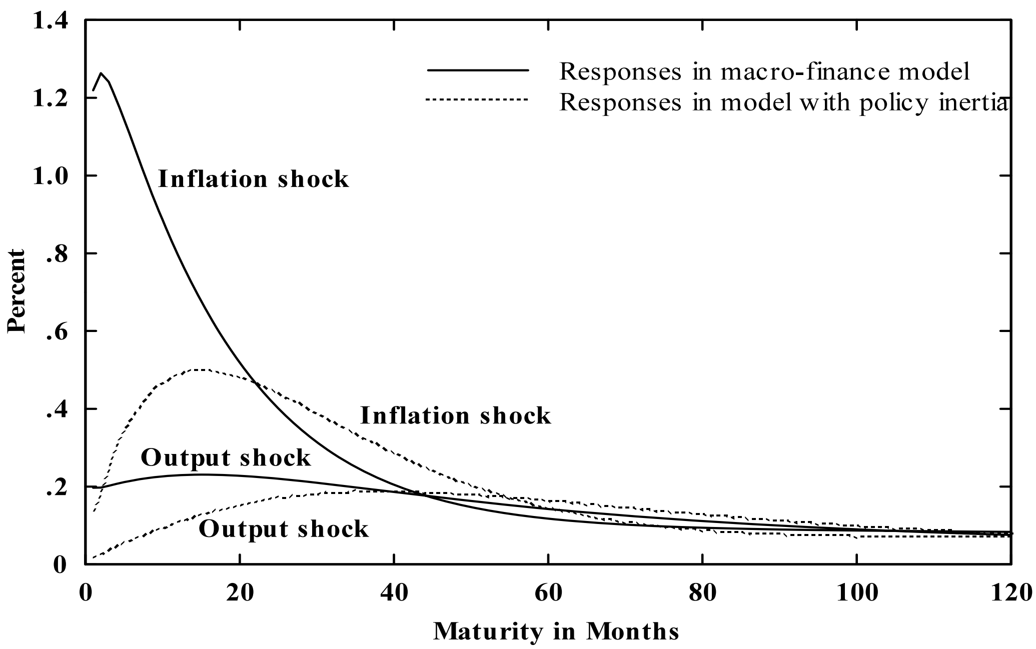


Figure 10
Initial Yield Curve Response to Output and Inflation Shocks



Note: The solid lines show the impact responses on the entire yield curve from a 1 percentage point increase in inflation or output in the estimated macro-finance model in Rudebusch and Wu (2004). The dashed lines give similar responses in a macro-finance model that assumes substantial monetary policy inertia ($\rho_S = 0.9$) and serially uncorrelated policy shocks ($\rho_u = 0$).