How stable are Monetary Policy Rules: Estimating the Time-Varying Coefficients in A Monetary Policy Reaction Function for the U.S.

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We consider the relation among the federal funds rate and the Federal Reserve's expectations for future inflation, the future gap between actual and potential output, and the future foreign exchange value of the U.S. dollar. The coefficients of this relation are biased when relevant explanatory variables are omitted and/or when the included explanatory variables are measured with error. This presents obstacles to verifying the conditions under which monetary policies can be effective which, as we show, can only be stated in terms of the relation's bias-free coefficients. To deal with this problem, we demonstrate how auxiliary variables, called concomitants, can be used to remove omitted-variable and measurement-error biases without assuming the "true" functional form of the relation to be known. Numerical algorithms for enacting this procedure are presented and an illustration is given using the U.S. quarterly data for 1960Q1-2000Q4.

I. INTRODUCTION

Considerable recent research on monetary policy has focused attention on examining the extent to which the conduct of monetary policy can be characterized by a simple relationship between a policy instrument and a small set of variables.[1] This work was largely inspired by John Taylor (1993), who suggested a rule whereby the central

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[1] There could be more than one instrument of monetary policy. For example, the Federal Reserve sometimes changes both the federal funds rate target and the discount rate and it some other times changes only the federal funds rate target.

bank sets its policy-determined interest rate in response to deviations of actual inflation from target and to the gap between actual and potential output. Taylor compared the actual federal funds rate with the rate given by a specified simple policy rule where the parameters were imposed rather than estimated, and found that the suggested rule captured the behavior of the funds rate quite well. This work spawned considerable research aimed at examining the extent to which monetary policy can be characterized as following a simple policy rule, where the parameters of the rule are estimated.[2]

Such a rule is usually represented by an equation of the following form:

$$r_t^* = \bar{r} + \alpha_1(\pi_t - \pi^*) + \alpha_2(y_t - y_t^*) + \varepsilon_t \tag{1}$$

where r_t^* = policy-determined interest rate, $\bar{r}$ = the long-term equilibrium nominal rate, π = inflation rate, π^* = target inflation rate, y = real output, y^* = potential output, and ε = disturbance term.

This approach assumes that the monetary authority reacts only to two variables--inflation and output--and that the response in the interest rate instrument to these variables is invariant over the sample period. In particular, there is assumed to be a linear relationship between the interest rate and its determinants and the coefficients in this relationship are fixed. There have been attempts to estimate changes in the coefficients in the periods before and after Paul Volcker was Chairman of the Federal Reserve Board (see Clarida, Gali and Gertler (2000), Judd and Rudebusch (1998), Taylor (1999), and Orphanides (2001a)), but this work has remained in the context of a fixed-coefficient linear model.

However, this is a rather restrictive assumption, as it is plausible that the degree to which monetary policy responds to its determinants varies over the business cycle. In other words, it would seem reasonable to allow for the possibility that the central bank adjusts the interest rate it controls more rapidly to a given gap between actual or expected

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[2] See, for example, Richard Clarida, Jordi Gali and Mark Gertler (1998, 2000), John Judd and Glenn Rudebusch (1998), Athanasios Orphanides (2001a) and Taylor (1999).

inflation, and to the gap between actual and potential output, in light of the existing cyclical situation. This highlights a more general issue of the functional form of the monetary policy reaction function. A linear function is generally used on the grounds of simplicity, but in principle one cannot rule out that the true model may be nonlinear. Thus, there is a problem of unknown functional forms. In addition, there are also the problems of interpretation of ε and the appropriateness of assumptions one might make about it.

Moreover, equation (1) assumes that the formulation and implementation of monetary policy can be described in terms of achieving two objectives relating to inflation and output. While this would appear to be a plausible overall approach to the conduct of monetary policy, it is a severely restricted characterization of the set of factors that determines the actually observed policy instrument. There are a host of other factors that affect the monetary authority's decision to raise or lower its interest rate and that are unlikely to be fully incorporated in the variables in equation (1). For example, actual and prospective developments in money supply growth, exchange rates, and commodity prices may influence the setting of the policy instrument over and above the extent to which they are reflected in current and projected inflation. Moreover, the fact that the disturbance term in (1) tends to be highly auto-correlated, which often leads researchers to introduce a lagged dependent variable into the equation, suggests that it is mis-specified (see Imke Brueggemann and Daniel Thornton (2001)). Thus, the estimation of equation (1) is subject to biases arising from omitted variables.

Finally, the estimation of equation (1) typically assumes that all of the right-hand side variables are observed without error. In particular, the assumption that the real interest rate is constant is questionable, as the rise in U.S. productivity growth in the second half of the 1990s, for example, probably indicates that the real interest rate increased over this period. More generally, it needs to be realized that in deciding on the level of the federal funds rate in a given period, the Federal Open Market Committee (FOMC)--the policy-making body of the Federal Reserve--does not have accurate data at the time of decision-making on inflation and output. While current-period inflation can be estimated fairly accurately, the stance of policy may be based on the projection of future inflation. This was clearly the case in 1994, for example, and hence the use of current-period inflation would be inappropriate. One can attempt to mimic how the monetary authority would forecast inflation, but errors would inevitably remain in the measure used for expected inflation. The problem is even more difficult with regards to the output gap, as there is considerable uncertainty at the time the policy stance is determined regarding both the current level of output, given the typically large data revisions in this series, and the level of potential output, given the conceptual and measurement issues involved (see Orphanides (2001b)). Therefore, there are ample reasons to believe that serious errors-in-variables problems are present with equation (1).

The major objective of this paper is to provide an integrated approach for dealing with all three problems identified above: unknown functional form, omitted-variable biases, and errors in variables. The next section describes our suggested method for achieving this objective. This is followed by the application of this approach to the estimation of a monetary policy reaction function for the U.S. Federal Reserve. Some concluding remarks are given at the end of the paper.

II. Specification and Estimation of a Monetary Policy Reaction Function

A. Embedding a True model in a Class of Models

We assume that the Federal Reserve has a target for a nominal short-term interest rate that is based on its expectations about future inflation, expected potential output, and other variables affecting the state of the economy. This target affects the actual interest rate. Algebraically,

$$r_t^* = \bar{r}_t + \alpha_{1t} E[(\pi_{t+f} - \pi_t^*) \mid \Omega_t] + \alpha_{2t} E[(y_t - y_t^*) \mid \Omega_t] + \sum_{\ell=3}^{n_t} \alpha_{\ell t} x_{\ell t}^* \qquad (2)$$

where r_t^* is an actual nominal short-term interest rate, $\bar{r}_t$ is the long-term equilibrium nominal rate, $E(\pi_{t+f} \mid \Omega_t)$ is the expected inflation rate for a future period t + f based on the information Ω_t available to the central bank at the time it sets interest rates, π_t^* is the central bank's target inflation rate, y_t is real output and y_t^* is potential output so that $E[(y_t - y_t^*) \mid \Omega_t]$ measures the expected output gap. The $x_{\ell t}^*$s with $\ell > 2$ are all other factors that influence the policy-determined interest rate besides the first three variables on the right-hand side of equation (2). We treat the variables with an asterisk as unobservable true measurements, known only to the central bank, t indexes time, and the total number of explanatory variables n_t depends on time if the set of the determinants of r_t^* changes over time. Note that there is no need to have an error term in equation (2), since we included all the determinants of r_t^* on the right-hand side of (2). To simplify our notation, we write $\bar{r}_t - \alpha_{1t}\pi_t^* = \alpha_{0t}$, $E(\pi_{t+f} \mid \Omega_t) = x_{1t}^*$, and $E[(y_t - y_t^*) \mid \Omega_t] = x_{2t}^*$. Suppose that data on $x_{\ell t}^*, \ell = K, \ldots, n_t$, are not available. Such variables are called excluded variables. With this notation, (2) can be rewritten as

$$r_t^* = \alpha_{0t} + \sum_{j=1}^{K-1} \alpha_{jt} x_{jt}^* + \sum_{\ell=K}^{n_t} \alpha_{\ell t} x_{\ell t}^* \tag{3}$$

Excluded variables are not unique and can have several representations, as John Pratt and Robert Schlaifer (1984, p. 13) have shown. We may not know anything about some of these variables. Even the data that are available on a subset of the variables in (3) may contain measurement errors. For example, our proxies for the unobservable central bank expectations that appear in equation (2) are necessarily approximations. We now show how we deal with these problems. The variables x_{jt}^*, j = 1,…, K-1, are labeled the included explanatory variables. The intercept, α_{0t}, is also of interest, since it is a function of both the long-run equilibrium nominal rate and the central bank's target inflation rate, both of which may change over time.

B. Three Fundamental Problems with Equation (1) and Their Solutions

Unknown-functional-form problem: P.A.V.B. Swamy and George Tavlas (2001a) define as true

(I) any variable or value that is not incorrectly measured and

(II) any economic relationship (i) with the true functional form, (ii) without any omitted explanatory variables, and (iii) without incorrectly measured variables.

Equation (3) satisfies conditions II(ii) and II(iii) because by construction, it has no omitted or incorrectly measured variables. These are the two good theoretical properties that are built into (3). Consequently, equation (3) is true in the sense of (II) if we can prove that it has the true functional form. But we are unable to do so because equation (3)'s true functional form is unknown. To be sure, there is no such thing as the true functional form of (3) unless there is a real-world relationship underlying (3). Suppose that such a relationship exits.[3] Then one way of embedding the unknown "true" functional form of equation (3) in a class without assuming a specific, possibly false, form is not to restrict the pattern of variation in its coefficients in any way. Different paths of variation in these coefficients generate various functional forms and the class of functional forms equation (3) represents is unrestricted as long as its coefficients' pattern of variation is not restricted in any way. This supports the notion that a member of such an unrestricted class can be assumed to be true in the sense of (II) above. For certain (unknown) paths of variation in its coefficients, (3) coincides with the underlying "true" economic relationship. Clearly, the coefficients, α_{jt}, j = 1, …, n_t, are constants if the functional form of the "true" version of (3) is linear. This linear form is unlikely to be true because the stance of monetary policy changes with the position of the economy in the business cycle phase. For this reason, we prefer not to assume that the α_{jt}'s are constant.

Thus, the possible non-linearity of the "true" economic relationship underlying equation (3) is one justification for letting the coefficients, α_{jt}, j = 0, 1, …, n_t, of the linear--in variables--form of (3) vary over time. The α coefficients that follow the "true"

[3] We use the term "true" in this paper to mean "the assumed truth" precisely because of this assumption. It is clear that the term "assumed true model" does not refer to a model that is absolutely true, but refers to a model that is assumed to be true.

paths of variation represent the "true" effects on r_t^* of its determinants. These are what we refer to as the bias-free coefficients on the x_{jt}^* s and $x_{\ell t}^*$ s. The adjective "bias-free" is appropriate for these coefficients because the biasing effects of omitted variables and misspecifications of the "true" functional form are not present on the coefficients of (3) with the "true" pattern of variation. However, the partial derivatives of r_t^* with respect to the x_{jt}^* s and $x_{\ell t}^*$ s cannot be determined without knowing the functional form of the "true" economic relationship underlying (3).

To the extent that the "true" functional form of equation (3) is different for different central banks, it is not possible to obtain a single solution to the unknown-functional-form problem that is appropriate for all applications of equation (3). Any solution to the problem is application specific. A solution to the problem that is appropriate for a particular application we make of equation (3) in this paper is given in Section III below.

Omitted-variables-bias problem: Pratt and Schlaifer (1984) proved that if the net effect of excluded variables ($\sum_{\ell=K}^{n_t} \alpha_{\ell t} x_{\ell t}^*$) in equation (3) is replaced by an error term, then it is "either meaningless or false" to assume that this error term is uncorrelated with the included explanatory variables. Their proof involves in showing that the α 's are altered when equation (3) is rewritten in terms of the x_{jt}^* s and a function of the x_{jt}^* s and $x_{\ell t}^*$ s with $j \neq \ell$. This property of the α s shows that the error term written in place of $\sum_{\ell=K}^{n_t} \alpha_{\ell t} x_{\ell t}^*$ is non-unique and assumes different forms in different representations of equation (3). The error term in some of these representations is correlated with the included explanatory variables (see Swamy, Jatinder Mehta and Rao Singamsetti 1996, p. 124). It is also true, however, that some of the $x_{\ell t}^*$ s cannot be identified. Under these circumstances, an assumption that the included explanatory variables are uncorrelated with the net effect of unidentified excluded variables is "meaningless," and the stronger assumption that the included explanatory variables are not correlated with any excluded variable is false, a result (and terminology) due to Pratt and Schlaifer (1984, p. 12). Thus, what we know about equation (3) is not enough to prove that excluded variables are uncorrelated with the included explanatory variables.

Solution: Do not assume that excluded variables are uncorrelated with the included explanatory variables, but assume that they are related to the included explanatory variables according to:

$$x^*_{\ell t} = \varphi_{0\ell t} + \sum_{j=1}^{K-1} \varphi_{j\ell t} x^*_{jt} \qquad (\ell = K, \ldots, n_t) \tag{4}$$

Again, all of the coefficients of equation (4) are allowed to vary freely so that equation (4) coincides with the underlying "true" economic relationship for a certain pattern of variation in its coefficients. The non-linearities of the "true" economic relationship underlying equation (4) are our justifications for letting the coefficients of equation (4) vary over time. In order to account for the correlations among the included and excluded explanatory variables, substitute equation (4) into equation (3). This gives

$$r^*_t = \beta_{0t} + \sum_{j=1}^{K-1} \beta_{jt} x^*_{jt} \tag{5}$$

where $\beta_{0t} = (\alpha_{0t} + \sum_{\ell=K}^{n_t} \alpha_{\ell t}\varphi_{0\ell t})$ and for j = 1,…, K-1, $\beta_{jt} = (\alpha_{jt} + \sum_{\ell=K}^{n_t} \alpha_{\ell t}\varphi_{j\ell t})$. The coefficients of this equation are unaltered when equation (3) is rewritten in terms of the x^*_{jt} s and a function of the x^*_{jt} s and $x^*_{\ell t}$ s with $j \neq \ell$ (see Swamy, Mehta and Singamsetti 1996, p. 124).

The elimination of the last n_t - K explanatory variables (or so-called "excluded variables") from the "true" economic relationship underlying equation (3) comes at a cost. This cost is the contamination of the first K coefficients of (3) caused by the addition of an unwanted term of the type $\sum_{\ell=K}^{n_t} \alpha_{\ell t}\varphi_{j\ell t}$ to each one of those coefficients, as in equation (5). It is not possible to avoid this contamination as long as the coefficients of equation (4) are nonzero. Because of this contamination the explanatory variables in equation (5) and the same explanatory variables in equation (3) cannot have the same coefficients, even though these two equations have the same dependent variable. Further differences between the corresponding coefficients on the included explanatory variables in equations (3) and (5) arise if measurement errors are present in our data, as we now show.

Errors-in-variables problem: Observed (as opposed to the "true") values of the variables are likely to contain measurement errors. More specifically, the differences between the unobservable central bank expectations in equation (2) and their proxies can be viewed as measurement errors. We can only use some proxies for these expectations because we have no idea of how these expectations are formed. So the proxies we use for these expectations cannot be based on any knowledge the central bank has but we do not have. They can only be based on our "coherent" beliefs about the central bank expectations and are different from the rational expectations of variables considered in much of the literature.

Solution: Suppose that $r_t = r_t^* + v_{0t}$ and $x_{jt} = x_{jt}^* + v_{jt}$, j = 1,…, K-1, where r_t and x_{jt} are the observed counterparts of r_t^* and x_{jt}^*, respectively, and the vs represent measurement errors, which may not have means zero. Some of the x_{jt}^*s denote the state of policy-maker's knowledge of the economy at the time of decision-making on inflation and output. They are different from the corresponding x_{jt}s because of data revisions, measurement errors, and of the differences between equation (5) and the models that give the forecasts of future variables for the policy-maker's use. We indicate our choice of r_t and x_{jt} in Section III below. To bring us closer to estimation, substitute into equation (5) the observable counterparts of its dependent and explanatory variables to obtain

$$r_t = \gamma_{0t} + \sum_{j=1}^{K-1} \gamma_{jt} x_{jt} \tag{6}$$

where $\gamma_{0t} = (\alpha_{0t} + \sum_{\ell=K}^{n_t} \alpha_{\ell t}\varphi_{0\ell t} + v_{0t})$ and for j = 1,…, K-1,

$\gamma_{jt} = (\alpha_{jt} + \sum_{\ell=K}^{n_t} \alpha_{\ell t}\varphi_{j\ell t})(1 - \frac{v_{jt}}{x_{jt}})$.[4] The method used to derive equation (6) may be called a model-consistent method of dealing with errors in variables in general and with individuals' expectations in particular.

C. The Appropriate Interpretations of the Coefficients of Equation (6) and Their Implications

(a) The intercepts of equation (4), i.e., $\varphi_{0\ell t}$, are the portions of excluded variables remaining after the effects of the included explanatory variables have been removed. Only these portions appear in the intercepts of equations (5) and (6). This shows that the interpretation that the error terms of econometric models represent the net effects of omitted variables is generally inappropriate.

(b) For j = 0, γ_{jt} of equation (6) is the sum of three parts: (i) the bias-free intercept of equation (3), (ii) a combination of the portions of excluded variables ($\varphi_{0\ell t}$) with the coefficients on excluded variables ($\alpha_{\ell t}$) acting as its coefficients, and (iii) the measurement error in the dependent variable of equation (6). This means that the intercepts of equations (3) and (4), the coefficients on excluded variables, and the measurement error in the dependent variable are the sources of the intercept of equation (6).

[4] Note that dummy variables should not be used as explanatory variables in equation (6) because they are not explanatory variables. The consequence of including a dummy variable as an explanatory variable with a time-varying coefficient is that the coefficient times the variable is zero whenever the variable takes the value zero regardless of the value of the coefficient. Hence the coefficient values corresponding to the zero values of the variable are arbitrary. Even when an explanatory variable is not a dummy variable, the value of its time-varying coefficient is also arbitrary whenever the variable takes the value zero.

(c) For j = 1,…, K-1, γ_{jt} of equation (6) is also the sum of three components, α_{jt}, $\sum_{\ell=K}^{n_t}\alpha_{\ell t}\varphi_{j\ell t}$, $(\alpha_{jt}+\sum_{\ell=K}^{n_t}\alpha_{\ell t}\varphi_{j\ell t})(-\frac{v_{jt}}{x_{jt}})$, which have the following economic interpretations: (i) the term α_{jt} represents the bias-free coefficient on the "true" value of the jth included explanatory variable (x^*_{jt}) in the "true" economic relationship underlying equation (3). (ii) The term $\sum_{\ell=K}^{n_t}\alpha_{\ell t}\varphi_{j\ell t}$ represents an "indirect" effect due to the fact that the "true" value of the jth included explanatory variable affects the "true" values of excluded variables that, in turn, affect the "true" value of the dependent variable of equation (3). This term is also called an "omitted-variables bias" in econometrics. (Recall that $\alpha_{\ell t}$ is the effect of the "true" value of the ℓth omitted variable ($x^*_{\ell t}$) on the "true" value of the dependent variable (r^*_t) and $\varphi_{j\ell t}$ is the effect of the "true" value of the jth included explanatory variable (x^*_{jt}) on the "true" value of the ℓth excluded variable ($x^*_{\ell t}$).) The term, $\sum_{\ell=K}^{n_t}\alpha_{\ell t}\varphi_{j\ell t}$, is the same as a simultaneous-equations bias. Under our assumptions, equations (3) and (4) hold simultaneously and they are combined into one in (5). In other words, whenever the $\varphi_{j\ell t}$s are not zero, i.e., the relations in (4) hold, they give nonzero values to the simultaneous-equations bias. (iii) Finally, the term $[-(\alpha_{jt}+\sum_{\ell=K}^{n_t}\alpha_{\ell t}\varphi_{j\ell t})(v_{jt}/x_{jt})]$ captures a "measurement-error bias" due to mismeasuring the jth included explanatory variable (recall that v_{jt} is the measurement error in x_{jt}).

(d) Omitted-variable biases are unlikely to have the same pattern of variation that excluded variables have and hence are likely to be different from excluded variables. No procedure other than that of converting some of excluded variables into included variables can reduce the magnitude of omitted-variable biases. With respect to measurement-error biases, their magnitudes are reduced whenever the absolute values of the errors of measurement in our data are reduced. We cannot go from equation (3) to equation (6) without contaminating the bias-free coefficients, α_{jt}, j = 1,…, K-1. Omitted-variable and measurement-error biases contaminate them. Similarly, the intercept of equation (3) becomes the intercept of equation (6) only after it is contaminated by a combination of the intercepts of equations (4) and by the measurement error in the dependent variable of equation (6). For these reasons, we call the γ_{jt} s the "biased" coefficients.

In the discussion that follows, it should be remembered that only the coefficients, α_{jt}, j = 0, 1,…, K-1, that have the "true" pattern of variation provide information about the "true" economic relationship underlying equation (3). Unfortunately, we do not have the necessary data to estimate (3). We only have data on the dependent and explanatory variables of equation (6). Estimates of the coefficients of the latter equation cannot provide information about the assumed true pattern of variation in α_{jt}, j = 0, 1,…, K-1, unless we have a method of separating the bias-free coefficients (i.e., the coefficients of equation (3) having the "true" pattern of variation) from the omitted-variable and measurement-error biases contained in the coefficients of equation (6). We present such a method below.

The bias-free coefficients cannot be constant unless equation (3) is linear, as we have already pointed out. The remaining two components of the coefficients of equation (6) (the omitted-variable and measurement-error biases) cannot be constant if (i) the set of omitted variables changes over time, (ii) the ratios (v_{jt} / x_{jt}) vary over t, and (iii) the "true" functional forms of equations (3) and (4) are nonlinear. Thus, the premises of fixed-coefficient models are inconsistent with the appropriate interpretations of the coefficients of equation (6) if any one of (i) to (iii) holds and hence fixed-coefficient models cannot coincide with the corresponding "true" economic relationships.[5]

[5] I-Lok Chang, Swamy, Charles Hallahan and Tavlas (2000), Swamy and Tavlas (2001a) and Sophocles Brissimis, George Hondroyiannis, Swamy and Tavlas (2001) show that in general, stationarity, linearity, and differencing of variables to induce stationarity are inconsistent with the appropriate interpretations of the coefficients of equation (6).

In addition, (i) for j = 1,…, K-1, γ_{jt} is a function of x_{jt}, as can be seen from its measurement-error-bias component, $[-(\alpha_{jt} + \sum_{\ell=K}^{n_t} \alpha_{\ell t}\varphi_{j\ell t})(v_{jt} / x_{jt})]$, (ii) all the coefficients of equation (6) are functions of the time-varying coefficients, $\alpha_{\ell t}$, ℓ = K, …, n_t, on excluded variables and (iii) α_{0t} is a function of α_{1t}. These properties demonstrate that it is inappropriate to assume that in equation (6), the coefficients are constant and uncorrelated with each other and with the included explanatory variables. Any method of estimation of equation (6) that ignores these correlations can lead to inconsistent estimators of its parameters.

D. A Consistent Method of Estimating Equation (6)

One question that needs to be answered before estimating equation (6) is that of parametrization: which features of equation (6) ought to be treated as constant parameters? We should be aware that inconsistencies arise if we adopt a parameterization that is not consistent with the interpretations of the coefficients of equation (6) given above. To avoid such a parameterization, we make the following assumptions:

Assumption I: *The bias-free and other components of the coefficients of equation (6) satisfy the stochastic equations*

$$\alpha_{jt} = \sum_{m \in P_j} \pi_{jm} z_{mt} + \varepsilon_{j1t} \tag{7}$$

$$\gamma_{jt} - \alpha_{jt} = \alpha_{jt}(-\frac{v_{jt}}{x_{jt}}) + (\sum_{\ell=K}^{n_t} \alpha_{\ell t}\varphi_{j\ell t})(1 - \frac{v_{jt}}{x_{jt}}) = \sum_{m \notin P_j} \pi_{jm} z_{mt} + \varepsilon_{j2t}, \tag{8}$$

which can be combined into

$$\gamma_{jt} = \sum_{m=0}^{p-1} \pi_{jm} z_{mt} + \varepsilon_{jt} \qquad \textit{(j = 0, 1,…, K-1)} \tag{9}$$

where $z_{0t} = 1$ *for all t, the* z_{mt} *with m > 0 are called "concomitants,"* $P_j \subset P$ = *{0, 1, 2,…,* $p-1$*}, which is the "index set" formed by the values of m, and* $\varepsilon_{jt} = \varepsilon_{j1t} + \varepsilon_{j2t}$ *with* *E(*ε_{j1t} $|$ z's) = 0 *and E(*ε_{j2t} $|$ z's) = 0*. The concomitants are assumed to be mean independent of* ε_{j1t} *and* ε_{j2t}*, which are assumed to satisfy the stochastic equation:*

$$\varepsilon_{jt} = \phi_{jj}\varepsilon_{jt-1} + a_{jt} \tag{10}$$

where for j, j' *= 0, 1,…, K-1,* $-1 < \phi_{jj} < 1$*, and the* a_{jt}*s are serially uncorrelated with E(*a_{jt}*) = 0 and* $E(a_{jt}a_{j't}) = \sigma_{jj'}$ *for all t.*[6, 7]

In equation (9), p concomitants including z_{0t} are used to explain the variation in γ_{jt}. Of these concomitants, some are used in equation (7) to explain the variation in the bias-free component of γ_{jt} and the remaining are used in equation (8) to explain the variation in the indirect effects and measurement-error biases contained in γ_{jt}. The coefficient π_{j0} on z_{0t} represents a constant portion of α_{jt} if it has the right sign and of γ_{jt} - α_{jt} otherwise. The portion of α_{jt} that depends on z_{mt} with m > 0 is a variable. Also, note that the sets of concomitants used in equations (7) and (8) may be different for different coefficients of equation (6). Equations (7) and (8) relate the time-varying components of the coefficients of equation (6) to the time-varying concomitants and error terms. The greater the proportion of the variation in γ_{jt} explained by the concomitants in equation (9), the better. Equating time-varying coefficients to fixed coefficients is a specification error and this error is avoided in equations (7) and (8). These equations have error terms because the concomitants included in them may not explain all the variation in the components of the coefficients of equation (6). Equation (10) is introduced to allow for the possibility that the unexplained portion of γ_{jt} is serially correlated.

It should be noted that the decompositions of γs in (9), unlike those in (6), depend on our choice of concomitants. This difference in decompositions will prove useful later in this analysis. Also, the explanatory variables of equation (6) and their coefficients

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6 The term "concomitants" is borrowed from Pratt and Schlaifer (1988) who do not use equation (9) in their work. They also do not deal with errors of measurement.

7 Equation (10) can be generalized to include nonzero correlations between ε_{jt} and $\varepsilon_{j't-1}$ with $j \neq j'$ (see Chang et al. 2000). We write the covariance matrix of $(a_{0t}, a_{1t}, \ldots, a_{K-1,t})'$ as $\sigma_a^2\Delta_a$, where Δ_a is the K × K matrix having $(\sigma_{jj'} / \sigma_a^2)$ as its (j, j')th element.

cannot have the same pattern of variation and the concomitants are introduced to explain the variation in the coefficients. For this reason, the concomitants cannot be the same as the explanatory variables and should not be included in equation (6) as explanatory variables. For the same reason, the explanatory variables should not be included in equations (7) and (8) as concomitants.

Inserting equation (9) into equation (6) gives the appropriate conditional moments of r_t given the values of x_{jt} s and z_{mt} s if equations (7) and (8) include those concomitants that satisfy the following assumption.

Assumption II: (i) *Given the values of the concomitants,* z_{mt}*, the explanatory variables of equation (6) are independent of the* ε_{jt}*, and*

(ii) $\mathrm{pr}(x_{jt} \in S_j, j = 1, \ldots, K-1 \mid z_{1t}, \ldots, z_{p-1,t}, r_t) = \mathrm{pr}(x_{jt} \in S_j, j = 1, \ldots, K-1 \mid z_{1t}, \ldots, z_{p-1,t})$,

where the S_j *'s are the intervals containing the realized values of the* x_{jt}*s to which the realized values of* r_t *are connected by the "true" relationship underlying equation (3).*[8, 9]

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[8] Assumption II(i) captures the idea that the explanatory variables of equation (6) can be independent of the ε_{jt} s conditional on the given values of concomitants, even though they are not unconditionally independent of the γ_{jt} s. In other words, Assumption II(i) says that x_{jt} is correlated with its coefficient γ_{jt} because of the terms $\pi_{j0} + \sum_{m=1}^{p-1} \pi_{jm} z_{mt}$, but once these terms are subtracted from γ_{jt}, the remainder (ε_{jt}) is independent of x_{jt}. That is, by using the decomposition (9) of γ_{jt} that is different from that given in equation (6), we find a solution to the problem of the correlation between γ_{jt} and x_{jt}. If the decomposition of γ_{jt} in (9) were the same as that in (6), then Assumption II(i) would be false. As regards Assumption II(ii), it says that r_t is related to the x_{jt}, j = 1,…, K-1, and the z_{mt}, m = 1,…, p-1, according to the "true" relationship underlying equation (3), even though the coefficients of this relationship are changed in the presence of omitted variables and measurement errors (see Pratt and Schlaifer 1988, pp. 34-37). In this statement, the quotes added to the word true are unnecessary if the true economic relationship underlying equation (3) actually exists. Arnold Zellner (1988), Robert Basmann (1988) and Pratt and Schlaifer (1984, 1988) describe the properties of such true economic relationships. Assumption II(ii) is false if, for all t and j = 1,…, K-1, both $E(r_t \mid x_{jt}$, j = 1,…, K-1, and z_{mt}, m = 1,…, p-1) and $E(x_{jt} \mid z_{mt}$, m = 1,…, p-1, $r_t)$ exist and are finite and if $E(x_{jt} \mid z_{mt}$, m = 1,…, p-1, $r_t) \neq E(x_{jt} \mid z_{mt}$, m = 1,…, p-1).

The a priori choice of concomitants depends on intuitions about how the information contained in the concomitants is relevant to the information about the bias-free coefficients being sought. One such intuition is that the sources of variation in the bias-free components of the coefficients of equation (6) are the non-linearities of the "true" true economic relationship underlying equation (3). From this intuition it follows that the concomitants to be included in equation (7) are those that capture such non-linearities. Thus, equations (7) and (8) play a crucial role of maintaining the connection between the coefficients of equations (3) and (6).[10]

Substituting (9) into (6) gives an equation in a consistently estimable form

$$\mathrm{r}_t = \pi_{00} + \sum_{m=1}^{p-1} \pi_{0m} z_{mt} + \pi_{10} x_{1t} + \sum_{m=1}^{p-1} \pi_{1m} z_{mt} x_{1t} + \cdots + \pi_{K-1,0} x_{K-1,t} + \sum_{m=1}^{p-1} \pi_{K-1,m} z_{mt} x_{K-1,t}$$
$$+ \varepsilon_{0t} + \varepsilon_{1t} x_{1t} + \cdots + \varepsilon_{K-1,t} x_{K-1,t} \qquad (t = 1, 2, \ldots, T) \qquad (11)$$

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[9] There are similarities as well as differences between Assumption II(i) and the econometrician's definition of instrumental variables. Like the instrumental variables, the concomitants are highly correlated with the explanatory variables of equation (6) but unlike the instrumental variables, the concomitants are only independent of the remainders ε_{jt} of the coefficients of equation (6) and are not independent of the intercept, γ_{0t}. Consequently, Assumption II(i) is consistent and the definition of instrumental variables is inconsistent with the appropriate interpretations of the coefficients of equation (6). Also, Chang et al. (2000, p. 117) show that the instrumental variables do not exist as long as the error term of equation (9) for j = 1,…, K-1 is not degenerate at zero. To choose among different sets of concomitants, we need additional criteria, such as those presented in Swamy and Tavlas (2001a).

[10] We have shown in footnote 4 that the coefficients of equation (6) are arbitrary if the corresponding explanatory variables are dummy variables. This problem does not arise if some of the concomitants are dummy variables because the constant coefficients of equation (9) take the same values whether or not the corresponding z variables take zero values. By making the coefficients of equation (6) functions of both dummy and continuous variables, we can allow both discrete and continuous changes in the coefficients of equation (6). The functional form of equation (6) when some of the z's in (7) and (8) are dummy variables, is more general than that of the fixed-coefficient analog of equation (6) with dummy variables as some of its explanatory variables. The inclusion of dummy variables in the fixed-coefficient analog of equation (6) introduces only discrete changes into its coefficients. This is a very restrictive functional form that can be false. Moreover, the fixed-coefficient analog of equation (6) with dummy variables is not consistent with the appropriate interpretations of the coefficients of equation (6).

Note that this equation is not obtained by adding an arbitrary error term to an ad hoc mathematical equation. It is derived from equations (3)-(10), which together form a consistent set. This consistency is relative to the appropriate interpretations of the coefficients of equation (6). The existence of equations (3)-(6) follows from that of the "true" economic relationship underlying equation (3).[11] Equation (11) has K error terms, as many as there are coefficients in equation (6). Of these, K-1 errors are the products of εs and the included explanatory variables. Consequently, the sum of the K error terms in (11) is both heteroscedastic and serially correlated. Thus, our derivation of equation (11) is justified by its producing such errors with no appeal to any arbitrary heteroscedasticity assumption. Since the pattern of variation in x_{jt} is different from that in γ_{jt}, the x_{jt}s cannot be proper concomitants. For choices of proper concomitants satisfying Assumptions I and II, equation (11) provides a better second-order approximation to equation (6) than equation (11) with $z_{jt} = x_{jt}$ or the equations given by another approach known as the hierarchical Bayes procedure.[12]

Under Assumptions I and II, the right-hand side of equation (11) with the last K terms suppressed gives the conditional expectation of r_t as a nonlinear function of the

[11] By contrast, cointegrating and error-correction models exist only if a particular term made explicit in Swamy and Mehta (1996) is added and subtracted on the right-hand side of a vector moving average model with fixed coefficients, which itself does not exist unless a stationarity condition is satisfied. This condition is inconsistent with the appropriate interpretations of the coefficients of equation (6), as the above argument shows. Omitted-variable and measurement-error biases contained in the coefficients of cointegrating and error-correction models are ignored.

[12] Differences between our and the hierarchical Bayes modeling procedures can be seen if our derivation of equation (11) is compared with G. S. Datta, P. Lahiri, T. Maiti and K. L. Lu's (1999) derivation of a time series generalization of a cross-sectional model used in small-area estimation. The principle of estimating models under assumptions that are consistent with the appropriate interpretations of their coefficients applies not only to equation (6) but also to hierarchical Bayes models. The reason is that the coefficients of hierarchical Bayes models cannot be free of omitted-variable and measurement-error biases. The estimates of those coefficients may not be consistent unless appropriate corrections for such biases are applied to them.

conditioning variables.[13] This result explains why the addition of a single error term to a mathematical equation and the exclusion of the interaction terms on the right-hand side of equation (11) yield an inappropriate conditional expectation of r_t in the usual situations where measurement-error and omitted-variable biases are present and the "true" functional forms are unknown.[14]

The procedure of identifying the coefficients of equation (11) is the same as that of identifying the coefficients of generalized linear regression models. Following the latter procedure, we can find the conditions under which the coefficients of equation (11) are identifiable and consistently estimable. What is novel here is our method of identifying the time average ($\frac{1}{T}\sum_{t=1}^{T}\alpha_{jt}$) of each of the bias-free coefficients, α_{jt}, j = 1,…, K-1, in equation (3). We now present this method.[15] Let x_{0t} be equal to 1 for all t. Then the conditional expectation, $E(r_t \mid xs, zs)$, implied by equation (11) is equal to

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[13] Chang et al. (2000) develop a numerically stable algorithm for estimating equation (11) subject to equality and inequality constraints on its parameters. J. Thomas Yokum, Albert Wildt and Swamy (1998) show that equation (11) has better forecasting properties than some of its special cases.

[14] Two of the conditions under which equation (11) has Diewert's flexible functional form are (i) $z_{jt} = x_{jt}$ for all j and t and (ii) the distribution of ε_{jt} is degenerate at zero for j = 1,…, K-1 and all t (see W. Erwin Diewert and T.J. Wales 1987). Neither condition is consistent with the appropriate interpretations of the coefficients of equation (6), as we have shown above.

[15] Making a distributional assumption about the a_{jt} in (10) gives the distributions of the random coefficients of equation (6). Random coefficients are also called random parameters. In this regard, Zellner (1989) asks: "Is the distribution of the random parameters a prior or is it part of the model?" (p. 302). One can perform a Bayesian analysis of model (11) by putting a prior distribution on its parameters. Without a doubt this distribution is a prior. The appropriate likelihood function for this Bayesian analysis is given by equation (11), which cannot be derived without inserting (9) into (6). Therefore, the distribution of the random coefficients of equation (6) based on equation (9) is part of model (11). This does not mean that in specifying equation (9) we did not use our prior information in the form of the appropriate interpretations of the coefficients of equation (6). We doubt that coherent inferences can be obtained by putting a distribution on the coefficients of equation (6) that is inconsistent with the coefficients' appropriate interpretations.

$\sum_{j=0}^{K-1}\left(\sum_{m=0}^{p-1}\pi_{jm}z_{mt}\right)x_{jt}$, which, in view of equations (7) and (8), can be written as $\sum_{j=0}^{K-1}\left(\sum_{m\in P_j}\pi_{jm}z_{mt}+\sum_{m\notin P_j}\pi_{jm}z_{mt}\right)x_{jt}$. Using the connection, established by equation (7), between $E(r_t \mid xs, zs)$ and equation (3) gives $E(r_t^* \mid xs, zs)=\sum_{j=0}^{K-1}\left(\sum_{m\in P_j}\pi_{jm}z_{mt}\right)x_{jt}$.

Therefore, for $t = 1, 2, \ldots, T,\ j = 1, \ldots, K-1,$ and $m = 1, \ldots, p-1$:

$$\text{(i)}\ \frac{\partial E(r_t^* \mid xs, zs)}{\partial x_{jt}}=\sum_{m\in P_j}\pi_{jm}z_{mt};\ \text{(ii)}\ \frac{\partial^2 E(r_t^* \mid xs, zs)}{\partial x_{jt}\partial z_{mt}}=\pi_{jm}\ \text{ if } m\in P_j \text{ and } = 0 \text{ otherwise} \tag{12}$$

The zs included in equation (9) are all the right concomitants only if they completely explain all the variation in the coefficients of equation (6) and the derivatives in (12) have the right signs and the right time profiles.[16] If the coefficients of equation (11) are identifiable and consistently estimable, then so are these partial derivatives.

We would also like to identify and consistently estimate the bias-free component (α_{jt}) of the “biased” effect (γ_{jt}) of x_{jt} on r_t because it is the quantity of interest. We do not obtain this quantity unless we add the error term (ε_{j1t}) of equation (7) to the partial derivative in (12)(i). Unfortunately, this error term is not identified, even though the partial derivative is identifiable. For this reason, the bias-free coefficient, α_{jt}, is only partially identifiable. That is, it is identifiable except for its random error term. More generally, the coefficients of equation (6) and their bias-free components in each period are only partially identifiable because the error terms of equations (7) and (8) are not

[16]In a recent study, Kashyap and Stein (2000) test their hypotheses about the signs of the second- and third-order partial derivatives of a variable with respect to other variables specified by them. They do the usual: estimate these derivatives without correcting for omitted-variable and measurement-error biases. They also do not correct for the inaccuracies in their specified functional forms. Furthermore, these derivatives may not exist because some of the variables with respect to which they are taken are discrete. Swamy and Tavlas (2001b) compare Kashyap and Stein’s method with the method adopted in this paper. Unlike Kashyap and Stein (2000), Laurentius Marais and William Wecker (1998) do the unusual: correct for omitted-variables and measurement- error bias, but their method ignores the unknown functional-form problem.

identified. This shows that there are limits to what we can learn about equation (2) from equation (11).

A key item of interest from equation (2) is the bias-free component, α_{1t} (see, e.g., Clarida et al. 1998, p. 1037). Yet, if α_{1t} is only partially identifiable, what can we learn about this key item? We address this issue.

It follows from equation (10) that the distribution of ε_{jt} is degenerate at zero (or at a nonzero value) if ϕ_{jj} = 0 (or 1) and the distribution of a_{jt} is degenerate at 0. If, in addition, ε_{j1t} takes values in a small interval around zero with probability 1, then the partial derivative in (12)(i) gives a good approximation to α_{jt}. We give below a method of verifying these conditions. Whether or not these conditions hold depends on the number and the appropriateness of concomitants included in equations (7) and (8).

We estimate all the parameters of equation (11) using an iterative re-weighted generalized least squares (IRWGLS) method developed by Chang et al. (2000). In this procedure, the estimated covariance matrix of the error term, $\varepsilon_{0t} + \varepsilon_{1t}x_{1t} + \cdots + \varepsilon_{K-1,t}x_{K-1,t}$, changes from one iteration to the next.[17] Hence the term "re-weighted" appears in IRWGLS. The minimum-norm solutions of equations that connect the residuals of equation (11) to the error term of equation (9) are used to estimate the ε_{jt} (see Chang, Hallahan and Swamy 1992). The estimated ε_{jt} are used to estimate the error covariance matrix of equation (11). Some authors would rather use the convenient a priori value of 1 in place of the unknown "true" value of ϕ_{jj} than estimate ϕ_{jj} from sample data subject to the restrictions that (i) $-1 < \phi_{jj} < 1$ and (ii) the variance of a_{jt} is nonnegative (see, for example, Datta et al. 1999). Our view is that it is a mistake to use a convenient a priori value in place of the unknown "true" value of a parameter without knowing whether or

[17] In the applications of this procedure reported in the next section, convergence of the iterated scheme to one or more sets of parameter estimates did not require over 50 iterations.

not the a priori value is compatible with the sample information. The IRWGLS method can be used to verify the compatibility of the a priori value of 1 for ϕ_{jj}. Because of our restriction, $-1 < \phi_{jj} < 1$, equation (10) cannot cover the case where $\phi_{jj} = 1$ as a special case. However, our experience with the IRWGLS method tells us that for some models and some data sets, the estimates of some of the ϕ_{jj} would be as close to 1 as 0.995 and the estimates of some of the variances of a_{jt}'s would be as close to zero as 0.0001. The IRWGLS estimate of 0.995 for ϕ_{jj} may mean that the restriction $\phi_{jj} = 1$ is compatible with the sample data. This estimate together with the IRWGLS estimate of 0.0001 for the variance of a_{jt} may mean that the distribution of ε_{jt} is close to a degenerate distribution. These estimates imply that the concomitants included in equation (9) explain most of the variation in γ_{jt}.

Let $\hat{\pi}_{jm}$ and $\hat{\varepsilon}_{jt}$ be the IRWGLS estimators of π_{jm} and ε_{jt}, respectively. Then inserting them into equation (9) gives the estimate $\hat{\gamma}_{jt} = \sum_{m=0}^{p-1} \hat{\pi}_{jm} z_{mt} + \hat{\varepsilon}_{jt}$ of the "biased" effect of x_{jt} on r_t. We take $(1/T)\sum_{t=1}^{T} \sum_{m \in P_j} \hat{\pi}_{jm} z_{mt}$ as the IRWGLS estimator of $\bar{\alpha}_{jT} = (1/T)\sum_{t=1}^{T} \alpha_{jt}$, a time average of the bias-free component of γ_{jt}, provided that the sign and time profile of $\sum_{m \in P_j} \hat{\pi}_{jm} z_{mt}$ and the signs of the estimates of the second-order partial derivatives in (12)(ii) agree with our prior beliefs based on the relevant economic theory. Similarly, the IRWGLS estimator of an average of the sum of omitted-variable and measurement-error biases contained in the coefficients of equation (6) is $(1/T)\sum_{t=1}^{T} \sum_{m \notin P_j} \hat{\pi}_{jm} z_{mt}$. The condition under which the estimator $(1/T)\sum_{t=1}^{T} \sum_{m \in P_j} \hat{\pi}_{jm} z_{mt}$ is consistent is that $|(1/T)\sum_{t=1}^{T} \sum_{m \in P_j} \hat{\pi}_{jm} z_{mt} - (1/T)\sum_{t=1}^{T} \alpha_{jt}|$ converges in probability to zero as $T \to \infty$.

A necessary condition that the right concomitants are included in equations (7) and (8) is that the IRWGLS estimates of $\sum_{m \in P_j} \pi_{jm} z_{mt}$ and the IRWGLS estimates of the second-order derivatives in (12)(ii) have the right signs. The source of these signs is, of course, the economic theory that has suggested the variables to be included in equation (3).

III. MONETARY POLICY RULES IN THE UNITED STATES DURING 1960Q3-2000Q4

A. A Model of the Federal Reserve's Reaction Function

We assume that the following version of equation (6) adequately represents the monetary-policy reaction function for the U.S. Federal Reserve:

$$r_t = \gamma_{0t} + \gamma_{1t} x_{1t} + \gamma_{2t} x_{2t} + \gamma_{3t} x_{3t}, \qquad (13)$$

where r_t = the federal funds rate in percent, x_{1t} = inflation in percent = log $(P_t/P_{t-4}) \times 100$, P_t = GDP implicit price deflator for the U.S., x_{2t} = output gap = (deviation of the logarithm of U.S. GDP from the least squares estimates of a quadratic trend) $\times$ 100, x_{3t} = the exchange rate between the U.S. dollar and a basket of other currencies. It is assumed that for j = 0, 1. 2, 3,

$$\gamma_{jt} = \pi_{j0} + \pi_{j1} z_{1t} + \pi_{j2} z_{2t} + \pi_{j3} z_{3t} + \varepsilon_{jt} \qquad (14)$$

with z_{1t} = the reciprocal of the U.S. unemployment rate in percent, z_{2t} = 5-year moving average of inflation (x_{1t}), and z_{3t} = $\log(M_{t-1}/M_{t-5}) \times 100$, M_t = M2 measure of money for the U.S.,

$$\varepsilon_{jt} = \phi_{jj} \varepsilon_{jt-1} + a_{jt}, \qquad (15)$$

and t indexes quarters. Equations (13)-(15) are analyzed under Assumptions I and II with K and p equal to 4. The data used cover the period 1960Q3-2000Q4, and 1979Q3 is the period in which we believe that the Federal Reserve changed the course of its monetary policy.[18]

Equations (13) and (14) are much more elaborate than the equation suggested by the following argument of Taylor (1999, p. 323): "a function relating the interest rate to the price level and real output will still emerge if the money stock is not growing at a fixed rate, but rather responds in a systematic way to the interest rate or to real output; *the response of money will simply change the parameters of the relationship*" [our emphasis].

A necessary and sufficient condition for macroeconomic stability: A real rate can be obtained by subtracting $x_{1t}^{*} = E(\pi_{t+f} \mid \Omega_t)$ from both sides of equation (3) with K = 4, provided r_t^{*} represents the realized value of the federal funds rate in quarter t and all the right-hand side variables of equation (3) represent its determinants. Doing so gives

$$r_t^{*} - x_{1t}^{*} = \bar{r}_t - \alpha_{1t}\pi_t^{*} + (\alpha_{1t} - 1)x_{1t}^{*} + \alpha_{2t}x_{2t}^{*} + \alpha_{3t}x_{3t}^{*} + \sum_{\ell=4}^{n_t}\alpha_{\ell t}x_{\ell t}^{*} \tag{16}$$

All the coefficients of equation (16) are bias-free, as shown in Section IIC. Suppose that equation (16) is a central bank's reaction function. Then the central bank increases r_t^{*} to reduce inflation whenever it expects inflation to go up. Taylor (1999, p. 331) argues that this is a good policy if it brings about an increase in the real interest rate and a wrong policy if it brings about a fall in the real interest rate. Decreases in r_t^{*} that decrease the real interest rate represent good policy actions only in periods of decreasing inflation. Now we ask the question: What are the conditions under which equation (16) gives such good policy rules? The answer is as follows. The real rate, $r_t^{*} - x_{1t}^{*}$, increases if an

[18] We also experimented with several other specifications that excluded the exchange rate variable from equation (13), excluded the 5-year moving average of inflation and/or the reciprocal of the U.S. unemployment rate from equation (14), included in equation (13) additional explanatory variables including the lagged dependent variable, r_{t-1}, and included additional concomitants in equation (14). None of these experiments gave us plausible results. In particular, the inclusion of r_{t-1} with a fixed or variable coefficient on the right-hand side of equation (13) introduced high spurious inter-correlations among the time-varying coefficients. This experience shows that equation (13) gives spurious results if it has a lagged value of its dependent variable as one of its explanatory variables.

increase in r_t^* exceeds an increase in x_{1t}^* and decreases if a decrease in r_t^* is less than a decrease in x_{1t}^*. From equation (16) we can derive that $\Delta r_{t+1}^* - \Delta x_{1,t+1}^* =$

$$\Delta(\bar{r}_{t+1} - \alpha_{1,t+1}\pi_{t+1}^* + \alpha_{2,t+1}x_{2,t+1}^* + \alpha_{3,t+1}x_{3,t+1}^* + \sum_{\ell=4}^{n_t}\alpha_{\ell,t+1}x_{\ell,t+1}^*) + (\alpha_{1,t+1} - 1)x_{1,t+1}^* - (\alpha_{1t} - 1)x_{1t}^*,$$

where Δ is the difference operator, i.e., for any x_t, $\Delta x_{t+1} = x_{t+1} - x_t$. Therefore, an increase in r_t^* exceeds an increase in x_{1t}^* if and only if

$$\Delta(\bar{r}_{t+1} - \alpha_{1,t+1}\pi_{t+1}^* + \alpha_{2,t+1}x_{2,t+1}^* + \alpha_{3,t+1}x_{3,t+1}^* + \sum_{\ell=4}^{n_t}\alpha_{\ell,t+1}x_{\ell,t+1}^*) >$$
$$\Delta x_{1,t+1}^*[1 - \alpha_{1t} - (\Delta\alpha_{1,t+1})x_{1,t+1}^* / \Delta x_{1,t+1}^*]. \qquad (17)$$

Alternatively, a decrease in r_t^* is less than a decrease in x_{1t}^* if and only if the strict opposite of condition (17), i.e., condition (17) with its inequality sign reversed, holds. Condition (17) is satisfied if

(i) $\Delta(\bar{r}_{t+1} - \alpha_{1,t+1}\pi_{t+1}^* + \alpha_{2,t+1}x_{2,t+1}^* + \alpha_{3,t+1}x_{3,t+1}^* + \sum_{\ell=4}^{n_t}\alpha_{\ell,t+1}x_{\ell,t+1}^*) \geq 0$ and (ii)

$$[(\Delta\alpha_{1,t+1})x_{1,t+1}^* / \Delta x_{1,t+1}^*] + \alpha_{1t} > 1 \qquad (18)$$ [19]

The strict opposite of condition (17) holds if condition (18)(ii) and the opposite of condition (18)(i), i.e., condition (18)(i) with its '$\geq$' sign changed to '$\leq$', hold. An expected increase in inflation will bring about an increase in the real interest rate if it results in an increase in r_t^* that satisfies conditions (i) and (ii) in (18). Thus, in the case where the central bank's expectation of an increase in inflation comes true and its policy action of increasing r_t^* leads to the satisfaction of conditions (i) and (ii) in (18), the real rate adjusts to stabilize the economy. Using strong prior information, Taylor (1999) and Clarida et al. (1998, p. 1037), among others, assume that $\bar{r}_t$ and π_t^* are constant and that they know the dates at which the coefficients on the included explanatory variables in condition (18)(i) change. They also ignore the biases introduced by excluded variables,

[19] $(\alpha_{1,t+1} - 1)x_{1,t+1}^* - (\alpha_{1t} - 1)x_{1t}^* = \alpha_{1,t+1}x_{1,t+1}^* - \alpha_{1t}x_{1,t+1}^* + \alpha_{1t}x_{1,t+1}^* - \alpha_{1t}x_{1t}^* - \Delta x_{1,t+1}^*$ $= (\Delta\alpha_{1,t+1})x_{1,t+1}^* + \alpha_{1t}\Delta x_{1,t+1}^* - \Delta x_{1,t+1}^* > 0$ if and only if condition (18)(ii) with positive $\Delta x_{1,t+1}^*$ holds. Conditions (i) and (ii) in (18) are stronger than the necessary and sufficient condition (17).

measurement errors, and the inaccuracies in their specified functional forms. Consequently, in their case, the term $\Delta(\bar{r}_{t+1} - \alpha_{1,t+1}\pi^*_{t+1})$ in condition (18)(i) reduces to zero and condition (18)(ii) simplifies to $\alpha_1 > 1$ in the periods in which α_1 is a constant. Any policy that does not satisfy these simplified conditions need not necessarily be bad because Taylor's and Clarida et al.'s assumptions could be false. For example, in the general case, where (i) α_{1t} is a positive variable and (ii) $[(\Delta\alpha_{1,t+1})x^*_{1,t+1} / \Delta x^*_{1,t+1}] > 0$, not only the values of α_{1t} greater than 1, but also some of the values of α_{1t} less than 1 satisfy condition (18)(ii). What values of α_{1t} lying between 0 and 1 satisfy condition (18)(ii) depends on the value of $[(\Delta\alpha_{1,t+1})x^*_{1,t+1} / \Delta x^*_{1,t+1}$. The values of $[(\Delta\alpha_{1,t+1})x^*_{1,t+1} / \Delta x^*_{1,t+1}]$ for which some of the negative values of α_{1t} satisfy condition (18)(ii) may not occur. Even when condition (18)(ii) is satisfied, condition (18)(i) may not be satisfied. Condition (17) can be satisfied even when conditions (18)(i) and (18)(ii) are not satisfied. A decrease in expected inflation will bring about a decrease in the real interest rate if and only if it leads to a decrease in r^*_t that satisfies the strict opposite of (17). All these are clearly dynamic conditions where the bias-free effect on the federal funds rate of expected inflation varies over time, and from these conditions it is not obvious that a hallmark of "good monetary policy" is the satisfaction of only condition (18)(ii). These interpretations extend the previous studies' interpretation of fixed γ_1 to the variable α_{1t} with the "true" pattern of variation.

Good monetary policy rules: Increases in r^*_t represent good monetary-policy rules if and only if they satisfy condition (17) in periods of rising inflation. Alternatively, decreases in r^*_t represent good monetary-policy rules if and only if they satisfy the strict opposite of (17) in periods of falling inflation. We define policy mistakes as big departures from such good policy rules. This definition, though not operational, is the right one if omitted variables and measurement errors are present and the "true" functional forms of monetary-policy reaction functions are unknown, as they usually are. Also, from a satisfaction of condition (18)(ii) we cannot infer that a policy mistake has not been made because without condition (18)(i) (or its opposite) it is not a sufficient condition. Alternatively, from a failure of condition (18)(ii) we cannot conclude that a policy mistake has been made because with condition (18)(i) (or its opposite) it is not a necessary condition. We obtain a fallacious argument if we do so. Even if a central bank's actions satisfy condition (18)(ii), they may not satisfy condition (17) in periods of rising inflation or the strict opposite of (17) in periods of falling inflation. Taylor's (1969) claim is that a monetary policy that stays close to his two *baseline* monetary policy rules would be a good policy. It should be noted that the parameter values Taylor (1999) uses to define his baseline policy rules are not corrected for omitted-variables and measurement-error bias and for inaccuracies in his specified functional form. Therefore, in the next section, we use the method of Section II to correct for such biases and inaccuracies. At a minimum, the above argument shows that it is not easy to find the correct interpretation of monetary history. Condition (17) or its strict opposite contains too many unknowns and our methodology in Section IID gives only the estimates of the bias-free components of the coefficients on the explanatory variables included in equation (13). With these estimates, we cannot verify any of condition (17), its strict opposite, and condition (18)(i).

B. Empirical Results

Before proceeding to describe the estimates of the time-varying coefficients of equation (13), it is useful to first estimate equation (13) using the standard fixed-coefficient assumption. The least squares estimates of these coefficients are given in Table 1 for the entire sample period, 1960Q3-2000Q4, as well as for the two sub-periods, 1960Q3-1979Q2 and 1979Q3-2000Q4, described above. It is interesting to note that the coefficient on inflation more than doubles between the two periods, which is consistent with the view that the Federal Reserve under the chairmanship of Paul Volcker gave much greater weight to reducing inflation than did previous chairmen. There is also some evidence that the importance accorded to fluctuations in output in the conduct of monetary policy was lower in the second period, but the estimation results for the entire sample period indicate that deviations of output from potential played essentially no role in monetary policy process. The empirical results for the full sample also suggest that the foreign exchange value of the dollar did not play a role in monetary policy decisions. These results will change if we correct for omitted-variables and measurement-error bias and for the inaccuracies in the linearity of the fixed-coefficients version of equation (13). Furthermore, the statistical consistency of the least squares estimator used to obtain the estimates in Table 1 requires the conditions that contradict the appropriate interpretations of the coefficients of equation (13) given in Section IIC. For these reasons, we cannot stop our investigation here.

We turn now to the estimation results shown in Table 2 and Charts 1-5 for the time-varying coefficients of equation (13) and the fixed coefficients of equation (14). We obtained these results by applying the IRWGLS method to equations (13) and (14) under Assumptions I and II. The estimated time profile of the intercept (γ_{0t}) is given in Chart 1. This time profile does not permit the separation of the time profiles of the long-run equilibrium nominal federal funds rate ($\bar{r}_t$) and the Federal Reserve's target inflation rate (π_t^*) from those of the other components of γ_{0t}. The only conclusion that we can draw from Chart 1 is that during 1974Q3-1986Q1, $\hat{\gamma}_{0t}$ is negative and in all other quarters of our sample period, it is positive. This means that during some of the quarters of the period 1974Q3-1986Q1, π_t^* might have exceeded $\bar{r}_t$. The Federal Reserve's target inflation rate (π_t^*) implies a negative real rate if it exceeds the long-term equilibrium nominal rate ($\bar{r}_t$). A comparison of the estimated time profiles of the biased effects (γ_{jt}, j = 1, 2, 3) of expected inflation, expected output gap, and expected exchange rate is given in Chart 2. The signs of these biased effects are ambiguous, since they are contaminated by omitted-variable and measurement-error biases.

The IRWGLS estimate of ϕ_{jj} in equation (15) is equal to 0.995 for j = 0 and 1 and is equal to -0.995 and 0.932 for j = 2 and 3, respectively. The maximum of the IRWGLS estimates of the variances of the a_{jt} in equation (15) is equal to 0.00004. An implication of these estimates is that the three concomitants included in equation (14) explain most of the variation in the coefficients of equation (13), provided that ϕ_{jj} in equation (15) is not constrained to be zero.

In equation (16), the bias-free effects on the federal funds rate of the Federal Reserve's expectations about the future inflation, the future output gap, and the future foreign exchange value of the U.S. dollar are denoted by $\alpha_{1t}, \alpha_{2t},$ and α_{3t}, respectively. Economic theory predicts that the signs of α_{1t} and α_{2t} are positive and the sign of α_{3t} is negative.[20] These bias-free effects are the weights the Federal Reserve assigns to the expected values of inflation, output gap, and the dollar exchange rate when taking monetary-policy decisions. We do not know of any variables that have the same behavior over time as these weights. It is possible that none of the $\alpha_{1t}, \alpha_{2t},$ and α_{3t} has the same behavior over time as any observable variables. We believe that the dummy variable that takes the value zero before 1979Q3 and 1 during and after this quarter represents a very restrictive form of variation over time. This variation may not agree with the changes the Federal Reserve actually made in the $\alpha_{1t}, \alpha_{2t},$ and α_{3t} during our sample period. Any method that splits a sample period into a finite number of sub-periods and fits a different fixed-coefficients analog of equation (13) in each sub-period, not only ignores both omitted-variables and measurement-error bias and inaccuracies in the specified functional form of equation (13), but also makes the strong assumption that the dates at which the parameters of the fixed-coefficients analog of equation (13) have changed are exactly known. Such methods have questionable relevance to any past behavior and may have little or no relevance to any future behavior of the coefficients of equation (13). Therefore, we decided to experiment with several sets of concomitants and to accept only one of those sets that gave us the most plausible estimates of the bias-free effects with the right signs and to interpret the results carefully.

[20] For our measure of the exchange rate between the U.S. dollar and a basket of other currencies, an increase in the exchange rate means an appreciation of the dollar, which leads to a decline in the federal funds rate.

Based on the signs of the IRWGLS estimates of the coefficients of equation (14), we take the functions, $\pi_{11}z_{1t}$, $\pi_{20}+\pi_{22}z_{2t}+\pi_{23}z_{3t}$, and $\pi_{30}+\pi_{31}z_{1t}+\pi_{33}z_{3t}$, as our measures of the bias-free components of the coefficients (γ_{jt}, j = 1, 2, 3), respectively.[21] We could not take any other function of the terms on the right-hand side of equation (14) as a measure of the bias-free component of γ_{jt} because its IRWGLS estimate has the wrong sign in at least one quarter of the sample period and/or has an implausible time profile. Our measures of the bias-free effects are true if (i) the distributions of $\varepsilon_{11t}, \varepsilon_{21t}$, and ε_{31t} in equation (7) are tight around zero, (ii) the second-order partial derivatives, $\partial^2 E(r_t^* \mid x_{1t}, x_{2t}, x_{3t}, z_{1t}, z_{2t}, z_{3t})/\partial x_{1t}\partial z_{1t}$, $\partial^2 E(r_t^* \mid x_{1t}, x_{2t}, x_{3t}, z_{1t}, z_{2t}, z_{3t})/\partial x_{2t}\partial z_{2t}$, $\partial^2 E(r_t^* \mid x_{1t}, x_{2t}, x_{3t}, z_{1t}, z_{2t}, z_{3t})/\partial x_{2t}\partial z_{3t}$, $\partial^2 E(r_t^* \mid x_{1t}, x_{2t}, x_{3t}, z_{1t}, z_{2t}, z_{3t})/\partial x_{3t}\partial z_{1t}$, and $\partial^2 E(r_t^* \mid x_{1t}, x_{2t}, x_{3t}, z_{1t}, z_{2t}, z_{3t})/\partial x_{3t}\partial z_{3t}$ exist and are equal to π_{11}, π_{22}, π_{23}, π_{31}, and π_{33}, respectively, and (iii) the second-order partial derivatives of $E(r_t^* \mid x_{1t}, x_{2t}, x_{3t}, z_{1t}, z_{2t}, z_{3t})$ with respect to (x_{1t}, z_{2t}), (x_{1t}, z_{3t}), (x_{2t}, z_{1t}), and (x_{3t}, z_{2t}) exist and are equal to zero. Note that our measure of the bias-free component of the coefficient on inflation in equation (13) is based on the assumption that the weight the Federal Reserve attaches to inflation in its formulations of the targets for the federal funds rate is proportional to the reciprocal of the U.S. unemployment rate. The lower the unemployment rate, the higher the weight it gives to inflation. This assumption is reasonable in view of the available trade off between unemployment and inflation in the absence of shifts in the short-run Philips curve.

The IRWGLS estimates of the coefficients of the above partial derivatives are given in Table 2. The estimates of the second-order partial derivatives of $E(r_t^* \mid x_{1t}, x_{2t}, x_{3t}, z_{1t}, z_{2t}, z_{3t})$ with respect to (x_{1t}, z_{1t}) and (x_{2t}, z_{3t}) have the positive sign and those with respect to $(x_{2t}, z_{2t}), (x_{3t}, z_{1t})$, and (x_{3t}, z_{3t}) have the negative sign.

[21] Before adopting these measures, we tried several other measures that did not yield plausible results.

In Table 2, the IRWGLS estimates of π_{11}, π_{20}, π_{23}, and π_{30} are positive. Of these, only the estimate of π_{11} is significant. The IRWGLS estimates of π_{22}, π_{31}, and π_{33} are negative. All these estimates are insignificant. It can be seen from Charts 3-5 that these estimates imply the estimates of "bias-free" effects that have the right sign throughout the sample period.

Table 2 also gives the time averages of the estimates of the "bias-free" effects over the different sample periods, as well as the constant and the time-varying components of each average. Recall that the bias-free effect of an explanatory variable excludes the influence of omitted variables, measurement errors, and of the deviation of any specified functional form from the "true" functional form. Thus, the estimates of average "bias-free" effects given in Table 2 should be more accurate than the least squares estimates given in Table 1, provided our measures of the bias-free effects are appropriate. One argument that favors our measures is that there is no apparent contradiction between our measures and the appropriate interpretations of the coefficients of equation (13) given in Section IIC. There are also other differences between the estimates in Tables 1 and 2, which we now discuss.

Turning first to the "bias-free" effect of inflation on the federal funds rate, implied by the estimates in Table 2, it can be seen that on average, the IRWGLS estimate of this effect is significantly higher in the earlier period than in the later period, in contrast to the least squares results. Moreover, this average effect is greater than unity over the entire sample as well as in the two sub-samples. Put differently, the estimated average "bias-free" effects of inflation on the federal funds rate for the two sub-periods are significant, greater than 1, and are also significantly different from each other. The least squares estimates of the coefficient on inflation in Orphanide's (2001b) model of the Federal Reserve's reaction function for the two sub-periods are also greater than 1. The results that are similar to the least squares results in Table 1 are that the IRWGLS estimates of the average "bias-free" effects of the output gap have t-ratios that are slightly less than 2, with the magnitude of the reaction somewhat lower in the Volcker sub-period. Alternatively stated, the estimated average "bias-free" effects of the output gap on the federal funds rate for the two sub-periods are significant at the 10% level and are not significantly different from each other. Finally, the results in Table 2 and Chart 5 indicate the expected average "bias-free" effect of exchange rate changes, but the effect is quite small and insignificant. With the exception of the second sub-period, this finding is similar to that reported in Table 1 for the least squares results. The least squares estimate of γ_3 for the second sub-period has the wrong sign.

Chart 3 shows the variation over time in the coefficient on inflation in equation (13). As discussed above, the IRWGLS estimation procedure yields two estimates for this coefficient in each period: the *biased* effect, which includes omitted-variable biases and the impacts of measurement errors and misspecifications of the "true" functional form of equation (13), and the *bias-free* effect, which is purged of these factors. It is readily apparent from Chart 3 that omitted-variables and measurement-error bias and misspecifications of the "true" functional form generate a substantial bias in the estimate of the response to inflation on the part of the Federal Reserve, which appears to be particularly pronounced during the 1960s and the period 1984-2000. Thus, if measurement errors, omitted variables, and misspecifications of the "true" functional form are not taken into account in the estimation of the monetary policy reaction function, one can get a seriously distorted impression of the responsiveness of the Federal Reserve to inflation.

As discussed above, the "bias-free" effect of inflation on the federal funds rate is estimated as a single time-varying term, which is a fixed coefficient, times the reciprocal of quarterly unemployment rate (z_{1t}). Thus, in our model of the Federal Reserve's reaction function, a decrease in the unemployment rate increases this effect. By virtue of this definition, the time profile of the estimated "bias-free" effect of inflation on the federal funds rate is the same as that of the reciprocal of the unemployment rate. We thus associate the decreasing values of the unemployment rate with the increasing estimates of this effect, and vice versa. What this means is that the weight, which the Federal Reserve attaches to inflation when taking the monetary-policy decisions, varies inversely with the unemployment rate. An increase in money growth over time could raise inflation.

Therefore, a central bank that is concerned about inflation will not allow money growth to continue to increase over time, as this will be inflationary. Now, using the reciprocal of the unemployment rate as a single determinant of the "bias-free"-effect component of the coefficient on inflation makes the IRWGLS estimates of this effect exceed 1 throughout our sample period, as can be seen from Chart 3. From the discussion in Section IIIA it follows that these greater than 1 estimates of α_{1t} do not necessarily mean that policy mistakes have not been made at any time during our sample period. These estimates trended upward during 1961Q2-1969Q1, 1971Q4-1973Q4, and 1975Q2-1979Q2. During most of the quarters of these periods inflation was rising (Chart 3). Under a more responsive monetary policy, inflation would not have risen as much as it did during 1963Q3-1970Q2, 1972Q3-1975Q1, and 1976Q4-1981Q1. Taylor (1999) could be right in saying that the monetary policy was too tight during the early 1960s and was too easy during the late 1960s and 1970s. The factors, such as increased government spending arising from the Vietnam War and the two oil price shocks in the 1970s, were the driving forces behind rising inflation during the late 1960s and 1970s.

Chart 3 shows that as inflation was brought under control in the 1980s, the estimates of the "bias-free" effect of inflation declined during 1979Q3-1983Q1 and 1989Q2-1992Q3. These estimates increased during 1983Q2-1989Q1, in which period inflation first decreased and then increased. The monetary tightness followed during the period 1983Q2-1986Q2 could be excessive because during this period inflation was falling. Chart 3 also shows that during 1992Q3-2000Q4 the estimated "bias-free" effect of inflation rose sharply even though inflation increased very little over this period.

Turning now to Chart 4, which shows the estimated biased and "bias-free" effects of the output gap on the fed funds rate, we see again that the presence of omitted variables, mismeasurements, and misspecifications of the "true" functional form generates a substantial bias in the estimated impact of the output gap on the interest rate. This may reflect the significant mismeasurement of the output gap during the 1970s and early 1980s, as documented by Orphanides (2001b); the output gap was thought at the time to be much larger than revised data subsequently indicated. It should also be noted that the estimated "bias-free" effect of the output gap on the interest rate is not only much larger than the biased effect, but is also much smoother. This estimated "bias-free" effect increased during 1960Q3-1964Q1, 1981Q4-1987Q2, and 1993Q3-1999Q4. It trended downward during 1964Q2-1981Q3 and 1990Q3-1993Q2. Some of the increases in the estimated "bias-free" effect of the output gap on the interest rate are consistent with the fact that inflation was brought under control during the 1980s and 1990s. These results are produced by our assumption that the function $\pi_{20} + \pi_{22}z_{2t} + \pi_{23}z_{3t}$ provides an adequate approximation to the bias-free component of the coefficient on the output gap.

It may be wrong to exclude the exchange rate variable from equation (13) because its estimated "bias-free" effects on the fed funds rate are small and insignificant. The reason is that it is the conjunction of the variables in equations (13) and (14) that produced the results in Table 2 and Charts 1-5. These results changed when we changed this conjunction. In any case, the rationale for dropping a variable because its coefficient estimate happened to have a low *computed* t-ratio in a given sample is ordinarily very weak, as Pratt and Schlaifer (1984, 1988) point out.

IV. Conclusions

The biasing effects of measurement errors, omitted variables, and misspecifications of "true" functional forms are a pervasive problem in econometrics. A necessary and sufficient and a sufficient condition under which a monetary-policy rule is a good policy can only be stated in terms of the bias-free effects on the federal funds rate of a number of variables including the Federal Reserve's expectations about future inflation, the future gap between actual and potential output, the future foreign exchange value of the dollar, etc. Not all of these "bias-free" effects are identifiable on the basis of the available data. The estimates of the identifiable "bias-free" effects presented in this paper leave open the possibility that the Federal Reserve may have made policy mistakes during 1960Q3-2000Q4. Under a more responsive monetary policy, inflation would not have risen as much as it did during 1963Q3-1970Q2, 1972Q3-1975Q1, and 1976Q4-

1981Q1. The paper also shows that it is not easy to find the correct interpretation of monetary history.

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Table 1
Least Squares Estimates of a Federal Reserve Interest Rate Reaction Function with Fixed Coefficients
Dependent variable = the U.S. Federal Funds rate

Coefficient	Coefficient Estimates		
	1960Q3-1979Q2	1979Q3-2000Q4	1960Q3-2000Q4
γ_0	4.8105 (1.5565)	-3.6671 (-3.9612)	3.8055 (2.8364)
γ_1	0.6176 (6.7111)	1.3442 (21.294)	0.9530 (12.593)
γ_2	0.2811 (5.8505)	0.1742 (3.5194)	0.0051 (0.0890)
γ_3	-0.0172 (-0.7957)	0.0600 (7.1183)	-0.0075 (-0.6972)

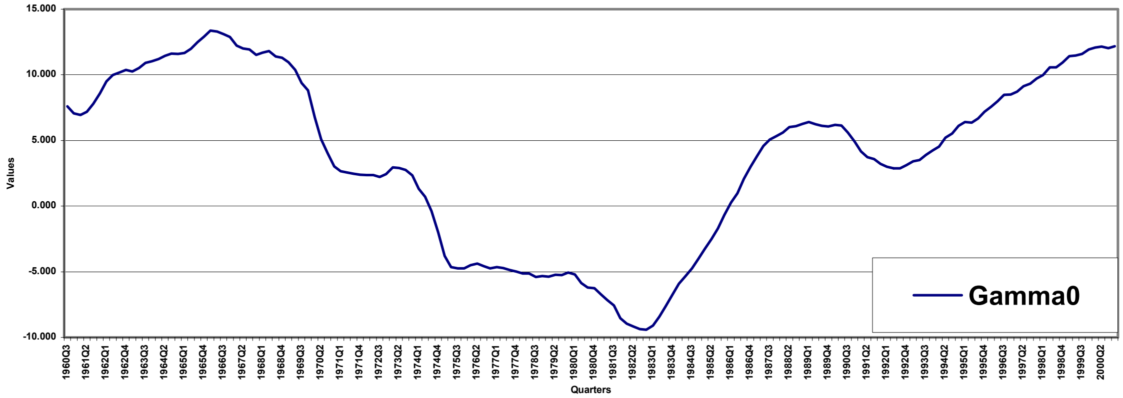
Note: Below each coefficient estimate is its t-ratio shown in parentheses. The coefficient γ_0 is the intercept, γ_1 is the coefficient on U.S. inflation (x_{1t}), γ_2 is the coefficient on U.S. output gap (x_{2t}), and γ_3 is the coefficient on the exchange rate between the U.S. dollar and a basket of other currencies (x_{3t}).

Table 2
IRWGLS Estimates of a Federal Reserve Interest Rate Reaction Function with Time-Varying Coefficients
Dependent variable = the U.S. Federal Funds rate

Average Bias-Free Effects, Their Coefficients, and Their Differences Between the Periods 1960Q3-1979Q2 and 1979Q3-2000Q4	Estimates		
	1960Q3-1979Q2	1979Q3-2000Q4	1960Q3-2000Q4
$\bar{\alpha}_{1T_i}{}^{a} = \hat{\pi}_{11}\left(\frac{1}{T_i}\sum_{t=1}^{T_i} z_{1t}\right)$	2.4933 (4.2157)	2.1308 (4.2157)	2.3008 (4.2157)
$\hat{\pi}_{11}$			12.8800 (4.2157)
$\bar{\alpha}_{1T_1} - \bar{\alpha}_{1T_2}$	0.3625 (4.2157)		
$\bar{\alpha}_{2T_i}{}^{b} = \hat{\pi}_{20} + \hat{\pi}_{22}\left(\frac{1}{T_i}\sum_{t=1}^{T_i} z_{2t}\right) + \hat{\pi}_{23}\left(\frac{1}{T_i}\sum_{1}^{T_i} z_{3t}\right)$	0.6368 (1.9342)	0.6027 (1.9777)	0.6187 (1.9624)
$\hat{\pi}_{20}$			0.7987 (1.8498)
$\hat{\pi}_{22}$			-0.0567 (-1.1570)
$\hat{\pi}_{23}$			0.0062 (0.2927)
$\bar{\alpha}_{2T_1} - \bar{\alpha}_{2T_2}$	0.0341 (0.6193)		
$\bar{\alpha}_{3T_i}{}^{c} = \hat{\pi}_{30} + \hat{\pi}_{31}\left(\frac{1}{T_i}\sum_{t=1}^{T_i} z_{1t}\right) + \hat{\pi}_{33}\left(\frac{1}{T_i}\sum_{t=1}^{T_i} z_{3t}\right)$	-0.0832 (-1.5369)	-0.0656 (-1.1615)	-0.0738 (-1.3429)
$\hat{\pi}_{30}$			0.0216 (0.1969)
$\hat{\pi}_{31}$			-0.4388 (-0.9599)
$\hat{\pi}_{33}$			-0.0025 (-0.6383)
$\bar{\alpha}_{3T_1} - \bar{\alpha}_{3T_2}$	-0.0176 (-1.3441)		

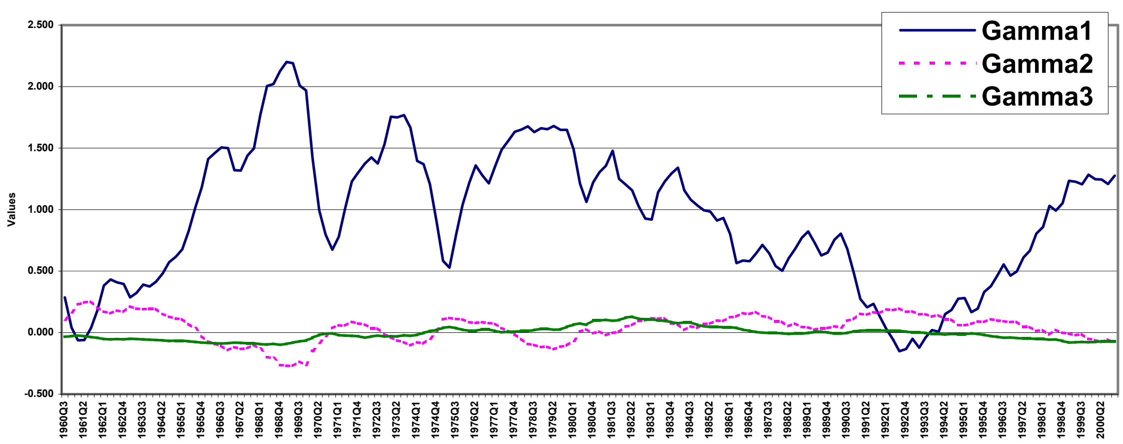
[a] $\bar{\alpha}_{1T_i}$ = Average bias-free component of γ_{1t}, the coefficient on U.S. inflation; $T_i = T_1, T_2$, or T_3, where T_1, T_2, and T_3 are the numbers of observations in the periods 1960Q3-1979Q2, 1979Q3-2000Q4, and 1960Q3-2000Q4, respectively. [b] $\bar{\alpha}_{2T_i}$ = Average bias-free component of γ_{2t}, the coefficient on U.S. output gap. [c] $\bar{\alpha}_{3T_i}$ = Average bias-free component of γ_{3t}, the coefficient on the exchange rate between the U.S. dollar and a basket other currencies. Below each coefficient estimate is its t-ratio shown in parentheses.

Chart 1: Time Profile of the Estimated Intercept of a Federal Reserve Interest Rate Reaction Function



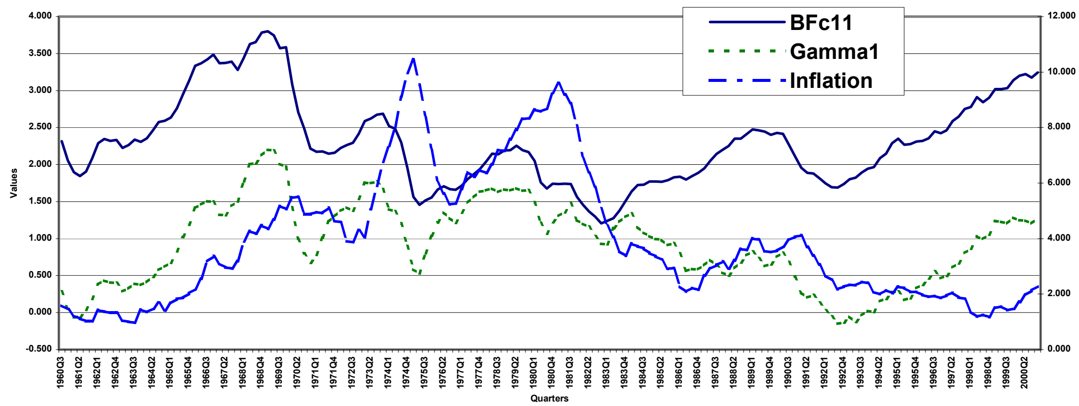
Note: $\hat{\gamma}_{0t}$ = the estimated intercept.

Chart 2: Time Profiles of the Estimated "Biased" Effects of Inflation, Output Gap, and the Exchange Rate on the Federal Funds Rate for the U.S.



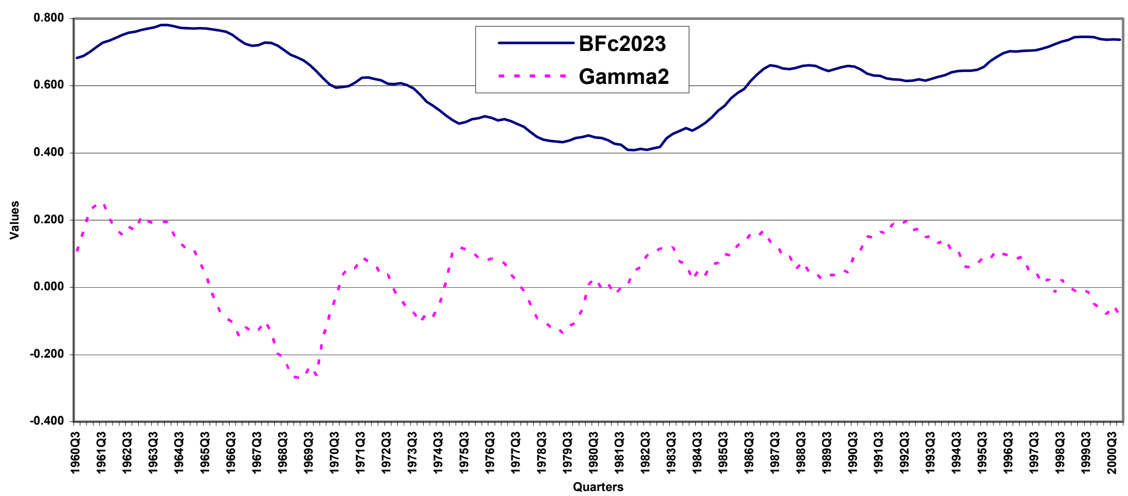
Note: Gamma1 = $\hat{\gamma}_{1t}$ = the estimated "biased" effects of inflation, Gamma2 = $\hat{\gamma}_{2t}$ = the estimated "biased" effects of output gap, and Gamma3 = $\hat{\gamma}_{3t}$ = the estimated "biased" effects of the U.S. dollar exchange rate.

Chart 3: Time Profiles of U.S. Inflation and Its Estimated “Biased” and “Bias-Free” Effects on the U.S. Federal Funds Rate



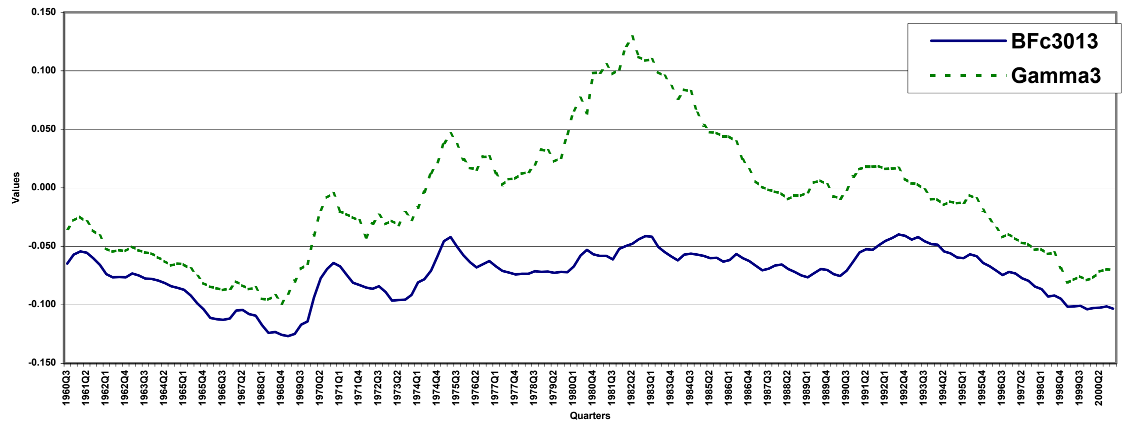
Note: Gamma1 = $\hat{\gamma}_{1t}$ = the estimated “biased” effects and BFc11 = $\hat{\pi}_{11}$(1/U.S. unemployment rate) = the estimated “bias-free” effects.

Chart 4: Time Profiles of the Estimated “Biased” and “Bias-Free” effects of U.S. Output Gap on the U.S. Federal Funds Rate



Note: Gamma2 = $\hat{\gamma}_{2t}$ = the estimated “biased” effects and BFc2023 = $\hat{\pi}_{20} + \hat{\pi}_{22}$ 5yrMAInf + $\hat{\pi}_{23}\log(M2_{t-1}/M2_{t-5})100$ = the estimated “bias-free” effects.

Chart 5: Time Profiles of the Estimated "Biased" and "Bias-Free" Effects of the Exchange Rate between the U.S. Dollar and A Basket of Other Currencies on the U.S. Federal Funds Rate



Note: Gamma3 = $\hat{\gamma}_{3t}$ = the estimated "biased" effects and BFc3013 = $\hat{\pi}_{30} + \hat{\pi}_{31}$(1/U.S. unemployment rate) + $\hat{\pi}_{33}\log(M2_{t-1}/M2_{t-5})100$ = the estimated "bias-free" effects.