MEASURING MONETARY POLICY

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Abtract

We develop and implement a general, VAR-based methodology for measuring innovations in the stance of monetary policy and their effect on the economy. Within this framework we are able to (1) perform statistical comparisons of existing proposed policy indicators, (2) derive a new indicator which is "optimal", in the sense of being most consistent with estimates of the central bank's operating procedures, and (3) compare and analyze the dynamic responses of economic variables to alternative measures of policy shocks. We also propose a new measure of the overall stance of policy (including the endogenous or anticipated part) which is consistent with our approach.

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I. Introduction

Accurate measurement of the effects of changes in monetary policy on the economy is essential, both for good policy-making and for choosing among alternative macroeconomic theories. Unfortunately, attempts to quantify the links between central bank actions and the economy quickly run into a major roadblock: There is no consensus on how to measure the size and direction of changes in monetary policy. The traditional approach, which identifies changes in monetary policy with changes in the stock of money, is not adequate, since in practice the growth rates of monetary aggregates depend on a variety of non-policy influences. For example, because the Fed's operating procedures have typically involved some smoothing of short-term interest rates, and hence accommodation of money demand shocks, observed money growth rates in the U.S. reflect changes in money demand as well as changes in money supply.[1] Secular changes in velocity brought about by financial innovation, deregulation, and other factors are a further barrier to using money growth rates alone as a measure of the direction of policy.

As the deficiencies of money stock growth as a measure of the stance of monetary policy have become widely recognized, many researchers have tried to find alternative indicators. Recent attempts have largely fallen into two general categories: First, following the example of Friedman and Schwartz (1963), Romer and Romer (1989) re-introduced the "narrative approach" to the study of monetary policy. Based on a reading of the minutes of the Federal Open Market Committee, Romer and Romer determined a set of dates at which policy-makers appeared to shift to a more anti-inflationary stance. An appealing aspect of the Romers' approach is that it uses additional information--specifically, policy-makers' statements of their own intentions--to try to disentangle money supply from money demand shocks. A disadvantage of this approach, besides inherent problems of reproducibility and subjectivity, is that it does not clearly distinguish between the endogenous and exogenous components of policy change (Dotsey and Reid, 1992; Leeper, 1993; Sims and Zha, 1993; Shapiro, 1994; Hoover and Perez, 1994), which is necessary for identifying the effects of monetary policy on the economy. The Romers' methodology also yields a rather limited amount of information: They give dates only for contractionary shifts, not expansionary shifts, and their method provides no indication of the severity or duration of each episode. Building on the Romers' work, Boschen and Mills (1991) used FOMC documents to rate monetary policy in each month as "very tight", "tight", "neutral", "easy", or "very easy", depending on the relative weights the policy-makers assigned to reducing unemployment and reducing inflation. Although Boschen and Mills provide a more continuous and possibly more informative measure of policy than do Romer and Romer, their indicator likely also suffers relatively more severe problems of subjectivity and policy endogeneity.

The second general strategy for measuring monetary policy stance--which is the focus of the present article--is to use prior information about central bank operating procedures, in conjunction with vector autoregressive (VAR) estimation techniques, to develop data-based indexes of policy. For example, Bernanke and Blinder (1992) argued that over much of the past thirty years the Fed has implemented policy changes through changes in the federal funds rate (the overnight rate in the market for commercial bank reserves). They concluded that the funds rate may therefore be used as an indicator of policy stance (see also Laurent, 1988, and Bernanke, 1990); in particular, they interpreted VAR innovations to the funds rate as innovations to the Fed's policy. In a similar vein, Sims (1992) used short-term rates as monetary indicators in a multi-country study. Not all researchers working in the VAR-based literature have adopted short-term interest rates as their preferred indicator of policy, however: Following a suggestion of Thornton (1988b), Christiano and Eichenbaum (1992) have made the case for using the quantity of nonborrowed reserves as the primary measure of monetary policy (see also Eichenbaum, 1992). Strongin (1995) proposed as a policy indicator the portion of nonborrowed reserve growth that is orthogonal to total reserve growth: He motivated this measure by arguing that the Fed is constrained to meet total reserve demand in the short run but can effectively tighten policy by reducing nonborrowed reserves and forcing banks to borrow more from the discount window. Cosimano and Sheehan (1994) characterized Fed policy after 1984 as borrowed-reserves targeting, which suggests that borrowed reserves might be a useful indicator for the more recent period.

Both the narrative and VAR-based methods for measuring monetary policy have been widely used in applied work.[2] Unfortunately, there is evidently little agreement on which of the various measures most accurately captures the stance of policy, leading many authors to hedge by using a variety of indicators. Eichenbaum and Evans's (1995) study of the effect of monetary policy on exchange rates is fairly typical in employing three alternative policy measures--in their case, Strongin's measure, innovations to the federal funds rate, and the Romer dates. However, although using alternative measures allows the researcher to claim robustness when the results for each indicator are similar, this strategy provides no guidance for cases when the results for different indicators are inconsistent. (Indeed, we show below that alternative indicators can lead to quite different inferences.) Moreover, simply using a variety of alternative measures of monetary policy cannot guarantee that some more accurate indicator has not been excluded; that the best indicator is not perhaps some combination of the various "pure" indicators; or that the best indicator is the same for all countries or for all periods. Thus it would be quite useful to have a systematic method of comparing alternative candidate indicators of policy.

Eichenbaum (1992, p. 1010) has stressed the importance of finding a means of choosing among indicators, noting that in his particular application "inference depends very sensitively on which of the two candidate measures [short-term interest rates or nonborrowed reserves] we work with." He also suggests that "further progress on these issues can be made only by carefully studying the institutional details of how monetary policy is actually carried out in the different countries..." Following Eichenbaum's suggestion, in this article we develop and implement a general, VAR-based methodology in which the indicator of monetary policy stance is not assumed but is rather derived from an estimated model of the central bank's operating procedure. More specifically, we employ a "semi-structural" VAR model which leaves the relationships among macroeconomic variables in the system unrestricted but imposes contemporaneous identification restrictions on a set of variables relevant to the market for commercial bank reserves.

Our method has several advantages over previous approaches: First, because our specification nests the best-known quantitative indicators of monetary policy used recently in VAR modelling, including all those mentioned above, we are able to perform explicit statistical comparisons of these and other potential measures, including hybrid measures that combine the basic indicators. Second, our analysis leads directly to estimates of a new policy indicator that is "optimal", in the sense of being most consistent with the estimated parameters describing the central bank's operating procedure and the market for bank reserves. Third, by estimating the model over different sample periods we are able to allow for changes in the structure of the economy and in operating procedures, while imposing a minimal set of identifying assumptions. Finally, although we consider only the post-1965 U.S. case in this paper, our method is applicable to other countries and periods, and to alternative institutional setups.[3]

A frequently-heard criticism of the VAR-based approach is that it focuses on monetary policy innovations rather than the arguably more important systematic or endogenous component of policy. We believe this criticism to be misplaced: The emphasis of the VAR-based approach on policy innovations arises not because shocks to policy are intrinsically important, but because (as we discuss further below) tracing the dynamic response of the economy to a monetary policy innovation provides a means of observing the effects of policy changes under minimal identifying assumptions. However, although we disagree with the view that analysis of policy shocks is uninteresting, we also recognize that it would be useful to have an indicator of the overall stance of monetary policy, including the endogenous component. An additional contribution of this article is to propose just such an indicator, one which is both consistent with our underlying approach and, we believe, intuitively appealing.

The rest of the article proceeds as follows. Section II briefly describes our general methodology for identifying innovations to monetary policy. Section III lays out a standard model of the market for bank reserves that nests some common alternative descriptions of Fed operating procedures. Estimation of this model by GMM (Section IV) allows us both to evaluate the leading candidate indicators of policy innovations and to develop an alternative measure. Section V discusses how the choice of policy measure affects our conclusions about the impact of monetary policy on the economy. Section VI introduces our total policy measure, inclusive of both the systematic and random components of policy, and compares it to narrative measures of monetary policy. Section VII concludes.

II. Methodology

Bernanke and Blinder (1992) proposed the following strategy for measuring the dynamic effects of monetary policy shocks. Suppose the "true" economic structure is

$$(1) \quad \mathbf{Y}_t = \sum_{i=0}^{k} \mathbf{B}_i \mathbf{Y}_{t-i} + \sum_{i=0}^{k} \mathbf{C}_i p_{t-i} + \mathbf{A}^y \mathbf{v}_t^y$$

$$(2) \quad p_t = \sum_{i=0}^{k} \mathbf{D}_i \mathbf{Y}_{t-i} + \sum_{i=1}^{k} g_i p_{t-i} + v_t^p$$

Equations (1) and (2) define an unrestricted linear dynamic model which allows both contemporaneous values and up to k lags of any variable to appear in any equation.[4] Boldface letters are used to indicate vectors or matrices of variables or coefficients. In particular, $\mathbf{Y}$ is a vector of macroeconomic variables, and p is a variable indicating the stance of policy. Note that for the moment p is taken to be a scalar measure, e.g., the federal funds rate. Equation (2) predicts current policy stance given current and lagged values of macroeconomic variables and lagged policy variables, while equation (1) describes a set of structural relationships in the rest of the economy. The vector $\mathbf{v}^y$ and the scalar v^p are mutually uncorrelated "primitive" or "structural" error terms. As in Bernanke (1986), the structural error terms in equation (1) are pre-multiplied by a general matrix $\mathbf{A}^y$, so that shocks may enter into more than one equation; hence the assumption that the elements of $\mathbf{v}^y$ are uncorrelated imposes no restriction. The assumption that the policy shock v^p is uncorrelated with the elements of $\mathbf{v}^y$ is also not restrictive, in our view; we think of independence from contemporaneous economic conditions as part of the definition of an exogenous policy shock.[5]

The system (1)-(2) is not econometrically identified in general. Bernanke and Blinder point out that to identify the dynamic effects of exogenous policy shocks on the various macro variables $\mathbf{Y}$, *without* necessarily having to identify the entire model structure, it is sufficient to assume that policy shocks do not affect the given macro variables within the current period, i.e., $\mathbf{C}_0 = 0$.[6] Under this assumption the system (1)-(2) can be written in VAR format by projecting the vector of dependent variables on k lags of itself. Estimation of the resulting system by standard VAR methods, followed by a Choleski decomposition of the covariance matrix (with the policy variable ordered last) yields an estimated series for the exogenous policy shock v^p (see Bernanke-Blinder). Impulse response functions for all variables in the system with respect to the policy shock can then be calculated, and can be interpreted as the true structural responses to policy shocks.

The Bernanke-Blinder method assumes that a good scalar measure of policy is available. However, it may be the case that we have only a *vector* of policy indicators, $\mathbf{P}$, which contain information about the stance of policy but are affected by other forces as well. For example, if the Fed's operating procedure is neither pure interest-rate targeting nor pure reserves targeting, then both interest rates and reserves will contain information about monetary policy; but in that case, both variables may also be affected by shocks to the demand for reserves and other factors. In this more general case the structural macroeconomic model (1)-(2) may be written:

$$(3) \quad \mathbf{Y}_t = \sum_{i=0}^{k} \mathbf{B}_i \mathbf{Y}_{t-i} + \sum_{i=1}^{k} \mathbf{C}_i \mathbf{P}_{t-i} + \mathbf{A}^y \mathbf{v}_t^y$$

$$(4) \quad \mathbf{P}_t = \sum_{i=0}^{k} \mathbf{D}_i \mathbf{Y}_{t-i} + \sum_{i=0}^{k} \mathbf{G}_i \mathbf{P}_{t-i} + \mathbf{A}^p \mathbf{v}_t^p$$

Equation (4) states that the set of policy indicators $\mathbf{P}$ depend on current and lagged values of $\mathbf{Y}$ and $\mathbf{P}$, and on a set of disturbances $\mathbf{v}^p$. We assume that one element of the vector $\mathbf{v}^p$ is a money supply shock or policy disturbance v^s; the other elements of $\mathbf{v}^p$ may include shocks to money demand or whatever disturbances affect the policy indicators. Equation (3) allows the non-policy variables $\mathbf{Y}$ to depend on current and lagged values of $\mathbf{Y}$ and on lagged values (only) of $\mathbf{P}$; allowing the non-policy variables to depend only on lagged values of policy variables ($\mathbf{C}_0 = 0$) is analogous to the identifying assumption made above in the scalar case. There is no need to impose any other restrictions on the coefficients $\mathbf{C}_i$, $i > 0$, as the non-policy variables could depend on the policy indicators through mechanisms not mediated through the stance of policy.

In analogy to the case of a scalar policy indicator, we would like to find a way to measure the dynamic responses of variables in the system to a policy shock v^s. As before, we can re-write the system (3)-(4) in VAR form (with only lagged variables on the right-hand side) and estimate by standard methods. Let $\mathbf{u}_t^p$ be the portion of the VAR residuals in the policy block that are orthogonal to the residuals in the non-policy block. Then straightforward calculation shows that $\mathbf{u}_t^p$ satisfies

$$\text{(5)} \quad \mathbf{u}_t^p = (\mathbf{I} - \mathbf{G}_0)^{-1} \mathbf{A}^p \mathbf{v}_t^p$$

where the variables on the right-hand side of (5) are as defined in (4). Alternatively, dropping subscripts and superscripts, we can rewrite (5) as:

$$\text{(6)} \quad \mathbf{u} = \mathbf{G}\mathbf{u} + \mathbf{A}\mathbf{v}$$

Equation (6) is a standard structural VAR (SVAR) system, which relates observable VAR-based residuals $\mathbf{u}$ to unobserved structural shocks $\mathbf{v}$, one of which is the policy shock v^s. This system can be identified and estimated by conventional methods, allowing recovery of the structural shocks, including v^s. The policy shock v^s is analogous to the innovation to the federal funds rate in the scalar case analyzed by Bernanke and Blinder (1992). As in the scalar case, the structural responses of all variables in the system to a policy shock can be measured by the associated impulse response functions. Further, the historical sequence of policy shocks can be recovered from the VAR residuals by means of (6). In the remainder of this article we apply this approach to the measurement of U.S. monetary policy stance since 1965.

III. Monetary policy and the market for bank reserves

To implement the methodology described in Section II, we need a specific model relating the VAR residuals and the structural shocks in the policy block. We employ a standard model of the market for commercial bank reserves and Federal Reserve operating procedures.[7] Although simple, this model is rich enough to nest all the VAR-based policy indicators mentioned in the introduction, as well as other plausible measures.

Continuing to use u to indicate an (observable) VAR residual and v to indicate an (unobservable) structural disturbance, we assume that the market for bank reserves is described by the following set of equations:

(8) $$u_{TR} = -\alpha u_{FF} + v^{d}$$

(9) $$u_{BR} = \beta(u_{FF} - u_{DISC}) + v^{b}$$

(10) $$u_{NBR} = \phi^{d} v^{d} + \phi^{b} v^{b} + v^{s}$$

Equation (8) is the banks' total demand for reserves, expressed in innovation form; it states that the innovation in the demand for total reserves u_{TR} depends (negatively) on the innovation in the federal funds rate u_{FF} (the price of reserves) and on a demand disturbance v^{d}. Equation (9) determines the portion of reserves that banks choose to borrow at the discount window: As is conventional, the demand for borrowed reserves (in innovation form), u_{BR}, is taken to depend positively on the innovation in the federal funds rate u_{FF} (the rate at which borrowed reserves can be relent) and negatively on the discount rate u_{DISC} (the cost of borrowed reserves); v^{b} is a disturbance to the borrowing function.[8] The innovation in the demand for nonborrowed reserves, the difference between total and borrowed reserves, is $u_{TR} - u_{BR}$.

Equation (10) describes the behavior of the Federal Reserve. We assume that the Fed observes and responds to shocks to the total demand for reserves and to the demand for borrowed reserves within the period, with the strength of the response given by the coefficients ϕ^{d} and ϕ^{b}. That the Fed observes reserve demand shocks within the period is reasonable, since it monitors total reserves (except vault cash) and borrowings continuously; however, the case in which the Fed does not observe (or does not respond to) one or the other of these disturbances can be accommodated by setting the relevant coefficients to zero. The disturbance term ν^s is the shock to policy that we are interested in identifying. Note that the system (8)-(10) is in the form of equation (6).

It will also be useful to write the reduced-form relationship between the VAR residuals **u** and the structural disturbances **v**, as in equation (5). To do so, we first make the simplifying assumption that the innovation to the discount rate u_{DISC} is zero.[9] To solve the model we impose the condition that the supply of nonborrowed reserves plus borrowings must equal the total demand for reserves. Solving in terms of innovations to total reserves, nonborrowed reserves, and the federal funds rate, we have

(11) $\mathbf{u} = (\mathbf{I}-\mathbf{G})^{-1}\mathbf{A}\mathbf{v}$, where

$\mathbf{u}' = \begin{vmatrix} u_{TR} & u_{NBR} & u_{FF} \end{vmatrix}$ $\qquad$ $\mathbf{v}' = \begin{vmatrix} \nu^d & \nu^s & \nu^b \end{vmatrix}$ $\qquad$ and

$$(\mathbf{I}-\mathbf{G})^{-1}\mathbf{A} = \begin{vmatrix} -(\frac{\alpha}{\alpha+\beta})(1-\phi^d)+1 & \frac{\alpha}{\alpha+\beta} & (\frac{\alpha}{\alpha+\beta})(1+\phi^b) \\ \phi^d & 1 & \phi^b \\ (\frac{1}{\alpha+\beta})(1-\phi^d) & -\frac{1}{\alpha+\beta} & -(\frac{1}{\alpha+\beta})(1+\phi^b) \end{vmatrix}$$

One can also invert the relationship (11) to determine how the monetary policy shock ν^s depends on the VAR residuals:

$$(12) \qquad \nu^s = -(\phi^d + \phi^b)u_{TR} + (1+\phi^b)u_{NBR} - (\alpha\phi^d - \beta\phi^b)u_{FF}$$

The model described by equation (11) has seven unknown parameters (including the variances of the three structural shocks) to be estimated from six covariances; hence it is underidentified by one restriction. However, as was noted earlier, this model nests some previous attempts to measure policy innovations, each of which implies additional parameter restrictions. We consider five alternative identifications of our unrestricted model, corresponding to four indicators of policy proposed in the literature (each of which implies over-identification) and one just-identified variant. These are:

Model FFR (federal funds rate). The Bernanke-Blinder assumption that the Fed targets the federal funds rate corresponds to the parametric assumptions $\phi^d = 1$, $\phi^b = -1$, i.e., the Fed fully offsets shocks to total reserves demand and borrowing demand. From (12), we see that the monetary policy shock implied by these restrictions is $v^s = -(\alpha+\beta)u_{FF}$, i.e., the policy shock is proportional to the innovation to the federal funds rate, as expected.

Model NBR (nonborrowed reserves). Christiano and Eichenbaum's assumption is that nonborrowed reserves respond only to policy shocks. In our context this assumption implies the restrictions $\phi^d = 0$, $\phi^b = 0$ in (10). With these restrictions the policy shock becomes $v^s = u_{NBR}$.

Model NBR/TR ("orthogonalized" nonborrowed reserves). Strongin's key assumption is that shocks to total reserves are purely demand shocks, which the Fed has no choice in the short run but to accommodate (either through open-market operations or the discount window). His specification also ignores the possibility that the Fed responds to borrowing shocks. Hence the parametric restrictions imposed by Strongin's model are $\alpha = 0$, $\phi^b = 0$. Note in particular that the assumption that $\phi^b = 0$ (the Fed does not respond to borrowing shocks) is key to distinguishing Strongin's specification from the Bernanke-Blinder model, which assumes $\phi^b = -1$. For Strongin's model innovations to monetary policy are given by $v^s = -\phi^d u_{TR} + u_{NBR}$.

Model BR. Several authors have provided evidence for borrowed-reserves targeting by the Fed during certain periods (see, e.g., Cosimano and Sheehan, 1994), implying that the quantity of borrowed reserves is another potential indicator of policy. It is straightforward to see that borrowed-reserves targeting corresponds to the restrictions $\phi^d = 1$, $\phi^b = \alpha/\beta$. The implied policy shock is $v^s = -(1+\alpha/\beta)(u_{TR} - u_{NBR})$, which is proportional to the negative of the innovation to borrowed reserves.

Model JI ($\alpha = 0$, just-identification). Each of the four models above imposes two restrictions and hence is overidentified by one restriction (recall that the base model is underidentified by one restriction). Tests of these models thus take the form of a test of the overidentifying restriction. An alternative strategy is to estimate a just-identified model and check how well the parameter estimates correspond to the predictions of the alternative models. Strongin makes plausible institutional arguments for his identifying assumption that the demand for total reserves is inelastic in the short run ($\alpha = 0$); and, as we see below, this assumption is not contradicted by estimates from the (non-rejected) overidentified models. Hence we also consider as a separate case the just-identified model that imposes only that restriction.[10]

IV. Data, estimation, and results

An important practical issue in measuring policy stance is that the preferred indicator of monetary policy may change over time, as operating procedures or other factors change. A useful feature of our approach is that it can accommodate such changes, by allowing for changes in the values of parameters (e.g., those describing the Fed's behavior). In the estimates presented in this section we take two alternative approaches to dealing with possible regime shifts: First, we present estimates for both the entire 1965-1994 sample and various subsamples, using break dates chosen on a priori grounds (based, e.g., on announced changes in operating procedure). Second, we estimate a Hamilton (1989)-style regime switching model, which bases inference about the dates at which Fed behavior changed entirely on the data. As we show, the results from these two approaches are qualitatively consistent.

Because our identifying assumption is that there is no feedback from policy variables to the economy within the period[11], the length of "the period" is potentially important. For our first approach, with exogenous break dates, we report results based on monthly and biweekly data (to conserve space, results from the regime-switching model, below, are reported for monthly data only).[12] Estimates using quarterly data generated qualitatively similar conclusions, but it is more difficult to defend the identification assumption of no feedback from policy to the economy at the quarterly frequency.[13]

As we discussed in Section II, our procedure accommodates the inclusion of both policy variables and non-policy variables in the VARs. At both frequencies the policy variables we use are total bank reserves, nonborrowed reserves[14], and the federal funds rate. At the monthly frequency the non-policy variables used were real GDP, the GDP deflator, and the Dow-Jones index of spot commodity prices. Real GDP and the GDP deflator were chosen because presumably they are better indicators of broad macroeconomic conditions than are more conventional monthly indicators like industrial production and the CPI. Monthly data for real GDP and the GDP deflator were constructed by the Chow-Lin (1971) interpolation method, with a correction for first-order monthly serial correlation (see Bernanke and Mihov, 1995, for details).[15] At the biweekly frequency the non-policy variables included the *Business Week* production index and the index of spot commodity prices (broader weekly price indices are unavailable).[16]

The index of commodity prices was included as a non-policy variable in order to capture additional information available to the Fed about the future course of inflation. As is now well-known, exclusion of the commodity price index tends to lead to the "price puzzle", the finding that monetary tightening leads to a rising rather than falling price level.[17] If the commodity price index is included to capture the Fed's information about future inflation, then a parallel argument suggests putting an indicator of future output movements into the system as well. For this reason in our initial estimation we included the index of leading indicators (short horizon) in the quarterly and monthly systems. However, unlike the case of the commodity price index, inclusion of the index of leading indicators had little effect on model estimates or implied impulse response functions. For comparability with earlier results, therefore, we excluded that variable when deriving the estimates presented here.

For estimation of the model with fixed break dates we used a two-step efficient GMM procedure (maximum likelihood estimates were similar). The first step of the procedure was equation-by-equation OLS estimation of the coefficients of the VAR system.[18] The second step involved matching the second moments implied by the particular theoretical model being estimated to the covariance matrix of the "policy sector" VAR residuals. We performed two types of tests of the various models: (1) tests of overidentifying restrictions based on the minimized value of the sample criterion function (Hansen's J test); and (2) tests of hypotheses on the estimates of the structural parameters.

We turn now to the results. Estimates based on monthly data are given in Table 1, estimates from biweekly data in Table 2. Each table reports parameter estimates, with standard errors in parentheses, for the five models introduced in Section III. Parameter restrictions associated with each model are indicated in boldface. The final two columns of Tables 1 and 2 show, for each of the four over-identified models, (1) a p-value corresponding to the test of the single overidentifying restriction (OIR); and (2) a p-value for the two parameter restrictions of the model, relative to the just-identified model. P-values greater than 0.05 are shown in bold, indicating that the particular model cannot be rejected at the 5% level of significance. As indicated above, to allow for possible regime switches, we present estimates for the entire sample period, for the pre- and post-1979 samples, and for a variety of sub-periods chosen on prior grounds. Both the reduced-form VAR and the structural model parameters are re-estimated within each sub-sample. The sub-periods we have chosen correspond broadly to periods identified by Federal Reserve insiders and observers as possibly distinct operating regimes (see, e.g., Strongin, 1995). For example, 1979:10, when Chairman Volcker announced dramatic changes in the operating procedure, is a conventional break date; 1984:2 reflects both the end of the Volcker experiment and the beginning of contemporaneous reserve accounting; and 1988:11 roughly marks the beginning of the Greenspan regime, excluding the stock market crash and its aftermath.[19]

A useful starting point for reviewing Tables 1 and 2 is the estimates of ϕ^d, the coefficient that describes the Fed's propensity to accommodate reserves demand shocks (see eq. 10). In monthly data this coefficient is always estimated to lie between 0.78 and 1.05, with high statistical significance, implying something close to full accommodation of reserves demand shocks ($\phi^d = 1$). This result is inconsistent with the nonborrowed reserves (NBR) model, which assumes $\phi^d = 0$; consequently, the NBR model is strongly rejected in all sample periods. Essentially, the NBR model is rejected because it treats innovations to nonborrowed reserves as policy shocks, whereas the estimates imply that much of the variation in nonborrowed reserves reflects instead accommodation of reserves demand shocks by the Fed. The biweekly data (Table 2) generally show even a greater degree of accommodation of demand shocks, and hence rejection of the NBR model. The important exception is the subperiod 1979:10-1982:10, during which accommodation of reserves demand shocks is estimated to be quite low, around 0.1, and the restrictions of the NBR model are far from being rejected. This last result is quite interesting, since 1979-82 was the only period in which the Fed indicated publicly that it was using a nonborrowed-reserves targeting procedure. We take this correspondence of our results to the conventional wisdom as support for our approach.

A second set of interesting results from Tables 1 and 2 relates to the comparison of the funds rate (FFR) and borrowed reserves (BR) models. It is well known that operating procedures based on targeting the funds rate and on targeting borrowed reserves are quite similar in practice (see, e.g., Thornton, 1988a), and indeed, the tests of overidentifying restrictions (OIR) give essentially identical results, sub-period by sub-period, for the FFR and BR models. Formally, the models differ only in that the BR model predicts $\phi^b = -\alpha/\beta$, while the FFR model predicts $\phi^b = -1$;

and, unfortunately, ϕ^b is not always well-identified in our data. Much stronger differentiation between the two models is found under the just-identified model, however, in which we impose Strongin's assumption of inelastic reserves demand ($\alpha = 0$)[20]. Indeed, both parameter estimates and p-values under the JI model point to federal-funds-rate targeting as the overall best description of the Fed's operating procedure prior to 1979 (including both the 1965-72 and 1972-79 subperiods separately), as well as during 1984-94 (in monthly data) and 1984-88 (in biweekly data). In particular, the finding that the Fed targeted the federal funds rate prior to 1979 is consistent with both many standard accounts and the analysis of Bernanke and Blinder (1992).

What of the NBR/TR model, proposed by Strongin (1995)? This model has the "advantage" that, unlike the other identifications, it treats the degree to which the Fed accommodates reserves demand shocks (ϕ^d) as a free, rather than imposed, parameter. Also, like our JI model, the NBR/TR model assumes $\alpha = 0$. This flexibility is probably the reason that the NBR/TR model has the highest p-values under the JI model (0.021 in monthly data, 0.076 in biweekly data) for the sample as a whole. The NBR/TR model is also consistently the best single model for the post-1979 period taken as a whole. However, our results strongly suggest that using the same model of Fed operating procedures over long periods, particularly over the entire 1965-94 sample, is not warranted. In particular, as we have noted, under the JI identification the funds-rate (FFR) model appears to be the best simple indicator of policy prior to 1979, in 1984-88 (biweekly data), and in 1984-94 (monthly data), while the nonborrowed reserves model (NBR) is probably the best choice for 1979-82 (biweekly data). The NBR/TR model has the highest p-values during the 1988-94 sub-period, in both monthly and biweekly data. However, because ϕ^b is poorly identified and ϕ^d is found to be close to one during 1988-94, the estimates do not discriminate strongly among the NBR/TR, FFR, or BR models for the Greenspan era.

In the estimates reported in Tables 1 and 2, we imposed sample break points based on prior knowledge of recent monetary history. An alternative approach is to allow the sample breaks to be determined entirely by statistical procedures. To do this, we began by looking for evidence for breaks in the reduced-form parameters of the VAR system. Next, we applied Hamilton's (1989) regime-switching model to the estimation of the structural VAR model of the bank reserves market, focusing particularly on possible switches in the parameters describing Fed behavior.

Conceptually, the appropriate test for stability of the reduced-form VAR system should allow for an arbitrary break point or points, as in Andrews (1993). However, such a test applied to our six-variable, twelve-lag system as a whole would no doubt have minimal power, given the large number of estimated parameters. To increase power, and to focus attention on qualitative changes in the dynamics, we tested for breaks in the coefficients on all lags of each variable in each equation; e.g., we tested all twelve lags of GDP in the nonborrowed reserves equation simultaneously, allowing arbitrary break points.

Since there are six equations in the system, each with six sets of lagged variables, our procedure involved 36 separate tests. We used the LM variant of the test proposed by Andrews; this involves calculating the LM statistic for every possible break point and then comparing the highest value in the sequence against the tabulated critical values. Of the 36 tests conducted, only one was significant at the 5% level, and then only marginally so. We did not consider this to be very strong evidence against stability of the reduced-form VAR system, and so for the purposes of this exercise we proceeded under the assumption of no breaks in the reduced-from coefficients of the VAR over the 1965-94 sample.[21]

Given the residuals from a VAR estimated for the whole sample period, our second step was to apply Hamilton's regime-switching approach to the (just-identified) structural VAR model of the market for bank reserves. We focused on the two parameters describing the Fed's response to shocks in the reserves market, ϕ^d and ϕ^b. We allowed each of these parameters to switch between two values, for a total of four states.[22] Note that, although we key regime switches to changes in the reaction parameters, since the model is just-identified we must allow for the possibility that other structural parameters change when the state changes.[23]

Figure 1 shows the results of this exercise graphically. The two possible values for each of the parameters are estimated to be (0.884, 0.405) for ϕ^d (-0.824, -0.215) for ϕ^b, respectively. The solid line in the figure shows the probability that ϕ^d takes its first estimated value, and the dashed line shows the analogous probability for ϕ^b. The probabilities are smoothed in the sense that full-sample information is used in inferring the state probabilities at each date. The vertical lines in the figure indicate the dates of the "Volcker experiment", 1979:10 to 1982:10.

The impression conveyed by Figure 1 is quite striking. The post-1965 period appears to be sharply demarcated, between a period beginning in 1978 or 1979 and ending in 1982 on the one hand, and the rest of the sample on the other. Between 1979 and 1982, our estimates show that, with probability near one, the Fed was employing an operating procedure in which the absolute values of ϕ^d and ϕ^b were relatively low. This finding is consistent with the estimates for 1979-82 shown in Table 2, which indicate that the Fed's behavior was best described during that period as nonborrowed reserves targeting.[24] Outside the 1978-82 window, Figure 1 indicates that the Fed was primarily engaged in interest-rate smoothing: During this period we find that ϕ^d and ϕ^b are close to 1 and -1, respectively, a result also broadly consistent with those in Tables 1 and 2. The absolute values of these parameters are, however, a bit less than 1, implying that the Fed may have deviated from pure interest-rate smoothing to do some smoothing of reserves growth as well.

Summarizing the results of this section, we conclude first that the common practice of using only one policy indicator for the entire 1965-94 period in the United States is not justified. In particular, there is considerable evidence that the Fed did indeed switch to targeting nonborrowed reserves during the 1979-82 period, as many have claimed. During the rest of the sample, in contrast, the Fed was largely accommodating shocks to the demand for reserves. Treating the federal funds rate as the policy indicator for this greater portion of the sample is therefore probably a reasonable approximation, although the results of Tables 1 and 2 suggest that it is difficult to choose among the funds-rate, borrowed-reserves, and Strongin models for the most recent period (1988-94). This finding likely reflects some degree of reserves smoothing by the Fed, as well as the relatively poor identification of its response to borrowing shocks.

The comparisons we have drawn in this section thus far presuppose a choice among the existing simple indicators. However, as discussed in Section II, our analysis also suggests an alternative strategy, which is to use the just-identified model to calculate the policy shock in each period which is most consistent with our estimates of the Fed's operating procedure in that period (see eq. 12). This strategy makes it possible to analyze the effects of monetary policy, despite differences or changes in regime, in a unified framework; in addition, this approach can accommodate "hybrid" as well as "pure" operating procedures. In the next section we show that using our framework may substantially affect the inferences one draws about monetary policy.

V. Implications for estimated responses to monetary policy shocks

As the introduction discussed, our reason for estimating the parameters of the model for bank reserves is not our interest in that market, or in the Fed's operating procedures, per se. Rather, the goal is to isolate relatively "clean" measures of monetary policy shocks. Given these shocks, standard impulse response functions can be used to provide a quantitative measure of the dynamic effects of policy changes on the economy. As we have noted, many recent studies have used this approach; and methods of this type have been used as an input to monetary policy-making both in the U.S. and other countries.

Figure 2 shows estimated dynamic responses of output, the price level, and the federal funds rate to a monetary policy shock, as derived from alternative models. For comparability, in each case we consider an expansionary shock with an impact effect on the funds rate of -25 basis points. The left column shows impulse responses, with 95% confidence bands, implied by our just-identified model.[25] The right panel of Figure 2 shows impulse responses derived using alternative policy indicators (standard error bands are omitted for legibility). Following most of the recent literature, for these indicators we use identifications based on the standard Choleski decomposition; i.e., we use just-identified specifications, rather than the overidentified models discussed in the previous section. Each set of impulse responses shown in the right panel of Figure 2 was derived from a six-variable VAR with GDP, the GDP deflator, and the index of commodity prices in the non-policy block (ordered first), followed by a policy block consisting of total reserves, nonborrowed reserves (or their difference, borrowed reserves), and the federal funds rate. The specifications differ only in the ordering of the policy block and in which innovation is treated as the policy shock. In particular, to obtain the impulse responses labeled FFR (resp., NBR), we order the federal funds rate (resp., nonborrowed reserves) first in the policy block and interpret innovations to that variables as shocks to monetary policy. The impulse response labelled NBR/TR, following Strongin's (1995) original proposal, is based on an ordering with total reserves first and nonborrowed reserves second in the policy block, with the (orthogonalized) innovation to nonborrowed reserves taken to be the policy shock. Finally, the impulse response labelled BR comes from putting borrowed reserves (total reserves less nonborrowed reserves) first in the policy block, and identifying the innovation to monetary policy with the innovation to that variable.

Qualitatively, the results from all five indicators are reasonable, in the sense of conforming to the predictions of standard models.[26] In each case, an expansionary monetary policy shock increases output relatively rapidly (the peak effect is typically at about 18 months), and raises the price level more slowly but more persistently. However, quantitatively, the results differ noticeably among indicators: For example, relative to the case in which the federal funds rate is the policy indicator, using nonborrowed reserves as the indicator implies a peak response of output more than twice as large (holding fixed the initial interest-rate movement), and a long-run response of the price level about five times greater.[27] The funds-rate and nonborrowed reserves indicators also imply rather different dynamic behavior of the funds rate itself to a monetary shock.

Our analysis provides a simple framework for understanding the differences in results among indicators. Let z_{FF}, z_{NBR}, $z_{NBR/TR}$, and z_{BR} be the innovations identified with policy shocks in the four standard specifications described above, with signs chosen so that a positive innovation corresponds to an expansionary policy shock. Then, under our just-identified model (with $\alpha = 0$), and using eq. (11), we can write these innovations in terms of the "true" structural shocks, as follows:

$$(13)\quad z_{FF} = -\frac{1-\phi^d}{\beta}\nu^d + \frac{1}{\beta}\nu^s + \frac{1+\phi^b}{\beta}\nu^b$$

$$(14)\quad z_{NBR} = \phi^d\nu^d + \nu^s + \phi^b\nu^b$$

$$(15)\quad z_{NBR/TR} = \nu^s + \phi^b\nu^b$$

$$(16)\quad z_{BR} = -(1-\phi^d)\nu^d + \nu^s + \phi^b\nu^b$$

From (13)-(16) we see that--if the just-identified model of this paper is correct--each of the putative policy indicators z is contaminated, in the sense of placing some weight on reserves demand or borrowings shocks. The degree of contamination depends, of course, on the values of the parameters. For example, if ϕ^d is close to one (as we have found to be the case, except during the Volcker period), then the nonborrowed reserves indicator z_{NBR} puts considerably more weight on reserves demand shocks relative to policy shocks than does the federal funds indicator z_{FF}

(compare eqs. 13 and 14). This greater degree of contamination by reserves demand shocks may well explain the much stronger response of output and prices found when nonborrowed reserves are used as the policy indicator, compared to when the federal funds rate is the policy measure. Of course, conditional on our estimated model, the least contaminated impulse responses are those shown in the left panel of Figure 2.

As another application of our approach to the analysis of empirical impulse responses, consider the recent debate on the "vanishing liquidity effect". A number of economists, using nonborrowed reserves as an indicator of policy, have found that the liquidity effect--the impact of a given increase in nonborrowed reserves on the interest rate--has become much smaller or even disappeared since 1982 (Pagan and Robertson, 1995; Christiano, 1995). If correct, this finding has important practical implications for policy-making. However, our approach suggests that this result is largely due to the bias associated with using nonborrowed reserves as the policy indicator.

To understand this bias in more detail, note first that, according to the model developed in the present paper, the magnitude of the liquidity effect is given by the (3,2)-element in the matrix $(\mathbf{I}-\mathbf{G})^{-1}\mathbf{A}$, which is $-1/(\alpha+\beta)$ (see eq. 11). However, if nonborrowed reserves are used as the policy indicator, and the interest rate (say, the federal funds rate) follows immediately in the ordering, then the liquidity effect will be measured as the projection (regression coefficient) of the funds rate on nonborrowed reserves. Calculating this projection, we find that

(17) Estimated liquidity effect = $-\dfrac{1}{(\alpha+\beta)}[1+\dfrac{\phi^b\sigma_b^2-\phi^d\sigma_d^2}{(\phi^d)^2\sigma_d^2+\sigma_s^2+(\phi^b)^2\sigma_b^2}]$

where the σ_i^2 are the variances of the structural shocks.

If $\phi^d = \phi^b = 0$, so that the Fed is targeting nonborrowed reserves, then the bias term (the second term in the brackets in eq. 17) is zero. However, if $\phi^d > 0$ and $\phi^b < 0$, as our estimates always imply, the bias term is negative, i.e., the magnitude of the liquidity effect is understated by using the nonborrowed reserves indicator. Indeed, using the parameter estimates for the just-identified model for 1984-94 (Table 1), we evaluate the bias term to be -1.028! Using the estimate of β for the same period (and maintaining the just-identified model's assumption that $\alpha = 0$), we calculate the liquidity effect implied by our model (for a 1% change in nonborrowed reserves) to be about 32 basis points. The liquidity effect estimated from the formula in eq. (17), in contrast, is -2 basis points (i.e., the "wrong" sign). This latter finding is close to what is found by Christiano (1995). We conclude that the vanishing liquidity effect may well be the result of using a biased indicator of monetary policy, rather than of a change in the economy. Note also that the understatement of the liquidity effect in the NBR model may help to explain the large output and price responses to NBR shocks shown in Figure 2 above; because the liquidity effect is biased downward in the NBR model, a larger innovation to nonborrowed reserves is necessary to achieve the given (25-basis-point) drop in the funds rate.

The empirical analysis of this section is meant only to be illustrative. Nevertheless, it demonstrates the potential of our method to clarify important debates about the quantitative effects of changes in monetary policy.

VI. A measure of the overall stance of monetary policy

Our focus thus far has been on modelling innovations to the stance of monetary policy, as opposed to the anticipated or endogenous part of policy (the "policy rule"). As we have discussed, the advantage of studying shocks to policy is that it allows us to gauge (at least roughly) the effects of monetary policy on the economy, with minimal identifying assumptions. In contrast, empirical analysis of the effects of different monetary rules on the economy is much more difficult; such an analysis requires either observations on a large number of monetary regimes, or else a structural model identified by strong prior restrictions.

Nevertheless, it would be interesting to have an indicator of monetary policy stance that includes the endogenous as well as the exogenous component of policy. Such an indicator might be useful in characterizing the overall behavior of the Fed--e.g., the degree to which it accommodates various types of shocks--and in providing a general measure of current monetary conditions. Indeed, central banks in a number of countries currently use "monetary conditions indices", intended to provide assessments of overall tightness or ease, in their day-to-day policy-making (see, e.g., Freedman, 1994).

It is not difficult to devise a monetary conditions index, or measure of overall policy stance, consistent with the framework of this paper. For example, using the framework and notation of Section II, consider the vector of variables $\mathbf{A}^{-1}(\mathbf{I}-\mathbf{G})\mathbf{P}$. This vector, which is observable given estimates of the structural VAR system, is a full-rank linear combination of the policy indicators $\mathbf{P}$, with the property that the orthogonalized VAR innovations of its elements correspond to the structural disturbances $\mathbf{v}$. In particular, one element of this vector, call it p, has the property that its VAR innovations correspond to the monetary policy shocks derived by our approach. In analogy to the scalar case, in which there is a single observable variable (e.g., the funds rate) whose innovations correspond to the policy shock, one might consider using p as an overall measure of policy. (Indeed, under the FFR model's restrictions, $\phi^d = 1$ and $\phi^b = -1$, p equals the funds rate; similarly, under the NBR model's restrictions, p equals nonborrowed reserves, etc.) Bernanke and Mihov (1995) show that a measure constructed in this way correlates well (in the full sample and in sub-samples) with other candidate indicators of policy, such as the Boschen-Mills (1991) index discussed in the introduction.

However, as a measure of overall policy stance p has some shortcomings: First, this indicator is not even approximately continuous over changes in regime (e.g., the funds rate and nonborrowed reserves growth are not in comparable units). Second, this measure does not provide a natural metric for thinking about whether policy at a given time is "tight" or "easy". Finally, p's property that its innovations correspond to the policy shocks is not unique to it; indeed, given a VAR that includes the non-policy variables $\mathbf{Y}$ and the elements of $\mathbf{A}^{-1}(\mathbf{I}-\mathbf{G})\mathbf{P}$ other than p, any linear combination of the current and lagged variables in the VAR will have that property, if it is ordered last in a standard Choleski decomposition. This last point suggests that we look for an indicator that is a linear combination of the variables in the base VAR (so that its innovations correspond to policy shocks), but that avoids the problems of interpretation associated with p.

Our suggestion of a measure that meets these criteria is as follows: Specify the measure of overall monetary policy stance in any month t to be the forecasted growth of nominal GDP (based on the estimated VAR system's implied forecasts of GDP and the GDP deflator) from one year to three years in the future (months $t+12$ to $t+36$), *minus* the average (actual) growth rate of nominal GDP over the two previous years (months $t-24$ to t). We choose forecasted nominal GDP growth as the basis of our measure of monetary policy because if velocity were stable, forecasted nominal GDP growth would equal expected money growth; thus nominal GDP is a sort of "velocity-adjusted" measure of money. (However, as a practical matter, forecasted inflation can be substituted for forecasted nominal GDP without changing the results much). We look at the one- to three-year horizon to reduce the relative importance of non-monetary influences (such as supply shocks) on our measure, and we compare the forecast to recent nominal GDP growth because it seems reasonable to define a a "neutral" monetary policy (for which the indicator is zero) as one that keeps expected nominal GDP growth equal to recent nominal GDP growth. Defining the indicator in terms of the expected change in nominal GDP growth has the additional advantage of putting it into an easily understandable metric.

Historical values for our suggested measure are shown in Figure 3. Also shown in the figure, for comparison, are the two narrative-based measures of monetary policy, the Romer-Romer (1989) dates and the Boschen-Mills (1991) index (scaled to have the same mean and variance as our proposed measure). Examination of Figure 3 suggests that our measure conforms well with the Boschen-Mills index (the monthly correlation is 0.57), as well as with other historical accounts of U.S. monetary policy. However, contractionary turns in our indicator appear to lead rather than to coincide with the Romer dates; by our measure, Romer dates look more like points of maximum tightness in monetary policy, rather than points at which policy changed from expansionary to contractionary.

Various exercises can be conducted using this indicator, in conjunction with the basic VAR estimated in this paper. For example, it is economically interesting and computationally straightforward to break down the error in forecasting the indicator (conditional on information in $t-1$) into the component proportional to the monetary policy shock ν^s in period t (interpretable as the exogenous innovation in policy stance) and the portion which can be written as a linear combination of the other disturbances (which may be thought of as the endogenous response of policy to other developments in the economy).[28] Alternatively, this framework could be used (possibly with additional structural identification) to characterize the monetary policy rule, in terms of the dynamic response of the policy indicator to innovations to non-policy variables. We leave these and other analyses using our measure of monetary conditions to future research.

VII. Conclusion

We have used a "semi-structural VAR" approach to evaluate and develop measures of monetary policy based on reserve market indicators. A principal conclusion is that no simple measure of policy is appropriate for the entire 1965-94 period; changes in operating procedure, such as those that occurred during the 1979-82 Volcker experiment, imply changes in the preferred indicator. For practitioners looking for a simple indicator of policy stance, our results suggest that using the federal funds rate prior to 1979; nonborrowed reserves from 1979 to 1982; and either the funds rate or Strongin's measure in the most recent period, will give reasonable results. However, a more general and only slightly more complicated alternative is to base the policy measure on an estimated model of the market for bank reserves, along the lines of our just-identified model.[29] The latter approach has the advantage of being able to incorporate the effects of subtle changes in reserve-market structure and in the Fed's operating procedures. Unlike the simpler indicators, our method can also be generalized to other countries or periods. Finally, associated (and consistent) with our approach is a measure of overall monetary conditions, which we hope will prove useful.

Overall, VAR-based methods seem a most promising approach for measuring monetary policy. In future work these methods should be applied to detailed analyses of the response of the economy to policy shocks, and to the development of quantitative aids to Federal Reserve policy-making.

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[1]The fact that innovations in the money stock reflect demand as well as supply influences helps explain the "liquidity puzzle", the finding that innovations in money are not reliably followed by declines in interest rates; see Reichenstein (1987), Leeper and Gordon (1992), and Strongin (1995) for discussions.

[2]There are dozens of examples. A sampling of better known studies includes Kashyap, Stein, and Wilcox (1993); Friedman and Kuttner (1993); Ramey (1993); Gertler and Gilchrist (1994); Cochrane (1994); Eichenbaum and Evans (1995); Cecchetti (1995); and Burnside, Eichenbaum, and Rebelo (1995). Applications to countries other than the U.S., which use the VAR methodology almost exclusively, include Sims (1992), Cushman and Zha (1994), Clarida and Gertler (1995), Gerlach and Smets (1995), Kim and Roubini (1995), Tsatsaronis (1995), and Barran et al. (1996), among others.

[3]Bernanke and Mihov (1996) apply these methods to a study of recent German monetary policy.

[4]Expectations variables are not explicitly included in (1)-(2), but these can be accommodated by replacing expected future values of variables occurring in the model by the linear projections of these expectations on current and lagged variables in the system. Alternatively, think of some of the component equations of (1)-(2) as reduced-form decision rules relating choice variables to observable state variables.

[5]The idea of an exogenous policy shock has been criticized as implying that the Fed randomizes its policy decisions. Although the Fed does not explicitly randomize, it seems reasonable to assert that, for a given objective state of the economy, many random factors affect policy decisions. Such factors include the personalities and intellectual predilections of the policymakers, politics, data errors and revisions, and various technical problems.

[6]This assumption is not plausible for all macro variables, notably for various types of asset prices. See footnote 17 below.

[7]Similar models are estimated by Brunner (1994) and Leeper and Gordon (1994), among many others. A stochastic general equilibrium model of the federal funds market is provided by Coleman, Gilles, and Labadie (1993).

[8]Various sanctions and restrictions imposed by the Fed on banks' use of the discount window make the true cost of borrowing greater than the discount rate; hence, banks do not attempt to borrow infinite quantities when the

funds rate exceeds the discount rate. Peristiani (1991) gives an extensive empirical analysis of the borrowing function.

[9] We make this assumption to conform with the previous studies being examined, all of which ignore the discount rate. The discount rate, which is an infrequently-changed administered rate, may also not be well-modelled by the linear VAR framework. An alternative to assuming that the innovation to the discount rate is zero, but which has essentially the same effect, is to treat the discount rate innovation as part of the innovation to the borrowings function. For estimates of the model with a non-zero discount rate innovation, see Bernanke and Mihov (1995) or Bernanke and Mihov (1996); their results are quite consistent with those reported here.

[10] Alternatively, the model could be identified by imposing a "long-run" restriction, e.g., that monetary policy shocks have only price-level effects in the long run; see Mihov (1996).

[11] As Bernanke and Blinder (1992) discuss, there are actually two alternative timing assumptions that can be used for identifying the effects of policy, which may be appropriate under different circumstances: either that policy-makers have contemporaneous information about the non-policy variables (implying that the policy variables should be ordered last in the VAR), or that policy-makers know only lagged values of the non-policy variables (implying that the policy variables should be ordered first).

[12] Weekly data are available prior to the change in reserve accounting procedures in 1984, but subsequently only biweekly data are available. For comparability we report only biweekly results for the whole sample period.

[13] However, in related work using only reserves-market data, Geweke and Runkle (1995) find that time aggregation from biweekly to quarterly intervals is not a problem for the identification of monetary policy.

[14] We use nonborrowed reserves plus extended credit, in order to eliminate the effects of a bulge of borrowings associated with the Continental Illinois episode in 1984. To induce stationarity, we normalize total bank reserves and nonborrowed reserves by a long (36-month) moving average of total reserves; this normalization is preferable to taking logs because the model is specified in levels. Strongin (1995) normalized by a short moving average of total reserves; however, we found that this procedure creates "jerky" impulse response functions and does not cleanly separate the dynamics of total reserves and nonborrowed reserves.

[15] James Stock pointed out to us that the moving average interpolation error created by the Chow-Lin procedure could in principle invalidate our

identifying assumption, that policy shocks do not feed back to the economy within the period. As a check for robustness, we repeated our monthly estimates using industrial production in place of real GDP and the CPI (less shelter) in place of the GDP deflator. The resulting parameter estimates were virtually identical to those reported here. Because the results using real GDP and the GDP deflator yield impulse response functions that are more easily interpretable and useful for policy-making, we continue to focus on the estimates using the interpolated variables.

[16]The non-policy variables are measured Saturday to Saturday and the policy variables are measured Wednesday to Wednesday. We use values of the non-policy variables corresponding to the week lying in the middle of the two-week reserve accounting period.

[17] For further discussion, see Sims (1992) and Christiano, Eichenbaum, and Evans (1994a,b). In particular, the latter show that the price puzzle is largely an artifact of not controlling for oil supply shocks. Sims and Zha (1993) argue that treating commodity prices as predetermined for monetary policy shocks is inappropriate, since these prices may well respond within the period to monetary surprises. As a check for robustness, we re-estimated the just-identified model allowing commodity prices to be determined simultaneously with policy innovations (results available on request); this change did not significantly affect our conclusions.

[18]The lag structure used in the VARs is described in the tables. We omitted some intermediate lags in short sample periods to conserve degrees of freedom. Lag lengths are conventional; the results were not changed when we used formal statistical criteria to select the lags to be included in the first-stage VARs.

[19] Using the test for structural change with unknown break dates proposed by Andrews (1993), we did confirm that our imposed break dates are not wildly inconsistent with the data. The Andrews test found evidence for a change in the structural parameters in 1980 and in 1988 or 1989, and marginal evidence for a break in 1973. As we discuss below, there is not much evidence for a break in the reduced-form VAR parameters.

[20]Note from Tables 1 and 2 that estimates of α from non-rejected models are small and often of the wrong sign (negative), which suggests that imposing $\alpha = 0$ does little violence to the data.

[21] We also conducted approximate Andrews-type tests for each complete equation in the VAR, finding no evidence against stability.

[22] The FFR, NBR, and BR models can all be accommodated by allowing ϕ^d to switch between 0 and 1 and ϕ^b to switch between 0 and -1. The NBR/TR model imposes $\phi^b = 0$ but leaves ϕ^d unrestricted. Thus a four-state model seems sufficient to capture the key changes in Fed operating procedures.

[23] Estimates were obtained by the EM algorithm, derived for this approach by Hamilton (1990). For computational reasons, the parameters of the regime-switching model were estimated sequentially rather than simultaneously: In particular, unconditional regime switches for ϕ^d were found first. Conditional on these estimates the regime switches for ϕ^b were determined.

[24] Figure 1 is a bit ambiguous about when this regime began. We suspect that the early switch of ϕ^d, in 1978, is associated with large negative policy shocks around that period; note that Romer and Romer (1989) indicate that a period of contractionary policy began in that year. The end of this regime is, however, sharply defined.

[25] We use parameter estimates from our regime-switching model of the Fed's operating procedure, discussed above, to construct the impulse responses. Specifically, we use the parameter estimates for the pre-1978 or post-1982 subsample (the "non-Volcker experiment period"); however, since we were not able to reject stability of the reduced-form coefficients for the entire sample, and since the estimated impact on the funds rate of a monetary shock (the "liquidity effect", equal in this model to $-1/\beta$) is very similar in the Volcker and non-Volcker periods, impulse responses for the just-identified model based on 1978-82 parameter estimates are virtually identical to those shown in Figure 2.

[26] That indicators which are standard in the literature give qualitatively reasonable results should not be surprising; presumably, if they did not give plausible results they would not have become standard.

[27] Except for the first few months after the shock, the response of prices to a nonborrowed reserves shock lies entirely outside the two-standard-error band for the price response generated by the federal funds rate model, and vice versa.

[28] Performing this exercise reveals that the variance of the policy shock is quite small relative to the variance of the overall measure. However, the cumulative effect of policy shocks, including the induced dynamics, plays an important role in the overall stance of policy.

[29] A RATS procedure that estimates the model, constructs the resulting policy indicators, and calculates impulse response functions (with standard errors) for arbitrary sets of non-policy variables is obtainable from the authors.

Table 1: Parameter Estimates for All Models (Monthly)

Sample	Model	α	β	ϕ^d	ϕ^b	Tests: For OIR	Tests: Restrictions under JI model
1965:1 - 1994:12	FFR	-0.004 (0.001)	0.012 (0.001)	1	-1	**0.052**	0.002
	NBR	0.029 (0.009)	0.014 (0.001)	0	0	0.000	0.000
	NBR/TR	0	0.054 (0.015)	0.841 (0.058)	0	0.044	0.021
	BR	-0.004 (0.001)	0.041 (0.009)	1	α/β	**0.052**	0.000
	JI	0	0.021 (0.016)	0.802 (0.058)	-0.588 (0.254)	-	-
1965:1 - 1979:9	FFR	-0.005 (0.002)	0.012 (0.003)	1	-1	**0.173**	**0.080**
	NBR	0.033 (0.008)	0.067 (0.010)	0	0	0.000	0.000
	NBR/TR	0	0.102 (0.025)	0.828 (0.103)	0	0.019	0.004
	BR	-0.005 (0.002)	0.083 (0.018)	1	α/β	**0.173**	0.000
	JI	0	0.024 (0.010)	0.779 (0.101)	-0.718 (0.248)	-	-
1979:10 - 1994:12	FFR	-0.002 (0.001)	0.013 (0.001)	1	-1	**0.104**	0.003
	NBR	0.308 (0.669)	0.277 (0.590)	0	0	0.000	0.000
	NBR/TR	0	0.044 (0.011)	0.835 (0.079)	0	**0.722**	**0.749**
	BR	-0.002 (0.001)	0.034 (0.007)	1	α/β	**0.104**	**0.109**
	JI	0	0.034 (0.022)	0.826 (0.083)	-0.117 (0.365)	-	-
1966:1 - 1972:10	FFR	-0.007 (0.004)	0.007 (0.004)	1	-1	**0.650**	**0.999**
	NBR	0.014 (0.007)	0.047 (0.012)	0	0	0.001	0.000
	NBR/TR	0	0.192 (0.158)	1.025 (0.166)	0	**0.086**	0.000
	BR	-0.008 (0.004)	0.175 (0.123)	1	α/β	**0.663**	0.000
	JI	0	0.000 (0.014)	0.998 (0.118)	-1.001 (0.075)	-	-
1972:11 - 1979:9	FFR	-0.008 (0.002)	0.017 (0.005)	1	-1	**0.691**	0.024
	NBR	0.021 (0.008)	0.094 (0.018)	0	0	0.001	0.000
	NBR/TR	0	0.106 (0.024)	0.780 (0.187)	0	0.011	0.000
	BR	-0.008 (0.002)	0.101 (0.023)	1	α/β	**0.691**	0.000
	JI	0	0.013 (0.011)	0.823 (0.157)	-1.038 (0.064)	-	-
1984:2 - 1994:12	FFR	-0.007 (0.006)	0.007 (0.002)	1	-1	**0.166**	**0.207**
	NBR	0.037 (0.016)	0.132 (0.043)	0	0	0.000	0.000
	NBR/TR	0	0.454 (0.387)	0.880 (0.092)	0	**0.233**	0.034
	BR	-0.008 (0.006)	0.133 (0.106)	1	α/β	**0.167**	0.002
	JI	0	0.030 (0.024)	0.861 (0.079)	-0.690 (0.326)	-	-
1988:11 - 1994:12	FFR	0.003 (0.004)	0.002 (0.002)	1	-1	**0.359**	**0.547**
	NBR	0.271 (0.199)	0.002 (0.002)	0	0	0.009	0.000
	NBR/TR	0	0.190 (0.190)	1.049 (0.048)	0	**0.531**	**0.709**
	BR	0.003 (0.004)	0.158 (0.149)	1	α/β	**0.360**	**0.305**
	JI	0	0.026 (0.045)	1.046 (0.046)	-0.276 (0.740)	-	-

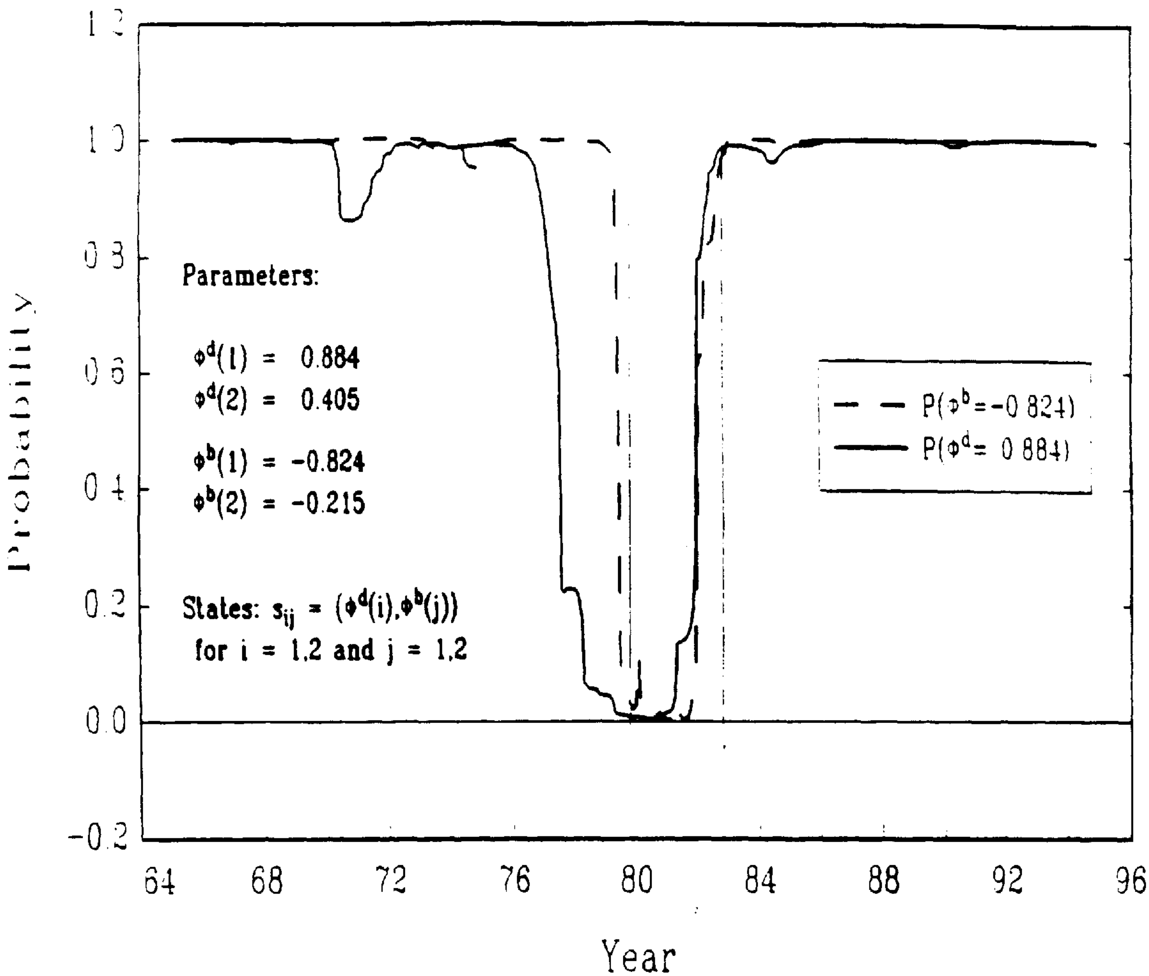
Notes: The estimates come from a 6-variable monthly VAR (see text for explanations). The last four subsamples are estimated with 6 non-consecutive lags (1,2,3,6,9,12) while the rest of the VARs are using 12 consecutive lags. The next-to-the-last column presents p-values from tests of overidentifying restrictions based on the minimized value of the criterion function. The last column gives p-values from tests of the implied restrictions under the just-identified model ($\alpha = 0$). In the last two columns the values in bold indicate that the restrictions implied by the particular model cannot be rejected at the 5% level of significance. The figures in the brackets are standard errors.

Table 2: Parameter Estimates for All Models (Biweekly)

Sample	Model	α	β	ϕ^d	ϕ^b	Tests: For OIR	Tests: Restrictions under JI model
1967:1 - 1994:12	FFR	-0.003 (0.001)	0.014 (0.002)	1	-1	**0.114**	0.048
	NBR	0.063 (0.019)	0.016 (0.002)	0	0	0.000	0.000
	NBR/TR	0	0.093 (0.021)	0.894 (0.050)	0	**0.111**	**0.076**
	BR	-0.003 (0.001)	0.083 (0.017)	1	α/β	**0.114**	0.000
	JI	0	0.031 (0.014)	0.883 (0.047)	-0.602 (0.340)	-	-
1967:1 - 1979:9	FFR	-0.005 (0.002)	0.017 (0.003)	1	-1	**0.427**	**0.213**
	NBR	0.010 (0.004)	0.077 (0.021)	0	0	0.000	0.000
	NBR/TR	0	0.137 (0.025)	0.877 (0.081)	0	**0.074**	0.036
	BR	-0.005 (0.002)	0.133 (0.020)	1	α/β	**0.427**	0.000
	JI	0	0.032 (0.020)	0.871 (0.082)	-0.781 (0.373)	-	-
1979:10 - 1994:12	FFR	-0.003 (0.002)	0.013 (0.002)	1	-1	**0.068**	0.031
	NBR	0.007 (0.005)	0.059 (0.025)	0	0	0.000	0.000
	NBR/TR	0	0.081 (0.011)	0.863 (0.063)	0	**0.428**	**0.495**
	BR	-0.003 (0.002)	0.062 (0.018)	1	α/β	**0.068**	0.036
	JI	0	0.041 (0.026)	0.858 (0.061)	-0.293 (0.429)	-	-
1967:1 - 1972:10	FFR	0.001 (0.004)	0.022 (0.003)	1	-1	**0.555**	**0.971**
	NBR	0.016 (0.007)	0.045 (0.004)	0	0	0.001	0.000
	NBR/TR	0	0.065 (0.008)	0.970 (0.122)	0	**0.469**	**0.714**
	BR	0.001 (0.003)	0.066 (0.008)	1	α/β	**0.555**	**0.163**
	JI	0	-0.024 (0.144)	0.971 (0.122)	-0.643 (1.753)	-	-
1972:11 - 1979:9	FFR	-0.002 (0.005)	0.022 (0.004)	1	-1	**0.995**	**0.894**
	NBR	0.024 (0.008)	0.099 (0.007)	0	0	0.001	0.000
	NBR/TR	0	0.182 (0.035)	0.985 (0.110)	0	**0.659**	**0.135**
	BR	-0.002 (0.004)	0.180 (0.034)	1	α/β	**0.995**	0.000
	JI	0	0.021 (0.113)	0.981 (0.110)	-1.004 (0.673)	-	-
1979:10 - 1982:10	FFR	-0.002 (0.001)	0.011 (0.003)	1	-1	**0.102**	0.013
	NBR	-0.000 (0.016)	0.061 (0.025)	0	0	**0.627**	**0.905**
	NBR/TR	0	0.058 (0.022)	0.138 (0.390)	0	**0.657**	**0.722**
	BR	-0.002 (0.001)	0.047 (0.016)	1	α/β	**0.102**	**0.075**
	JI	0	0.039 (0.031)	0.119 (0.398)	-0.144 (0.405)	-	-
1984:2 - 1988:10	FFR	-0.009 (0.006)	0.008 (0.007)	1	-1	**0.456**	**0.520**
	NBR	0.054 (0.017)	0.048 (0.014)	0	0	0.000	0.000
	NBR/TR	0	0.629 (0.885)	0.952 (0.129)	0	**0.152**	0.000
	BR	-0.009 (0.006)	0.485 (0.534)	1	α/β	**0.466**	0.000
	JI	0	0.023 (0.022)	0.887 (0.114)	-0.911 (0.187)	-	-
1988:11 - 1994:12	FFR	0.001 (0.008)	0.009 (0.003)	1	-1	**0.420**	0.041
	NBR	-0.024 (0.020)	-0.040 (0.029)	0	0	0.000	0.000
	NBR/TR	0	0.105 (0.026)	1.025 (0.030)	0	**0.963**	**0.967**
	BR	0.001 (0.008)	0.106 (0.027)	1	α/β	**0.420**	**0.697**
	JI	0	0.090 (0.278)	1.025 (0.030)	-0.017 (0.419)	-	-

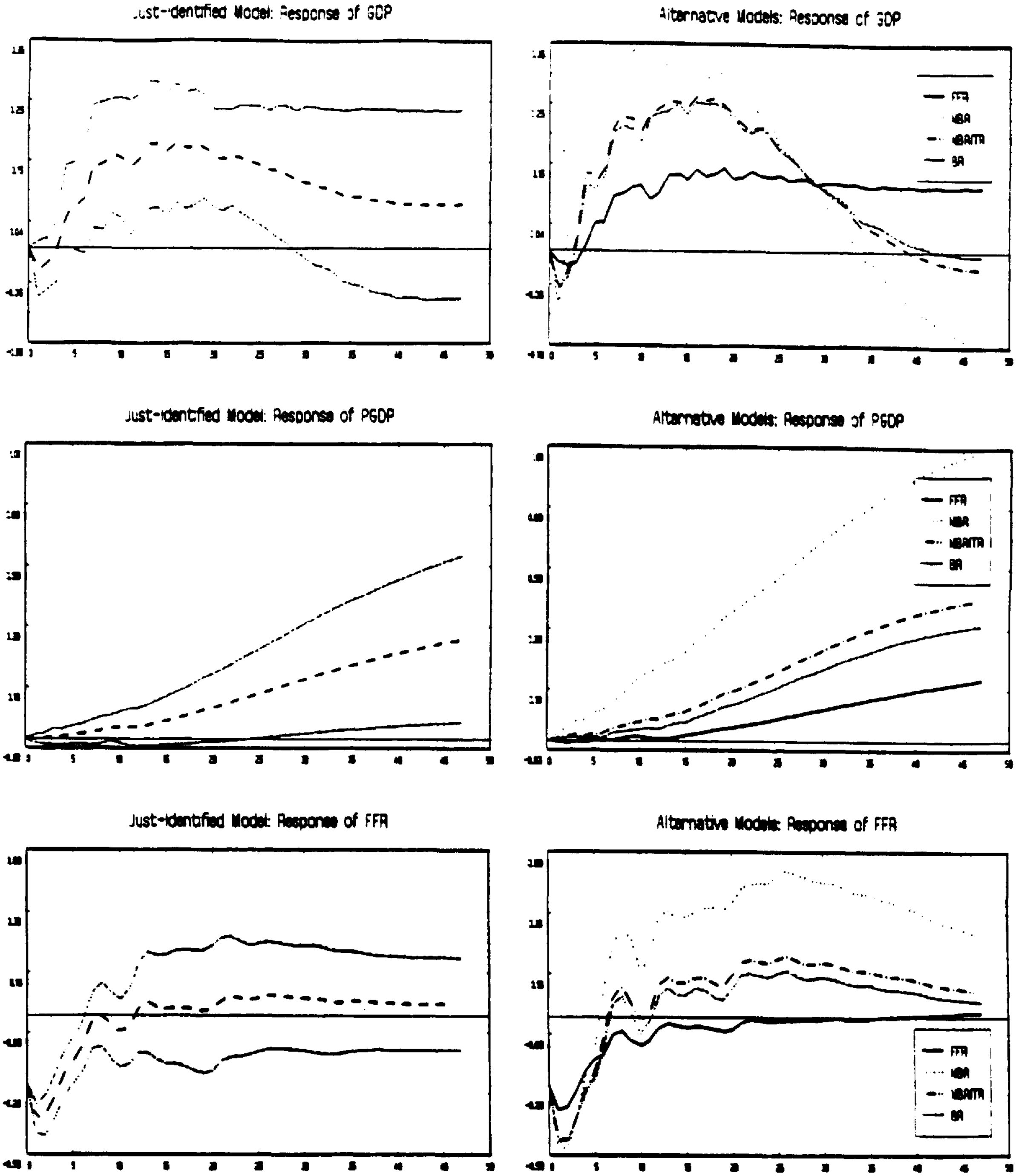
Notes: The estimates come from a 5-variable biweekly VAR. The first three samples in the table are estimated with 26 consecutive lags, the 1979 - 82 subsample is estimated with 11 non-consecutive lags (first 6 and then every fourth one); the rest of the VARs use 16 non-consecutive lags (the first 6 and then every other one). The next-to-the-last column presents p-values from a test of overidentifying restrictions based on the minimized value of the criterion function. The last column gives p-values from tests of the implied restrictions under the just-identified model ($\alpha = 0$). In the last two columns the values in bold indicate that the restrictions implied by the particular model cannot be rejected at the 5% level of significance. The figures in the brackets are standard errors.

Figure 1: Structural Parameter Estimates and Smoothed Probabilities from a Regime-Switching Model of Federal Reserve Operating Procedures



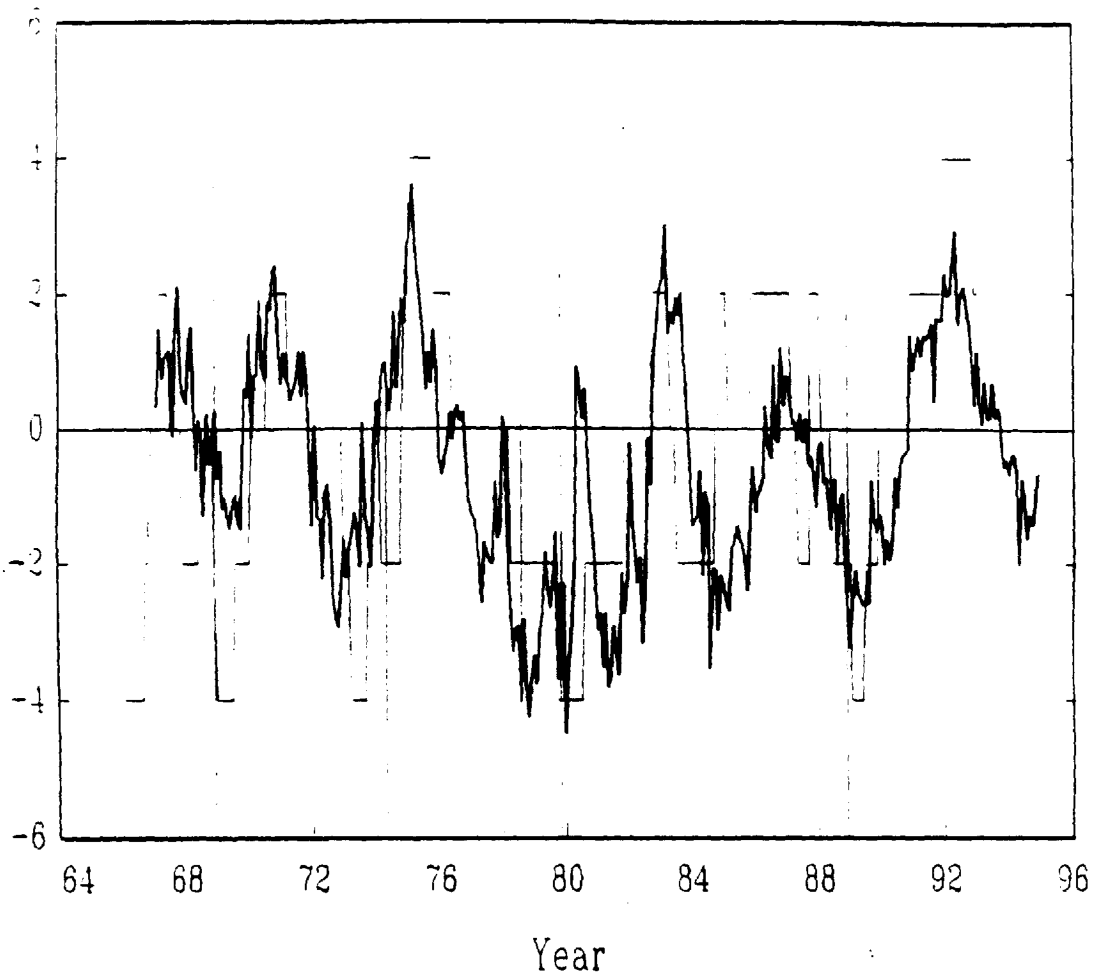
Notes: The parameters and the state probabilities are estimated using a 6-variable vector autoregression under the assumption of independent switches of the two parameters between different states. The picture shows the probabilities of being in state 1 of each of the two parameters. The two vertical lines correspond to 1979:10 and 1982:10.

Figure 2: Responses of Output, Prices, and the Federal Funds Rate to Monetary Policy Shock in Alternative Models (1965:1 - 1994:12)



Notes: The impulse responses are based on estimates from 6-variable monthly VARs with interpolated real GDP, interpolated GDP deflator, commodity price index, total reserves, nonborrowed reserves plus extended credit, and federal funds rate. Twelve lags of each variable are included in the VAR. The first column presents impulse responses with 95% confidence intervals from the just-identified model described in the text. The second column gives the responses under alternative identification schemes. In all specifications the monetary policy shock has been normalized to induce a 25 basis points drop in the federal funds rate.

Figure 3: An Indicator of Overall Monetary Policy Stance



Notes: The figure shows an indicator of monetary policy stance defined as the forecast of nominal GDP growth one to three years in the future minus the nominal GDP growth in the last two years. The solid line is the Boschen-Mills index. The vertical lines are Romer dates.