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MEASURING MONETARY POLICY

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ABSTRACT

Extending the approach of Bernanke and Blinder (1992), Strongin (1992), and Christiano, Eichenbaum, and Evans (1994a, 1994b), we develop and apply a VAR-based methodology for measuring the stance of monetary policy. More specifically, we develop a "semi-structural" VAR approach, which extracts information about monetary policy from data on bank reserves and the federal funds rate but leaves the relationships among the macroeconomic variables in the system unrestricted. The methodology nests earlier VAR-based measures and can be used to compare and evaluate these indicators. It can also be used to construct measures of the stance of policy that optimally incorporate estimates of the Fed's operating procedure for any given period. Among existing approaches, we find that innovations to the federal funds rate (Bernanke-Blinder) are a good measure of policy innovations during the periods 1965-79 and 1988-94; for the period 1979-94 as a whole, innovations to the component of nonborrowed reserves that is orthogonal to total reserves (Strongin) seems to be the best choice. We develop a new measure of policy stance that conforms well to qualitative indicators of policy such as the Boschen-Mills (1991) index. Innovations to our measure lead to reasonable and precisely estimated dynamic responses by variables such as real GDP and the GDP deflator.

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I. Introduction

Is the stance of monetary policy currently more expansionary, or more contractionary, than in the recent past? How does the stance of policy today compare to what it was in earlier periods with similar macroeconomic conditions? These questions are of great practical importance for good policy-making, yet answering them can be surprisingly difficult and contentious: During the 1990-91 recession, for example, Federal Reserve officials claimed that policy was becoming more expansionary, pointing to repeated cuts in the federal funds rate. At the same time, some critics cited the unusually slow growth of M2 as evidence that monetary policy was a drag on the recovery. It would be useful to have methods for addressing such issues quantitatively.

Measuring monetary policy accurately is also important for scientific reasons: There is little hope that economists can evaluate alternative theories of monetary policy transmission, or obtain quantitative estimates of the impact of monetary policy changes on various sectors of the economy, if there exists no reasonably objective means of determining the direction and size of changes in policy stance.

Traditionally, the stance of monetary policy was measured by the rate of growth of one or more monetary aggregates (e.g, M2, M1, or the monetary base). However, it is now well recognized that the growth rates of monetary aggregates typically depend on a variety of non-policy influences. For example, if the Fed's operating procedure involves some smoothing of short-term interest rates, as has been the case for most of the past thirty years, then shocks to money demand will be partially accommodated by the monetary authorities. As a result, money growth rates will reflect changes in money demand as well as changes in policy.[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.

In this paper we discuss a general econometric methodology for measuring the stance of monetary policy, and use this methodology both to formally evaluate existing measures and to develop new ones. Our method builds on the "structural VAR" approach, utilized in this context by Bernanke and Blinder (1992), Strongin (1992), and Christiano, Eichenbaum, and Evans (1994a, 1994b), among others. More specifically, we develop 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. By estimating the model over different sample periods we are able to allow for changes in the structure of the economy and in Fed operating procedures, while at the same time imposing a minimal set of identifying assumptions. Although we spend most of the paper using our methods to analyze policy measures derived from the structural VAR tradition, we also consider briefly the relationship of our derived policy measures to more qualitative indicators suggested by Romer and Romer (1989) and Boschen and Mills (1991).

Much of the recent work using the structural VAR approach has focused on measuring exogenous *innovations* to policy, rather than developing an overall measure of policy stance. Measuring policy innovations is useful,

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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 (1992) for discussions of the liquidity puzzle.

since (with some additional assumptions) the dynamic responses of other variables in the VAR to policy innovations can be identified with the structural responses of those variables to monetary policy shifts. We elaborate this point below. However, it is also desirable to have indicators of the overall thrust of policy, including the endogenous or anticipated portion of policy. We consider the measurement of both policy innovations and overall policy stance in this paper.

The rest of the paper proceeds as follows. Section II briefly reviews the recent literature on indicators of monetary policy stance. Section III presents our VAR-based methodology for measuring the direction of policy. Section IV develops a simple model of the market for bank reserves that nests some common alternative descriptions of Fed operating procedures. Estimation of this model by GMM (Section V) allows us both to evaluate some existing proposed measures of policy and to develop an alternative "optimal" (given our framework) measure. Section VI presents estimates from an extended version of the model that uses discount-rate information, and Section VII concludes.

II. Some existing indicators of monetary policy

Recent attempts to provide measures of the stance of monetary policy (or the change in stance) have 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. The Romers read the minutes of the Federal Open Market Committee and 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 attempts to use additional information--in this case, statements of policy-makers' intentions--to try to disentangle money supply from money demand shocks. The Romers' strategy also has the advantage of being "nonparametric", in that its implementation does not require any modelling of the details of the Fed's operating procedure or of the financial system and is potentially robust to changes in those structures. A disadvantage of this approach, besides its inherent subjectivity, is its difficulty in distinguishing between 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. Another weakness is that the Romers' approach provides a rather limited amount of information: They provide only dates on which policy shifted in a contractionary direction, not dates on which policy became more expansionary. Further, no distinction is made between mildly contractionary and severely contractionary episodes.

An interesting attempt to increase the amount of information extracted from the narrative approach has been made by Boschen and Mills (1991). Based on their reading of FOMC documents, they provide a monthly index that rates Fed policy in each month as "strongly expansionary", "mildly expansionary", "neutral", "mildly contractionary", or "strongly contractionary", depending on the relative weights the policy-makers assign to reducing unemployment and reducing inflation. Although Boschen and Mills provide a more continuous measure of policy than do Romer and Romer, the tradeoff is that the problems of subjectivity and policy endogeneity are probably relatively more severe for the Boschen and Mills approach than for that of the Romers.

A second general strategy for measuring monetary policy stance is to use information about Federal Reserve operating procedures to develop data-based indexes of policy. For example, Bernanke and Blinder (1992) argue that over much of the past thirty years (particularly prior to 1979) the Fed has implemented policy changes primarily through changes in the federal funds rate (the overnight rate in the market for commercial bank reserves). They conclude that the funds rate (or, alternatively, the spread between the funds rate and a long-term bond rate) may therefore be used as an indicator of policy stance (see also Laurent, 1988; Bernanke, 1990).

Another measure of policy was suggested by Christiano and Eichenbaum (1992) (see also Eichenbaum, 1992). Noting the endogeneity problems with broad-money measures of policy, Christiano and Eichenbaum proposed using the quantity of nonborrowed reserves--the instrument which is perhaps the most directly controlled by the Fed--as an indicator of policy. They found that the responses of interest rates and other macro variables to innovations in nonborrowed reserves matched prior notions of how monetary shocks are supposed to affect the economy.

A potential problem with the Bernanke-Blinder and Christiano-Eichenbaum measures of policy is that each presumes a more-or-less constant set of operating procedures by the Fed--in the former case, a funds-rate-based procedure, and in the latter case a procedure that targets nonborrowed reserves. In an innovative paper, Strongin (1992) proposed a measure that could accommodate some changes in operating procedure. Strongin's measure of policy is the portion of nonborrowed reserve growth that is orthogonal to total reserve growth: He motivates this measure by arguing that the Fed is constrained to meet total reserve demand in the short run (failure to do so would lead to wild swings in the funds rate), but it can effectively tighten policy by reducing nonborrowed reserves and forcing banks to borrow more from the discount window. Econometrically, an important advantage of Strongin's approach is that--because it allows the projection coefficient of nonborrowed reserves on total reserves to vary over subperiods--it is able to nest alternative operating procedures: For example, a policy in which the Fed fully accommodates shocks to the demand for reserves (i.e., the projection coefficient of nonborrowed on total reserves is unity) approximates a funds-rate-targeting strategy, as in Bernanke-Blinder. Alternatively, a strategy of targeting nonborrowed reserves, as suggested by Christiano and Eichenbaum, can be represented in the Strongin framework by a zero response of nonborrowed reserves to total reserves innovations. In this paper we adopt an approach similar to Strongin's in that we employ a specification of the bank reserves market that can accommodate a variety of alternative operating procedures. However, our approach is broader than Strongin's in that it can accommodate a larger set of operating procedures, institutional features, and identifying restrictions.

The articles by Bernanke-Blinder, Christiano-Eichenbaum, and Strongin all focused primarily on the measurement of monetary policy *innovations*. Nevertheless, their work suggests some potentially interesting indicators of the *overall* stance of monetary policy as well: Figures 1-3 graph historical monthly data for the federal funds rate--ten-year bond rate spread (Bernanke-Blinder), the growth rate of nonborrowed reserves (Christiano-Eichenbaum), and the portion of nonborrowed reserves that is orthogonal to total reserves (Strongin).[2] The latter two measures are

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[2]The funds rate--bond rate spread is inverted, so that higher values correspond to easier policy as in Figures 2 and 3. Nonborrowed and total reserves are both normalized by a 36-month moving average of total reserves, consistent with the approach used by Strongin (although he employed a shorter moving average) and the practice we adopt below. Figure 3 assumes a single projection coefficient for the entire sample, which doesn't do justice to Strongin's point that the responsiveness of

smoothed to reduce high-frequency noise. Also shown on each graph are the Boschen-Mills index and the Romer dates. The Boschen-Mills index has been rescaled in all figures to have a variance comparable to the other policy indicator being displayed.

The visual impression is that the three data-based indexes correspond fairly well with each other, with the qualitative indicators, and with conventional wisdoms about the historical evolution of U.S. monetary policy. Table 1 shows the correlations of the various measures. The spread measure (Bernanke-Blinder) and the orthogonal component of nonborrowed reserves (Strongin) seem particularly closely related; the correlation of these two measures over the whole sample is 0.70. The three measures have similar correlations with the Boschen-Mills index--0.60 for the funds-rate spread, 0.56 for nonborrowed-reserve growth, and 0.55 for the Strongin measure.

Although it is reassuring to know that alternative measures of policy give qualitatively similar answers, it would be useful to have a methodology for comparing and evaluating these measures. Our suggested approach is outlined in the next section.

III. Methodology

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

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nonborrowed to total reserves may vary over time. Figures 1-3 are intended to be illustrative only; we provide more formally motivated measures later in the paper.

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

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

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. Boldface letters are used to indicate vectors or matrices of variables or coefficients. In particular, $\mathbf{Y}$ is a vector of (nonpolicy) macroeconomic variables, and p is a variable indicating the stance of policy. (Note that for the moment we take p to be a scalar measure; in Bernanke and Blinder's application, p was the federal funds rate or funds-rate spread.) Equation (2) may be interpreted as the policy-maker's reaction function, 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), we allow the structural error terms in equation (1) to be 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.[3]

[3]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, given the state of the economy, there are many random factors affecting policy decisions. Such factors include the personalities and intellectual predilections of the policymakers, political factors, data revisions, and various technical problems. In any case, if there are no exogenous policy shocks, then as

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., $\mathbf{C}_0 = 0$.[4] Under this assumption the system (1)-(2) can be written in standard VAR format as

$$(1') \quad \mathbf{Y}_t = (\mathbf{I}-\mathbf{B}_0)^{-1}\sum_{i=1}^{k}\mathbf{B}_i\mathbf{Y}_{t-i} + (\mathbf{I}-\mathbf{B}_0)^{-1}\sum_{i=1}^{k}\mathbf{C}_i p_{t-i} + (\mathbf{I}-\mathbf{B}_0)^{-1}\mathbf{A}^y\mathbf{v}_t^y$$

$$(2') \quad p_t = \sum_{i=1}^{k}[\mathbf{D}_i + \mathbf{D}_0(\mathbf{I}-\mathbf{B}_0)^{-1}\mathbf{B}_i]\mathbf{Y}_{t-i} + \sum_{i=1}^{k}[g_i + \mathbf{D}_0(\mathbf{I}-\mathbf{B}_0)^{-1}\mathbf{C}_i]p_{t-i} + [\mathbf{D}_0(\mathbf{I}-\mathbf{B}_0)^{-1}\mathbf{A}^y\mathbf{v}_t^y + v_t^p]$$

Equations (1') and (2') show explicitly how reduced-form VAR coefficients are related to the underlying structural coefficients (which are not separately identifiable). Estimation of the above 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. Impulse response functions for all variables in

far as we can see the effects of policy on the economy are econometrically unidentifiable by any method.

[4]This assumption is obviously more plausible if the time period is short (our results below are based on monthly and biweekly data), and if the list of macro variables excludes variables such as interest rates that are likely to respond quickly to policy changes. An alternative identifying assumption discussed by Bernanke and Blinder is that the policy-maker does not respond to contemporaneous information, i.e., $D_0 = 0$. Christiano, Eichenbaum, and Evans (1994b) compare the results obtained by these two alternative assumptions; they prefer the assumption that policy does not feed back to the economy within the period, although most of their results do not depend strongly on which assumption is used. As we report below, we also find our estimates to be relatively invariant to the identifying assumption, suggesting that the policy shock is not highly contemporaneously correlated with other disturbances in the system.

the system with respect to the policy shock can then be calculated, and can be interpreted as the true structural responses to policy changes (assuming that the linear structure is invariant). Further, the policy variable p_t itself, which is the sum of the "forecasted" portion of policy[5] and the policy shock, provides an overall indicator of the stance of policy.

The Bernanke-Blinder method assumes that a good scalar measure of policy (e.g., the federal funds rate) 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, additionally, 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 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 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

[5]That is, the portion predicted by both lagged and contemporaneous macroeconomic variables and by lagged policy variables.

or whatever disturbances affect the policy indicators. Equation (3) allows the nonpolicy variables $\mathbf{Y}$ to depend on current and lagged values of $\mathbf{Y}$ and on lagged values (only) of $\mathbf{P}$; allowing the nonpolicy 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.[6] There is no need to impose any other restrictions on the coefficients $\mathbf{C}_i$, $i > 0$, as the nonpolicy 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 ν^s. As before, write the system in VAR form (with only lagged variables on the right-hand side):

$$\text{(3')} \quad \mathbf{Y}_t = (\mathbf{I}-\mathbf{B}_0)^{-1}\sum_{i=1}^{k}\mathbf{B}_i\mathbf{Y}_{t-i} + (\mathbf{I}-\mathbf{B}_0)^{-1}\sum_{i=1}^{k}\mathbf{C}_i\mathbf{P}_{t-i} + (\mathbf{I}-\mathbf{B}_0)^{-1}\mathbf{A}^y\mathbf{v}_t^y$$

$$\text{(4')} \quad \begin{aligned}\mathbf{P}_t = {} & (\mathbf{I}-\mathbf{G}_0)^{-1}\sum_{i=1}^{k}[\mathbf{D}_i + \mathbf{D}_0(\mathbf{I}-\mathbf{B}_0)^{-1}\mathbf{B}_i]\mathbf{Y}_{t-i} + (\mathbf{I}-\mathbf{G}_0)^{-1}\sum_{i=1}^{k}[\mathbf{G}_i + \mathbf{D}_0(\mathbf{I}-\mathbf{B}_0)^{-1}\mathbf{C}_i]\mathbf{P}_{t-i} \\ & + [(\mathbf{I}-\mathbf{G}_0)^{-1}\mathbf{D}_0(\mathbf{I}-\mathbf{B}_0)^{-1}\mathbf{A}^y\mathbf{v}_t^y + (\mathbf{I}-\mathbf{G}_0)^{-1}\mathbf{A}^p\mathbf{v}_t^p]\end{aligned}$$

More compactly, write the VAR system made up of equations (3') and (4') as

$$\text{(5)} \quad \mathbf{Y}_t = \sum_{i=1}^{k}\mathbf{H}_i^y\mathbf{Y}_{t-i} + \sum_{i=1}^{k}\mathbf{H}_i^p\mathbf{P}_{t-i} + \mathbf{u}_t^y$$

$$\text{(6)} \quad \mathbf{P}_t = \sum_{i=1}^{k}\mathbf{J}_i^y\mathbf{Y}_{t-i} + \sum_{i=1}^{k}\mathbf{J}_i^p\mathbf{P}_{t-i} + [(\mathbf{I}-\mathbf{G}_0)^{-1}\mathbf{D}_0\mathbf{u}_t^y + \mathbf{u}_t^p]$$

[6]Actually, the assumption is a bit stronger than in the scalar case, as we are also assuming that non-policy innovations to the policy indicators do not feed back into the rest of the economy during the current period.

where the coefficients and error terms are defined by the obvious correspondence with equations (3') and (4'). Imagine now that we estimate (5) and (6) by standard VAR methods, then extract the component of the residual of (6) that is orthogonal to the residual of (5). Comparing (6) with (4'), we see that this orthogonal component, $\mathbf{u}^p$, is given by

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

Alternatively, dropping subscripts and superscripts, we can rewrite (7) as:

(8) $$\mathbf{u} = \mathbf{G}\mathbf{u} + \mathbf{A}\mathbf{v}$$

Equation (8) is a standard "structural VAR" 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 standard methods (Blanchard and Watson, 1986; Bernanke, 1986; Sims, 1986), 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, given the estimated coefficients of the structural VAR, the following vector of variables is observable:

(9) $$(\mathbf{I} - \mathbf{G})\mathbf{A}^{-1}\mathbf{P}$$

The variables described by (9), which are linear combinations of the policy indicators $\mathbf{P}$, have the property that their orthogonalized VAR innovations correspond to the structural disturbances $\mathbf{V}$. In particular, one of these variables, call it p, has the property that its VAR innovations correspond to innovations in the monetary policy shock. In analogy to the scalar case, in which there is a single observable variable (e.g., the fed funds rate) whose innovations correspond to policy shock, we propose using the estimated linear combination of policy indicators p as a measure of overall monetary policy stance.[7] Although not in itself a normative measure, a total measure of policy stance is potentially useful for evaluating the overall direction of policy, and for making comparisons of current policy stance with policies chosen under similar circumstances in the past. Thus the total policy measure may be useful input to the policy-making process.

IV. Monetary policy and the market for bank reserves

To implement the general strategy described in Section III, we use as policy indicators several variables bearing on the market for bank reserves; and we employ a simple and conventional model of the bank reserve market to extract the policy innovation.[8] An advantage of our model, despite its simplicity, is that it nests the Bernanke-Blinder, Christiano-Eichenbaum, and Strongin measures of policy described earlier, as well as

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[7]The requirement that the total policy indicator have innovations corresponding to the policy shock determines the total indicator only up to a constant term. An interesting issue is whether the constant can be set so that a zero value for the indicator implies that policy is "neutral" in some sense. Note also that our definition of the total indicator relies heavily on the linearity of the system.

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

measures based on yet other possible operating procedures such as borrowed-reserves targeting.

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 (also see Figure 4):

$$(10) \qquad u_{TR} = -\alpha u_{FF} + v^{d}$$

$$(11) \qquad u_{BR} = \beta(u_{FF} - u_{DISC}) + v^{b}$$

$$(12) \qquad u_{NBR} = \phi^{d} v^{d} + \phi^{b} v^{b} + v^{s}$$

Equation (10) 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 (11) 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.[9] The innovation

[9]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.

in the demand for nonborrowed reserves, the difference between total and borrowed reserves, is $u_{TR} - u_{BR}$.

Equation (12) 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 v^s is the shock to policy that we are interested in identifying. Note that the system (10)-(12) is in the form of equation (8).

It will also be useful to write the reduced-form relationship between the VAR residuals $\mathbf{u}$ and the structural disturbances $\mathbf{V}$, as in equation (7). To do so, we first make the simplifying assumption that the innovation to the discount rate u_{DISC} is zero; we relax this assumption in estimates presented in Section VI.[10] 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

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

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[10]The principal reason for this assumption is to conform with the three previous studies being examined, all of which also 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.

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

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

One can also invert the relationship (13) to determine how the monetary policy shock v^s depends on the VAR residuals:

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

The model described by equation (13) 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:

Case I (Bernanke-Blinder). 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. In this case the matrix relating the VAR residuals to the structural disturbances is

(15) $$(\mathbf{I}-\mathbf{G})^{-1}\mathbf{A} = \begin{vmatrix} 1 & \frac{\alpha}{\alpha+\beta} & 0 \\ 1 & 1 & -1 \\ 0 & -\frac{1}{\alpha+\beta} & 0 \end{vmatrix}$$

and the policy shock (from equation (14)) is

(16) $$v^s = -(\alpha+\beta)u_{FF}$$

i.e., the policy shock is proportional to the innovation to the federal funds rate, as expected.

Case II (Christiano-Eichenbaum). Christiano and Eichenbaum's assumption is that nonborrowed reserves respond only to policy shocks. From equation (12), this assumption implies the restrictions $\phi^d = 0$, $\phi^b = 0$. The matrix relating the VAR residuals to the structural disturbances is

(17) $$(\mathbf{I}-\mathbf{G})^{-1}\mathbf{A} = \begin{vmatrix} \frac{\beta}{\alpha+\beta} & \frac{\alpha}{\alpha+\beta} & \frac{\alpha}{\alpha+\beta} \\ 0 & 1 & 0 \\ \frac{1}{\alpha+\beta} & -\frac{1}{\alpha+\beta} & -(\frac{1}{\alpha+\beta}) \end{vmatrix}$$

and the policy shock, obviously enough, is

(18) $$v^s = u_{NBR}$$

Case III (Strongin). 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 on our model that yield 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 we have:

$$(19)\quad (\mathbf{I}-\mathbf{G})^{-1}\mathbf{A} = \begin{vmatrix} 1 & 0 & 0 \\ \phi^d & 1 & 0 \\ (\frac{1}{\beta})(1-\phi^d) & -\frac{1}{\beta} & -\frac{1}{\beta} \end{vmatrix}$$

$$(20)\quad v^s = -\phi^d u_{TR} + u_{NBR}$$

Equation (20) is the same as used by Strongin; it states that the policy shock is the innovation to nonborrowed reserves less its projection on the innovation to total reserves.

Case IV (Borrowed-reserves targeting). It has been suggested that at certain times the Fed has targeted borrowed reserves (see, e.g., Cosimano and Sheehan, 1994, for some evidence), and so for completeness we consider this case as well. It is straightforward to see that borrowed-reserves targeting corresponds to the restrictions $\phi^d = 1$, $\phi^b = \alpha/\beta$. Here the matrix relating the VAR residuals to the structural disturbances is

$$(21)\quad (\mathbf{I}-\mathbf{G})^{-1}\mathbf{A} = \begin{vmatrix} 1 & \frac{\alpha}{\alpha+\beta} & \frac{\alpha}{\beta} \\ 1 & 1 & \frac{\alpha}{\beta} \\ 0 & -\frac{1}{\alpha+\beta} & -\frac{1}{\beta} \end{vmatrix}$$

and the policy shock is

$$(22)\quad v^s = -(u_{TR} - u_{NBR})$$

which equals the negative of the innovation to borrowed reserves. Borrowed-reserves targeting is well-known to be close to funds-rate targeting when shocks to the borrowing function are small: Comparing (21) to (15) we note that the matrices defining the two cases differ only in the last column, which gives the responses of the innovations in the observable variables to borrowing shocks; hence, if borrowing shocks are small, the two operating procedures are not distinguishable.

Case V ($\alpha=0$, just-identification). Cases I-IV each impose two restrictions, leading to models that are 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. As we will see below, Strongin's identifying assumption that shocks to total reserves may be identified with reserve demand shocks ($\alpha=0$) seems to perform well in the overidentified models; and so we also consider as a separate case the just-identified model that imposes only that restriction.[11]

V. Data, estimation, and results

Because our identifying assumption is that there is no feedback from policy variables to the economy within the period, the length of "the period" is potentially important. In this paper we report results based on

[11]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. We expect to pursue this approach in future research.

monthly and biweekly data.[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 (or the alternative assumption that policy-makers do not respond to contemporaneous information) at the quarterly frequency.[13]

As we discussed in Section III, our procedure requires the inclusion of both policy variables and nonpolicy 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 an 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 Appendix 1 for more details).[15] At the biweekly frequency the non-policy variables

[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. Weekly results for the pre-1984 period did not differ substantially from the biweekly results.

[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 econometric 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 (1992) normalized by a short moving average of total reserves; however, 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

included were the Business Week production index and the index of commodity prices (a broader weekly price index does not seem to be available).[16]

The index of commodity prices was included in the various VAR systems 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. (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.) If the commodity price index is included to capture the Fed's information about the future of 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.

As was also noted in Section III, for every system there are two possible identifying assumptions: That policymakers have contemporaneous

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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 production index is measured Saturday to Saturday, while other variables are measured Wednesday to Wednesday (the commodity price index is available daily). We used averaged values of the production index for the two-week period ending the Saturday prior to settlement Wednesday; other variables are averaged over the two-week settlement period.

information about the nonpolicy variables (implying that the policy variables should be ordered last in the VAR) and that policymakers know only lagged values of the nonpolicy variables (implying that the policy variables should be ordered first). Our results turned out to be relatively insensitive to the ordering chosen. Thus, rather than give all permutations, in all results reported below we assume that policymakers have contemporaneous information, i.e., the policy variables are ordered last.[17] We also estimated three-variable systems including policy variables only (not reported), again obtaining relatively similar results.[18]

For estimation 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.[19] 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 (suitably orthogonalized to purge correlation with nonpolicy variables). Appendix 2 provides more information about our procedure. 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 of the alternative overidentified models based on monthly and biweekly data are given in

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[17]Interestingly, the assumption of contemporaneous information may be more nearly literally true at the biweekly frequency, as the production index and commodity price index are available with very little lag.

[18]Strongin (1992) similarly considers systems with and without the nonpolicy variables, finding qualitatively similar results in the two cases.

[19]The lag structure used in the VARs is described in Tables 2-6. 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.

Tables 2 and 3 respectively. Each table reports parameter estimates from the four models described in the previous section: Bernanke-Blinder (Model B), Christiano-Eichenbaum (Model C), Strongin (Model S), and borrowed-reserves targeting (Model BR). Also shown for each model and sample period is a p-value corresponding to the test of the overidentifying restriction (OIR). Estimates are presented for the entire sample period and a variety of sub-periods. The sub-periods chosen correspond broadly to periods identified by Federal Reserve insiders and observers as possibly distinct operating regimes (see, e.g., Strongin (1992)). For example, 1979 is a conventional break date, corresponding to the change in operating procedures announced by Chairman Volcker; 1984 reflects both the end of the Volcker experiment and the beginning of contemporaneous reserve accounting; and 1988 marks the beginning of the Greenspan regime and the period following the stock market crash. We conducted more formal tests for structural breaks in the model parameters, which largely supported our a priori choices of break dates.[20] Below we also report some results using a rolling sample period, an approach that requires no prior assumptions about break dates.

A convenient way to organize the discussion is by considering each model of the Fed's operating procedure in turn. We begin with the Bernanke-Blinder model (Model B), which assumes that the federal funds rate is the best indicator of policy. The conventional wisdom is that funds-

[20]We ran tests for structural breaks with unknown change dates as proposed by Andrews (1993). Essentially, the test consists of calculating a Wald-type statistic for each possible break point, finding the dates for which the statistic takes high values, and comparing these values to the critical values tabulated by Andrews. In monthly data, the Andrews test applied to the whole sample suggested a break in May 1980 or shortly thereafter. We cannot reject the hypothesis of no break for the 1965-1979 period, although there is marginal evidence of a break in late 1973. For the latter part of the sample, there is strong evidence of a break in 1988 or 1989.

rate targeting was a good description of Federal Reserve policy prior to 1979, and possibly also during the more recent period (under Chairman Greenspan). Our findings in monthly data (Table 2) are quite consistent with this impression: Although Model B is not rejected at the .05 level for the sample as a whole or for either half of the sample, its performance is clearly best in the 1966-72, 1972-79, and 1988-94 subsamples, and worst in the 1984-88 subsample.[21] Similarly, in the biweekly data (Table 3), Model B is not rejected for the sample as a whole or in any subsample, but performs particularly well in the pre-1979 period and in the more recent portion of the sample (including 1984-88, in this case).

The Christiano-Eichenbaum model (Model C) takes nonborrowed reserves to be the policy indicator. Model C does generally poorly in monthly data, being rejected at $p = 0.000$ for the sample as a whole and at $p < .05$ for all but one subsample (1984-94).[22] In biweekly data, Model C is rejected in every subperiod except 1979-82, during which Models C and S are the preferred models. This last result is quite interesting, since 1979-82 was the only period in which the Fed officially adhered to a nonborrowed-reserves targeting procedure.

The Strongin model (Model S) is relatively more flexible than the first two models in that it treats the response of nonborrowed reserves to shocks to the demand for total reserves as a free parameter, rather than assuming that that response necessarily equals one (Model B) or zero (Model C). However, Model S imposes the restriction that the demand curve for total reserves is vertical ($\alpha=0$) and it does not allow nonborrowed

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[21]The pattern is similar in quarterly data (results not shown), for which Model B does particularly well on the p-value criterion for 1972-79 and 1988-94.

[22]Model C performs somewhat better in quarterly data, particularly in the second half of the sample period.

reserves to respond to shocks to borrowings. In the monthly data, Model S is rejected for the sample as a whole. However, although Model S does not perform quite as well as Model B in the post-1988 period, for the post-1979 period taken as a whole it appears to be the preferred model. In biweekly data, the Strongin model does relatively well, its overidentifying restrictions not being rejected for any subsample. The general pattern of the monthly results again holds in the biweekly data, as Model S appears to be the best for the post-1979 period as a whole on the p-value criterion but is dominated by Model B for the pre-1979 and more recent periods.

Based on these results, the borrowed-reserves model (BR) turns out to be impossible to distinguish from the Bernanke-Blinder funds-rate-based model; the p-values of the two models in Tables 2 and 3 are virtually identical. Evidently, in this framework the data cannot pin down the value of ϕ^b, the response of the Fed to borrowing shocks, well enough to distinguish the two models. However, we see below that in a nested testing framework we are able to reject the borrowed-reserves model in favor of the funds-rate-based model.

We have not yet discussed the parameter estimates themselves, as reported in Tables 2 and 3. The estimates of most parameters seem reasonable, with the exception that the parameter α (the negative of the slope of the reserves demand equation) although small is generally incorrectly signed (negative), notably for Model B. In terms of the variance-covariance matrix of the VAR residuals, the "problem" for Model B is that the contemporaneous covariance of innovations to the funds rate and to total reserves is typically positive, rather than negative as implied by a model in which funds rate shocks reflect policy changes only. An advantage of Model C is that it can accommodate this positive covariance, and negative estimates of α occur less frequently for Model C.

Of the various models, only Model S (Strongin's model) allows for a free estimate of the responsiveness of the Fed to shocks to the demand for total reserves, the parameter ϕ^d. These results give an interesting insight on the performance of the Bernanke-Blinder model, Model B. Recall that Model B assumes that $\phi^d = 1$, i.e., that (in order to stabilize the funds rate) the Fed fully accommodates shocks to the demand for total reserves. Table 2 shows that, for most subperiods (especially 1966-72 and 1988-94) the estimate of ϕ^d in monthly data is close to 1, implying substantial accommodation of demand shocks. The exception in the monthly data is the period 1984-88, in which the Fed appears to have accommodated demand shocks to a lesser degree; this result indicates why the federal funds rate may be a poor proxy for the stance of policy in that particular period. In the biweekly data (Table 3), the estimate of ϕ^d is close to one in all subsamples, except the nonborrowed-reserves-targeting period, 1979-82.

So far we have compared existing models in a non-nested fashion, by testing each model's overidentifying restriction. This approach has two drawbacks: It gives only the statistical, and not economic, significance of model rejections, and it does not allow for direct comparison of models. An alternative strategy is to impose only one restriction on the model of the bank reserves market, to get just-identification; and then, conditional on that restriction, to observe how closely the estimated parameter values of the more general model correspond to those assumed by the more restricted models discussed above.

As we noted, Strongin (1992) argues that the Fed has no choice but to accommodate reserves demand shocks in the short run, either through provision of nonborrowed reserves or through the discount window, so that shocks to total reserves reflect only changes in the demand for reserves in the short run. In terms of the parameters of the model, Strongin's assumption implies that $\alpha=0$, i.e., the innovation in total reserves depends only on the demand shock and not on the current funds rate. As we are using high-frequency data, for which Strongin's assumption seems most plausible; and as the restriction $\alpha=0$ does not seem grossly inconsistent with the estimates discussed thus far[23], we imposed that restriction and re-estimated the model of the reserves market. We refer to the resulting model as Model A. Conditional on the estimates of Model A, we are then able to test the parameter restrictions implied by Models B, C, S, and BR.

The basic results of estimating the just-identified model and the associated hypothesis tests are given in Tables 4 (monthly data) and 6 (biweekly data). The findings largely reinforce the conclusions drawn from Tables 2 and 3. In the monthly data, we find that Model B (funds-rate smoothing) is clearly the best in the periods 1966-72, 1972-79, and 1988-94, as per the conventional wisdom; in particular, the estimates for the policy response parameters ϕ^d and ϕ^b in those periods are close to their theoretical values of 1 and -1, and the equality of the estimates to the theoretical parameters is never rejected statistically. In contrast, the evidence for interest-rate smoothing is much weaker in 1984-88. Note that in the latter period, we find that ϕ^d and ϕ^b, although less than one in absolute value, remain approximately equal and opposite. This suggests an

[23]As just noted, most estimates of α were small and of the wrong sign. Also, the Strongin model, which imposes $\alpha=0$, was not rejected for the biweekly sample as a whole or for the second half of the monthly sample.

operating regime in which both nonborrowed reserves and the funds rate should receive weight as indicators of Fed policy (see equation 14). Model S continues to perform relatively well in the second half of the sample, presumably because it allows for a policy that lies between pure interest-rate smoothing and pure nonborrowed-reserves targeting. Interestingly, the borrowed-reserves model (Model BR) generally performs poorly under the restriction that $\alpha=0$, in contrast to Table 2 where we were not able to distinguish it from Model B.

To this point we have not reported the estimates of the variances of the structural shocks (to total reserves demand, policy, and borrowings demand); since these are not measured in comparable units, the raw estimates are not particularly useful. To present the variance estimates in a more interpretable way, Table 5 shows the share of federal funds rate innovations attributable to each of the three structural shocks in each sub-period (similar exercises could be performed for innovations to total and nonborrowed reserves). Note that these variance shares depend on the parameters describing the Fed's operating procedure as well as the shock variances themselves. Consistent with the finding that the funds rate is a good policy indicator for the subperiods 1966-72, 1972-79, and 1988-94, Table 5 shows that policy shocks accounted for 95%, 91%, and 99%, respectively, of the variance of funds-rate innovations in those three subperiods.[24] In 1984-88, by contrast, borrowings shocks accounted for 45%, and policy shocks for only 42%, of the variance of the funds rate. Shocks to the demand for total reserves generally account for a negligible fraction of the variance of the funds rate, not so much because the

[24]Brunner (1994) obtains a similar finding.

variance of those shocks is small but rather because of the Fed's policy of offsetting the effects of reserve demand shocks on interest rates.

In the biweekly data (Table 6), we again find that Model B is dominant prior to 1979 and in the recent periods, and that Model S is the best model for the post-1979 period taken as a whole. The parameter estimates suggest that the Fed offset reserves demand shocks quite consistently throughout the sample period (i.e., ϕ^d is close to 1), except in 1979-1982. Borrowings shocks were also offset at the biweekly frequency but less completely, which is why the data sometimes have difficulty in choosing between the Bernanke-Blinder ($\phi^b = -1$) and Strongin ($\phi^b = 0$) specifications. Again Model C (nonborrowed-reserves targeting) is consistently rejected, except during the 1979-82 period where (as before) it is the preferred model. Model BR (borrowed-reserves targeting) is inferior on the p-value criterion to Model B in all subsamples, except 1979-82.

To this point we have allowed for parameter instability by estimating the model within fixed sub-periods. An alternative way to see what is going on is to estimate the model repeatedly for rolling samples. This strategy is implemented in Figure 5, which shows the evolution of estimated Fed reaction coefficients ϕ^d and ϕ^b obtained from estimating Model A (in monthly data) for every possible nine-year sample. In the figure, dates given on the horizontal axis represent the end-point of the sample, and dashed lines indicate two-standard-deviation error bands. Figure 5 shows that the Model B (funds-rate targeting) restrictions, $\phi^d = 1$ and $\phi^b = -1$, fit the data extremely well before 1979 and in the most recent period. In the period around 1982, the Fed's behavior was best described by nonborrowed-reserves targeting (Model C), that is, $\phi^d = \phi^b = 0$. These results fit quite comfortably with our previous findings and with the conventional wisdom. During the mid-1980s the Fed's behavior seems best described by $\phi^d = 1$, $\phi^b = 0$, which is consistent with either the Strongin model or the borrowed-reserves-targeting model.

To summarize, if the choice is restricted to the four simple models compared here (Models B, C, S, and BR), the results suggest using the Bernanke-Blinder funds-rate measure of monetary policy for the periods prior to 1979 and after 1988 (i.e., at all times except during the tenure of Chairman Volcker), and the Strongin measure otherwise. Conditional on the assumption that $\alpha = 0$, however, another possibility is to use the policy measures implied by Model A, which nests each of the simpler models as special cases. In the rest of this section we consider measures based on the just-identified Model A.

In general, there are two ways to use these model-based policy measures. First, we can construct an overall measure of policy stance, including the endogenous or anticipated component of policy. Second, we can look at the innovations to the implied policy measure (the "policy shocks") as well as at the dynamic responses of other variables in the VAR system to policy shocks. The three panels of Figure 6 show, respectively, the total measure of policy stance as derived from Model A; the shocks or innovations to the policy measure; and the anticipated or endogenous component of the monetary policy measure. Units have been chosen so that changes in the various measures can be interpreted analogously to changes in the federal funds rate (e.g., a rise of 2.0 in the total measure, which is a monetary easing, is comparable to a two-percentage point decline in the federal funds rate). The pictures in Figure 6 are based on estimates for the whole sample; total policy measures for some important subsamples are given in Figure 7. Romer dates and the Boschen-Mills index are included for comparison to our derived total measure of policy stance.

Comparison of our total measure to the qualitative indicators provides a useful check on the plausibility of both types of measures. As can be seen in Figures 6 and 7, the similarity of our estimated measure of policy stance to the Boschen-Mills index--both in the whole sample and in sub-samples--is particularly striking, considering that the two measures are derived from different information by very different methods. Comparison of our policy measure to the Romer dates also show general agreement about which periods represent episodes of monetary tightening. Note though that our policy measure peaks considerably before the occurrence of Romer dates. In particular, Romer dates seem to correspond to periods of maximum tightness of policy (troughs in the indicator) rather than to points at which the thrust of policy changed from easing to tightening.

Table 7 reports various subsample correlations of our total measure of policy stance with the Boschen-Mills index and the raw policy measures shown in Figures 1-3. (Our policy measure and the Strongin measure are re-estimated within each subsample.) Consistent with our earlier findings, the funds-rate-based measure (here taken to be the funds-rate Treasury-bond spread) is highly correlated with the measure based on Model A prior to 1979 and after 1988. In the 1984-88 period, however, the correlation of the two measures is negative! Strongin's measure, the orthogonal component of nonborrowed reserves, has the second-highest correlation overall with the total measure and is the most closely correlated of the various measures in 1984-88 (although the correlation is only 0.339). Nonborrowed reserves growth has a generally low correlation with the total measure, except, somewhat surprisingly, in 1988-94.

What does our measure say about the historical performance of monetary policy? Comparison of the top two panels of Figure 6 suggest a few tentative points: First, monetary policy was unusually tight (relative to the normal reaction function) in 1966 and 1969; quite loose during the 1970s; and very tight in 1980-81. These conclusions are consistent with standard historical accounts. Since about 1983, monetary policy appears to have been both more stable and predictable on average, although some "stop-go" pattern is still apparent in the measure of policy shocks. Most recently, although the total measure suggests that policy was expansionary, negative policy shocks in 1992-93 suggest that policy was less expansionary than might have been expected, given the state of the economy. Thus critics who argue that the Fed did not ease aggressively enough during the recovery period receive some support from our measures.

A second way to use the policy measure generated by our technique is to examine the behavior of the economy following policy "shocks". Much previous VAR-based work has displayed the dynamic responses of a long list of variables to monetary policy innovations, as measured by various indicators. Figure 8 replicates this exercise by showing the 48-month impulse responses of the six variables in the monthly VAR to a positive shock to the policy measure based on Model A.[25] For ease of interpretation, the shock has been normalized so that it produces a 25-basis point decline

[25]An innovation to the policy measure based on Model A is simply a linear combination of innovations to total reserves, nonborrowed reserves, and the funds rate; see equation (14). The impulse responses shown in Figure 8 are for the whole sample period, which is somewhat illegitimate since we cannot reject the existence of structural breaks. Subsample impulse responses yield similar pictures.

in the funds rate on impact. Two-standard-deviation confidence intervals calculated by Monte Carlo methods are included in the figures.

The results are largely in line with expectations. A positive monetary shock lowers the funds rate, raises total and nonborrowed reserves, and raises output and prices. All responses are statistically significant. Note that there is no "price puzzle": Both the GDP deflator and commodity prices rise quite reliably following a positive policy shock. There is a bit of an "output puzzle", as output dips slightly before beginning to rise. The output puzzle may reflect inventory decumulation at the beginning of the expansion; this phenomenon is not apparent in impulse responses drawn in quarterly data. An interesting result, which has been noticed previously in similar exercises, is that the effects of the policy shock on output and prices last much longer than the effects on reserve-market variables. We will not speculate here on whether this finding is due to a misspecification or to the existence of propagation mechanisms that transform short-lived policy shocks into longer-lived economic effects.[26]

How do these variables' dynamic responses to an innovation to our policy measure correspond to their responses to alternative measures of policy shocks? Figure 9 shows that, for the whole sample and various subsamples, dynamic responses to alternative measures of policy shocks are similar in shape but sometimes differ in amplitude. Quantitative measures of these differences are given by Table 8, which shows the average absolute deviation (over a 48-month horizon) of the impulse responses generated by various models, relative to the impulse responses generated by Model A.

[26]Bernanke and Gertler (forthcoming) argue that the "balance-sheet" channel of monetary policy can rationalize persistent effects of transitory policy shocks.

Deviations are normalized by the standard errors of the impulse responses from Model A at each horizon and for each variable; this normalization both put deviations in interpretable units and has the effect of downweighting deviations at horizons where the impulse responses are estimated imprecisely. The results suggest that Model B (funds-rate-targeting) yields impulse response functions most similar to those of Model A in the whole sample, in 1965-79, and in 1984-94. The Strongin model yields the most similar impulse responses for 1979-94 subsample. These results are not surprising, given our findings discussed above. Note that if we treat Model A as the "truth", it appears that using the best alternative simple model in any period yields errors that are not negligible, but neither are the deviations so great as to call the qualitative results into question. Hence there is some justification for using these computationally simpler measures in applications.

In VAR-based exercises, impulse response functions are often supplemented by decompositions of forecast variance. We do not find variance decomposition analysis to be particularly informative in the present context, for at least two reasons: First, calculating the share of the forecast variance of non-policy variables due to policy shocks tells us nothing about whether policy is stabilizing or not, since the effects of the systematic portion of policy are excluded. Second, changes in variance decomposition results over subperiods may simply reflect changes in the variances of the structural shocks; "instability" of variance decomposition results over time does not necessarily imply anything about the potency or stability of the monetary policy transmission mechanism. For studying that mechanism, impulse response functions appear to us to be much more useful.

VI. The model with discount rate innovations

In the prior section we simplified our model of the market for bank reserves by ignoring innovations to the discount rate. Here we briefly examine the effects of including discount rate innovations in the model.

To extend the model to incorporate the discount rate, we combine equations (10) and (11), the total reserves demand and borrowing equations, with an extended version of the Fed's reaction function and an equation describing discount rate behavior:

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

(24) $$u_{DISC} = \theta^d v^d + \theta^b v^b + \theta^s v^s + v^{disc}$$

Equation (23), the Fed's reaction function, is the same as equation (12) except that it is augmented by a term ($\phi^{disc}v^{disc}$) that allows the supply of nonborrowed reserves to be adjusted in response to discount rate shocks. Equation (24), a new equation, can also be thought of as a reaction function; it describes how the Fed adjusts the discount rate in response to reserve-market innovations, and it allows for a discount-rate-specific innovation, v^{disc}. Equation (24) is (so far) unrestricted.

The solution of the discount rate model can be written as:

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

$$\mathbf{u} = \begin{vmatrix} u_{TR} \\ u_{NBR} \\ u_{FF} \\ u_{DISC} \end{vmatrix} \qquad \mathbf{v} = \begin{vmatrix} v^d \\ v^s \\ v^b \\ v^{disc} \end{vmatrix}$$

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

Models B, C, S, and BR can all be nested in the above more general model (see Table 9 for the implied restrictions). To conserve space, however, we report results only of a just-identified version of the model.

Just-identification of this model requires two restrictions. As before, we impose the restriction that $\alpha = 0$. To obtain a second restriction, we consider the discount-rate reaction function, equation (24). Informal descriptions suggest that the discount rate is used primarily in two ways: "passively", to adjust to the market level of interest rates, and "actively", as a signal or instrument of policy (Thornton, 1994). We interpret this characterization as saying that the discount rate is likely to be responsive to reserve demand shocks (which influence market interest rates) and policy shocks. Thus θ^d and θ^s should be non-zero. However, because of the discrete and highly publicized nature of discount rate changes, it seems reasonable that the Fed would adjust to borrowing shocks by changing the supply of nonborrowed reserves rather than by changing the discount rate. Hence we achieve just-identification by imposing the restriction $\theta^b = 0$.

Estimates of this model for monthly data are reported in Table 9. The results seem most supportive of a funds-rate indicator for monetary policy. When discount-rate innovations are accounted for, Model B is not rejected for any of the five short sub-periods (although it is rejected for the 1979-94 sub-period and for the sample as a whole). In particular, the 1984-88 subsample, for which Model B was strongly rejected in the earlier results, now seems to fit Model B's restrictions ($\phi^d = 1$, $\phi^b = -1$) quite closely. Overall, the estimates for ϕ^d and ϕ^b now suggest a good deal of interest-rate smoothing throughout the sample period.

The estimates of the discount-rate model are not entirely satisfactory. The parameter θ^s is generally negative and large in absolute value, suggesting that discount rate movements are driven primarily by policy shocks; the value of the parameter is unstable, however, and it is not estimated to be negative in all sub-periods. The parameter θ^d is relatively smaller and of mixed sign, suggesting no systematic effect of reserve demand shocks on the discount rate. A few other "wrong signs" are scattered throughout Table 9. The results are sufficiently interesting, though, to motivate further analysis of the discount-rate model.

VII. Conclusion

We have used a "semi-structural VAR" approach to evaluate and develop measures of monetary policy based on reserve market indicators. For practitioners looking for a simple indicator of policy stance, our results suggest that the federal funds rate, which was the best indicator of policy prior to 1979, has become so again during the tenure of Chairman Greenspan.

However, the funds rate was not necessarily a good indicator of policy during the early to mid-1980s; for studies of the post-1979 period as a whole, Strongin's measure of policy may be the most robust. A more general though slightly more complicated approach is to base the policy measure on an estimated model of the market for bank reserves, along the lines of our Model A.[27] The latter approach has the advantage of being able to incorporate the effects of subtle changes in reserve-market structure and the Fed's operating procedure.

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.

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

Appendix 1. Development of monthly GDP components data

Monthly GDP data were constructed for an ongoing research project by Bernanke with Mark Gertler, with assistance from Peter Simon.

We constructed monthly GDP components and GDP deflator data for the period January 1964-March 1994 using the method of Chow and Lin (1971). To illustrate the method briefly, let y be the (unobserved) monthly values of one of the GDP components, and suppose that there exists a set of variables X that are available monthly and contain information about y. For example, if the GDP component in question is residential investment, then a potentially informative variable that is available monthly is new housing starts. We assume that, at the monthly frequency

(A1) $y = X\beta + u$

where $V = E(uu')$. In our application we assumed that the monthly error term u is AR(1) with unknown serial correlation coefficient ρ. The error covariance matrix V then takes a standard form (see Chow-Lin, eqn. (20)).

If there are t quarters and hence 3t months of data, let C be the t x 3t matrix that converts monthly observations to quarterly averages. Taking quarterly averages of eq. (A1), we have

(A2) $Cy = CX\beta + Cu$

or

(A3) $y_. = X_.\beta + u_.$

where a dot subscript denotes a quarterly average. Define the quarterly error covariance matrix $V_. = E(u_.u_.') = E(Cuu'C') = E(CVC')$. If the monthly errors are AR(1), then of course the implied process for the quarterly average errors will be something more complicated.

Because quarterly values of the GDP components are observable, eq. (A3) can be consistently estimated by OLS. (To reduce heteroscedasticity, we first scale y and X by a 24-quarter moving average of real GDP). Solving a nonlinear equation (Chow-Lin, eq. (23)), we obtain a consistent estimate of the monthly serial correlation coefficient ρ from the estimated first-order correlation coefficient of the quarterly residuals. Using this estimate, we construct the implied covariance matrix of the quarterly residuals $\hat{V}$, then reestimate eq. (A3) by GLS. Finally, the estimated monthly values for the GDP component $\hat{y}$ are computed by Chow-Lin's formula

(A4) $\hat{y} = X\hat{\beta} + \hat{V}C'(C\hat{V}C')^{-1}\hat{u}_.$

As Chow and Lin note, this formula has the property that, in each quarter, the average of monthly imputed values of the GDP component equals the actual quarterly value. (Monthly GDP components are rescaled by the moving average of GDP and expressed at annual rates).

Interpolation for the components of consumption spending was unnecessary as these are available monthly. Monthly inventory-stock data are available back only to 1967:1; for the previous three years, we simply assumed that quarterly inventory investment was distributed equally among the months of each quarter. In a few cases, the monthly interpolating variables were not available for the whole period; in that situation, we ran subperiod regressions corresponding to data availability. Monthly GDP is constructed as the simple sum of monthly GDP components.

A list of interpolated GDP components (all in real terms) and the associated monthly interpolating variables is given below. Data series with start dates after 1964:1 are noted. All interpolating equations included constants.

Interpolated variable (CITIBASE name)	Monthly interpolators (CITIBASE names)
1. GDP deflator (GD)	PCE deflator (GMDC) PPI, finished goods (PWFSA) PPI, finished goods-- capital equipment (PWFPSA) PPI, intermediate materials, supplies, and components (PWIMSA) PPI, crude materials (PWCMSA)
2. Producers' durables investment (GIPDQ)	IP index, business equipment (IPE) Shipments of non-defense capital goods (MSNDF) (1968:1-) Shipments of machinery and equipment (MSMAE) (1968:1-)
3. Structures investment (GISQ)	IP index, intermediate products-- construction supplies (IPIC) Shipments of construction supplies (MMCON) New construction--industrial (CONIC) New construction--commercial (CONCC) (1972:1-)
4. Residential investment (GIRQ)	IP index, intermediate products-- construction supplies (IPIC) Shipments of construction supplies (MMCON) New construction of private residential buildings (CONFRC) New housing units started (HSF)
5. Inventory investment (GVQ)	Inventory investment, manufacturing (DIVMFM) (1967:2-) Inventory investment, retail trade (DIVRM) (1967:2-) Inventory investment, wholesale trade (DIVWM) (1967:2-)
6. Government purchases (GGEQ)	New construction--public (CONQC) IP index, defense and space equipment (IPH) Federal expenditures (net) (FBO) (seasonally adjusted by monthly dummies) (1968:1-)

7. Exports (GEXQ)	Merchandise exports (excluding military aid shipments) (FSE602) Exports of nonelectronic machinery (FTE71) (1965:1-) Exports of agricultural products (FTEF) (1965:1-)
8. Imports (GIMQ)	General imports (FSM612) Imports, petroleum and petroleum products (FTM333) (1965:1-) Imports, autos and parts (FTM732) (1965:1-)

Appendix 2. GMM estimation

Switching to a standard notation, we write the complete model (e.g., equations (5)-(6) in the text) in the form:

$$Y_t = BX_t + u_t$$
$$u_t = Av_t$$
$$E(u_t u_t') = \Omega$$

We are estimating a $nx1$ parameter vector θ_0 which can be partitioned as

$$\theta_0 = \begin{pmatrix} \beta_0 \\ \lambda_0 \end{pmatrix}$$

where β_0 is the n_1x1 vector of VAR coefficients and λ_0 is the n_2x1 vector of coefficients of the structural model of the bank reserves market.

Let $Z_t = (Y_t, X_t)$. Then $\theta_0 \in R^n$ satisfies the moment restrictions $E[m(Z_t, \theta_0)] = 0$ where $m: \mathbf{Z} x R^n \rightarrow R^r$. To perform two-step estimation we have to be able to partition the vector of moment restrictions as:

$$m(Z_t, \theta_0) = \begin{bmatrix} m_1(Z_t, \beta_0) \\ m_2(Z_t, \beta_0, \lambda_0) \end{bmatrix}$$

In the case we are considering the two sets of moment restrictions are

$$E\begin{bmatrix} vec(X_t u_t') \\ vech(\Omega) - vech(u_t u_t') \end{bmatrix} = 0$$

which satisfies the partioning criterion.

The object of particular interest here is the vector of GMM estimates of the structural parameters, $\hat{\lambda}_{GMM}$, which is estimated in the second stage. Consistency of this estimator follows from Theorem 2.3 in Hansen (1982). Asymptotic normality of the estimator is a direct extension of the proof of Hansen's (1982) Theorem 3.1, after partitioning conformably the derivatives of the vectors of moment restrictions with respect to the vector of parameters and accounting for the first-stage estimation.

This leaves the question of efficiency of the estimator. Normally, to achieve efficient estimates one would have to adjust the weighting matrix to take into account the first-stage estimation. In the remainder of this appendix we show that, under the specific moment restrictions we have imposed, no adjustment to the asymptotic variance-covariance matrix of the GMM estimators or the optimal weighting matrix is needed, i.e., uncertainty about the estimates of the variance-covariance matrix of VAR residuals can be ignored.

Let

$$M_0 \equiv E\left[\frac{\partial m(Z_t, \theta_0)}{\partial \theta'}\right]$$

and

$$V_0 \equiv E[m(Z_t,\theta_0)m'(Z_t,\theta_0)]$$

be rxn and rxr matrices respectively. Partition these matrices analogously to our partitions of the parameter vector and the moment restrictions and denote the appropriate submatrices by $M_{1\beta}$, $M_{2\beta}$, $M_{2\lambda}$, V_{11}, V_{12}, V_{21}, and V_{22}. Then it can be shown in the general case of sequential estimation that $\sqrt{T}(\hat{\lambda}_{GMM} - \lambda_0)$ converges in distribution to $N(0, M_{2\lambda}'\Xi M_{2\lambda})^{-1}$, where

$$\Xi = (-M_{2\beta}(M_{1\beta}'W_1M_{1\beta})^{-1}M_{1\beta}'W_1 \quad I)\begin{pmatrix} V_{11} & V_{12} \\ V_{21} & V_{22} \end{pmatrix}\begin{pmatrix} -W_1'M_{1\beta}(M_{1\beta}'W_1M_{1\beta})^{-1'}M_{2\beta}' \\ I \end{pmatrix}$$

where W_1 is the weighting matrix used in the first step (see, for example, Ogaki, 1993). For the specific moment restrictions of our application it turns out that $M_{2\beta} = 0$, since

$$M_{2\beta} = E[\tfrac{\partial m_2(Z_t,\beta_0,\lambda_0)}{\partial\beta}] = E[\tfrac{\partial(vech(\Omega) - vech(u_tu_t'))}{\partial\beta}] = E[Q] = 0$$

where Q is a matrix whose elements are either zero or of the form $X_{jt}u_{it}$, where j indexes regressors and i indexes equations. The last equality is ensured by the fact that in a symmetric VAR the regressors are the same in each equation, so that regressors from any equation are orthogonal to the error terms of all equations. Hence

$$\sqrt{T}(\hat{\lambda}_{GMM} - \lambda_0) \xrightarrow{d} N(0,(M_{2\lambda}'V_{22}M_{2\lambda})^{-1})$$

It follows that we do not have to adjust standard errors to allow for the fact that we are using a two-step procedure; and that the optimal weighting matrix is just the inverse of the estimated covariance matrix of the sample moment restrictions, or V_{22}^{-1}.

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**Table 1: Correlation Among the Raw Measures of Monetary Policy Stance
1965 - 1994**

	Spread(inverted)	NBR growth	NBR ⊥ TR
Spread(inverted)	1.000		
NBR growth	0.242	1.000	
NBR ⊥ TR	0.701	0.185	1.000
Boschen-Mills	0.600	0.557	0.546

Table 2: Parameter Estimates for the Over-Identified Models (Monthly)
Information assumption: The Fed reacts to contemporaneous information

	Model	α	β	ϕ^d	ϕ^b	Test for OIR
1965:1 - 1994:3	B	-0.004	0.012	1	-1	0.078
	C	0.029	0.014	0	0	0.000
	S	0	0.057	0.848	0	0.024
	BR	-0.004	0.042	1	α/β	0.078
1965:1 - 1979:9	B	-0.005	0.012	1	-1	0.234
	C	0.031	0.066	0	0	0.000
	S	0	0.099	0.834	0	0.013
	BR	-0.005	0.081	1	α/β	0.234
1979:10 - 1994:3	B	-0.002	0.013	1	-1	0.055
	C	0.043	0.013	0	0	0.000
	S	0	0.045	0.802	0	0.591
	BR	-0.002	0.035	1	α/β	0.055
1966:1 - 1972:10	B	-0.008	0.007	1	-1	0.937
	C	0.013	0.059	0	0	0.001
	S	0	0.170	0.958	0	0.104
	BR	-0.008	0.169	1	α/β	0.950
1972:11 - 1979:9	B	-0.006	0.018	1	-1	0.647
	C	0.018	0.076	0	0	0.002
	S	0	0.093	0.833	0	0.028
	BR	-0.006	0.095	1	α/β	0.647
1984:2 - 1994:3	B	-0.008	0.007	1	-1	0.134
	C	-0.572	0.007	0	0	0.057
	S	0	0.480	0.868	0	0.191
	BR	-0.009	0.123	1	α/β	0.136
1984:2 - 1988:10	B	-0.014	0.009	1	-1	0.021
	C	-0.140	0.003	0	0	0.008
	S	0	0.430	0.344	0	0.088
	BR	-0.014	0.137	1	α/β	0.021
1988:11 - 1994:3	B	0.002	0.002	1	-1	0.967
	C	0.406	0.003	0	0	0.030
	S	0	0.148	1.007	0	0.644
	BR	0.002	0.143	1	α/β	0.967

Notes: The estimates come from a 6-variable monthly VAR which includes interpolated real GDP (GDPM), interpolated GDP deflator (PGDPM), commodity price index (PCOM), total reserves (TR), nonborrowed reserves plus extended credit (NBR) and federal funds rate (FFR). GDPM, PGDPM, and PCOM are in logs, TR and NBR are normalized by 36-month MA of TR. The last two 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 last column presents p-values from a test of overidentifying restrictions based on the minimized value of the criterion function.

Table 3: Parameter Estimates for the Over-Identified Models (Biweekly)

Information assumption: The Fed reacts to contemporaneous information

	Model	α	β	ϕ^d	ϕ^b	Test for OIR
1967:1	B	-0.003	0.014	1	-1	0.136
-	C	0.064	0.015	0	0	0.000
1994:2	S	0	0.093	0.890	0	0.087
	BR	-0.003	0.083	1	α/β	0.136
1967:1	B	-0.005	0.017	1	-1	0.427
-	C	0.011	0.077	0	0	0.000
1979:9	S	0	0.137	0.877	0	0.074
	BR	-0.005	0.133	1	α/β	0.427
1979:10	B	-0.003	0.013	1	-1	0.103
-	C	0.090	0.012	0	0	0.000
1994:2	S	0	0.081	0.852	0	0.305
	BR	-0.003	0.062	1	α/β	0.103
1967:1	B	0.001	0.022	1	-1	0.555
-	C	0.016	0.045	0	0	0.000
1972:10	S	0	0.065	0.970	0	0.469
	BR	0.001	0.065	1	α/β	0.555
1972:10	B	-0.002	0.022	1	-1	0.995
-	C	-0.037	-0.088	0	0	0.000
1979:9	S	0	0.182	0.985	0	0.659
	BR	-0.002	0.180	1	α/β	0.995
1984:2	B	-0.007	0.006	1	-1	0.085
-	C	0.592	0.014	0	0	0.014
1994:2	S	0	1.265	0.881	0	0.319
	BR	-0.007	0.280	1	α/β	0.085
1984:2	B	- 0.008	0.007	1	-1	0.369
-	C	0.059	0.043	0	0	0.000
1988:10	S	0	0.732	0.936	0	0.143
	BR	-0.009	0.614	1	α/β	0.380
1988:11	B	-0.008	0.010	1	-1	0.349
-	C	1.028	0.015	0	0	0.000
1994:2	S	0	0.094	1.022	0	0.299
	BR	-0.007	0.098	1	α/β	0.352
1979:9	B	-0.002	0.011	1	-1	0.102
-	C	-0.000	0.061	0	0	0.627
1982:10	S	0	0.058	0.139	0	0.657
	BR	-0.002	0.047	1	α/β	0.102

Notes: The estimates are from 5-variable biweekly VARs which include Business Week Production Index (BWPI), commodity price index (PCOM), total reserves (TR), nonborrowed reserves plus extended credit (NBR), federal funds rate (FFR). BWPI and PCOM are in logs, TR and NBR are normalized by 52-week MA of TR. The first three VARs use 26 lags; the VAR for the 1979-82 sub-sample is estimated with 11 non-consecutive lags (1 to 6 and 10,14,18,22,26); the rest of the VARs use 16 non-consecutive lags - skipping every other lag after the 6th one. The last column presents p-values from a test of overidentifying restrictions based on the minimized value of the criterion function..

Table 4: Parameter Estimates for Just-Identified Model A (Monthly)
Information assumption: The Fed reacts to contemporaneous information

	Parameter Estimates			Single Parameter Tests				Joint Tests		
	β	ϕ^d	ϕ^b	$\phi^d = 0$	$\phi^b = 0$ (S)	$\phi^d = 1$	$\phi^b = -1$	$\phi^d = 0$ $\phi^b = 0$ (C)	$\phi^d = 1$ $\phi^b = -1$ (B)	$\phi^d = 1$ $\phi^b = 0$ (BR)
1965:1 - 1994:3	0.020	0.801	-0.661	0.000	0.005	0.001	0.151	0.000	0.003	0.000
1965:1 - 1979:9	0.023	0.785	-0.752	0.000	0.002	0.039	0.317	0.000	0.091	0.000
1979:10 - 1994:3	0.034	0.786	-0.156	0.000	0.628	0.010	0.009	0.000	0.000	0.035
1966:1- 1972:10	0.006	0.962	-1.005	0.000	0.000	0.681	0.941	0.000	0.415	0.000
1972:11 - 1979:9	0.012	0.870	-1.052	0.000	0.000	0.398	0.353	0.000	0.049	0.000
1984:2 - 1994:3	0.031	0.846	-0.712	0.000	0.010	0.054	0.296	0.000	0.150	0.001
1984:2 - 1988:10	0.066	0.491	-0.460	0.008	0.058	0.007	0.026	0.030	0.020	0.000
1988:11 - 1994:3	0.004	1.004	-0.975	0.000	0.265	0.928	0.977	0.000	0.949	0.000

Notes: The estimates come from 6-variable monthly VARs which include interpolated real GDP (GDPM), interpolated GDP deflator (PGDPM), commodity price index (PCOM), total reserves (TR), nonborrowed reserves plus extended credit (NBR), and federal funds rate (FFR). GDPM, PGDPM, and PCOM are in logs, TR and NBR are normalized by 36-month MA of TR. The first three VARs are estimated with 12 lags while the rest of the VARs are using 6 non-consecutive lags (1, 2, 3, 6, 9, 12). The last seven columns present p-values.

Table 5: Contributions of Structural Shocks to the Variance of the Federal Funds Rate Innovations
Model A (Monthly)

	Structural Shock		
	v^d	v^s	v^b
1965:1 – 1994:3	3.69%	82.48%	13.83%
1965:1 - 1979:9	5.17	83.21	11.62
1979:10 - 1994:3	1.63	44.47	53.90
1966:1 - 1972:10	5.33	94.60	0.07
1972:11 - 1979:9	6.40	90.95	2.65
1984:2 - 1994:3	4.99	71.82	23.19
1984:2 - 1988:10	12.44	42.29	45.28
1988:11 - 1994:3	0.25	98.61	1.13

Notes: The estimates, which are used in this decomposition, come from 6-variable monthly VARs which include interpolated real GDP (GDPM), interpolated GDP deflator (PGDPM), commodity price index (PCOM), total reserves (TR), nonborrowed reserves plus extended credit (NBR), and federal funds rate (FFR). GDPM, PGDPM, and PCOM are in logs, TR and NBR are normalized by 36-month MA of TR. The first three VARs are estimated with 12 lags while the rest of the VARs are using 6 non-consecutive lags (1, 2, 3, 6, 9, 12).

Table 6: Parameter Estimates for Just-Identified Model A (Biweekly)

Information assumption: The Fed reacts to contemporaneous information

	Parameter Estimates			Single Parameter Tests				Joint Tests		
	β	ϕ^d	ϕ^b	$\phi^d=0$	$\phi^b=0$ (S)	$\phi^d=1$	$\phi^b=-1$	$\phi^d=0$ $\phi^b=0$ (C)	$\phi^d=1$ $\phi^b=-1$ (B)	$\phi^d=1$ $\phi^b=0$ (BR)
1967:1 - 1994:2	0.030	0.880	-0.642	0.000	0.047	0.014	0.269	0.000	0.049	0.000
1967:1 - 1979:9	0.032	0.871	-0.781	0.000	0.036	0.119	0.558	0.000	0.213	0.000
1979:10 - 1994:2	0.037	0.847	-0.390	0.000	0.357	0.024	0.150	0.000	0.059	0.017
1967:1- 1972:10	-0.024	0.971	-0.643	0.000	0.714	0.814	0.839	0.000	0.971	0.163
1972:11 - 1979:9	0.021	0.981	-1.004	0.000	0.135	0.863	0.995	0.000	0.894	0.000
1979:9 - 1982:10	0.039	0.119	-0.144	0.765	0.722	0.027	0.035	0.905	0.013	0.075
1984:2 - 1988:10	0.025	0.874	-0.890	0.000	0.000	0.274	0.579	0.000	0.484	0.000
1988:11 - 1994:2	-0.044	1.021	-0.377	0.000	0.689	0.487	0.508	0.000	0.775	0.185
1984:2 - 1994:2	0.055	0.869	-0.504	0.000	0.259	0.079	0.267	0.000	0.179	0.045

Notes: The estimates come from 5-variable biweekly VARs which include Business Week Production Index (BWPI), commodity price index (PCOM), total reserves (TR), nonborrowed reserves plus extended credit (NBR), and federal funds rate (FFR). BWPI and PCOM are in logs, TR and NBR are normalized by 52-week MA of TR. The first three VARs use 26 lags; the 1979-82 VAR uses 11 lags (1 through 6 and then skipping three lags); the rest of the VARs use 16 lags (1 through 6 and then skipping every other lag). The last seven columns present p-values.

Table 7: Correlations Between the Total Measure and the Raw Measures of Monetary Policy Stance

	1965-94	1965-79	1979-94	1966-72	1972-79	1984-88	1988-94
Spread(inverted)	0.757	0.845	0.753	0.860	0.908	-0.432	0.976
NBR growth	0.202	0.010	0.379	0.109	-0.201	0.240	0.930
NBR $\perp$ TR	0.621	0.797	0.735	0.896	0.801	0.339	0.302
Boschen-Mills	0.532	0.412	0.654	0.348	0.581	0.530	0.927

Note: The parameter estimates used to construct the total measure of monetary policy stance are from 6-variable monthly VARs which include interpolated real GDP (GDPM), interpolated GDP deflator (PGDPM), commmodity price index (PCOM), total reserves (TR), nonborrowed reserves plus extended credit (NBR) and federal funds rate (FFR). GDPM, PGDPM, and PCOM are in logs, TR and NBR are normalized by 36-month MA of TR. The identification of the model is secured by imposing $\alpha = 0$.

Table 8: Average Absolute Deviations of the Responses to a Monetary Policy Shock Relative to Baseline Model A

	Responses of	Model B	Model C	Model S
1965 - 1994	GDPM	0.356	0.605	1.373
	PGDPM	0.597	2.464	1.235
	PCOM	0.336	1.660	0.558
	TR	0.394	2.645	1.520
	NBR	0.467	1.776	1.795
	FFR	0.479	1.537	1.368
1965 - 1979	GDPM	0.280	1.469	1.033
	PGDPM	0.665	3.816	0.768
	PCOM	0.301	1.677	1.019
	TR	0.258	1.452	0.898
	NBR	0.323	1.550	1.671
	FFR	0.294	1.398	1.600
1979 - 1994	GDPM	1.033	1.285	0.587
	PGDPM	1.266	1.634	0.719
	PCOM	0.526	1.974	0.306
	TR	0.934	2.847	0.532
	NBR	0.896	2.890	0.527
	FFR	0.896	1.115	0.510
1984 - 1994	GDPM	0.278	0.569	0.649
	PGDPM	0.164	0.308	0.536
	PCOM	0.111	0.225	0.279
	TR	0.190	0.407	0.369
	NBR	0.171	0.368	0.335
	FFR	0.213	0.427	0.549

Notes: The impulse responses are constructed from estimates which come from 6-variable monthly VARs which include interpolated real GDP (GDPM), interpolated GDP deflator (PGDPM), commodity price index (PCOM), total reserves (TR), nonborrowed reserves plus extended credit (NBR), and federal funds rate (FFR). GDPM, PGDPM, and PCOM are in logs, TR and NBR are normalized by 36-month MA of TR. The first three VARs are estimated with 12 lags while the last VAR uses 6 non-consecutive lags (1, 2, 3, 6, 9, 12). The impulse responses are constructed at 48 months horizon. The absolute deviations of Models B, C, and S responses from the responses of Model A are normalized by the standard errors of the impulse responses of Model A.

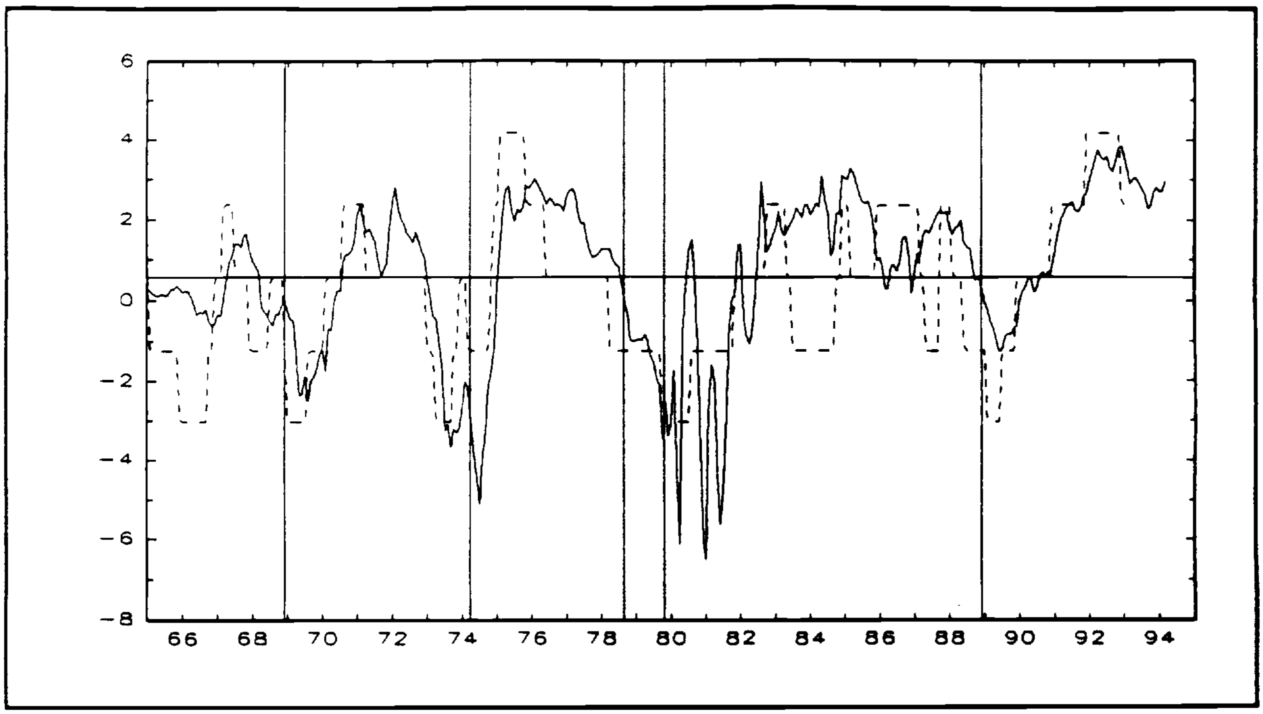
Table 9: Parameter Estimates for a Just-Identified Model with Discount Rate (Monthly)

Information assumption: The Fed reacts to contemporaneous information

	Parameter Estimates						Joint Tests			
	β	ϕ^d	ϕ^b	ϕ^{disc}	θ^d	θ^s	$\phi^d=0$ $\phi^b=0$ $\phi^{disc}=0$ (C)	$\phi^{disc}=0$ $\phi^b=0$ (S)	$\phi^{disc}=\beta$ $\phi^b=-1$ $\phi^d-\beta\theta^d=1$ (B)	$\phi^d=1$ $\phi^b=0$ $\phi^{disc}=0$ (BR)
1965:1 - 1994:3	0.019	0.844	-0.794	0.009	-0.359	-9.245	0.000	0.000	0.011	0.000
1965:1 - 1979:9	0.026	0.862	-0.778	0.038	1.940	-12.57	0.000	0.000	0.172	0.000
1979:10 - 1994:3	0.034	0.905	-0.300	0.006	-2.265	-7.240	0.000	0.016	0.002	0.026
1966:1- 1972:10	0.056	1.082	-0.359	0.010	-2.178	-1.013	0.000	0.709	0.517	0.634
1972:11 - 1979:9	-0.004	1.028	-0.953	0.018	2.567	113.5	0.000	0.002	0.312	0.000
1984:2 - 1994:3	0.024	0.827	-0.880	0.007	-1.672	-16.37	0.000	0.000	0.169	0.000
1984:2 - 1988:10	0.005	0.907	-1.010	0.018	-3.173	-86.39	0.000	0.000	0.297	0.000
1988:11 - 1994:3	-0.015	0.955	-0.753	-0.011	-0.512	41.41	0.000	0.000	0.881	0.000

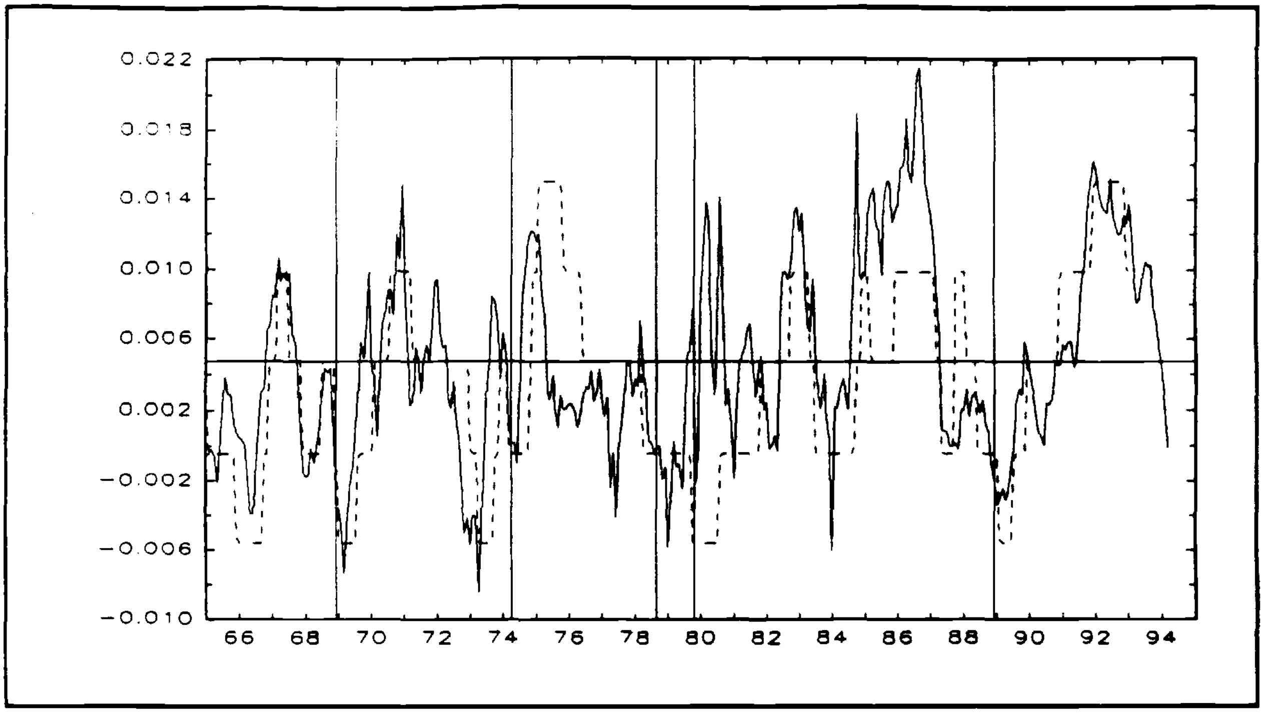
Notes: The estimates come from 7-variable monthly VARs which include interpolated real GDP (GDPM), interpolated GDP deflator (PGDPM), commodity price index (PCOM), total reserves (TR), nonborrowed reserves plus extended credit (NBR), federal funds rate (FFR) and the discount rate (DR). GDPM, PGDPM, and PCOM are in logs, TR and NBR are normalized by 36-month MA of TR. The first three VARs are estimated with 12 lags while the rest of the VARs are using 6 non-consecutive lags (1, 2, 3, 6, 9, 12). The last four columns present p-values.

Figure 1: Federal Funds Rate - Ten-Year Bond Spread
1965 - 1994



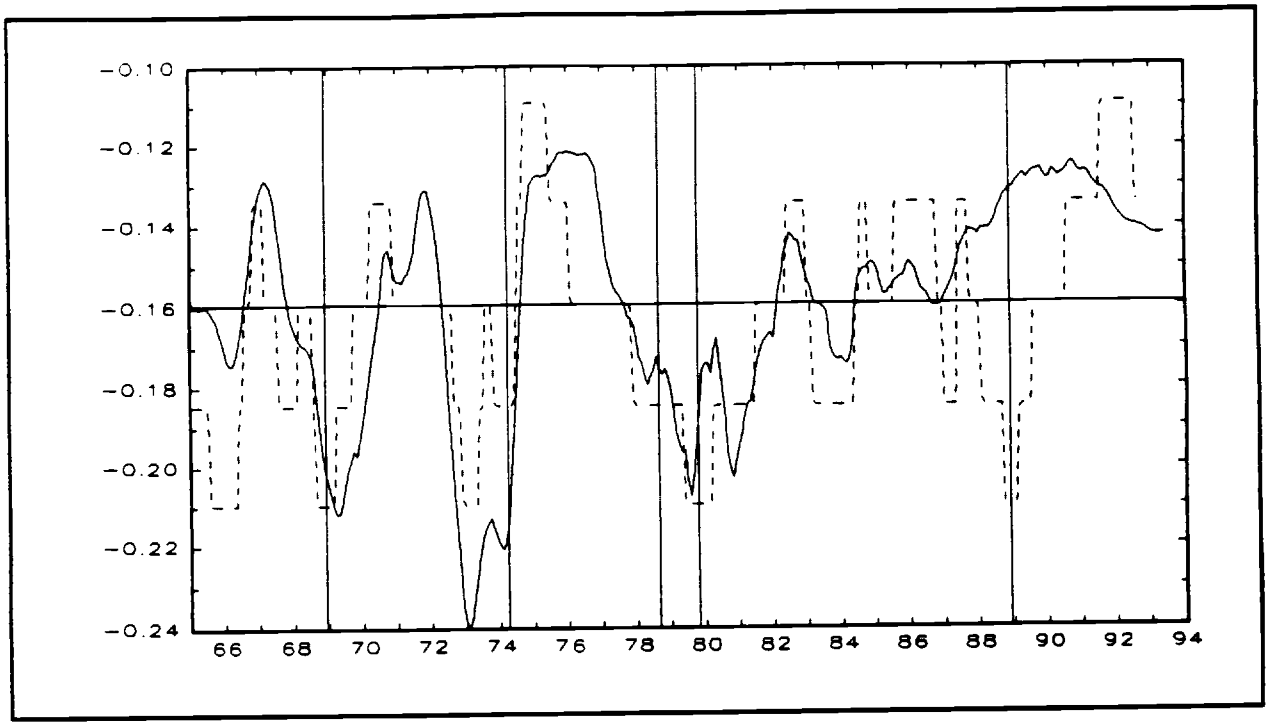
Note: The solid line is the spread between the federal funds rate and the rate on a 10-year Treasury bond (inverted).
The dashed line is the Boschen-Mills index.
The vertical lines are the Romer dates.

Figure 2: Growth Rate of Nonborrowed Reserves
1965 - 1994



Note: The solid line is the growth rate of nonborrowed reserves, 9-month centered average.
The dashed line is the Boschen-Mills index.
The vertical lines are the Romer dates.

Figure 3: The Component of Nonborrowed Reserves Orthogonal to Total Reserves
1965 - 1994



Note: The solid line is the 9-month centered average of the component of nonborrowed reserves (NBR) which is orthogonal to total reserves (TR). Before regressing NBR on TR both series are normalized by a 36-month moving average of total reserves.
The dashed line is the Boschen-Mills index.
The vertical lines are the Romer dates.

Figure 4: The Market for Bank Reserves

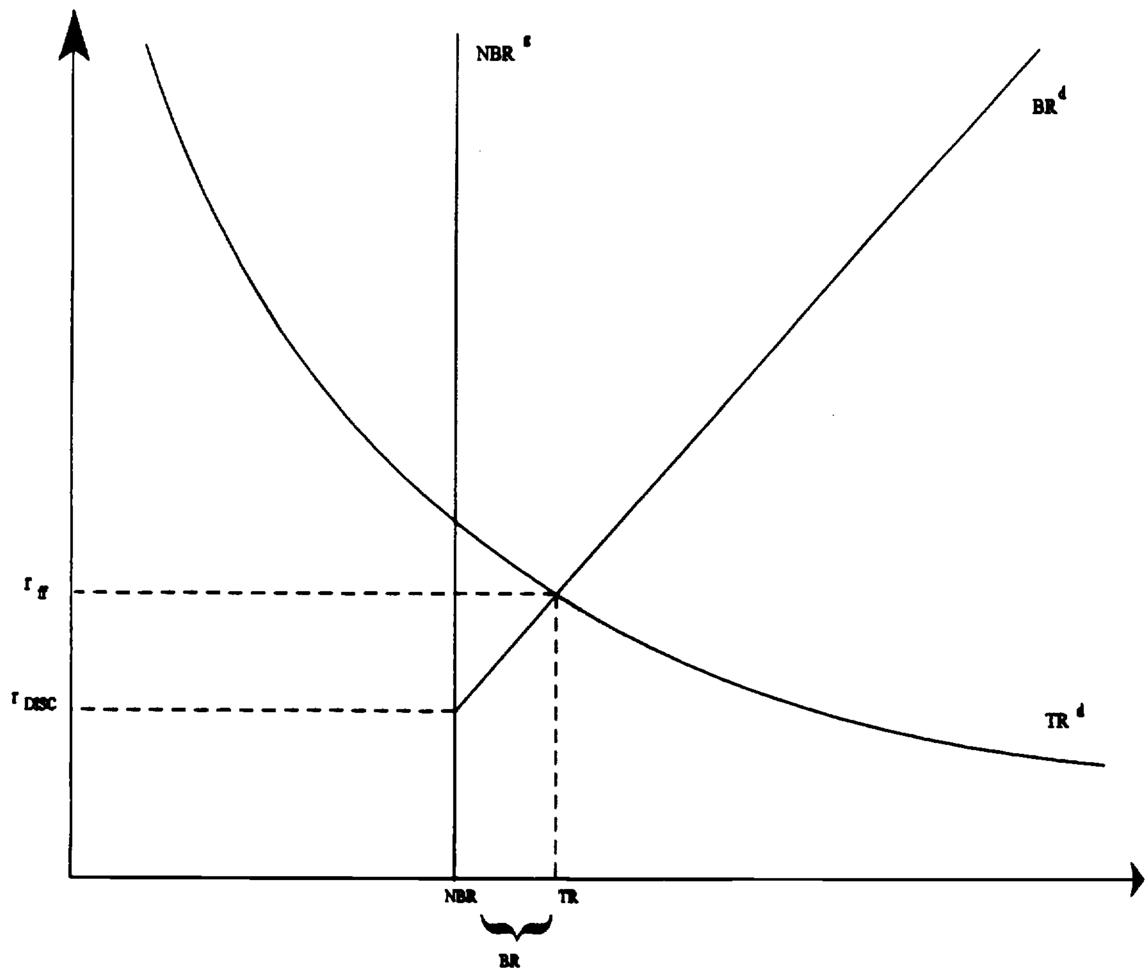
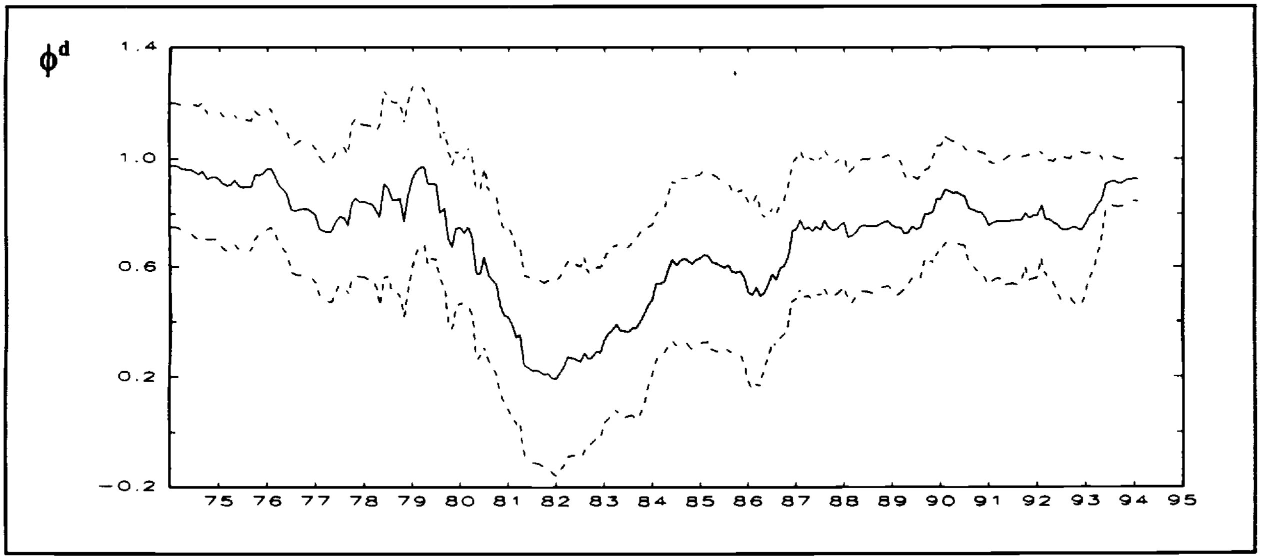
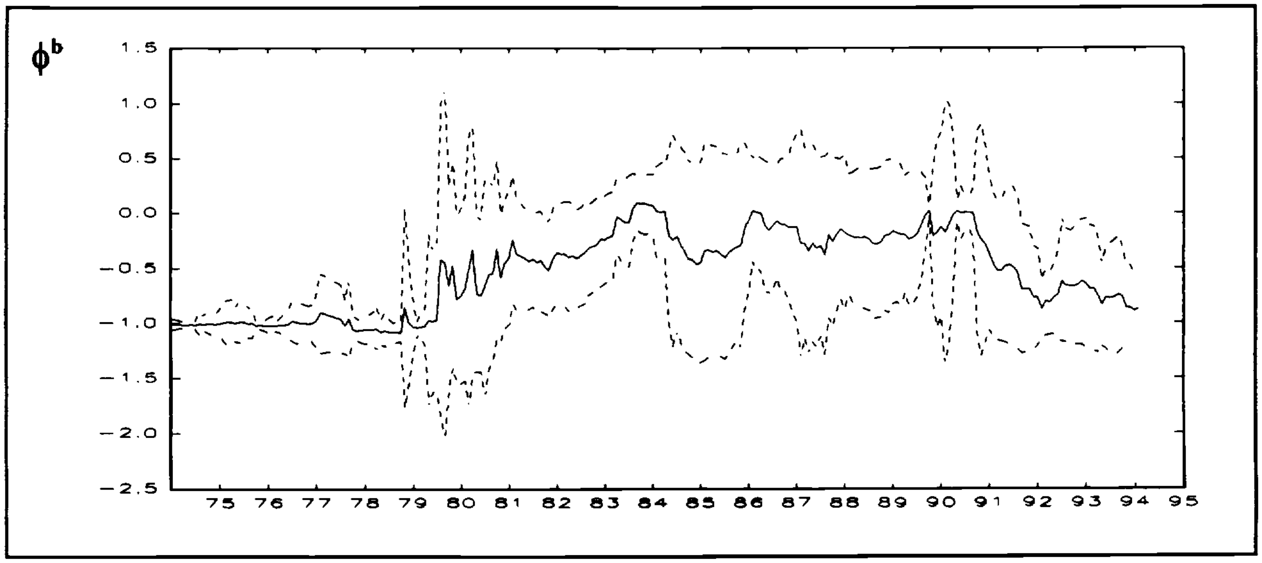


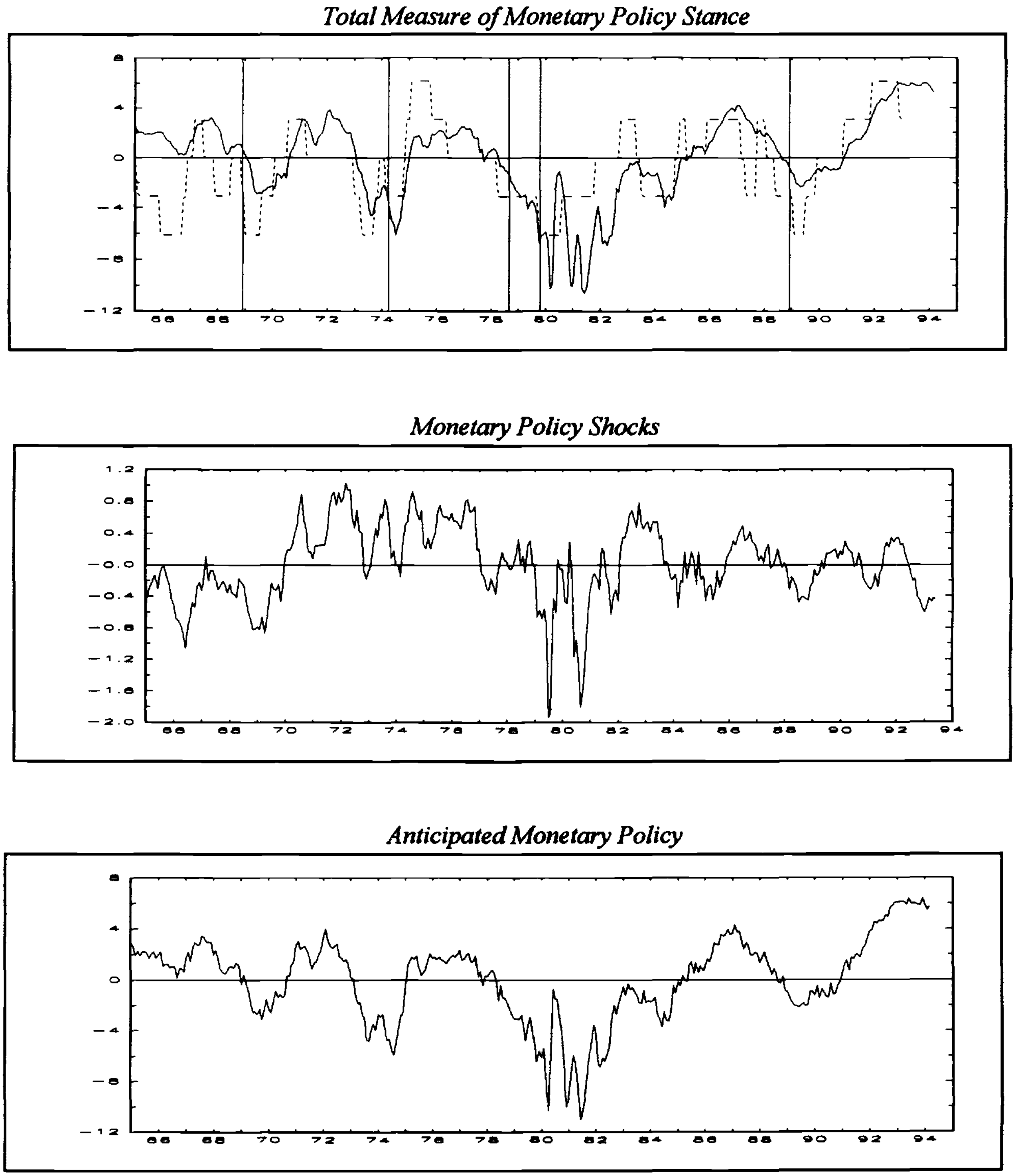
Figure 5: Parameters Describing the Federal Reserve's Operating Procedures As Estimated from a Rolling Regression with a 9-Year Window





Notes: The parameter estimates come from 6-variable monthly VARs which include interpolated real GDP (GDPM), interpolated GDP deflator (PGDPM), commodity price index (PCOM), total reserves (TR), nonborrowed reserves plus extended credit (NBR), and federal funds rate (FFR). GDPM, PGDPM, and PCOM are in logs, TR and NBR are normalized by 36-month MA of TR. The rolling VARs are estimated with 6 non-consecutive lags (1, 2, 3, 6, 9, 12) over a 9-year window. The dates on the graphs are the end-dates for the VARs. Dashed lines are two standard error bands.

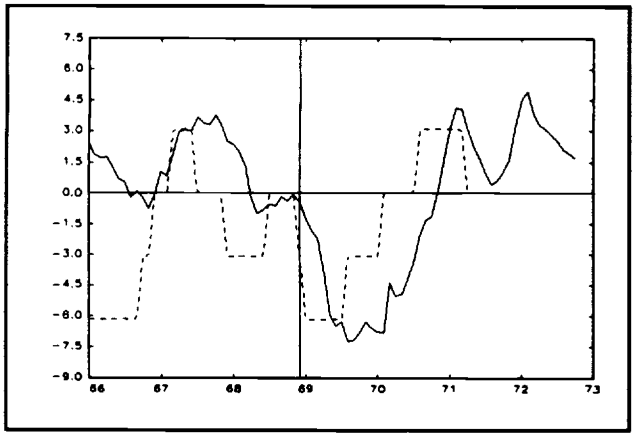
Figure 6: Anticipated and Unanticipated Components of Monetary Policy (Model A)



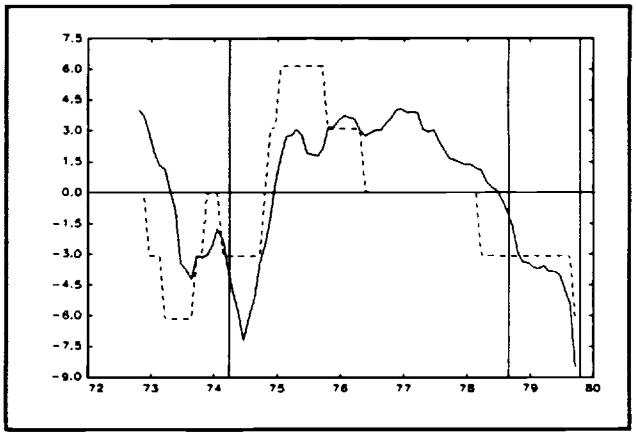
Notes: In the top panel the dashed line is the Boschen-Mills index and the vertical lines are the Romer dates. The solid line in the first panel is the total measure of monetary policy stance rescaled to have the same variance as the federal funds rate and zero mean. The line in the second panel displays monetary policy shocks rescaled to have the same variance as innovations to the federal funds rate. The anticipated component of monetary policy, displayed in the third panel, is rescaled to have the variance of the anticipated component of the federal funds rate.

Figure 7: Total Measure of Monetary Policy Stance
Just-Identified Model A ($\alpha = 0$)

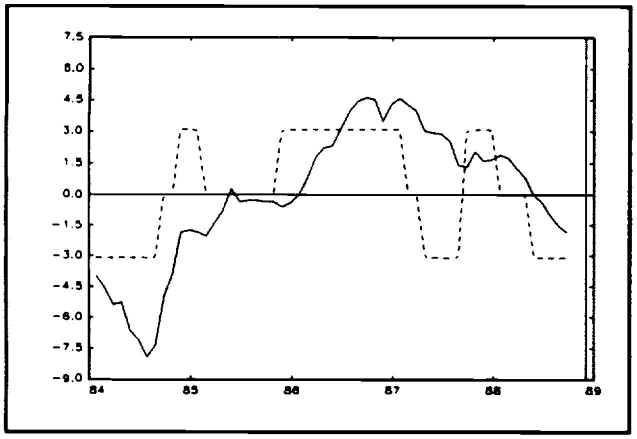
1966 - 1972



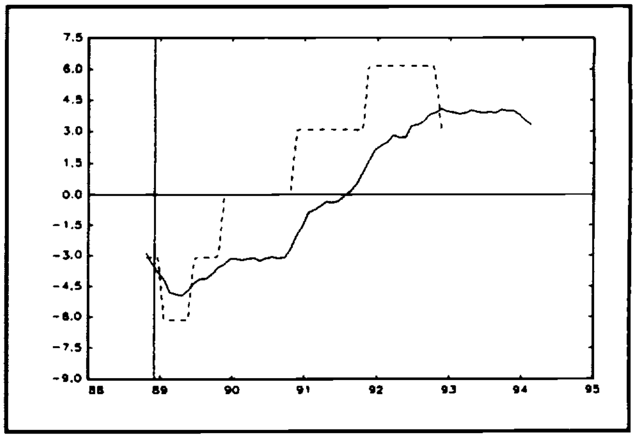
1972 - 1979



1984 - 1988

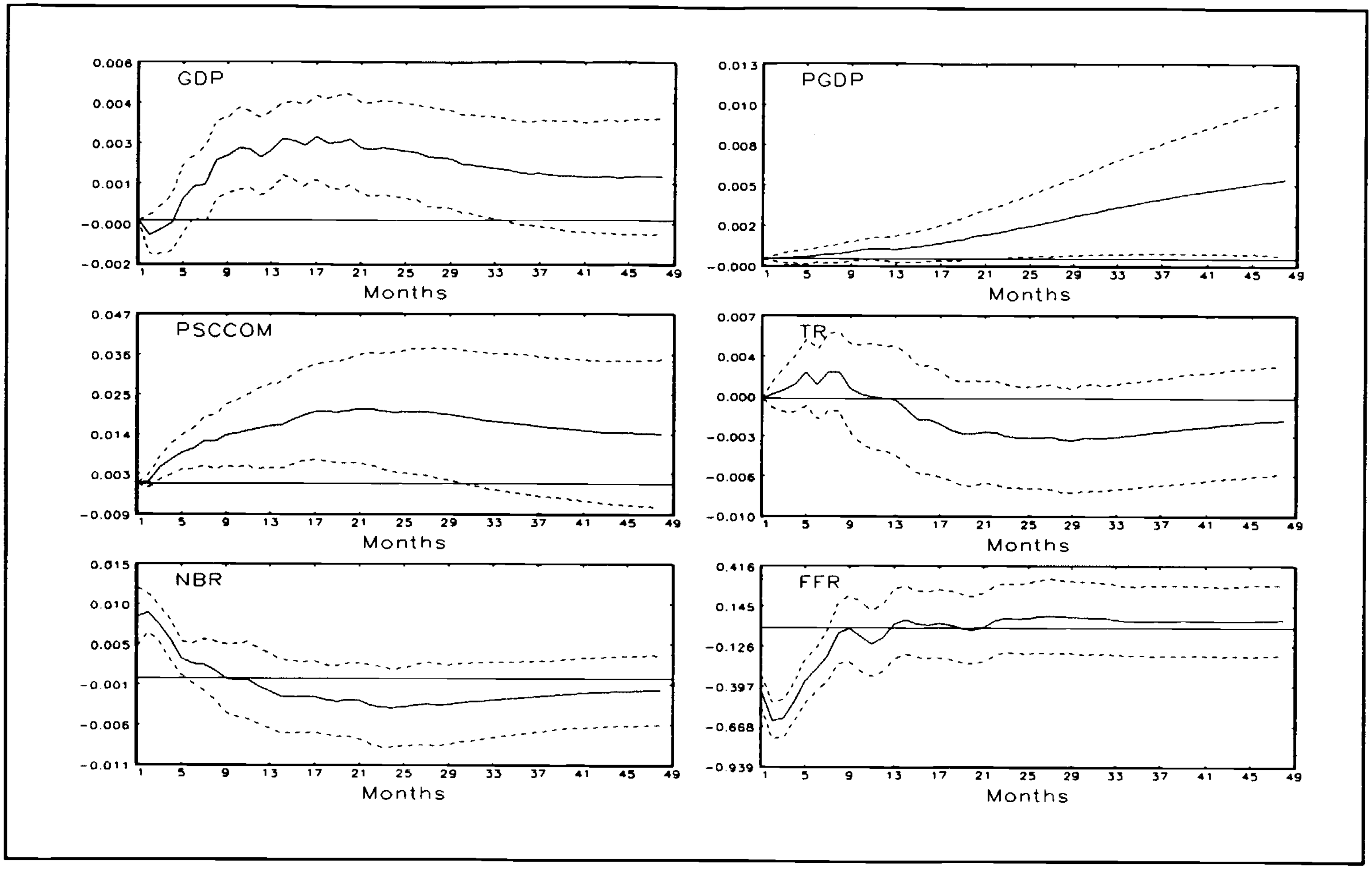


1988 - 1994



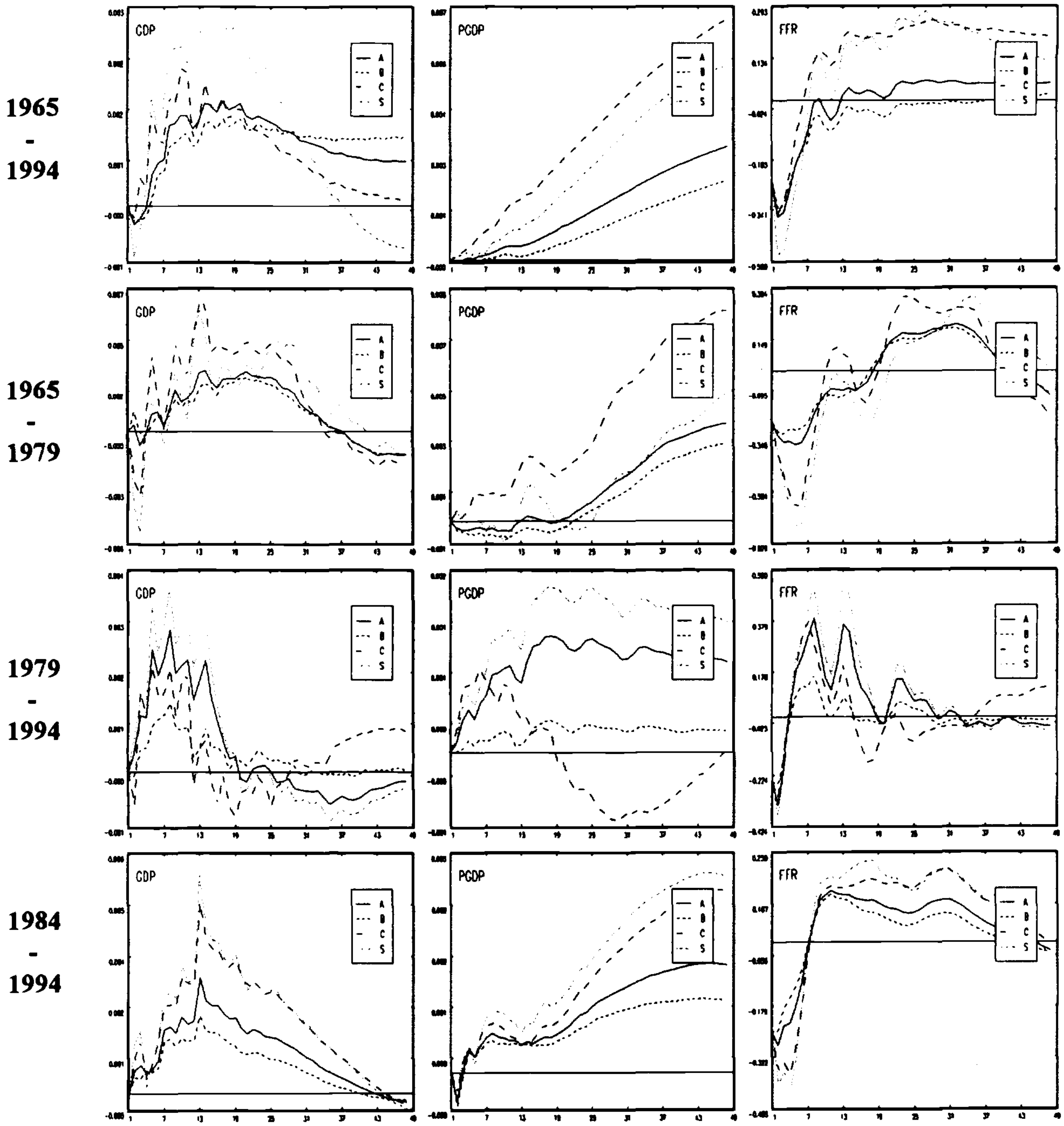
Note: The solid line is the total measure of monetary policy stance rescaled to have the same variance as the federal funds rate and zero mean. The dashed line is the Boschen-Mills index and the vertical lines are the Romer dates.

Figure 8: Responses to a Monetary Policy Shock, Monthly, 1965 - 1994



Note: The impulse responses are based on estimates from a 6-variable monthly VAR which includes interpolated GDP (GDPM), interpolated GDP deflator (PGDPM), commodity price index (PCOM), total reserves (TR), nonborrowed reserves plus extended credit (NBR), and federal funds rate (FFR). GDPM, PGDPM, and PCOM are in logs, TR and NBR are normalized by 36-month MA of TR. The shock is constructed under the identification of model A. The dashed lines are two standard error bands.

Figure 9: Responses of Real GDP, the GDP Deflator, and the Federal Funds Rate to Alternative Monetary Policy Shocks



Notes: The impulse responses are constructed from estimates which come from 6-variable monthly VARs which include interpolated real GDP (GDPM), interpolated GDP deflator (PGDPM), commodity price index (PCOM), total reserves (TR), nonborrowed reserves plus extended credit (NBR), and federal funds rate (FFR). GDPM, PGDPM, and PCOM are in logs, TR and NBR are normalized by 36-month MA of TR. The first three VARs are estimated with 12 lags while the last VAR uses 6 non-consecutive lags (1, 2, 3, 6, 9, 12). The impulse responses are constructed at 48 months horizon.