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Source: *Southern Economic Journal*, Vol. 67, No. 3 (Jan., 2001), pp. 618-636
Published by: Southern Economic Association
Stable URL: http://www.jstor.org/stable/1061454

*Southern Economic Journal* 2001, 67(3), 618–636

# The Effects of Monetary Policy Shocks: Comparing Contemporaneous versus Long-Run Identifying Restrictions

W. Douglas McMillin*

This study compares the effects of monetary policy shocks on the macroeconomy using four different procedures for identifying policy shocks that use contemporaneous restrictions and a procedure that uses long-run restrictions. Impulse response functions are computed using the same vector autoregressive (VAR) model and sample period. The comparison is done for a model that includes only a short-term interest rate and for a model that adds a long-term rate as well. Sources of differences in the magnitude of effects across identification schemes are examined.

## 1. Introduction

Vector autoregressive (VAR) models have been widely used in recent years to analyze the effects of monetary policy shocks. However, estimates of the macroeconomic effects of monetary policy often differ across studies with regard to both timing and magnitude. The studies generating these estimates frequently differ in terms of the variables constituting the model, the sample period for estimation, and the method of identifying policy shocks (see, e.g., Christiano, Eichenbaum, and Evans 1994, 1996, 1998; Gordon and Leeper 1994; Lastrapes and Selgin 1995; Pagan and Robertson 1995, 1998; Leeper, Sims, and Zha 1996).

Certainly, a critical element in the estimation of the effects of policy shocks is the identification of these policy shocks, that is, the determination of exogenous shocks to monetary policy. Two methods have been widely used in the VAR literature to identify structural shocks to monetary policy. One general approach employs restrictions on the contemporaneous relations among the variables of the VAR model, while the second general approach imposes restrictions on the long-run relations among the variables. Although economic and institutional arguments can be used to rationalize each identification scheme, there is no consensus as to which approach to identifying shocks is preferred, and the weaknesses of both approaches have been discussed in the literature.[1] Keating (1992), Lastrapes and Selgin (1995), and McCarthy (1995) consider limitations of the use of contemporaneous identifying restrictions. Faust and Leeper (1997) discuss potential drawbacks of imposing long-run restrictions.

The aim of this study is to examine the implications of contemporaneous versus long-run

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The author thanks two anonymous referees, James Fackler, Bill Lastrapes, and Prosper Raynold for helpful comments and suggestions.

Received May 1998; accepted April 2000.

[1] Lastrapes (1998) suggests a Bayesian approach to dealing with uncertainty about the appropriate identification scheme. The framework for his analysis is the Gordon and Leeper (1994) model.

identification schemes for estimating the effects of monetary policy shocks within the VAR model used by Christiano, Eichenbaum, and Evans (1994, 1996, 1998; hereafter CEE) and Bernanke and Mihov (1998; hereafter BM) over a particular sample period. Holding constant the variables in the VAR model and the sample period allows one to clearly observe the effect of the identification scheme in estimating the timing and magnitude of the effects of monetary policy actions. The model employed comprises output, the price level, commodity prices, and three reserves market variables: total reserves, nonborrowed reserves, and the federal funds rate. The focus on the reserves market is important since it allows a more thorough consideration of how policy actions are implemented than does a model that includes only a reserve aggregate or the federal funds rate as the policy variable. Following BM (1998), monthly data are used in estimating the model; use of monthly data reduces problems that may arise with temporal aggregation (see Christiano and Eichenbaum 1987). The effects of monetary policy shocks for different identification schemes are evaluated by computing impulse response functions.

The approach in this paper is similar in spirit to Keating (1992) and Lastrapes (1998). However, these studies focused on the effects of *money supply* shocks, while the focus of the current study is on *monetary policy* shocks. It is generally thought that, since money supply shocks typically confound policy actions and nonpolicy events, they are not a good measure of monetary policy shocks. For example, consider a textbook model of the money supply process in which the money supply equals the product of a money multiplier and a reserve aggregate like nonborrowed reserves. The money multiplier is affected by portfolio decisions of the nonbank public as reflected in changes in the currency/checkable deposit ratio and, depending on the definition of money considered and whether reserves are imposed on time deposits, the time deposit/checkable deposit ratio. The money multiplier is also affected by bank behavior as embodied in the ratio of excess reserves to checkable deposits, by reserve requirements set by the central bank, and in some formulations by the discount rate set by the central bank. A change in either the money multiplier or the reserve aggregate will alter the money supply, and, since changes in the money multiplier and reserve aggregates frequently occur in the same period, changes in the money supply will often reflect the behavior of the central bank, banks, and the nonbank public. Fackler and McMillin (1998) demonstrated the importance of separating money supply shocks into reserve aggregate shocks and money multiplier shocks within the context of a VAR model that used long-run restrictions to identify structural shocks to the money multiplier, a reserve aggregate, and money demand, as well as structural shocks to aggregate supply and the IS curve. They found differences in the timing and magnitude of the effects of the money multiplier and reserve aggregate shocks on macro variables. This suggests that considering just money supply shocks may yield a distorted picture of the effects of monetary policy actions.

Although BM (1998) and CEE (1998) compared the effects of alternative monetary policy shocks identified using contemporaneous restrictions within a common model and sample period, no comparison was made with monetary policy shocks identified using long-run restrictions. In their study of alternative approaches to estimating the liquidity effect, Pagan and Robertson (1995) explicitly considered the CEE model, but, within this specific framework, they did not consider the Strongin, Bernanke-Mihov, Bernanke-Blinder, or long-run restrictions identification schemes. They impose CEE-type and Strongin-type restrictions within other models that comprise a subset of the CEE model variables but do not consider long-run restrictions schemes or the Bernanke-Blinder or Bernanke-Mihov schemes within these models. They also compare estimates of the impact liquidity effect for money supply shocks within a four-variable model that includes money, price, output, and an interest rate for a long-run restrictions scheme, a scheme that blends long-run and contemporaneous restrictions, and a scheme that uses only contemporaneous identifying restrictions.

Pagan and Robertson (1998) compared estimates of the liquidity effect of a shock to a reserve or monetary aggregate within three different VAR models. One model used only contemporaneous restrictions to identify a shock to total reserves; one model used only long-run restrictions to identify a shock to either the monetary base, M1, or M2; and the third used a blend of contemporaneous and long-run restrictions to identify a money supply (M1) shock. The variables in each model differ, and the same sample period is not used for all models.

Although these previous studies have provided valuable information about estimating the macro effects of either the money supply or monetary policy, it seems important to compare the effects of contemporaneous versus long-run restrictions within a model that contains the major reserve market variables over a common sample period, something not done in previous studies. Section 2 of the paper discusses the model and the alternative identification schemes in more detail. Section 3 presents the impulse response functions, while section 4 provides a brief summary and conclusion.

## 2. Model Specification and Identification of Monetary Policy Shocks

As noted earlier, the model consists of output, the price level, a commodity price index, total reserves, nonborrowed reserves, and the federal funds rate. The commodity price index is included in light of the "price puzzle" often generated in VAR models that do not include a variable that proxies for information about future inflation. The reserves market variables are the ones generally considered critical in specifying a model of this market.

The model is estimated using monthly data for the period 1962:1–1996:12. Data from 1962:1–1964:12 are used as presample data, and estimation is done for 1965:1–1996:12. The three-year gap between the beginning of the data and the start of the estimation period is necessitated by the manner in which the reserve variables are constructed. Following CEE (1994), a lag of 12 months is used in all VAR models. All data are from the DRI Basic Economics database, and the database name is enclosed in parentheses after the variable description. Following BM (1998), output is measured by the log of real GDP (*gdpq* [chain-weighted real GDP]) interpolated from quarterly data).[2] The price level is measured by the log of the interpolated chain-weighted price index for GDP (*gdpdfc*). The commodity price index is the log of the Commodity Research Bureau's spot market price index for all commodities (*psccom*).

Total reserves (*fmrra*) are adjusted for reserve requirement changes, as are nonborrowed

[2] The interpolation of real GDP and the chain-weighted price index for GDP is done using the distrib.src procedure in RATS. The random walk option is selected in this procedure. The distrib procedure ensures that the average of the three months' interpolated data for a quarter equals the quarterly figure.

To check robustness of results, the commonly used industrial production index is also considered along with the index of coincident indicators. Walsh and Wilcox (1995) argue that the index of coincident indicators is a more comprehensive and hence better measure of aggregate output than is industrial production alone since the index of coincident indicators is a weighted average of industrial production, nonagricultural employment, real income minus transfers, and real manufacturing and trade sales. When these alternative output measures were used, the log of the personal consumption deflator was used as the price variable. Since the results for industrial production and the index of coincident indicators were very similar to those in Figure 1, all subsequent analysis was done using real GDP.

reserves (*fmrnbc*). The nonborrowed reserves measure includes extended credit; the series with only nonborrowed reserves exhibits a sharp drop at the time of the Continental Illinois crisis in 1984. Following BM (1998), both total reserves and nonborrowed reserves are normalized by a 36-month moving average of total reserves. They do this rather than take logs since they employ a linear model of the reserves market in their identification scheme. Since the BM scheme is considered in this paper, their method of constructing the reserves variables is used. The level of the federal funds rate (*fyff*) is employed.

As noted earlier, this study focuses on the implications of using contemporaneous restrictions versus long-run restrictions to identify monetary policy shocks for the estimation of the effects of monetary policy on the macroeconomy. Four alternatives using contemporaneous restrictions are employed. Three rely solely on the Choleski decomposition, while the other uses the Choleski decomposition in conjunction with the estimation of a structural model of the reserves market.

The first identification scheme was suggested by CEE (1994, 1996) and employs the following order for the decomposition: output ($y$), price level ($p$), commodity price ($cp$), nonborrowed reserves ($nbr$), federal funds rate ($ffr$), and total reserves ($tr$). The $nbr$ are taken as the policy variable. Since all contemporaneous correlation between two variables is attributed to the variable higher in the ordering with the Choleski decomposition, this scheme implies that monetary policy actions affect $y$, $p$, and $cp$ only with a lag. It also implies that the Federal Reserve responds to contemporaneous movements in these three variables; that is, the Federal Reserve's reaction function includes the contemporaneous values of these three variables as well as lagged values of these variables and lagged values of $nbr$, $tr$, and $ffr$. The assumption that monetary policy affects $y$ and $p$ only with a lag and that it has a contemporaneous effect on a short-term market interest rate is uncontroversial; however, the assumption that monetary policy affects an auction market variable like $cp$ only with a lag has been questioned (McCarthy 1995). McCarthy (1995) and Rudebusch (1998) have also criticized the assumption that the Federal Reserve responds to the current period values of $y$ and $p$. They point out that the Fed is likely to have only noisy preliminary information about the current period values of these variables. Depending on the nature of the revision to the preliminary estimates, the use of the current period value of revised data for $y$ and $p$ may have important effects on the estimates of the structural monetary policy shocks and impulse response functions, although Sims (1998) questions the quantitative importance of this criticism. Thus, this method of identifying monetary policy shocks has some unappealing as well as appealing features.

The second identification scheme involving contemporaneous restrictions is in the spirit of Strongin (1995). It employs the Choleski decomposition with the ordering $y$, $p$, $cp$, $tr$, $nbr$, $ffr$. Strongin argues that shocks to $nbr$ are mixtures of reserve demand shocks and policy shocks. He contends that under the policy procedure followed over the sample used here, the level of $tr$ was determined primarily by Fed accommodation of the demand for reserves. Thus, in this view, shocks to $tr$ reflect reserve demand shocks, and ordering $tr$ before $nbr$ purges $nbr$ shocks of reserve demand effects. The contemporaneous causal link between $nbr$ and $tr$ is the reverse in the Strongin identification approach (hereafter STR) of what it was in the CEE approach. The critique of the CEE scheme carries over to STR as well.

The third procedure considered that uses contemporaneous restrictions is that of BM (1998). This procedure blends the Choleski decomposition with the estimation of a small structural model of the reserves market. The estimation of the reserves market model is done with VAR residuals for $nbr$, $tr$, and $ffr$ that are orthogonalized with respect to $y$, $p$, and $cp$. Thus, as in CEE and STR, it is assumed that monetary policy actions affect the macro variables only with a lag and that policy makers respond to contemporaneous movements in $y$, $p$, and $cp$.

The structural model has the following specification:

$$e_{TR} = -\alpha e_{FFR} + \mu^d \qquad \text{(total reserve demand)}$$

$$e_{BR} = \beta e_{FFR} + \mu^b \qquad \text{(borrowed reserve demand)}$$

$$e_{NBR} = \phi^d \mu^d + \phi^b \mu^b + \mu^s \quad \text{(Federal Reserve reaction function)},$$

where the $e$'s represent the VAR residuals from the $tr$, $nbr$, and $ffr$ equations orthogonalized with respect to $y$, $p$, and $cp$ and the $\mu$'s are structural shocks with $\mu^s(\mu^d)(\mu^b)$ representing the structural shock to monetary policy (total reserve demand) (borrowed reserve demand). Equilibrium is defined by equality between $tr$ demand and $tr$ supply. Conceptually, $tr$ demand is assumed to depend negatively on $ffr$, while borrowed reserve demand is assumed to depend positively on $ffr$.[3] The Fed is assumed to react to contemporaneous shocks to both $tr$ demand and borrowed reserve demand in determining the supply of $nbr$. As in CEE and STR, structural shocks to $nbr$ are the measure of monetary policy shocks. Furthermore, BM note that a just-identified version of their model with $\alpha = 0$ performs well. Consequently, this assumption is employed in this paper as well; with $\alpha = 0$, shocks to $tr$ (orthogonalized with respect to $y$, $p$, and $cp$) are assumed to be shocks to $tr$ demand, as in STR. The model is estimated with a two-step GMM procedure; specifically, a RATS procedure (measure.src provided by BM) is used to estimate the reserves market model and obtain $\mu^s$.

Again, the critique of the CEE identification scheme with regard to $cp$ and current period knowledge of $y$ and $p$ is applicable to the BM procedure. CEE (1998) present an additional criticism of BM based on BM's assumption that there is no contemporaneous effect of $nbr$ on borrowed reserves. Although they allow shocks to borrowed reserve demand to affect $nbr$, it is assumed by BM that $nbr$ have no contemporaneous effects on borrowed reserves. CEE argue, using as an example Goodfriend's (1983) model of borrowed reserves, that theory suggests an effect of $nbr$ on borrowed reserves, and they present empirical evidence that $nbr$ affects borrowed reserves contemporaneously.

Following Bernanke and Blinder (1992; hereafter BB), the fourth scheme assumes that $ffr$ is the policy variable. A Choleski decomposition with the ordering $y, p, cp, ffr, nbr, tr$ is used. As before, it is assumed that monetary policy actions have only a lagged effect on $y$, $p$, and $cp$ and that the Fed responds to current period movements in these variables.

The final method of identifying monetary policy shocks examined imposes restrictions on the long-run relations among the variables in the model. No restrictions are placed on the contemporaneous relations among the variables. This procedure (hereafter referred to as LR) was introduced by Blanchard and Quah (1989) and Shapiro and Watson (1988) to identify shocks to aggregate demand and supply and has been used recently by Lastrapes and Selgin (1995) to identify money supply shocks and by Fackler and McMillin (1998) to identify monetary policy shocks.[4]

[3] BM (1998) specify borrowed reserve demand to depend on the gap between $ffr$ and the discount rate, but in most of their empirical work they make the simplifying assumption that discount rate shocks are zero. This is consistent with the studies of CEE and Strongin, who do not explicitly consider the discount rate.

[4] The model used in Fackler and McMillin (1998) is a good bit different from the CEE- and BM-type model used in this paper. However, the basic patterns of effects of a monetary policy shock on $y$ and $p$ are similar to those reported in this paper.

The key restrictions used to identify monetary policy shocks in this approach are neutrality restrictions. Prior to implementing this procedure, the model is transformed in the following way. The model is specified as comprising *y*, the log of real commodity prices ($cp - p$), *cp*, *nbr*, *tr*, and *ffr*. We note that *p* no longer enters as a separate variable, but the effect of monetary policy on *p* can be determined in a straightforward way from the separate effects of monetary policy on the relative price of commodities and on *cp*. The *nbr* are assumed to be the monetary policy variable.[5] All variables are first differenced prior to estimation; that is, a unit root is imposed. With the model in first differences, a Choleski decomposition of the long-run relations allows one to easily impose neutrality restrictions. With the model in first differences, the moving average representation indicates the effect of shocks to the variables on the changes in the variables. The effect on the level of a variable at a particular point is the cumulative effect of the changes up to and including that point. The long-run effect of a shock on the level of a variable is simply the cumulative sum of the relevant part of the entire moving average representation. Since the Choleski decomposition attributes all of the correlation between two variables to the one higher in the ordering, one can impose neutrality restrictions by placing real variables prior to the monetary policy variable in a Choleski decomposition of the long-run relations among the variables. This is demonstrated in Keating (1999).

The first restriction used to identify the monetary policy shock is that shocks to monetary policy have no long-run effects on *y*. A second restriction is that shocks to monetary policy have no long-run effects on ($cp - p$), and a third is that monetary policy shocks have no long-run effects on the interest rate. The first and third restrictions are familar results from a sticky-wage/price aggregate demand–aggregate supply-type model with IS-LM underlying aggregate demand. A positive shock to *nbr* initially raises real money balances, shifting the LM curve and the aggregate demand curves right and raising *y* above the natural level. The interest rate falls initially. However, as *p* adjusts and *y* returns toward its initial level, real balances begin to fall, and the interest rate begins to return to its initial level. In long-run equilibrium, real balances are back at their initial level, as are *y* and the interest rate; *p* is permanently higher.

No restrictions are placed on the long-run effects of monetary policy shocks on *tr*, *cp*, or *p*. As noted earlier, the structural shock to monetary policy can be identified by a Choleski decomposition of the long-run relations among the variables, with *y* ordered first, the relative price of commodities ordered second, *ffr* ordered third, *nbr* ordered fourth, *tr* ordered fifth, and *cp* ordered last.[6] Since *y*, ($cp - p$), and *ffr* precede *nbr* in the ordering, it is assumed that shocks to these variables can influence *nbr* and hence monetary policy in the long run. Placing *tr* and *cp* after *nbr* allows monetary policy to have long-run effects on these variables but also assumes that shocks to these variables have no long-run effects on *nbr*. If one interprets *tr* shocks as shocks to *tr* demand, then ordering *tr* after *nbr* implies that the Fed does not accommodate shocks to *tr* demand in the long run, even though it may well do so in the short and intermediate runs. An alternative ordering with *tr* preceding *nbr* has the unappealing implication

[5] With *ffr* as the monetary policy variable, applying long-run restrictions to identify the policy shock implies that the central bank can set the *level* of *ffr* at any desired value in the long-run. This assumption is more questionable than is the analogous assumption that the central bank can set *nbr* at a desired level in the long-run when *nbr* is the monetary policy variable.

[6] Since the focus of this paper is on monetary policy shocks, what is critical to the identification of monetary policy shocks is that *nbr* is ordered after *y*, ($cp - p$), and *ffr* and before *tr* and *cp*. Within the block of variables before *nbr*, the relative ordering is not critical for estimating the effects of a shock to *nbr*; the same is true for the block following *nbr*.

that permanent shocks to *nbr* have no long-run effects on *tr*. Finally, the assumption that shocks to *cp* have no long-run effect on *nbr* in conjunction with a long-run effect of $(cp - p)$ on *nbr* implies that shocks to *p* can have long-run effects on the monetary policy variable. Ordering *cp* after *nbr* is consistent with the view that the Fed looks at *cp* as an indicator of future movements in *p*, which is the price variable of ultimate interest to the Fed, and not as a variable of fundamental concern to the Fed. Other interpretations of the ordering are, no doubt, possible. For a discussion of the conditions under which long-run recursive structures like that employed here identify structural shocks, see Keating (1999).

One advantage of the use of LR is that no restrictions are placed on the contemporaneous relations among the variables. Thus, a restriction that monetary policy shocks have no contemporaneous effects on *cp* is not imposed, as was done in the schemes previously considered. However, Faust and Leeper (1997) note the problematic nature of imposing infinite horizon restrictions in a VAR estimated with data from a finite sample. They argue that the estimate of the long-run effect is uncertain and that uncertainty about the long-run effect is transmitted to impulse response functions since long-run restrictions are used to identify structural shocks. It is apparent that each approach to identifying monetary policy shocks has its weaknesses, and no consensus on the best approach has emerged. Consequently, it is of interest to compare the effects of monetary policy shocks identified using contemporaneous and long-run restrictions, holding constant the model variables, lag length, and sample period.

## 3. Empirical Results

### *Impulse Response Functions*

The effects of monetary policy shocks are evaluated by computing impulse response functions (IRFs). The IRFs present the effects of a one-standard-deviation shock to the monetary policy variable and represent the "average" effect of a monetary policy shock over the sample period. The IRFs for *y*, *p*, and *ffr* are presented in Figure 1. The first column of this figure presents the effects of a shock identified using the CEE procedure. The remaining columns present analogous results for the STR, BM, BB, and LR restrictions approaches, respectively. In each diagram, the solid line is the point estimate, and the dotted lines represent a one-standard-deviation band around the point estimate. The confidence bands are derived from Monte Carlo simulations with 1000 draws. We note that the point estimates of the effects of a monetary policy shock vary somewhat in terms of magnitude, timing, and persistence, although the general pattern is similar for each variable. For *y*, we observe a hump-shaped pattern, with *y* eventually returning essentially to its initial value for the CEE, STR, BM, and LR approaches. The BB model indicates a very persistent positive effect even after 48 months. All identification schemes indicate a permanent effect of monetary policy on *p*. A liquidity effect is present in all cases. The *ffr* falls initially but rebounds close to its initial value within a year and remains at the initial value thereafter for the BM, BB, and LR procedures. For the STR procedure, the lower bound of the confidence interval is close to zero and eventually includes zero. The pattern

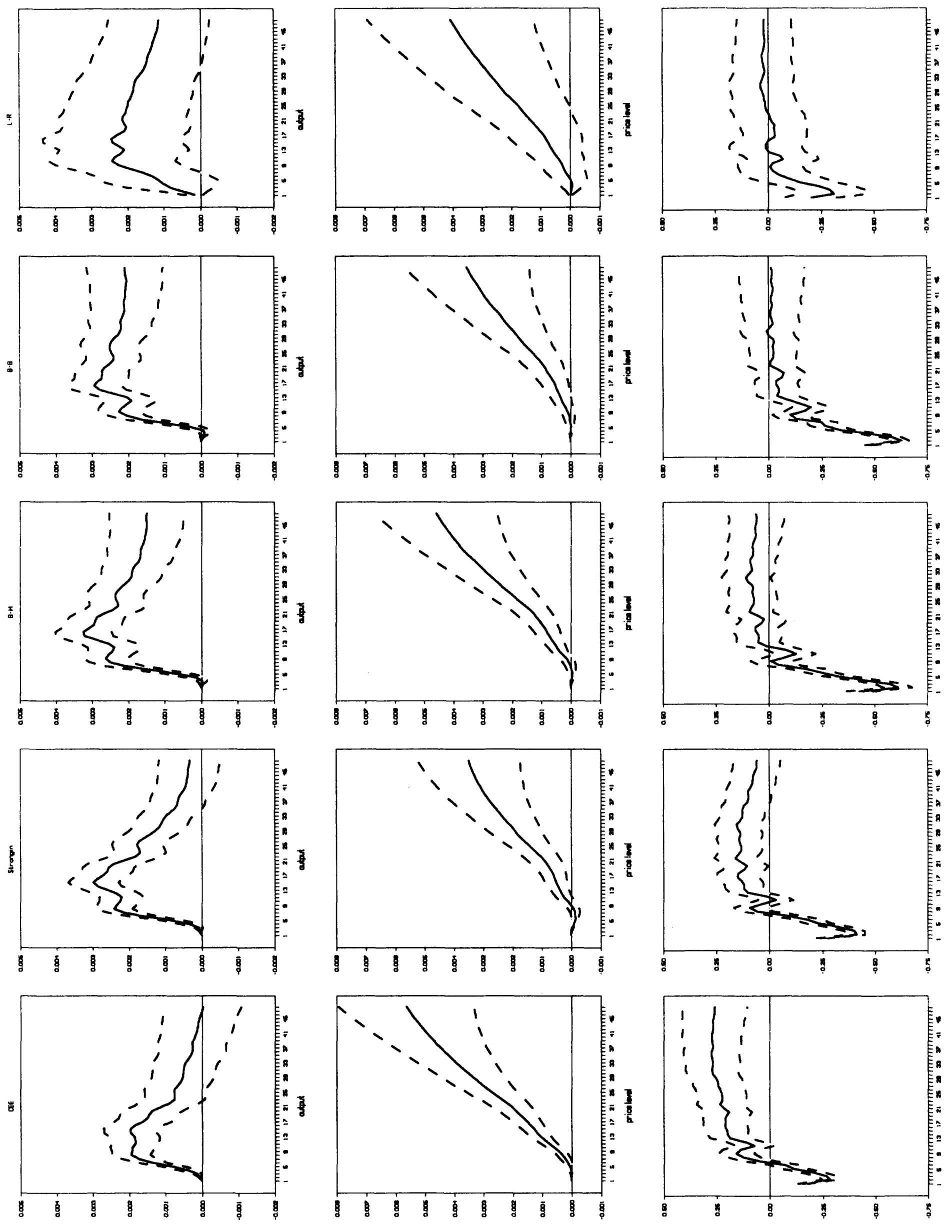


**Figure 1.** Shock to Monetary Policy (Basic Model)

for the CEE identification is troublesome, however. After about eight months, the confidence interval for *ffr* lies above zero for the remainder of the horizon reported.[7]

Although the general pattern of effects is similar across identification schemes, the magnitudes of the point estimates differ across schemes. Consequently, it is useful to determine whether these differences are substantial. This is done by first assuming that the LR approach is the appropriate way to identify shocks. The confidence bands for the LR approach are then plotted along with the point estimates from the other approaches. This provides information on whether the differences in magnitudes across schemes are substantial in the sense that the point estimates lie outside the confidence bands. Next it is assumed that a particular contemporaneous identifying restriction is appropriate. Confidence bands for this scheme are plotted along with the point estimates from the other schemes. This procedure could be repeated using the confidence bands from the other contemporaneous restrictions identification schemes, but doing this provides essentially no additional information. Consequently, the confidence bands for the BM procedure are plotted along with the point estimates of the other schemes.

Figure 2 plots the confidence bounds for the LR approach and the point estimates for the CEE, STR, BM, and BB approaches. For *y* and *p,* the point estimates of the approaches using contemporaneous restrictions lie within the LR confidence bands. However, for *ffr,* we observe that the point estimate for the CEE identification lies above the upper bounds of the confidence interval at horizons greater than a year. The point estimate for the STR identification essentially lies within the confidence bound, while the point estimates for the BM and BB identifications lie below the lower bound for the first six months and within the bounds thereafter.

Figure 3 plots the confidence bounds for the BM procedure and the point estimates for the other procedures. In the case of *y,* we observe that for approximately 12 months, the point estimates from the CEE and STR procedures lie within the confidence bounds. The point estimates for CEE drop below the lower bound after approximately 15 months and remain below the lower bound after that. The point estimates for the STR procedure lie within the confidence bounds until approximately 32 months, when they drop slightly below the lower bounds, while the BB point estimates lie within or on the confidence intervals at all horizons. The point estimates for the LR approach lie above the upper bound for the first six months but are within the bounds thereafter.

For *p,* the point estimates essentially lie within the confidence bounds, although there are some slight deviations above the upper bound for part of the horizon for the CEE identification procedure. There are some substantial differences for *ffr,* however. We observe that the point estimate for the CEE scheme lies entirely above the upper bound of the confidence interval. The point estimate for the STR procedure lies above the upper bound for approximately six months and is then close in value to the upper bound until about 13 months, when it falls entirely within the bounds. The point estimate for the BB scheme lies on or within the bounds at all horizons. The point estimate for the LR scheme lies above the upper bound for about

[7] Figure 1 presents results only for *y, p,* and *ffr* since these variables have been the focus of attention in the literature estimating the effects of monetary policy shocks. The effects on the *cp, nbr,* and *tr* will be described briefly and figures are available on request. For all identification schemes, there is a long-lived positive effect on the commodity price level. For the procedures using contemporaneous restrictions, there are transitory positive effects on both *tr* and *nbr,* with the level of these variables returning to the initial value in the long run. For LR, a monetary policy shock has only a transitory effect on the change in *nbr* and *tr* but has a permanent positive effect on the level of both *tr* and *nbr.* This is not surprising since this identification scheme used restrictions on the long-run effects of policy shocks, and long-run effects on these variables were explicitly allowed for in the identification procedure.

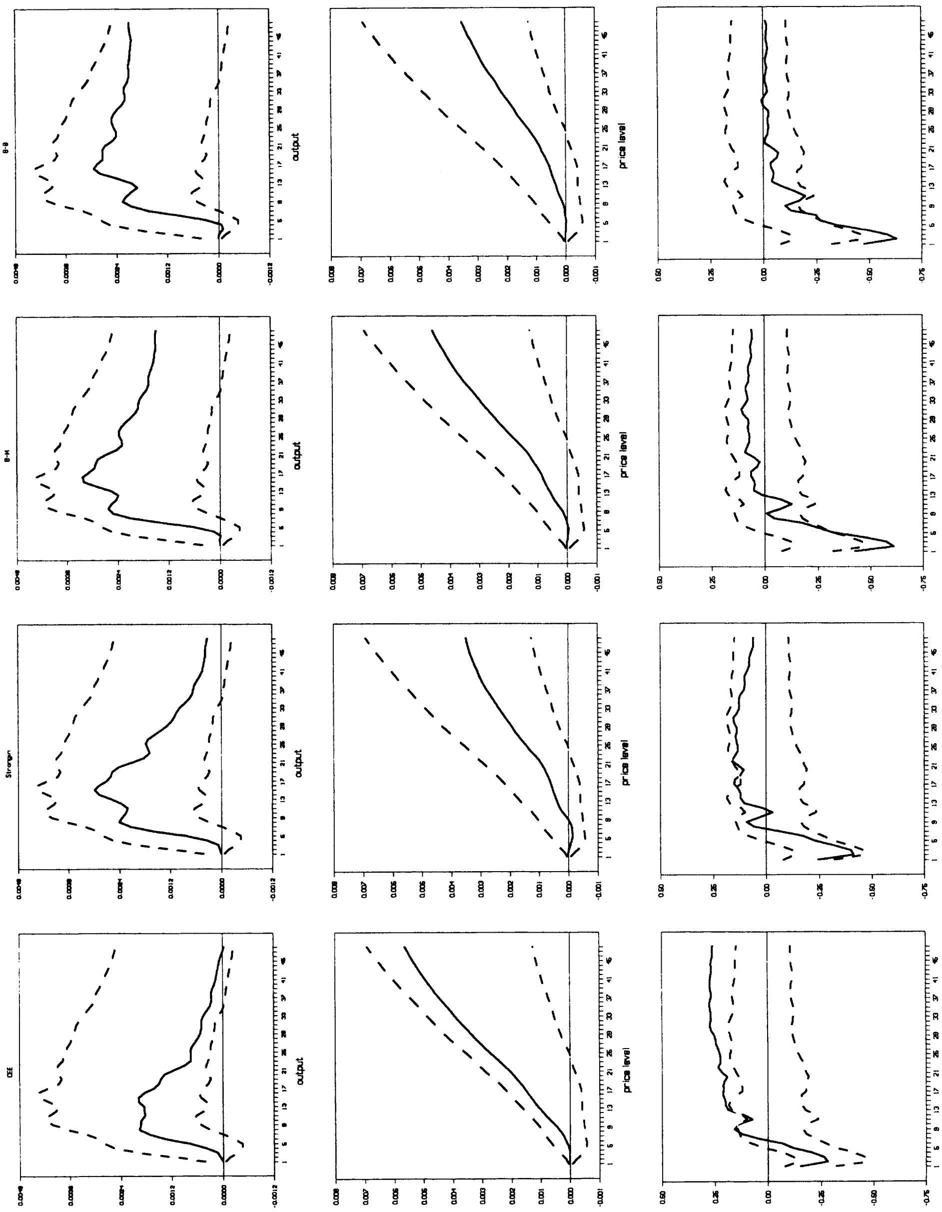


**Figure 2.** Long-Run Restrictions Confidence Intervals Point Estimates from Other Identification Procedures

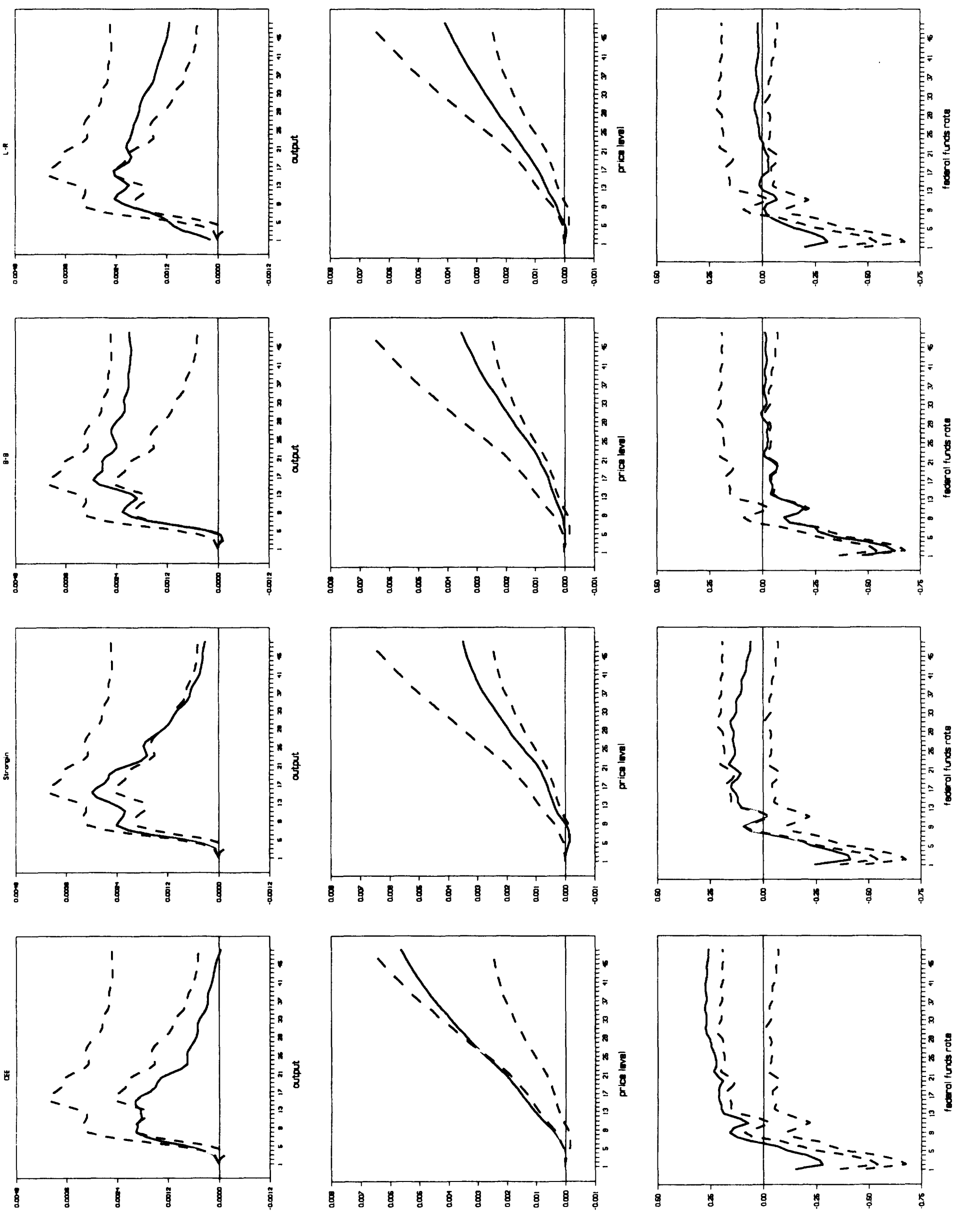


**Figure 3.** Bernanke-Mihov Restrictions Confidence Intervals Point Estimates from Other Identification Procedures

eight months but then lies within the bounds thereafter. Clearly, the BM and BB identification procedures indicate a stronger liquidity effect than do the other identification schemes.

### *Why Do the Magnitudes of the IRFs Differ?*

Figure 3 suggests that the CEE procedure generates results for $y$ and $ffr$ that differ substantially from the other contemporaneous identification schemes. It is useful to explore why this occurs. Consider the following structural model:

$$y_t = A_0 y_t + A_1 y_{t-1} + \cdots + A_q y_{t-q} + \mu_t$$

where $y_t$ = vector of model variables, $A_0$ = coefficient matrix of contemporaneous effects, $A_i$, $i = 1, \ldots, q$ = coefficient matrices for lagged effects of $y$, $q$ = maximum lag, and $\mu_t$ = vector of structural shocks (which are assumed to be uncorrelated) with variance-covariance matrix $\Omega$. Solving for $y_t$, we obtain

$$y_t = B_1 y_{t-1} + \cdots + B_q y_{t-q} + e_t$$

where $B_i = (I - A_0)^{-1}A_i$ and $e_t = (I - A_0)^{-1}\mu_t$. The moving average representation is

$$y_t = (I - B_1 L - \cdots - B_q L^q)^{-1} e_t \quad \text{or} \quad y_t = C(L)e_t$$

where $C(L) = (I - B_1 L - \cdots - B_q L^q)^{-1}$. In terms of structural shocks, we have $y_t = C(L)(I - A_0)^{-1}\mu_t$. The term $C(L)$ is identical for all the contemporaneous identification schemes employed in this paper (except, of course, the order of the variables differs). It is different for the LR scheme since the model variables are transformed when this scheme is employed. Of course, both $(I - A_0)^{-1}$ and $\mu_t$ differ across identification schemes and is the only source of difference in the IRFs for the contemporaneous schemes.

Part A of Table 1 presents the correlation coefficients for the structural monetary policy shocks ($\mu_t$) generated by the different schemes. We note that the correlations are high between some policy shock measures (e.g., CEE-STR and BM-BB) and low between others (e.g., CEE-BB, LR-STR, and LR-BB). This has been noted by Rudebusch (1998) in his critique of VAR measures of policy shocks. Sims (1998) argues that as long as appropriate instruments are used to identify monetary policy shocks, qualitatively similar effects on macro variables may be obtained, even though the monetary policy shocks themselves may not be highly correlated across schemes.[8]

To obtain the effects of one-standard-deviation shocks, one can replace $\mu_t$ by $\Omega^{1/2}$, where $\Omega^{1/2}$ is a diagonal matrix of standard deviations of the structural shocks. The term $(I - A_0)^{-1}\Omega^{1/2}$ is generated by the Choleski decomposition of the variance-covariance matrix for the CEE, STR, and BB schemes. For the BM scheme, $(I - A_0)^{-1}\Omega^{1/2}$ is a "hybrid" matrix in which the GMM estimates of the reserves market structural parameters and variances replace the relevant elements of a regular Choleski decomposition. In the case of the LR scheme, $(I - A_0)^{-1}\Omega^{1/2}$ is a transformation of the Choleski decomposition based on the long-run restrictions for the LR scheme. These results are demonstrated in an appendix available on request.

Part B of Table 1 presents the contemporaneous effects of a monetary policy shock in each of the identification schemes. This is the column of $(I - A_0)^{-1}\Omega^{1/2}$ corresponding to the monetary policy variable. The differences in the magnitudes of the effects are propagated forward through

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[8] See Evans and Kuttner (1998) for an insightful discussion of Rudebusch's critique of VARs.

**Table 1**

A. Correlations among Structural Shocks

| | CEE | STR | BM | BB | LR |
|---|---|---|---|---|---|
| CEE | 1.0 | | | | |
| STR | .82 | 1.0 | | | |
| BM | .68 | .83 | 1.0 | | |
| BB | .33 | .52 | .90 | 1.0 | |
| LR | .74 | .35 | .53 | .41 | 1.0 |

B. $(I - A_0)^{-1}\Omega^{1/2}$: Basic Model

| | Identification Scheme | | | | |
|---|---|---|---|---|---|
| Variable | CEE | STR | BM | BB | LR |
| $y$ | 0 | 0 | 0 | 0 | .00016 |
| $p$ | 0 | 0 | 0 | 0 | −.00002 |
| $cp$ | 0 | 0 | 0 | 0 | −.0023 |
| $nbr$ | .0129 | .0105 | .0087 | .0042 | .0101 |
| $tr$ | .0053 | 0 | 0 | −.0015 | .0072 |
| $ffr$ | −.1550 | −.2469 | −.4282 | −.4759 | −.2028 |

C. $(I - A_0)^{-1}\Omega^{1/2}$: Extended Model

| | Identification Scheme | | | | |
|---|---|---|---|---|---|
| Variable | CEE | STR | BM | BB | LR |
| $y$ | 0 | 0 | 0 | 0 | .00015 |
| $p$ | 0 | 0 | 0 | 0 | −.00003 |
| $cp$ | 0 | 0 | 0 | 0 | −.0020 |
| $nbr$ | .0123 | .0101 | .0078 | .0035 | .0065 |
| $tr$ | .0049 | 0 | 0 | −.0017 | .0047 |
| $ffr$ | −.1292 | −.2180 | −.4109 | −.4492 | −.0663 |
| r10 | −.0485 | −.0690 | −.0544 | −.0707 | −.1999 |

time by the moving average coefficients in $C(L)$. In the CEE scheme, a one-standard-deviation shock to the monetary policy variable, *nbr*, is 0.0129. This shock induces a contemporaneous change in *ffr* of −0.155 and in *tr* of 0.005. Since the CEE scheme assumes that monetary policy affects *y, p,* and *cp* only with a lag, the entries for these variables are 0. We see that for the STR scheme, the one-standard-deviation shock to *nbr,* 0.0105, is smaller than for CEE. The *tr* are ordered before *nbr* in this scheme, and the contemporaneous correlation between *tr* and *nbr* (0.6) is attributed to *tr,* so there is no contemporaneous effect on *tr* of a shock to *nbr.* The change in *ffr,* −0.2469, is larger than for CEE. One interpretation of these relative effects in the spirit of Strongin is that the larger structural shock to *nbr* in the CEE scheme is contaminated by shocks to *tr* demand. When one controls for *tr* demand shocks, the structural shock to *nbr* is smaller but the contemporaneous effect on *ffr* is larger (since this shock now omits positive *tr* demand shocks, which tend to raise *ffr*).

The standard deviation of structural shocks to *nbr* in the BM scheme is smaller than for CEE (two-thirds the size) or STR (80% the size). This is expected, of course, since the BM scheme purges *nbr* shocks of the effects of demand shocks to both *tr* and borrowed reserves. The contemporaneous decline in *ffr* is larger than for CEE or STR, again as expected. In the BB scheme, *ffr* is the policy variable. A one-standard-deviation shock to *ffr* is larger in absolute value than for the other schemes using contemporaneous restrictions, and the change in *nbr* is much smaller since *ffr* precedes *nbr* in the Choleski decomposition, and hence *ffr* is given credit for all contemporaneous correlation between the two variables. Surprisingly, there is a negative effect on *tr* for which there is no obvious explanation. This counterintuitive result raises some concern about the appropriateness of this identification scheme.

In the LR scheme, there are no constraints on contemporaneous effects. A one-standard-deviation shock to the monetary policy variable, *nbr,* is 0.0101, approximately the same value as for STR. The change in *ffr* is −0.2028, again approximately the same size as for the STR scheme. The *tr* rise by 0.007, a larger value than for CEE. We also note there is a small contemporaneous positive movement in *y*. The sign of the contemporaneous effect on *p* and *cp* is puzzling since one would normally expect a positive sign. However, in the case of *p,* the contemporaneous effect is essentially zero.

From Figure 3, we see that the CEE scheme generates results for *y* that are substantially below those for the other schemes. To the extent that monetary policy effects on output are transmitted through a liquidity effect, the results for *y* are explicable in terms of the much weaker liquidity effect for the CEE scheme. As seen in Table 1, the initial decline in *ffr* for the CEE scheme is less than half the decline in *ffr* for the BM and BB schemes and is only about 60% of the decline for the STR scheme. Even though $C(L)$ is the same for the four contemporaneous schemes, the effects of the smaller initial decline in *ffr* for CEE are carried forward, and the path of *ffr* is above the path of *ffr* for the other schemes. The initial effects on *ffr* in the STR scheme are weaker that in the BM or BB schemes, but after approximately a year and a half, the point estimate for *ffr* for the STR scheme begins to move away from the upper bound of the confidence interval. We note that, for the STR scheme, *y* moves toward the lower bound or is actually slightly below the lower bound after about 20 months. Thus, the two contemporaneous schemes with the weakest liquidity effects also display the weakest effects on *y*. Furthermore, when the impact liquidity effect from the CEE scheme, −0.155, is substituted for the impact effect on *ffr* in the other contemporaneous schemes, the point estimates for *y* drop below the lower bound of the BM confidence intervals.

We also note in Figure 3 that the initial effects on *ffr* for the LR scheme are weaker than for the BM scheme; they are similar in magnitude to those of the STR scheme. However, the effects on *ffr* quickly move within the confidence bounds and stay there. Even though the initial liquidity effect is weaker in the LR scheme than in BM, the initial effects on *y* are somewhat stronger. Recall that there is a positive contemporaneous effect of a monetary policy shock on *y* in the LR scheme. This apparently causes *y* to rise above the upper bounds on the BM confidence interval initially even though the liquidity effect is weaker than for BM.

Figure 3 is more suggestive of substantial differences across schemes than is Figure 2, which plots the relatively wide confidence bounds of the LR scheme. The only sustained departure from the confidence bounds in this figure is *ffr* for the CEE scheme; even though *ffr* is initially within the confidence bounds, it rises, and remains, above the upper bound after about a year. The point estimate of *y* for CEE remains within the confidence bounds at all horizons, although it drops toward the lower bound after 18 months.

## *Robustness of IRF Results*

### *Nonborrowed Reserve Targeting*

The estimates in Figures 1 to 3 assume that monetary policy was implemented in essentially the same way over the entire sample. As has been widely discussed (see Strongin 1995; BM

1998), there were several changes in operating regimes over the period considered here. Perhaps the most substantive changes were the switch in October 1979 from targeting short-term interest rates to targeting nonborrowed reserves and the return to a primary focus on short-term interest rates in October 1982. In order to deal with the possibility that inclusion of the October 1979–October 1982 period substantially affected the IRFs presented thus far, the following was done. A dummy variable that takes on the value of 1 over 1979:10–1982:10 and 0 in all other periods was created. The reserve market variables—*tr, nbr,* and *ffr*—were multiplied by this dummy variable. Lagged values (12) of these interaction dummy variables were then added to each equation of the VAR. This allows the reserves market variables to have effects that differ over the periods of focus on short-term interest rates from the period of focus on nonborrowed reserves. The VAR with the interaction dummy variables was estimated, and the coefficients on the dummy variables were then set to zero. The identification procedures were then applied and IRFs computed. To conserve space, the figure for this exercise is not presented here but is available on request. With only a few minor exceptions in the case of *ffr,* the IRFs are within the confidence bounds from the initial estimates.

### *Extension of the Model*

The basic model studied in this paper contains only one interest rate, *ffr.* At the suggestion of a referee, a long-term interest rate was added to the basic system. Since most discussions of the interest rate channel of the monetary transmission process focus on long-term interest rates as the main determinant of interest-sensitive spending, it is important to consider what happens when a long-term interest rate is added to the system. Gordon and Leeper (1994), Pagan and Robertson (1995), and Edelberg and Marshall (1996) are among the relatively few studies to consider the effect of monetary policy shocks on long-term interest rates within VAR models. In this paper, the constant maturity 10-year Treasury bond yield (DRI basic series fygt10, hereafter referred to as *r*10) is added to the model. Following Pagan and Robertson (1995) and Edelberg and Marshall (1996), *r*10 is added as the last variable in the ordering for the CEE, STR, and BB identification schemes, thereby implying that monetary policy affects *r*10 contemporaneously but does not respond to current period movements in *r*10.[9] For the BM scheme, *r*10 is assumed to respond contemporaneously to shocks to *y, p, cp,* and *tr* demand shocks, borrowed reserve demand shocks, and monetary policy shocks but is assumed to have no contemporaneous effects on the other model variables. For the LR scheme, *r*10 is ordered before *nbr* and after *ffr*; that is, the ordering is *y*, $(cp - p)$, *ffr, r*10, *nbr, tr,* and *cp.* This implies that monetary policy actions have no long-run effect on either short-term or long-term interest rates [or *y* or $(cp - p)$], but these variables can have long-run effects on *nbr*. The LR scheme does allow contemporaneous and intermediate-term effects of *nbr* on *r*10, and, of course, contemporaneous as well as long-run effects of *r*10 on *nbr* are possible in the LR scheme.

Figure 4 presents results analogous to Figure 1 for the models with *r*10. The inclusion of *r*10 has essentially no impact on the magnitude and pattern of monetary policy effects on *y, p,* and *ffr* for all schemes, with the exception of *ffr* in the LR approach. There is no current period effect on *ffr* in this scheme. All schemes indicate that *r*10 falls immediately following a monetary policy shock. For the approaches using contemporaneous restrictions, *r*10 quickly rebounds to its initial value after only a small decline. For CEE, the confidence bands lie above

[9] In contrast, Gordon and Leeper (1994) allow a contemporaneous effect of a long-term rate on *ffr.*

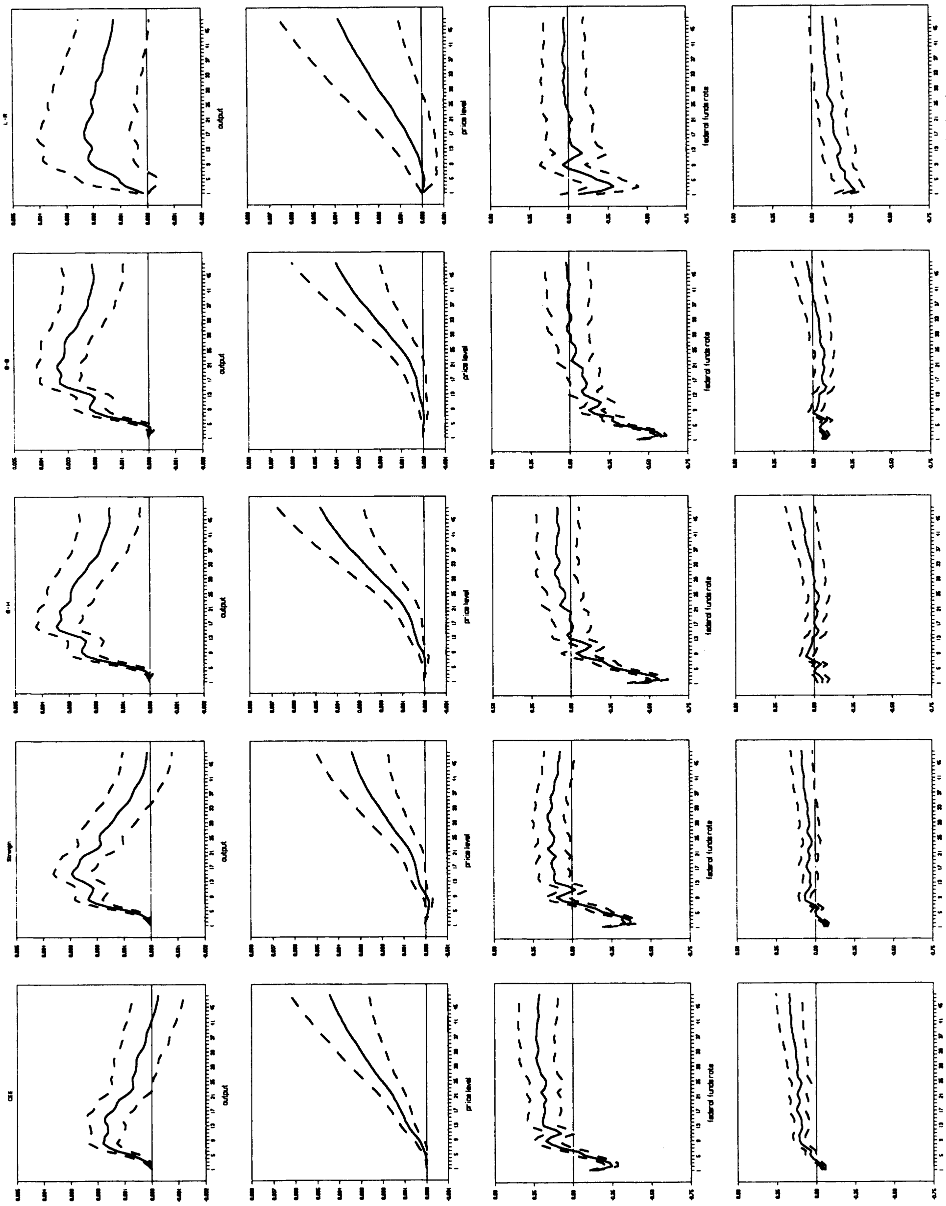

**Figure 4.** Shock to Monetary Policy (Extended Model)

zero after about six months (similar to the case for *ffr*), while the confidence bands essentially span zero after six months for the STR, BM, and BB schemes. In contrast, the LR approach suggests a very long-lived decline in *r*10 that has no obvious explanation.

Figures for the model with *r*10 analogous to Figures 2 and 3 are available on request but are not presented in order to conserve space. For *y, p,* and *ffr,* these figures are very similar to Figures 2 and 3. For *r*10, only the BB point estimate lies anywhere within the LR confidence bands, and it lies near the upper bound for months 14 to 40. When the BM confidence intervals for *r*10 are plotted, the STR and BB point estimates lie within or on the confidence bands, while the CEE point estimate lies above the upper bound and the LR point estimate lies below the lower bound.

Adding *r*10 thus reinforces the earlier conclusions about CEE relative to the other schemes that use contemporaneous restrictions. Adding *r*10 has the most impact for the LR approach. Although the effects of monetary policy shocks on *y* and *p* using the LR scheme are essentially the same as for the basic model, the LR scheme generates results for *r*10 that differ sharply from the other schemes and that are difficult to understand. It thus appears that the LR approach is much more sensitive to the extension of the basic model than are the contemporaneous approaches.

Part C of Table 1 presents the relevant entries of $(I - A_0)^{-1}\Omega^{1/2}$ for the model with *r*10. As might be expected from Figure 4, the biggest differences from the basic model occur for the LR approach. The one-standard-deviation shock to *nbr* is a good bit smaller than in the basic model, and the contemporaneous effect on *ffr* is much smaller as well. The effect on *r*10 is much larger than the effect on *ffr*. The results for the contemporaneous restrictions schemes are very similar in magnitude to those for the basic model, and the effects on *r*10 are much smaller than the effects for *ffr*.

## 4. Summary and Conclusion

Many previous studies of the effects of monetary policy shocks in VAR models have used alternative methods of identifying these policy shocks and have employed different VAR models and different sample periods in the analysis. The use of alternative models and sample periods complicates isolating the effect of the identification scheme on the differing estimates of the effects of monetary policy shocks on the macroeconomy. Holding constant the VAR model and sample period, this study has compared the implications of four different procedures for identifying monetary policy shocks that use contemporaneous restrictions with a procedure that uses long-run restrictions. The four identification procedures employed that use contemporaneous restrictions are those of Christiano-Eichenbaum-Evans, Strongin, Bernanke-Mihov, and Bernanke-Blinder. The long-run restrictions approach is based on that of Blanchard-Quah. The effects of monetary policy shocks identified using each procedure are evaluated by computing impulse response functions.

The impulse response functions for the basic model reveal that monetary policy shocks identified by all procedures considered have a similar pattern of effect on output, the price level, and the federal funds rate. However, the magnitude and timing differ to some degree. It appears that the contemporaneous identification schemes of Strongin, Bernanke-Mihov, and Bernanke-Blinder and the long-run restrictions identification procedure generate impulse response functions of essentially the same magnitude for output and the price level. The Bernanke-Mihov and Bernanke-Blinder procedures do seem to generate somewhat stronger liquidity effects than do either the Strongin procedure or the long-run restrictions procedure. The results for the method of CEE differ more substantially from the others. The effects on output appear to peak sooner and die out more quickly than for the other contemporaneous identification schemes. The liquidity effect is weaker than for the Bernanke-Mihov or Bernanke-Blinder schemes, and this appears to generate the difference in results from the other schemes. A troubling aspect of the Christiano-Eichenbaum-Evans scheme is the observation that the confidence interval for the federal funds rate lies entirely above zero after a year, unlike all the other procedures. Thus, the results are quite similar for the Strongin, Bernanke-Mihov, Bernanke-Blinder, and long-run restrictions procedures, and there is little basis for selecting one of these as the preferred procedure.[10]

When the basic model is extended to include a long-term interest rate, similar results for output and the price level are found for all schemes, and similar results for the federal funds rate are found for the contemporaneous identification schemes. All the contemporaneous identification schemes indicate a small, short-lived drop in the long-term rate following an expansionary monetary policy shock. This change is smaller than for the federal funds rate, and the long-term rate returns to the initial level quicker than for the federal funds rate. However, the confidence interval for the Christiano-Eichenbaum-Evans scheme for the long-term interest rate lies entirely above zero after about six months in contrast to the other schemes, where the confidence intervals include zero after about six months. The results for the long-run restrictions scheme differ substantially for the interest rate variables. The liquidity effect on the federal funds rate is much smaller than in the basic model, and the effect on the long-term interest rate is much larger than for the contemporaneous identification schemes. Furthermore, the confidence interval for the long-term interest rate lies below zero for over two years, a very puzzling result. The results for the long-run restrictions procedure are thus much more sensitive to the addition of a long-term interest rate than are the other schemes.

When the results for both the basic model and the extended model are considered, it is difficult to choose between the Strongin and the Bernanke-Mihov scheme as a preferred approach to identification of policy shocks. These schemes share the features that total reserve shocks are assumed to be shocks to total reserve demand and that there is one way contemporaneous causality from total reserve demand shocks to nonborrowed reserves shocks. Although the Bernanke-Blinder scheme produces similar impulse response functions for output, the price level, and interest rates to those for Strongin and Bernanke-Mihov, it generates the counterintuitive result that an expansionary monetary policy shock is associated with a contemporaneous decline in total reserves. The Christiano-Eichenbaum-Evans and long-run restrictions procedures have some undesirable features. The Christiano-Eichenbaum-Evans scheme suggests a long-run positive effect on both short- and long-term interest rates of a shock to the level of nonborrowed reserves. The long-run restrictions scheme results for the federal funds rate are sensitive to the addition of a long-term rate to the model, and a monetary policy shock generates a very long-lived negative effect on the long-term rate in this scheme.

## References

Bernanke, Ben S., and Alan S. Blinder. 1992. The Federal funds rate and the channels of monetary transmission. *American Economic Review* 82:901–21.

[10] An alternative to choosing one of the identification procedures would be to use the Bayesian approach to combining the impulse response functions suggested by Lastrapes (1998).

Bernanke, Ben S., and Ilian Mihov. 1998. Measuring monetary policy. *Quarterly Journal of Economics* 113:869–902.

Blanchard, Olivier Jean, and Danny Quah. 1989. The dynamic effects of aggregate demand and supply disturbances. *American Economic Review* 79:655–73.

Christiano, Lawrence J., and Martin Eichenbaum. 1987. Temporal aggregation and structural inference in macroeconomics. *Carnegie-Rochester Conference Series on Public Policy* 26:63–130.

Christiano, Lawrence J., Martin Eichenbaum, and Charles L. Evans. 1994. Identification and the effects of monetary policy shocks. Federal Reserve Bank of Chicago Working Paper No. 94-7.

Christiano, Lawrence J., Martin Eichenbaum, and Charles L. Evans. 1996. The effects of policy shocks: Evidence from the flow of funds. *Review of Economics and Statistics* 78:16–34.

Christiano, Lawrence J., Martin Eichenbaum, and Charles L. Evans. 1998. Monetary policy shocks: What have we learned and to what end. NBER Working Paper No. 6400.

Edelberg, Wendy, and David Marshall. 1996. Monetary policy shocks and long-term interest rates. Federal Reserve Bank of Chicago. *Economic Perspectives* 20:2–17.

Evans, Charles L., and Kenneth N. Kuttner. 1998. Can VARs describe monetary policy? Federal Reserve Bank of New York Research Paper No. 9812.

Fackler, James S., and W. Douglas McMillin. 1998. Historical decomposition of aggregate demand and supply shocks in a small macro model. *Southern Economic Journal* 64:648–64.

Faust, Jon, and Eric M. Leeper. 1997. When do long-run identifying restrictions give reliable results? *Journal of Business and Economic Statistics* 15:345–53.

Goodfriend, Marvin. 1983. Discount window borrowing, monetary policy, and the post-October 6, 1979 Federal Reserve operating procedure. *Journal of Monetary Economics* 12:343–56.

Gordon, David B., and Eric M. Leeper. 1994. The dynamic impacts of monetary policy: An exercise in tentative identification. *Journal of Political Economy* 102:1228–47.

Keating, John W. 1992. Structural approaches to vector autoregressions. Federal Reserve Bank of St. Louis. *Review* 74:37–57.

Keating, John W. 1999. Structural inference with long-run recursive empirical models. Unpublished paper, University of Kansas.

Lastrapes, William D. 1998. The dynamic effects of money: Combining short-run and long-run identifying restrictions using Bayesian techniques. *Review of Economics and Statistics* 80:588–99.

Lastrapes, William D., and George Selgin. 1995. The liquidity effect: Identifying short-run interest rate dynamics using long-run restrictions. *Journal of Macroeconomics* 17:387–404.

Leeper, Eric M., Christopher A. Sims, and Tao Zha. 1996. What does monetary policy do? *Brookings Papers on Economic Activity* 2:1–78.

McCarthy, Jonathan. 1995. VARs and the identification of monetary policy shocks: A critique of the Fed reaction function. Unpublished paper, Federal Reserve Bank of New York.

Pagan, Adrian R., and John C. Robertson. 1995. Resolving the liquidity effect. Federal Reserve Bank of St. Louis. *Review* 77:33–54.

Pagan, Adrian R., and John C. Robertson. 1998. Structural models of the liquidity effect. *Review of Economics and Statistics* 80:202–17.

Rudebusch, Glenn D. 1998. Do measures of monetary policy in a VAR make sense? *International Economic Review* 39:907–31.

Shapiro, Matthew, and Mark Watson. 1988. Sources of business cycle fluctuations. *NBER Macroeconomics Annual* 3: 111–48.

Sims, Christopher A. 1998. Comment on Glenn Rudebusch's 'Do measures of monetary policy in a VAR make sense?'. *International Economic Review* 39:933–41.

Strongin, Steven. 1995. The identification of monetary policy disturbances: Explaining the liquidity puzzle. *Journal of Monetary Economics* 35:463–97.

Walsh, Carl E., and James A. Wilcox. 1995. Bank credit and economic activity. In *Is bank lending important for the transmission of monetary policy?,* edited by Joe Peek and Eric S. Rosengren. Federal Reserve Bank of Boston Conference Series No. 39, pp. 83–112.