Empirical Analysis of Policy Interventions

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Abstract:
We construct linear projections of macro variables conditional on hypothetical paths of monetary policy, using as an example an identified VAR model. Hypothetical policies are restricted to ones where both the policy intervention and its impacts are consistent with history—otherwise the linear projections are likely to be unreliable. We use the approach to interpret Federal Reserve decisions, modeling their frequent reassessment in light of new information about the tradeoffs policymakers face. The interventions we consider matter: they can shift projected paths and probability distributions of macro variables in economically meaningful ways.

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1.   Introduction

This paper offers a framework for interpreting the kinds of policy choices that central banks make at regular policy meetings. The approach is positive, adopting the perspective of an outside observer—the econometrician—to (i) address counterfactual questions with linear projections of macro variables conditional on exogenous policy interventions; (ii) assess how reliable linear projections are likely to be for a given intervention; (iii) account for uncertainty about model parameters and realizations of shocks over the forecast horizon. The framework is suited to consider interventions that would not be construed as changes in policy behavior that bring forth systematic changes in private decision rules.

Our approach differs from two dominant approaches to counterfactual policy evaluation—the rational expectations program that Lucas and Sargent (1981) laid out and the reduced-form forecasting approach that Doan, Litterman and Sims (1984) developed. The rational expectations program identifies the mapping from policy behavior to private expectations and decision rules to evaluate the impacts of adopting a new policy process. This paper does not address questions associated with changing the policy process. Instead, it analyses policy choices that are made sequentially and that are not wholly different from past choices.

Doan, Litterman and Sims (1984) use forecasting models to project the effects of alternative sequences of policy variables, taking private expectations and decision rules as given by history. Their approach is designed to address routine policy interventions. Like us, they argue that if the interventions do not constitute regime changes, then linear projections are likely to be reliable.

Unlike us, Doan, Litterman and Sims focus on reduced-form forecasting models in which monetary policy behavior is not identified. We extend Doan, Litterman and Sims's methods to environments in which policy behavior is identified, so exogenous variation in policy is separated from systematic responses of policy to the economy. Identification allows us to check whether the hypothetical policies we consider are consistent with linear projections, which hold estimated private decision rules fixed over the projection period.

We emphasize the reliability of linear projections because we interpret Lucas's (1976) critique as pointing toward a source of nonlinearity that may be important in practice.[1] When policy interventions create sustained dynamic patterns of change in policy variables, agents may grow to believe policy regime has shifted and adjust their expectations and decision rules accordingly. Learning dynamics can produce smoothly evolving private decision rules that undermine the accuracy of linear projections: a model that is linear if regime is fixed, becomes nonlinear, with the nonlinearity produced by the economic behavior that Lucas emphasizes. Linear projections will remain reliable, however, if the interventions and their effects remain in line with historical fluctuations.[2]

An outgrowth of identifying policy behavior is the decomposition of policy choice into two components: a deterministic function of observable variables and an exogenous random variable. All empirical work on policy performs a similar decomposition—work ranging from Taylor's (1993) simple rules to Romer and Romer's (1989) dummy variables to Fair's (1984) simultaneous equations, as well as identified VARs. The presence of a random exogenous term in the policy rule implies that private agents put probability mass on more than one policy action, even though their best estimate of the action is uniquely dictated by the deterministic part of the rule. This leaves room for counterfactual policy analysis even in a linear model. It takes the form of contemplating alternative interventions that a priori agents think could occur; and if those actions are taken, agents are not shocked into thinking the underlying policy process has changed. We show that the impacts of this class of interventions can be reliably projected using a linear model. And these interventions matter: they can shift the projected paths and probability distributions of macro variables in economically meaningful ways. How much they matter

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[1] Sims (1998) puts forth this interpretation of the Lucas critique.
[2] Leeper and Zha (2001b) explore these points in a theoretical model.

depends on both the nature of the intervention and the elasticities of private behavior with respect to policy.

The efficacy of our approach hinges on whether it can be applied without implying policy choices that would force the economy into a new policy regime. We offer some statistical guards against this. We compute an ex-ante metric for assessing whether a contemplated intervention is consistent with the prevailing policy regime. The metric is designed to flag any hypothetical intervention that would trigger a change in private agents' beliefs about the policy process, shift their decision rules, and undermine the linear projections. By that metric we find that many interventions the Federal Reserve actually considers fall within the prevailing regime. We also compute an ex-post version of the metric, which reports whether the choices actually taken affected macro variables in ways that are consistent with historical effects.

We illustrate the approach in a small model of the U.S. economy that identifies monetary policy behavior. We assess the estimated model's in-sample and out-of-sample fits, examine the model's stability, and evaluate the model's suitability for the policy questions we put to it.

We focus on two recent episodes. The first episode is the 1990-1991 recession, which was a time of aggressive easing of policy and one in which the Fed perceived that it faced difficult tradeoffs between inflation and real activity. The analysis offers a probabilistic ex-ante assessment of the tradeoffs, conditional on alternative policy actions. The second episode is the Fed's 1994 "pre-emptive strike" against inflation, during which the federal funds rate increased 300 basis points in a year. Sequential forecasts from the model suggest that the incremental increases in the funds rate that occurred through the year are a natural outcome when policy choices are reappraised in light of new information arriving in early 1994. The analysis shows how uncertainty about future exogenous disturbances leads to conservative policymaking, as Brainard (1967) advocates, and it formalizes Blinder (1997) description of how central banks reappraise their policy decisions.

2.   An Econometric Framework for Policy Analysis

This section specifies and estimates a small structural model of American monetary policy behavior. The model's fit, stability, and suitability for policy analysis are scrutinized. The section also reports the model's impulse response functions, which underlie the conditional policy projections reported below, and develops a metric for assessing policy interventions.

2.1   The Model

Actual policy behavior is a complicated function of a high-dimensional vector of variables. Policymakers choose R_t, the vector of policy choices at date t, as a function of their information set, Ω_t. Actual policy behavior is a function g such that

$$R_t = g(\Omega_t). \tag{1}$$

(1) describes historical policy behavior.

We assume that private agents are not privy to the details of the policymakers' decision problems, including the policymakers' incentives and constraints. That is, they observe the information set $S_t \subset \Omega_t$. Agents perceive that policy is composed of a regular response to the state of the economy that they observe at time t, S_t, and a random part, ε_{Pt}. The econometric model of policy is:

$$R_t = f(S_t) + \varepsilon_{Pt}. \tag{2}$$

We take f to be linear. ε_{Pt} is exogenous to the econometric model. A policy regime is a choice of g, which implies an f and a stochastic process for ε_P.[3]

The econometric model embeds the policy behavior in (2) in a system of equations. If y_t is an $(m \times 1)$ vector of time series, the structural form is

$$\sum_{s=0}^{p} A_s y_{t-s} = \varepsilon_t, \tag{3}$$

where ε_t is a vector of i.i.d. structural disturbances that are exogenous to the model. Those disturbances hit both nonpolicy and policy sectors of the economy, so

$$\varepsilon_t = \begin{bmatrix} \varepsilon_{Nt} \\ \varepsilon_{Pt} \end{bmatrix}, \tag{4}$$

where ε_{Nt} is the vector of nonpolicy disturbances.

Counterfactual policy questions are addressed using the structure in (3) to project y conditional on hypothesized paths for ε_P, as Marschak (1953) and the Cowles Commission

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[3] It is beyond the scope of this paper to rationalize randomness in policy behavior. See Sims (1987) for one detailed rationale and Leeper, Sims and Zha (1996) for further discussion.

instructed. We impose that ε_{Pt} is a vector of exogenous random variables, uncorrelated with all the nonpolicy exogenous disturbances in the economy. The errors are Gaussian with

$$E\left(\varepsilon_t \varepsilon_t' \,|\, y_{t-s}, s > 0\right) = I, \quad E\left(\varepsilon_t \,|\, y_{t-s}, s > 0\right) = 0, \quad \text{all } t. \tag{5}$$

The A_s matrices and the probability distribution of ε define the model's structure.

Assuming the matrix of contemporaneous coefficients, A_0, is non-singular, there is a representation of y in terms of the impulse responses functions:

$$y_t = \sum_{s=0}^{t-1} C_s \varepsilon_{t-s} + E_0 y_t. \tag{6}$$

The elements of C_s report how each variable in y responds over time to the behavioral disturbances in ε. $E_0 y_t$ is the projection of y_t conditional on initial conditions. The reduced form of (3) is

$$\sum_{s=0}^{p} B_s y_{t-s} = u_t, \tag{7}$$

with $B_0 = I$ and the covariance of the reduced-form errors, u, is $\Sigma = A_0^{-1} A_0^{-1\prime}$.

Expressions (3) and (7) imply a linear mapping from the reduced-form errors to the behavioral disturbances:

$$u_t = A_0^{-1} \varepsilon_t. \tag{8}$$

Identification of the structural form follows from imposing sufficient restrictions on A_0 so that there are no more than $m(m-1)/2$ free parameters in A_0.

2.2   Identification

We estimate a version of the model in (3) that contains six variables and three sectors. The identification scheme follows the general approach in Gordon and Leeper (1994) and Sims and Zha (1998b).[4] Table 1 shows the restrictions placed on the contemporaneous coefficient matrix A_0. There are no restrictions on lagged variables.

Three goods market variables—real GDP (y), consumer prices (P), and the unemployment rate (U)—compose the production sector, and are the ultimate objectives of monetary policy.

We do not model the markets for reserves and a broad monetary aggregate, opting for compactness to treat the *M2* money stock as the aggregate and the federal funds rate (R^f)—the monetary policy instrument—as the price that clears the money market. The third sector describes an "information variable"—commodity prices (*CP*)—that is available at high frequencies and reacts instantaneously to shocks from all sectors of the economy.

The data are monthly from January 1959 to September 1998 and described in Appendix B. Monthly GDP is interpolated from quarterly GDP using the procedure that Leeper, Sims and Zha (1996) describe. All data are logarithmic except for the federal funds rate and the unemployment rate. We estimate with 13 lags.

The identification treats the production sector as predetermined for the rest of the system, reflecting the view that production, pricing, and employment decisions do not respond immediately to shocks from outside the sector. Production sector variables interact only with each other within the period. Money market variables and information variables do not enter this sector, reflecting sluggishness in the goods market due to contracts and advance planning of production. Distinct behavioral equations within the production sector are not identified; instead, the coefficients are arranged in lower triangular form in the order *y, P,* and *U*.

The monetary and information sectors interact simultaneously, with the strongest simultaneity determining the money stock and the federal funds rate. The demand for nominal money balances depends on the short-term nominal interest rate, real income (proxied by real GDP), and the price level. We do not impose short-run homogeneity in prices. Changes in ε_{MD} reflect exogenous shifts in money demand.

We base the specification of monetary policy behavior in (2) on the information available to the Federal Reserve within the month. During the month, the Federal Reserve sets its interest rate instrument based on current observations on the money stock and commodity prices.[5] We set to zero the coefficients on *y, P*, and *U* in the policy function because within the month the

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[4] For further discussion of the methods, see Christiano, Eichenbaum and Evans (1999) or Leeper, Sims and Zha (1996). Appendix A describes the Bayesian methods used to estimate the model.
[5] We allow policy to choose not to react strongly to commodity prices by shrinking the prior standard deviation on the coefficient of *CP* toward the zero prior mean by a factor of .05.

Federal Reserve does not observe these variables directly.[6] The error term in the monetary policy equation, ε_P, represents exogenous policy interventions.

An efficient markets assumption guides the specification of the information sector, so commodity prices may respond to all variables immediately. The error term ε_I is the exogenous information disturbance.

2.3   Parameter Estimates and In-Sample Fit

Table 2 reports the contemporaneous coefficients along with 68 percent equal-tailed probability intervals for the behavioral coefficients, estimated over the full sample period. The money demand and monetary policy equations have reasonable economic interpretations. The interest elasticity of demand is negative and the output elasticity is positive. The price elasticity is small and imprecisely estimated. Monetary policy responds strongly to the money stock: disturbances that raise the money stock induce the Fed to increase the federal funds rate. Although the estimates seem to suggest the Fed does not react much to information contained in commodity prices, this interpretation may be misleading. In the policy equation the coefficients on R^f and *CP* are highly correlated (.97), as are the coefficients on *M2* and *CP* (.72); these correlations muddy inferences about individual coefficients.

All but two coefficients are tightly estimated in the information sector. The coefficients on output and *M2* are highly correlated with the coefficients on the price level and unemployment, so the separate influences cannot be discerned. The estimates imply a quick reaction of commodity prices to exogenous disturbances from elsewhere in the economy.

We base a probabilistic assessment of the model's overall fit on the shape of the likelihood surface, rather than on tests of whether individual coefficients are different from zero [Sims and Zha (1999)]. Table 2 nonetheless displays the (marginal) .68 probability intervals for individual parameters to show that they are skewed, with most of the probability mass concentrated around the maximum likelihood estimates. Because we care about the equilibrium effects of exogenous policy actions, inferences do not rest on an individual parameter in A_0. Equilibrium effects,

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[6] The Fed has some contemporaneous information about these variables. We experimented with soft zeroes on *P* and *U* and found that the more we relaxed the zero restrictions the more likely the identification was to produce nonsensical responses to exogenous monetary policy actions.

which section 2.4 discusses, depend on the joint distribution of all the parameters, not the marginal distribution of an individual parameter.

The model is overidentified. To evaluate whether the data favor the restricted model relative to the unrestricted (reduced-form) model, we conduct an exact small-sample comparison using a Bayes factor.[7] Given the data, Y, we compare the probability densities of Y under the two models, denoted by $p(Y|M_R)$ and $p(Y|M_U)$ where M_R represents the restricted model and M_U the unrestricted model. Let Θ_R be the restricted parameter space and Θ_U be the unrestricted parameter space. The Bayes factor is

$$\frac{p(Y|M_R)}{p(Y|M_U)} = \frac{\int_{\theta \in \Theta_R} p(Y|\theta) p(\theta) d\theta}{\int_{\theta \in \Theta_U} p(Y|\theta) p(\theta) d\theta}, \tag{9}$$

where $p(\theta)$ is our reference prior probability density function, described in Appendix A. The Bayes factor is a standard reporting procedure in Bayesian hypothesis testing. In our case, since Θ_R is a lower dimensional parameter space $\Theta_R \subset \Theta_U$, the Bayes factor compensates effectively for over-parameterization by integrating the likelihood $p(Y|\theta)$ over the two different parameter spaces Θ_R and Θ_U. Recent developments in Bayesian analysis and computer technology, surveryed in Geweke (1999), allow us to calculate the Bayes factor accurately as

$$\log p(Y|M_R) - \log p(Y|M_U) = 0.9. \tag{10}$$

The data weakly favor the restricted model, suggesting that to determine the model's consistency with data it is important to compensate for over-parameterization.[8]

To use the model to compute projections conditional on alternative interventions, the policy disturbances must be uncorrelated with the other shocks. We use a small-sample procedure to check whether ε_{Pt} and ε_{Nt} are uncorrelated. For each simulated draw from the posterior

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[7] A number of statistics are used to approximate Bayes factors, including the likelihood ratio criterion and the Schwarz criterion. The method of Laplace and the Schwarz criterion are employed as higher-order corrections to the likelihood ratio criterion as an approximate Bayes factor [Schervish (1995, chapters 4 and 7) and Sims (1999b)].

[8] The Schwarz criterion computes the chi-square statistic with degrees of freedom multiplied by the log of the sample size. The chi-square statistic, which is twice the difference in log likelihood values of the unrestricted and the restricted models at their peaks, is 18.56. With a critical value of 18.11, the Schwarz criterion weakly rejects the restricted model.

distribution of the model's parameters, we compute the sequences of exogenous disturbances consistent with the data and calculate the correlation matrix for these sequences. Table 3 reports .68 probability intervals for the correlations among the exogenous shocks, along with the correlations calculated at the maximum likelihood estimates of the parameters. ε_{Pt} is uncorrelated with ε_{Nt}, so restriction (5) holds well.[9] We also compute a posterior .68 probability region for the correlations of policy shocks with five non-policy shocks. Zero values of all five correlations fall jointly inside the region.[10] With the policy disturbance uncorrelated with other shocks, it is reasonable to condition on a path of exogenous policy interventions and draw nonpolicy disturbances independently.

2.4   Dynamic Impacts of an Exogenous Monetary Policy Action

Figure 1 displays the C_s's in (6) over 48 months for the six variables in the model when there is a unit exogenous monetary policy contraction. The solid lines are the maximum likelihood estimates of responses and the dashed lines are equal-tailed error bands containing .68 probability, following Sims and Zha (1999).

The contraction raises the funds rate initially and immediately decreases the money stock and commodity prices, both of which continue to decline smoothly over the four-year horizon. After a brief delay, output falls and stays lower, while unemployment rises. Six months after the exogenous action, both output and unemployment are likely to differ from their initial levels. Consumer prices adjust more slowly and are unlikely to be appreciably lower for about a year. After a year prices decline smoothly and remain well below their initial level.

The response of the interest rate to an exogenous policy contraction stands out. In Figure 1 the initial liquidity effect lasts about eight months, but by 18 months the funds rate lies well

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[9] The table suggests, however, that the model does not fit the correlation between the money demand shock and the unemployment shock well, affecting some behavioral inferences, but not the ones we pursue.

[10] The probability region is computed from the small-sample distribution of the random variable $\left(\delta - \hat{\delta}\right)' \Xi^{-1} \left(\delta - \hat{\delta}\right)$ where $\hat{\delta}$ is the ML estimate of a vector of the five correlations, δ, and Ξ is the (simulated) covariance matrix of δ. Based on our simulations,

$$\text{Prob}\left(\left(\delta - \hat{\delta}\right)' \Xi^{-1} \left(\delta - \hat{\delta}\right) \leq 6.06\right) = .68 \text{ and } \left(\mathbf{0} - \hat{\delta}\right)' \Xi^{-1} \left(\mathbf{0} - \hat{\delta}\right) = 5.06.$$

below its initial level. The decline in the funds rate then persists. Friedman (1968) and Cagan (1972) describe this path following a monetary contraction as a short-lived liquidity effect followed by income and expected inflation effects. After four years the declines in inflation and the federal funds rate are the same size, as predicted if expected inflation is the dominant source of fluctuations in nominal rates over long periods. The responses in the figures suggest that to lower inflation persistently the Fed should raise the funds rate only briefly. Because lower inflation is ultimately associated with a lower funds rate, the Fed must begin to reduce the rate within about a year, and then keep it lower.

2.5   Parameter Stability

Because many analyses work from the premise that U.S. monetary policy has shifted substantially over the post-World War II period, some readers may object to our treating the entire sample as a fixed policy regime. We think the issue is unsettled. Bernanke and Mihov (1998a, 1998b) carefully test the stability of the reduced-form coefficients and the residual covariance matrices in VARs containing many of the same time series in our model. They conclude that the reduced-form parameters are stable.[11] In contrast, tests of the covariance matrices of VAR innovations—from which the structural model's A_0 parameters are obtained in (8)—find evidence of breaks in late 1979 or early 1980 and between early 1982 and early 1988. The results conform to Sims's (1999a) reaction function estimates. Using a hidden Markov chain approach, he finds strong evidence that the Fed's responsiveness to commodity price inflation deviates from a linear, Gaussian reaction function. But in terms of model fit, he reports that variations in the size of the errors in the policy rule are more important than variation in the coefficients of the rule.

Much recent attention focuses on the finding that coefficients in simple specifications of the Fed's policy rule shift over time [e.g., Clarida, Gali and Gertler (2000) and Taylor (1999)]. The

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[11] They use a Lagrange multiplier variant of Andrew's (1993) test, which checks every possible break point and tests the significance of the associated LM statistics. Using quarterly data, Sims and Zha (1998b) corroborate Bernanke and Mihov's results by comparing a model fit from 1964:1 to 1994:1 to one fit separately to 1964:1-1979:3 and 1983:1-1994:1. Both the Akaike and Schwartz criteria favor the restricted model.

simple rules emerge from restrictions on the policy behavior embedded in VARs.[12] For our

purposes, what matters is whether the dynamic responses to exogenous policy interventions are

stable. Those responses depend on the system of equations, not just the policy rule. To explore

the system properties, we compute the dynamic effects of a one-period unit intervention using

models estimated over a variety of sub-periods.

Figure 2 reports responses to all six exogenous disturbances for models estimated over four

sub-periods. The periods covered are 1959:1-1998:9 (entire sample); 1959:1-1979:9 (pre-

Volcker); 1959:1-1982:12 (including non-borrowed reserves targeting); 1959:1-1998:9 with

1979:10-1982:12 eliminated (excluding non-borrowed reserves targeting). The reserves

targeting period from 1979:10 to 1982:12 is too brief to obtain reliable estimates from just those

years, so we treat the three years as potentially anomalous relative to the entire sample.

Although it is popular to treat the Greenspan era (from late 1987 on) as a distinct policy regime,

we do not. The last 12 years of the sample contain only one mild recession and few interesting

exogenous disturbances, except possibly from the stock market.

The qualitative responses in Figure 2 to exogenous policy interventions are robust across

sub-periods.[13] There is some tendency for the responses estimated by eliminating the reserves

targeting episode to be slightly weaker and for those estimated only through the end of 1982 to

be somewhat stronger. Otherwise, the responses to policy are even quantitatively very similar.

Two implications flow from finding that parameters are stable. First, over the post-1959

period either monetary policy has resided in a single regime or the various regimes have been too

close to be detected statistically. Second, there is no evidence that agents' beliefs about regime

have changed in ways that matter quantitatively for decision rules. This justifies our use of the

entire sample period to estimate the model and to form the history against which we evaluate

whether linear projections are likely to reliably predict the impacts of an intervention.

Overall, the model's fit is very good. The overidentifying restrictions do not invalidate our

uses of the model. Those uses rely on whether exogenous policy interventions are uncorrelated

with nonpolicy disturbances. Tests confirm that this assumption holds well. An exact small-

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[12] Leeper and Zha (2001a) study connections between models with simple rules and identified
VARs.
[13] Qualitative differences do appear in responses to some non-policy disturbances, especially
money demand and goods market shocks. This finding is consistent with Hanson (2001).

sample comparison of the restricted and unrestricted models using a Bayes factor somewhat favors the restricted model.[14] Finally, the dynamic responses of the economy to exogenous monetary policy interventions are quite stable over time. Taken together, these indicators of fit and robustness lend credibility to the model's projections conditional on policy.

2.6   Measuring Policy Interventions

In the VAR the K-period forecast errors, given information at T, that arise from the intervention $I_T = \{\tilde{\varepsilon}_{PT+1},...,\tilde{\varepsilon}_{PT+K}\}$ are

$$\eta_P(T,K) = \sum_{s=0}^{K-1} C_s(:,i)\tilde{\varepsilon}_{PT+K-s}, \tag{11}$$

where the $C_s(:,i)$ are the impulse response matrices associated with the monetary policy disturbance in equation i. The vector $\eta_P(T,K)$ is normally distributed with mean zero and variance vector $\sum_{s=0}^{K-1} C_s^2$. We transform $\eta_P(T,K)$ to a vector of standard normal variables, $\eta_P^*(T,K)$ by scaling by the standard error of the η_P 's.

Linear projections are reliable if both the intervention and its impacts are within two standard deviations of their historical fluctuations. Specifically, we deem projections reliable if over a specified forecast horizon K, for variable i, we find that

$$\left| e_i \eta_P^*(T,K) \right| < 2, \tag{12}$$

where e_i is a row vector of zeros with unity in the ith column. η^* reports how unusual a conditional projection is relative to the historical impacts of exogenous shifts in policy. Large values of the statistic suggest the intervention is predicted to generate impacts on macro variables that lie outside the historical range of policy effects. When an intervention violates (12), the reliability of linear projections is called into doubt. The metric when $K > 0$ is ex-ante in the sense that it assesses a hypothetical intervention. When $K < 0$ the metric is ex-post, computed from the realized $\hat{\varepsilon}_P$'s, and it assesses actual interventions.

The metric isolates exogenous policy changes as being most pertinent to evaluating policy and assessing the impact of the Lucas critique. This improves on the reduced-form methods that

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[14] The model's out-of-sample fit compares favorably to other econometric models [Appendix C].

Doan, Litterman and Sims (1984) employ. The statistic at horizon K reflects the full impacts of interventions up to time $T + K$. The metric emphasizes the dynamic impacts, rather than the magnitude of the intervention itself because if policy impacts are small, even very large and unprecedented interventions that shift beliefs about prevailing regime may not generate significant changes in decision rules.

3.   Some Practical Analysis of U.S. Monetary Policy

This section addresses some questions that Federal Reserve officials may have asked during the 1990s. To do so we compute policy projections under several alternative scenarios tied to actual U.S. policy experience. In each case we check the contemplated interventions to ensure that the linear projections are likely to be reliable. We focus on two periods of aggressive policy moves: the 500 basis point drop in the funds rate from late 1990 through 1993 and the 300 basis point increase in the rate between January 1994 and early 1995.

3.1   "Exogenizing" Policy Can Produce Incredible Results

The pure forecasting exercise that Doan, Litterman and Sims (1984) discuss, and which is implemented in many practical applications, can generate implausible results. We illustrate the point with the dramatic period beginning in September 1990. The period clarifies a problem that applies generally. There are many ways to "exogenize" policy. One could change the systematic part of policy—the function f in (11)—but this can be couched as a particular ε_P sequence. Bernanke, Gertler and Watson (1997) conduct the counterfactual exercise of changing f. Hamilton and Herrera (2000) convert the change in f to a sequence of $\hat{\varepsilon}_P$'s and compute our metric to conclude the exercise is problematic.[15] Here we provide two more examples.

We mimic conventional out-of-sample forecasting by "exogenizing" the federal funds rate, forcing the actual path of R^f over the next 48 months to be produced solely by exogenous policy actions. For each month of the forecast period we calculate the $\tilde{\varepsilon}_{Pt}$ that produces a forecast that equals the actual funds rate, R_t^f, given the value of $\tilde{\varepsilon}_{Pt}$. Instead of conditioning on reduced-

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[15] Other recent papers that "exogenize" policy in this manner include Bernanke, Gertler and Watson (1997), Sims (1998), Sims and Zha (1998b), and Dungey and Pagan (2000).

form errors, our interpretation of Doan, Litterman and Sims derives the sequence of exogenous policy actions consistent with the R^f path.

Table 4 reports the central tendencies for the paths of the funds rate, output growth, inflation, and unemployment over the four-year forecast horizon. The actual funds rate fell dramatically over the period, from an annual average of almost 8 percent in 1990 to an average of less than 3 percent in 1993. It takes a sequence of same-signed exogenous impulses to generate that decline solely from exogenous policy actions. The impulses drive output growth and inflation into double digits and the unemployment rate down to 3.5 percent. Given the actual path for the policy instrument, the implausible paths for macro variables suggest that much of the movement in R^f over the period was an endogenous response of policy to nonpolicy disturbances.[16]

The results in Table 4 come from hypothesizing policy behavior that deviates substantially from any behavior observed over the sample. The forecast errors for each variable are extremely unlikely, as shown by the statistics below:

Variable	$\eta_P^*(90:9,48)$
R^f	-2.93
y	18.95
P	9.64
U	-18.96

The metric warns against placing much confidence in the linear projections.

It is commonplace for central banks to condition forecasts on an unchanged path of the policy instrument. The Bank of England (2000), for example, projects GDP growth and inflation over horizons exceeding two years, under the assumption that the official interest rate is constant.[17] As the state of the economy is forecasted to change, however, an unchanged path of the instrument always requires some pattern of intervention. Whether that intervention is likely to shift private decision rules depends both on how persistent the intervention is and on how large the dynamic impacts are.

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[16] Altig, Carlstrom and Lansing (1995) assume a path of exogenous monetary policy disturbances to use their general equilibrium model to simulate the effects of reducing inflation by a percentage point over a two-year period and holding it at the lower value for three more years. The model predicts implausible paths for the nominal interest rate and output growth.
[17] Central banks frequently interpret a constant instrument path an unchanged stance of policy [Board of Governors of the Federal Reserve System (1999) or Sveriges Riksbank (1999)].

In September 1990, the federal funds rate stood at 8.20 percent. Suppose the Fed were to hold the rate fixed at 8.20 percent over the next 4 years. Doing this requires a sequence of positive $\tilde{\varepsilon}_P$'s that increase nearly monotonically over the forecast horizon, and average 1½ standard deviations in year 3 and over 2 in the last year. This intervention implies large forecast errors:

Variable	$\eta_P^*(90\!:\!9,48)$
y	-7.09
P	-4.51
U	6.69

Linear forecasts are unlikely to be reliable.

3.2   Routine Policy Interventions Can Matter

We now address whether conditioning on the class of interventions we consider is informative about routine policy decisions. It turns out that this class of interventions is rich: it can generate economically meaningful shifts in the distributions of forecasted macro variables and clarify the tradeoffs facing policymakers. The analysis demonstrates that a complete probability model, which quantifies the uncertainties that Brainard (1967) emphasizes, can answer complex joint probability questions about those tradeoffs.

We conduct the analysis through the eyes of an econometrician who has information about the economy through September 1990. Minutes of the October 2, 1990 FOMC meeting reveal that the Fed predicted a mild downturn in economic activity followed by a rapid resumption of moderate growth. The minutes report that "insofar as could be judged on the basis of traditional indicators, the available data did not point to cumulating weakness and the onset of a recession."[18] Political developments in the Middle East, however, generated concerns about future oil prices and created uncertainty about the outlook for inflation. Although the domestic policy directive that emerged from the meeting sought "to maintain the existing degree of pressure on reserve position," several FOMC members dissented. One member favored immediate easing and three members opposed the FOMC's perceived leaning in favor of easing.

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[18] Board of Governors of the Federal Reserve System (1990, p. 135).

In light of the dissension among FOMC members, we consider two scenarios. The first scenario conditions forecasts on the actual path of the federal funds rate from October 1990 to January 1991. An alternative scenario considers tighter policy over those four months.

Figure 3 reports the actual time series, the out-of-sample forecasts conditional on the actual path of the funds rate, and 68 percent probability bands for the forecasts. The actual funds rate was 8.11 percent in October, 7.81 percent in November, 7.31 percent in December, and 6.91 percent in January. With that path of the funds rate, there is substantial probability that inflation will rise above 5½ or 6 percent through 1993, real growth will fall below 1 percent in 1991, and unemployment will rise to near 7 percent through 1993. Based on the path of the funds rate, it may appear that the Fed was concerned primarily about recession. As it happened, inflation fell to 3 percent by 1992, a recession occurred from July 1990 to March 1991 (according to the NBER dating), and unemployment hit 7½ percent in 1992. Of course, as the model's forecasts confirm, policymakers were unaware in October that the recession began three months earlier.

The FOMC minutes report that some policymakers were concerned about higher inflation. Those FOMC members might want to see forecasts conditional on tighter policy. The forecasts assume an intervention that raises the funds rate by 50 basis points in October (to 8.70 percent) and an additional 25 basis points over the period from November 1990 to January 1991. Ex-ante it appears that tighter policy would reduce the likelihood of higher inflation, but at the cost of raising the probability of negative real growth in 1991 (Figure 4). The point forecasts of output growth and unemployment, conditional on tighter monetary policy, come very close to the actual paths of the variables in 1991 and 1992. In spite of the sharp decline in the projected funds rate in 1991-93, the intervention exerts a contractionary influence.[19]

Debate during the October 1990 FOMC meeting centered on the tradeoffs associated with alternative policy choices. The tradeoffs can be framed as joint probability statements. To

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[19] To hit the actual path of the funds rate the intervention is $I_T = \{0.5, 0.1, -0.7, -0.7, 0, ..., 0\}$ and for tightening it is $I_T = \{2.3, 1.7, 1.0, 0.9, 0, ..., 0\}$. The resulting metrics are

Variable	$\eta_P^*(90:9,48)$ Actual R^f	$\eta_P^*(90:9,48)$ Tighter Policy
y	0.10	-0.69
P	0.19	-1.49
U	-0.02	0.04

answer the concerns over economic slowdown and higher inflation, Table 5 reports a variety of joint probabilities involving real GDP growth in 1991 or 1992 or 1993 and inflation in 1992 and 1993, conditional on two alternative policies. "Tighter" policy assumes the same counterfactual policy behavior as in Figure 4, while "Actual R^f" adjusts policy to be consistent with Figure 3.

The probabilities put a sharp point on the tradeoffs Federal Reserve officials perceived they faced. In terms of the marginal probabilities, tighter policy makes it very likely that inflation will remain low in 1992 and 1993 (below 5½ percent), but it also produces a better than 50 percent chance of a recession in 1991 (negative real GDP growth for the year). More relevant to policymakers is the apparent tradeoff: tighter policy creates a one-third chance of a recession in 1991 or 1992 or 1993 and low inflation in 1992 and 1993.

For an intervention to generate the actual path of the funds rate, the Fed would have to tighten slightly in October and November and then ease in December and January. The column in the table labeled "Actual R^f" reports these results. This policy reduces by half the marginal probability of a recession in 1991 while lowering the marginal likelihoods of low inflation in 1992 and 1993. It also greatly reduces the joint probability of a recession in 1991 or 1992 or 1993 and low inflation in 1992 and 1993. Again the tradeoff is clear: the probability of no recession coupled with inflation over 5½ percent now exceeds 40 percent, compared to 18 percent when policy is tighter.

3.3   *Appraising and Reappraising Policy with Sequential Interventions*

Blinder's (1997) description of the appraisal/reappraisal process inherent in routine policymaking is echoed by Kohn (1995, p. 235) who observes that policymakers must "be flexible in revising forecasts and the policy stance in response to new information." This perspective helps to understand the Federal Reserve's "preemptive strike" against inflation in 1994. Our analysis shows that in February the tighter policy looked to be sufficient to offset higher inflation in 1996 and 1997; it was not sufficient by April, once three months of new information arrived. When we reappraise policy in April, a further tightening appears necessary to preempt inflation. The analysis puts empirical flesh on Brainard's (1967) argument for gradualism.

Figure 5 displays actual data and out-of-sample forecasts made in January under two alternative policy scenarios for February through May. Given that the federal funds rate had

been nearly constant at 3 percent over the previous year, a natural baseline maintains this constancy through May. The first intervention holds the funds rate constant for 4 months. That policy portends rising inflation over the next several years, exceeding 3½ percent in 1996 and 1997; but policymakers do not seem to face an unpleasant tradeoff between inflation and real activity. Real GDP is expected to grow at least 3 percent annually from 1995 to 1997, while the unemployment rate is projected to continue to decline.

Key policy questions at the time were when to raise the rate and how much to raise it. To address these questions we consider an alternative policy of moderate tightening that raises the funds rate along its actual path. Although the actual funds rate rose a full percentage point from January to May, that path requires a relatively small exogenous intervention.[20] Even moderate tightening in January shifts the projected inflation path down without severely affecting real activity (Figure 5).

The Fed reappraises policy in April with three additional months of news about the economy. Now the baseline of a constant funds rate at 3.75 percent leads policymakers to expect inflation will once again drift toward the 3½ to 4 percent range in 1996 and 1997 (Figure 6). The outlooks for output and unemployment remain promising, so the Fed still does not face a difficult tradeoff. A somewhat stronger tightening move to match the actual path of the funds rate from May through August pushes the funds rate to 4.47 percent in August. Tighter policy shifts the mean forecast of inflation down below 3 percent through 1997 without risking recession.[21]

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[20] Holding the funds rate constant requires the intervention $I_T = \{-.3, -.2, -.2, -.2, 0, ..., 0\}$, while tracking the actual funds rate calls for the intervention $I_T = \{.5, 0, .5, 1.1, 0, ..., 0\}$. The η^* statistics associated with the interventions are:

Variable	$\eta_P^*(94:1,48)$ Constant R^f	$\eta_P^*(94:1,48)$ Actual R^f
y	0.12	-0.28
P	0.23	-0.53
U	-0.02	0.07

[21] A constant R^f requires $I_T = \{.1, -.3, 0, ..., 0\}$ and the actual path of R^f requires $I_T = \{.8, .4, .1, .8, 0, ..., 0\}$. The associated η^*'s are:

By reappraising their decisions in light of updated forecasts, policymakers move cautiously against inflation. So long as forecasts extend far enough into the future to capture monetary policy's lagged effects, the gradual approach can be successful. This analysis formalizes Blinder's (1997) description of policymaking. It also illustrates why uncertainty about future exogenous disturbances may lead policymakers to move cautiously, as Brainard (1967) instructs.

Some readers might worry that if policy makers use the kind of on-going policy analysis that we describe as a basis for policy decisions, agents' beliefs about regime might shift. We perform an ex-post check of how likely this situation was in January 1994. We estimate the model through January 1994, back out the realized policy disturbances, $\hat{\varepsilon}_P$'s, over the preceding 48 months, and compute an ex-post metric using the realized disturbances. The metric yields

Variable	$\eta_P^*(90:2,-48)$ Ex-post
y	-1.96
P	-1.76
U	1.94

for forecasts beginning in February 1990 and ending in January 1994. These values for η^* warn that over this period agents' beliefs may have shifted in favor of a new, tighter monetary policy regime. This result would be consistent with Taylor's (1999) finding that during the Greenspan era the Federal Reserve responded more strongly to inflation than in previous periods.

Although this ex-post metric raises a warning flag, we judge that the conditional projections of interventions in January and April 1994 are reliable. The judgment call is based on three considerations. First, a change in beliefs about regime is necessary but not sufficient for undermining the reliability of the linear approximation. Second, as Appendix D shows, the responses to policy disturbances are extremely stable across estimation periods ending in September 1990, January 1994, and April 1994. If changes in beliefs generated quantitatively important shifts in decision rules, one would expect to observe greater instability in the response

Variable	$\eta_P^*(94:4,48)$ Constant R^f	$\eta_P^*(94:4,48)$ Actual R^f
y	0.03	-0.28
P	0.05	-0.54
U	-0.01	0.05

functions. Finally, the large η^* 's arise from only a few positive interventions larger than one standard deviation and there is no clear sign pattern in the interventions that might systematically shift beliefs. These observations underscore the bias that can arise from mechanically applying the metric without factoring in the dynamic patterns an intervention implies.[22]

4.   Concluding Remarks

Our framework unifies identification and forecasting. Identifying exogenous variation in policy is essential for addressing counterfactual questions. A careful examination of both in-sample and out-of-sample fits and the model's stability allow the econometrician to assess the model's statistical properties. We showed how the approach can address practical counterfactual policy questions. The approach frames answers in probabilistic terms, quantifying uncertainty about the model and about future exogenous events.

This is a framework for interpreting actual central bank behavior. We illustrated the approach with an identified VAR model of U.S. monetary policy. We probed the range of policy interventions for which an identified linear model is likely to be reliable. We checked that this range of policies is consistent with an unchanged policy regime and showed that it is sufficiently rich to address questions considered by central banks at regular policy meetings. The interventions studied matter: they can shift the projected paths and probability distributions of macro variables in economically meaningful ways.

In line with recent work, the paper has not found compelling evidence that Federal Reserve behavior has been unstable over the past 40 years. Although some readers may dispute our inference that the Fed has operated in approximately one regime since 1959, this does not diminish the usefulness of our approach. Researchers can simply estimate the model over any sample period they deem constitutes a single regime, and then apply our approach to compute linear projections. The approach assesses the projections' reliability conditional on policy

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[22] For the 4 years preceding September 1990, the example in section 3.2, the ex-post metrics for the three variables are below 0.5 in absolute value.

regime: if two researchers can agree that a data sample constitutes a single regime, they should also agree on whether a linear projection is reliable.

Our methodology has broad applicability. We have stressed its usefulness for the kinds of practical analyses central banks conduct. Closely related techniques are being used to study the behavior of the central bank of Sweden as it implements inflation targeting [Jansson and Vredin (2000)]. Our method also yields insights about the plausibility of answers to counterfactual questions. Hamilton and Herrera (2000) apply the methodology to assess the plausibility of Bernanke, Gertler, and Watson's (1997) counterfactual analysis of the Federal Reserve's response to oil price shocks.

Table 1. Structure of Contemporaneous Variables

Money demand

$$a_1 M2 + a_2 R^f + a_3 y + a_4 P = \varepsilon_{MD}$$

Monetary policy

$$a_5 R^f + a_6 M2 + a_7 CP = \varepsilon_P$$

Information sector

$$a_8 CP + a_9 M2 + a_{10} R^f + a_{11} y + a_{12} P + a_{13} U = \varepsilon_I$$

Production sector

This subsystem is arranged in the lower-triangular order y, P, and U.

Table 2. Maximum Likelihood Estimates of Contemporaneous Coefficients

Money demand

$$310.33\, M2 \;+\; 161.89\, R^f \;-\; 28.47\, y \;+\; 6.84\, P = \varepsilon_{MD}$$

$(49.65, 434.46) \quad (45.33, 184.55) \quad (-36.44, -9.11) \quad (-23.95, 36.95)$

Monetary policy

$$100.66\, R^f \;-\; 336.84\, M2 \;-\; 3.10\, CP \;=\; \varepsilon_P$$

$(-12.08, 169.12) \quad (-444.63, -40.42) \quad (-11.29, 4.90)$

Information

$$49.84\, CP \;-\; 35.84\, M2 \;+\; 20.63\, R^f \;-\; 7.09\, y \;-\; 41.64\, P \;+\; 48.43\, U \;=\; \varepsilon_I$$

$(45.85, 50.83) \quad (-65.95, 36.56) \quad (6.60, 51.59) \quad (-19.57, 3.40) \quad (-65.79, -15.82) \quad (21.31, 85.24)$

68 percent probability intervals for the maximum likelihood estimates are reported in parentheses. Those intervals are based on exact finite-sample results computed by a Gibbs sampler algorithm with 360,000 Monte Carlo draws. See Waggoner and Zha (1998) for details.

Table 3. Correlations Among Exogenous Disturbances

Maximum likelihood estimates and 68% probability intervals, in parentheses, based on 360,000 draws from the posterior distribution of the model coefficients

	ε_P	ε_{MD}	ε_I	ε_y	ε_{CPI}	ε_U
ε_P	1.0 (1,1)					
ε_{MD}	.013 (-.048,.052)	1.0 (1,1)				
ε_I	-.016 (-.064,.032)	-.015 (-.060,.032)	1.0 (1,1)			
ε_y	.020 (-.056,.084)	-.001 (-.052,.044)	.003 (-.048,.044)	1.0 (1,1)		
ε_{CPI}	.042 (-.004,.052)	-.0023 (-.056,.044)	-.009 (-.060,.036)	-.002 (-.052,.040)	1.0 (1,1)	
ε_U	-.041 (-.140,.068)	-.192 (-.192,-088)	.001 (-.048,.044)	.001 (-.052,.044)	.002 (-.048,.044)	1.0 (1,1)

(grouped into bins of size .002 to keep storage demands manageable)

Table 4. Out-of-Sample Forecasts Conditional on Actual Path of *R*

Forecasts assume the path of R^f is produced by exogenous policy actions only
Forecasts from October 1990 to September 1994

Maximum likelihood estimates; annual average growth rates or percentage points.

	R^f	*y* forecast (actual)		*π* forecast (actual)		*U* forecast (actual)	
1990	8.1	1.4	(1.2)	5.4	(5.4)	5.6	(5.6)
1991	5.7	1.4	(-0.9)	6.0	(4.2)	6.7	(6.9)
1992	3.5	7.8	(2.7)	7.2	(3.0)	6.0	(7.5)
1993	3.0	15.0	(2.3)	12.7	(3.0)	3.5	(6.9)

Table 5. Joint and Marginal Probabilities Conditional on Alternative Policies

Outcomes Based on Out-of-Sample Forecasts from September 1990.

"Tighter" policy raises R^f to 8.70% in October and to 8.95% in November 1990-January 1991 and is produced by the sequence of exogenous actions $\tilde{\varepsilon}_P = (2.3, 1.7, 1.0, 0.9)$.

"Actual R^f" sets R^f at 8.11% in October, 7.81% in November, 7.31% in December, 6.91% in January 1991 and is produced by the sequence of exogenous actions $\tilde{\varepsilon}_P = (0.5, 0.1, -0.7, -0.7)$.

	Tighter	Actual R^f
$P(low\ \pi\ in\ 1992)$	.67	.47
$P(low\ \pi\ in\ 1993)$	.66	.46
$P(low\ \pi\ in\ 1992\ and\ 1993)$	.57	.36
$P(recession\ in\ 1991)$	.53	.27
$P(recession\ in\ 1992)$	.12	.05
$P(recession\ in\ 1993)$	.05	.06
$P(recession\ and\ low\ \pi)$	.33	.11
$P(recession\ and\ high\ \pi)$	.25	.22
$P(no\ recession\ and\ low\ \pi)$	.24	.25
$P(no\ recession\ and\ high\ \pi)$	.18	.42

$P(recession)$ is the probability of negative real GDP growth in 1991 or 1992 or 1993.

$P(low\ \pi)$ is the probability of inflation below 5½ percent in 1992 and 1993.

$P(recession\ and\ low\ \pi)$ is the probability of negative real GDP growth in 1991 or 1992 or 1993 *and* inflation below 5½ percent in 1992 and 1993.

Figure 1. Responses to an Exogenous Monetary Policy Contraction: Monthly

Maximum likelihood estimates (solid) and 68% probability bands (dashed). Log levels or percentage points.
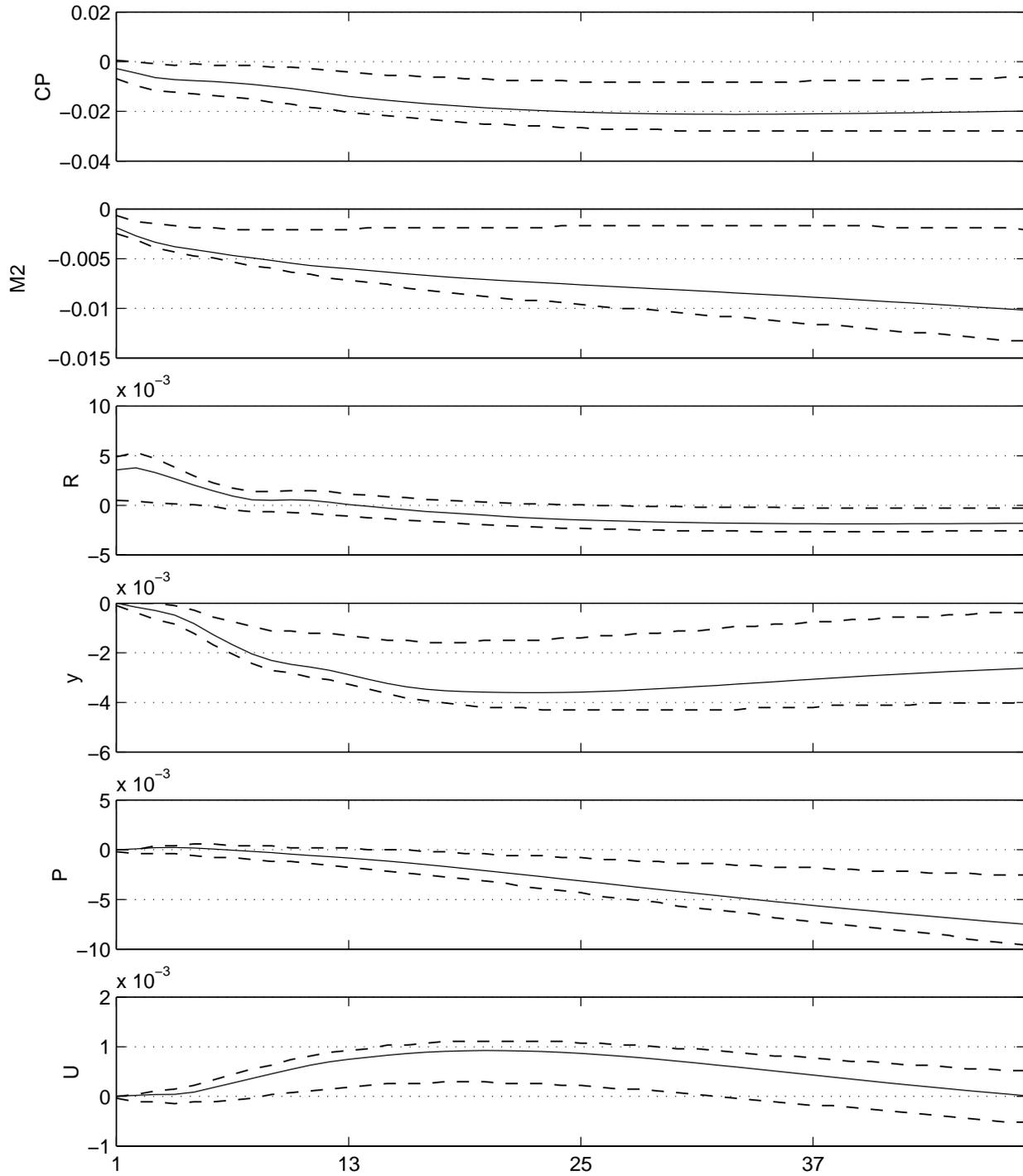
Figure 2. Full Model Responses Estimated Over Various Sub-Periods

Impulse response functions estimated over various sub-periods. Solid line is entire sample period, 1959:1-1998:9; dashed line is 1959:1-1979:9; dashed/dotted line is 1959:1-1982:12; dotted line is 1959:1-1998:9 with 1979:10-1982:12 eliminated.

ε_x denotes exogenous disturbance; x = I (commodity prices – information sector), = P (monetary policy), = MD (money demand), = y (GDP – goods sector), = CPI (price level – goods sector), U = (unemployment – goods sector)
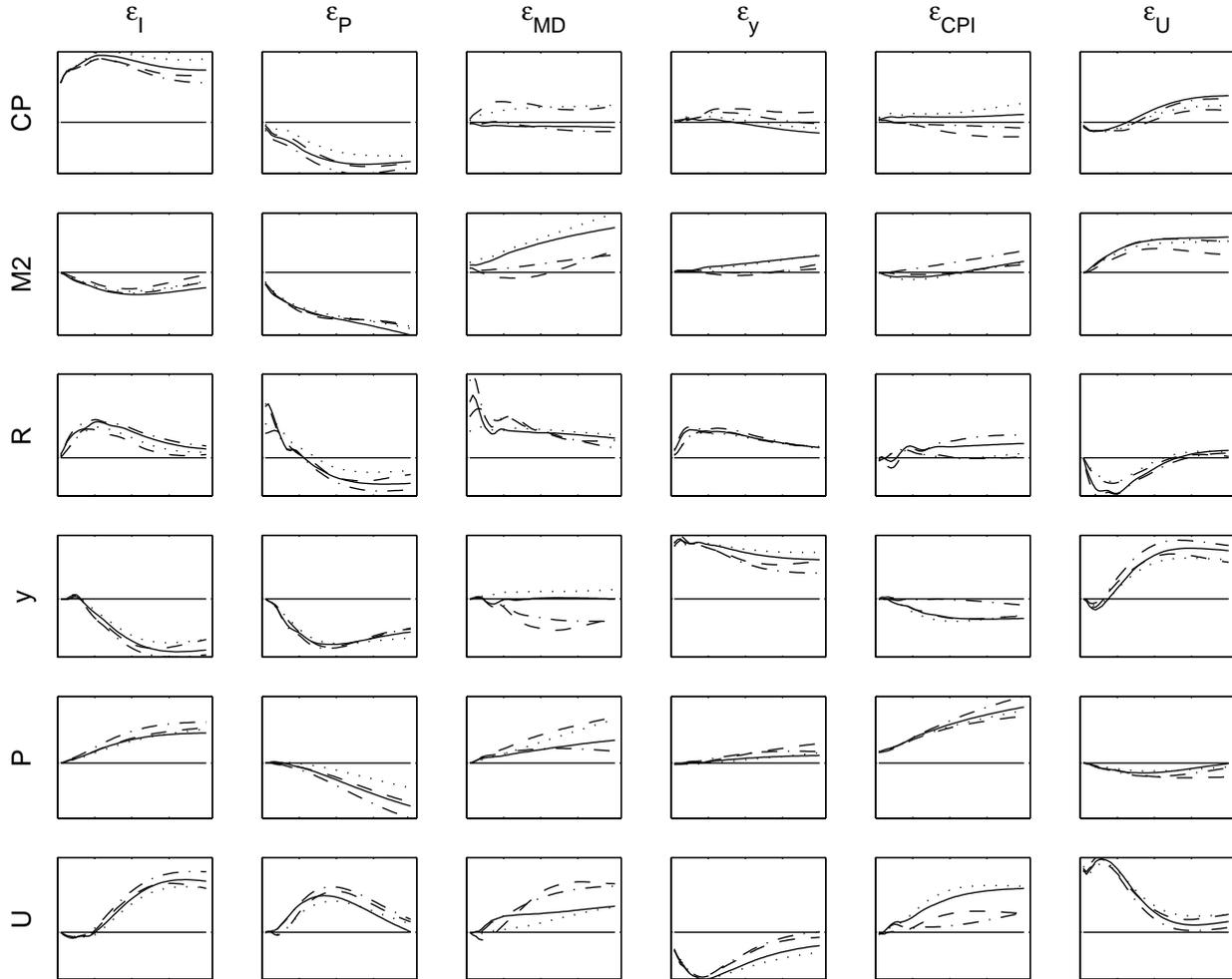
Figure 3. Forecasts Conditional on Actual Path of Funds Rate

Actual (solid) and out-of-sample forecast (dashed). Out-of-sample forecasts conditional on actual path of the federal funds rate from October 1990 to January 1991. Forecasts include maximum likelihood estimates (dashed-dot) and 68% probability bands (dashed). Annual average growth rates or percentage points.
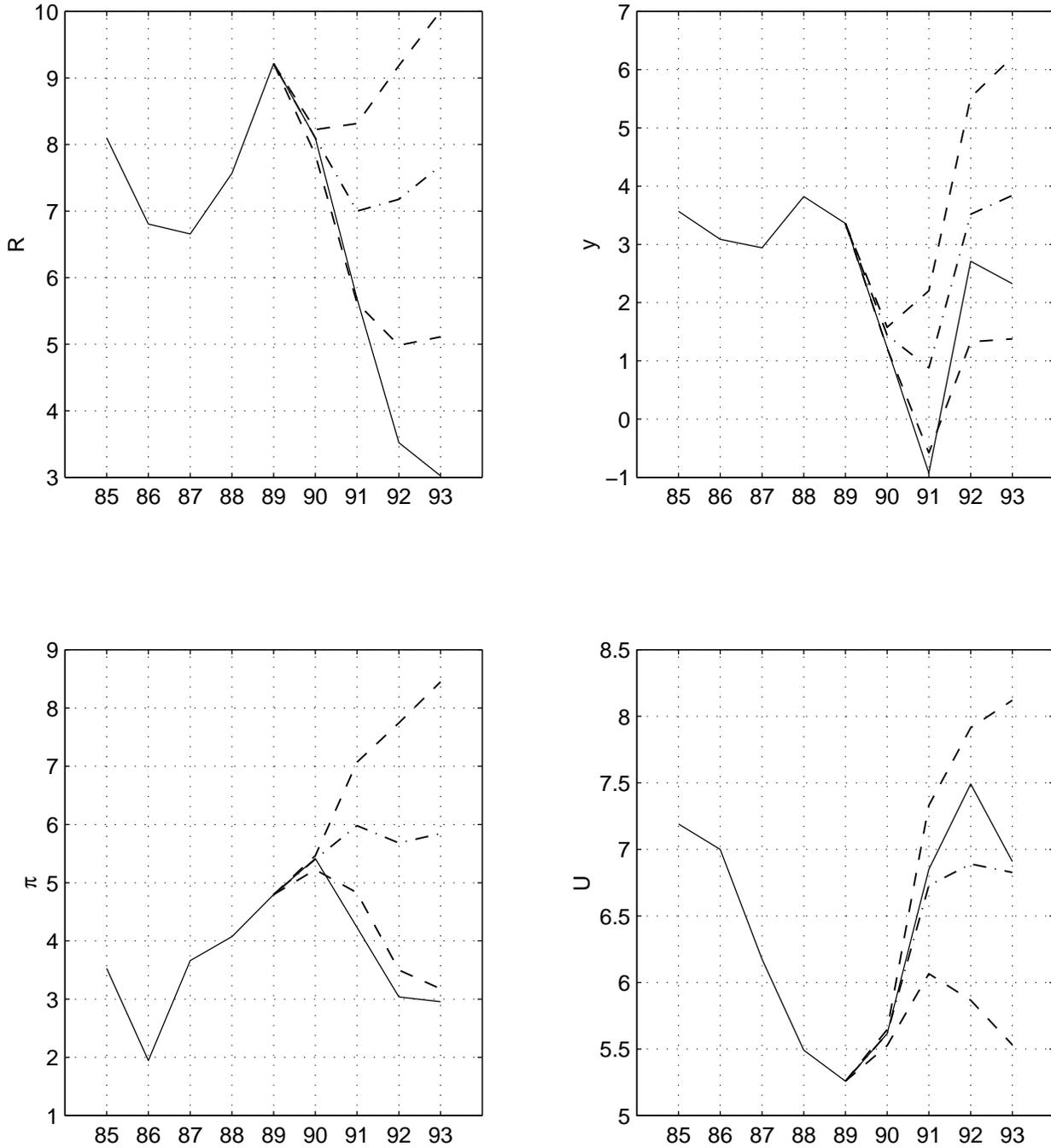
Figure 4. Forecasts Conditional on Tighter Policy

Actual (solid) and out-of-sample forecast (dashed). Out-of-sample forecasts conditional on tighter policy, raising the federal funds rate 50 basis points in October 1990 (to 8.70%) and another 25 basis points over November 1990 to January 1991 (to 8.95%). Forecasts include maximum likelihood estimates (dashed-dot) and 68% probability bands (dashed). Annual average growth rates or percentage points.
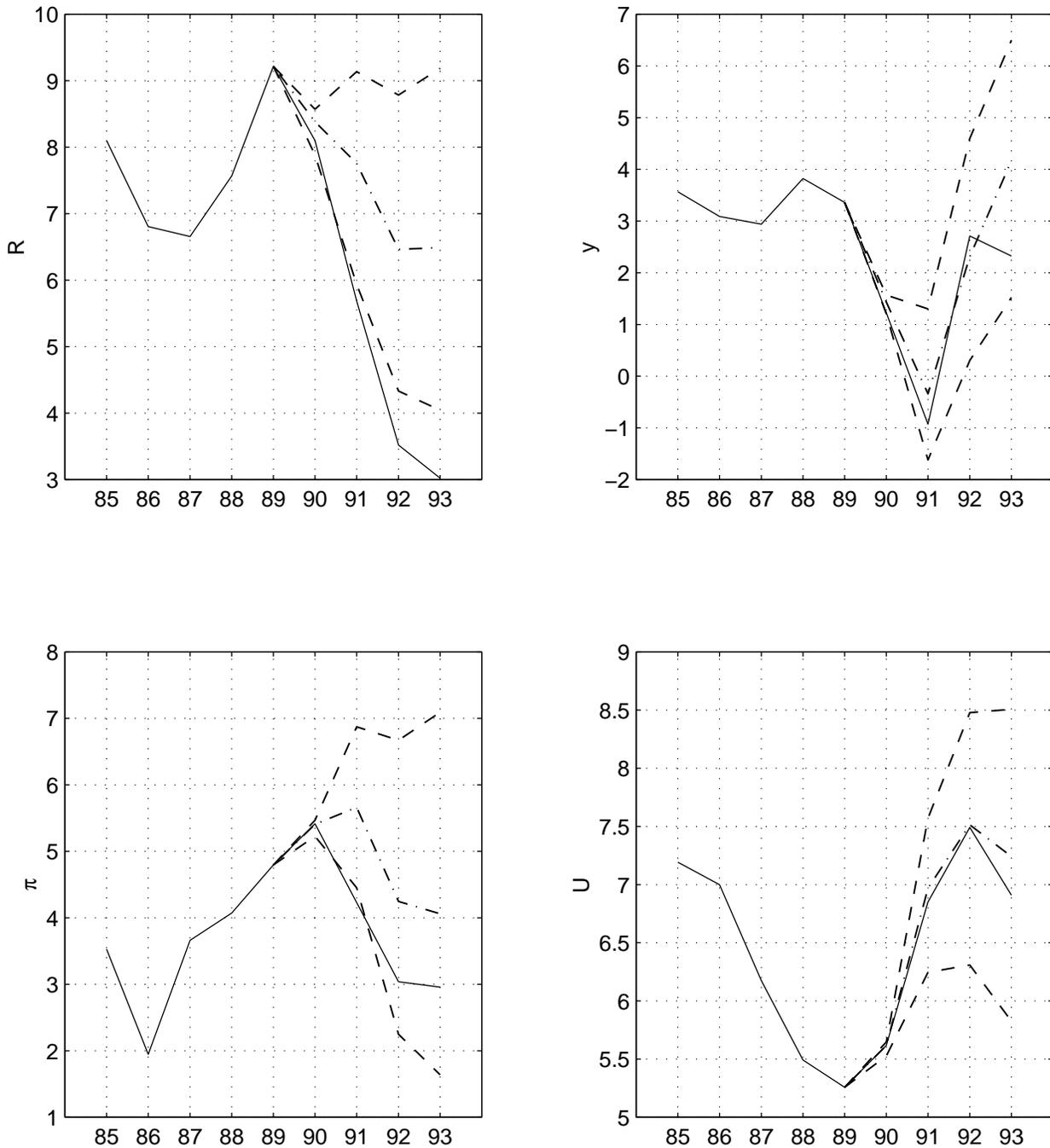
Figure 5. Forecasts Conditional on Constant and Actual Funds Rate: 94:2

Actual (solid) and out-of-sample forecasts conditional on a constant funds rate (+) and on the actual path of the funds rate (*) from February to May 1994. Constant rate is 3.00%; actual path is 3.25%, 3.34%, 3.56%, 4.01%. Annual average growth rates or percentage points.
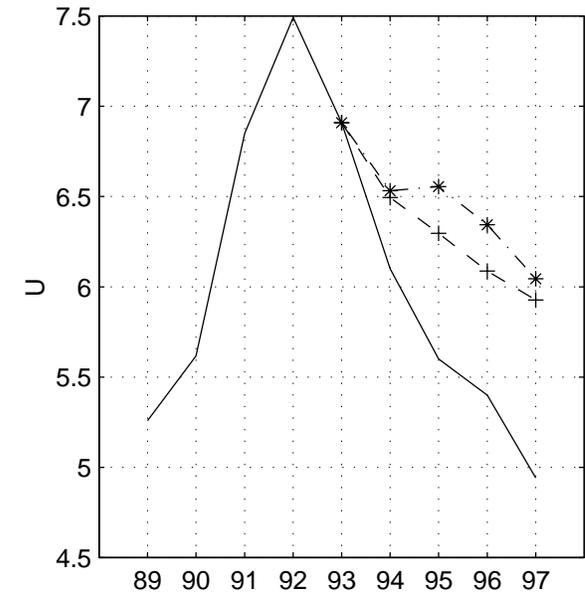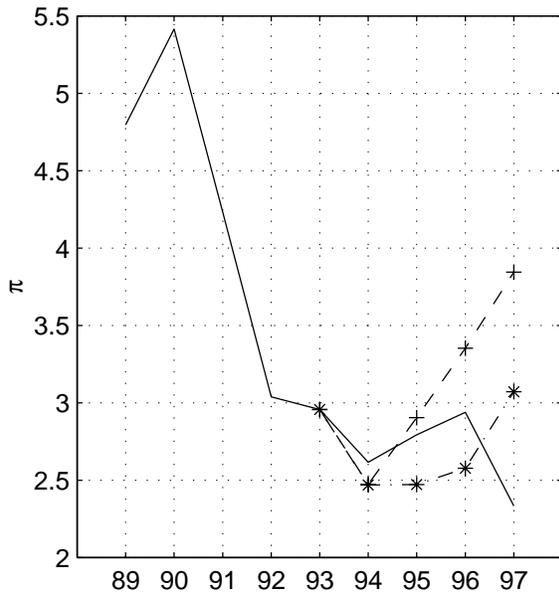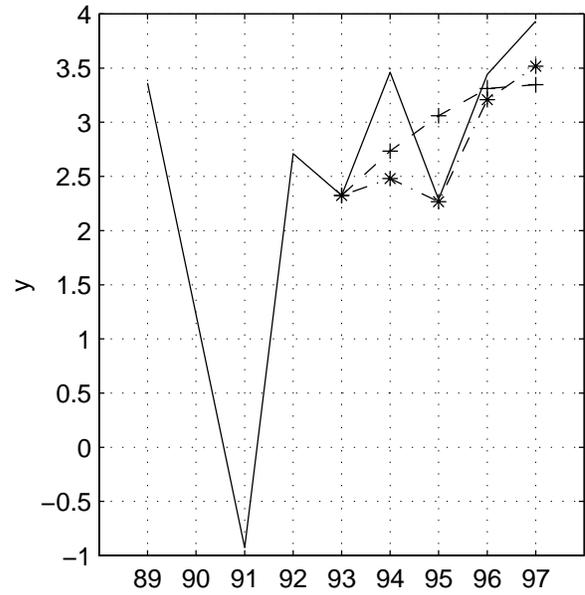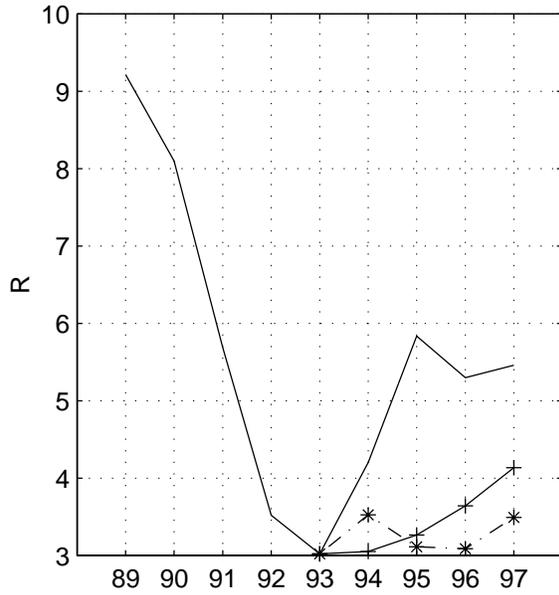
Figure 6. Forecasts Conditional on Constant and Actual Funds Rate: 94:5

Actual (solid line) and out-of-sample forecasts conditional on a constant funds rate (+) and on the actual path of the funds rate (*) from May to August 1994. Constant rate is 3.75%; actual path is 4.01%, 4.25%, 4.26%, 4.47%. Annual average growth rates or percentage points.
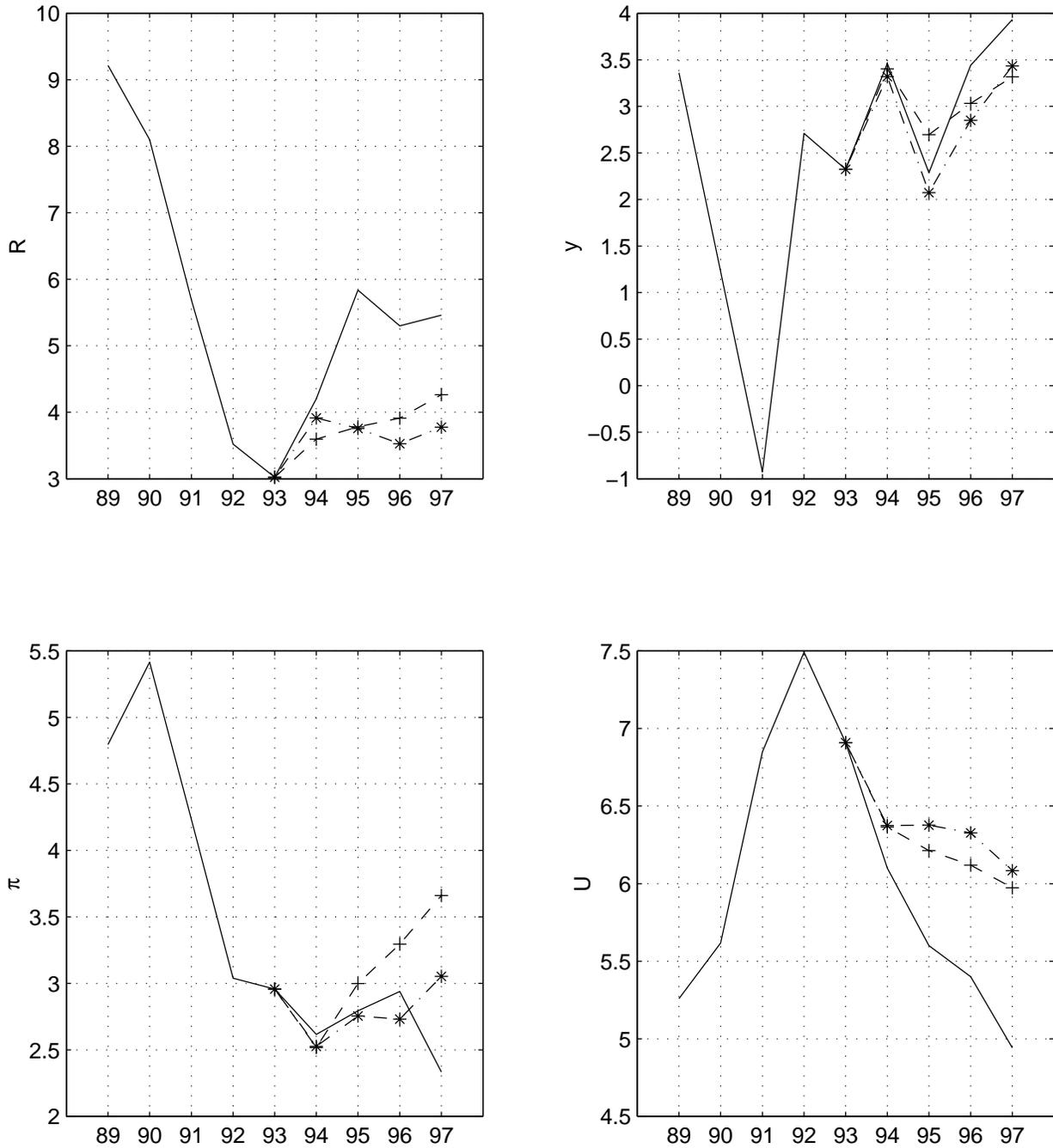
Appendix A: Estimation

Litterman (1986) shows that reduced-form Bayesian VARs forecast out-of-sample better than other time series methods or commercial models do. We adopt the Bayesian procedures for identified VARs that Sims and Zha (1998a) develop.

There are two layers to the prior information. The first layer is a reference prior that reduces the sampling error that produces erratic results in large models under a diffuse prior. We postulate a joint normal prior, with a diagonal covariance matrix on the elements of A_0 that the identification does not constrain to equal zero. We then specify a joint normal prior on all the coefficients in A_s, for $s > 0$, conditional on A_0. To implement Litterman's random walk specification, we assume the conditional mean of $A_1|A_0$ is A_0, while the conditional mean of $A_s|A_0$, for $s > 1$, is zero. The prior standard deviations of the elements of the A_s matrices shrink as s increases, dampening unreasonably large lagged coefficients.

The second layer of the prior corrects the overfitting problem endemic to VARs. Because the number of parameters in A_s grows with the square of the number of variables, model (3) tends to fit the data unrealistically well in sample but can fail badly when projecting post-sample. This manifests as deterministic components of the model explaining implausibly large fractions of the observed variation in the data. To express our belief that such outcomes are implausible, we add to the reference prior parameters that control beliefs about the number of unit roots and the number of cointegrating relationships in the m-variable system. The data determine the exact number of unit roots and cointegrating vectors. In Sims and Zha's (1998a) notation, the tightness of the prior is set as $\lambda_0 = 0.6$, $\lambda_1 = 0.1$, $\lambda_2 = \lambda_3 = 1$, $\lambda_4 = 0.1$, $\mu_5 = 5$, and $\mu_6 = 5$.[23]

The prior accommodates an idea Hall (1996) proposed to relax the dogmatic zero restrictions typical in simultaneous equations models. Instead of restricting a coefficient to be exactly zero or freely estimating it, we specify a "soft zero" restriction. The restriction shrinks a coefficient's prior standard deviation around a zero mean without forcing the coefficient to be zero.

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[23] These priors do not change the character of the results [Robertson and Tallman (2000)].

Estimation and inference explore the overall shape of the likelihood surface for A_0 and A_s.

Maximum likelihood estimates occur at the peak of the posterior density.

Appendix B: Data

The data, from 1959:1 to 1998:9, are collected from the Bureau of Economic Analysis, the Department of Commerce unless otherwise stated.

Federal Funds Rate: effective rate, monthly average. Source: Board of Governors of the Federal Reserve System (BOG).

M2: M2 money stock, seasonally adjusted, billions of dollars. Source: BOG.

CPI: consumer price index for urban consumers (CPI-U), seasonally adjusted.

GDP: Real GDP, seasonally adjusted, billions of chain 1992 dollars. Monthly real GDP is interpolated using the procedure described in Leeper, Sims and Zha (1996).

Unemployment: civilian unemployment rate (16 and over), seasonally adjusted. Source: Bureau of Labor Statistics.

Commodity Prices: International Monetary Fund's index of world commodity prices. Source: *International Financial Statistics*.

Appendix C: Out-of-Sample Forecast Performance

To have confidence in the model's projections conditional on alternative future policies, it is important to evaluate the model's out-of-sample forecast performance. This appendix compares the annual forecast performance of our model under the Sims and Zha (1998a) prior to forecasts with a VAR with a diffuse prior and a VAR with Litterman's (1986) prior.

Table 6 lists the root-mean-square errors (RMSEs) for unconditional forecasts over horizons of one to four years using the three estimation methods. The models are estimated recursively using data from January 1959 up to the beginning of each forecast period. The RMSEs are computed from forecasts over four years beginning in December 1979 and running through September 1994.

The table shows that, except in the case of *M2*, the structural model outperforms the diffuse-prior VAR and the Litterman-prior VAR. The biggest improvement occurs in the inflation rate and the federal funds rate. This is encouraging for our purposes: inflation is the Fed's primary concern and the funds rate is the policy instrument.[24]

The forecasts compare favorably to the ones that Cecchetti (1995) computes and to the DRI forecasts of inflation that Cecchetti reproduces. Kohn (1995) reports that the Federal Reserve's one-year forecasts of inflation have RMSEs of from 1 to 1¼ percentage points. Our structural model's RMSE for one-year inflation forecasts is .95 percentage points. The model's compactness relative to the Fed's models makes its performance especially remarkable.

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[24] Robertson and Tallman (1999) study the forecast performance of this model relative to others.

# Table 6. Summary of Out-of-Sample Forecast Performance

The following root-mean-square errors are computed using forecast horizons from 1 to 4 years. Forecasts begin in December 1979 and run through September 1994, 4 years before the end of the sample.

Years ahead		1	2	3	4
Federal Funds Rate (R^f)	Diffuse Prior	2.534	6.478	9.443	12.074
	Litterman	1.945	4.212	5.310	5.747
	Sims-Zha	1.572	2.531	2.784	3.287
Real GDP Growth (y)	Diffuse Prior	2.142	3.038	2.831	4.067
	Litterman	1.613	2.022	2.283	2.561
	Sims-Zha	1.384	2.324	2.013	1.706
Inflation (π)	Diffuse Prior	1.199	3.397	4.987	4.571
	Litterman	1.138	2.959	4.201	4.532
	Sims-Zha	0.951	2.066	2.825	3.238
Unemployment (U)	Diffuse Prior	0.760	1.794	2.070	2.961
	Litterman	0.535	1.109	1.338	2.116
	Sims-Zha	0.537	1.241	1.525	1.832
Commodity Prices (CP)	Diffuse Prior	10.599	21.103	18.504	16.481
	Litterman	8.593	16.829	15.633	13.063
	Sims-Zha	7.316	11.279	10.461	10.969
Money Stock (M2)	Diffuse Prior	1.792	3.455	3.352	3.187
	Litterman	1.553	2.643	2.369	2.638
	Sims-Zha	1.534	3.103	3.885	4.359

Appendix D: Figure 7

Impulse response functions estimated over various sub-periods. Solid line is entire sample period, 1959:1-1998:9; dashed line is 1959:1-1990:9; dashed/dotted line is 1959:1-1994:1; dotted line is 1959:1-1994:4.

ε_X denotes exogenous disturbance; X = I (commodity prices – information sector), = P (monetary policy), = MD (money demand), = y (GDP – goods sector), = CPI (price level – goods sector), U = (unemployment – goods sector)
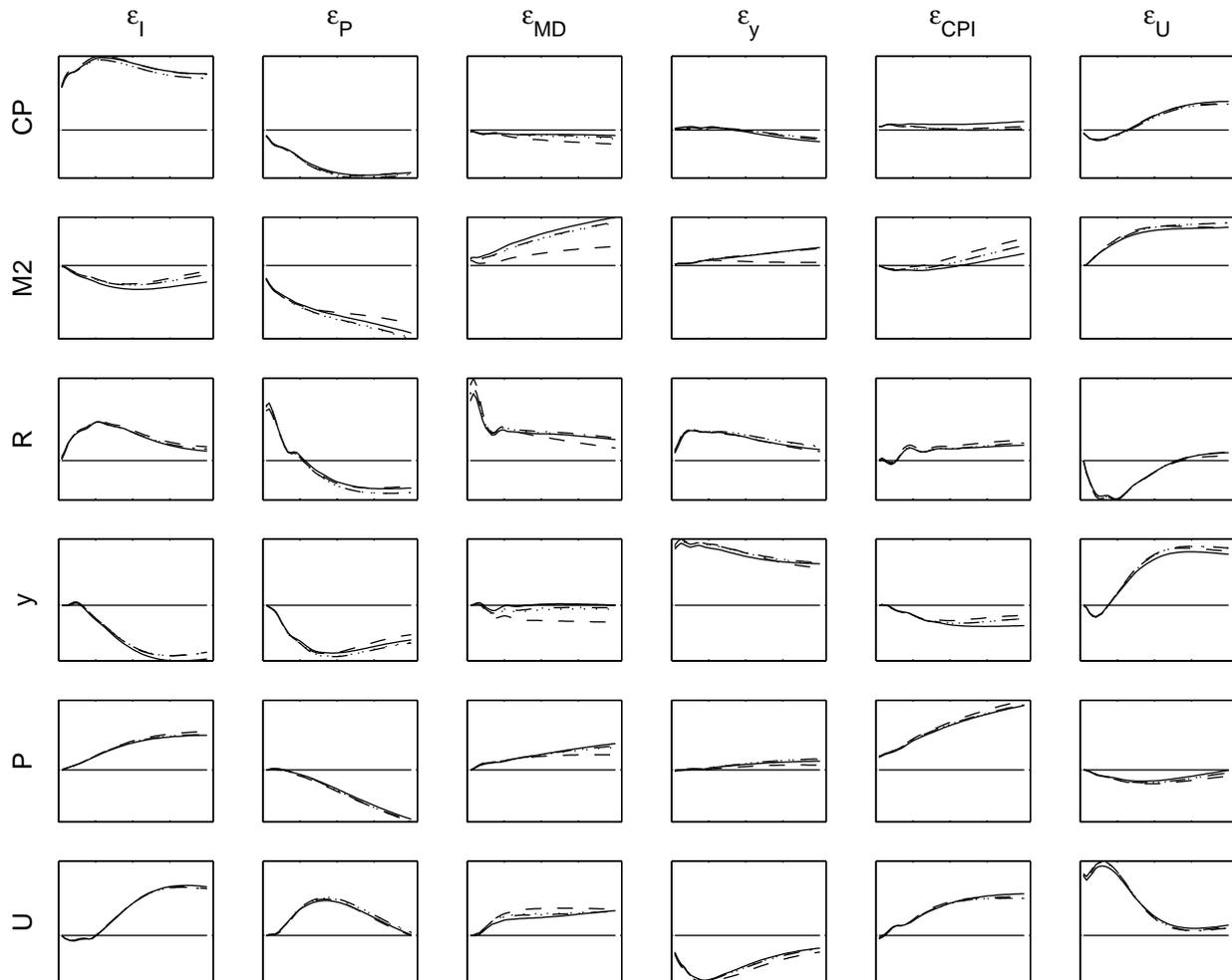
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