ECONOMIC RESEARCH
FEDERAL RESERVE BANK OF ST. LOUIS

WORKING PAPER SERIES

# Modeling Volcker as a Non-Absorbing State: Agnostic Identification of a Markov-Switching VAR

| | |
|---|---|
| **Authors** | Michael T. Owyang |
| **Working Paper Number** | 2002-018A |
| **Creation Date** | August 2002 |
| **Citable Link** | https://doi.org/10.20955/wp.2002.018 |
| **Suggested Citation** | Owyang, M.T., 2002; Modeling Volcker as a Non-Absorbing State: Agnostic Identification of a Markov-Switching VAR, Federal Reserve Bank of St. Louis Working Paper 2002-018. URL https://doi.org/10.20955/wp.2002.018 |

Federal Reserve Bank of St. Louis, Research Division, P.O. Box 442, St. Louis, MO 63166

# MODELING VOLCKER AS A NON-ABSORBING STATE: AGNOSTIC IDENTIFICATION OF A MARKOV-SWITCHING VAR

**Michael T. Owyang**

Federal Reserve Bank of St. Louis*
411 Locust Street
St. Louis, MO 63102, U.S.A.

July 26, 2002

## ABSTRACT

Recently, models of monetary policy have been constructed to include structural breaks to account for changes in policymaker preferences or operating procedures. These models typically assume that when changes occur, they happen once and for all. In this paper, we allow the policymaker and the economy to switch freely between regimes. We find that not only does the nature and effect of innovations to monetary policy change, but switching the policy rule and the economy's subsequent response can in and of itself alter the path of the economy. We find the switch itself can generate disinflationary dynamics.

**Keywords:** Monetary policy, regime switching, vector autoregression

**JEL classification:** E52, C51

---

* The views expressed are those of the author and do not necessarily reflect official positions of the Federal Reserve Bank of St. Louis or the Federal Reserve System.

# 1 Introduction

Attempting to model the effects of exogenous, unanticipated innovations to monetary policy, a vast literature has developed around restricted vector autoregressions (structural VARs). The endogenous component of monetary policy has been examined more formally in the Taylor rule literature beginning with Taylor (1993) and more recently Clarida, Gali and Gertler (1999, 2000). These papers explicitly model the Federal Reserve's reaction to non-policy variables. They focus on Fed behavior and require significant assumptions regarding the impact of monetary policy on prices and output.

We are interested in the effects of both the exogenous (shocks) and endogenous (response) components of monetary policy. Specifically, we ask two questions: (1) Does monetary policy, both shocks and responses, change over time and (2) does the effect and effectiveness of monetary policy change over time? To answer these questions, we consider a model in which the economy is subject to structural changes. A number of studies have also explored the possibility that the economy experiences regime switches over the sample period. Recent studies have identified differences in the effect and behavior of monetary policy around structural breaks usually corresponding to changes in Federal Reserve leadership or changes in operating procedures (e.g., Boivin and Giannoni (2002), Dennis (2001), and Hanson (2001)). This approach typically assumes that only a single break exists–equivalent to an underlying Markov process with an absorbing state. A standard assumption is that Volcker forever changed Fed objectives to a more stable, inflation averse policy rule. The drawback of this approach is that incorrectly choosing the break dates or, perhaps more importantly, the number of breaks can lead to a misspecification that leaves the conclusions about the effects of policy in question.

An alternative is to estimate the model assuming an underlying Markov-switching process (Bernanke and Mihov (1998), Sims (1999, 2001), Owyang (2001), Owyang and Ramey (2001), and Sims and Zha (2002)). The breaks are determined by estimating the posterior probabilities for a hidden state variable in each period. In theory, the two methods of identifying the breakpoint are equivalent if the econometrician has some (correct) prior knowledge of the break process. We choose the latter approach to ensure that we do not misspecify the transition probabilities in the Markov process and average the effects of policy over regimes that go unidentified. In essence, we allow shifts in policy to be temporary, returning to a baseline state after some stability is restored.

This paper incorporates Markov-switching into the estimation of the covariance and coefficient matrices of the monetary VAR. We employ an agnostic technique to identify the monetary policy shock. In a typical structural VAR, restrictions are imposed on the covariance matrices (see Christiano, Eichenbaum and Evans (1999) and Bernanke and Mihov (1998) for reviews of the literature) and rely on *ex ante* assumptions about the relationships between components of the economic system. An alternative to the restricted VAR approach is employed by Uhlig (1999, 2001), Faust (1998) and Roush (2002). These papers estimate reduced-form, unrestricted VARs and impose restrictions on the shapes (magnitudes, directions, and durations) of the impulse responses. Shocks are then sampled and characterized based on the set of restrictions.

We argue that it is not clear that standard identification schemes are invariant across regimes. That is, the exclusion restrictions sometimes employed in structural VARs may not be consistent across regimes. To assume that they are may lead to misidentification of the monetary shock and confound any resulting inference. The model offers two innovations to the literature. First, we examine the combination of switching in both the endogenous and exogenous components of monetary policy–that is, we allow both the implied policy rule and the covariance matrix generating the monetary shocks to switch. Second, we consider the implication of shifts not only to policy but to the economy's response to both endogenous and exogenous policy.

The outline of this paper is as follows: Section 2 discusses the model with regime-switching. Section 3 reviews the characterization of the monetary shock. Section 4 discusses the estimation results and compares the impulse responses of the two states. Section 5 discusses the policy implications of the model and compares the results to VAR estimates over sample periods with a fixed break in October, 1979 or in November, 1982. In Section 6, we discuss the transition dynamics between states of the VAR and show that the state change can actually mimic a policy tightening. Section 7 concludes.

## 2 Model

We consider a reduced-form VAR representation that *ex ante* imposes no structural relationships between variables and allows both the monetary innovations and the economy's response to vary governed by an underlying Markov process. We incorporate variables commonly associated with a monetary policymaker's objective function (i.e., output and inflation) and assume that the policymaker employs the Federal Funds rate as a policy instrument. In addition, we include a short-term interest rate that reflects the agents' short-run inflationary expectations and short-run risk premium. We do not explicitly write out a Taylor rule for the policymaker but implicitly allow the model to freely estimate the rule governing the policymaker's rule governing the Federal Funds target.

A typical regime-switching structural VAR can be written as

$$A_{0,S_t} y_t = \sum_{j=1}^{k} A_{j,S_t} y_{t-j} + \nu_t, \tag{1}$$

where $y_t$ is the period $t$ vector of $n$ variables of interest, $k$ is the number of lags and $\nu_t \sim N(0, I)$ assumes that the system shocks are uncorrelated. The hidden state variable $S_t = \{1, 2\}$ governs the coefficient and covariance matrices in the model and has transition matrix $\Pi$. In this paper, we will assume that there are two states of the world–in principle the number of states can be expanded to match the number that maximizes the criteria but for ease of estimation, we limit the model to a two-state system.[1]

We rewrite (1) as a reduced-form Markov-switching VAR (MSVAR)

$$y_t = \sum_{j=1}^{k} \beta_{j,S_t} y_{t-j} + \varepsilon_{S_t}, \tag{2}$$

where $\varepsilon_{S_t} \sim N(0, \Sigma_{St})$ and $\Sigma_{St} = A_{0,S_t}^{-1} A_{0,S_t}^{-1\prime}$. Exact identification of $A_{0,S_t}$ requires $\frac{n(n+1)}{2}$ restrictions. A few restrictions in a structural VAR can be drawn from economic theory; often, there is insufficient theory to provide exact identification and additional (sometimes *ad hoc*) assumptions must be imposed. In our alternative approach, we estimate the reduced-form model (2) and impose no ex ante restrictions on either the coefficient or the covariance matrix. We identify the distribution to the column of $A_{0,S_t}^{-1\prime}$ attributable to monetary policy.

To estimate (2), we are interested in characterizing the joint distribution

$$p(\widetilde{\beta}_k, \widetilde{\Sigma}, \tilde{S}_T | \tilde{y}_T), \tag{3}$$

where $\widetilde{\beta}_k = \left\{\beta_1^{[1]}, \beta_2^{[1]}, ..., \beta_k^{[1]} | \beta_1^{[2]}, \beta_2^{[2]}, ..., \beta_k^{[2]}\right\}'$, $\widetilde{\Sigma} = \left\{\Sigma^{[1]}, \Sigma^{[2]}\right\}$, $\tilde{S}_T = \{S_1, S_2, ..., S_T\}'$, and $\tilde{y}_T = \{y_1, y_2, ..., y_T\}'$. To accomplish this, we employ the Gibbs sampler using a Normal-Wishart distribution with a flat prior to generate the parameters $\widetilde{\beta}_k$ and $\widetilde{\Sigma}$. The posterior distributions for the states is a product of Hamilton's (1989, 1990) filter. Details of the estimation procedure can be found in the

[1] Adding Markov switching to the model complicates the estimation algorithm, doubling the number of restrictions on the covariance matrix required for exact identification.

appendix.

Once the joint distribution (3) is characterized, the state-dependent impulse response to any vector innovation $\hat{\alpha}$ can be defined as

$$\Delta y_{t+j}|S_t = \Gamma^j_{S_t}\overline{\alpha}, \tag{4}$$

where $\overline{\alpha} = \left[\hat{\alpha}', 0_{1\times n(k-1)}\right]'$ and the $(nk \times nk)$ matrix $\Gamma_{S_t}$

$$\Gamma_{S_t} = \begin{bmatrix} \multicolumn{2}{c}{B'_{S_t}} \\ I_{n(k-1)} & \widetilde{0}_{n(k-1)xn} \end{bmatrix}, \tag{5}$$

where $B_{S_t}$ is the stacked matrix of $\beta_{j,S_t}$, $j = 1, 2, ...k$, $I_m$ is the $m \times m$ identity matrix and $\widetilde{0}_{i\times j}$ is an $i \times j$ matrix of zeros. Since the impulse vector will be characterized by the impulse response, the change in the coefficient matrix resulting from a change in state causes the monetary shock to vary over states.

# 3 Identification of the Monetary Shock

The characterization of a shock to the VAR is merely the decision of how to interpret the vector components of the time series of residuals. Generally, shocks are interpreted by imposing economic structure on either the causal ordering or the covariance matrix. These restrictions imply a basis rotation for the space spanned by the VAR residuals. The orthogonal components of the basis represent model shocks and the component of the residual that lies in the direction of any basis vector represents the magnitude of that shock for that time period. Since we allow the components of the model to switch, standard identification restriction may not hold across regimes. We choose to take an agnostic approach to identifying the distributions for the monetary shocks.[2]

[2] Sims and Zha (2002) point out that imposing restrictions on the contemporaneous coefficient matrix complicates estimation by making the parameter distributions nonstandard.

Instead of imposing economic structure *ex ante*, we define the shock based on its *ex post* impulse response for certain variables. We define the shock in terms of the reduced-form model (2) and not the full model (1). Thus, the identified shocks will have contemporaneous non-policy impacts. Any shock in the space spanned by the residuals that satisfies these restrictions is a candidate shock. The distribution for these candidates converges asymptotically as the distribution for the joint density of the VAR parameters converge. The algorithm for identifying the monetary shock is as follows.

Once the joint distribution for the parameters $p(\widetilde{\beta}_k, \widetilde{\Sigma}, \tilde{S}_T | \tilde{y}_T)$ has converged, a set of candidate parameters $\widehat{\beta}_k$, $\widehat{\Sigma}$ are drawn conditional on each $S_t$. The impulse response, conditional on the state of the economy, $S_t$ to any shock $\hat{a} \in \Re^n$ can then be calculated using (4) and (5). The monetary shock $\hat{a}$ is associated with a restriction matrix of $R$ that is invariant to the state of the economy (e.g., a contractionary monetary policy shock always increases the federal funds rate). $R$ is an $n \times l$ matrix that represents the priors that we impose on the response of model variables to the incidence of a shock $\hat{a}$ out to a horizon $l$. Our identification of monetary policy shocks revolves around selection of the shock $\hat{a}$ that produces an impulse response satisfying the restriction matrix $R$. For example, we can impose the condition that a contractionary monetary policy shock causes the federal funds rate to rise (weakly) for twelve months. We would then impose this condition by assigning positive values to the relevant row of $R$ up through the twelfth column.

Regime-dependent heteroskedasticity implies generating qualifying candidate impulse vectors is state-dependent. We must characterize a state-dependent impulse vector by

Rather than impose these restrictions, we estimate the reduced-form freely. Identification can then be achieved by the agnostic technique which does not require ex ante restrictions on the reduced-form VAR's covariance matrix.

$$\hat{a}_{S_t} = \widehat{A}_{S_t}\alpha,$$

where $\widehat{A}_{S_t}\widehat{A}'_{S_t} = \Sigma_{S_t}$ is the Cholesky decomposition of the state-dependent covariance matrix. Thus, the definition of the impulse vector changes over states. A shift in the coefficient matrix results in a change in the impulse response generating function $\Gamma_{S_t}$. To satisfy the restrictions imposed on the impulse responses, the impulse vector must change when the state of the economy changes. The switch in the covariance matrix changes the set of eligible impulse vectors.[3] We then apply (4) and (5) to generate state-dependent impulse responses and test them against the restriction matrix $R$. Any $\hat{a}_{S_t}$ that satisfies the restrictions on $y_{t+j}$ is stored. Multiple iterations over a single set of sampled model parameters yields $\mathcal{A}_{S_t}(B, \Sigma, R)$.[4] The algorithm generates multiple draws from the posterior joint density for the model parameters; convergence of the model parameters is equivalent to convergence of the space containing the impulse vectors.

We then calculate the historical incidence of the monetary shocks for the sample period. The amount $v_t$ of the shock $\varepsilon_t$ attributable to the identified monetary shock $\hat{a}_{S_t}$. Since the distributions for the monetary shocks, $\mathcal{A}_{S_t}(B, \Sigma, R)$, are state-dependent, we construct the following state-dependent model of the shocks:

$$\varepsilon_t = a\overline{S}_t v_{a,t}, \tag{6}$$

[3] We draw $\alpha$ from the relevant portion of the unit sphere but the rotation into the space that defines the impulse vector changes when the state changes.

[4] Our characterization of the impulse vector space is slightly different from Uhlig (2001). He implicitly assumes that the sign restrictions on the impulse response functions hold out to horizon $l$ and he characterizes the space as $\mathcal{A}(B, \Sigma, l)$. Since we will impose long run restrictions, it is beneficial to denote the impulse vector space as dependent on a restriction matrix $R$.

where $a = \begin{bmatrix} \widehat{a}_{S_t=1} & \widehat{a}_{S_t=2} \end{bmatrix}$, $\overline{S}_t = \begin{bmatrix} I_{S_t=1} \\ I_{S_t=2} \end{bmatrix}$ and $I_{S_t=i} = 1$ if $S_t = i$. We will adopt the assumption that the shocks are uncorrelated and effect-orthogonal. The scalar multipliers $v_{a,t} = b'\varepsilon_t$ can be calculated by realizing that the vector $b'$ is the normalized eigenvector associated with the zero eigenvalue of $\Sigma - a_i a_i'$. The vector of states $\widetilde{S}_T$ is generated previously from the estimation of the reduced form VAR.

# 4 Empirical Results

The data used in the VAR are the federal funds rate (FFR), the one year Treasury rate (10YR), the composite index of coincident indicators (CII) and the change in the personal consumption expenditure price chain index (PCE) from 1959:01 to 2001:06. Data are taken from the Federal Reserve Board (FFR and 1YR), the Conference Board (CII), and the Bureau of Economic Analysis (PCE). We specify flat priors (see appendix) for the conditional distributions governing the selection of the model parameters.

We estimate the two-state Markov process governing the model (2) and generate the distributions for the monetary policy shocks implied by the restriction matrix $R$ shown in Table 1.[5] The identification scheme puts no restriction on output at any horizon but requires the federal funds rate to increase and the inflation rate to decrease immediately following the contractionary shock. We assume the expansionary shock is the exact negative of the contractionary shock. The algorithm generating the joint distribution $p(\widetilde{\beta}_k, \widetilde{\Sigma}, \tilde{S}_T|\tilde{y}_T)$ is run 5000 times, discarding the first 3000 iterations.

[5] A three state version of the model was estimated but the posterior probabilities of a third stae in any period was negligible. Sims (2001) and Sims and Zha (2002) argue that more than two states are required to achieve fit in their models. These models, however, restrict switching to the policy rule only.

### 4.1 Responses to the Monetary Policy Shock

For each of the 2000 saved iterations, 2000 policy shock candidates are drawn. Conditional on $S_t = \{1, 2\}$ and the combination of $\widehat{B}$ and $\widehat{\Sigma}$, a candidate shock $\hat{a}_{S_t}$ is drawn and saved if it satisfies the restriction matrix $R$. The distributions for the monetary policy shock $\hat{a}_{S_t}$ for each $S_t$ are given in Table 2.

The primitive VAR monetary shock, the component of $\nu_t$ associated with the policy instrument, is an innovation to the federal funds rate. Here, the impulse vector includes the contemporaneous response to non-policy variables into the reduced-form shock. The impulse vector has a non-zero impact effect for non-policy variables. For example, we constrain inflation to be non-positive six periods subsequent to a contractionary monetary innovation.[6]

Figure 1 about here

Figure 1 shows the response of the identified monetary shock for the State 1 out to 60 months. The impact of the contractionary shock on the federal funds rate is reversed after about twenty-four months. In this case, the policy reversal is smooth, a gradual decline in the funds rate from a high of 0.18 percentage points which occurs after three months. The contemporaneous impact of the contractionary monetary shock on inflation is temporary drop in the inflation rate. The decrease in inflation is statistically significant out to 36 months with the central tendency reaching a minimum value of -0.25 percentage points after three months.

The response of the one-year interest rate to the monetary shock is transitory. Here, the one-year rate rises on impact but then gradually decreases back to nearly the original level after about three years. This might imply that after

[6] Since we do not include a commodity price, we do not constrain inflation immediately after a monetary contraction to avoid the price puzzle.

the initial decrease in inflation, inflationary expectations stabilize and the one-year rate adjusts accordingly. After the initial shock to the funds rate is past, there is no further impact on inflation and interest rates return to their steady-state values. The output response in this case is ambiguous. While the mean response shows a recession out to 60 months, there exist iterations in which the output response is zero or even poistive at a 60 month horizon. These results are consistent with other studies of the effects of contractionary monetary policy on output (Uhlig (1999)).

Figure 2 about here

The response to the identified monetary shock for the State 2 is shown in Figure 2. The contemporaneous shock to the funds rate is reversed quickly, beginning almost immediately and is almost fully reversed by twenty months. One difference between states is that the contemporaneous impact on the funds rate is much larger here than in the previously discussed state. One possible explanation to this phenomenon is that policy is less credible in this state and that a larger impulse is required to initiate the same contractionary action on the inflation rate.

In this state, the contractionary policy lowers inflation in the short run. The inflation response to the shock in this state is a decline (usually to about -0.4 percentage points) but the effect at 60 months is nearly dissipated. Additionally, a contractionary monetary shock causes an increase in the one-year interest rate out to a horizon of 36 months.

Lastly, consider the identified contractionary monetary shock's effect on output. An examination of the monetary contraction leads one to belive that the shock produces an ambiguous effect on output. The central tendancy, however, is strictly positive at all horizons and examination of individual iterations suggests that nearly 70 percent of the iterations produce a long run increase in output. These results are counter to the textbook treatment of contractionary monetary policy. We believe that during this state, expectations become uncoupled with policy objectives; the large contractionary shock can restore expectations and produce a net gain in output in the long run (see Orphanides and Williams (2002) for a discussion of how expectations might become uncoupled with policy objectives). The extent that these monetary contractions may be more unanticipated could change their effectiveness. A discussion of the historical incidence of the states and the phenomena surrounding them appears in the following subsection.

## 4.2 State Process

Figure 3 about here

Consider the posterior probabilities for State 2 for each period in Figure 3. Note the switch in regime occurring in November 1979 and the subsequent switch back in October 1982. Models that assume a single breakpoint implicitly assume that the model switches to an absorbing state.[7] Since we do not assume this state is absorbing, we are able to distinguish between a transitory policy period and permanent shift in the underlying response of the economy to innovation in policy. Our model shows that the Volcker period was, in fact, a temporary event and the economy is able to resume "normal operation" once the Volcker period concluded in 1982.[8]

The regime shift that occurs from September 1973 to March 1975 coincides

[7] Many of the models that assume an absorbing state are allowing a switch in the Fed's reaction function only. Clarida, Gali and Gertler (2000), for example, omit the 1979 to 1982 period and allow a switch in the Fed's post-1982 reaction to the inflation rate. Hanson (2001) estimates a structural VAR using pre-1979 and post-1982 samples.

[8] The model was initially run for three states in order to identify a pre- and post-Volcker state in addition to the period 1979 to 1982. The filter was unable to detect the existence of a third state.

with the NBER recession around the January, 1974 oil shock and the shift occurring during 1969-70 is associated with the oil price shock of March, 1969. It is important to note that all NBER identified recessions do not switch the economic state; in fact, most of the NBER identified recessions do not appear to affect the model state at all. This implies that the states are not picking up (model-forecastable) business cycles. In this paper, the state switches are changes in both the policy rule (the response of the federal funds rate to lagged variables) and the response of the economy to monetary policy. Not all recessions are a result of a switch in regime (i.e., not all change the economy's response to monetary policy); some are just part of the business cycle dynamic embedded in the State 1 VAR. When the economy experiences a technology or supply shock of great enough magnitude, there is a change in the response to policy (and in the response of policy) that forces the switch in the underlying state. The resulting state is correlated with periods of rising inflation (the result of accommodative policy as a response to the supply shock) and falling output.

Two important elements of the identified State 2 are its duration and the economic conditions that underlie its existence. First, this state is short-lived, indicating that the economy is relatively resilient and rapidly reverts back to State 1. Second, the state appears to be associated with recessions that involve a greater than $100 billion fall in GDP (in 1996 dollars) and with periods in which there is falling output and rising inflation. These periods seem to be correlated with oil price shocks identified by Hamilton (1983, 2001) and Hoover and Perez (1994). During these periods, the economy's response to monetary stimulus and the nature of the innovations themselves change. We contend that the State 1 of the model is usually sufficient to explain both the endogenous and exogenous components of monetary policy. However, when the economy experiences inflationary recessions, it is necessary to impose a second state of the world, $S_t = 2$.

Figure 4 about here

Figure 4 shows the estimated incidence, $\nu_t^a$, of the monetary innovation and the state of the economy over the sample. The interpretation of the incidence is the magnitude of the primitive shock to the policy variable in the economic model (1). When the economy switches to State 2, the innovation becomes more volatile. The responsiveness of the funds rate to non-policy variables, especially inflation, increases dramatically. Also, the funds rate component of the identified monetary innovation is larger in State 1.[9]

# 5 Implications for Policy

The preceding analysis has the following implication for policy: The effect of an innovation in monetary policy is dependent on the state of the economy. The state of the economy can change such that the resulting economic response was different than might would be anticipated if we assume there were no state change. Since the state change alters both the impact effect and resulting response to a monetary shock, a policymaker operating under the previous regime would not forecast the same response from the economy to an equivalent contractionary shock. While these switches are infrequent, the magnitude of the change to the economy can be great. In addition, the switches tend to occur when the monetary policy may be useful as a stabilization device.

Figure 5 about here (Entire Sample)

[9] Recall that the initial impulse $\alpha$ is normalized but the resulting $\widehat{a}$ is defined by a state-dependent rotation.

To assess the model's implication for policy, consider some thought experiments in which the policymaker attempts to forecast the effect of a surprise monetary intervention. First, consider the case in which the forecaster believes the world is governed by the reduced-form VAR estimated over the entire sample, 1959:01 to 2001:06. Table 4 gives the identified monetary shock estimated for a single state model over the entire sample and Figure 5 shows the impulse responses. We will call this alternative Model A.

The Model A shock and response is qualitatively similar to the base state in the MSVAR for the two interest rates. The shock to the funds rate is reversed after about twenty months and overshoots, showing a persistent decrease even after 60 months. The contemporaneous impact of the contraction is to lower inflation and, perhaps through inflation expectations, the one-year interest rate. This response is qualitatively similar to that of State 1 in the Markov-switching model but the magnitude of the long-run impact of the contractionary shock is smaller. The output response to the contractionary shock, however, is also similar to State 1. In both cases, the monetary contraction causes output to rise in the very short run. This is caused by the contemporaneous impact of the shock. Output falls, though, after approximately six months, a lag similar to the anecdotal monetary policy lag. The long-run output response to the shock is a permanent (although insignificant) reduction in output. While these responses are qualitiatively consistent with State 1 of the MSVAR, they are dramatically different from State 2.

Figure 6 about here (Post-79 Sample)

Next, consider an economy that has permanently switched after Volcker took office in 1979. Figure 6 shows the impulse responses to the monetary contraction for the VAR estimated with post-1979 data. Again, we find a rapid policy reversal–the funds rate overshoots and falls below its original level. The shock is completely removed from the funds rate after about 40 months. Similar to the impact of the shock in the other cases, a contraction permanently lowers inflation and temporarily lowers the one year rate via inflation expectations. However, in this case, the output response is positive. The contemporaneous impact dominates the effect of the rise in the funds rate, perhaps because there is a rapid reversal in the policy instrument, and there is no cost to the disinflation.

Finally, suppose that the economy experiences a state switch but that the forecaster assumes that the switch post-1982 was permanent. She would surmise the effect of a contraction would be to cause a permanent reduction in output. Price stabilization might become a secondary objective to stabilizing output by accommodating the oil shock with a monetary injection. Figure 1 shows that this may be a mistake–increasing the funds rate in this state can be a stabilizing force in the long run. An adverse shock that raises prices and lowers output can be essentially undone by undertaking a price stabilization policy. The alternative, attempting to accommodate the shock by focusing on stabilizing output only, may lead to further inflation and a longer path to recovery.

# 6 Transition Dynamics

The above experiments indicate the differences in the contemporaneous and subsequent responses to monetary policy shocks. This section explores the transition between the two states. In particular, we are interested in the effect of a state change on the paths of both output and inflation in the absence of monetary innovations. It is important to keep in mind that these experiments are not conducted in a mentary vaccuum. Indeed, the change in states reflects a change in the policy rule. Thus, the policymaker continues to react to the economic environment. We are simply abstracting from the surprise innovations to policy that are a component of $\varepsilon_t$. In this light, we conduct the following two experiments:

Set $y_t$ equal to the historical values for the four periods preceding beginning of the Volcker disinflation, November, 1979. We then impose State 2 and allow $y_t$ to transition through $\beta_{i,2}$. The system remains in State 2 for 24 periods, roughly the average duration of a State 2 episode. After 24 periods, the system switches to State 1. This experiment (henceforth experiment A) is concluded after 48 periods. We compare this to an economy that begins with the same initial conditions. Instead of imposing State 2, the system remains in State 1 for all 48 periods (henceforth experiment B). The two tranistion paths are shown in figure 7.

Figure 7 about here

Surprisingly, the implied path absent monetary innovations for experiment A results in a lower inflation rate and a shallower recession than experiment B. This does not imply that State 2 is "better" because it sacrafices less output in order to lower inflation. Since this state is associated with a generally higher inflation rate to begin with, results suggest that lowering inflation in a high (at least by U.S. standards) inflation state may not be as costly as measured by full sample VARs. Additionally, the cost of disinflating at relatively low levels of inflation may be more costly than anticipated.

Figure 8 about here

Finally, Figure 8 compares the actual path of the four economic variables to the path implied by experiment A. Specifically, note the two actual increases in the federal funds rate. These two contractionary monetary shocks are associated with an increase in output and a decrease in mean inflation. While switching to the State 2 can alter the implied output and inflation paths, the monetary shocks lower the inflation rate further, without a tradeoff in output. This adds further confirmation that disinflationary policy during moderate inflations can be stabilizing in the short run as well as the long run. The policymaker's restoration of credibility, in this case, can be sufficient to outweigh the effect of contraction on output.

# 7 Conclusion

Two of the fundamental questions underlying the analysis of a structural VAR are what is a monetary shock and what effects does it generate? Most structural VAR models define restrictions on the covariance matrix, essentially specifying the answer to the first question, to answer the second. We take as given the effect of a contractionary policy shock on inflation and the policy instrument to verify both the effects on other variables and to answer the question: what is a monetary shock? We find that the contemporaneous effect of and subsequent response to the shock can differ across regimes. The output response to the contractionary shock is different in different regimes: positive when the economy is in one state and persistently negative when the economy is in normal times.

We also find that evidence that the shift in the underlying structural process governing the VAR is not a single breakpoint process as assumed by much of the literature. We show that the so-called Volcker break at the end of 1979 is not an absorbing state–i.e., that the process reverts back after a number of periods. The estimated coefficient matrix for the Volcker subperiod can dominate post-1979 and simulate the presence of a single structural break with an absorbing state. We conclude that forecasting the effect of policy under the assumption of an absorbing mistake can lead to drastically different conclusions about the impact of a proposed policy.

While it is clear that addressing models with predetermined structural breaks can be misleading, further study of the timing and especially the forecasting of switches in the underlying process are necessary. Long run restrictions, perhaps in the form of specifying restrictions on the long-run impact matrix of a Markov-switching VECM, could provide additional insight (Sarno, Thornton and Valente (2002) and Francis and Owyang (2002)).

# 8 Appendix: Estimation via Gibbs sampling

The algorithm can be summarized by the following process:

- Given a $\tilde{S}_T$ and $\tilde{y}_T$, generate $\widetilde{\beta}_k$ and $\widetilde{\Sigma}$ from the conditional distribution $p(\widetilde{\beta}_k, \widetilde{\Sigma}|\tilde{y}_T, \tilde{S}_T)$.

- Given a $\widetilde{\beta}_k$ and $\widetilde{\Sigma}$ and $\tilde{y}_T$, generate $\tilde{S}_T$ from the conditional distribution $p(\tilde{S}_T|\tilde{y}_T, \widetilde{\beta}_k, \widetilde{\Sigma})$.

- If the number of iterations exceeds that required to meet a predetermined convergence criteria:

  Draw $\hat{a}_{S_t} = \widehat{A}\alpha$ 2000 times.

  Calculate the impulse response to $\hat{a}_{S_t}$.

  Compare to the restriction matrix $R$. Save if the restrictions are satisfied.

  Return to step 1.

- If the number of iterations does not exceed that required to meet a predetermined convergence criteria:

  Return to step 1.

## 8.1 Parameter Sampling

The objective is to characterize the joint distribution

$$p(\widetilde{\beta}_k, \widetilde{\Sigma}, \tilde{S}_T | \tilde{y}_T), \tag{7}$$

where $\widetilde{\beta}_k = \left\{\beta_1^{[1]}, \beta_2^{[1]}, ..., \beta_k^{[1]} | \beta_1^{[2]}, \beta_2^{[2]}, ..., \beta_k^{[2]}\right\}'$, $\widetilde{\Sigma} = \left\{\Sigma^{[1]}, \Sigma^{[2]}\right\}$ $\tilde{S}_T = \{S_1, S_2, ..., S_T\}'$, and $\tilde{y}_T = \{y_1, y_2, ..., y_T\}'$.[10] Deriving $p(\widetilde{\beta}_k, \widetilde{\Sigma}, \tilde{S}_T | \tilde{y}_T)$ directly is not feasible; instead we will employ the method of Gibbs sampling to generate the posterior distributions $p(\widetilde{\beta}_k, \widetilde{\Sigma} | \tilde{y}_T, \tilde{S}_T)$ and $p(\tilde{S}_T | \tilde{y}_T, \widetilde{\beta}_k, \widetilde{\Sigma})$. Iterative sampling from these posterior distribution is equivalent to sampling from the ergodic distribution $p(\widetilde{\beta}_k, \widetilde{\Sigma}, \tilde{S}_T | \tilde{y}_T)$. With this structure, consider the $k$th order system of equations (2), stacked in the following manner:

$$\mathbf{Y} = \mathbf{XB} + \varepsilon, \tag{8}$$

where

$$\mathbf{B} = \begin{bmatrix} \beta_1^{[1]}, & \beta_2^{[1]}, & ..., & \beta_k^{[1]} \\ \beta_1^{[2]}, & \beta_2^{[2]}, & ..., & \beta_k^{[2]}, \end{bmatrix}',$$

$$\mathbf{Y} = \begin{bmatrix} y_1 & y_2, & ..., & y_T \end{bmatrix}',$$

$$X_t = \begin{bmatrix} y'_{t-1}, & y'_{t-2}, & ..., & y'_{t-l} \end{bmatrix}',$$

$$\mathbf{X} = \begin{bmatrix} X_1, & X_2, & ..., & X_T \\ S_1X_1, & S_2X_2, & ..., & S_TX_T \end{bmatrix}'$$

[10] For an overview of techniques used to estimate Markov-switching VARs, see Krolzig (1997, 1998).

and

$$\boldsymbol{\varepsilon} = \begin{bmatrix} \varepsilon_1, & \varepsilon_2, & ..., & \varepsilon_T \end{bmatrix}'.$$

Then, assuming normality of the coefficient matrix and no contemporaneous correlation, candidate parameters' conditional joint posterior distribution $p(\widetilde{\beta}_k, \widetilde{\Sigma}|\tilde{y}_T, \tilde{S}_T)$ is a Normal-Wishart distribution with priors $\nu_0$, $N_0$, $Z_0$, and $W_0$. Thus, $\widetilde{\beta}_k$ and $\widetilde{\Sigma}$ can be drawn from a Normal-Wishart distribution with degrees of freedom $\upsilon$, precision matrix $N$, parameter means $Z$, and and covariance matrix $W$ defined by

$$\begin{aligned} \nu &= \widehat{T} + \nu_0, \\ N &= N_0 + \mathbf{X}'\mathbf{X}, \\ \overline{\mathbf{Z}} &= N^{-1}\left(N_0 Z_0 + \mathbf{X}'\mathbf{X}\widehat{Z}\right), \\ W &= \frac{\nu_0}{\nu} W_0 + \frac{\widehat{T}}{\nu}\hat{\Sigma} + \frac{1}{\nu}\left(\widehat{Z} - Z_0\right)' N_0 N^{-1}\mathbf{X}'\mathbf{X}\left(\widehat{Z} - Z_0\right), \end{aligned} \tag{9}$$

where $\widehat{T}$ is the number of periods in the state, $\widehat{Z} = (\mathbf{X}'\mathbf{X})^{-1}\mathbf{X}'\mathbf{Y}$ and $\hat{\Sigma} = \left(\mathbf{Y} - \mathbf{X}\widehat{Z}\right)'\left(\mathbf{Y} - \mathbf{X}\widehat{Z}\right)$.

### 8.2 State Sampling

The distribution $f(\tilde{S}_T|\tilde{y}_T, \widetilde{\beta}_k, \widetilde{\Sigma})$ can be generated by first determining the conditional likelihood for $S_t$ at each $t$. Recall that, conditional on $\widetilde{y}_T$, $\widetilde{\beta}_k$ and $\widetilde{\Sigma}$, (2) is linear in $\widetilde{S}_T$. Given $p(S_0|\widetilde{y}_0)$, a prior probability for the initial state, Hamilton's (1989, 1990) filter generates the conditional density $p(S_T|\widetilde{y}_T, \widetilde{\beta}_k, \widetilde{\Sigma})$. Then, following Carter and Kohn (1994) and Kim and Nelson (1998, 1999), the density $p(\widetilde{S}_T|\widetilde{y}_T, \widetilde{\beta}_k, \widetilde{\Sigma})$ is obtained from

$$p(\widetilde{S}_T|\widetilde{y}_T, \widetilde{\beta}_k, \widetilde{\Sigma}) = p(S_T|\widetilde{y}_T, \widetilde{\beta}_k, \widetilde{\Sigma}) \prod_{t=1}^{T-1} p(S_t|\widetilde{y}_T, \widetilde{\beta}_k, \widetilde{\Sigma}, S_{t+1}). \tag{10}$$

Each density $p(S_t|\widetilde{y}_T, \widetilde{\beta}_k, \widetilde{\Sigma}, S_{t+1})$ is generated from a filtering algorithm and Bayes' Law:

$$p(S_t|\widetilde{y}_T, \widetilde{\beta}_k, \widetilde{\Sigma}, S_{t+1}) = \frac{p(S_{t+1}|\widetilde{y}_t, S_t, \widetilde{\beta}_k, \widetilde{\Sigma})p(S_t|\widetilde{y}_t, \widetilde{\beta}_k, \widetilde{\Sigma})}{p(S_{t+1}|\widetilde{y}_t, \widetilde{\beta}_k, \widetilde{\Sigma})}$$

$$= \frac{p(S_{t+1}|S_t)p(S_t|\widetilde{y}_t, \widetilde{\beta}_k, \Sigma)}{\sum_{S_t} p(S_{t+1}|\widetilde{y}_t, S_t, \widetilde{\beta}_k, \Sigma p(S_t|\widetilde{y}_t, \widetilde{\beta}_k, \Sigma)}$$

$$= \frac{p(S_{t+1}|S_t)p(S_t|\widetilde{y}_t, \widetilde{\beta}_k, \widetilde{\Sigma})}{\sum_{S_t} p(S_{t+1}|S_t)p(Z_t|\widetilde{y}_t, \widetilde{\beta}_k, \widetilde{\Sigma})}, \qquad (11)$$

where $p(S_{t+1}|S_t)$ is the transition probability and the filter determines the density $p(S_t|\widetilde{y}_t, \widetilde{\beta}_k, \widetilde{\Sigma})$. The final two inequalities arise from the Markov property of $S_t$: in determining the density for $S_{t+1}$, the only relevant information in the available set is the previous state $S_t$. The numerator in (11) is calculated from the Hamilton filter as

$$p(S_t|\widetilde{y}_t, \widetilde{\beta}_k, \widetilde{\Sigma}) = \frac{f(y_t|\widetilde{y}_{t-1}, S_t, \widetilde{\beta}_k, \widetilde{\Sigma})p(S_t|\widetilde{y}_{t-1}, \widetilde{\beta}_k, \widetilde{\Sigma})}{f(y_t|\widetilde{y}_{t-1}, \widetilde{\beta}_k, \widetilde{\Sigma})}$$

$$= \frac{f(y_t|\widetilde{y}_{t-1}, S_t, \widetilde{\beta}_k, \widetilde{\Sigma})p(S_t|\widetilde{y}_{t-1}, \widetilde{\beta}_k, \widetilde{\Sigma})}{\sum_{S_t} f(y_t|\widetilde{y}_{t-1}, S_t, \widetilde{\beta}_k, \widetilde{\Sigma})p(S_t|\widetilde{y}_{t-1}, \widetilde{\beta}_k, \widetilde{\Sigma})}, \qquad (12)$$

where

$$p(S_t|\widetilde{y}_{t-1}, \widetilde{\beta}_k, \widetilde{\Sigma}) = \sum_{S_{t-1}} p(S_t|S_{t-1})p(S_{t-1}|\widetilde{y}_{t-1}, \widetilde{\beta}_k, \widetilde{\Sigma}).$$

The density $p(S_{t-1}|\widetilde{y}_{t-1}, \widetilde{\beta}_k, \widetilde{\Sigma})$ is taken from the previous iteration.

The transition matrix $\Pi$ is assumed to have a prior distribution of the form $\Pi_{ii} \sim \beta(u_{ii}, u_{ji})$ for $j \neq i$, the distribution for the transition probabilities $\Pi_{ii}$ is determined by

$$\Pi_{ii} = \Pr[S_t = i | S_{t-1} = i] \sim \beta(u_{ii} + n_{ii}, u_{ji} + n_{ji}),$$

where $n_{ii}$ is the number of periods that $S_t$ remained in state $i$, and $n_{ji}$ is the number of periods that $S_t$ switched to state $j \neq i$ after beginning in state $i$. Then the other elements of $\Pi$ can be determined by $\Pi_{ji} = 1 - \Pi_{ii}$.

Estimation of the model (2) amounts to repeated iterations of sequential draws from the conditional distributions (7) and (10), constituting a Markov chain. After a sufficient number of iterations, the set of samples drawn from each conditional distribution, the ergotic distribution of this chain of conditional densities, is the joint density $p(\widetilde{\beta}_k, \widetilde{\Sigma}, \tilde{S}_T | \tilde{y}_T)$.[11]

### 8.3 Priors

Estimation of the system specified above requires priors on the posterior distribution parameters (9) and the transition matrix governing the states in the Markov process. We specify flat priors for the parameter priors by choosing $\nu_0 = 0$, $N_0 = 0$ and $Z_0 = I_{nm \asymp n}$, and $W_0 = I_{n \asymp n}$. This choice of priors makes the posterior distribution parameters

$$\begin{aligned}
\nu &= \widehat{T}, \\
N &= \mathbf{X}'\mathbf{X}, \\
\overline{\mathbf{Z}} &= \left(\mathbf{X}'\mathbf{X}\right)^{-1} \mathbf{X}'\mathbf{Y}, \\
W &= \frac{T}{\nu} \hat{\Sigma},
\end{aligned}$$

which can be used to generate the draws from the sampling algorithm.

The priors for generating the transition matrix probabilities are weighted toward persistent states. Any values $u_{ii} > u_{ji}$ achieve this prior. We choose the

[11] For an analysis of the convergence criteria, see Geweke (1992) and Gamerman (1998).

persistent states prior to maximize the economic interpretation of the posterior state distribution. States that are highly transitiory (e.g., states lasting only a few months) leave little room for historical economic interpretation.

# 9 References

## References


[1] Bernanke, B. and Mihov, I., 1998, "Measuring Monetary Policy," *Quarterly Journal of Economics* 113, 869-902.

[2] Boivin, J. and Giannoni, M., 2002, "Assessing Changes in the Monetary Transmission Mechanism: A VAR Approach," Unpublished Manuscript.

[3] Carter, C. and Kohn, R., 1994, "On Gibbs Sampling for State Space Models," *Biometrika* 81, 541-553.

[4] Christiano, L. Eichenbaum, M., and Evans, C., 1999, "Monetary Policy Shocks: What Have We Learned and to What End?" in Taylor, J., and Woodford, M., eds., *Handbook of Macroeconomics, Volume 1*, Elsevier, 65-148.

[5] Clarida, R., Gali, J., and Gertler, M., 2000, "Monetary Policy Rules and Macroeconomic Stability: Evidence and Some Theory," *Quarterly Journal of Economics* 115, 147-180.

[6] Dennis, R., 2001, "The Policy Preferences of the US Federal Reserve," Federal Reserve Bank of San Francisco Working Papers. No 2001-08.

[7] Faust, J., 1998, "The Robustness of Identified VAR Conclusions About Money," Board of Governors of the Federal Reserve System. International Finance Discussion Papers. No. 610.

[8] Gamerman, D., 1997, *Markov Chain Monte Carlo: Stochastic Simulations for Bayesian Inference*, Chapman and Hall.

[9] Geweke, J., 1992, 'Evaluating the Accuracy of Sampling-Based Approaches to the Calculation of Posterior Moments', in J.O. Berger, J.M. Bernardo,

A.P. Dawid, and A.F.M. Smith (eds.) *Proceedings of the Fourth Valencia International Meeting on Bayesian Statistics*, pp. 169-194, Oxford University Press.

[10] Hamilton, J., 1983, "Oil and the Macroeconomy Since World War II," *Journal of Political Economy*, 91, pp. 228-248.

[11] Hamilton, J., 1989, "A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle," *Econometrica*, 57, pp. 357-384.

[12] Hamilton, J., 1990, "Analysis of Time Series Subject to Changes in Regime," *Journal of Econometrics,* 45, 39-70.

[13] Hamilton, J., 2001, "What is an Oil Shock?" Unpublished Manuscript.

[14] Hanson, M., 2001, "Varying Monetary Policy Regimes: A Vector Autoregressive Investigation," Unpublished Manuscript.

[15] Hoover, K. and Perez, S., 1994, "Post Hoc Ergo Propter Once More: An Evaluation of 'Does Monetary Policy Matter?' in the Spirit of James Tobin," *Journal of Monetary Economics*, 34, 47-73.

[16] Kim, C.J., and Nelson, C., 1998, "Business Cycle Turning Points, a New Coincident Index, and Tests of Duration Dependence Based on a Dynamic Factor Model with Regime Switching," *Review of Economics and Statistics,* 80, pp. 188-201.

[17] Kim, C.J., and Nelson, C., 1999, *State-Space Models with Regime Switching: Classical and Gibbs-Sampling Approaches with Applications*, Cambridge, MA: MIT Press.

[18] Krolzig, H., 1997, *Markov-Switching Vector Autoregressions. Modelling, Statistical Inference and Application to Business Cycle Analysis*, Lecture

Notes in Economics and Mathematical Systems, Volume 454, New York: Springer.

[19] Krolzig, H., 2000, "Predicting Markov-Switching Vector Autoregressive Processes," Nuffield College, Economics Discussion Paper W31.

[20] Leeper, E., Sims, C., and Zha, T., 1996, "What Does Monetary Policy Do?" *Brookings Papers on Economic Activity*, 2, pp. 1-63.

[21] Orphanides, A., Williams, J. 2002, "Imperfect Knowledge, Inflation Expectations, and Monetary Policy." Federal Reserve Board of Governors Finance and Economics Discussion Papers No. 27.

[22] Owyang, M., 2001, "Persistence, Excess Volatility and Volatility Clusters in Inflation," *Federal Reserve Bank of Saint Louis Review*, November/December 2001, 41-51.

[23] Owyang, M. and Ramey, G., 2001, "Regime Switching and Monetary Policy Measurement," Federal Reserve Bank of Saint Louis Working Papers. No. 2001-002A.

[24] Rousch, J., 2001, " Evidence Uncovered: Long-Term Interest Rates, Monetary Policy, and the Expectations Theory," Federal Reserve Board of Governors International Finance Discussion Working Paper No. 2001-712.

[25] Sims, C., 1999, " Drift and Breaks in Monetary Policy," Unpublished Manuscript.

[26] Sims, C., 2001, "Stability and Instability in US Monetary Policy Behavior," Unpublished Manuscript.

[27] Sims, C. and Zha, T., 1999, "Error Bands for Impulse Responses," *Econometrica*, 67, pp. 1113-55.

[28] Sims, C. and Zha, T., 2002, "Macroeconomic Switching," Unpublished Manuscript.

[29] Taylor, J., 1993, "Discretion Versus Policy Rules in Practice" Carnegie-Rochester Conference Series on Public Policy, 39, 195-214.

[30] Thornton, D., 2001, "The Federal Reserve's Operating Procedures, Nonborrowed Reserves, Borrowed Reserves and the Liquidity Effect," *Journal of Banking and Finance*, 31, pp. 1717-39

[31] Uhlig, H., 1999, "What Are the Effects of Monetary Policy on Output? Results From an Agnostic Identification Procedure," Harald Centre for Economic Policy Research Discussion Paper 2137.

[32] Uhlig, H., 2001, "Did the Fed Surprise the Markets in 2001? A Case Study for VAR's with Sign Restrictions," CESifo Working Paper No. 629.

[33] Waggoner, D. and Zha, T., 2000, "A Gibbs Simulator for Restricted VAR Models," Federal Reserve Bank of Atlanta Working Paper No. 2000-3.

# 10 Tables

Table 1: Restriction Matrix, $R$

| Variable | Restriction | Periods |
|---|---|---|
| CII | $n/a$ | $n/a$ |
| PCE | $-$ | $6-12$ |
| 1YR | $n/a$ | $n/a$ |
| FFR | $+$ | $1-8$ |

Table 2: Monetary Shocks

| Variable | $a_1$ | $a_2$ |
|---|---|---|
| CII | $-0.0016$ (0.1278) | 0.1164 (0.1362) |
| PCE | $-0.8855$ (0.4078) | $-0.2776$ (0.8673) |
| 1YR | 0.0976 (0.0557) | 0.4264 (0.1582) |
| FFR | 0.1195 (0.0701) | 0.4277 (0.3120) |

Table 3: Monetary Shocks

| Variable | $a_1$ | $a_2$ | $a_{59}$ | $a_{79}$ |
|---|---|---|---|---|
| CII | $-0.0016$ (0.1278) | 0.1164 (0.1362) | 0.0543 (0.0503) | 0.0311 (0.0366) |
| PCE | $-0.8855$ (0.4078) | $-0.2776$ (0.8673) | $-0.0448$ (0.0169) | $-0.0539$ (0.0255) |
| 1YR | 0.0976 (0.0557) | 0.4264 (0.1582) | $-0.1016$ (0.0543) | $-0.1073$ (0.0571) |
| FFR | 0.1195 (0.0701) | 0.4277 (0.3120) | 0.2217 (0.0760) | 0.2639 (0.0659) |

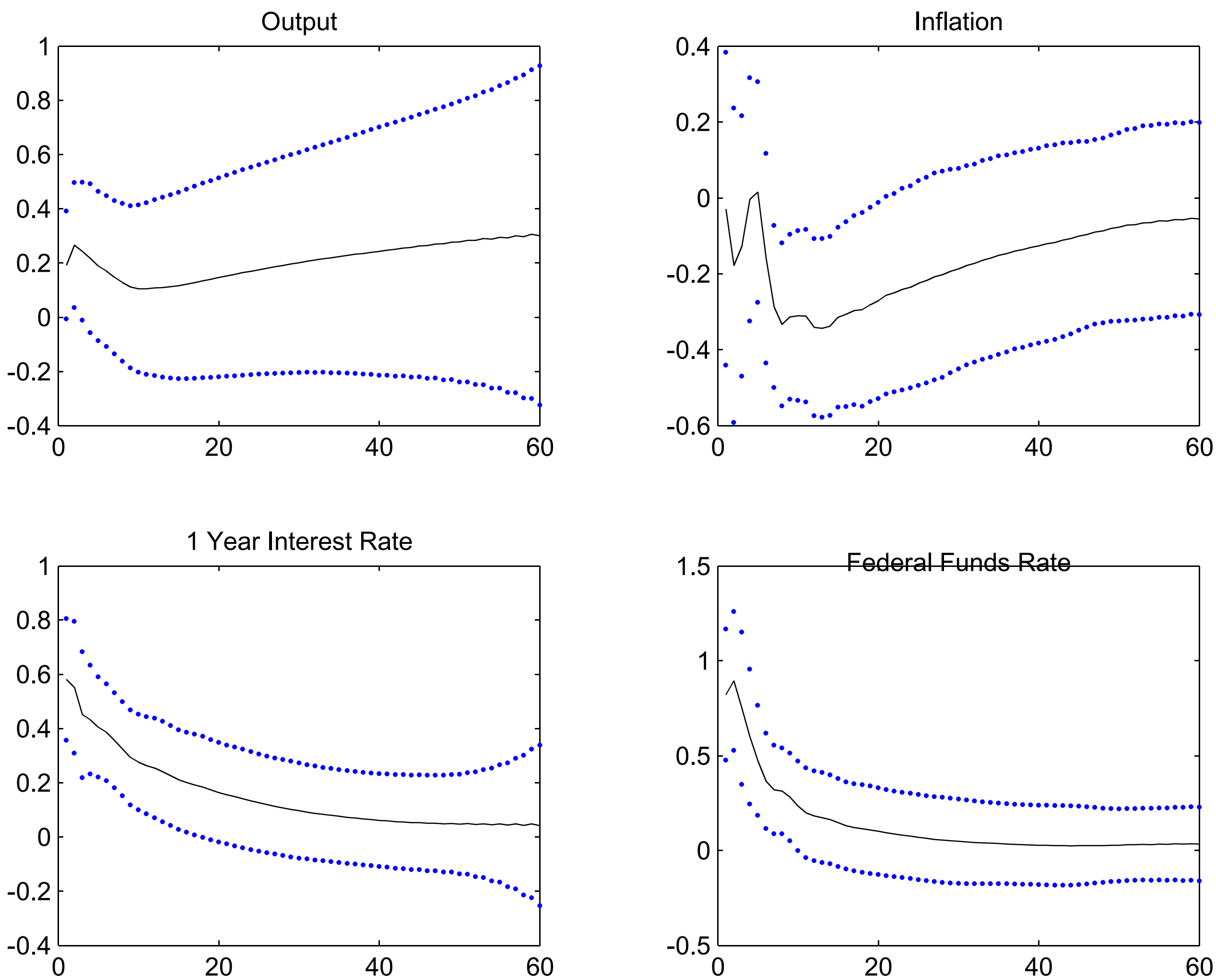


Figure 1: Impulse Responses for State 1

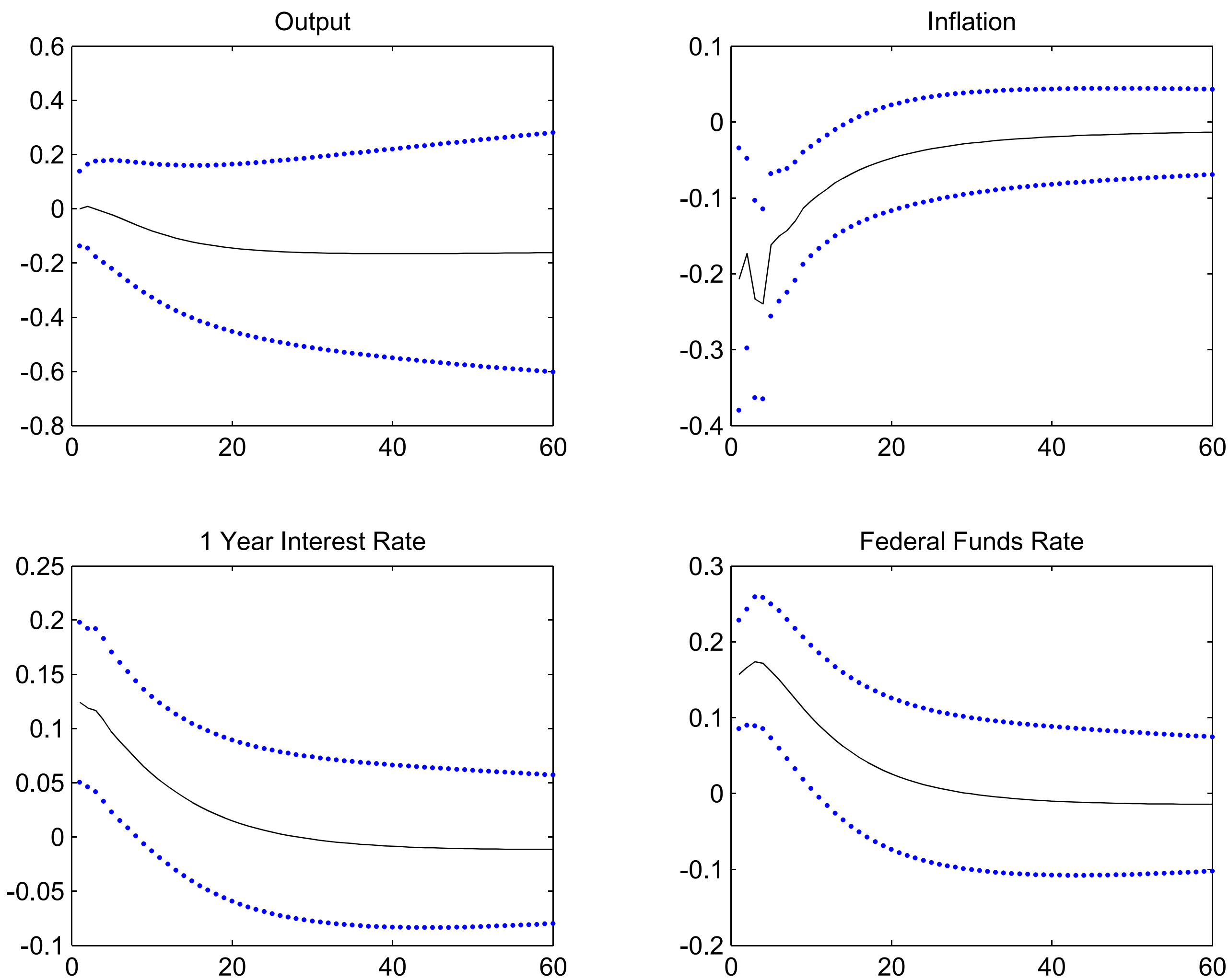


Figure 2: Impulse Responses for State 2

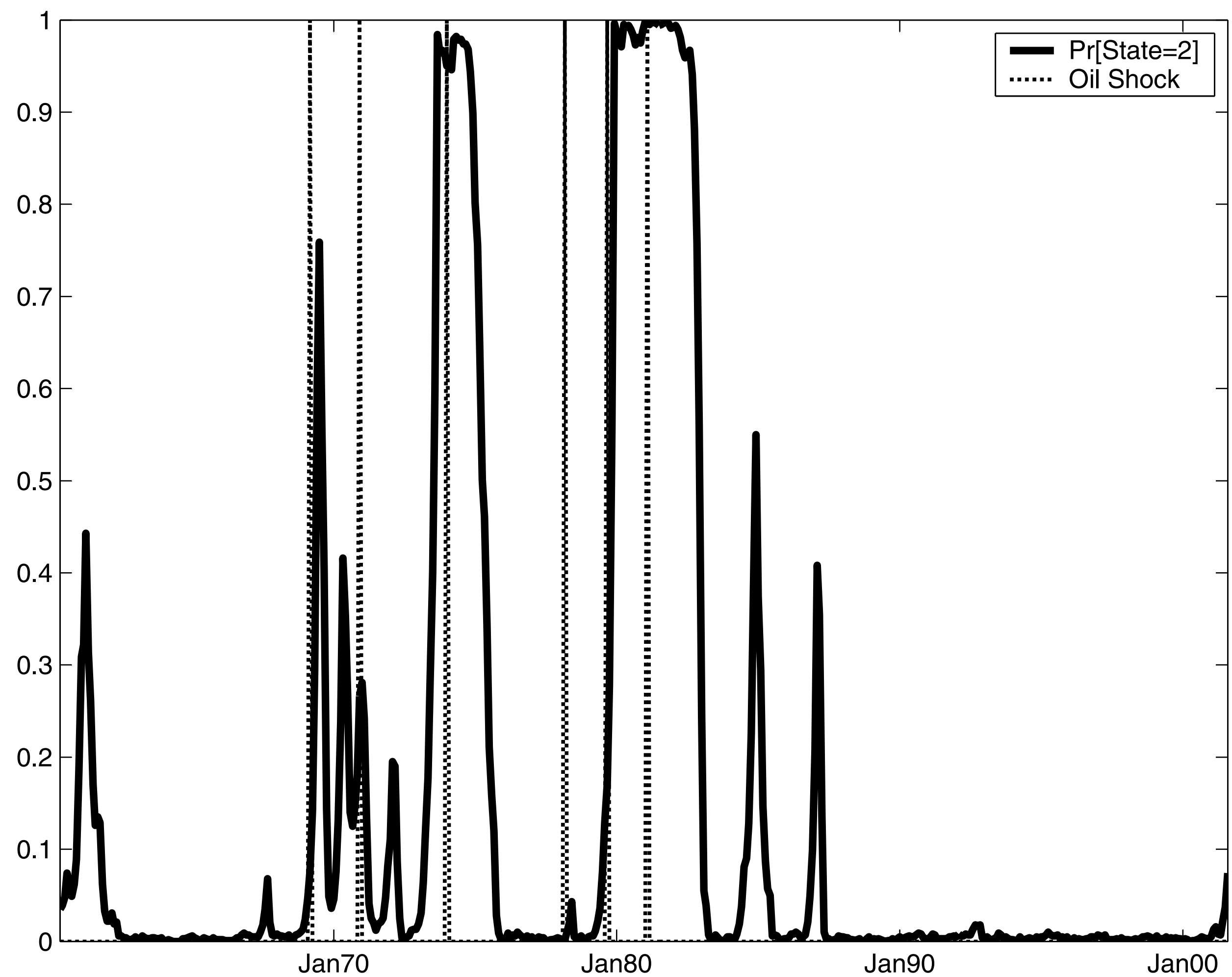


Figure 3: States and Oil Dates

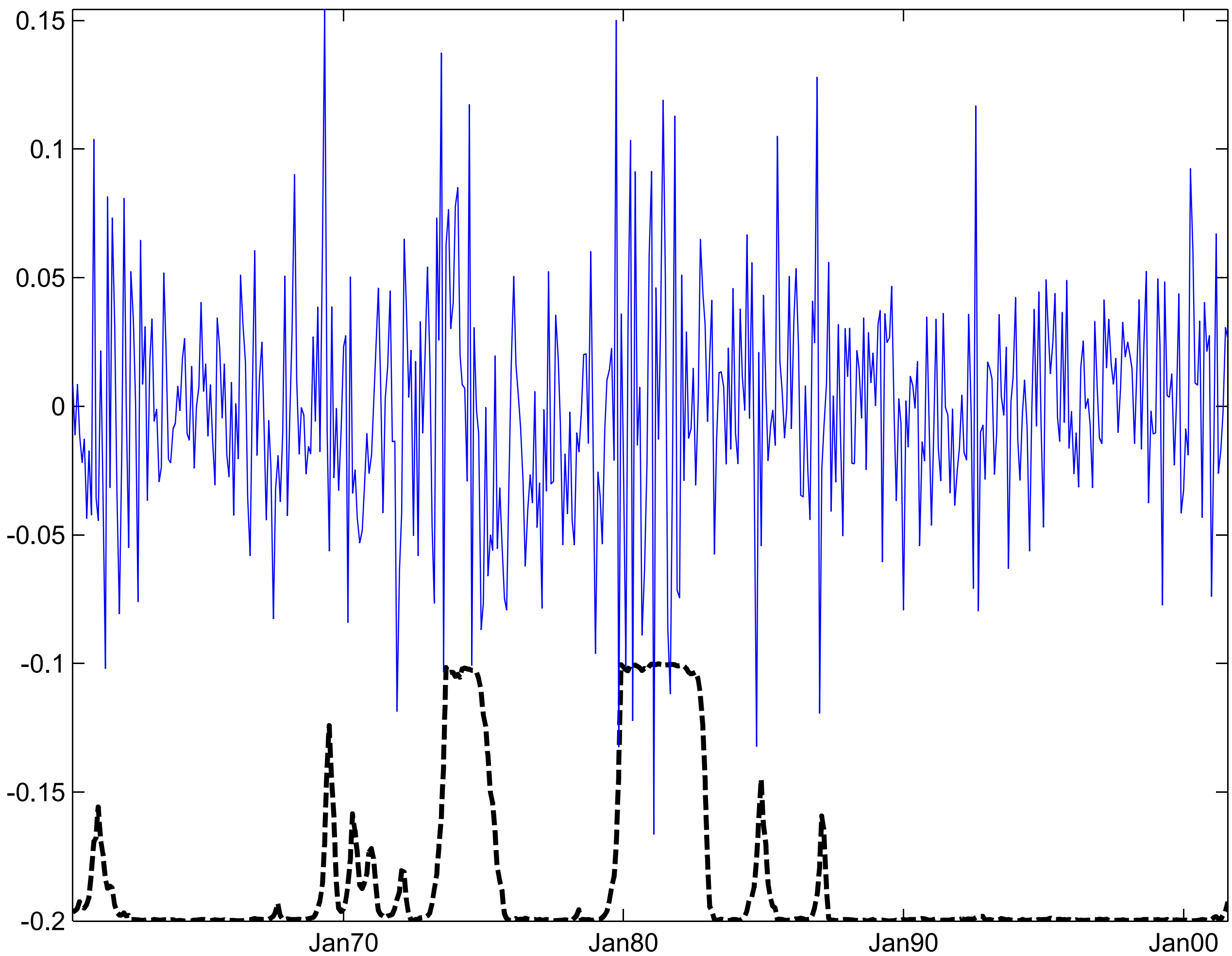


Figure 4: Monetary Policy Incidence

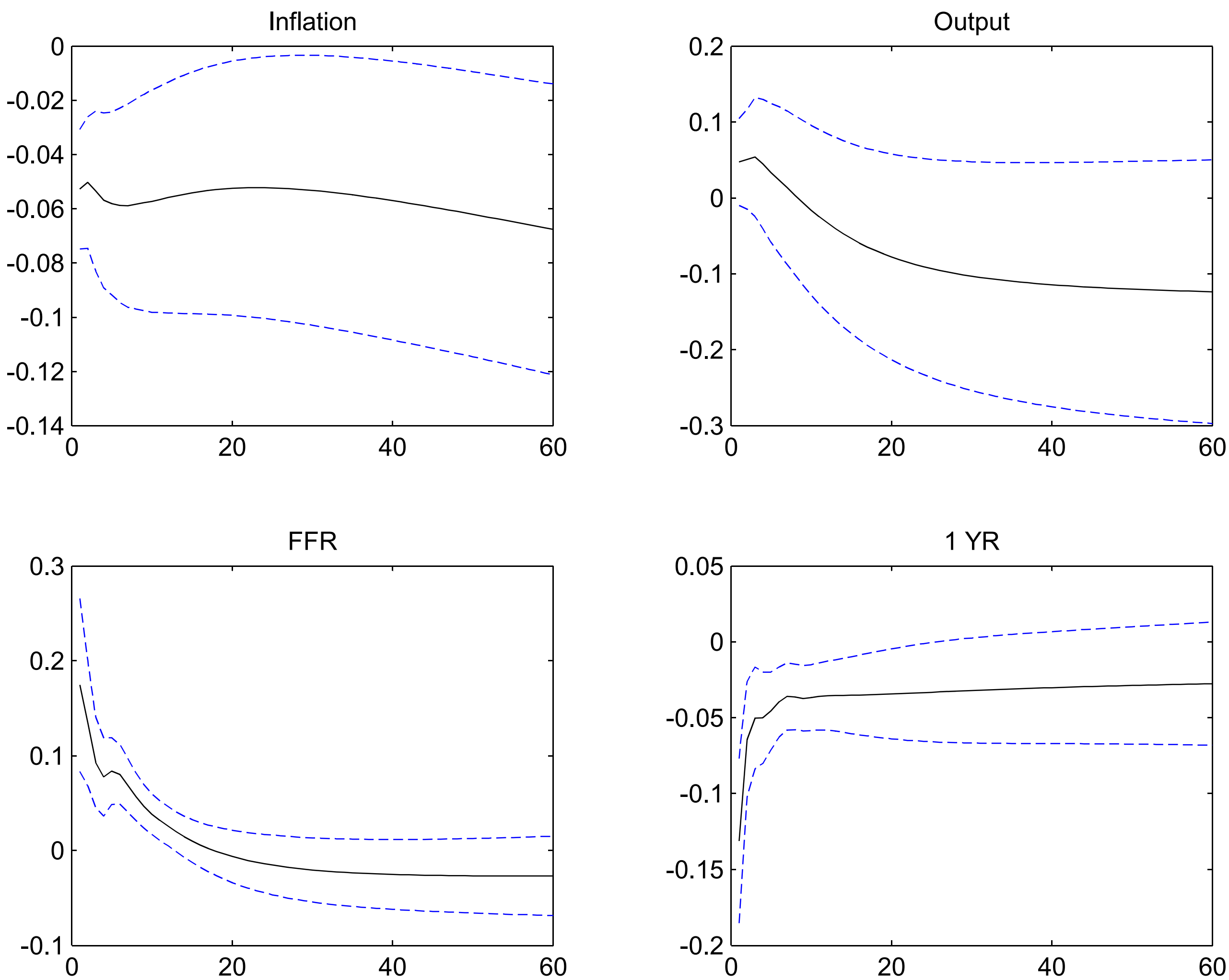


Figure 5: Impulse Responses for 1959:01 to 2001:06

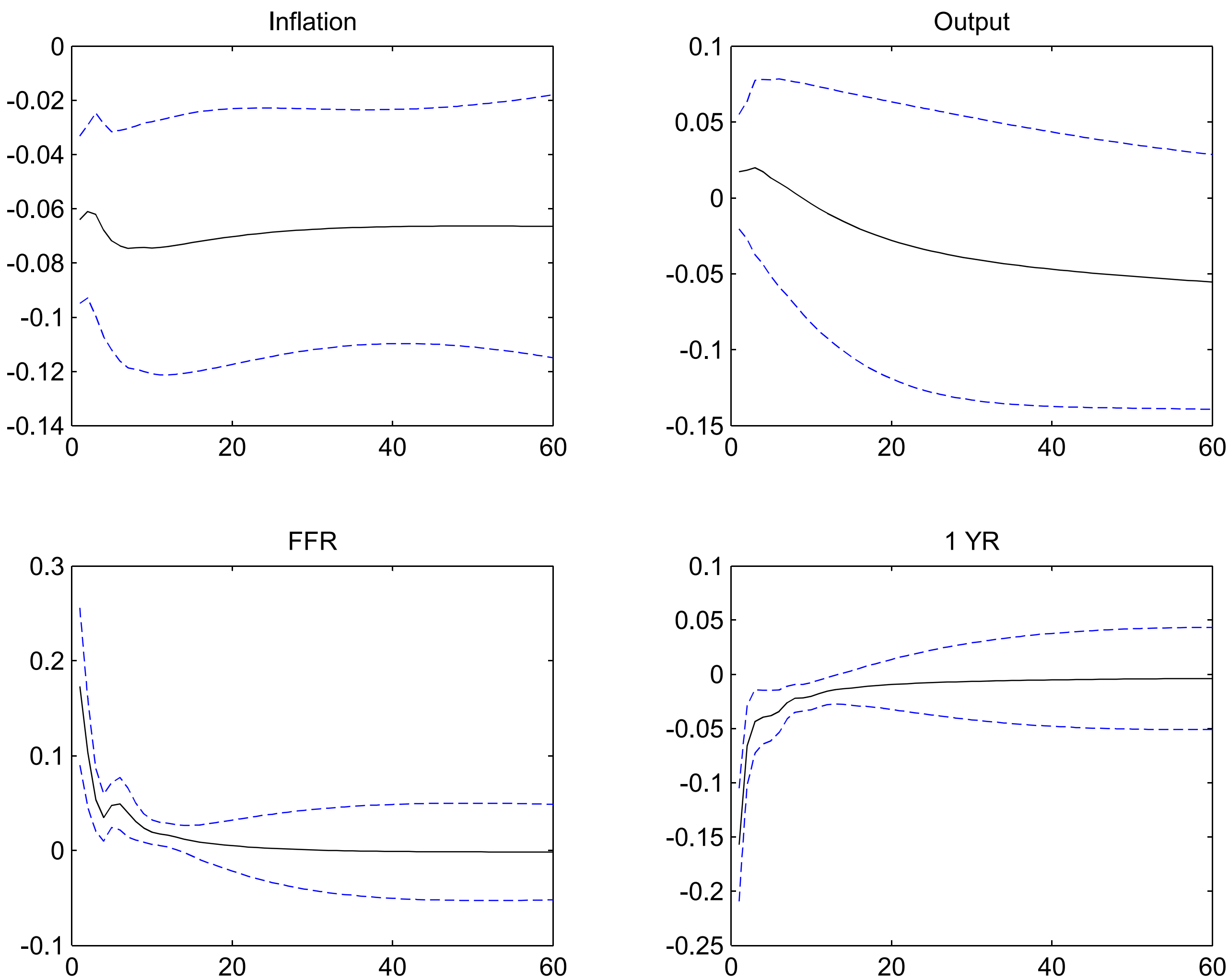


Figure 6: Post 1979 Sample

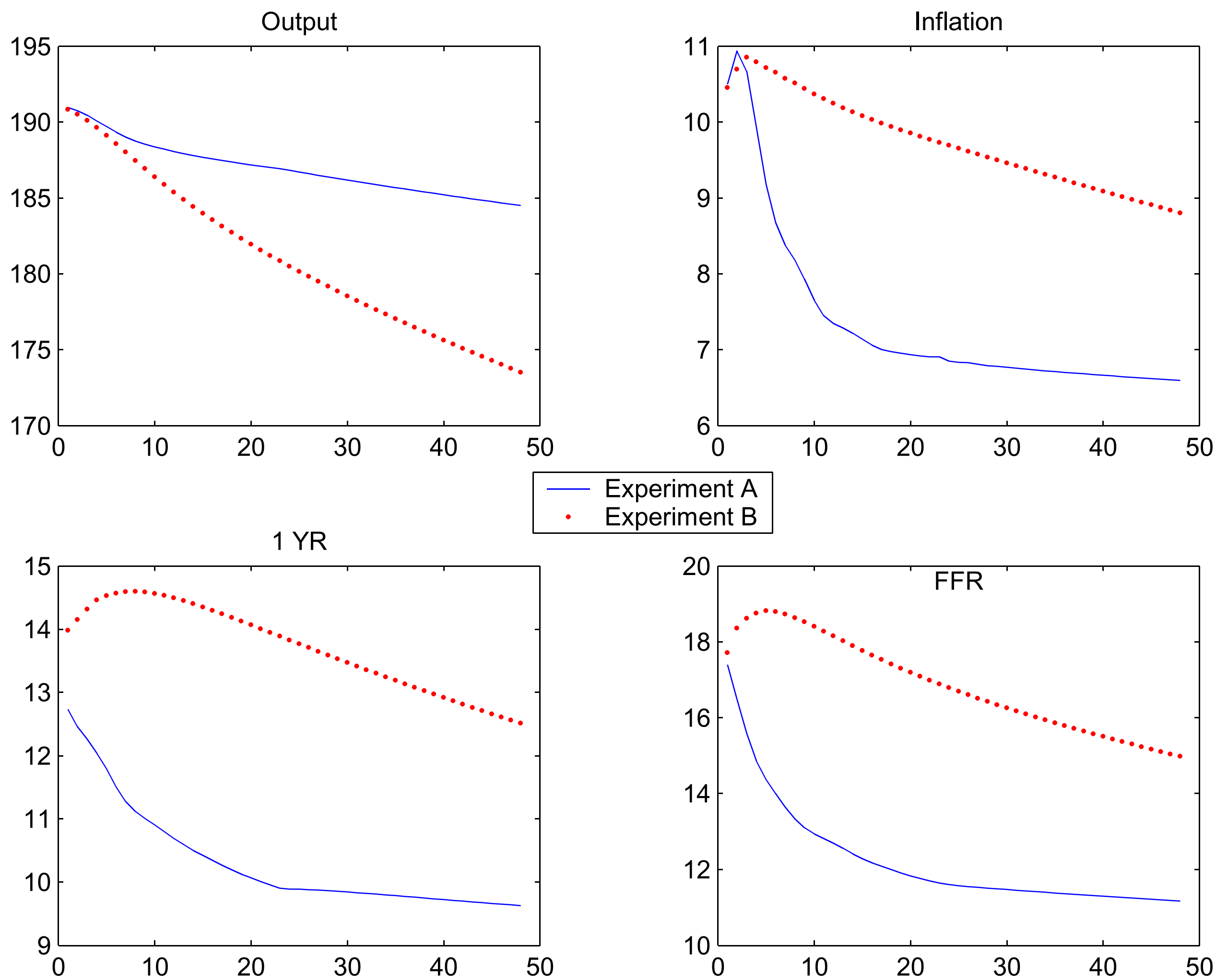


Figure 7: State Transition Dyanamics

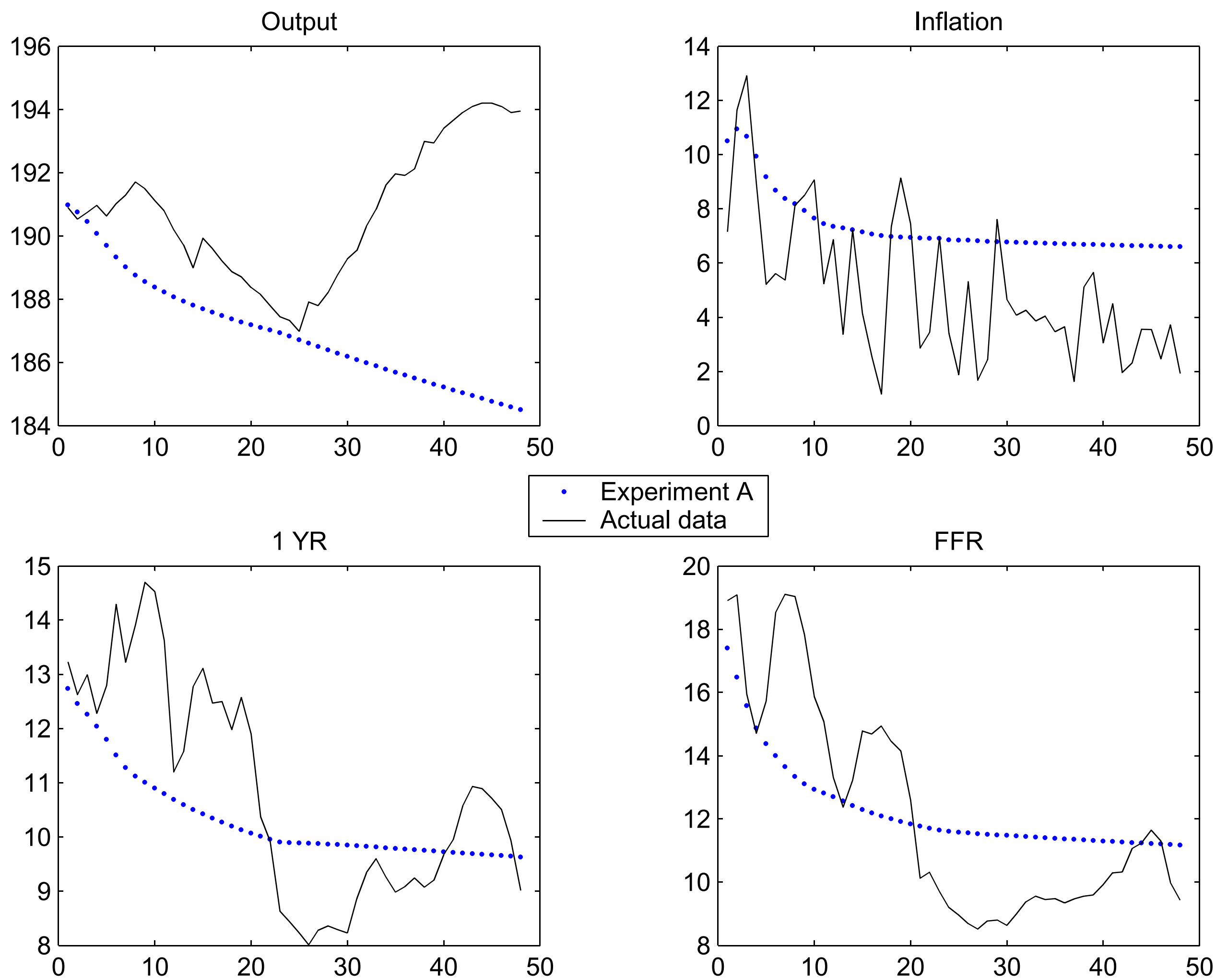


Figure 8: Experiment A v. Actual