



ECONOMIC RESEARCH
FEDERAL RESERVE BANK OF ST. LOUIS
WORKING PAPER SERIES

Monetary Policy in a Markov-Switching VECM: Implications for the Cost of Disinflation and the Price Puzzle

Authors	Neville Francis, and Michael T. Owyang
Working Paper Number	2003-001D
Citable Link	https://doi.org/10.20955/wp.2003.001
Suggested Citation	Francis, N., Owyang, M.T.; Monetary Policy in a Markov-Switching VECM: Implications for the Cost of Disinflation and the Price Puzzle, Federal Reserve Bank of St. Louis Working Paper 2003-001. URL https://doi.org/10.20955/wp.2003.001

Published In	Journal of Business & Economic Statistics
Publisher Link	https://doi.org/10.1198/073500104000000325

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

The views expressed in this paper are those of the author(s) and do not necessarily reflect the views of the Federal Reserve System, the Board of Governors, or the regional Federal Reserve Banks. Federal Reserve Bank of St. Louis Working Papers are preliminary materials circulated to stimulate discussion and critical comment.

Monetary Policy in a Markov-Switching VECM: Implications for the Cost of Disinflation and the Price Puzzle

Neville Francis
Lehigh University
Dept. of Economics
621 Taylor Street
Bethlehem, PA 18015

Michael T. Owyang
Federal Reserve Bank of St. Louis
Research Department
411 Locust Street
Saint Louis, MO 63102
(owyang@stls.frb.org)

JEL Classification: C32, E31, E52

Keywords: Monetary Policy, Markov Switching, Cointegration,
Price Puzzle, Sacrifice Ratio

June 21, 2004

Abstract

Monetary policy VARs typically presume stability of the long-run outcomes. We introduce the possibility of switches in the long-run equilibrium in a cointegrated VAR by allowing both the covariance matrix and weighting matrix in the error-correction term to switch. We find that monetary policy alternates between sustaining long-run growth and disinflationary regimes. Allowing state changes can also help explain the price puzzle and justify the use of commodity prices as a corrective measure. Finally, we show that regime-switching has implications for disinflationary monetary policy and can explain the variety of sacrifice ratio estimates that exist in the literature.

1 Introduction

Changes in monetary policy can occur in either the implementation of policy (shocks) or the objectives of policy (regimes). The former are typically modeled as vector innovations to a vector autoregression (VAR) in which monetary policy is identified by structural restrictions on the contemporaneous impacts of the variables (e.g., Sims 1992) or restrictions on the long-run effects of shocks (e.g., Blanchard and Quah 1989). This so-called structural VAR literature has identified a number of stylized facts resulting from the implementation of monetary policy and has spawned a vast literature (see Bernanke and Mihov 1998; Christiano, Eichenbaum, and Evans 1999 for surveys).

More recently, switching monetary policy regimes have garnered some attention (Clarida, Galí, and Gertler 2000; Dennis 2001; Hanson 2002a; Boivin and Giannoni 2002). Policy regimes involve switches in the policy rule that reflect, for example, changes in the policymaker's reaction to deviations from the target inflation rate or output growth rate. These studies are aimed at finding persistent changes in policy which result, for example, from changes in central bank leadership or transparency. These regime changes can have a large effect on the volatility of money, interest rates, and output. For example, Clarida, Gali, and Gertler show that a switch in the objectives of monetary policy post-1982 has resulted in a more stable, inflation-controlling policy. Dennis argues that a change in policymaker preferences has shifted the post-1979 inflation target down from over 7 percent to under 2 percent. Hanson examines whether a change in Fed policy was the cause of increased instability in the late 1970s and early 1980s. Boivin and Giannoni consider whether the Fed's effectiveness has changed in the postwar period.

A new branch of literature has begun to simultaneously examine both regime changes and policy shocks (see, for example, Bernanke and Mihov 1998; Owyang 2002; Sims and Zha 2002). These authors show, among other things, that the stance of monetary policy is important not only to the policymaker's response to the exogenous economic shocks (e.g., the Taylor rule) but also to the contemporaneous effects of the monetary policy innovations (i.e., the monetary shock itself). What these papers do not address are the long-run objectives and impacts of monetary policy. We investigate those long-run impacts here. Our long-run identification is achieved through the long-run impact matrix of a vector error-correction model (VECM). Short-run identification is achieved by making standard assumptions of how monetary policy impacts other economic variables of interest and by similar assumptions about the information

set of the Fed.

Our approach incorporates regime switches in the long-run relationships through the weighting matrix of the error-correction term. Gregory and Hansen (1996) also looked at regime switches in the cointegrating vector but made no distinction between the weighting matrix and the long-run equilibrium (see also Hall, Psaradakis, and Sola 1997; Clarida, Sarno, Taylor, and Valente 2003; Paap and van Dijk 2003 for recent papers using Markov switching in an error-correction framework). We find it natural to assume that the long-run relationships between the cointegrated variables remain the same across states but that the weighting term is state-dependent. Modeling the weighting matrix as state-dependent allows variables to respond differently to monetary policy shocks, even in the long run. This lends itself to plausible economic interpretation and at the same time preserves the Engle and Granger (1987) notion of cointegration.

The paper proceeds as follows: Section 2 provides a brief motivation for our approach. Section 3 presents a VECM for monetary policy, with Markov switching in the weighting matrix for long-run impacts and regime-dependent heteroskedasticity, and outlines the estimation technique. Section 4 discusses the results of the estimation. Section 5 considers the implications of the switching model in the context of the price puzzle. Section 6 examines the sacrifice ratio and the consequences for disinflationary policy brought about by the presence of the switching process governing the weighting matrix. Section 7 concludes.

2 Motivation

To motivate our approach, we briefly present our empirical model—a more detailed version will be presented below. Our model is the following VECM that allows for different states of the economy:

$$\Delta y_t = c + \sum_{i=1}^k \Gamma_i \Delta y_{t-i} + \Pi_{S_t} y_{t-1} + \varepsilon_t,$$

where S_t denotes the period- t state. In principle, we could allow part or all of the coefficient matrix to switch independently or with the error-correction term. However, we are interested in changes in the adjustment to the long-run equilibrium and, thus, restrict our attention to switching in the error-correction term.

This modeling approach partially overcomes the rational expectations critique of models of this nature. The basic rational expectations argument is that in producing impulse responses, VAR (and structural VAR) models look at responses to shocks that are outside the realm of such models. That is, these models look at responses to shocks assumed not to have happened at the time of modeling. To compound matters, these models further assume that the nonpolicy component of the economy is naive about any change in policy that may have taken place—this is so because the coefficient matrix in the VAR is invariant to any switches in policy regime. Our approach does not suffer the same fate. We not only allow the nature of the shock to vary across states through the state-dependency of the error term, but we also allow individuals’ responses to the structural shocks identified to be state-dependent by allowing for a different coefficient matrix in each of the states. This way individuals “correctly” respond to monetary or any other shocks. Sims (1986) compares the rational expectations approach and the VAR methodology when it comes to forecasting and policy analysis.

Our approach allows for a variety of interpretations regarding the response to monetary policy and is not limited to changes that result only from switches in the policymaker’s decision rule. While we will often, for the purposes of exposition, attribute changes in the “nonpolicy” response to switches in the speed of adjustment of inflation expectations, we allow for alternative interpretations. Specifically, changes in adjustment speed can result from sectoral shifts (Ramey and Shapiro 1998) or changes in policymaker credibility (Faust and Svensson 1998).

3 Model and estimation

This section presents a more detailed version of our modeling approach along with our identification scheme. We adopt some of the more common assumptions used in the identification of the impact matrix of VAR monetary models. The result is a benchmark recursive model. Thus, long-run dynamics obtain (solely) through the error-correction term, Π_{S_t} , which includes both the regime-switching weighting matrix, α_{S_t} , and the regime-invariant cointegrating vector, β . We combine both the short-run dynamics, adopted from standard models in the monetary policy literature, and the aforementioned long-run dynamics in a Markov-switching framework. This allows us to examine the state-dependent responses to monetary policy shocks.

3.1 Model

Consider the following Markov-switching vector error-correction model (MSVECM):

$$\Delta y_t = c + \sum_{i=1}^k \Gamma_i \Delta y_{t-i} + \Pi_{S_t} y_{t-1} + \varepsilon_t, \quad (1)$$

where $\Delta y_t = [\Delta y_{1t}, \Delta y_{2t}, \dots, \Delta y_{mt}]'$ is an m -dimensional vector of differenced variables of interest, c is a vector of intercepts, the Γ_i 's are $m \times m$ parameter matrices, and Σ_{S_t} are state-dependent covariance matrices. Π_{S_t} are the state-dependent long-run impact matrices defined by the $r \times m$ matrix of cointegrating vectors, β , and the $m \times r$ state-dependent weighting matrix, α_{S_t} . Thus, we have

$$\Pi_{S_t} = \alpha_{S_t} \beta'.$$

We assume that S_t is a two-state first-order Markov process in which $S_t \in \{0, 1\}$ is governed by the transition kernel P , where $P_{ij} = \Pr[S_t = i | S_{t-1} = j]$. Rewrite the model (1) as

$$\Delta y_t = c + \sum_{i=1}^k \Gamma_i \Delta y_{t-i} + \alpha_{S_t} \beta' y_{t-1} + \varepsilon_t, \quad (2)$$

where any switch in the cointegrating relationship is restricted to the weighting matrix, α_{S_t} .

The framework (1) can be readily expanded to incorporate switches in the intercept term or the coefficient matrices. We forgo analysis of switching in these components to focus on long-run dynamics. (We considered simultaneous switching in the weighting matrix and the intercept term. We found that including switching in the intercept term reduces the estimated number of periods in the transitory state and prevents reasonable estimates of the model dynamics.)

Switching in Π_{S_t} can be interpreted as switching in the cointegrating vectors, the weighting matrix, or both. We note that these approaches are de facto equivalent. However, our interpretation of switches in the error-correction term implies a single set of long-run relationships and preserves the Engle-Granger notion of cointegration. In our framework, switches can be interpreted as differences in the rate at which the long-run relationships obtain.

We propose that only allowing switches in the weighting matrix is not overly restrictive and provides some further interpretation for our approach. The first reason for confining the nature of the switches is computational convenience. We find that the model becomes intractable if we allow all the coefficients

in the MSVECM to be state-dependent. Allowing for switching in more parameters also confines us to models having fewer variables. We argue that such a sacrifice is not too great since additional variables provide us with more dynamics that would otherwise be absent from a model allowing for more switches but fewer variables. Second, if one were to map back to the reduced form of the model, it would become apparent that differences between states arise through the coefficient on the first lag of the data matrix (this is how the weighting matrix, α_{St} , can be state-dependent while Γ_i is not). For lower-order systems this should prove not too restrictive, since A_1 , the lag-one coefficient from the reduced form of the model, would figure prominently in the error-correction term. The error-correction term is $I - \sum_{i=1}^k A_i$ from the structural model.

We offer as further motivation some other potential interpretations of the form of the model. Our preferred interpretation is that the long-run relationship between the variables, $\beta' y_{t-1}$, is invariant to the state of the economy; however, the weights given to each relationship, α_{St} , are state-dependent. This implies that shocks to the system could potentially have different long-run effects across states, through α_{St} , while maintaining any long-run relationship among the variables. For example, a shock to monetary policy will have different long-run effects on, say, output growth, depending on whether the Fed targeted inflation or output. There are two things at work here: the long-run response to shocks and the cointegrated relationship. We can have the cointegrated relationship, $\beta' y_{t-1}$, unchanged while having different long-run responses to shocks. This is due to the fact that the long-run response coefficient is $\Pi_{St} = \alpha_{St} \beta'$, which is a function of the switching elements (see Hamilton 1994, pp. 579-581, for the long-run response matrix).

The nature of the impulse responses thus depends on the Fed's preference at the time of the shock. A second interpretation is that α_{St} could reflect the rate at which people learn the nature of the shock to the Fed's policy rule. The slower the rate of learning, the greater the impact of shocks to the Fed's policy rule, implying more persistence in the impulse responses. The weighting coefficient could also be interpreted as indicating the amount of relative frictions that exist in each state. In states where there are more frictions, the responses to monetary shocks will be curtailed.

Finally, we cannot rule out the possibility that α_{St} could capture the nature of nonpolicy instruments across states. That is, we should expect different responses to monetary policy if the mixture of nonpolicy instruments changes. This latter interpretation is outside the realm of this paper, as it would

require us to redefine the number of states. That is, the states would be a mixture of both policy and nonpolicy instruments resulting in $s_{policy} * s_{nonpolicy}$ states, instead of s_{policy} states. We ignore this latter interpretation for now and treat all nonpolicy instruments as if they were one.

3.2 Estimation

The data are the monthly coincident indicators index, the personal consumption expenditure (PCE) chain price index, and the federal funds rate from 1960:01 to 2003:08. Each of the first two variables are entered in log levels. The model (2) can be estimated using an iterative three-step Gibbs sampling procedure (e.g., Krolzig 1996, 1997). First, we determine the number of cointegrating relationships using Johansen's maximum eigenvalue procedure. This two-step procedure is adopted from the work of Saikkonen (1992) and Saikkonen and Luukkonen (1997) to obtain estimates of the cointegrating vectors, β . They show that the Johansen procedure estimates consistent cointegrating vectors even in the presence of switching. Then, conditional on these cointegrating vectors, define the $(km + 1 + 2r) \times 1$ vector

$$X_t = \begin{bmatrix} 1 & \Delta y_{t-1} & \cdots & \Delta y_{t-k} & \beta' y_{t-1}(1 - S_t) & \beta' y_{t-1} S_t \end{bmatrix}'.$$

Then, the system (2) can be stacked

$$\mathbf{Y} = \Phi \mathbf{X} + \boldsymbol{\varepsilon}, \tag{3}$$

where

$$\begin{aligned} \mathbf{Y} &= \begin{bmatrix} \Delta y_1 & \Delta y_2 & \cdots & \Delta y_T \end{bmatrix}, \\ \mathbf{X} &= \begin{bmatrix} X_1 & X_2 & \cdots & X_T \end{bmatrix}, \\ \boldsymbol{\varepsilon} &= \begin{bmatrix} \varepsilon_1 & \varepsilon_2 & \cdots & \varepsilon_T \end{bmatrix}, \end{aligned}$$

and the $m \times (km + 1 + 2r)$ parameter matrix is

$$\Phi = \begin{bmatrix} c & \Gamma_1 & \cdots & \Gamma_k & \alpha_1 & \alpha_2 \end{bmatrix}.$$

Given (3), a matrix of cointegrating vectors β , and a series of states $\tilde{S}_T = \left\{ S_1, S_2, \dots, S_T \right\}$, we draw the parameter values from the posterior normal-inverted Wishart distribution with priors ν_{01} , ν_{02} , N_0 , Z_0 , W_{01} , and W_{02} . The priors we use are uninformative. Alternative priors that could take advantage of the Bayesian methodology might employ, for example, the Sims-Zha (1998) prior. In that case, the prior accounts for possible (unestimated) cointegrating vectors. Since we employ uninformative priors, we model the cointegrating vectors explicitly.

At each iteration, Φ , Σ_1 , and Σ_2 can be drawn from a distribution with degrees of freedom ν , precision matrix N , parameter means Z , and covariance matrices W_1 and W_2 , defined by

$$\begin{aligned} \nu_1 &= \nu_{01} + \hat{T}_1, \\ \nu_2 &= \nu_{02} + \hat{T}_2, \\ N &= N_0 + \mathbf{X}'\mathbf{X}, \\ Z &= N^{-1} \left(N_0 Z_0 + \mathbf{X}'\mathbf{X}\hat{Z} \right), \\ W_1 &= \frac{\nu_0}{\nu} W_{01} + \frac{\hat{T}_1}{\nu_1} \hat{\Sigma} + \frac{1}{\nu} \left(\hat{Z} - Z_0 \right)' N_0 N^{-1} \mathbf{X}'\mathbf{X} \left(\hat{Z} - Z_0 \right), \\ W_2 &= \frac{\nu_0}{\nu} W_{02} + \frac{\hat{T}_2}{\nu_2} \hat{\Sigma} + \frac{1}{\nu} \left(\hat{Z} - Z_0 \right)' N_0 N^{-1} \mathbf{X}'\mathbf{X} \left(\hat{Z} - Z_0 \right), \end{aligned} \tag{4}$$

where \hat{T}_1 and \hat{T}_2 are the number of periods in each state, $\hat{Z} = (\mathbf{X}'\mathbf{X})^{-1} \mathbf{X}'\mathbf{Y}$, and $\hat{\Sigma} = \left(\mathbf{Y} - \mathbf{X}\hat{Z} \right)' \left(\mathbf{Y} - \mathbf{X}\hat{Z} \right)$.

The states $\tilde{S} = \left\{ S_1, S_2, \dots, S_T \right\}$ can be drawn from the posterior distribution $p(\tilde{S}_T | \tilde{y}_T, \Phi, \Sigma_1, \Sigma_2)$, which is conditional on the data \tilde{y}_T and the drawn parameters Φ , Σ_1 , and Σ_2 . The posteriors are obtained from

$$p(S_t | \tilde{y}_t, \Phi, \Sigma_1, \Sigma_2) = \frac{f(y_t | \tilde{y}_{t-1}, S_t, \Phi, \Sigma_1, \Sigma_2) p(S_t | \tilde{y}_{t-1}, \Phi, \Sigma_1, \Sigma_2)}{\sum_{S_t} f(y_t | \tilde{y}_{t-1}, S_t, \Phi, \Sigma_1, \Sigma_2) p(S_t | \tilde{y}_{t-1}, \Phi, \Sigma_1, \Sigma_2)}, \tag{5}$$

where

$$p(S_t | \tilde{y}_{t-1}, \Phi, \Sigma_1, \Sigma_2) = \sum_{S_{t-1}} p(S_t | S_{t-1}) p(S_{t-1} | \tilde{y}_{t-1}, \Phi, \Sigma_1, \Sigma_2),$$

and $p(S_{t-1} | \tilde{y}_{t-1}, \Phi, \Sigma_1, \Sigma_2)$ is taken from each previous iteration (see Hamilton 1989 or Kim and Nelson 1999).

The transition probabilities, $p_{ij} = \Pr[S_t = i | S_{t-1} = j]$, are also derived from the estimation algorithm. Though we forgo formal discussion of their estimation, we note that they are drawn from

posteriors formed from beta conjugate distributions.

To satisfy the Lucas critique, we require the state process to depend on the underlying stochastic process, ε (e.g., see Hamilton 1995). We partially accomplish this in two ways. First, the posterior distribution for the state process (5) is a function of the state-dependent variance-covariance matrix. Second, the data used in the filtering of the state process is the same data used in the estimation of the state-dependent coefficient matrices in the VECM.

3.3 Identification

Identification of the model implies two steps: identification of the long-run relationships through the set of cointegrating vectors and identification of the short-run effects through the contemporaneous impact matrix. The former was discussed above. In a three-variable model, our short-run identification consists of a Cholesky ordering of the system with the variables ordered: output, prices, and federal funds rate. This implies the monetary authority takes both output and prices into consideration when setting policy but that policy does not impact output and prices contemporaneously.

We utilize the recursive form of identification in order to evaluate the reasonableness of the results using our methodology with the most basic form of identification. More complicated forms of identification came about because simpler identifying assumptions could not deliver impulse responses that were consistent, or considered consistent, with monetary policy. Our aim is to see how much mileage we get from our methodology with as few structural assumptions as possible.

4 Empirical Results

In this section, we examine monetary shocks in a three-variable cointegrated extension of common VAR models, which account for, among other factors, the changes in inflation expectations. We impose a two-state restriction on the underlying Markov process for tractability. While we recognize this is limiting, we believe it is illustrative of how a model of this nature can partially characterize a solution to the rational expectations critique. As a test, we estimated a three-state model. Under our specification, the filter did not identify enough periods in the third state to estimate that state's parameters with any confidence. We therefore leave higher-order Markov models for future research.

4.1 States

The state process is shown in Figure 1, where the probability that the economy is in state 2 is indicated on the y axis. This regime could be interpreted as a high inflation target or high inflation expectations regime. This figure corresponds to the posterior probabilities governing the weighting matrix in the error-correction term and shows that this cointegration relationship undergoes five significant periods of change. The timing of switches to the high inflation expectations state tends to be correlated with events such as oil shocks and recessions, although not exclusively so. For example, the 1973-75 recession and the Volcker disinflation are clearly identified by the underlying state process.

The majority of the turning points are related to events coincident with large increases in prices. These include the CPI reaching a new peak in 1960, the devaluation of the dollar and a spike in inflation in early 1969, and a farm recession and oil price shock in the mid-1980s. These posterior state probabilities are consistent with the findings of other monetary policy models (Owyang 2002; Sims and Zha 2002) and recession-dating models (Hamilton 1989).

Figure 1 about here

Table 1 about here

Table 1 shows that there are two long-run relationships, with cointegrating vectors β_1 and β_2 , that are fixed across regimes; it also provides the weighting matrices for these relationships that vary across regimes. The absolute size of the weights is greater in state 1 compared with state 2, implying the rate of convergence (or rate of learning) differs across states. Moreover, the nature of the long-run response—or to be more precise, the transition to the long-run equilibrium—associated with the first cointegrating relationship differs across states. Finally, Table 1 provides estimates of the transition probabilities for each state. We note, in particular, that each state, characterized by transition probabilities below 0.1, is relatively persistent.

4.2 Transition Dynamics

To evaluate the effect of a change in state, we conduct the counterfactual state-switching experiments to demonstrate the model's transition dynamics. These switches might be viewed as either the economy's assimilation of a change in the Taylor rule (e.g., Clarida, Gali, and Gertler 2000) or the inflation objective (e.g., Dennis 2001). We initialize the model with 1978 data and set the state to 1. We then switch to state 2 and observe the transition dynamics in the absence of shocks. After 24 periods, we switch the regime back to state 1. This exercise (experiment A) is shown as the solid line in Figure 2. This contrasts with the model remaining in state 1 for all time (experiment B), the dashed line in Figure 2. Although the recession in A is deeper, the policymaker is able to lower prices without exogenous monetary shocks. Moreover, in A, the Fed's policy appears more anti-inflationary at the outset and, thus, produces a steeper recession and more rapid disinflation. Since this outcome is produced in the absence of monetary shocks, there are a limited number of interpretations that can explain the results. Changes in Fed preferences or changes in the inflation target can explain the switch in the expected path of the funds rate.

Figure 2 about here

Figure 3 about here

A second set of experiments demonstrates the effect of remaining in state 2 for all time. Figure 3 shows the results of these experiments—the solid line indicating the case in which the long-run relationship switches to state 1 after 24 periods (experiment C), and the dashed line representing the path governed by state 2 for all time (experiment D). This set of experiments further verifies the nature of the long-run relationships in this model. State 2 represents an upward trend in both prices and output as well as a sustained expansionary monetary policy exhibited by an ever-decreasing funds rate. We can interpret the inflation regime (state 2) as one in which the Fed has either established a new, higher inflation target or is sacrificing price stability in order to increase growth. Further, we can interpret the growth regime (state 1) as the Fed adopting either sustained inflation-neutral or actively contractionary monetary policy under low inflation expectations.

While, in the preceding experiments, it may appear that the regimes are simply manifestations of contractionary and expansionary policy *shocks*, we reiterate that the state changes reflect changes in the *stance* of policy. The idea of a *sustained* policy stance is important to our interpretation. These regimes are not supported by contractionary or expansionary shocks, but by a revision of preferences or expectations. We explore the effect of monetary policy shocks in the next subsection.

4.3 Responses to Policy Shocks

Consider the short-run response to a one-standard-deviation shock to the federal funds rate. These impulse responses are generated conditional on the state—that is, we assume that if the shock is generated in state 1, it is transmitted through state 1. Although we acknowledge this is a restriction, the state-dependent impulse responses and the state transition experiments are sufficient to describe the majority of the model’s short-run dynamics.

Figure 4 about here

The effect of a contractionary monetary policy shock (a 100-basis-point increase in the federal funds rate) is shown in Figure 4. In state 1, the contractionary shock has the anticipated effect on output—the increase in the funds rate induces a recession. The effect on output is relatively weak. The recession is deflationary, causing a reduction in prices over the four-year period.

The contractionary monetary shock in state 2 keeps the funds rate strictly positive for one year. Contrary to state 1, the central tendency is to fully reverse the shock after approximately 30 months. Also in state 2, output is bolstered by the policy reversal and rises (weakly) over 48 months. This causes prices to continue to rise; the policy shock stems inflation for the first few years but the reversal allows prices to continue to increase.

5 The Price Puzzle

Originally identified by Sims (1992), VAR studies of monetary policy have shown that a contractionary shock to the federal funds rate results in a temporary (often lasting a year or more) increase in the aggregate price level. This “price puzzle” has remained a question mark in the identification of monetary

policy shocks, as most researchers believe that identifications that exhibit the phenomenon are incorrect. Sims recognized that including a commodity price index (PCOM) in the estimation eliminates the increase in prices associated with a contractionary monetary shock. Hanson (2002b), however, argues that the addition of PCOM to solve the price puzzle is *ad hoc* and theoretically unappealing.

Proponents of including PCOM might argue that they capture inflation expectations. Our approach is to model inflationary expectations (or at least take them into account) by allowing both shocks (authority side) and responses (agents side) to change accordingly. Modeling the VAR this way allows inflationary expectations to vary across states. Thus, we do not need PCOM to reflect expectations. Since we model both the monetary authority's and respondents' expectations, we have a more theoretically interpretable explanation. In a recent paper, Owyang and Ramey (2004) discovered that switches from an inflation-hawk policy regime to a dove policy regime can lead to the hump-shaped price response that characterizes the price puzzle. We conjecture that the price puzzle may, in fact, stem from the standard VAR's inability to effectively model these types of switches.

Figure 5 about here

We also contend that commodity prices model inflation expectations corresponding to Fed policy. To illustrate this point, we plot the detrended industrial commodity price index against the state-switching process (results were similar when using the Dow Jones commodity price index for a smaller sample). The first panel of Figure 5 shows that the majority of the state switches coincide with (or predate) major commodity price changes in either direction. For example, the switch in the early 70s coincides with a fall in commodity prices, whereas the switch in the early 80s coincides with a rise in prices. The correlation between the state-switching process and the detrended PCOM is 0.34. We view this as significant, given that we restricted our model to only two states of nature and also recognize that not every (major) price change necessarily implies a new state. That is, price changes could also be due to, say, demand factors and are not necessarily limited to policy actions. The second panel of Figure 5 plots the absolute value of the detrended price series against the state-switching process to highlight the fact that switches in states invariably coincide with major price movements.

The conclusion we draw from this simple exercise is that including PCOM in structural VARs was appropriate because they proxy for policy-related changes in inflationary expectations. However,

including PCOM does not lead to an obvious economic interpretation. The problem remains that responses to monetary policy vary depending on the goals of the Fed and how quickly such goals are realized by respondents. That is, traditional VARs provide us with only one response for, say, inflation, even though there are as many responses as there are states (as in the MSVECM framework). It comes as no surprise to us that PCOM plays such a role since, in our opinion, they are the most responsive, and thus the most likely, candidates to reflect the state process. Therefore, considering monetary policy in a state-switching framework resolves the price puzzle without the inclusion of a commodity price index and, at the same time, provides impulse responses that are more accurate representations of the nature of policy and the goals of the Fed.

6 The Cost of Disinflation

A number of studies have used structural models to assess the output or employment loss caused by disinflationary monetary policy (for example, see Okun 1978; Fuhrer 1994, 1995; Cecchetti and Rich 2001). These studies have constructed “sacrifice ratios” that measure the cumulative increase in unemployment or loss of output associated with each percentage point of policy-induced inflation reduction. In particular, there is a rough consensus that a 1 percent reduction in inflation increases cumulative unemployment by about 2 percentage points per year. Cecchetti and Rich (2001) found sacrifice ratios in terms of output loss, over a two-year horizon, ranging from 0.62 to 3.71 using three structural VARs of monetary policy. Recently, Filardo (1998) concluded that the sacrifice ratio varied across different regimes. He interpreted these regimes as growth states in which monetary policy produced different effects.

To calculate the sacrifice ratio, we can assess the output cost of a temporary disinflationary monetary shock within a single regime. Our model has the additional advantage that we can measure the cost of disinflation occurring as a result of switches between regimes. This also allows us to reconcile a number of different, seemingly contrary, facts derived from the literature. First, our model supports multiple “within-regime” sacrifice ratios, which may explain the different numbers others have estimated using structural models with alternative assumptions. This is also consistent with Fuhrer’s (1994, 1995) and Filardo’s (1998) claims of multiple structural breaks in sacrifice ratio estimation. Further, our model reconciles, in an intuitive manner, a low estimated sacrifice ratio without forfeiting the high-output-loss

recessions that were seemingly driven by the Volcker disinflation (see Owyang and Ramey 2004).

We posit two distinct disinflationary episodes: one driven by a policy shock and one driven by a change in regime. If the state truly is a reflection of the underlying inflation expectations, the difference in the two disinflationary forces is apparent. Within regime, the policymaker disinflates under fixed inflation expectations. Fixed expectations build in inflation persistence and, as a result, the disinflation is costly. Across regimes, the policymaker has the benefit of shifting expectations, which can, in and of themselves, assist disinflation. Lower expectations can support a lower inflation rate, and the disinflation is less costly.

Table 3 contains the estimated disinflation costs both within-regime and across states. These disinflationary losses exclude the effects of disinflationary price shocks, which can confound the analysis of the effects of policy. Within-regime sacrifice ratios of 2.16 and 3.60 cumulative percent output lost per percentage point inflation reduced in states 1 and 2, respectively, are consistent with estimates from the aforementioned studies. Moreover, the disinflation costs associated with a switch from state 2 to state 1 are even smaller—1.19 percent output growth lost per percentage point of inflation reduced per year.

The policy implications of the above analysis remain ambiguous, depending on the policymaker’s belief about his effect on the underlying state. In the model, the state follows a Markov process; a natural extension is to assume that the regime follows from a credible switch to a price-stability objective. When the objective is credible, inflation expectations become more downwardly flexible and disinflationary policy can be conducted at a lower cost. However, in practice, a change in inflation expectations may not be easily accomplished (e.g., the change in expectations following the Volcker disinflation may have occurred only after a number of contractionary monetary shocks at a high output cost).

7 Conclusion

We examine monetary policy shocks in a Markov-switching vector error-correction (MSVECM) framework, in which the long-run responses to such shocks (and any other shock for that matter) can vary across states. Long-run variation is achieved by allowing switches in the weighting matrix of the error-correction term, while leaving the cointegrating relationship between variables intact. We suggest that

such an approach to monetary policy is theoretically appealing and goes to the heart of the rational expectations critique of models of this nature.

In particular, we find that a contractionary monetary shock generates different impulses in each state. The nature of the impulse responses is suggestive of the presence of two (unique) types of states, one a growth state and the other a state in which the policymaker cannot credibly commit to low inflation. In the former state, a contractionary shock to monetary policy leads to the usual fall in output and an eventual fall in prices. However, in the latter state the contractionary shock is quickly reversed, resulting in a persistent rise in prices and output.

One implication of our methodology is the absence of the price puzzle commonly found in monetary VAR models. This we accomplished without the need to include a commodity price index in the VAR. We offer as explanation that our state process directly captures inflationary expectations, the role hitherto played by PCOM in previous (one-state) specifications.

Our model also has implications for examining disinflationary monetary policy. Specifically, our model allows us to interpret the cost of disinflation in the context of both contractionary shocks to policy and changes in the stance of policy. We find that although the contractionary shocks produce within-regime sacrifice ratios consistent with the existing literature, regime changes can reduce the inflation rate at a lower cost.

8 Acknowledgements

The authors thank Clive Granger, Jim Hamilton, Mike Hanson, Jeremy Piger, Hailong Qian, Bob Rasche, Lucio Sarno, Martin Sola, Dan Thornton, and an anonymous referee for comments and discussions. This paper benefited from discussions at the Western Economics Association and the Northwestern Econometrics Society Annual Meetings. Abbigail J. Chiodo, Kristie M. Engemann, and Athena T. Theodorou provided invaluable research assistance. The views herein are the authors' alone and do not represent the views of the Federal Reserve Bank of St. Louis or the Federal Reserve System.

References

- [1] Bernanke, B. S., and Mihov, I. (1998), “Measuring Monetary Policy,” *Quarterly Journal of Economics*, 113, 869-902.
- [2] Blanchard, O. J., and Quah, D. (1989), “The Dynamic Effects of Aggregate Demand and Supply Disturbances,” *American Economic Review*, 79, 655-673.
- [3] Boivin, J., and Giannoni, M. (2002), “Has Monetary Policy Become Less Powerful?” Staff Reports No. 144, Federal Reserve Bank of New York.
- [4] Cecchetti, S. G., and Rich, R. W. (2001), “Structural Estimates of the U.S. Sacrifice Ratio,” *Journal of Business & Economic Statistics*, 19, 416-427.
- [5] Christiano, L. J., Eichenbaum, M., and Evans, C. L. (1999), “Monetary Policy Shocks: What Have We Learned and to What End?” in *Handbook of Macroeconomics* (Vol. 1A), eds. J. B. Taylor and M. Woodford, New York: Elsevier Science, pp. 65-148.
- [6] Clarida, R., Galí, J., and Gertler, M. (2000), “Monetary Policy Rules and Macroeconomic Stability: Evidence and Some Theory,” *Quarterly Journal of Economics*, 115, 147-180.
- [7] Clarida, R. H., Sarno, L., Taylor, M. P., and Valente, G. (2003), “The Out-of-Sample Success of Term Structure Models as Exchange Rate Predictors: A Step Beyond,” *Journal of International Economics*, 60, 61-83.
- [8] Dennis, R. (2001), “The Policy Preferences of the US Federal Reserve,” Working Paper No. 2001-08, Federal Reserve Bank of San Francisco.
- [9] Engle, R. F., and Granger, C. W. J. (1987), “Co-integration and Error Correction: Representation, Estimation, and Testing,” *Econometrica*, 55, 251-276.
- [10] Faust, J., and Svensson, L. E. O. (1998), “Transparency and Credibility: Monetary Policy with Unobservable Goals,” International Finance Discussion Papers No. 1998-605, Board of Governors of the Federal Reserve System.
- [11] Filardo, A. J. (1998), “New Evidence on the Output Cost of Fighting Inflation,” Federal Reserve Bank of Kansas City *Economic Review*, 83, 33-61.

- [12] Fuhrer, J. C. (1994), "Optimal Monetary Policy and the Sacrifice Ratio," in *Goals, Guidelines, and Constraints Facing Monetary Policymakers: Proceedings from Conference Series No. 38*, Federal Reserve Bank of Boston, pp. 43-69.
- [13] ——— (1995), "The Persistence of Inflation and the Cost of Disinflation," *New England Economic Review*, 3-16.
- [14] Gregory, A. W., and Hansen, B. E. (1996), "Residual-Based Tests for Cointegration in Models with Regime Shifts," *Journal of Econometrics*, 70, 99-126.
- [15] Hall, S. G., Psaradakis, Z., and Sola M. (1997), "Cointegration and Changes in Regime: The Japanese Consumption Function," *Journal of Applied Econometrics*, 12, 151-168.
- [16] Hamilton, J. D. (1989), "A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle," *Econometrica*, 57, 357-384.
- [17] ——— (1994), *Time Series Analysis*, Princeton: Princeton University Press.
- [18] ——— (1995), "Rational Expectations and the Economic Consequences of Changes in Regime," in *Macroeconometrics: Developments, Tensions and Prospects*, ed. K. D. Hoover, Boston: Kluwer Academic, pp. 325-344.
- [19] Hanson, M. S. (2002a), "Varying Monetary Policy Regimes: A Vector Autoregressive Investigation," unpublished manuscript, Wesleyan University.
- [20] ——— (2002b), "The 'Price Puzzle' Reconsidered," unpublished manuscript, Wesleyan University.
- [21] Kim, C.-J., and Nelson, C. (1999), *State-Space Models with Regime Switching: Classical and Gibbs-Sampling Approaches with Applications*, Cambridge, MA: MIT Press.
- [22] Krolzig, H.-M. (1996), "Statistical Analysis of Cointegrated VAR Processes with Markovian Regime Shifts," SFB 373 Discussion Paper 25/1996, Humboldt University Berlin.
- [23] ——— (1997), *Markov-Switching Vector Autoregressions: Modelling, Statistical Inference, and Application to Business Cycle Analysis*, Lecture Notes in Economics and Mathematical Systems (Vol. 454), New York: Springer.

- [24] Okun, A. M. (1978), "Efficient Disinflationary Policies," *American Economic Review*, 68, 348-352.
- [25] Owyang, M. T. (2002), "On the Robustness of Policy Asymmetries," Working Paper No. 2002-018, Federal Reserve Bank of St. Louis.
- [26] Owyang, M. T., and Ramey, G. (2004), "Regime Switching and Monetary Policy Measurement," *Journal of Monetary Economics*, forthcoming.
- [27] Paap, R., and van Dijk, H. K. (2003), "Bayes Estimates of Markov Trends in Possibly Cointegrated Series: An Application to U.S. Consumption and Income," *Journal of Business & Economic Statistics*, 21, 547-563.
- [28] Ramey, V. A., and Shapiro, M. D. (1998), "Capital Churning," unpublished manuscript, University of California at San Diego.
- [29] Saikkonen, P. (1992), "Estimation and Testing of Cointegrated Systems by an Autoregressive Approximation," *Econometric Theory*, 8, 1-27.
- [30] Saikkonen, P., and Luukkonen, R. (1997), "Testing Cointegration in Infinite Order Vector Autoregressive Processes," *Journal of Econometrics*, 81, 93-126.
- [31] Sims, C. A. (1986), "Are Forecasting Models Usable for Policy Analysis?" *Federal Reserve Bank of Minneapolis Quarterly Review*, 10, 2-16.
- [32] ——— (1992), "Interpreting the Macroeconomic Time Series Facts: The Effects of Monetary Policy," *European Economic Review*, 36, 975-1011.
- [33] Sims, C. A., and Zha, T. A. (1998), "Does Monetary Policy Generate Recessions?" Working Paper No. 98-12, Federal Reserve Bank of Atlanta.
- [34] ——— (2002), "Macroeconomic Switching," unpublished manuscript, Princeton University.

Table 1: Estimation Results

		Y	P	R
cointegrating vectors ^a	β_1^b	-2.137	0.1099	0.1904
	β_2^c	0.3584	-0.1958	0.2755
weighting matrices ^d	$\alpha_1(S_t = 1)$	0.0067 (0.037)	0.1044 (0.062)	-0.2729 (0.217)
	$\alpha_1(S_t = 2)$	-0.022 (0.041)	-0.0971 (0.069)	0.4275 (0.244)
	$\alpha_2(S_t = 1)$	-0.0002 (0.043)	-0.0441 (0.068)	0.1101 (0.227)
	$\alpha_2(S_t = 2)$	-0.0161 (0.042)	-0.0267 (0.069)	0.099 (0.239)
transition probabilities	Pr [$S_t = 1 \mid S_{t-1} = 2$]		0.0218	
	Pr [$S_t = 2 \mid S_{t-1} = 1$]		0.0874	

^a unnormalized cointegrating vectors from Johansen's max eigenvalue procedure

^b null of no cointegrating vector rejected at 5%

^c null of only one cointegrating vector rejected at 5%

^d 60% coverage intervals in parentheses

Table 2: Costs of Disinflation

	<i>Sacrifice Ratio*</i>
Within State 1	2.16
Within State 2	3.6
From State 1 to State 2	1.33
From State 2 to State 1	1.19

* percent growth lost per percentage point inflation
reduction per year at 5-year horizon

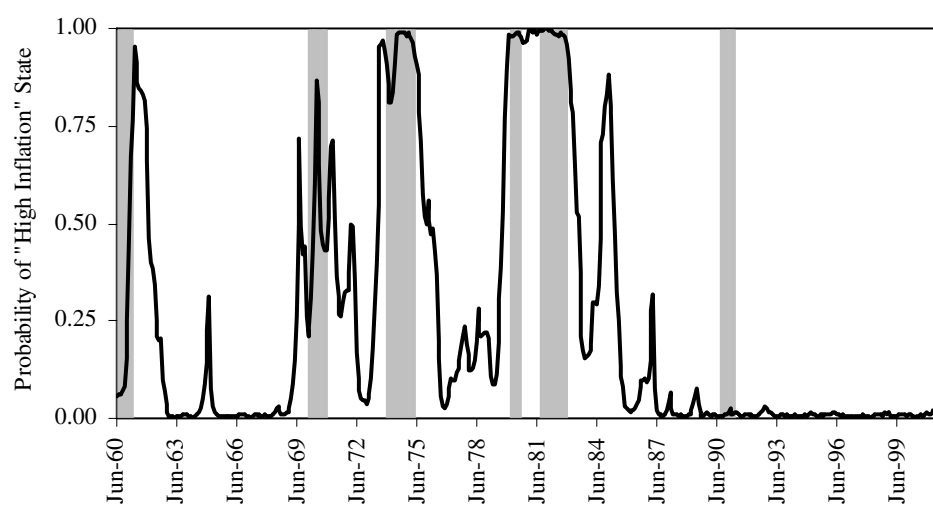


Figure 1. State 2 Posterior Probabilities: Smoothed Probabilities, $Pr [S_t=2 \mid Y_1, \dots, Y_t]$ (solid line); NBER recession (shaded regions).

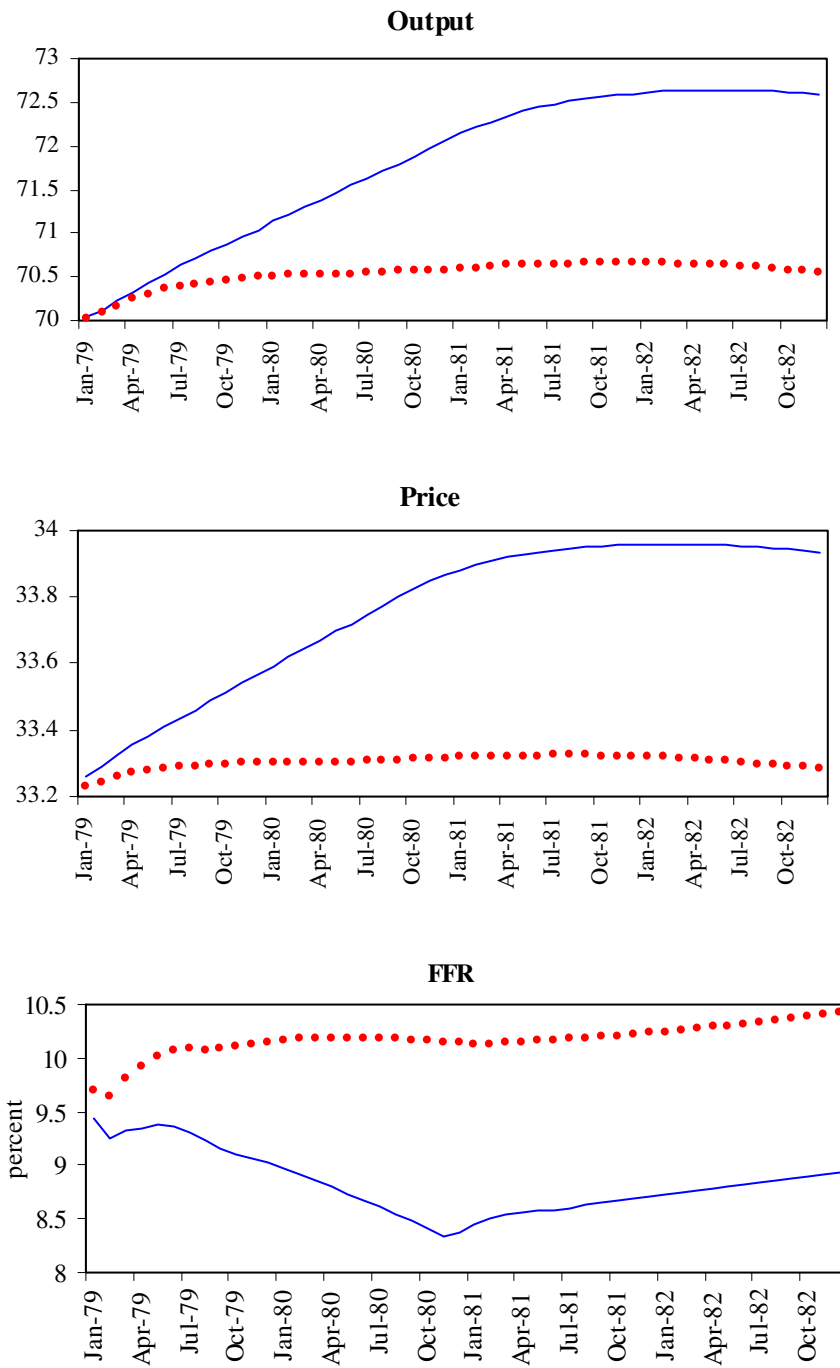


Figure 2. State Transition Experiments 1: Experiment A (switches from State 2 to State 1 in January 1981, solid line); Experiment B (remains in State 1 for all time, dotted line).

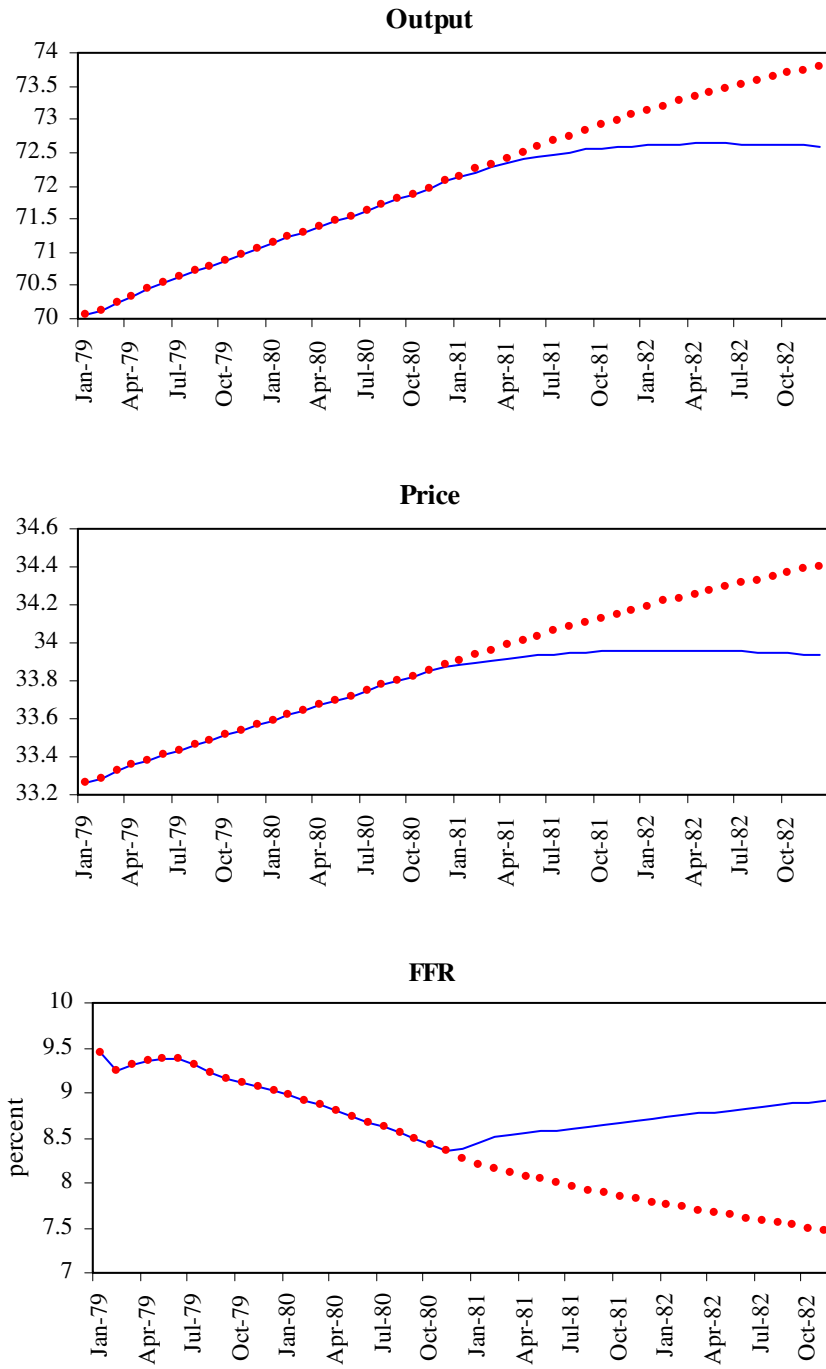


Figure 3. State Transition Experiments 2: Experiment C (switches from State 2 to State 1 in January 1981, solid line); Experiment D (remains in State 2 for all time, dotted line).

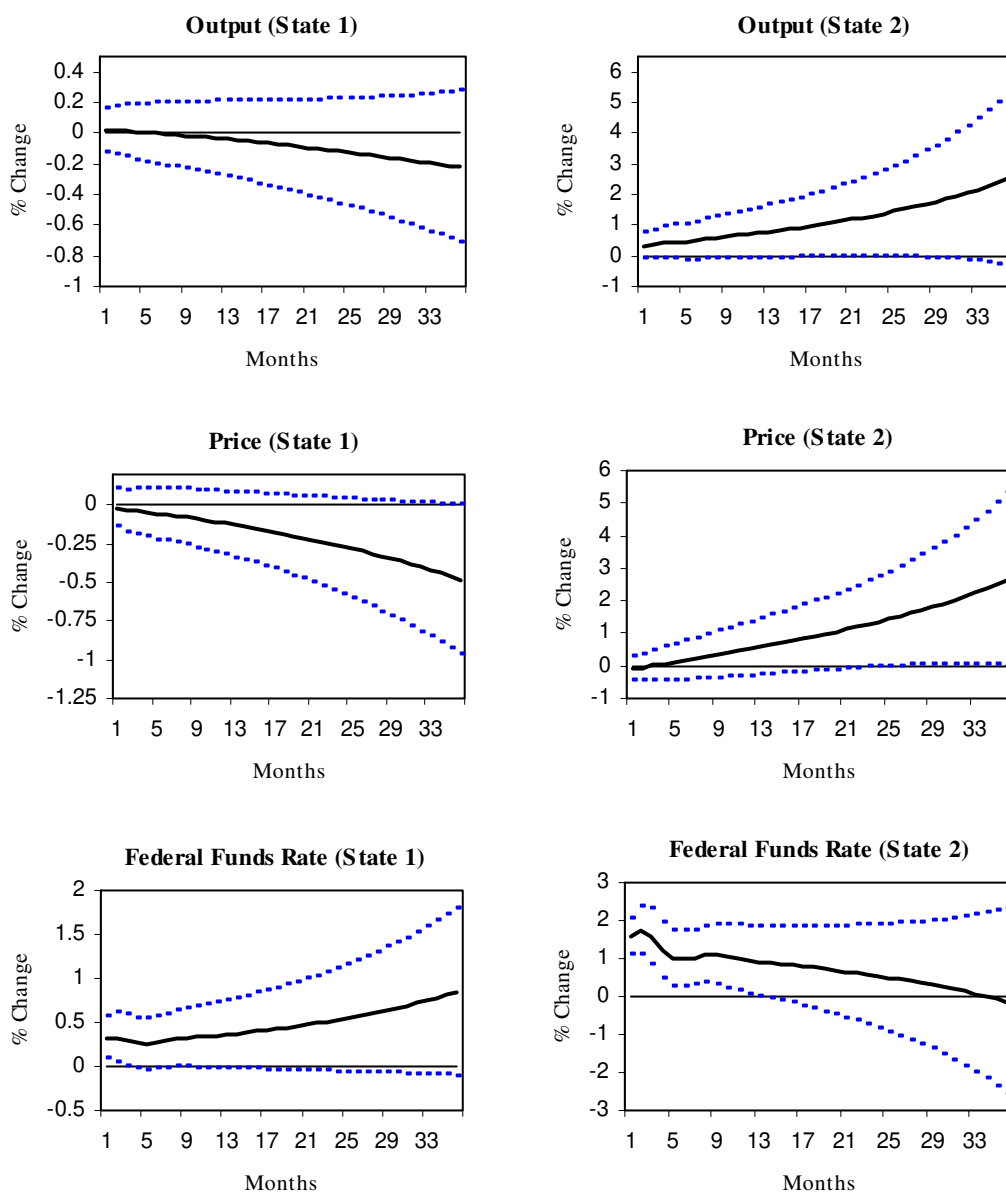


Figure 4. Impulse Responses: 60% coverage intervals are shown.

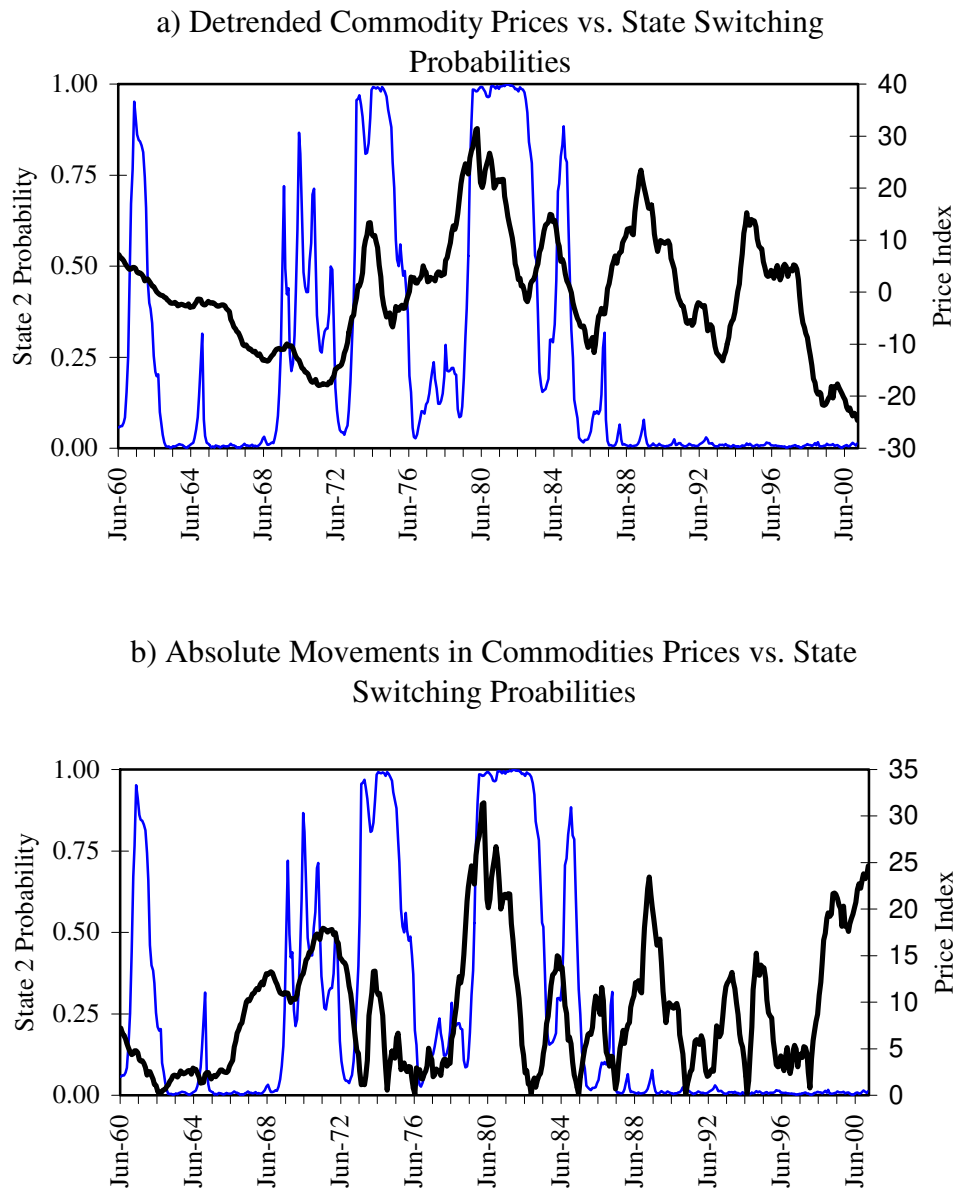


Figure 5. Graphs plotting the posterior probability of State 2 against a) (linear) detrended commodity prices (bold line) with Correlation = 0.34, and b) the absolute movements in commodities prices around its trend (bold line), with Correlation = 0.14. Commodities prices are taken from the Journal of Commerce's industrial price index.