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# Is the Response of Output to Monetary Policy Asymmetric? Evidence from a Regime-Switching Coefficients Model

This paper investigates regime switching in the response of U.S. output to a monetary policy action. We find substantial, statistically significant, time variation in this response that corresponds to "high response" and "low response" regimes. We then investigate whether the timing of the regime shifts are consistent with three particular manifestations of asymmetry by modeling the transition probabilities governing the switching process as functions of state variables. We find strong evidence that policy actions taken during recessions have larger effects than those taken during expansions. We find less evidence of asymmetry related to the direction or size of the policy action.

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SINCE AT LEAST the Great Depression, economists have argued that the impact of monetary policy actions on the real economy is not symmetric with respect to economic conditions or the nature of the policy action. In recent years, interest in such asymmetry has experienced a resurrection, as evidenced by a growing body of empirical work.[1] This literature has focused on three

1. See, for example, Belongia (1996), Cover (1992), DeLong and Summers (1988), Evans (1986), Garcia and Schaller (2002), Karras (1996), Karras and Stokes (1999), Kaufmann (2002), Morgan (1993), Parker and Rothman (2004), Peersman and Smets (2002a, 2002b), Ravn and Sola (1999), Rhee and Rich (1995), Senda (2001), Thoma (1994), and Weise (1999).

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particular manifestations of asymmetry: (1) asymmetry related to the direction of the monetary policy action, (2) asymmetry related to the existing business cycle phase, and (3) asymmetry related to the size of the policy action.[2]

Such asymmetry implies time variation in the coefficients measuring the response of output to a monetary policy action. In this paper, we use a Markov regime-switching model to capture this time variation. To investigate whether the time variation is consistent with specific asymmetries, we define state variables by which side of a dividing line some characteristic of an indicator variable, which could be the monetary policy action itself, falls. We then allow the transition probabilities governing the regime-switching process to be a function of these state variables. We use this time-varying transition probability (TVTP) framework to evaluate the evidence for each of the three asymmetries mentioned above.

The results suggest substantial, statistically significant, time variation in the coefficients describing the response of output to a monetary policy action, and that this time variation corresponds to "high response" and "low response" regimes. We find strong evidence that the time variation can be explained by a state variable indicating whether the economy is in a recession at the time the policy action was taken. In particular, policy actions taken during recessions seem to have larger effects than those taken during expansions. We find that this result is robust across the historical record of business cycles, that is, it does not appear to be driven by only a small subset of recessions. We find much less evidence of any asymmetry related to the nature of the policy action, such as its direction or size.

This paper attempts to advance the existing empirical literature in three ways. First, that literature tests for asymmetry by allowing the coefficients of an equation linking real activity to policy variables to be state dependent, where the states are linked to a particular asymmetry.[3] Our approach instead focuses on modeling the time variation in the coefficients linking output to policy, without forcing that time variation to correspond to a particular asymmetry. Indeed, the TVTP framework we use is capable of capturing the coefficient time variation in the data, even if all of the state variables we use to explain this time variation are statistically insignificant. An advantage of this approach is that it allows us to evaluate the robustness of the evidence for any particular asymmetry, as the estimated pattern of time variation can be used to determine whether that evidence is coming from only a small subset of episodes, or is robust across the historical record.

A second contribution of this paper is its focus on several manifestations of asymmetry and measures of monetary policy. As opposed to much of the literature, which investigates only one asymmetry at a time, we investigate all three asymmetries discussed above. This is useful as it is likely that these asymmetries are correlated,

2. A smaller literature investigates asymmetry in the effects of monetary policy depending on the level of economic activity relative to trend. We do not investigate this type of asymmetry in detail in this paper.

3. For example, Cover (1992) and DeLong and Summers (1988) investigate asymmetry in the effects of policy stimulus vs. contraction by regressing output growth on positive and negative policy shocks. They then test the null hypothesis that the parameters on the two types of shocks are equal.

making it difficult to determine the "true" asymmetry from independent investigations of each. We also consider different measures of monetary policy actions, including monetary policy "shocks" obtained from identified VARs and an endogenous measures of monetary policy, namely the change in the real federal funds rate.

Finally, we investigate these issues using an unobserved-components decomposition of real output into trend and cyclical components. The structural representation in terms of trend and transitory components allows for the introduction of monetary policy variables such that policy actions have only short-run effects on the economy. This is as opposed to the majority of the literature, which generally proceeds by regressing output growth on measures of policy actions.

Our approach is closest in spirit to Garcia and Schaller (2002), Kaufmann (2002), Peersman and Smets (2002a), Ravn and Sola (1999), Thoma (1994), and Weise (1999). As in our model, the first four papers employ a regime-switching framework to investigate asymmetry in the effects of monetary policy. However, each ties the regime switching on the coefficients of the policy variables to a particular manifestation of asymmetry. In the first three papers, the regime switching is connected to switching of the economy from boom to recession, in order to investigate asymmetry related to the business cycle. In Ravn and Sola (1999), the regime switching is connected to switching in the variance of monetary policy shocks, in order to capture large vs. small monetary policy shocks. By contrast, in this paper, we do not force the regime switching to correspond to any particular asymmetry. Weise's (1999) is the only paper we are aware of to jointly evaluate evidence for all three asymmetries discussed above. While Weise considers money-based indicators of monetary policy, here we also evaluate evidence for multiple asymmetries using interest-rate-based measures.[4] Thoma (1994) uses plots of the *p*-values from rolling sample Granger causality tests to document the timing of instabilities in the relationship between monetary policy indicators and future output.[5] Here, we alternatively use a regime-switching model to investigate this issue. Finally, none of the above studies incorporate the unobserved-components framework employed here to constrain monetary policy actions to have only short-run effects.

Asymmetry in the real effects of monetary policy can be motivated by a variety of theoretical models. First, models generating asymmetry in the rigidity of prices, specifically prices that are more rigid downward than upward, are capable of generating asymmetries in the effects of contractionary and expansionary policy.[6] Here, a positive shift in aggregate demand is primarily reflected in prices, while a negative shift

4. In addition, Weise (1999) uses a smooth transition threshold VAR, whereas the regime-switching model we use in this paper assumes discrete regime shifts. For the three specific types of asymmetry we discuss in this paper, the substantial majority of the literature has focused on sharp regime definitions: positive monetary policy shocks vs. negative, expansions vs. recessions, "big" monetary policy shocks vs. "small." Thus, the regime-switching framework we use here is partly motivated as an effort to be consistent with this literature. Also, using a Monte Carlo experiment, the results of which are available from the authors on request, we have found that the regime-switching framework provides sensible estimates of the within-regime parameter estimates even when the data are generated from Weise's smooth transition model.

5. See also Thoma and Gray (1998).

6. See, for example, Ball and Mankiw (1994) and Senda (2001).

is primarily reflected in output, implying that a policy contraction is more effective than policy stimulus. Second, theories of a credit channel through which monetary policy affects output can predict larger effects of monetary policy in unfavorable growth states. Bernanke and Gertler (1995) review one strain of this literature, which they call the "balance-sheet" channel. Here, changes in short-term interest rates affect not only the cost of capital, but also the external finance premium. Monetary policy then has its largest effects at those times when the balance-sheet channel augments the traditional interest rate channel to the greatest extent. For a variety of reasons, this is likely to be during periods of unfavorable growth states, for example during recessions. During such periods, a greater proportion of firms rely on external financing, and the external finance premium is larger. Finally, many models predict asymmetry related to the size of the monetary policy action. For example, menu cost models predict that only small policy actions have large effects, since a large shock makes paying the menu cost optimal.

In the next section, we formally describe the model to be estimated. Section 2 discusses the results and their implications for the nature of asymmetry in the effects of monetary policy actions. Section 3 concludes.

## 1. EMPIRICAL MODEL

The model we consider is an unobserved-components model with regime switching:

$$y_t = y_t^P + y_t^T , \tag{1}$$

$$y_t^P = \mu_t + y_{t-1}^P + v_t , \tag{2}$$

$$\mu_t = \mu_{t-1} + \omega_t , \tag{3}$$

$$\phi(L)\cdot y_t^T = \gamma_0(L)\cdot x_t + \gamma_1(L)\cdot x_t \cdot S_t + \varepsilon_t , \tag{4}$$

$$\phi(L) = \sum_{k=0}^{K} \phi_k \cdot L^k ; \ \phi_0 = 1; \ \gamma_i(L) = \sum_{j=1}^{J} \gamma_{j,i} \cdot L^j , \tag{5}$$

where $y_t$ is the log level of output, $x_t$ is a scalar variable measuring monetary policy, and $S_t$ is a regime-indicator variable taking on the values 0 or 1.

Ignoring $x_t$ in Equation (4), the model in Equations (1–5) is simply the unobserved-components decomposition of real output into stochastic trend component, $y_t^P$, and transitory component, $y_t^T$, discussed in Clark (1987) and Watson (1986). The stochastic trend is specified as a random walk with a time-varying drift term, $\mu_t$, which evolves as a driftless random walk. This specification for the drift is used to capture low frequency innovations to the stochastic trend such as structural breaks in trend growth rate. The transitory component is modeled as an autoregressive process in which all roots of $\phi(L)$ lie outside the unit circle. The innovations $v_t$, $\omega_t$, and $\varepsilon_t$ are assumed to be normally distributed, i.i.d. random variables.

As in Gerlach and Smets (1999), we augment this standard unobserved-components model with a monetary policy variable, $x_t$. We assume that monetary policy has no long run real effect, and thus allow $x_t$ to affect only the transitory component. As is discussed below, the primary measure of monetary policy we consider is a policy shock from a recursive VAR in which the policy variable is ordered after output. In order to be consistent with this identifying restriction, $x_t$ does not enter Equation (4) contemporaneously.

To capture time variation in the response of the transitory component to the policy variable, we allow for time variation in the coefficients linking $y_t^T$ and $x_t$, hereafter called the "response coefficients." In particular, the response coefficients vary between two regimes, with the regime indexed by the indicator variable, $S_t$. We assume that $S_t$ is unobserved by the econometrician, and thus must be filtered from the data. This requires an assumption regarding the evolution of $S_t$, which we satisfy by specifying $S_t$ as a first-order Markov process as in Hamilton (1989). The simplest version of this process specifies that $S_t$ switches between 0 and 1 in accordance with the following fixed transition probabilities (FTP):

$$
\begin{aligned}
&P(S_t = 0 \,|\, S_{t-1} = 0, S_{t-2}, S_{t-3}...) = P(S_t = 0 \,|\, S_{t-1} = 0) = \frac{\exp(c_0)}{(1 + \exp(c_0))} \\
&P(S_t = 1 \,|\, S_{t-1} = 0) = 1 - P(S_t = 0 \,|\, S_{t-1} = 0) \\
&P(S_t = 1 \,|\, S_{t-1} = 1, S_{t-2}, S_{t-3}...) = P(S_t = 1 \,|\, S_{t-1} = 1) = \frac{\exp(c_1)}{(1 + \exp(c_1))} \\
&P(S_t = 0 \,|\, S_{t-1} = 1) = 1 - P(S_t = 1 \,|\, S_{t-1} = 1)
\end{aligned}
\tag{6}
$$

The model in Equations (1–6) is capable of capturing shifts in the response coefficients. However, a primary goal of this exercise is to determine not just the timing of the shifts in the response coefficients, but also whether the three asymmetries discussed in the introduction to this study—asymmetry related to the direction of the policy action, asymmetry related to the size of the policy action, and asymmetry related to the position of the business cycle—are able to explain these shifts. To this end, we augment the model to allow $S_t$ to depend on state variables linked to each of these asymmetries. Specifically, we modify Equation (6) to allow the transition probabilities of the regime-switching process to be time varying, where the time variation depends on state variables:

$$
\begin{aligned}
&P(S_t = 0 \,|\, S_{t-1} = 0) = \frac{\exp(c_0 + \mathrm{z}_t^{'} \cdot a_0)}{(1 + \exp(c_0 + z_t^{'} \cdot a_0))} \\
&P(S_t = 1 \,|\, S_{t-1} = 1) = \frac{\exp(c_1 + \mathrm{z}_t^{'} \cdot a_1)}{(1 + \exp(c_1 + z_t^{'} \cdot a_1))}
\end{aligned}
\tag{7}
$$

Here, $z_t$ is a $q \times 1$ vector of state variables, $(z_{1t}, z_{2t}, ..., z_{qt})'$, while $a_0$ and $a_1$ are $q \times 1$ vectors of coefficients, $(a_{01}, a_{02}, \ldots, a_{0q,})'$ and $(a_{11}, a_{12}, \ldots, a_{1q},)'$. The time-varying

transition probability (TVTP) specification in Equation (7) has been used in a variety of contexts (see, for example, Diebold, Lee, and Weinbach, 1994, Filardo, 1994).

The state variables we include in $z_t$ are designed to be consistent with the existing literature testing for asymmetry. Specifically, we include dummy variables in $z_t$ that describe the size and sign of the policy action, as well as the business cycle phase at the time the policy action is taken. Because these dummy variables are meant to describe conditions prevailing at the time of the policy action, we include $J$ lags of each dummy variable in $z_t$, where $J$ is the number of lags of the policy variable that enter Equation (4). As is the case for the policy variable in Equation (4), the contemporaneous values of the dummy variables are not included in $z_t$.

To specify the state variable related to the direction of the policy action, we define a dummy variable, $D\text{sign}_t$, which is zero if the policy action taken at time $t$ was accommodative, and one if the policy action was contractionary. This categorization will be defined by the sign of the policy action—for example, if the policy action is an interest-rate-based monetary policy shock, a negative shock is defined as accommodative and a positive shock is defined as contractionary. The state variable for asymmetry related to the size of the policy action is given by the dummy variable, $D\text{size}_t$, which is zero if the policy action taken at time $t$ is within one standard deviation of its historical mean, and one otherwise. Finally, the state variable for asymmetry related to business cycle phase is given by the dummy variable $D\text{rec}_t$, which is zero if the economy is in an expansion, as defined by the National Bureau of Economic Research (NBER), and one if the economy is in a recession.

## 2. ESTIMATION RESULTS

In this section, we discuss estimation results for the model presented in Section 1. For maximum likelihood estimation, we cast the model in state-space form and apply the Kim (1994) filter. This procedure is described in detail in Kim and Nelson (1999).

### *2.1 Measurement of Output and Monetary Policy*

We measure real output, $y_t$, as the logarithm of quarterly U.S. industrial production.[7] This data series, as well as all others, is from the Haver Analytics Database. For the monetary policy variable, $x_t$, we construct an interest-rate-based monetary

7. This is a narrow measure of output, made up of the output of the manufacturing, mining, and public utilities sectors, with manufacturing composing approximately 85% of the index. We use industrial production for two reasons. The first is to more tightly identify the effects of monetary policy on the economy. As Christiano, Eichenbaum, and Evans (1997) point out, the manufacturing sector tends to react to a greater extent to a monetary policy shock than economy-wide measures of output. For this reason, much of the recent literature searching for asymmetry in the effects of monetary policy focuses on industrial production. The desire to compare our results with that of this literature provides a second reason to employ industrial production as the measure of output.

policy shock from an identified VAR.[8] The VAR contains three variables: the Federal Funds rate, the logarithm of real Gross Domestic Product (GDP), and the logarithm of the GDP price deflator.[9] To identify the policy shock, we make the assumption that monetary policy shocks do not affect real output or the price level contemporaneously, that is the policy variable is ordered after the output variable in the VAR.[10] The VAR is estimated over the period from the third quarter of 1954 to the fourth quarter of 2002. Four lags of each variable are used, meaning the first policy shock recovered is for the third quarter of 1955. Because the transition equation for the state-space representation of the model in Equations (1–5) is non-stationary, an unconditional expectation of the transition equation to initialize the Kalman filter portion of the Kim (1994) filter is not available. We thus initialize the filter with guesses on which we place high variance, and begin computing the likelihood function only after five years of data have passed to allow the effects of these initial guesses to dissipate.[11] Thus, all the outputs from the model will cover the third quarter of 1960 to the fourth quarter of 2002.

As a robustness check on our results, we consider two alternative measures of $x_t$: an endogenous measure of policy, namely the change in the ex-post real interest rate, and a monetary policy shock where the M1 money supply is the monetary policy instrument. As is discussed in Section 2.5, both of these alternative policy variables yield results similar to those based on the interest-rate-based monetary policy shock.

### *2.2 Is the Regime Switching Significant?*

In this section, we use statistical tests to determine whether the model presented in Section 1 provides a significant improvement in model fit over a model with constant response coefficients. To begin, we must first specify the lag orders, $K$ and $J$. For all the versions of the model that we consider, we employ a backward lag order selection methodology in which we set a maximum lag order of four for both $K$ and $J$ and work backwards until a likelihood ratio test finds a significant value of either $\phi_k$ or $\gamma_{j,i}$. This procedure never chooses a lag order greater than two, thus, for the results presented in the remainder of the paper, $K$ and $J$ are both set equal to two. As a supplement to the likelihood ratio tests, we also performed diagnostic tests

8. The monetary policy shock from the VAR is a generated regressor in the estimation of the unobserved-components model in Equations (1–5). The standard errors of our parameter estimates ignore this fact, and thus are likely to understate the true uncertainty associated with our parameter estimates. In Section 2.5, we present results for an alternative policy measure that is not a generated regressor, and the results are very similar.

9. We also considered a policy shock from a four variable VAR that includes the logarithm of the M1 money supply. Results for this policy shock were very similar to those obtained from the three variable VAR.

10. The policy rule equation of the VAR from which the policy shock is recovered is linear. Thus, while we are allowing for asymmetry in the response of the economy to a policy shock, we do not allow for any asymmetry in the response of monetary policy to economic conditions.

11. Kim and Nelson (1999) provide detailed discussion of the issues surrounding initialization of the Kalman filter.

on the residuals of the estimated model. These diagnostics produced no evidence of remaining serial correlation or heteroskedasticity when $K = J = 2$.[12]

We next turn to evaluating the significance of the regime-switching model over one with constant response coefficients. That is, we test the null hypothesis that $\gamma_{j,0} = \gamma_{j,1}$ for all $j$. It is well known that the standard likelihood ratio test of this null hypothesis does not have the usual $\chi^2$ distribution, as there are nuisance parameters that are unidentified under the null hypothesis. Several authors, for example, Garcia (1998) and Hansen (1992), have developed alternative tests of the null hypothesis of parameter constancy against the alternative of a model with regime switching, where the regime switching is characterized by fixed transition probabilities such as those given in Equation (6). Here, we use the testing procedure developed by Hansen (1992) to test the significance of the FTP model given in Equations (1–6) against the null hypothesis of constant response coefficients. The Hansen (1992) procedure provides an upper bound of the $p$-value for this null hypothesis, and as a result is generally thought to provide a conservative test of the null hypothesis.[13]

When applied to our model, the Hansen test yields a $p$-value of 0.11. Given that this is an upper bound, we interpret this as a relatively strong evidence in favor of the model with regime-switching response coefficients. As will be discussed further in Section 2.5, the alternative measures of $x_t$ that we consider provide even stronger evidence in favor of the regime-switching model. For example, when $x_t$ is measured by the change in the ex-post real Federal Funds rate, the $p$-value is 0.03, while when $x_t$ is measured by a money-based monetary policy shock, the $p$-value is 0.04.

### *2.3 Which Asymmetries Explain the Timing of the Regime Shifts?*

Given the evidence of instability in the response coefficients, we are next interested in whether this instability corresponds to one or more of the three specific manifestations of asymmetry described above. To this end, we estimate the TVTP model in Equations (1–5) and (7) and investigate the significance of the dummy variables described in Section 1 when they are included in $z_t$.

First, as will be discussed further in the next section, estimation results from all the versions of the model that we consider suggest that $S_t$ is equal to one in only short bursts, generally lasting only a single quarter. Given this, there is very little variation in $P(S_t = 1 \mid S_{t-1} = 1)$ for $z_t$ to explain, suggesting that $P(S_t = 1 \mid S_{t-1} = 1)$ is best modeled as a fixed parameter. Indeed, in none of the estimated models was the coefficient vector $a_1$ significant using a likelihood ratio test. Thus, in all of the versions

12. Specifically, neither the Breusch–Godfrey test for the null of no serial correlation nor a Lagrange multiplier test for the null of no ARCH effects rejects at any reasonable significance level when applied to the estimated residuals.

13. The Hansen procedure involves the evaluation of the constrained likelihood over a grid of values for the nuisance parameters of the model, which in this case are $c_0$, $c_1$, $\gamma_{1,1}$, $\gamma_{2,1}$. The grid was chosen by searching in a range around the maximum likelihood estimates. The grid for $c_0$ and $c_1$ is such that $P(S_t = 0 \mid S_{t-1} = 0)$ varies from 0.6 to 0.9 and $P(S_t = 1 \mid S_{t-1} = 1)$ varies from 0.05 to 0.35, each in increments of 0.15. The grid for $\gamma_{1,1}$ varies from 0 to 0.15 and for $\gamma_{2,1}$ it varies from $-0.25$ to $-0.05$, each in increments of 0.05.

of the model that we discuss below, $a_1$ is set equal to zero and we focus only on modeling time variation in the transition probability $P(S_t = 0 \mid S_{t-1} = 0)$.

Table 1 presents summary statistics of the model's fit for various specifications of $z_t$. The first part of Table 1 describes the model in which $z_t$ is empty, which is the FTP model, as well as models in which the three dummy variables described in Section 1 are included in $z_t$ one at a time. We first consider the case where $z_t$ contains the dummy variable capturing the direction of the policy action, that is $z_t = (D\text{sign}_{t-1}, D\text{sign}_{t-2})'$. There is very little evidence that the direction of the policy action is helpful for explaining regime shifts. The likelihood ratio statistic for a test of the null of the FTP model is 0.65, with an associated $p$-value of 0.86. Further, both the SIC and AIC prefer the FTP model. Thus, these results do not provide much evidence of asymmetry related to the direction of the policy action.

We next consider the case where $z_t$ contains the dummy variable capturing the size of the policy action, that is $z_t = (D\text{size}_{t-1}, D\text{size}_{t-2})'$. Here, there is more evidence of asymmetry. The likelihood ratio statistic for a test of the null of the FTP model has a $p$-value of 0.10, and the AIC, but not the SIC, prefers the TVTP model. Finally, we consider the case where $z_t$ contains the dummy variable indicating NBER recession dates, that is $z_t = (D\text{rec}_{t-1}, D\text{rec}_{t-2})'$. Here, there is a very strong evidence of asymmetry. The likelihood ratio statistic for a test of the null of the FTP model has a $p$-value that is zero to the third decimal place, and both the SIC and AIC prefer the TVTP model.

To investigate the robustness of these results, we conduct two additional model specification experiments. First, we investigate whether the *D*size dummy variable retains its significance when it is included jointly in the model with *D*rec, that is we investigate the significance of $D\text{size}_{t-1}, D\text{size}_{t-2}$ when $z_t = (D\text{rec}_{t-1}, D\text{rec}_{t-2}, D\text{size}_{t-1}, D\text{size}_{t-2})'$. The second part of Table 1 presents the details for this model. The $p$-value for the likelihood ratio test of the null hypothesis that $z_t = (D\text{rec}_{t-1}, D\text{rec}_{t-2})'$, is 0.28 and the SIC and AIC both prefer this simpler model.

TABLE 1

MODEL SELECTION FOR TIME-VARYING TRANSITION PROBABILITY SPECIFICATION (MONETARY POLICY MEASURE: FEDERAL FUNDS RATE BASED MONETARY POLICY SHOCK)

| Elements in $z_t$ | SIC | AIC | Log likelihood |
|---|---|---|---|
| None | −5.742 | −5.945 | 516.282 |
| *D*sign | −5.686 | −5.926 | 516.713 |
| *D*size | −5.709 | −5.949 | 518.619 |
| *D*rec | −5.777 | −5.998 | 521.844 |
| *D*rec, *D*sign | −5.688 | −5.965 | 522.000 |
| *D*rec, *D*size | −5.701 | −5.978 | 523.121 |
| *D*rec, *D*sign * *D*rec | −5.689 | −5.966 | 522.079 |
| *D*rec, *D*size * *D*rec | −5.703 | −5.979 | 523.250 |

NOTES: This table contains model selection statistics for the estimated model in Equations (1–5) and (7), under various specifications for the vector of explanatory variables, $z_t$. The monetary policy variable, $x_t$, is measured as a monetary policy shock from an identified VAR in which the monetary policy instrument is the federal funds rate. The sample is the third quarter of 1960 to the fourth quarter of 2002.

Thus, once asymmetry related to the business cycle phase is accounted for, it appears that asymmetry related to the size of the policy shock is no longer significant.

Finally, the third part of Table 1 investigates the extent to which asymmetry related to the direction and size of a policy action might be significant when considered *within* an NBER recession. In particular, we are interested in whether *D*sign and *D*size, while not unconditionally significant, might be significant conditional on being in a recession. To answer this question, we estimate two models in which $z_t = (D\text{rec}_{t-1}, D\text{rec}_{t-2}, D\text{size}_{t-1} * D\text{rec}_{t-1}, D\text{size}_{t-2} * D\text{rec}_{t-2})'$ and $z_t = (D\text{rec}_{t-1}, D\text{rec}_{t-2}, D\text{sign}_{t-1} * D\text{rec}_{t-1}, D\text{sign}_{t-2} * D\text{rec}_{t-2})'$. The results provide little evidence of any such conditional asymmetry. In particular, the likelihood ratio test of the null hypothesis of the model including only *D*rec is not rejected at any reasonable significance level. Also, the SIC and AIC prefer this simpler model in both cases.

In summary, the above model specification exercise suggests that our preferred model is one in which the response of the transitory component of output to lagged policy shocks varies between two regimes, with the probability that the $S_t = 0$ regime will continue, or switch to the $S_t = 1$ regime, dependent on whether the economy is in an NBER recession at the time the policy action is taken. In the next section, we discuss the estimation results for this preferred model in more detail.

### *2.4 Estimation Results for Preferred Model*

This section presents detailed estimation results for our preferred model, which is the model in Equations (1–5) and (7) with $z_t = D\text{rec}_{t-1}, D\text{rec}_{t-2}$. Table 2 presents the maximum likelihood estimates of the parameters. The parameters of the trend

TABLE 2

PARAMETER ESTIMATES FOR PREFERRED MODEL ($z_t = (D\text{rec}_{t-1}, D\text{rec}_{t-2})'$)(MONETARY POLICY MEASURE: FEDERAL FUNDS RATE BASED MONETARY POLICY SHOCK)

| Parameter | Estimate |
|---|---|
| $\sigma_v$ | 0.000 (0.000) |
| $\sigma_\varepsilon$ | 0.010 (0.001) |
| $\sigma_\omega$ | 0.001 (0.000) |
| $\phi_1$ | 1.452 (0.026) |
| $\phi_2$ | −0.501 (0.017) |
| $\gamma_{1,0}$ | −0.002 (0.000) |
| $\gamma_{2,0}$ | −0.002 (0.000) |
| $\gamma_{1,1}$ | 0.011 (0.002) |
| $\gamma_{2,1}$ | −0.028 (0.004) |
| $c_0$ | 6.098 (—) |
| $a_{01}$ | −5.886 (1.060) |
| $a_{02}$ | 0.189 (1.397) |
| $c_1$ | −1.623 (0.037) |
| Log likelihood | 521.844 |

NOTES: This table contains maximum likelihood estimates of the parameters of the model in Equations (1–5) and (7), when $z_t = (D\text{rec}_{t-1}, D\text{rec}_{t-2})'$ and the monetary policy variable, $x_t$, is measured as a monetary policy shock from an identified VAR in which the monetary policy instrument is the federal funds rate. The sample is the third quarter of 1960 to the fourth quarter of 2002. Standard errors are in parentheses. The estimate of $c_0$ implies that when $z_t = 0$, $P(S_t = 0 \mid S_{t-1} = 0)$;si1. This is a boundary value for the transition probability and creates difficulties in inverting the information matrix. Thus, when computing standard errors for the other parameters we constrained $c_0$ to its maximum likelihood estimate.

component, $y_t^P$, suggest that trend growth of industrial production is well characterized as being largely constant, with occasional shifts. Specifically, $\sigma_\omega$ is estimated to be non-zero, suggesting that the trend component is characterized by low frequency shocks, which permanently change the trend growth rate. However, $\sigma_v$ is estimated to be very close to zero, suggesting that once these low frequency shocks are accounted for, there are no other permanent shocks to industrial production.[14] The transitory component, $y_t^T$, shown in Figure 1, is large, with declines that correspond closely to NBER-dated recessions and expansions.[15] There is also significant

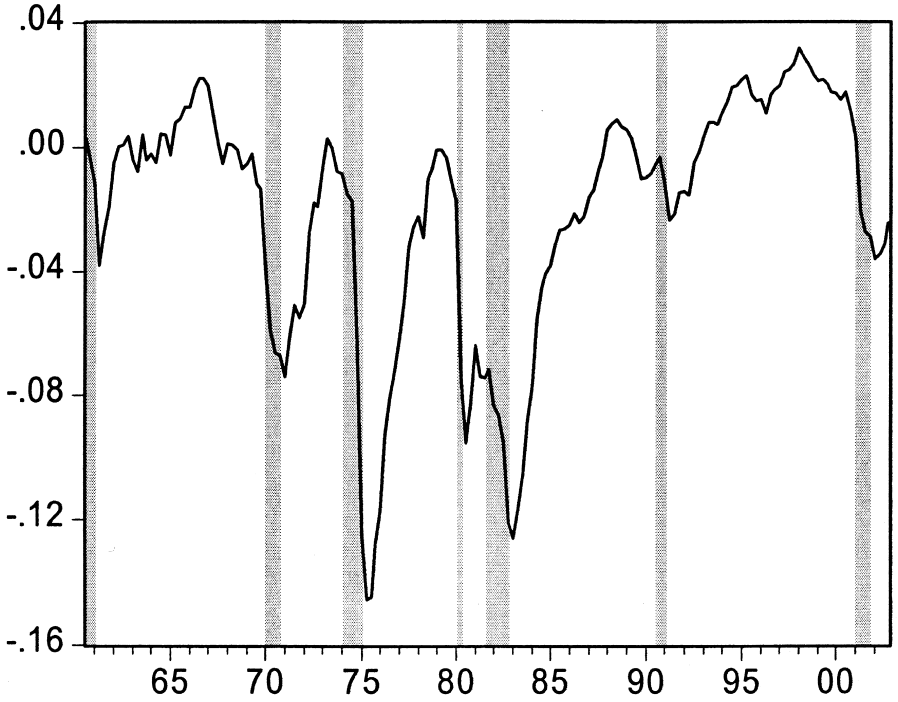


FIG. 1. Estimated transitory component, $y_t^T$ (monetary policy measure: federal funds rate based monetary policy shock) (Notes: This figure shows the estimated transitory component, $y_t^T$, from the model in Equations (1–5) and (7), when $z_t = (D\text{rec}_{t-1}, D\text{rec}_{t-2})'$ and the monetary policy variable, $x_t$, is measured as a monetary policy shock from an identified VAR in which the monetary policy instrument is the federal funds rate. The sample is the third quarter of 1960 to the fourth quarter of 2002. Shaded areas indicate NBER recession dates.)

14. These results relate to the debate on whether the trend in U.S. log output is best characterized as a broken time trend or as a stochastic trend. Of course, it is well known that distinguishing these two views of trend in the data is very difficult in practice (Campbell and Perron 1991). Indeed, Clark (1987) finds evidence of a large $\sigma_v$ for U.S. real GDP using a univariate unobserved-components model. In this paper, we are interested in the cyclical component of output and, given that these two views of trend yield similar estimated cycles, we will not be concerned with this debate further.

15. However, Morley, Nelson, and Zivot (1999) demonstrate that allowing for correlation between innovations to the stochastic trend and innovations to the transitory component can yield a transitory component that is both small and at odds with NBER-dated cycles.

negative skewness in the transitory component. That is, negative deviations from trend are deeper, shorter episodes than positive deviations from trend. This deepness of recessions, documented by Sichel (1993), is consistent with Friedman's (1964, 1993) "plucking" view of economic fluctuations.

We now turn to the regime-switching response coefficients, $\gamma_{1,0}$, $\gamma_{2,0}$, $\gamma_{1,1}$, and $\gamma_{2,1}$. The parameter estimates for these coefficients suggest that the indicator variable, $S_t$, divides policy shocks that have relatively small effects from those that have much larger effects. To see this, we compute state-dependent impulse response functions. Specifically, we set $x_{t-1}$ equal to its historical standard deviation (approximately 0.80) and simulate the path of $y^T_{t+j}$ using Equation (4) and the maximum likelihood point estimates of the parameters. Because $J = 2$ in Equation (4), the impulse response functions will depend only on the values of $S_t$ and $S_{t+1}$. We thus compute impulse response functions under the four possible realizations of these indicator variables: $S_t = S_{t+1} = 0$, $S_t = 1$ and $S_{t+1} = 0$, $S_t = 0$, and $S_{t+1} = 1$, $S_t = S_{t+1} = 1$. In computing the responses, we assume that $y^T_{t-1} = y^T_{t-2} = 0$, $\varepsilon_{t+j} = 0$, $\forall j$ and $x_{t-j} = 0, j \neq 1$. From Figure 2, which holds the impulse response functions, we can see that the response of $y^T_t$ to a monetary policy shock is much larger when either

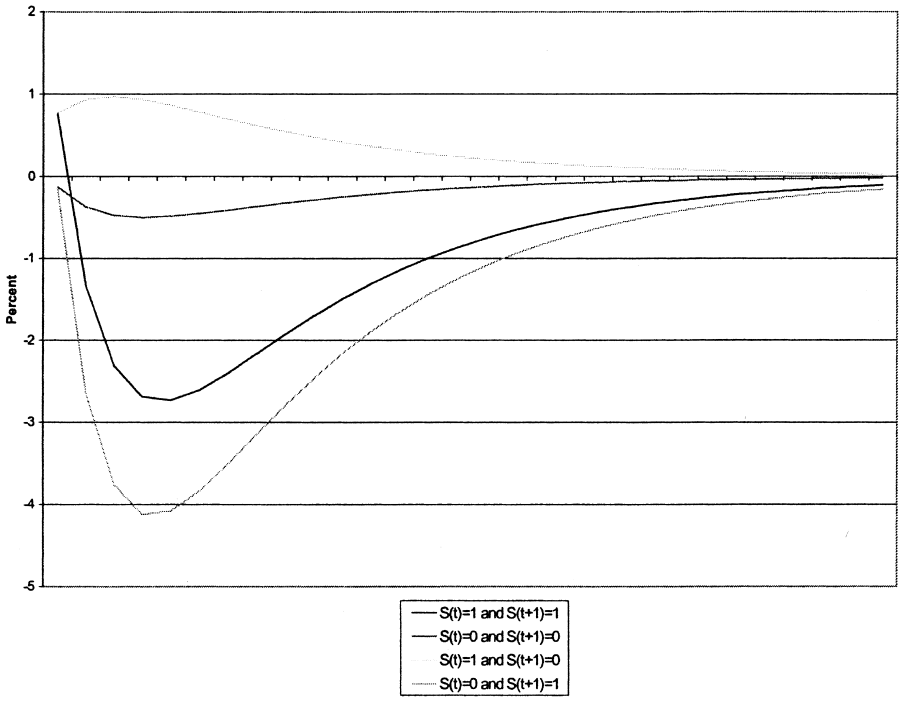


FIG. 2. Impulse response function of $y^T_t$ (monetary policy measure: federal funds rate based monetary policy shock) (Notes: This figure shows regime dependent impulse response functions of the transitory component, $y^T_t$, to a positive shock to the federal funds rate at time $t - 1$, computed as described in Section 2.4. The size of the shock is equal to the standard deviation of historical federal funds rate shocks, computed from an identified VAR.)

$S_t$ or $S_{t+1}$ are equal to one. For example, a one standard deviation realization of $x_t$ lowers industrial production by a maximum amount of $-0.5\%$ when $S_t = S_{t+1} = 0$. However, when $S_t = S_{t+1} = 1$ or $S_t = 0$, $S_{t+1} = 1$, the maximum response of industrial production is much larger, reaching –2.7% and –4.1%, respectively. Finally, when $S_t = 1$, $S_{t+1} = 0$, the maximum response is 1.0%.

Note that this final combination contains the counter-intuitive implication that a positive shock to the federal funds rate yields an increase in the cyclical component of industrial production. This result is an example of a case where models that assume a constant response of output to a policy shock mask interesting features of the data. For example, when we estimate the model in Equations (1–5) assuming that the response coefficients are constant, which can usefully be thought of as averaging the responses in Figure 2, the estimated response of industrial production to a positive federal funds rate shock is negative. This is consistent with the vast literature based on linear VAR models. However, the results in Figure 2 suggest that when $S_t = 1$, $S_{t+1} = 0$, the correlation between these policy shocks and future output is positive. Recall from Section 2.3 that the timing of the $S_t = 1$ regime is significantly correlated with the dates of NBER recessions. Given this, one explanation for the shape of the IRF when $S_t = 1$, $S_{t+1} = 0$ is that the Federal Reserve's information set in the standard VAR used here to extract policy shocks, particularly with regard to expected future output growth, is not well specified during recessions. In this case, the policy shocks obtained during recessions from this VAR would contain an endogenous component.

We now turn to the estimated coefficients determining the transition probabilities in Equation (7). To begin with, we consider the parameters determining $P(S_t = 0 \mid S_{t-1} = 0)$: $c_0$, $a_{01}$, and $a_{02}$. First, consider the case in which $D\text{rec}_{t-1} = D\text{rec}_{t-2} = 0$. Here, the estimate of $P(S_t = 0 \mid S_{t-1} = 0)$ is determined by $\hat{c}_0 = 6.1$ and is equal to $\exp(\hat{c}_0)/(1 + \exp(\hat{c}_0)) = 0.99$. Thus, if $S_{t-1} = 0$ and the economy has not been in a recession in the recent past, $S_t = 0$ with near certainty. However, the estimates $\hat{a}_{01}$ and $\hat{a}_{02}$ suggest that when the economy has been in a recession in the recent past, the probability that $S_t = 1$ rises drastically. From Table 2, $\hat{a}_{02}$ is small and statistically insignificant, while $\hat{a}_{01}$ is large, negative, and statistically significant. From Equation (7), this implies that if $D\text{rec}_{t-1} = 1$, that is the economy is in an NBER recession at time $t-1$, $P(S_t = 0 \mid S_{t-1} = 0)$ declines to $\exp(\hat{c}_0 + \hat{a}_{01})/(1 + \exp(\hat{c}_0 + \hat{a}_{01})) = 0.55$. Given that the $S_t = 1$ regime is one in which $x_{t-1}$ has large effects, this suggests that policy actions taken during NBER-dated recessions will be much more likely to have large output effects than those taken outside of NBER recessions. The parameter determining $P(S_t = 1 \mid S_{t-1} = 1)$, $\hat{c}_1$, is equal to –1.6, suggesting that $P(S_t = 1 \mid S_{t-1} = 1) = 0.17$. This implies that $S_t = 1$ only in short bursts.

Finally, we turn to the estimated timing of the regime switches, which can be viewed graphically using the estimated probability that $S_t = 1$, which we denote $P(S_t = 1 \mid t)$ .[16] There are several items of note in these estimates, which are shown

16. This estimate is constructed using data from $(y_1 \ldots y_t)$, and is often called a "filtered" probability.

in Figure 3. First, the model is identifying two clear regimes, as $P(S_t = 1 \mid t)$ is generally very close to zero or far from zero. Second, the occurrence of the $S_t = 1$ regime is both infrequent and brief. In particular, $S_t$ is generally equal to zero, suggesting that monetary policy shocks usually have small effects. When $S_t$ does switch on, it tends to remain on for only a very brief period of time, usually no more than two or three quarters. Finally, as should not be surprising, given the significance of *D*rec in explaining time variation in the transition probabilities, there is a strong correspondence between the $S_t = 1$ regime and NBER recession dates, which are the shaded areas on the graph. Every period that $P(S_t = 1 \mid t)$ is high is in close proximity to one of the NBER recessions in the sample. Also, around every NBER recession in the sample, there is at least one episode in which $P(S_t = 1 \mid t)$ spikes up.

The striking correspondence between the $S_t = 1$ regime and the business cycle provides a useful robustness check on a growing literature finding that U.S. monetary policy actions tend to predict output much more significantly during recessions than during expansions (see, for example, Garcia and Schaller 2002). Figure 3 suggests

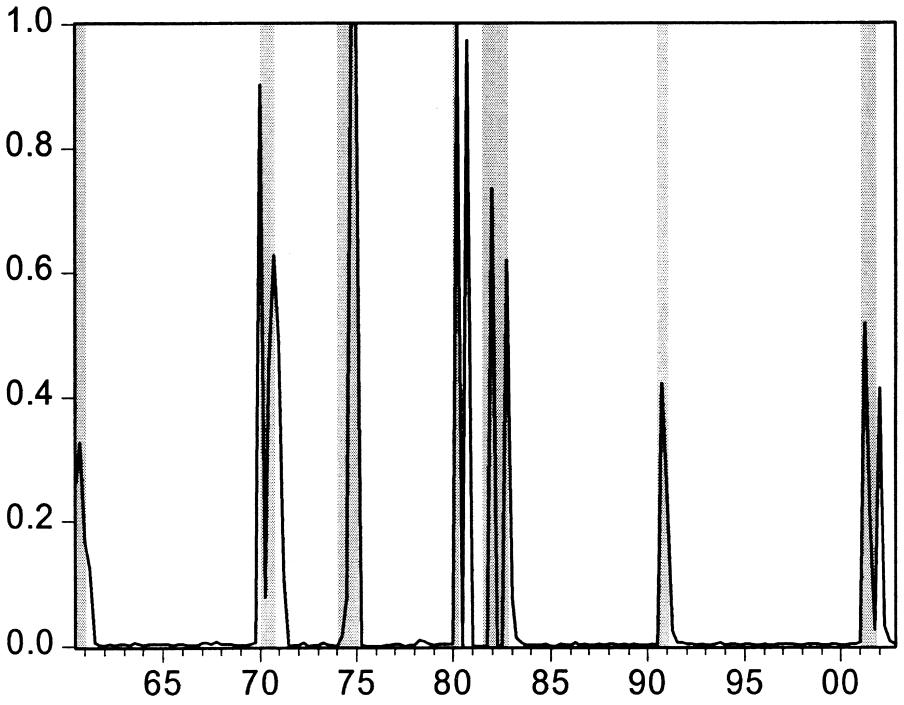


FIG. 3. Filtered probability, $P(S_t = 1 \mid t)$ (monetary policy measure: federal funds rate based monetary policy shock) (Notes: This figure shows the filtered probability that $S_t = 1$, $P(S_t = 1 \mid t)$, from the model in Equations (1–5) and (7), when $z_t = (D\text{rec}_{t-1}, D\text{rec}_{t-2})'$ and the monetary policy variable, $x_t$, is measured as a monetary policy shock from an identified VAR in which the monetary policy instrument is the federal funds rate. The sample is the third quarter of 1960 to the fourth quarter of 2002. Shaded areas indicate NBER recession dates.)

that this result is fairly robust across the historical record of business cycles. In other words, this result is not being driven by only a small subset of recessions.

*2.5 Robustness Checks: Alternative Measures of Monetary Policy*

In this section, we investigate the robustness of the results obtained above to two alternative measures of $x_t$ . The first is an endogenous measure of monetary policy, namely the change in the ex-post real interest rate, measured as the quarterly average federal funds rate less the four quarter percentage change in the GDP price deflator. The second is a monetary policy shock in which the Federal Reserve's monetary policy instrument is the M1 money supply. This is obtained from a recursively identified four variable VAR in which M1 is ordered after the log of real GDP and the log of the GDP price deflator and before the federal funds rate. Due to data limitations on the M1 variable, the model is estimated beginning in the first quarter of 1965 for this measure of $x_t$.[17]

The response coefficients linking the cyclical component of industrial production to the alternative measures of $x_t$ appear well characterized by regime switching. The Hansen (1992) test of the null hypothesis of constant regime coefficients against the alternative of the FTP model in Equations (1–6) yields a $p$-value of 0.03 for the model in which $x_t$ is the change in the ex-post real federal funds rate, and 0.04 for the model in which $x_t$ is the money-based monetary policy shock. In Table 3,

TABLE 3

MODEL SELECTION FOR TIME-VARYING TRANSITION PROBABILITY SPECIFICATION

| Elements in $z_t$ | SIC | AIC | Likelihood |
|---|---|---|---|
| *$x_t$ = Change in Ex-Post Real Federal Funds Rate* | | | |
| None | −5.719 | −5.923 | 511.489 |
| *D*sign | −5.660 | −5.901 | 511.640 |
| *D*size | −5.679 | −5.920 | 513.211 |
| *D*rec | −5.739 | −5.980 | 518.294 |
| *D*rec, *D*sign | −5.681 | −5.958 | 518.478 |
| *D*rec, *D*size | −5.685 | −5.962 | 518.822 |
| *D*rec, *D*sign ∗ *D*rec | −5.689 | −5.967 | 519.205 |
| *D*rec, *D*size ∗ *D*rec | −5.686 | −5.964 | 518.934 |
| *$x_t$ = Money-Based Monetary Policy Shock* | | | |
| None (FTP model) | −5.640 | −5.859 | 456.289 |
| *D*sign | −5.583 | −5.842 | 456.980 |
| *D*size | −5.579 | −5.837 | 456.628 |
| *D*rec | −5.627 | −5.886 | 460.341 |
| *D*rec, *D*sign | −5.570 | −5.868 | 460.952 |
| *D*rec, *D*size | −5.570 | −5.866 | 460.828 |
| *D*rec, *D*sign ∗ *D*rec | −5.638 | −5.936 | 466.167 |
| *D*rec, *D*size ∗ *D*rec | −5.561 | −5.860 | 460.345 |

NOTES: This table contains model selection statistics for the estimated model in Equations (1–5) and (7), under various specifications for the vector of explanatory variables, $z_t$. The monetary policy variable, $x_t$ , is measured either as the change in the ex-post real federal funds rate (top panel) or a monetary policy shock from an identified VAR in which the monetary policy instrument is the M1 money supply (bottom panel). The sample is the fourth quarter of 1960 to the fourth quarter of 2002 for the top panel and the first quarter of 1965 to the fourth quarter of 2002 for the bottom panel.

17. The first quarter for which M1 data were available was the first quarter of 1959. After forming the monetary policy shock and discarding an additional five years of data to mitigate the effects of initial conditions on the maximum likelihood estimation (discussed in Section 2.1), the model is estimated beginning in the first quarter of 1965.

we show the results of a model comparison exercise to determine which asymmetry dummy variables are able to explain the regime switching. The results for the model in which $x_t$ is measured as the change in the ex-post real federal funds rate are very similar to those for the federal funds rate based monetary policy shock in Table 1. In particular, the SIC and AIC choose the model in which $z_t$ includes the NBER recession dummy variable only. When $x_t$ is measured using the money-based monetary policy shock, the results are a bit different. The preferred model based on the AIC is one in which $z_t$ includes the NBER dummy variable in addition to the sign dummy variable interacted with the NBER dummy variable. The SIC, on the other hand, slightly prefers the FTP model. Going forward, we use as our preferred model the one in which $z_t = D\text{rec}_{t-1}, D\text{rec}_{t-2}$ when $x_t$ is the change in the ex-post real federal funds rate and $z_t = D\text{rec}_{t-1}, D\text{rec}_{t-2}, D\text{sign}_{t-1} * D\text{rec}_{t-1}, D\text{sign}_{t-2} * D\text{rec}_{t-2}$ when $x_t$ is the M1-based monetary policy shock.

We now consider the estimation results from these preferred models. Figures 4 and 5 show the regime-dependent impulse response functions, computed as discussed in Section 2.3, to a one standard deviation tightening of $x_t$. For the change in the ex-post real federal funds rate, this is equal to a 95 basis point increase in the real federal funds rate, while for the M1-based monetary policy shock this is equal to

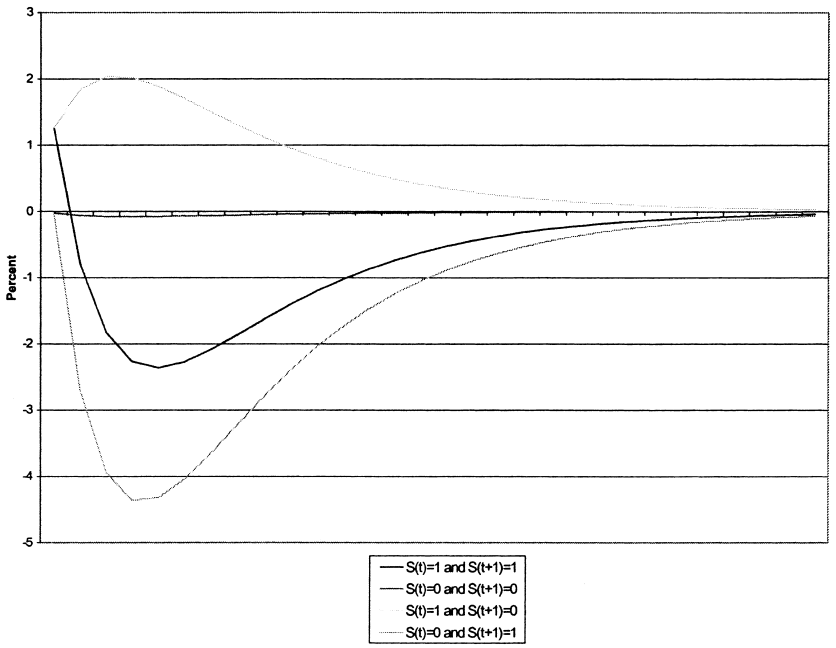


FIG. 4. Impulse response function of $y_t^T$ (monetary policy measure: change in ex-post real federal funds rate) (Notes: This figure shows regime dependent impulse response functions of the transitory component, $y_t^T$, to a positive change in the ex-post real federal funds rate at time $t - 1$, computed as described in Section 2.4. The size of the change is equal to the standard deviation of historical real federal funds rate changes.)

a 0.7% decrease in the M1 money supply. In both cases, the impulse response is much larger when either $S_t$ or $S_{t+1}$ are equal to one, consistent with the results for the federal funds rate based monetary policy shock shown in Figure 2. When $x_t$ is measured using the change in the ex-post real federal funds rate, the estimated coefficients $\hat{c}_0$, $\hat{a}_{01}$, and $\hat{a}_{02}$ (not reported), are again suggestive that the occurrence of an NBER recession in the recent past significantly increases the probability that $S_t = 1$. When $x_t$ is measured using the money-based monetary policy shock, the variables $D\text{sign}_{t-1} * D\text{rec}_{t-1}$, $D\text{sign}_{t-2} * D\text{rec}_{t-2}$ also enter the $z_t$ vector in addition to $D\text{rec}_{t-1}$ and $D\text{rec}_{t-2}$. The estimated coefficients on the $z_t$ vector are such that the probability that $S_t = 1$ increases considerably when the policy shock observed during a recession is a stimulus rather than a tightening. This is consistent with the results of Weise (1999), who, using a three variable threshold vector autoregression consisting of the consumer price index, industrial production, and the M1 money supply, finds that policy stimulus taken when output is declining has a larger effect than a policy contraction.

Finally, we consider the timing of the regime switches for the alternative policy measures. Figures 6 and 7, which show the filtered probabilities $P(S_t = 1 \mid t)$, demonstrate that the timing is similar to that obtained when $x_t$ is measured using the federal

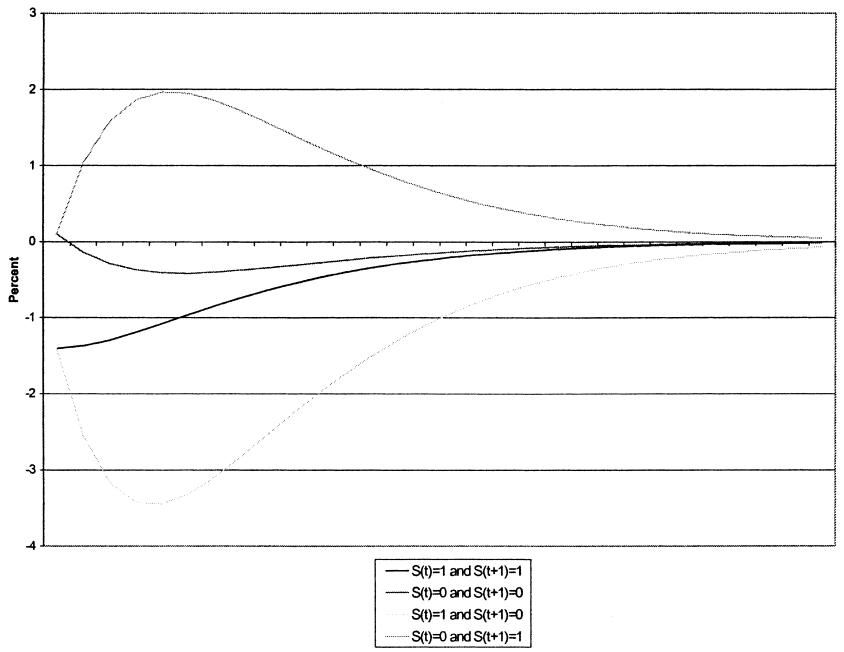


FIG. 5. Impulse response function of $y_t^T$ (monetary policy measure: M1 based monetary policy shock) (Notes: This figure shows regime dependent impulse response functions of the transitory component, $y_t^T$, to a negative shock to the M1 money supply at time $t - 1$, computed as described in Section 2.4. The size of the shock is equal to the standard deviation of historical M1 shocks, computed from an identified VAR.)

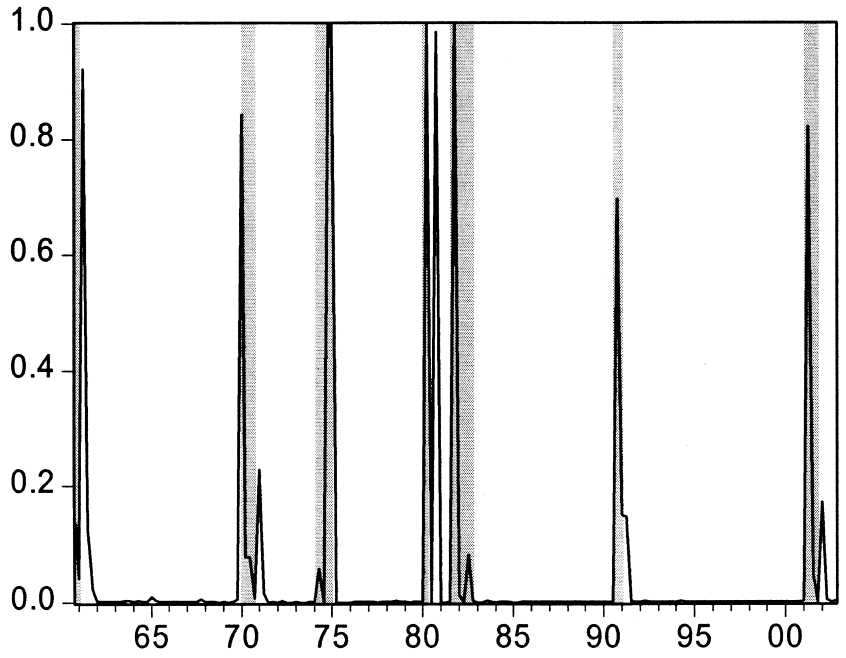


Fig. 6. Filtered probability, $P(S_t = 1 \mid t)$ (monetary policy measure: change in ex-post real federal funds rate) (Notes: This figure shows the filtered probability that $S_t = 1$, $P(S_t = 1 \mid t)$, from the model in Equations (1–5) and (7), when $z_t = (D\text{rec}_{t-1}, D\text{rec}_{t-2})'$ and the monetary policy variable, $x_t$, is measured as the change in the ex-post real federal funds rate. The sample is the fourth quarter of 1960 to the fourth quarter of 2002. Shaded areas indicate NBER recession dates.

funds rate based monetary policy shock. Specifically, as was the case in Figure 3, $P(S_t = 1 \mid t)$ is high infrequently and briefly. Also, the periods in which $P(S_t = 1 \mid t)$ spike up are highly correlated with NBER recessions.

## 3. CONCLUSIONS

A growing literature has investigated asymmetries in the effects of U.S. monetary policy on the real economy. In this paper, we have used a Markov regime-switching model to investigate time variation in the response of the cyclical component of output to monetary policy actions. A time-varying transition probability specification allows us to explain the regime shifts using state variables that are linked to three particular manifestations of asymmetry: asymmetry related to the direction of the monetary policy action, asymmetry related to the existing business cycle phase, and asymmetry related to the size of the policy action.

The results suggest substantial, statistically significant, time variation in the coefficients describing the response of output to a monetary policy action, and that this time variation corresponds to "high response" and "low response" regimes. The time-varying transition probability model yields strong evidence that the time variation

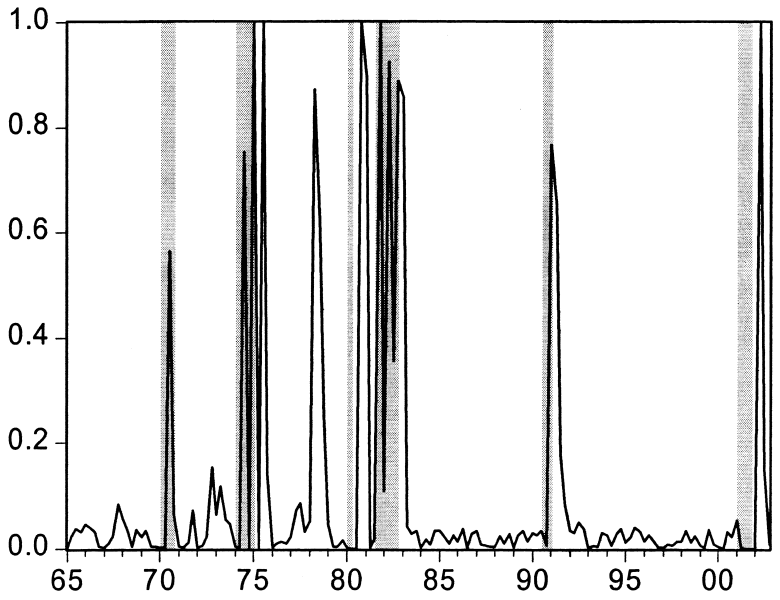


FIG. 7. Filtered probability, $P(S_t = 1 \mid t)$ (monetary policy measure: M1 based monetary policy shock) (Notes: This figure shows the filtered probability that $S_t = 1$, $P(S_t = 1 \mid t)$ from the model in Equations (1–5) and (7), when $z_t = (D\text{rec}_{t-1}, D\text{rec}_{t-2}, D\text{rec}_{t-1} * D\text{sign}_{t-1}, D\text{rec}_{t-2} * D\text{sign}_{t-2})'$ and the monetary policy variable, $x_t$, is measured as a monetary policy shock from an identified VAR in which the monetary policy instrument is the M1 money supply. The sample is the first quarter of 1965 to the fourth quarter of 2002. Shaded areas indicate NBER recession dates.)

in the response of output can be explained by a dummy variable indicating whether the economy is in a recession at the time the policy action is taken. In particular, policy actions taken during recessions seem to have larger effects than those taken during expansions. This result appears to be robust across the historical record of business cycles, that is, it does not appear to be driven by only a small subset of recessions. We find much less evidence of any asymmetry related to the direction or size of the policy action.

Our finding that output responds more to policy actions taken during recessions than those taken during expansions is consistent with a growing literature, including Garcia and Schaller (2002) for the United States, Peersman and Smets (2001) for the euro area, and Kaufmann (2002) for Austrian data. However, the results are not supportive of the literature, for example, Cover (1992) and DeLong and Summers (1988), that finds that output responds more to a policy contraction than to a policy stimulus. For policy variables based on interest rates, we find no evidence of any asymmetry related to the direction of the policy actions. When the policy measure is based on the M1 money supply, we find that a policy stimulus has larger output effects than a policy contraction, but only when taken during recessions. This is consistent with the results of Weise (1999) who, using a three variable threshold

vector autoregressive model consisting of the consumer price index, industrial production, and the M1 money supply, finds that policy stimulus taken when output is declining has a larger effect than a policy contraction. Finally, our results are in general not supportive of a literature, see for example Ravn and Sola (1999), that documents asymmetry related to the size of the policy action. While we do find some evidence of this sort of asymmetry when it is the only type of asymmetry considered, it does not retain its significance once it is considered jointly with asymmetry related to the business cycle.

The results presented here leave open the important question of why policy actions would have larger effects in recessions. As was noted in the introduction to this study, one theory posits that the "balance-sheet" channel of monetary policy augments the traditional interest rate channel to a greater extent during recessions than during recession. Peersman and Smets (2002) have provided some evidence in favor of this theory using industry level data for seven euro area countries. They find that those industries for which business cycle asymmetry in the effects of monetary policy is greatest tend to be those with firm size and financial structure characteristics that make them most susceptible to a "balance-sheet" channel.

## LITERATURE CITED

Ball, Laurence, and N. Gregory Mankiw (1994). "Asymmetric Price-Adjustment and Economic Fluctuations." *The Economic Journal* 423, 247–261.

Belongia, Michael (1996). "Measurement Matters: Recent Results from Monetary Economics Reexamined." *Journal of Political Economy* 104, 1065–1083.

Bernanke, Ben S., and Mark Gertler (1995). "Inside the Black Box: The Credit Channel of Monetary Policy Transmission." *Journal of Economic Perspectives* 9, 27–48.

Campbell, John Y., and Pierre Perron (1991). "Pitfalls and Opportunities: What Macroeconomists Should Know About Unit Roots." *NBER Macroeconomics Annual* 6, 141–201.

Christiano, Lawrence J., Martin Eichenbaum, and Charles L. Evans (1997). "Monetary Policy Shocks: What Have We Learned and to What End?" In *Handbook of Macroeconomics*, edited by John B. Taylor and Michael Woodford. Amsterdam: Elsevier Science.

Clark, Peter K. (1987). "The Cyclical Component of U.S. Economic Activity." *Quarterly Journal of Economics* 102, 797–814.

Cover, James P. (1992). "Asymmetric Effects of Positive and Negative Money-Supply Shocks." *Quarterly Journal of Economics* 107, 1261–1282.

DeLong, J. Bradford, and Lawrence H. Summers (1988). "How does Macroeconomic Policy Affect Output?" In *Brookings Papers on Economic Activity*, edited by William C. Brainard and George L. Perry, pp. 433–494. Washington, DC: The Brookings Institution.

Diebold, Francis X., Joon-Haeng Lee, and Gretchen C. Weinbach (1994). "Regime Switching with Time-Varying Transition Probabilities." In *Non-Stationary Time Series Analysis and Cointegration*, edited by Colin Hargreaves. Oxford, U.K.: Oxford University Press.

Evans, Paul (1986). "Does the Potency of Monetary Policy Vary with Capacity Utilization?" *Carnegie Rochester Conference Series on Public Policy* 24, 303–332.

Filardo, Andrew J. (1994). "Business-Cycle Phases and Their Transitional Dynamics." *Journal of Business and Economic Statistics* 12, 299–308.

Friedman, Milton (1964). "Monetary Studies of the National Bureau, the National Bureau enters its 45th Year." 44th Annual Report, 7-25, NBER, New York. Reprinted in Friedman, Milton (1969). " *The Optimum Quantity of Money and other Essays*," Aldine, Chicago.

Friedman, Milton (1993). "The 'Plucking Model' of Business Fluctuations Revisited." *Economic Inquiry* 31, 171–177.

Garcia, Rene (1998). "Asymptotic Null Distribution of the Likelihood Ratio Test in Markov Switching Models." *International Economic Review* 39, 763–788.

Garcia, Rene, and Huntley Schaller (2002). "Are the Effects of Interest Rate Changes Asymmetric?" *Economic Inquiry* 40, 102–119.

Gerlach, Stefan, and Frank Smets (1999). "Output Gaps and Monetary Policy in the EMU Area." *European Economic Review* 43, 801–812.

Hamilton, James D. (1989). "A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle." *Econometrica* 57, 357–384.

Hansen, Bruce E. (1992). "The Likelihood Ratio Test Under Nonstandard Conditions: Testing the Markov Switching Model of GNP." *Journal of Applied Econometrics* 7, S61–S82.

Karras, Georgios (1996). "Why are the Effects of Money-Supply Shocks Asymmetric? Convex Aggregate Supply or 'Pushing on a String'?" *Journal of Macroeconomics* 18, 605–619.

Karras, Georgios, and Houston Stokes (1999). "Why are the Effects of Money-Supply Shocks Asymmetric? Evidence from Prices, Consumption, and Investment." *Journal of Macroeconomics* 21, 713–727.

Kaufmann, Sylvia (2002). "Is there an Asymmetric Effect of Monetary Policy over Time? A Bayesian Analysis using Austrian Data." *Empirical Economics* 27, 277–297.

Kim, Chang-Jin (1994). "Dynamic Linear Models with Markov-Switching." *Journal of Econometrics* 60, 1–22.

Kim, Chang-Jin, and Charles R. Nelson (1999). *State Space Models with Regime Switching*. Cambridge, MA: The MIT Press.

Morgan, Donald (1993). "Asymmetric Effects of Monetary Policy." *Federal Reserve Bank of Kansas City Economic Review* 78, 21–33.

Morley, James, Charles R. Nelson, and Eric Zivot (1999). "Why are Unobserved Component and Beveridge Nelson Trend-Cycle Decompositions of GDP so Different?" *Review of Economics and Statistics* 85, 235–243.

Parker, Randall E., and Philip Rothman (2004). "An Examination of the Asymmetric Effects of Money Supply Shocks in the Pre-World War I and Interwar Periods." *Economic Inquiry* 42, 88–100.

Peersman, Gert, and Frank Smets (2002a). "Are the Effects of Monetary Policy in the Euro Area Greater in Recessions than in Booms?" In *Monetary Transmission in Diverse Economies*, edited by Lavan Mahadeva and Peter Sinclair, Cambridge, UK: Cambridge University Press.

Peersman, Gert, and Frank Smets (2002b). "The Industry Effects of Monetary Policy in the Euro Area." European Central Bank Working Paper No. 165.

Ravn, Morten, and Martin Sola (1999). "Asymmetric Effects of Monetary Policy in the US: Positive versus Negative or Big versus Small?" *Federal Bank of St. Louis Review*, forthcoming.

Rhee, Wooheon, and Robert W. Rich (1995). Inflation and the Asymmetric Effects of Money on Output Fluctuations." *Journal of Macroeconomics* 17, 683–702.

Senda, Takashi (2001). "Asymmetric Effects of Money Supply Shocks and Trend Inflation." *Journal of Money, Credit, and Banking* 33, 65–89.

Sichel, Daniel E. (1993). "Business Cycle Asymmetry: A Deeper Look." *Economic Inquiry* 31, 224–236.

Thoma, Mark A. (1994). "Subsample Instability and Asymmetries in Money-Income Causality." *Journal of Econometrics* 64, 279–306.

Thoma, Mark A., and Jo Anna Gray (1998). "Financial Market Variables do not Predict Real Activity." *Economic Inquiry* 36, 522–539.

Watson, Mark W. (1986). "Univariate Detrending Methods with Stochastic Trends." *Journal of Monetary Economics* 18, 49–75.

Weise, Charles (1999). "The Asymmetric Effects of Monetary Policy: A Nonlinear Vector Autoregression Approach." *Journal of Money, Credit, and Banking* 31, 85–108.