ASYMMETRIC EFFECTS OF MONETARY POLICY IN BRAZIL

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Abstract

In this paper, we check whether the effects of monetary policy actions on output in Brazil are asymmetric. Therefore, we estimate Markov-switching models that allow positive and negative shocks to affect the growth rate of output in an asymmetric fashion in expansion and recession states. Results show that: i) when monetary policy actions are measured by means of orthogonalized innovations for the Selic rate in a VAR model, the real effects of negative monetary shocks are larger than those of positive shocks in an expansion and the real effects of negative shocks are greater in an expansion than in a recession; ii) when the variation in the Selic rate is used to measure monetary policy, we also have asymmetries between the real effects of positive and negative variations in the Selic rate during a recession, and between the real effects of negative variations of the Selic rate between the states of the business cycle.

Keywords: Monetary policy; Asymmetries; Positive and negative shocks; Business cycle; Markov-switching models.

JEL Classification: E52, E32, C32.

1 Introduction

Asymmetric real effects of monetary policy may be observed in several theoretical models, and they are classified into three types. The first type is related to the direction of the monetary policy action, where a negative monetary shock affects the output more strongly than a positive shock.[1] This type of asymmetry is found in models that yield a convex aggregate supply curve due to a stronger nominal rigidity to reduce the price and/or wage levels.[2] An example of this is the asymmetric adjustment cost

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[1] Positive monetary shocks can be defined as unanticipated increases in money supply or unanticipated decreases in the interest rate of the monetary policy. Conversely, negative monetary shocks are unanticipated decreases in money supply or unanticipated increases in the interest rate of the monetary policy.

[2] The Keynesian model with downward nominal wage rigidity and the asymmetric adjustment cost models analyzed by Caballero and Engel (1993) and Tsiddon (1993) are included in this class of models.

model developed by Ball and Mankiw (1994). In their model, asymmetric price adjustment stems from the assumption that inflation has a positive trend. In the absence of shocks, inflation pushes the desired nominal price above the current nominal price chosen by the firm. The effect of a positive monetary shock is a sharper increase in the difference between the desired and current nominal prices, whereas a negative monetary shock brings the desired nominal price towards the current nominal price. Therefore, a positive monetary shock leads to a higher frequency of price adjustments, while a negative monetary shock produces a major effect on the output level of firms.

The second type of asymmetry is concerned with the size of the monetary policy action, where small monetary shocks affect the real side of the economy more strongly than larger shocks. The menu cost model proposed by Ball and Romer (1989, 1990) is a perfect epitome of this type of asymmetry. In this model, when a small monetary shock occurs, there is a change in output, whereas the price level remains unchanged. This is because, after the shock, the utility of each producer in maintaining his price at a fixed level is larger than the utility in adjusting prices. For large monetary shocks, the utility of each producer in adjusting the price is greater than the utility in maintaining prices fixed. This causes firms to adjust their prices, whereas output remains unchanged.

Thirdly, monetary policy may affect output in a different manner depending on whether the economy is in a recession or in an expansion. This may occur due to frictions in credit market generated by asymmetric information between lenders and borrowers.[3] As a result of these frictions, monetary policy affects not only the interest rate, but also the external finance premium, strengthening the impact of monetary shocks on borrowing costs, investment demand, and real output. Since the amount of collaterals is smaller in a recession, the financial stance of economic agents is weaker and the credit supply by commercial banks is more strict than during an expansion, the external finance premium is higher in this phase of the business cycle. As a result, monetary policy shocks may have greater effects on output during a recession.

Several works have provided empirical evidence of different types of asymmetry in the effects of monetary policy on output. A pioneering work was that by Cover (1992), who checked whether negative monetary shocks had larger real effects than positive shocks on the U.S. economy. According to Barro (1977) and Mishkin (1982), Cover uses a two-step procedure. The first step consists in estimating a money supply

[3] Gertler and Hubbard (1988), Bernanke and Gertler (1989), Gertler et al (1990), Gertler and Gilchrist (1994) and Bernanke and Gertler (1995) provide an in-depth analysis of credit market imperfections.

process and employing the residuals of this equation to measure monetary shocks. Afterwards, monetary shocks are included in the output equation, where the specification allows for different effects of positive and negative values of these shocks. The results indicate that negative monetary shocks affect output more strongly than positive ones. De Long and Summers (1988), Rhee and Rich (1995) and Karras and Stokes (1999) use the approach suggested by Cover (1992) and obtain similar results to those obtained by Cover.

Ravn and Sola (2004) investigate whether the asymmetric effects of monetary shocks in the USA are related to the direction and size of these shocks. The difference between small and large shocks is based on the variances of innovations of a Markov-switching model for the benchmark interest rate (Federal Funds rate). The results indicate non-neutrality of small and negative monetary shocks only.

Garcia and Schaller (2002), Dolado and Maria-Dolores (2001, 2006), Peersman and Smets (2001) and Kaufmann (2002) assess asymmetry in the business cycle phase by estimating Markov-switching models that allow monetary policy effects on output to be different during an expansion and a recession. These authors found plenty evidence that monetary policy actions have stronger effects on output in a recession in the U.S., Austrian, German, Spanish, Italian, French, and Belgian economies.

Weise (1999) and Lo and Piger (2005) jointly assess the three types of asymmetry in the real effects of monetary policy in the USA. While Weise (1999) investigates the types of asymmetry by analyzing the impulse-response functions generated by the estimation of an LSTVAR (*Logistic Smooth Transition Vector Autoregression*) model, Lo and Piger (2005) estimate an unobserved components model and check whether the time variation in the response of the cyclical component of output to monetary shocks is related to different types of asymmetry. Both works provide strong evidence of asymmetry in the business cycle phase.

The major aim of the present work is to investigate whether the effects of monetary policy on output are asymmetric in Brazil. Specifically, we seek to answer the following questions: i) in a given state of the business cycle, are the real effects of a contractionary monetary policy different from the effects of an expansionary policy? ii) do the real effects of a countercyclical monetary policy depend on the state of the business cycle in place at the time at which the policy was implemented? iii) are the real effects of a contractionary (or expansionary) policy different between the phases of the business cycle?

In order to achieve our goal, we first measured monetary policy shocks and then we built positive and negative monetary shock series. Thereafter, we assessed the different types of asymmetry by extending the Markov-switching model developed by Hamilton (1989). In particular, we considered a specification of the Markov-switching model that allows positive and negative shocks to asymmetrically affect the rate of GDP growth during expansions and recessions. This model has the following advantages: i) the selection of states is jointly determined with the estimation of the model parameters; ii) a greater relative weight is placed upon the observations that correspond more clearly to this state when estimating the coefficients of a given state; iii) the model allows jointly assessing the different types of asymmetric effects of monetary policy on output. To check whether the real effects of monetary shocks are asymmetric, we performed a set of Wald tests and imposed different restrictions on the estimated parameters. Finally, we built and analyzed the output paths in response to a positive or negative monetary shock for each state of the business cycle.

An important aspect of the analysis of monetary policy effects is the choice of the monetary policy instrument and consequently of the measure of policy actions. We chose the Selic rate as policy instrument because we believe this variable properly shows the main monetary policy actions taken by the Central Bank in the post-Real Plan period. After that, we measured monetary policy actions based on the innovations of a vector autoregressive model in which the Selic rate is included as one of the system variables. To check the robustness of results, we regarded the variation in the Selic rate as a monetary policy measure.

Two were our contributions to the existing empirical literature. Unlike most works in this line of research, we assessed the asymmetry related to the direction of the shock and also the asymmetry during expansions and recessions. Several macroeconomists tend to conflate these two types of asymmetry, thinking that monetary policy actions are typically countercyclical, in such a way that an analysis of the asymmetry between an expansionary and contractionary policy allows inferring on the asymmetry between expansions and recessions. However, some works demonstrate that this might not apply to the Brazilian economy. Kaminsky et al. (2004) provide a body of evidence suggesting that monetary policy is procyclical in emerging economies (e.g.: Brazil). Minella et al. (2002, 2003) estimate several reaction functions for the Central Bank of Brazil and show that, in some specifications, the coefficient that measures the response of the Selic rate to output variations is negative and statistically significant.

This suggests that the interest rate may have increased while the output was decreasing or may have decreased while the output was increasing.

Our second contribution consists in providing evidence of asymmetric real effects of monetary policy in Brazil for the period following the Real Plan. Many authors, such as Moreira et al. (1998), Rabanal and Schwartz (2001), Minella (2003), Cysne (2005), Céspedes et al. (2005) and Fernandes and Toro (2005) have provided empirical evidence that monetary policy actions affect the real side of the Brazilian economy. Nevertheless, none of these studies considered that such effects were likely to be asymmetric in terms of economic conditions and/or nature of the policy action.

The empirical evidence obtained points to three results. First, when monetary policy actions are measured by orthogonalized innovations for the Selic rate in a VAR model, the real effects of negative monetary shocks are larger than the positive shocks in an expansion and the real effects of negative shocks are larger in an expansion than in a recession. Secondly, when the variation in the Selic rate is used to measure monetary policy, we also have asymmetries between the real effects of positive and negative variations in the Selic rate during a recession, and between the real effects of positive variations in the Selic rate between the states of the business cycle. Finally, we did not find any evidence of asymmetries in countercyclical monetary policy actions implemented in different business cycle phases.

Besides the introduction, this paper is organized into four sections. Section 2 outlines the empirical model and the statistical tests we are going to use to assess the different types of asymmetry in the real effects of the monetary policy. Section 3 describes the time series used. Section 4 presents and analyzes the results. And finally, section 5 concludes and gives suggestions for future research studies.

2 Empirical methodology

2.1 Econometric model

An empirical strategy that can be used to check the nonlinearity and asymmetry of time series data is to estimate the Markov-switching (MS) model developed by Hamilton (1989). In this model, the regime shifts in the behavior of a time series depend on an unobservable random variable (denoted by S_t) that characterizes the regime or state the process was in at time t. This variable supposedly assumes value 0 or 1 and follows an *r-th* order Markov-switching process. Hamilton (1989) adjusts a univariate specification of this model for the real output growth, allowing the mean growth rate to depend on regime S_t, whose behavior is determined by a first-order Markov chain and two states (economic recession and expansion).

To check whether the real effects of monetary policy are asymmetric, we extend Hamilton's (1989) model, which allows positive and negative monetary shocks to asymmetrically affect the output growth rate between expansions and recessions. A more general specification of the Markov-switching model is given by:

$$\begin{aligned}
&\Delta y_t - \mu_{s_t} = \phi_1(\Delta y_{t-1} - \mu_{s_{t-1}}) + \ldots + \phi_p(\Delta y_{t-p} - \mu_{s_{t-p}}) + \\
&\qquad\qquad \gamma^-_{S_{t-1},1} u^-_{t-1} + \ldots + \gamma^-_{S_{t-p},p} u^-_{t-p} + \gamma^+_{S_{t-1},1} u^+_{t-1} + \ldots + \gamma^+_{S_{t-p},p} u^+_{t-p} + \varepsilon_t \\
&\varepsilon_t \sim i.i.d.\,N(0,\sigma^2) \\
&\mu_{s_t} = \mu_0(1-s_t) + \mu_1 s_t \\
&\gamma^-_{S_{t-1},1} = \gamma^-_{0,1}(1-s_{t-1}) + \gamma^-_{1,1} s_{t-1} \\
&\gamma^+_{S_{t-i},1} = \gamma^+_{0,1}(1-s_{t-1}) + \gamma^+_{1,1} s_{t-1}
\end{aligned} \qquad (1)$$

where Δy_t is the growth rate of the real Brazilian GDP, μ_{s_t} is the mean state-dependent growth rate of output, u^-_{t-i} is a negative monetary shock, $\gamma^-_{S_{t-i},i}$ is the state-dependent coefficient measuring the response of Δy_t to a negative monetary shock, u^+_i is a positive monetary shock and $\gamma^+_{S_{t-i},i}$ is the state-dependent coefficient measuring the response of Δy_t to a positive monetary shock. The state variable S_t supposedly assumes value 0 when the economy is in a recession and value 1 when it is in an expansion. Thus, the model parameters in a recession are μ_0, $\gamma^-_{0,i}$ and $\gamma^+_{0,i}$, whereas, in an expansion, these parameters are given by μ_1, $\gamma^-_{1,i}$ and $\gamma^+_{1,i}$. The generating process of regimes S_t involves a first-order Markov-switching process and two states, whose supposedly ergodic and irreducible transition probability matrix is given by:

$$P = \begin{bmatrix} p & 1-q \\ 1-p & q \end{bmatrix} \qquad (2)$$

where

$$p = \Pr[S_t = 0 \mid S_{t-1} = 0],$$

$$1-p = \Pr[S_t = 1 \mid S_{t-1} = 0],$$

$$q = \Pr[S_t = 1 \mid S_{t-1} = 1], \qquad (3)$$
$$1 - q = \Pr[S_t = 0 \mid S_{t-1} = 1].$$

We assume transition probabilities to be constant in time and determined by the following logistic functions:

$$p = \Pr[S_t = 0 \mid S_{t-1} = 0] = \frac{\exp(\theta_0)}{1 + \exp(\theta_0)} \qquad (4)$$

$$q = \Pr[S_t = 1 \mid S_{t-1} = 1] = \frac{\exp(\theta_1)}{1 + \exp(\theta_1)} \qquad (5)$$

where θ_0 and θ_1 are unrestricted parameters. Based on probabilities in (4) and (5), one can calculate the mean duration (d_{st}) of economic recession and expansion regimes by using expressions 1/(1-p) and 1/(1-q), respectively. The duration of each regime can be different, but it will be constant in time since the transition probability matrix is fixed.

Model (1)-(2) is estimated by the use of the filter described in Hamilton (1989). The filter is implemented by following the steps below:

i) one calculates

$$\begin{aligned} \Pr[S_t = s_t, S_{t-1} = s_{t-1}, \ldots, S_{t-p} = s_{t-p} \mid I_{t-1}] = \Pr[S_t = s_t \mid S_{t-1} = s_{t-1}] \times \\ \Pr[S_{t-1} = s_{t-1}, \ldots, S_{t-p} = s_{t-p} \mid I_{t-1}] \end{aligned} \qquad (6)$$

where $\Pr[S_t = s_t \mid S_{t-1} = s_{t-1}]$ is given by (3) and $I_{t-1}=\{\Delta y_{t-1}, \Delta y_{t-2}, \ldots, \Delta y_{-p+1}, u^-_{t-1}, \ldots, u^-_{-p+1}, u^+_{t-1}, \ldots, u^+_{-p+1}\}$ is the set of information available at time t-1;

ii) one calculates the joint density of Δy_t and ($S_t, \ldots, S_{t-p}$) using the expression

$$\begin{aligned} & f(\Delta y_t, S_t = s_t, S_{t-1} = s_{t-1}, \ldots, S_{t-p} = s_{t-p} \mid I_{t-1}) = \\ & \quad f(\Delta y_t \mid S_t = s_t, S_{t-1} = s_{t-1}, \ldots, S_{t-p} = s_{t-p}, I_{t-1}) \times \\ & \quad \Pr[S_t = s_t, S_{t-1} = s_{t-1}, \ldots, S_{t-p} = s_{t-p} \mid I_{t-1}] \end{aligned} \qquad (7)$$

where

$$\begin{aligned} & f(\Delta y_t \mid S_t = s_t, S_{t-1} = s_{t-1}, \ldots, S_{t-p} = s_{t-p}, I_{t-1}) = \\ & \frac{1}{\sqrt{2\pi}\sigma} \exp[-\frac{1}{2\sigma^2}((\Delta y_t - \mu_{s_t}) - \phi_1(\Delta y_{t-1} - \mu_{s_{t-1}}) - \ldots - \phi_p(\Delta y_{t-p} - \mu_{s_{t-p}}) - \\ & \quad \gamma^-_{S_{t-1},1} u^-_{t-1} - \ldots - \gamma^-_{S_{t-p},p} u^-_{t-p} - \gamma^+_{S_{t-1},1} u^+_{t-1} - \ldots - \gamma^+_{S_{t-p},p} u^+_{t-p})^2]; \end{aligned} \qquad (8)$$

iii) the density of Δy_t given I_{t-1} is obtained by

$$f(\Delta y_t \mid I_{t-1}) = \sum_{s_t=0}^{1} \ldots \sum_{s_{t-p}=0}^{1} f(\Delta y_t, S_t = s_t, S_{t-1} = s_{t-1}, \ldots, S_{t-p} = s_{t-p} \mid I_{t-1}); \tag{9}$$

iv) the joint density of states given I_t is calculated using the expression

$$\Pr[S_t = s_t, S_{t-1} = s_{t-1}, \ldots, S_{t-p} = s_{t-p} \mid I_t] = \frac{f(\Delta y_t, S_t = s_t, S_{t-1} = s_{t-1}, \ldots, S_{t-p} = s_{t-p} \mid I_{t-1})}{f(\Delta y_t \mid I_{t-1})} \tag{10}$$

v) finally, using (10), the joint density of states (S_t,..., S_{t-p+1}) is obtained by

$$\Pr[S_t = s_t, S_{t-1} = s_{t-1}, \ldots, S_{t-p+1} = s_{t-p+1} \mid I_t] = \sum_{s_{t-p}=0}^{1} \Pr[S_t = s_t, S_{t-1} = s_{t-1}, \ldots, S_{t-p} = s_{t-p} \mid I_t]. \tag{11}$$

Following Hamilton (1989), we implemented the filter at t=1 with the unconditional joint probability $\Pr[S_0=s_0, S_{-1}=s_{-1}, \ldots, S_{-p+1}=s_{-p+1}]$. To obtain this probability, we adjusted $\Pr[S_{-p+1}=0]=\pi$ and $\Pr[S_{-p+1}=1]=1-\pi$, where π and $1-\pi$ are ergodic probabilities of the Markov process, and we calculated

$$\Pr[S_\tau = s_\tau, \ldots, S_{-p+1} = s_{-p+1}] = \Pr[S_\tau = s_\tau \mid S_{\tau-1} = s_{\tau-1}] \times \Pr[S_{\tau-1} = s_{\tau-1}, \ldots, S_{-p+1} = s_{-p+1}] \tag{12}$$

for τ = -p+2, -p+3,...,0. [4] The interaction on the filter is repeated for t=1,...T.

Two by-products of filter (6)-(11) can be obtained. First, by adding the joint density (10) over the states, it is possible to obtain filtered probabilities of being in a recession or in an expansion, given the set of information available at time t. These probabilities are given by

$$\Pr[S_t = s_t \mid I_t] = \sum_{s_{t-1}=0}^{1} \ldots \sum_{s_{t-p}=0}^{1} \Pr[S_t = s_t, S_{t-1} = s_{t-1}, \ldots, S_{t-p} = s_{t-p} \mid I_t], \; s_t = 0,1 \tag{13}$$

and provide information about the regime in which the series Δy_t is most likely to be at each point of the sample. The second by-product of the filter is the assessment of the density of Δy_t given I_{t-1}. Since in the interaction on (6) and (11), the vector of parameters $\lambda = \{p, q, \sigma, \mu_0, \mu_1, \phi_1, \ldots, \phi_p, \gamma^-_{0,1}, \ldots, \gamma^-_{0,p}, \gamma^-_{1,1}, \ldots, \gamma^-_{1,p}, \gamma^+_{0,1}, \ldots, \gamma^+_{0,p}, \gamma^+_{1,1}, \ldots, \gamma^+_{1,p}\}$ is supposedly known and fixed, the log-likelihood function $L(\lambda)$ assessed at λ can be calculated as:

$$L(\lambda) = \sum_{t=1}^{T} \log f(\Delta y_t \mid I_{t-1}; \lambda). \tag{14}$$

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[4] See Hamilton (1989, 1994) for more detailed information.

The value of λ that maximizes the log-likelihood function $L(\lambda)$ can be found through a numerical optimization routine implemented from an initial given vector of parameters λ_0.

2.1.1 Determination of the number of regimes

For the specification of the Markov-switching model, it is important to know whether it provides a more adequate characterization of data comparatively to a model with constant coefficients. A formal procedure is to use the likelihood ratio (LR) statistic to test the null hypothesis that a one-regime process generates data that run counter to the alternative hypothesis that these data are generated by a two-regime model. However, this test is problematic because regularity conditions are not maintained under the null hypothesis due to the presence of nuisance parameters and singularity of the information matrix. Hansen (1992) and Garcia (1998) proposed some procedures for deriving the asymptotic distribution of the LR statistics for Markov-switching models. While the procedures proposed by Hansen (1992) require a heavy computational load and provide only the p-values that are an upper threshold for the real p values, the critical values for the LR test proposed by Garcia (1998) are not valid for our specification (1).

An alternative to the formal test of hypothesis is to choose the number of Markov states based on information criteria, such as Akaike (AIC), Schwarz (SC) and Hannan-Quinn (HQ). Psaradakis and Spagnolo (2003) assess the performance of these criteria when selecting the number of regimes in autoregressive Markov-switching models and show that the selection process based on AIC succeeds in determining the correct number of states if the sample size and the changes in parameters are not too small. Smith et al. (2006) derive a new criterion to select the number of regimes and variables in Markov-switching models, the Markov-switching criterion (MSC), and demonstrate that it has a good performance in regression and autoregressive models, with one and several states in large and small samples and with low and high noise. These authors also show that the MSC does not select a larger number of states as occurs in the data generating process, as the AIC generally does.

In this paper, we use three procedures to check whether the Markov-switching model is more appropriate than the linear model. The first two use the AIC and MSC, expressed by:

$$AIC = \frac{-2L(\Delta y_t, \hat{\lambda})}{T} + \frac{2N}{T} \tag{15}$$

$$MSC = -2L(\Delta y_t, \hat{\lambda}) + \sum_{l=0}^{1} \frac{\hat{T}_l(\hat{T}_l + 2K)}{\hat{T}_l - 2K - 2} \tag{16}$$

where $L(\Delta y_t, \hat{\lambda})$ is the value of the log-likelihood function in $\lambda = \hat{\lambda}$, N is the number of parameters of the model, T is the number of observations, $T_0 = \sum_t \Pr(S_t = 0 \mid I_T)$, $T_1 = \sum_t \Pr(S_t = 1 \mid I_T)$, $\Pr(S_t = s_t \mid I_T)$ is the smoothed probability of regime s_t (0,1) and K is the number of explanatory variables plus the constant term. The third procedure consists of the approximation to the asymptotic distribution of the LR statistics proposed by Ang and Bekaert (1998). As shown by these authors, the asymptotic distribution of the LR statistic between 1 and 2 regimes can be approximated by a chi-square distribution, where the number of degrees of freedom is given by the number of nuisance parameters of the two-regime model plus the number of linear restrictions imposed on the one-regime model by the two-regime model.

2.2 Symmetry tests for the real effects of monetary shocks

After estimating model (1), we tested the possible asymmetries in the real effects of the monetary policy by imposing restrictions on the sum of parameters $\gamma^{j}_{S_t,i}$, where j=-,+ and i=1,...,p. The statistical significance of the restrictions imposed on the model is assessed by the usual Wald test.[5] The Wald test is preferred to the LR test since the former allows testing the imposed restrictions by estimating only the unrestricted model.

As mentioned in the Introduction, we seek to determine the existence of three types of asymmetry in the effects of monetary policy on output: i) asymmetry between negative and positive monetary shocks in each phase of the business cycle; ii) asymmetry between countercyclical monetary shocks; iii) asymmetry of monetary

[5] See Greene (2000) for details about the Wald test.

shocks between the business cycle phases. For that purpose, we test the following null hypotheses of symmetry:

- Symmetry in the effects of positive and negative monetary shocks

$$H_0 : \sum_{i=1}^{p} \gamma_{0,i}^{-} = \sum_{i=1}^{p} \gamma_{0,i}^{+} \tag{17}$$

$$H_0 : \sum_{i=1}^{p} \gamma_{1,i}^{-} = \sum_{i=1}^{p} \gamma_{1,i}^{+} \tag{18}$$

- Symmetry in the effects of countercyclical monetary shocks

$$H_0 : \sum_{i=1}^{p} \gamma_{0,i}^{+} = \sum_{i=1}^{p} \gamma_{1,i}^{-} . \tag{19}$$

- Symmetry in the effects of monetary shocks between recessions and expansions

$$H_0 : \sum_{i=1}^{p} \gamma_{0,i}^{-} = \sum_{i=1}^{p} \gamma_{1,i}^{-} \tag{20}$$

$$H_0 : \sum_{i=1}^{p} \gamma_{0,i}^{+} = \sum_{i=1}^{p} \gamma_{1,i}^{+} \tag{21}$$

Under null hypotheses (17) through (21), the Wald test has a chi-square distribution with 1 degree of freedom.

3 Data

Model (1) is estimated using monthly frequency data for the period between July 1995 and August 2006. As the monthly data for the real GDP are not available for the study period, we used the seasonally adjusted monthly industrial production index as proxy for the real GDP and then calculated the growth rate of output (Δy_t) as $ln(y_t/y_{t-1})x100$, where *ln* denotes the natural logarithm. The series of the seasonally adjusted industrial production index was obtained from IBGE.[6]

The series of monetary shocks (u_t) is obtained through the estimation of a vector autoregressive (VAR) model of order *p*.[7] We estimate the VAR model using three variables: i) the natural logarithm of the industrial production index; ii) the monthly

[6] IBGE – Brazilian Institute of Geography and Statistics.
[7] For details on VAR models, see Sims (1980), Bernanke (1986) and Enders (1995).

inflation rate defined by $ln(IPCA_t / IPCA_{t-1})x100$, where IPCA is the Brazilian consumer price index calculated by IBGE; iii) monthly overnight interest rate (SELIC), regarded as a monetary policy instrument. The data used were obtained from the Central Bank of Brazil and from IBGE. The VAR is estimated using monthly data for the period between January 1995 and August 2006. We followed the recommendation of Sims (1980) and Doan (1992) and included all the variables in levels in the VAR model.[8] The selection of order p was based on the multivariate AIC, SC and HQ criteria, and also on the results of multivariate LM tests for autocorrelation of the system residuals, White test for the heteroskedasticity of system residuals and of each equation, LM-ARCH test for autoregressive conditional heteroskedasticity (ARCH) in the residuals of each equation and the Jarque-Bera (JB) test for normality of residuals.

Table 1 shows the information criteria for the VAR(p) models, with p $\in 1,...,16$. All the information criteria indicate that VAR(1) is the most appropriate specification. However, the specification tests for the system residuals and for each equation indicate first-order autocorrelation problems, heteroskedasticity, ARCH effect and non-normality (see Table 2). Given that autocorrelation and/or heteroskedasticity problems also persisted in models VAR(2) through VAR(4), we decided to use VAR(5) in order to obtain the monetary shock series (u_t).[9]

Table 1
Information criteria for VAR

p	AIC	SC	HQ
1	-4.3694	-4.0950	-4.2579
2	-4.3600	-3.8798	-4.1649
3	-4.2387	-3.5528	-3.9601
4	-4.1294	-3.2377	-3.7672
5	-4.1158	-3.0183	-3.6700
6	-4.0634	-2.7602	-3.5341
7	-3.9517	-2.4427	-3.3388
8	-3.9589	-2.2441	-3.2623
9	-3.8605	-1.9400	-3.0804
10	-3.7851	-1.6588	-2.9214
11	-3.7212	-1.3891	-2.7739
12	-3.6871	-1.1493	-2.6562
13	-4.3694	-0.9466	-2.5757
14	-4.3600	-0.8572	-2.6085
15	-4.2387	-0.5821	-2.4557
16	-4.1294	-0.3207	-2.3164

[8] Sims et al. (1990), Hendry (1996) and Bernanke and Mihov (1997) underscore that the specification of the VAR model with the variables in levels yields consistent estimates regardless of the existence of a cointegration relationship, where the specification in the first differences is inconsistent if the variables are cointegrated.
[9] All the estimated VAR models had problems with the normality of system residuals and in at least one equation.

Table 2
Specification tests for VAR residuals

Multivariate LM test for autocorrelation of residuals										
p	**VAR(1)**		**VAR(2)**		**VAR(3)**		**VAR(4)**		**VAR(5)**	
	Value	**P-value**	**Value**	**P-value**	**Value**	**P-value**	**Value**	**P-value**	**Value**	**P-value**
1	23.2	0.0058	19.2	0.0232	4.26	0.0001	6.60	0.0554	15.7	0.0742
2	6.05	0.7349	5.51	0.7882	36.3	0.0000	23.2	0.0059	9.82	0.3651
3	6.17	0.7228	9.25	0.4144	14.4	0.1079	13.7	0.1344	11.9	0.2220
4	7.59	0.5763	6.91	0.6460	10.2	0.3338	15.6	0.0768	6.29	0.7102
5	6.99	0.6393	3.83	0.9224	4.14	0.9019	4.86	0.8466	6.58	0.6803
6	5.50	0.7885	6.62	0.6764	6.90	0.6474	8.48	0.4861	6.76	0.6619
7	7.91	0.5428	8.38	0.4962	8.46	0.4890	9.60	0.3839	10.8	0.2909
8	4.96	0.8374	5.67	0.7727	5.70	0.7697	5.45	0.7932	4.48	0.8771
9	6.31	0.7081	7.53	0.5822	8.78	0.4574	8.68	0.4675	7.63	0.5714
10	6.42	0.6969	7.47	0.5879	7.08	0.6283	7.13	0.6232	8.72	0.4638

Heteroskedasticity, autoregressive conditional heteroskedasticity, and normality tests[1]					
	Test	**System**	**Output**	**Inflation**	**Selic**
VAR(1)	ARCH(1)	-	0.0000	0.9796	0.1649
	White	0.0000	0.0000	0.4930	0.0995
	JB	0.0000	0.0000	0.0000	0.0000
VAR(2)	ARCH(2)	-	0.0008	0.4832	0.3532
	White	0.0001	0.0000	0.2285	0.5333
	JB	0.0000	0.0000	0.0000	0.0000
VAR(3)	ARCH(3)	-	0.8766	0.5799	0.5184
	White	0.0000	0.0000	0.4277	0.5890
	JB	0.0000	0.0000	0.0004	0.0000
VAR(4)	ARCH(4)	-	0.9477	0.6947	0.7698
	White	0.1021	0.1675	0.6952	0.8777
	JB	0.0000	0.0000	0.2554	0.0000
VAR(5)	ARCH(5)	-	0.8984	0.6807	0.5918
	White	0.2388	0.3025	0.7066	0.9586
	JB	0.0000	0.0000	0.2741	0.0000

Note: [1] The values refer to p-values.

The residuals of the Selic rate equation do not necessarily represent a true monetary shock, since they can be correlated with the residuals of other equations in the VAR model. Thus, to estimate structural monetary shocks, we followed the recursive identification framework proposed by Sims (1980) – Choleski's orthogonalization. The ordering of variables in the VAR model was output, inflation, and Selic rate.[10] Since monthly data are used, it is reasonable to assume that the output and inflation rate are not contemporaneously affected by the Selic rate. The imposition that the monetary policy reacts contemporaneously to output shocks and to shocks to the inflation rate is quite arguable since the data on inflation and output are made publicly available with some delay. Nevertheless, it is plausible to assume that by deciding on the value of the

[10] This ordering has been used in several works, such as in Bernanke and Blinder (1992), Christiano et al. (1996), Bernanke and Mihov (1998) and Minella (2003).

Selic rate in a given month the Central Bank has access to some current indicators of aggregate output and inflation.

After obtaining the structural monetary shock series (u_t), the next step involved obtaining different types of the shocks taken into consideration in model (1). We followed Cover (1992) and obtained the series of positive and negative monetary shocks using the following definitions:

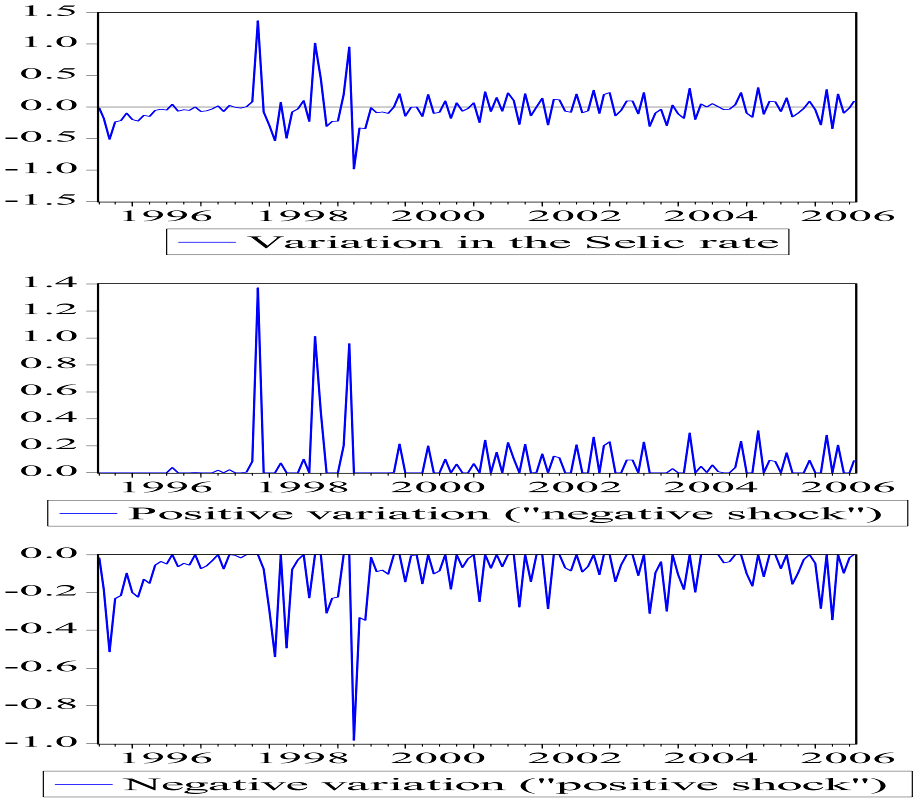
$$\begin{aligned} u_t^{+} &= \min(u_t, 0) \\ u_t^{-} &= \max(u_t, 0) \end{aligned} \qquad (22)$$

where u_t^{+} is a positive monetary shock and u_t^{-} is a negative monetary shock.

Figure 1 shows the behavior of monetary shocks u_t, u_t^{+} and u_t^{-} between July 1995 and August 2006. Monetary shocks had sudden increases in November 1997, August 1998 and March 1999. Whereas the first two increases occurred in periods in which external shocks (Asian and Russian crises) affected the Brazilian economy, jeopardizing the exchange rate band regime, the third shock took place in the month in which the Selic rate was increased in order to reduce the inflation expectations and the high volatility of the exchange rate following the adoption of a floating exchange rate regime in January 1999. Positive monetary shocks have been more and more frequent and seem to be more serially correlated than negative monetary shocks.

Figure 1
Monetary shocks obtained from the VAR model

Figure 2
Variations in the Selic rate



To check the robustness of results, we used the variation in the Selic rate as an alternative method to measure monetary policy. This variable is not appropriate for measuring monetary policy shocks since it is partially endogenous. However, its use is acceptable given that the monetary shocks detected in the VAR model are generated regressors in (1). As pointed out by Pagan (1984), the presence of generated regressors may imply inconsistent standard deviations of the estimated parameters, as well as of the tests based on these standard deviations. Therefore, the use of a monetary policy instrument that is not a generated regressor is an important way to test the robustness of results. Since we considered the variation in the Selic rate a monetary policy measure, the negative “monetary shock” series (positive variations in the interest rate) and the positive ones (negative variations in the interest rate) were obtained according to the definitions described in (22) and can be observed in Figure 2.

4 Results

In this section, we report the results of the Markov-switching models estimated to check whether the effects of the monetary policy on output are asymmetric. The estimates of the parameter vector λ were obtained by the maximization of the log-likelihood function (14). All estimates were made using the Optimum procedure of the Gauss software. The numerical optimization was made using the BFGS (Broyden, Fletcher, Goldfarb and Shanno) algorithm described in Gill et al. (1981).

A common difficulty in estimating Markov-switching models lies in the fact that the log-likelihood function does not have a global maximum (Hamilton, 1991). Additionally, there are often several local maxima that yield similar values for the log-likelihood function, but different parameter estimates. Owing to these problems, we followed the recommendations made by Boldin (1996), which consist in using different initial values for the optimization routine and checking whether each set of parameter estimates is plausible. For some estimated models, we noted that the local maximum with the largest log-likelihood function value was characterized by a transition probability *p* equal to zero and/or filtered recession probabilities (regime 0) that captured only outliers in the data. Since the models with these characteristics imply that only one state practically persists throughout the sample period, we consider an alternative local maximum whose estimate of the parameter vector allows assessing the periods of recession and expansion more properly.

4.1 Estimates for the MS model with VAR shocks

Initially, model (1) was estimated by taking into account positive and negative monetary shocks obtained from the VAR model. We refer to this model as MS(2)-ARX1(p).[11] Table 3 shows a set of LR tests used to determine the autoregressive order of the model. The results indicate that the optimal number of lags is p=7.

[11] MS(2)-ARX1(p) denotes a two-regime Markov-switching model and p lags of the endogenous variable and of the exogenous regressors (positive and negative monetary shocks) obtained from a VAR.

Table 3
LR test for the selection of p in the MS(2)-ARX1(p) model

Test	Value	No. of restrictions	P-value
7 against 6	16.25	5	0.0062
7 against 5	24.89	10	0.0056
7 against 4	39.8	15	0.0005
7 against 3	46.79	20	0.0006
7 against 2	50.96	25	0.0016
7 against 1	65.76	30	0.0002
6 against 5	8.64	5	0.1243
6 against 4	23.55	10	0.0089
6 against 3	30.54	15	0.0101
6 against 2	34.72	20	0.0216
6 against 1	49.51	25	0.0024
5 against 4	14.91	5	0.0108
5 against 3	21.91	10	0.0156
5 against 2	26.08	15	0.0372
5 against 1	40.87	20	0.0039
4 against 3	6.99	5	0.2214
4 against 2	11.17	10	0.3444
4 against 1	25.96	15	0.0384
3 against 2	4.17	5	0.5252
3 against 1	18.97	10	0.0406
2 against 1	14.79	5	0.0113

To check whether the selection of the MS(2)-ARX1(7) model is appropriate, we used two strategies. First, we used the AIC and MSC criteria and the LR test proposed by Ang and Bekaert (1998) in order to verify whether the Markov-switching model is more suitable than the model with constant coefficients. Additionally, we performed tests to check the absence of serial correlation and autoregressive conditional heteroskedasticity in the standardized prediction errors in the MS(2)-ARX1(7) model. These errors were defined as:

$$\xi_t = \sum_{s_t, s_{t-1}, \ldots, s_{t-7}} \frac{\Delta y_t - E[\Delta y_t \mid s_t, s_{t-1}, \ldots, s_{t-7}, I_{t-1}]}{\sigma} \times \Pr[s_t, s_{t-1}, \ldots, s_{t-7} \mid I_{t-1}]. \tag{23}$$

We tested the null hypothesis that there is no autocorrelation in ξ_t up to k lag in two different ways. First, we regressed ξ_t under a constant and k lags of ξ_t and used the F statistic to test the null hypothesis that the autoregressive coefficients of this auxiliary regression are jointly equal to zero.[12] Thereafter, we used the Ljung-Box (LB) test statistic.[13] Under the null hypothesis of lack of serial correlation up to k lag, the LB(k) statistic is asymptotically distributed as a χ^2 with k degrees of freedom. To test the null hypothesis that there is no ARCH effect up to order q, we used the ARCH(q) statistic,

[12] This procedure has been used by Goodwin (1993).
[13] See Ljung and Box (1978).

which was built using TR^2 (where T is the number of observations and R^2 is R-squared) of an auxiliary regression that models the squared ξ_t as an autoregressive process of order q. Under H_0, the ARCH(q) statistic is distributed asymptotically as a χ^2 with q degrees of freedom.[14]

Table 4 shows the results of the comparison between the Markov-switching model and the specification with constant coefficients. As can be observed, the MSC and the LR test carried out for a significance level of 2% suggest that the two-regime model represents the data more appropriately than the linear model. The specification test results shown in Table 5 suggest absence of significant autocorrelation and ARCH effects in the prediction errors of the model.

Table 4
Information criteria and LR test of the MS(2)-ARX1(7) model versus the linear model

	MS model	**Linear model**
AIC	3.9976	3.9786
MSC	632.24	643.01
Log-lik	-213.85	-229.64
LR test – Ang and Bekaert (1998)		
Value	**No. of restrictions**	**P-value**
31.60	17	0.0169

Table 5
Specification tests for the MS(2)-ARX1(7) model

F test for autocorrelation up to order k		
	Value	P-value
AR(12)	0.43	0.9497
AR(24)	0.45	0.9841
Ljung-Box test for autocorrelation up to order k		
	Value	P-value
LB(12)	5.30	0.9470
LB(24)	13.80	0.9510
Test for ARCH effects up to order q		
	Value	P-value
ARCH(8)	8.44	0.3917

Table 6 shows the maximum likelihood estimates for the parameters of the MS(2)-ARX1(7) model. The estimated values for μ_{St} indicate that the economy has a mean growth rate of -1.17% per month (-14.04% per annum) in recessions (state 0) and of 0.728% per month (8.74% per annum) in expansions (state 1). The probabilities

[14] In Markov-switching models, the LB and ARCH tests on the standardized errors have been used by several authors, such as Driffill and Sola (1998), Maheu and McCurdy (2000), Ravn and Sola (2004) and Kim et al. (2004).

of remaining in each regime are estimated at 0.735 for state 0 and at 0.946 for state 1. This implies that recessions last on average 3.8 months, where periods of expansion last on average 18.64 months. All of these results indicate that the estimated Markov-switching model captures asymmetries in the business cycles, characterizing recessions as short periods of sudden drops in output and expansions as periods of gradual and lasting increase in output.

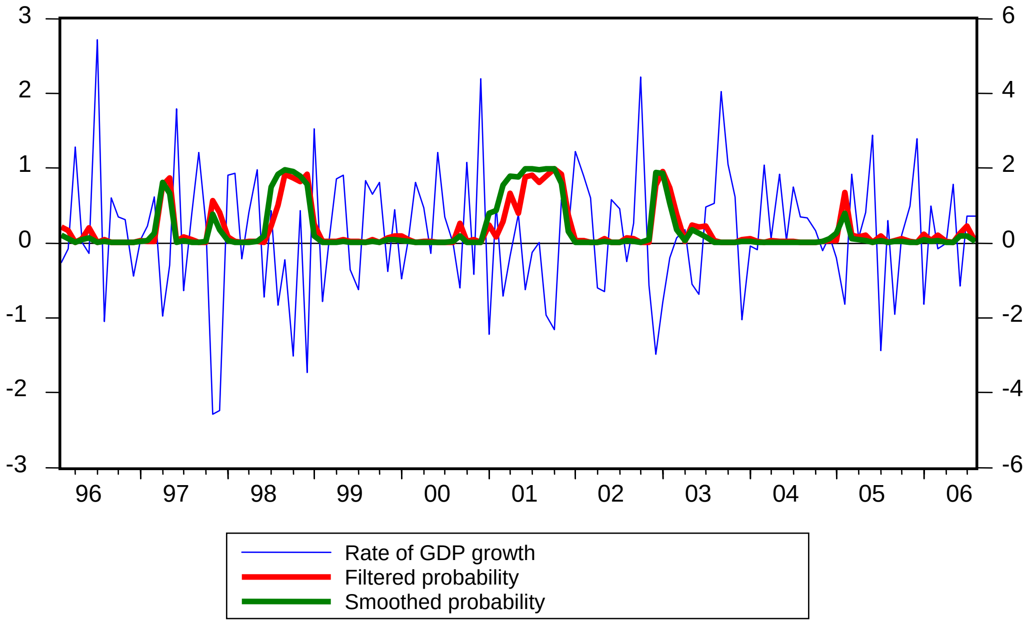
Table 6
Parameter estimation for the MS(2)-ARX1(7) model

Parameter	Regime 0		Regime 1	
	Estimate	Standard errors	Estimate	Standard errors
μ_{S_t}	-1.1738*	0.3164	0.7280*	0.1523
β_1	-0.3172*	0.0939	-0.3172*	0.0939
β_2	-0.1865**	0.0926	-0.1865**	0.0926
β_3	-0.3055*	0.0929	-0.3055*	0.0929
β_4	-0.0820 n.s	0.0849	-0.0820 n.s	0.0849
β_5	-0.1833**	0.0807	-0.1833**	0.0807
β_6	-0.1051 n.s	0.0795	-0.1051 n.s	0.0795
β_7	0.0225 n.s	0.0695	0.0225 n.s	0.0695
$\gamma^-_{S_{t-1},1}$	-2.1926 n.s	1.5060	-5.2377*	0.8658
$\gamma^-_{S_{t-2},2}$	3.3960***	1.7855	-1.0456 n.s	0.9335
$\gamma^-_{S_{t-3},3}$	-4.8438*	1.7631	0.2106 n.s	0.9238
$\gamma^-_{S_{t-4},4}$	3.0367***	1.6720	-2.4248*	0.8434
$\gamma^-_{S_{t-5},5}$	-5.0492*	1.7717	-0.5419 n.s	0.8532
$\gamma^-_{S_{t-6},6}$	2.6511 n.s	1.8676	-0.6683 n.s	0.8427
$\gamma^-_{S_{t-7},7}$	2.1596 n.s	1.8285	-2.3243 *	0.8598
$\gamma^+_{S_{t-1},1}$	-0.0909 n.s	1.2257	-0.3766 n.s	1.5483
$\gamma^+_{S_{t-2},2}$	-7.0426*	2.6244	-1.4112 n.s	1.5126
$\gamma^+_{S_{t-3},3}$	0.3926 n.s	2.8931	-3.8426 *	1.4271
$\gamma^+_{S_{t-4},4}$	-6.8694 *	2.2504	0.9764 n.s	1.5554
$\gamma^+_{S_{t-5},5}$	-4.2807 ***	2.5595	-2.1910 n.s	1.4725
$\gamma^+_{S_{t-6},6}$	4.4145***	2.3742	0.2470 n.s	1.4213
$\gamma^+_{S_{t-7},7}$	3.9618***	2.0644	3.1888**	1.5390
σ	1.0583*	0.0784	1.0583*	0.0784
p/q	0.7351*	0.1019	0.9463*	0.0257
d_{st}	3.7749	-	18.617	-
$\Sigma\gamma^-$	-0.84	-	-12.03	-
$\Sigma\gamma^+$	-9.51	-	-3.41	-
Log-lik.	-213.85			

Note: * Significant at 1%. ** Significant at 5%. *** Significant at 10%. n.s Not significant.

The behavior of the growth rate of output and of recession probabilities is shown in Figure 3. The filtered probability of recession can be understood as an optimal inference on this regime at time t using the information available up to time t, whereas the smoothed probability of recession is concerned with the inference on this state using all the available information.[15] By looking at Figure 3, one can perceive that the time paths of the filtered and smoothed probabilities are quite similar, suggesting that the estimates of recession periods can be obtained through recursive one-step ahead estimates (filtered probabilities) or by using all the available information (smoothed probabilities). The specific dating of recession periods can be obtained by the rule that connects observation t to state 0 if the smoothed probability of this regime is greater than 0.5 (Hamilton, 1989). The application of this rule to the estimated model allows making a distinction between two periods of longer recessions (1998:7-1998:12 and 2001:3-2001:11) and two recessions with a duration no longer than three months (1997:4-1997:5 and 2002:12-2003:2).

Figure 3
Behavior of the GDP growth rate and of filtered and smoothed probabilities of recession for the MS(2)-ARX1(7) model



[15] The smoothed probability of recession was obtained through Kim's (1994) algorithm.

The effects of monetary shocks on output are measured by parameters $\gamma^{j}_{S_t,i}$. The coefficients $\gamma^{-}_{0,i}$ ($\gamma^{-}_{1,i}$) can be interpreted as the effect on the growth rate of output of an unexpected one-percentage point increase in the Selic rate at *t-i* if the economy was in a recession (expansion) when this increase occurred. The coefficients $\gamma^{+}_{0,i}$ ($\gamma^{+}_{1,i}$) have analogous interpretations for the positive monetary shocks. As unexpected changes in the Selic rate in one direction move output towards the opposite direction, the coefficients $\gamma^{j}_{S_t,i}$ are expected to be negative.

The estimates for parameters $\gamma^{j}_{S_t,i}$ are shown in Table 6. Based on the signs and statistical significance of these coefficients, it is possible to notice that the predicted inverse relationship between monetary policy actions and the output level can only be directly observed for negative monetary shocks in an expansion phase. However, the sum of coefficients of each type of shock for each regime ($\Sigma_i \gamma^{j}_{0,i}$ and $\Sigma_i \gamma^{j}_{1,i}$, for j=-,+) suggests that an unexpected increase in the Selic rate lowers the output in both states of the business cycle, whereas an unexpected decrease in the Selic rate increases the aggregate output.

To determine the statistical significance of the effects of different monetary shocks, we performed a set of Wald tests and imposed different restrictions on the estimated parameters. The results of these tests are shown on the first eight lines of Table 7. On the first four lines, we test the null hypothesis that the coefficients related to a given monetary shock in a given regime are jointly equal to zero. In all cases, this null hypothesis is rejected at a 1% significance level. The subsequent four lines test whether the sum of the coefficients of each shock is equal to zero. Given a significance level of 10%, we can observe that this hypothesis is rejected only for the negative shocks in the state of expansion and positive shocks in the state of recession. This indicates that the effects of the procyclical monetary shocks on the output are neutral.

Now we check whether the effects of different monetary policy actions are asymmetric. To do that, we test the null hypotheses of symmetry explained in section 2.2. In Table 7, we highlight several results. First, the null hypothesis of symmetry between the effects of positive and negative shocks is not rejected at a 10% significance level in state 0, but it is rejected (in favor of $\Sigma\gamma^{-}_{1,i} = -12.03 > \Sigma\gamma^{+}_{1,i} = -3.41$) at a 6% significance level in state 1. This suggests that the real effects of negative monetary shocks are larger than the effects of positive shocks only in the expansion phase of the business cycle. Secondly, the non-rejection of the null hypothesis $\Sigma\gamma_{0,i}^{+} = \Sigma\gamma_{1,i}^{-}$ shows that the impact on the output of a given unexpected decrease in the Selic rate in a recession is equivalent to that of an increase during an expansion. This indicates that the real effects of a countercyclical monetary policy do not depend on the prevailing state of the business cycle when the monetary shock occurs. Thirdly, the Wald test result strongly rejects the null hypothesis that the real effects of negative shocks are the same between the Markov states in favor of a larger effect of these shocks in the expansion state ($\Sigma\gamma_{1,i}^{-} = -12.84 > \Sigma\gamma_{0,i}^{-} = -0.86$). Finally, we cannot reject the null hypothesis of symmetry of real effects of positive monetary shocks between the states of the business cycle.

Table 7
Wald tests for the MS(2)-ARX1(7) model

Null hypothesis	Value of χ^2 statistic	No. of restrictions	P-value
$\gamma_{0,i}^{-} = 0, \forall i$	20.92	7	0.0039
$\gamma_{1,i}^{-} = 0, \forall i$	52.42	7	0.0000
$\gamma_{0,i}^{+} = 0, \forall i$	23.44	7	0.0014
$\gamma_{1,i}^{+} = 0, \forall i$	19.42	7	0.0070
$\Sigma\gamma_{0,i}^{-} = 0$	0.08	1	0.7773
$\Sigma\gamma_{1,i}^{-} = 0$	22.04	1	0.0000
$\Sigma\gamma_{0,i}^{+} = 0$	2.80	1	0.0943
$\Sigma\gamma_{1,i}^{+} = 0$	0.98	1	0.3221
$\Sigma\gamma_{0,i}^{-} = \Sigma\gamma_{0,i}^{+}$	2.17	1	0.1407
$\Sigma\gamma_{1,i}^{-} = \Sigma\gamma_{1,i}^{+}$	3.60	1	0.0578
$\Sigma\gamma_{0,i}^{+} = \Sigma\gamma_{1,i}^{-}$	0.17	1	0.6801
$\Sigma\gamma_{0,i}^{-} = \Sigma\gamma_{1,i}^{-}$	8.26	1	0.0041
$\Sigma\gamma_{0,i}^{+} = \Sigma\gamma_{1,i}^{+}$	1.19	1	0.2753

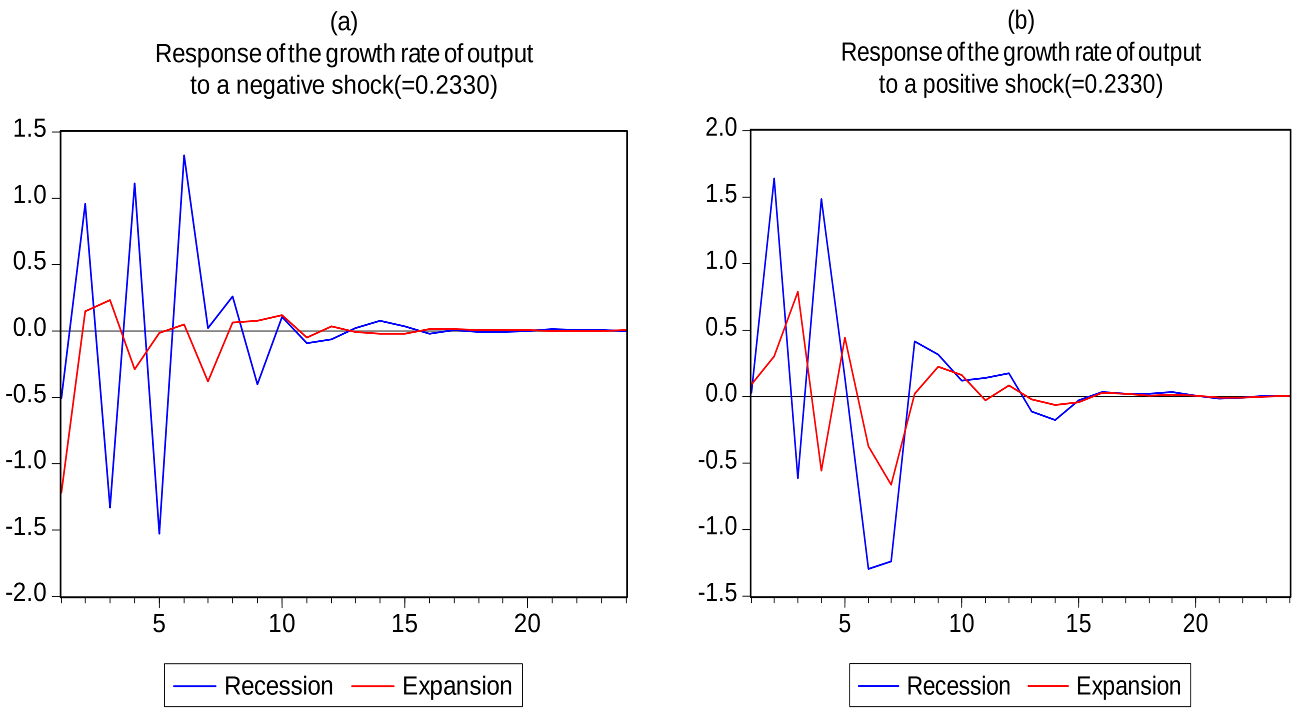
The differences in the real effects of the monetary policy are shown in Figure 4. In Panel a, we show the paths of the growth rate of output in a recession and in an expansion in response to a negative shock (or to an unexpected increase in the Selic rate) equal to 0.233 percentage points implemented at t=0.[16] In Panel b, the responses of the growth rate of output to a positive shock (or to an unexpected decrease in the Selic rate) of -0.233 percentage points are presented for the two Markov regimes. Initially, we

[16] The value 0.2330 is the estimate for the standard deviation of the structural monetary shock series u_t which was obtained from the VAR.

can perceive that the oscillation in output in response to different types of monetary shocks is a common characteristic in both panels. This probably results from the high noise level in the monthly series of the industrial production growth rate. When we compare the paths of output across the different regimes, we also note that the response of output has a greater variability in the recession phase of the business cycle for any of the types of shocks considered.

With regard to the impact of specific monetary shocks, Figure 4 shows that the effect of an unexpected increase in the Selic rate is the reduction in output at t=1, whereas an unexpected decrease in this rate increases output level one month after the shock. The largest decrease in output in response to a negative shock occurs in the fifth month if the economy is in a recession and in the first month if it is in an expansion. With respect to a positive monetary shock, the increase in output reaches its maximum in the second and third months during recession and expansion, respectively.

Figure 4
State-dependent effects of a monetary shock in the VAR model



Finally, one can observe that the response of the growth rate of output accumulated in the 24 months after a negative (positive) shock is equal to -0.09% (1.02%) in the recession regime and to -1.3% (0.37%) in the expansion regime. The differences in these values suggest that: i) in the expansion regime, the negative shocks affect output more strongly than positive shocks; ii) the real effect of a negative shock is larger during an expansion than in a recession; iii) in the recession regime, the real effect of a given positive shock outperforms, in absolute value, the effect of a negative shock; iv) the real effect of a positive shock is larger during an expansion; and v) the real effect of an unexpected increase in the Selic rate is larger than that of an unexpected decrease in this rate during recession. It is important that results iii to v be taken with caution, since the Wald tests did not reject the null hypotheses of symmetry for these cases.

4.2 Checking the robustness of results

To check the robustness of the results shown above, we measured the monetary policy actions by the variation in the Selic rate and we estimated the Markov-switching model (1) with seven lags of each explanatory variable. This model is referred to as MS(2)-ARX2(7).

Table 8
Information criterion and LR test of the MS(2)-ARX2(7) model versus the linear model

	MS model	**Linear model**
AIC	3.8068	3.9804
MSC	611.48	643.23
Log-lik	-201.73	-229.75
LR test – Ang and Bekaert (1998)		
Value	**No. of restrictions**	**P-value**
56.05	17	0.0000

Table 9
Specification tests for the MS(2)-ARX2(7) model

F test for autocorrelation up to order k		
	Value	P-value
AR(12)	0.91	0.5358
AR(24)	1.02	0.4529
Ljung-Box test for autocorrelation up to order k		
	Valor	P-value
LB(12)	12.99	0.3700
LB(24)	23.05	0.5170
Test for ARCH effects up to order q		
	Value	P-value
ARCH(8)	11.47	0.1765

The AIC and MSC values shown in Table 8 indicate that the Markov-switching model is superior to its linear counterpart. Moreover, the LR test proposed by Ang and

Bekaert (1998) rejects the null hypothesis that a one-regime model generates data against the alternative hypothesis that these data are generated by the two-regime model.

Table 9 presents the specification tests for the standardized prediction errors of the MS(2)-ARX2(7) model. The p-values for the calculated values of F statistics, LB and ARCH tests show that the null hypotheses of autocorrelation and of the ARCH effects on the prediction errors are rejected at a 10% significance level.

Table 10
Parameter estimation for the MS(2)-ARX2(7) model

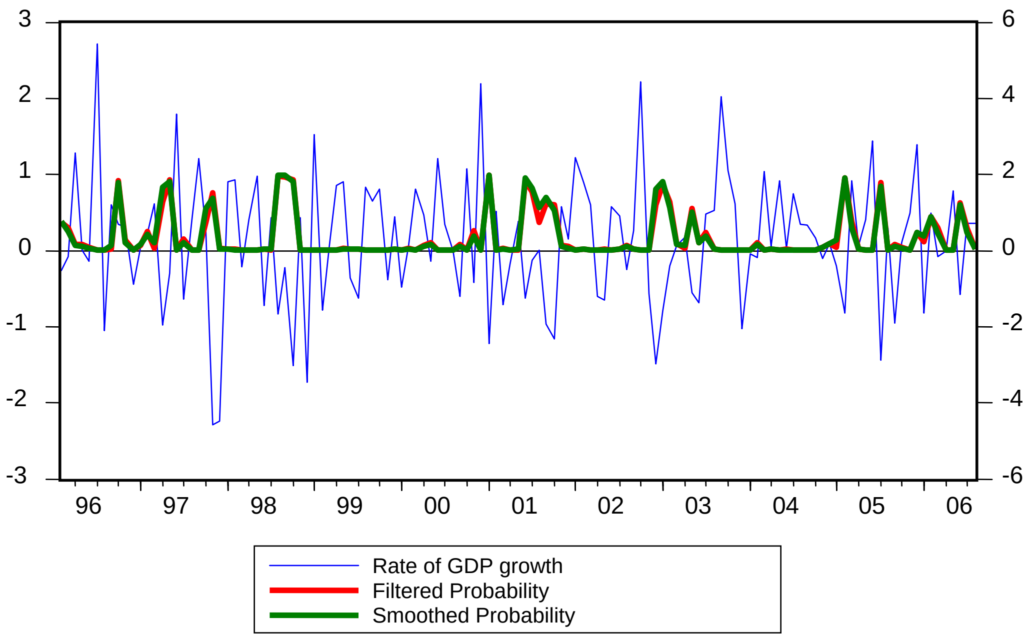
Parameter	Regime 0		Regime 1	
	Estimate	Standard errors	Estimate	Standard errors
μ_{S_t}	-1.2461*	0.2134	0.8601*	0.0897
β_1	-0.2852*	0.0689	-0.2852*	0.0689
β_2	-0.3859*	0.0774	-0.3859*	0.0774
β_3	-0.4209*	0.0853	-0.4209*	0.0853
β_4	-0.2015*	0.0708	-0.2015*	0.0708
β_5	-0.3078*	0.0676	-0.3078*	0.0676
β_6	-0.0575$^{n.s}$	0.0680	-0.0575$^{n.s}$	0.0680
β_7	0.1574**	0.0653	0.1574**	0.0653
$\gamma^-_{S_{t-1},1}$	-1.2658$^{n.s}$	1.1741	-6.2263*	0.6471
$\gamma^-_{S_{t-2},2}$	0.8755$^{n.s}$	1.3173	-2.5189*	0.7120
$\gamma^-_{S_{t-3},3}$	-3.4797**	1.3849	-0.9555$^{n.s}$	0.8413
$\gamma^-_{S_{t-4},4}$	-2.9437**	1.4114	-2.3527*	0.7311
$\gamma^-_{S_{t-5},5}$	2.1908$^{n.s}$	1.3927	-2.4040*	0.8693
$\gamma^-_{S_{t-6},6}$	-8.2666*	1.4873	-0.5373$^{n.s}$	0.8640
$\gamma^-_{S_{t-7},7}$	8.9452*	1.3709	-1.6894**	0.6574
$\gamma^+_{S_{t-1},1}$	-6.5526**	2.9220	1.9680***	0.9928
$\gamma^+_{S_{t-2},2}$	-1.5263$^{n.s}$	3.3452	-1.4762$^{n.s}$	0.9792
$\gamma^+_{S_{t-3},3}$	1.7698$^{n.s}$	3.6406	-1.7461**	0.7756
$\gamma^+_{S_{t-4},4}$	3.4051$^{n.s}$	3.3603	-1.3908***	0.7830
$\gamma^+_{S_{t-5},5}$	-22.069*	3.5623	0.0572$^{n.s}$	1.8025
$\gamma^+_{S_{t-6},6}$	21.830*	3.6777	-0.2326$^{n.s}$	0.9104
$\gamma^+_{S_{t-7},7}$	-19.995*	3.4128	1.9529*	0.7236
σ	0.8041*	0.0549	0.8041*	0.0549
p/q	0.3966*	0.0315	0.8904*	0.1105
d_{st}	1.6574	-	9.1208	-
$\Sigma\gamma^-$	-3.94	-	-16.68	-
$\Sigma\gamma^+$	-23.14	-	-0.87	-
Log-lik.	-201.73			

Note: * Significant at 1%. ** Significant at 5%. *** Significant at 10%. $^{n.s}$ Not significant.

Now we consider the parameter estimation of the Markov-switching model. As shown in Table 10, state 0 is characterized by a mean negative growth rate of output (-1.25% per month), whereas state 1 has a mean positive growth rate (0.86% per month). Again, this allows classifying regimes 0 and 1 as recession and expansion.

The estimated transition matrix shows that the probability of remaining in the recession state is equal to 0.40, which implies a mean length of 1.66 months for this regime. Conversely, the duration of the expansion state averages approximately 9 months (q=0.89). When compared to the model in the previous section, the smaller persistence and duration of both Markov states observed herein result from the adjustment of the MS(2)-ARX2(7) model to sudden and transient decreases in output. By observing the filtered and smoothed probabilities in Figure 5, it is possible to note that now the Markov-switching model regards the October-November 1997 dyad and January 2001 and January and July 2005, and June 2006 as periods in which the Brazilian economy was in a recession.

Figure 5
Behavior of the rate of GDP growth and of filtered and smoothed probabilities of recession for the MS(2)-ARX2(7) model



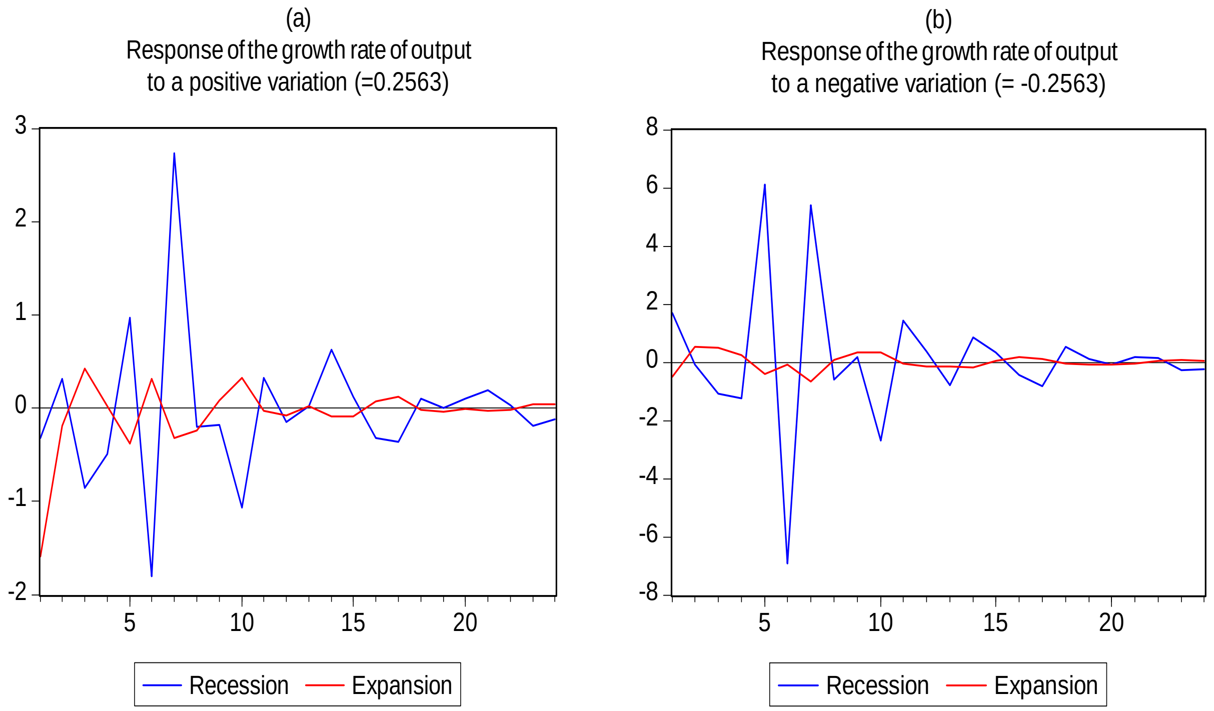
As to the real effects of the monetary policy actions, we can observe in Table 10 that the sum of the coefficients related to each "shock" suggests that a positive variation in the Selic rate ("negative monetary shock") reduces output, whereas a negative variation in the Selic rate ("positive monetary shock") increases output. To check the statistical significance of these effects, we tested the null hypothesis that the sum of the coefficients related to each "shock" is equal to zero. As shown in Table 11, the Wald test results support the non-neutrality of the effects of positive variations in the Selic rate in both states of the business cycle. As far as negative variations in the Selic rate are concerned, we found evidence that the sum of the coefficients related to this policy action is statistically different from zero only in the recession state.

Table 11
Wald tests for the MS(2)-ARX2(7) model

Null hypothesis	Value of χ^2 statistic	No. of restrictions	P-value
$\gamma^-_{0,i} = 0, \forall i$	79.87	7	0.0000
$\gamma^-_{1,i} = 0, \forall i$	114.1	7	0.0000
$\gamma^+_{0,i} = 0, \forall i$	91.65	7	0.0000
$\gamma^+_{1,i} = 0, \forall i$	19.58	7	0.0066
$\Sigma\gamma^-_{0,i} = 0$	2.99	1	0.0838
$\Sigma\gamma^-_{1,i} = 0$	36.03	1	0.0000
$\Sigma\gamma^+_{0,i} = 0$	11.83	1	0.0006
$\Sigma\gamma^+_{1,i} = 0$	0.26	1	0.6101
$\Sigma\gamma^-_{0,i} = \Sigma\gamma^+_{0,i}$	8.68	1	0.0032
$\Sigma\gamma^-_{1,i} = \Sigma\gamma^+_{1,i}$	23.91	1	0.0000
$\Sigma\gamma^+_{0,i} = \Sigma\gamma^-_{1,i}$	0.86	1	0.3537
$\Sigma\gamma^-_{0,i} = \Sigma\gamma^-_{1,i}$	19.97	1	0.0000
$\Sigma\gamma^+_{0,i} = \Sigma\gamma^+_{1,i}$	9.52	1	0.0020

Table 11 shows the set of tests of hypotheses on symmetry of the real effects of different types of monetary policy actions. As in the previous section, we found strong evidence of asymmetry in the real effects of a contractionary and expansionary monetary policy in the expansion regime, as well as in the real effects of a contractionary policy between the states of the business cycle. The difference now lies in our power to reject the null hypotheses of symmetry between the real effects of positive and negative variations in the Selic rate in the recession state, and between the real effects of negative variations in the Selic rate between the states of the business cycle.

Figure 6
State-dependent effects of a variation in the Selic rate



The regime-dependent paths for the growth rate of output in response to a positive (negative) variation of 0.2563% (-0.2563%) in the Selic rate can be seen in Panel a (b) of Figure 6. The differences in these graphs provide clear evidence in favor of asymmetries in the effects of the monetary policy. The accumulated response of the growth rate of output to a given positive (negative) variation in the Selic rate is now equal to -0.53% (2.13%), in a recession, and to -1.67% (0.16%) during an expansion. This corroborates the two findings described in the previous section. The first one is that, in periods of expansion, the impact on output from a negative shock is larger than that of a positive shock, whereas in periods of recession, the opposite applies. The second finding is that a positive (negative) variation in the Selic rate has a stronger effect on output when the economy is in the expansion (recession) state of the business cycle.

5 Conclusions

In this paper, we assessed whether the real effects of monetary shocks in Brazil are asymmetric. In particular, we investigated three types of asymmetry: i) asymmetry between the real effects of negative and positive monetary shocks in each state of the business cycle; ii) asymmetry between the effects of countercyclical monetary shocks; and iii) asymmetry in the real effects of a given monetary shock between the states of the business cycle. To do that, we extended the Markov-switching model proposed by Hamilton (1989) to allow positive and negative monetary shocks to asymmetrically affect the growth rate of output during expansions and recessions. Based on information criteria such as AIC and MSC, and on the LR test proposed by Ang and Bekaert (1998), we observed that the two-regime Markov-switching model represents the data more appropriately than its linear counterpart.

We assessed whether the effects of the monetary policy on output are asymmetric by using the Wald statistics to test a set of null hypotheses of symmetry in the effects of different shocks. The results show that, when monetary policy actions are measured through orthogonalized innovations for the Selic rate in a simple VAR model, there is strong evidence that: i) the real effects of negative monetary shocks are not the same as those of positive shocks in the expansion state; ii) the real effects of negative shocks in the expansion state are different from the effects during a recession. When the variation in the Selic rate was used to measure monetary policy, we also found asymmetries between the real effects of positive and negative variations in the Selic rate in the recession state, and between the real effects of negative variations in the Selic rate between the states of the business cycle. Under no circumstance did we find evidence in favor of asymmetries between the effects of a contractionary monetary policy implemented during an expansion and those of an expansionary policy implemented during a recession.

Finally, we simulated the paths for the growth rate of output in response to positive and negative monetary shocks in each state of the business cycle. The differences observed between these paths reflect the asymmetries in the real effects of the monetary policy. When we calculated the response of the growth rate of output accumulated within 24 months after a given monetary shock, we obtained three results. The first one shows that, during the expansion state, the real effect of a negative monetary shock outperforms, in absolute value, the effect of a positive shock. The second one indicates that, in periods of recession, positive monetary shocks affect the output more strongly than negative shocks. The third and last result shows that the real effect of a negative shock is larger during an expansion, whereas the real effect of a positive shock is larger in the recession state of the business cycle. This body of evidence clearly indicates that the real effects of monetary policy in Brazil depend upon the type of action and upon the state of the business cycle in which the action is taken by the Central Bank.

The results found here leave an important question unanswered: why are the effects of the monetary policy on output asymmetric? As described in the Introduction, the models that yield a convex aggregate supply curve predict that negative monetary shocks affect the output more intensely than positive shocks. For Brazil, this result was only observed when the economy was in an expansion. Additionally, the predicted result for theoretical models with frictions in the credit market that the real effects of monetary shocks are larger during a recession was only observed empirically for positive monetary shocks. Thus, an interesting suggestion for further research is the identification of what factors determine the asymmetry of the effects of monetary policy. Moreover, our study can be extended by considering the possibility that the real effects of large monetary shocks are different from the effects of small monetary shocks.

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