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Article

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Testing asymmetries in Central Bank preferences in a small open economy: A study for Brazil[☆]

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Abstract

In this paper, we extend Surico's (2007a) model to an open economy and test if the Brazilian Central Bank's loss function is asymmetric with regard to positive and negative deviations of the output gap and of the inflation rate from its target. Furthermore, we use tests for structural breaks to investigate changes in the conduct of policy and monetary authority's preferences. The results revealed that the Central Bank reacted more strongly to deviations of inflation from the target and the output gap after 2003. With regard to the monetary authority's preferences, estimates indicate only an asymmetric preference over an above-target inflation rate until mid-2003. After this period, the evidence obtained by estimating the reaction function with inflation expectations showed that the Central Bank has been more averse to positive deviations of inflation from its target.

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Resumo

Neste trabalho, nós estendemos o modelo de Surico (2007a) para uma economia aberta e testamos se a função perda do Banco Central brasileiro é assimétrica em relação a desvios positivos e negativos do hiato do produto e da inflação em relação à meta. Além disso, nós utilizamos testes de quebra estrutural para investigar alterações na condução da política e nas preferências da autoridade monetária. Os resultados revelaram que o Banco Central reagiu mais fortemente a desvios da inflação em relação à meta e ao hiato do produto após 2003. Em relação às preferências da autoridade monetária, as estimativas indicaram, em geral, uma assimetria em favor de uma inflação acima da meta até meados de 2003. Após esse período, as evidências obtidas através da estimação da função

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de reação com expectativas de inflação mostraram que o Banco Central tem sido mais avesso a desvios positivos da inflação em relação à meta.

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Palavras-chave: preferências assimétricas; economia aberta; regras de taxa de juros não lineares; mudança estrutural; Banco Central do Brasil

1. Introduction

Since the early 1990s, the economic literature on monetary policy actions estimated by reaction functions has increasingly expanded. [Taylor \(1993\)](#) rule is probably the most widely known reaction function specification in this literature. According to this rule, the monetary authority responds to deviations of output and inflation from the targets through nominal interest rate movements, regarded as a policy instrument. Another specification that has received considerable attention is the forward-looking reaction function proposed by [Clarida et al. \(1998, 2000\)](#). In this type of policy rule, the policymaker adjusts the current interest rate based on the expected future inflation and output gap rates. These two types of interest rate rules share a common characteristic: they are linear functions related to economic variables. This can be explained by the fact that both specifications are theoretically supported by the linear-quadratic paradigm, where the monetary authority's loss function is assumed to be quadratic while the equations for the economic structure are linear.

At recent times, however, two theoretical approaches have challenged the linear-quadratic framework underlying the linear reaction function. The first approach disproves the assumption that the economic structure is linear. [Orphanides and Wieland \(1999\)](#) derive optimal policy rules for the case where the monetary authority makes use of a quadratic loss function and a zone-linear Phillips curve that allows for nonlinearities in the short-term tradeoff between inflation and output. [Nobay and Peel \(2000\)](#) analyze an optimal discretionary monetary policy under a nonlinear Phillips curve and find out that the monetary authority can no longer eliminate the inflationary bias by defining an output target that is the same as the natural rate. [Dolado et al. \(2005\)](#) demonstrate that the central bank's optimal reaction function for an economy with a nonlinear Phillips curve is the forward-looking interest rate rule augmented to include the interaction between expected inflation and output gap.

The second theoretical approach considers that policymakers can have asymmetric preferences in terms of their goals. According to [Cukierman \(2000\)](#), politicians and the public at large are usually more averse to negative than to positive gap differences from the potential output. The vice chairman of the Federal Reserve, Alan Blinder, states that “in most situations the CB will take far more political heat when it tightens pre-emptively to avoid higher inflation than when it eases pre-emptively to avoid higher unemployment” [Blinder \(1998, pp. 19–20\)](#). Since in democracies independent central banks are not totally insensitive to political organs, this type of asymmetry may be present in the policymaker's loss function. In addition, in periods in which the monetary authority is more concerned with lending credibility to its disinflation policy, the loss ascribed to positive deviations of the inflation rate from the inflation target is likely greater than that from negative deviations with the same magnitude.

The consequences of introducing asymmetric preferences into the monetary authority's loss function have been investigated by several authors. [Cukierman \(2000\)](#) demonstrates that, when the policymaker is unsure about economic conditions and is more sensitive to negative output gaps, there is an inflationary bias even in the case where the effective output target is the potential output of the economy. This finding has been supported by empirical evidence gathered by [Cukierman and Gerlach \(2003\)](#) for a group of 22 OECD countries. [Gerlach \(2000\)](#) and [Surico \(2007a\)](#) show that the Federal Reserve was more concerned about negative than about positive output gaps in the pre-1980 period. [Bec et al. \(2002\)](#) verify that the business cycle phase, measured by output gap, has mattered in the conduct of monetary policy for Central Banks of Germany, the United States, and France. [Cukierman and Muscatelli \(2003\)](#) give evidence of nonlinearities with respect to inflation and to output gap in reaction functions estimated for Germany, the United Kingdom, and the USA. [Dolado et al. \(2004\)](#) observe that Federal Reserve preferences about inflation were asymmetric during the Volcker-Greenspan's term.

Bearing in mind what was presented earlier, the present paper aims to estimate the nonlinear reaction function for the Central Bank of Brazil so as to test asymmetries in its goals regarding the deviations of output gap and inflation from the targets. To achieve that, we extend the theoretical model proposed by [Surico \(2007a\)](#) for an open economy

and we obtain the optimal monetary policy for the monetary authority considering that the loss function is potentially asymmetric. Given that the presence of asymmetries in the goals leads to nonlinear responses of the interest rates to inflation and to output gap, we check whether the policymaker's preferences are symmetric by testing the null hypothesis of a linear reaction function. Moreover, we estimate the Central Bank asymmetric coefficients and test whether they are statistically significant.

Several research studies in the Brazilian literature seek to estimate monetary policy reaction functions.¹ Although a smaller number of these studies assess reaction function nonlinearities, only Aragón and Portugal, 2010 try to check whether there exist asymmetric preferences in the Central Bank of Brazil. The present paper shows some progress in relation to that of Aragón and Portugal (2010) in two points. First, we consider asymmetries in the policymaker's preferences and derive the optimal monetary rule based on a macroeconomic model for a small open economy. Second, we follow Bai and Perron (1998, 2003) and run structural break tests for nonlinear reaction function parameters. The analysis of structural breaks allows investigating possible changes in monetary policy conduct, as well as in the Central Bank asymmetric preference coefficients.

The major findings of this paper can be summarized as follows. First, the structural break tests strongly reject the hypothesis of stable reaction function parameters. For two out of the three monetary rule specifications analyzed, the date estimated for the break indicates that the change in the conduct of the Brazilian monetary policy occurred in the third quarter of 2003. Second, the estimates for the reaction function coefficients reveal that the Central Bank of Brazil has had stronger reactions to deviations of (current and expected) inflation from its target and to output gap in the more recent subperiod. Third, the evidence indicates, in general, that there is not an asymmetric response of the Selic rate to the output gap, suggesting that the loss function of the Central Bank of Brazil is quadratic relative to this state variable. Finally, the coefficient estimates of asymmetric preference concerning inflation show differences in the analyzed monetary rules. For the specification that regards the inflation target as constant, results indicate an increase in asymmetry in favor of above-target inflation after 2003. When we consider the deviation of inflation from a variable target, the asymmetry coefficient of inflation was not significant. Conversely, for the policy rule with deviation of the expected inflation from a variable target, results show that the Central Bank was more averse to a below-target inflation until mid-2003 and that it was averse to an above-target inflation thereafter.

Aside from this introduction, the paper is organized into another four sections. Section 2 introduces the basic theoretical model used in the study, as well as the extension to this model and the derivation of the optimal reaction function for the interest rate. Section 3 presents the reduced-form interest rate rule to be estimated in order to check for the existence of monetary authority's asymmetric goals. Section 4 describes and analyzes the results obtained from the estimations. Section 5 concludes.

2. The theoretical model

2.1. The economic structure for a closed economy

The basic theoretical model on which the present paper is based is shown in Surico (2007a). The model uses the new Keynesian framework analyzed by Clarida et al. (1999), allowing the monetary authority to have asymmetric preferences in terms of goals or targets. More specifically, the monetary authority is allowed to be more averse to negative deviations of the effective output from the potential output and to positive deviations of the inflation rate from its target. The presence of these types of asymmetries is the explanation for possible nonlinear responses of the monetary rule interest rate to variations in inflation and in output.

Following Clarida et al. (1999), we take into account an economy whose evolution can be described by the following equations:

$$x_t = -\varphi(i_t - E_t\pi_{t+1}) + E_t x_{t+1} + e_t \quad (1)$$

$$\pi_t = kx_t + \theta E_t\pi_{t+1} + u_t \quad (2)$$

¹ See, for instance, Minella (2003), Salgado et al. (2005), Bueno (2005), Holland (2005), Soares and Barbosa (2006), Teles and Brundo (2006), Lima et al. (2007) and Barcellos Neto and Portugal (2007).

where x_t is the output gap (that is, the difference between effective output and potential output), π_t is the inflation rate, $E_t x_{t+1}$ and $E_t \pi_{t+1}$ are the expected values for output gap and inflation rate based on the information available at t , i_t is the interest rate, e_t and u_t are a demand shock and a supply shock, respectively, and φ , k and θ are positive constants.²

The IS curve, given by Eq. (1), is a log-linearized version of Euler equation for consumption derived from the optimal decision of households about consumption and saving, after the imposition of the market clearing condition. The expected output gap value shows that, as households would rather reduce consumption over time, the expectation of a higher level of consumption leads to an increase in current consumption, thus boosting the current demand for products.

The Phillips curve, given by Eq. (2), contains overlapping nominal prices, and the firms show a probability ω of keeping the product price constant at any time (Calvo, 1983). Since probability ω is presumably constant and independent from the time of the last adjustment, the average time during which the price remains constant is given by $1/(1 - \omega)$. The discrete nature of price adjustment resulting from that encourages firms to set a higher price, the greater the expectation of future inflation.

Shocks e_t and u_t are given by autoregressive processes:

$$e_t = \rho_e e_{t-1} + \hat{e}_t \quad (3)$$

$$u_t = \rho_u u_{t-1} + \hat{u}_t \quad (4)$$

where $0 \leq \rho_e, \rho_u \leq 1$, \hat{e}_t and \hat{u}_t are random variables with zero mean and standard deviation σ_e and σ_u , respectively.

2.2. Asymmetric goals of the monetary authority

Suppose that monetary policy decisions are taken before the realization of shocks e_t and u_t . Hence, conditional on the information available at the end of the previous period, the monetary authority seeks to choose current inflation rate i_t and a sequence of future interest rates in order to minimize:

$$E_{t-1} \sum_{\tau=0}^{\infty} \delta^\tau L_{t+\tau} \quad (5)$$

subject to the economic structure, given by Eqs. (1) and (2), where δ is the fixed discount factor. The loss function at t is given by:

$$L = \frac{e^{\alpha(\pi_t - \pi^*)} - \alpha(\pi_t - \pi^*) - 1}{\alpha^2} + \lambda \frac{e^{\gamma x_t} - \gamma x_t - 1}{\gamma^2} + \frac{\mu_i}{2} (i_t - i^*)^2 + \frac{\mu_{\Delta i}}{2} (i_t - i_{t-1})^2 \quad (6)$$

where π^* is the inflation target, λ is the relative weight of the deviation of output from potential output, and μ_i and $\mu_{\Delta i}$ are the relative weights of interest rate stabilization around an implicit target, i^* , and of the interest rate at $t - 1$, i_{t-1} .³ The monetary authority is assumed to stabilize inflation around the constant inflation target, π^* , to maintain the output gap at zero and to stabilize the nominal interest rate around target i^* and the nominal interest rate at $t - 1$.⁴

The linear loss function, akin to specification (6), was initially discussed by Varian (1974) and Zellner (1986), and first used in the study on monetary policy by Nobay and Peel (1998). Ever since, this type of loss function has been applied in the analysis of optimal monetary policy by Cukierman and Gerlach (2003), Nobay and Peel (2003), Ruge-Murcia (2003a,b, 2004), Dolado et al. (2004), Surico (2007a,b) and Ikeda (2010).

² The aggregate behavioral Eqs. (1) and (2) are explicitly derived from the optimizing behavior of firms and households in an economy with money and nominal price rigidity (Clarida et al., 1999).

³ Interest rate smoothing is justifiable for several reasons, such as: (i) presence of uncertainties over the values of data and coefficients in the macroeconomic model; (ii) large changes in the interest rate could destabilize the exchange rate and financial markets; (iii) constant variations in the short-term interest rate, even if they are small, would have a remarkable impact on aggregate demand and on the inflation rate. For a theoretical and empirical study about monetary policy interest rate smoothing, see Clarida et al. (1998), Sack (1998), Woodford (1999, 2003) and Sack and Wieland (2000).

⁴ As pointed out by an anonymous referee, a weakness of loss function (6) is that the Central Bank will only be able to stabilize the interest rate when $i^* = i_{t-1}$. Given that i^* is constant, such stabilization will hardly be put into practice.

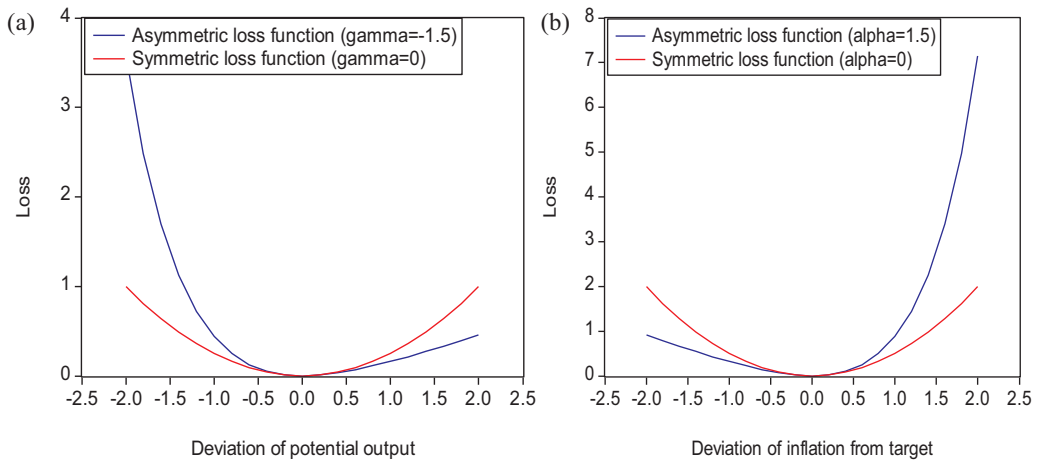


Fig. 1. Symmetric and asymmetric loss function with respect to the output gap (a) and inflation (b).

The great difference in expression (6) is that it allows the policymaker to make a distinction between the positive and negative deviations of output from the potential output and of inflation from its target. As shown in Fig. 1, a negative value of γ indicates that the marginal loss associated with a negative output gap is greater than that of a positive output gap with the same absolute value. This occurs because whenever the output gap is positive, the exponential component of loss function (6) predominates over the linear component, whereas the opposite is observed whenever the output gap is negative. In this case, the monetary authority is said to have a precautionary demand for economic expansion (Cukierman, 2000).

A positive value of α shows that the monetary authority has a precautionary demand for price stability, i.e., the marginal loss of a positive deviation of the inflation rate from its target is larger than that of a negative deviation of the same magnitude (see Fig. 1). This behavior is appropriate for describing a Central Bank concerned with lending credibility to its anti-inflationary policy. Notwithstanding, it is important to underscore that linear specification (6) does not prevent α from being negative, indicating that a below-target inflation rate is costlier than an above-target inflation.

If γ and α tend to zero in Eq. (6), we obtain a symmetric loss function given by:

$$L_t = \frac{1}{2} [(\pi_t - \pi^*)^2 + \lambda x_t^2 + \mu_i (i_t - i^*)^2 + \mu_{\Delta i} (i_t - i_{t-1})^2] \tag{7}$$

The solution to optimization (5) with asymmetric loss function (6) is obtained by assuming that the monetary policymaker takes the expectations of future variables as given and chooses the current interest rate in each period.⁵ Since there is no endogenous persistence in inflation and output gap, the intertemporal optimization problem can be reduced to a sequence of static optimization problems. Therefore, taking the first-order condition and solving for i_t , we get:

$$i_t = (1 - \rho_1) \left\{ i^* + c_1 E_{t-1} \left[\frac{e^{\alpha(\pi_t - \pi^*)} - 1}{\alpha} \right] + c_2 E_{t-1} \left[\frac{e^{\gamma x_t} - 1}{\gamma} \right] \right\} + \rho_1 i_{t-1} \tag{8}$$

and

$$c_1 = \frac{k\varphi}{\mu_i}; \quad c_2 = \frac{\lambda\varphi}{\mu_i}; \quad \rho_1 = \frac{\mu_{\Delta i}}{\mu_i + \mu_{\Delta i}}. \tag{9}$$

From Eq. (8), we note that the optimal nominal interest rate at t reacts nonlinearly to inflation and to the output gap at t . Because c_1 and c_2 are both positive, the monetary authority elevates the nominal interest rate in response to increases in the expected inflation rate and in the expected output gap.

⁵ Palma and Portugal (2011) provide evidence in favor of a discretionary monetary policy in Brazil for the 2000–2010 period.

When γ and α tend to zero, using the L' Hospital rule, we obtain the following reduced form for the interest rate rule (8):

$$i_t = (1 - \rho_1)[i_t^* + c_1 E_{t-1}(\pi_t - \pi^*) + c_2 E_{t-1}x_t] + \rho_1 i_{t-1} \quad (10)$$

In this case, the monetary policy interest rate responds linearly to the expected rate of inflation and to the expected output gap at t .⁶ By comparing Eqs (8) and (10), we note that the presence of monetary authority's asymmetric goals directly implies a nonlinear interest rate reaction function. Thus, a way to verify the hypothesis of symmetric preferences is by testing the functional form of the monetary authority's reaction function.

2.3. A model for an open economy

A limitation of the model introduced in the previous section is that it does not take into account the probable influences of the exchange rate on inflation rate and on the output gap, in addition to the effects caused by interest rate fluctuations on the exchange rate. This, however, differs from the current behavior of the Central Bank of Brazil, which considers the exchange rate to be an important tool for monetary policy transmission (Bogdanski et al., 2000; Muinhos, 2002).

Because of that, we follow Walsh (2003) and Galí (2008) and extend the basic theoretical model adding the exchange rate to the curves that represent the economic structure. As a result, the IS and Phillips curves with the exchange rate are given by:

$$x_t = -\varphi(i_t - E_t\pi_{t+1}) + E_t\pi_{t+1} + \varphi_1 s_t + e_t \quad (11)$$

$$\pi_t = kx_t + \theta_1 E_t\pi_{t+1} - \theta_1 \vartheta E_t(\Delta s_{t+1}) + \vartheta \Delta s_t + u_t \quad (12)$$

where s_t is the natural log of the real exchange rate; $\vartheta = \psi/(1 - \psi)$ (with $0 \leq \psi \leq 1$); and $1 - \psi$ is the share of domestic prices in the Brazilian consumer price index.

For determination of the expected exchange rate depreciation in (12), $E_t(\Delta s_{t+1})$, we use the uncovered interest rate parity (UIP) expressed, in nominal terms, by:

$$q_t = E_t(q_{t+1}) - (i_t - i_t^f) + \varepsilon_{1t} \quad (13)$$

where q_t is the nominal exchange rate, i_t^f is the interest rate in the foreign economy and ε_{1t} is the error term. From the definition of the real exchange rate, we have:

$$q_t = s_t - p_t^f + p_t \quad (14)$$

where p^f and p are the foreign and domestic price levels, respectively. Substituting (14) into (13) and isolating the real exchange rate, we have:

$$E_t(\Delta s_{t+1}) = E_t(\pi_{t+1}^f) - E_t(\pi_{t+1}) + i_t - i_t^f - \varepsilon_{1t} \quad (15)$$

where the term $E_t(\Delta s_{t+1})$ corresponds to the expected real exchange rate fluctuation at $t+1$, $E_t(\pi_{t+1}^f) - E_t(p_{t+1}^f - p_t^f)$ refers to the expected foreign inflation at $t+1$ and $E_t(\pi_{t+1}) = E_t(p_{t+1} - p_t)$ is the expected domestic inflation at $t+1$. Supposing that the foreign inflation rate at t and the expected foreign inflation at $t+1$ are equal to zero, we can write Eq. (15) as follows:⁷

$$E_t(\Delta s_{t+1}) = i_t - E_t\pi_{t+1} = \varepsilon_{1t} \quad (16)$$

⁶ This type of interest rate rule has been analyzed by Rudebusch (2002) and Clarida et al. (2000).

⁷ We follow some studies in the literature and adjust external (exogenous) variables to zero (see, for instance, Bonono and Brito, 2002; Leitemo and Söderström, 2008). Assuming that these external variables can vary over time would not affect the policy rule specification to be estimated because when they enter the model, they affect the exchange rate that is not explicitly observed in the policy rule. However, this does not mean that external shocks would not matter for monetary policy conduct. For example, in response to a positive shock to the foreign interest rate, the domestic interest rate should increase to keep the real exchange rate unchanged, thus avoiding the effects on inflation and on the output gap of the domestic economy. This result is similar to that obtained by Divino (2009), who analyzed the optimal monetary policy in a small open economy when the Central Bank's loss function is quadratic.

Substituting (16) into (12), the Phillips curve is given by:

$$\pi_t = kx_t + (1 + \vartheta)\theta_1 E_1 \pi_{t+1} + \vartheta \Delta s_t - \theta_1 \vartheta i_t + \eta_t \quad (17)$$

where $\eta_t = u_t + \theta_1 \vartheta \varepsilon_{1t}$ is an error term.

Considering the inclusion of the exchange rate in the IS and Phillips curves and minimizing the loss function given by Eq. (6), subject to expressions (11) and (17), we have the following specification for the Central Bank's reaction function:

$$i_t = (1 - \rho_1) \left\{ i^* + c_1 E_{t-1} \left[\frac{e^{\alpha(\pi_t - \pi^*)} - 1}{\alpha} \right] + c_2 E_{t-1} \left[\frac{e^{\gamma x_t} - 1}{\gamma} \right] \right\} + \rho_1 i_{t-1} \quad (18)$$

where

$$c_1 = \frac{k\varphi + \theta_1 \vartheta}{\mu_i}; \quad c_2 = \frac{\lambda\varphi}{\mu_i}; \quad \rho_1 = \frac{\mu_{\Delta i}}{\mu_i + \mu_{\Delta i}}. \quad (19)$$

By deriving specification (18), we note a difference in parameter c_1 in (19) compared to (9), which indicates the presence of positive terms that show, albeit not explicitly, the effect of the exchange rate on inflation. Specifically, it is possible to observe that with the addition of parameters θ_1 and ϑ to c_1 , the interest rate reacts more strongly to deviation of inflation from its target.

3. Empirical model

3.1. Reduced form of the Central's Bank reaction function

In this section, we derive the reduced form for the interest rate rule to be estimated in order to test the existence of asymmetries in the Central Bank of Brazil's reaction function in the inflation targeting regime. As pointed out by Surico (2007a), the procedures used to estimate the model and to test the null hypothesis of symmetric preferences ($H_0: \gamma = \alpha = 0$) are complex due to the lack of determination of important parameters and to the presence of nuisance parameters that could not be identified under the null. For instance, if $\gamma = \alpha = 0$, then the coefficients related to the inflation rate and to the output gap in reaction function (18) are indeterminate. In addition, when $\alpha = 0$, the inflation target is an unidentified nuisance parameter, implying that the conventional statistical theory is not available for obtaining the asymptotic distribution of the statistical tests under the null hypothesis (Luukkonen et al., 1988; van Dijk et al., 2002).

To solve these problems, we followed the suggestion made by Luukkonen et al. (1988) which consists in linearizing the exponential terms in (18) by means of a second-order Taylor expansion around $\pi_t - \pi^* = 0$ and $x_t = 0$. By doing that, we have the following monetary policy rule specification:

$$i_t = (1 - \rho_1)[c_0 + c_1 E_{t-1}(\pi_t - \pi^*) + c_2 E_{t-1}x_t + c_3 E_{t-1}(\pi_t - \pi^*)^2 + c_4 E_{t-1}x_t^2] + \rho_1 i_{t-1} + \frac{\zeta_t}{\mu_i} \quad (20)$$

where

$$c_0 = i^*; \quad c_1 = \frac{k\varphi + \theta_1 \vartheta}{\mu_i}; \quad c_2 = \frac{\lambda\varphi}{\mu_i}; \quad c_3 = \frac{\alpha(k\varphi + \theta_1 \vartheta)}{2\mu_i}; \quad c_4 = \frac{\lambda\varphi\gamma}{2\mu_i}; \quad \rho_1 = \frac{\mu_{\Delta i}}{\mu_i + \mu_{\Delta i}} \quad (21)$$

ζ_t/μ_i is the rest of the Taylor expansion.

To get the final reaction function specification estimated in this paper, we made two changes to Eq. (20). First, we introduced the nominal interest rate at $t - 2$ to avoid possible problems with serial autocorrelation.⁸ After that, we replaced the expected values with the realized ones for inflation and for the output gap. Having made these changes to specification (20), we obtained the following reaction function:

$$i_t = (1 - \rho_1 - \rho_2)[c_0 + c_1(\pi_t - \pi^*) + c_2x_t + c_3(\pi_t - \pi^*)^2 + c_4x_t^2] + \rho_1 i_{t-1} + \rho_2 i_{t-2} + v_t \quad (22)$$

⁸ This procedure was also adopted by Barcellos Neto and Portugal (2007), Lima et al. (2007), Aragón and Portugal (2010) and Minella and Souza-Sobrinho (2013). Theoretically, if we want to obtain a loss function that yields an optimal rule with term i_{t-2} , then we should add a concern of the Central Bank with interest rate deviations at t compared to the interest rate at $t - 2$.

where $v_t \equiv -(1 - \rho_1 - \rho_2)$
 $\left\{ c_1(\pi_t - E_{t-1}\pi_t) + c_2(x_t - E_{t-1}x_t) + c_3[(\pi_t - \pi^*)^2 - E_{t-1}(\pi_t - \pi^*)^2] + c_4[x_t^2 - E_{t-1}(x_t)^2] \right\} + \frac{\xi_t}{\mu}$ is the error term.

Two important characteristics of reaction function (22) can be highlighted. The first one of them is that the hypothesis of symmetry in the monetary authority's goals can be tested from the estimation of the coefficients c_i 's. It is easy to note that the imposition of constraint $\gamma = \alpha = 0$ is equivalent to $c_3 = c_4 = 0$. So, testing the null hypothesis of the Central Bank's symmetric preferences, $H_0: \gamma = \alpha = 0$, is the same as testing the null hypothesis of linearity, i.e., $H'_0: c_3 = c_4 = 0$.⁹ The statistical significance of the constraints imposed by H'_0 is checked by the Wald test. Under H'_0 , the Wald test statistic has almost a χ^2 distribution with r degrees of freedom, where r is the number of imposed constraints. The second characteristic is that the reduced form of the monetary policy rule allows obtaining estimates for the asymmetry parameters of the loss function, given that $\alpha = 2c_3/c_1$ and $\gamma = 2c_4/c_2$.

In addition to specification (22), we estimated five alternative specifications in order to make the empirical model more suitable to the Brazilian monetary policy conduct in the current inflation targeting regime. First of all, we considered a contradiction of the original assumption that the inflation target is constant. This change is necessary because the inflation targets set by the Brazilian Monetary Council in the 1999–2004 period changed on a yearly basis. That being said, the specification with a time-varying inflation target is given by:¹⁰

$$i_t = (1 - \rho_1 - \rho_2)[c_0 + c_1(\pi_t - \pi_t^*) + c_2x_t + c_3(\pi_t - \pi_t^*)^2 + c_4x_t^2] + \rho_1i_{t-1} + \rho_2i_{t-2} + v_t \quad (23)$$

In the second alternative specification, we considered that the Central Bank reacts to deviations of the expected inflation from the inflation target. Since, in the inflation targeting regime adopted in Brazil, the inflation target at T and $T+1$ are known by the policymaker at T , it is plausible to assume that monetary policy actions are taken based on the deviation of the expected inflation from the target for the current and subsequent years. Therefore, we followed the suggestion made by Minella (2003) and used variable Dj_t , which is a weighted average of the deviation of the expected inflation for years T and $T+1$ from the respective inflation targets, i.e.,:

$$Dj_t = \frac{(12-j)}{12}(E_j\pi_T - \pi_T^*) + \frac{j}{12}(E_j\pi_{T+1} - \pi_{T+1}^*) \quad (24)$$

where j is a monthly index, $E_j\pi_T$ is the inflation expectation in month j for year T , $E_j\pi_{T+1}$ is the inflation expectation in month j for year $T+1$, π_T^* is the inflation target for year T and π_{T+1}^* is the inflation target for $T+1$. The nonlinear reaction function with the variable Dj_t is expressed by:¹¹

$$i_t = (1 - \rho_1 - \rho_2)(c_0 + c_1Dj_t + c_2x_t + c_3Dj_t^2 + c_4x_t^2) + \rho_1i_{t-1} + \rho_2i_{t-2} + v_t \quad (25)$$

Finally, we considered nonlinear reaction functions in which the inflation rate reacts to the output gap at $t-2$ and to the deviation of inflation from its target at $t-1$. This assumption is justified by the fact that monthly data about inflation and economic activity are only available to the monetary authority with a lag of 1 and 2 periods, respectively. Hence, we estimated the following specifications:¹²

$$i_t = (1 - \rho_1 - \rho_2)[c_0 + c_1(\pi_{t-1} - \pi^*) + c_2x_{t-2} + c_3(\pi_{t-1} - \pi^*)^2 + c_4x_{t-2}^2] + \rho_1i_{t-1} + \rho_2i_{t-2} + v_t \quad (26)$$

$$i_t = (1 - \rho_1 - \rho_2)[c_0 + c_1(\pi_{t-1} - \pi_{t-1}^*) + c_2x_{t-2} + c_3(\pi_{t-1} - \pi_{t-1}^*)^2 + c_4x_{t-2}^2] + \rho_1i_{t-1} + \rho_2i_{t-2} + v_t \quad (27)$$

$$i_t = (1 - \rho_1 - \rho_2)(c_0 + c_1Dj_t + c_2x_{t-2} + c_3Dj_t^2 + c_4x_{t-2}^2) + \rho_1i_{t-1} + \rho_2i_{t-2} + v_t \quad (28)$$

⁹ The power of the test based on reaction function (22) depends on the finding that c_1 and c_2 are statistically different from zero because it is possible to not reject the null hypothesis of linearity when these coefficients are equal to zero.

¹⁰ This specification can be obtained if we solve the Central Bank's minimization problem with a time-varying inflation target. As in the Brazilian inflation targeting regime this target is explicit and known by the monetary authority and by private agents beforehand, the solution to the problem remains the same.

¹¹ This specification can be regarded as a forward-looking variant of monetary rule (20). Forward-looking variants of the optimal monetary rule were also investigated by Surico (2007b) and Ikeda (2010).

¹² These specifications are backward-looking variants of specifications (22), (23) and (25), respectively. Although they are not explicitly derived from the monetary authority's minimization problem, these specifications take into account the criticism that the rules under which the policy instrument responds to expected and/or current inflation and output values are not operational (see, for example, McCallum, 1999). In addition, Rotemberg and Woodford (1999) show that backward-looking monetary rules are good approximations to an optimal monetary policy.

Table 1
Unit root tests.

Variable	Exogenous regressors	ADF (k)	ERS (k)	MZ $_{\alpha}^{\text{GLS}}$ (k)
i_t	c, t	-2.26 ^{n.s} (9)	11.4 ^{n.s} (9)	-7.27 ^{n.s} (9)
x_t	c	-3.36 ^{**} (0)	1.92 [*] (0)	-14.7 [*] (0)
x_t^2	c	-3.03 ^{**} (2)	1.44 [*] (2)	-16.9 [*] (2)
$\pi - \pi^*$	c	-1.10 ^{n.s} (12)	11.8 ^{n.s} (12)	-1.98 ^{n.s} (12)
$(\pi - \pi^*)^2$	c	-2.18 ^{n.s} (8)	1.86 [*] (8)	-13.2 ^{**} (8)
$\pi - \pi_t^*$	c, t	-2.13 ^{n.s} (12)	14.7 ^{n.s} (12)	-4.67 ^{n.s} (12)
$(\pi - \pi_t^*)^2$	c	-2.42 ^{n.s} (5)	1.93 [*] (5)	-12.0 ^{**} (5)
Dj_t	c	-2.19 ^{n.s} (2)	2.64 ^{**} (2)	-9.27 ^{**} (2)
Dj_t^2	c	-3.02 ^{**} (4)	1.37 [*] (4)	-17.4 [*] (4)

Note: ^{n.s}Not significant.

* Significant at 1%.

** Significant at 5%.

4. Results

4.1. Data description

For the estimations of the Central Bank's nonlinear reaction functions shown in Section 3, we used the monthly series for the period between January 2000 and December 2010. The series were obtained from the websites of the Brazilian Institute of Applied Economic Research (IPEA) and of the Central Bank of Brazil. The dependent variable, i_t , is the annualized Selic interest rate accumulated on a monthly basis. This variable has been used as the major monetary policy instrument under the inflation targeting regime.

The inflation rate, π_t , stands for the inflation accumulated over the past 12 months, measured by the broad consumer price index (IPCA).¹³ For the specification that includes the deviation of inflation from a constant target, we used the average of annual inflation targets.¹⁴ In the case where inflation targets are time-varying, we interpolated the annual targets to obtain the series with a monthly frequency.

The variable Dj_t in specifications (25) and (28) is constructed from the inflation targets established for years T and $T + 1$, and from the inflation expectations series obtained from the survey carried out by the Central Bank with financial institutions and consultancy firms. In that survey, firms have to state what inflation rate they expect for years T ($E_j\pi_T$) and $T + 1$ ($E_j\pi_{T+1}$).

The output gap (x_t) is measured by the percentage difference between the seasonally adjusted industrial production index (y_t) and the potential output (yp_t), i.e., $x_t = 100(y_t - yp_t)/yp_t$. Here, an important problem arises because the potential output is an unobservable variable and, therefore, it has to be estimated. Bearing that in mind, we obtained the proxy variable for the potential output using the Hodrick–Prescott (HP) filter.

Before the estimations, we tested whether the variables described above are stationary. Initially, the order of integration of the variables was investigated by three tests, namely: ADF (Augmented Dickey–Fuller); ERS, proposed by Elliott et al. (1996); and the MZ $_{\alpha}^{\text{GLS}}$ test suggested by Perron and Ng (1996) and Ng and Perron (2001). The null hypothesis of the tests is that the series is nonstationary (or unit root). As demonstrated by Ng and Perron (2001), the choice of the number of lags (k) was based on the Modified Akaike Information Criterion (MAIC) considering a maximum number of lags of $k_{\max} = \text{int}(12(T/100)^{1/4}) = 12$. The tests are run with the inclusion of a constant term (c) as deterministic component, but a linear trend (t) is added if it is statistically significant.

The unit root tests are shown in Table 1. In general, results show that we can reject the unit root hypothesis in the output gap and squared output gap series, as well as in the squared deviation of inflation from its (constant and variable) target, and of Dj_t and Dj_t^2 . For the Selic rate, the deviation of inflation from a constant target and the deviation of

¹³ The IPCA is calculated by the Brazilian Institute of Geography and Statistics (IBGE) and the price index used by the Brazilian Monetary Council is used as reference for the inflation targeting regime.

¹⁴ In all years, except for 2003, we used central inflation targets determined by the Brazilian Monetary Council. In 2003, we used the target adjusted by the Central Bank of Brazil (8.5%).

Table 2
Estimates of reaction functions (22), (23) and (25).

Parameters	Specifications		
	(22)	(23)	(25)
c_0	12.46* (1.58)	10.77* (1.67)	8.478* (2.24)
c_1	4.171* (1.38)	3.920** (1.56)	9.438** (0.05)
c_2	0.731 ^{n.s.} (0.59)	1.761** (0.83)	1.482*** (0.86)
c_3	-0.318** (0.14)	-0.168 ^{n.s.} (0.12)	0.885 ^{n.s.} (0.97)
c_4	0.043 ^{n.s.} (0.08)	-0.204** (0.08)	-0.024 ^{n.s.} (0.04)
ρ_1	1.580* (0.07)	1.290* (0.14)	1.404* (0.05)
ρ_2	-0.597* (0.06)	-0.316** (0.13)	-0.424* (0.05)
$W(2) - prob$	0.075	0.049	0.390
$J(12) - prob$	0.541	0.331	0.254

Note: ^{n.s.} Not significant. The standard errors are in brackets.

* Significant at 1%.

** Significant at 5%.

*** Significant at 10%.

inflation from a variable target, test results show that we cannot reject the hypothesis that these variables have a unit root.

Since the non-rejection of the unit root null hypothesis in series i_t , $\pi_t - \pi^*$ and $\pi_t - \pi^*_t$ may result from the structural break in the trend function,¹⁵ we ran two unit root tests with structural break. Following Carrion-i-Silvestre et al. (2009), the $MZ^{GLS}_\alpha(\lambda^0)$ and $MZ^{GLS}_t(\lambda^0)$ statistics were used to test the unit root null hypothesis, allowing for structural breaks in the level and slope of the trend function at unknown dates under the null and alternative hypotheses.¹⁶ The values obtained for $MZ^{GLS}_\alpha(\lambda^0)$ and $MZ^{GLS}_t(\lambda^0)$ were, respectively, -366.5 and -13.5 for the Selic rate, -175.9 and -9.4 for the deviation of inflation from a constant target, and -32883.0 and -128.22 for the deviation of inflation from a variable target. These calculated values allow rejecting the null hypothesis that these series have a unit root at a 1% significance level.

4.2. Estimated reaction functions

First, we estimated reaction functions (22), (23) and (25) using the generalized method of moments (GMM) with the optimal weighting matrix that takes into account possible heteroskedasticity and serial autocorrelation in the residuals (Hansen, 1982).¹⁷ In practice, we used the method proposed by Newey and West (1987) with three lags to estimate the variance-covariance matrix. The set of instruments includes a constant term, lags (-3) and (-4) of the Selic rate, and four lags - of output gap, of the deviation of inflation from its target, of the squared output gap, and of the squared deviation of inflation from its target. These instruments imply 12 over-identification restraints. We tested the validity of these restraints using the Hansen's (1982) J test.

The estimation results are displayed in Table 2 (the standard errors are in brackets). Specifications (22), (23) and (25) refer, respectively, to the specifications with the fixed inflation target, with the variable target, and with the deviation of expected inflation from the inflation target. Initially, we observed that, for specification (22), the estimate of parameter c_3 , which measures the response of the Selic rate to the squared deviation of current inflation from the target, had a negative sign and was statistically significant. It is important to underline that a negative coefficient on $\pi_t - \pi^*$ indicates that the reduction of the Selic rate in response to a decrease in the inflation rate relative to the target of a given magnitude is larger than the increase of this interest rate caused by an elevation in the deviation of inflation of the same magnitude. When we took into account the deviation of the current inflation or of the expected inflation from a time-varying

¹⁵ See, for instance, Perron (1989).

¹⁶ The tests were run with up to five structural breaks at unknown dates.

¹⁷ Here, we follow Clarida et al. (1998, 2000), Surico (2007a,b) and Ikeda (2010) among others, and use the GMM because of the endogenous regressors in the monetary authority's reaction function. As advertised by an anonymous referee, it is necessary to be cautious about the interpretation of the standard errors as they are based on the asymptotic theory, whereas the number of observations in this paper (132) is small.

Table 3
Estimates of asymmetric preferences.

Parameters	Specifications		
	(22)	(23)	(25)
α	-0.153* (0.03)	-0.086** (0.04)	0.188 ^{n.s.} (0.22)
γ	0.118 ^{n.s.} (0.18)	-0.231** (0.11)	-0.033 ^{n.s.} (0.07)

Note: ^{n.s.} Not significant. The standard errors are in brackets.

* Significant at 1%.

** Significant at 5%.

inflation target, we could not find any evidence of nonlinearity in the response of the monetary policy instrument to these variables.

Due to the nonlinear structure of specification (22), the response of the monetary policy instrument to deviations of the current inflation from the inflation target is given by:

$$\frac{\partial i}{\partial(\pi - \pi^*)} = c_1 + 2c_3 E(\pi - \pi^*) \quad (29)$$

where $E(\cdot)$ indicates the sample mean. Using this expression and the coefficient values shown in Table 2, we estimated that the response of the Selic rate to a deviation of inflation from its target was equal to 3.60. This indicates that the nonlinear interest rate rule satisfies the Taylor (1993) principle. For specifications (23) and (25), where no evidence of nonlinearity was found, the responses of the interest rates to the variables $\pi_t - \pi_t^*$ and Dj_t were 3.92 and 9.44, respectively. The stronger reaction of the monetary policy to the expected inflation is in line with the results obtained by Holland (2005), Soares and Barbosa (2006) and Aragón and Portugal, 2010, and demonstrates the forward-looking nature of Central Bank decisions.

By and large, the reaction of the interest rate to output gap, measured by parameter c_2 , was significant at 10%. The coefficient on the squared output gap, c_4 , was not statistically different from zero in two out of the three estimated models. This means that there is no compelling empirical evidence in favor of a nonlinear response of the monetary policy instrument to output gap.

The last two rows in Table 2 show the p -values (*prob*) for the joint hypothesis of symmetric preferences and for the hypothesis of validity of over-identification restraints. Only in the case of specification (25), the hypothesis of a linear reaction function is not rejected at a 10% significance level. The J test results indicate that the over-identification restraints cannot be rejected at a 10% significance level.

Table 3 displays the estimates for the monetary authority's asymmetric preferences. The coefficients were obtained from expressions $\alpha = 2c_3/c_1$ and $\gamma = 2c_4/c_2$. The standard errors were calculated using the delta method (Greene, 2000). We can observe that, in general, the coefficients that measure asymmetric preferences on output gap, γ , were not statistically different from zero. Conversely, the values of the coefficient of asymmetric preferences on inflation, α , had a negative sign and were statistically negative for specifications (22) and (23). This suggests that negative deviations of inflation from the target of a given magnitude cause a larger loss to the Brazilian monetary authority than the positive deviations with the same magnitude.

In Table 4, we show the estimates of reaction functions (26)–(28), in which the monetary policy instrument depends on the deviation of inflation from the target at $t - 1$ and on output gap at $t - 2$. Due to the presence of outliers, two dummy variables, $D02$ (=1 for 2002:10 and 0, otherwise) and $D03$ (=1 for 2003:09 and 0, otherwise), were added to all specifications.¹⁸ Initially, we estimated the monetary policy rules by ordinary least squares (OLS). As the ARCH test indicated serious problems with conditional autoregressive heteroskedasticity, we estimated the reaction functions assuming that the conditional variance of the error terms follows an ARMA(p , q) process, where p is the order of the autoregressive part and q is the order of moving averages. The last row of Table 4 shows the orders p and q of the GARCH models estimated by maximum likelihood (ML).

¹⁸ The outliers observations were detected by using the criterion $|\hat{\epsilon}_t| > 3\hat{\sigma}$.

Table 4
Estimates of reaction functions (26)–(28).

Parameters	Specifications		
	(26)	(27)	(28)
c_0	13.24* (0.90)	12.26* (1.83)	11.67* (0.88)
c_1	2.464* (0.56)	2.402** (0.13)	5.618* (1.20)
c_2	0.830*** (0.47)	2.078** (1.05)	0.827** (0.37)
c_3	–0.158* (0.06)	–0.175 ^{n.s.} (0.13)	0.250 ^{n.s.} (0.19)
c_4	–0.035 ^{n.s.} (0.03)	–0.022 ^{n.s.} (0.05)	–0.019 ^{n.s.} (0.03)
ρ_1	1.631* (0.04)	1.660* (0.04)	1.556* (0.05)
ρ_2	–0.658* (0.04)	–0.673* (0.04)	–0.583* (0.05)
$D02$	72.19* (21.18)	162.5** (76.61)	48.06* (13.2)
$D03$	–41.33* (34.60)	–94.72** (43.07)	–39.82 ^{n.s.} (65.29)
α	–0.128* (0.02)	–0.146** (0.06)	0.089 ^{n.s.} (0.08)
γ	0.086 ^{n.s.} (0.11)	–0.021 ^{n.s.} (0.06)	–0.047 ^{n.s.} (0.09)
R^2 – adjusted	0.996	0.996	0.997
$W(2)$ – prob	0.008	0.383	0.287
$LB(4)$ – prob	0.823	0.810	0.977
$ARCH(4)$ – prob	0.971	0.897	0.632
JB – prob	0.651	0.642	0.840
$GARCH(p,q)$	0.1	0.1	1.1

Notes: The standard errors are in brackets. LB(4) refers to the Ljung-Box statistic for serial autocorrelation up to the fourth order. ARCH(4) refers to the LM-ARCH statistic for conditional autoregressive heteroskedasticity up to the fourth order. JB refers to the Jarque-Bera statistic.

* Significant at 1%.

** Significant at 5%.

*** Significant at 10%.

^{n.s.} Not significant.

The results are similar to those shown in Tables 2 and 3. The estimates of the coefficient of asymmetric preferences on output gap, γ , are not statistically different from zero, whereas the coefficient that measures asymmetric preferences on deviations of inflation from the target, α , is negative and significant in two out of the three specifications analyzed. This is consistent with Arag3n and Portugal (2010) and suggests that the Central Bank of Brazil has been more averse to a below-target than to an above-target inflation rate. It is important to underscore that this behavior is unexpected from a monetary authority that is more concerned with lending credibility to its disinflation policy.

4.3. Structural break tests for the estimated reaction functions

The concavity of the monetary rule related to inflation can result from policy decisions made at times of supply shocks and of fiscal dominance.¹⁹ To assess that, Arag3n and Portugal (2010) estimated reaction functions for a more stable period (2004–2007), but they did not find any evidence of asymmetric preferences by the Central Bank. This suggests the existence of a possible structural break in the reaction function parameters and, consequently, in the coefficients of asymmetric preferences by the monetary authority.²⁰ However, one should stress that Arag3n and Portugal (2010) choose the regime shift period and do not use any statistical test for structural break.

In this paper, we performed the structural break test proposed by Bai and Perron (1998, 2003) to investigate the existence of possible changes in the monetary rule and in the coefficients of asymmetric preferences by the Central Bank. This econometric procedure allows estimating the number and the time of possible structural changes in the monetary rule by means of sequential break tests with unknown dates.

¹⁹ The 2001 energy crisis can be regarded as the main supply shock after 2000. As to fiscal dominance, Blanchard (2004) provides empirical evidence in favor of this phenomenon for the Brazilian economy during the last quarter of 2002. On any of these occasions, the Central Bank may have adopted a more gradualistic behavior towards inflation than that which would be expected from a policymaker with asymmetric preference for a below-target inflation.

²⁰ Other pieces of evidence of changes in the Central Bank's reaction function parameters are provided by Barcellos Neto and Portugal (2007) and Lima et al. (2007).

Table 5
Estimates of reaction functions (26)–(28).

Parameters	(26)		(27)		(28)	
	2000:01–2003:07	2003:08–2010:12	2000:01–2005:07	2005:08–2010:12	2000:01–2003:09	2003:10–2010:12
c_0	11.85* (2.22)	12.58* (0.84)	16.99* (0.36)	9.964* (0.93)	14.14* (1.10)	8.690* (0.93)
c_1	3.208** (1.18)	2.705* (0.66)	0.547* (0.19)	0.574 ^{n.s.} (0.69)	5.442* (1.27)	2.244** (1.43)
c_2	0.208 ^{n.s.} (0.26)	0.745*** (0.40)	0.118 ^{n.s.} (0.20)	1.127* (0.38)	0.485 ^{n.s.} (0.35)	0.732* (0.26)
c_3	-0.182** (0.09)	-0.342** (0.15)	0.012 ^{n.s.} (0.02)	-0.063 ^{n.s.} (0.68)	-0.570* (0.19)	6.862* (1.59)
c_4	0.043 ^{n.s.} (0.06)	-0.051** (0.03)	-0.020 ^{n.s.} (0.04)	0.004 ^{n.s.} (0.02)	0.004 ^{n.s.} (0.06)	-0.016 ^{n.s.} (0.03)
ρ_1	1.489* (0.14)	1.610* (0.05)	1.611* (0.08)	1.368* (0.09)	1.419* (0.13)	1.422* (0.04)
ρ_2	-0.643* (0.12)	-0.632* (0.05)	-0.803* (0.08)	-0.402* (0.09)	-0.560* (0.13)	-0.451* (0.03)
α	-0.113* (0.01)	-0.253* (0.08)	0.045 ^{n.s.} (0.08)	-0.220 ^{n.s.} (2.37)	-0.210* (0.03)	6.117*** (5.06)
γ	0.413 ^{n.s.} (0.55)	-0.138 ^{n.s.} (0.13)	-0.330 ^{n.s.} (0.79)	0.008 ^{n.s.} (0.04)	-0.017 ^{n.s.} (0.27)	-0.043 ^{n.s.} (0.09)
$R^2 - adjusted$	0.986	0.996	0.984	0.997	0.986	0.998
$W(2) - prob$	0.113	0.025	0.690	0.981	0.013	0.000
$LB(4) - prob$	0.495	0.138	0.498	0.267	0.212	0.234
$ARCH(4) - prob$	0.554	0.707	0.181	0.586	0.918	0.378
$JB - prob$	0.000	0.706	0.004	0.483	0.254	0.954
$GARCH(p,q)$	-	1.2	-	-	-	1.1
$SupF_T(1)$	38.82*	38.03*	39.68*	-	-	-
T_j	03:07 (03:05–03:12)	05:07 (05:06–06:06)	03:09 (03:08–03:10)	-	-	-

Notes: The standard errors are in brackets. LB(4) stands for the Ljung-Box statistic for serial autocorrelation up to the fourth order. ARCH(4) represents the LM-ARCH statistic for autoregressive conditional heteroskedasticity up to the fourth order. JB stands for the Jarque-Bera statistic.

* Significant at 1%.

** Significant at 5%.

*** Significant at 10%.

^{n.s.} Not Significant.

Initially, we tested the existence of a structural break for specifications (22), (23) and (25). Although the results indicate changes in the parameters, the GMM estimates were quite distinct among the analyzed specifications.²¹ An explanation to this finding lies in the fact that GMM estimators can be strongly biased and widely dispersed in small samples (Tauchen, 1986; Fuhrer et al., 1995; Andersen and Sørensen, 1996). Therefore, we ran structural break tests and analyzed their implications only for specifications (26)–(28) that can be estimated by OLS or ML, as they do not contain endogenous explanatory variables.

The results of the structural break tests and the estimates of break dates (with the 95% confidence interval in brackets) are shown in the last two rows of Table 5.²² According to Bai and Perron (2003), we used the $supF(1|0)$ test to check the null hypothesis of no break ($m = 0$) against the alternative hypothesis of $m = 1$ break. For the three reaction function specifications, the $supF(1|0)$ test rejects the null hypothesis of no break at a 1% significance level. As to the time of the break, the results obtained for specifications (26) and (28) suggest a change in the monetary rule in the third quarter of 2003, whereas for specification (27), the structural break occurred in July 2005.

The estimates for coefficients c_i 's, $i = 1, 2, 3, 4$, reveal important changes in the monetary policy conduct. Except for specification (27), the values of c_1 and c_3 show that the Central Bank began to react more strongly to deviations of (current and expected) inflation from its target after 2003. Using expression (29), we found that the responses of the Selic rate to the variables $\pi_t - \pi^*$ and Dj_t increased from 1.79 to 2.19, and from 3.93 to 5.95, respectively. Moreover, the estimates obtained for c_2 indicate that the monetary policy was more sensitive to output gap movements after the structural break.

With respect to the coefficient of asymmetric preference on the stabilization of inflation, α , the evidence obtained for the specification with constant inflation target indicates an increase in asymmetric preference in favor of an above-target inflation, whereas it is not possible to reject the null hypothesis of symmetric preferences for the specification with a variable target. For the specification with the variable Dj_t , commonly used in the Brazilian literature on monetary rule,

²¹ The results can be obtained from the authors upon request.

²² Due to the sample size, we set the maximum number of breaks as 1 and used a 15% trimming. For further details, see Bai and Perron (2003).

the results show that the Central Bank of Brazil was more averse to a below-target inflation until mid-2003, but more averse to an above-target inflation thereafter. This finding suggests that the Central Bank asymmetric preferences in favor of an above-target inflation in the first subperiod may be related to monetary policy actions taken in periods in which domestic crises and external shocks strongly affected inflation and inflation expectations.

Finally, we observed that the estimates for parameter γ were not statistically different from zero in any of the specifications. This is in line with Aragón and Portugal (2010) and indicates that the Brazilian monetary authority has not had asymmetric preference for an output above or below the potential output.

5. Conclusion

In this paper, we estimated the nonlinear reaction function for the Central Bank of Brazil in order to check for asymmetries in its goals related to output gap and deviations of inflation from its target. To achieve that, we extended the theoretical model of Surico (2007a) for an open economy and we obtained an optimal monetary policy rule for the monetary authority considering that its loss function is potentially asymmetric. Given that the presence of asymmetries yields nonlinear responses of the interest rate to inflation and to the output gap, we investigated whether the policymaker's preferences are symmetric by testing the null hypothesis of linearity of the reaction function. In addition, we estimated the Central Bank asymmetric preference coefficients and tested whether they were statistically significant.

From an empirical standpoint, our paper followed Bai and Perron (1998, 2003) and we ran structural break tests for the nonlinear reaction function parameters. The analysis of structural changes allowed checking for possible changes in the monetary policy conduct, as well as in the Central Bank asymmetric preference coefficients.

The structural break tests strongly rejected the null hypothesis of no break. For two out of three reaction function specifications, the estimated period for the change in the monetary rule was the third quarter of 2003. In general, the estimates of reaction function coefficients revealed that the Central Bank of Brazil reacted more strongly to deviations of (current and expected) inflation from the inflation target and to the output gap after 2003.

The estimated values for the coefficient of asymmetric preference on the stabilization of inflation, α , indicated distinct changes in the Brazilian monetary authority's loss function. For the specification of the reaction function with a constant inflation target, the estimates showed an increase in asymmetry in favor of an above-inflation inflation after the date of the structural change. When we considered the deviation of inflation from a variable target, the asymmetry coefficient for the inflation was not significant in any of the analyzed subperiods. However, in the specification with the Dj_t variable, commonly used in the Brazilian literature on monetary rule, the results demonstrated that the Central Bank was more averse to a below-target inflation up to mid-2003, but more averse to an above-target inflation thereafter.

Finally, the results show that the estimates for parameter γ were not statistically different from zero in any of the specifications. This indicates that the Central Bank did not have an asymmetric preference for an output above or below the potential output.

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