Changes in the Brazilian monetary policy: Evidence of a reaction function with time-varying parameters and endogenous regressors

Abstract: This paper estimates a forward-looking reaction function with time-varying parameters to examine changes in the Brazilian monetary policy under the inflation-targeting regime. As the monetary policy rule has endogenous regressors, the conventional Kalman filter cannot be applied. Thus, a Heckman-type (1976) two-step procedure is used for consistent estimation of the hyper-parameters of the model. The results show that: i) there is strong empirical evidence of endogeneity in the regressors of the policy rule; ii) the response of the Selic rate to inflation varies considerably over time and has shown a decreasing trend; iii) since mid-2010, policy rule has violated the Taylor principle; iv) the implicit target for the Selic rate has shown a decline over time; v) the degree of interest rate smoothing has shown a relative stability.

Keywords: forward-looking monetary policy rule; time-varying parameter model; Brazil.

JEL Classification: C32, E52, C50.

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1 Introduction

In the last 20 years, several papers have estimated different specifications of the reaction function in order to study the decisions of the central banks regarding the interest rate of the monetary policy. A quite well known specification is Taylor (1993) rule, given by:

$$i_t = i^* + 1,5\left(\pi_t - \pi^*\right) + 0,5y_t$$

According to that rule, the central bank increases the nominal interest rate, i_t, as a response to inflation deviations from the target, π_t - π^*, and the output gap, y_t.

Another specification which has received considerable attention is the *forward-looking* reaction function proposed by Clarida et al. (1998, 2000):

$$i_t = (1-\rho)\left[\alpha + \beta E_t(\pi_{t+n}) + \delta E_t(Y_{t+n})\right] + \rho i_{t-1} + \nu_t$$

In this policy rule, the policymaker adjusts the current interest rate based upon the future values expected for inflation (π_{t+n}) and output gap (y_{t+n}).

Several works in the Brazilian literature seek to estimate reaction functions for the monetary policy.[1] Within that literature, some papers have highlighted important coefficient variations in the monetary policy rules. For example, Salgado et al. (2005) note the different dynamics of the Selic interest rate, during and out of periods of exchange rate crisis. Policano and Bueno (2006) show that the responses of the Selic rate to inflation, product and exchange rate, were different between the pre and post inflation target periods. Bueno (2005) and Lima et al. (2007) estimate a Markov-switching reaction function and point out the existence of different monetary policy regimes after the Real Plan. Barcellos Neto and Portugal (2007) found empirical evidence that, during the Henrique Meirelles administration, the Selic rate responded less to the deviations of expected inflation from the target inflation and more to exchange rate variations when compared to Arminio Fraga´s administration. Medeiros and Aragon (2011) note that the Brazilian monetary authority reacted more strongly to inflation deviations regarding the target and to the output gap after 2003.

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[1] See, for example Silva and Portugal (2001), Minella et al. (2003), Holland (2005), Soares and Barbosa (2006), Teles and Brundo (2006), and Aragón and Portugal (2010).

In this paper, we seek to estimate a reaction function with time-varying parameters in order to analyze possible changes in monetary policy by the Central Bank of Brazil (BCB) during the inflation targeting regime. Since the proposed monetary policy rule presents endogenous regressors, the conventional Kalman filter leads to invalid inferences regarding the model; therefore it must not be applied. Due to this, we follow Kim (2006) and use a two-step estimation procedure, similar to that of Heckman (1976). In this procedure, the terms of bias correction are entered in the second step. To correct possible regressor generated problems, the Augmented Kalman Filter is used. We also test the null hypothesis of non-endogeneity in the reaction function of the monetary authority.

The results found, indicate that the reaction function parameters of the BCB are time-varying and that the regressors of that function, are endogenous. Besides, we observed that: i) the Selic rate responses to current inflation and the inflationary expectations, present considerable changes and have diminished with the passing of time; ii) since mid-2010, policy rule has violated the Taylor principle; iii) the implicit target for the Selic rate has shown a decline over time; iv) the degree of interest rate smoothing has shown a relative stability. Finally, the policy instrument response to the output gap, presents an increasing trend/tendency over the 2010-2011 period.

Besides this introduction, this paper consists of four sections. The second section presents the theoretical model used in the study. In section 3, we have the specification of the reduced form of the reaction function, as well as the descriptions of the two-step estimation procedure and the Augmented Kalman filter. The fourth section presents the analysis of the results. The final conclusions of the paper are in the fifth section.

2 Optimal monetary policy in a forward-looking model

In order to analyze optimal decisions of monetary policy, we follow Clarida et al. (1999) and we consider a three component model. The first one relates to the constraints of the control problem faced by the *policymaker* and consists of two equations: an IS curve, which governs the dynamics of the output; and a Phillips curve, that describes the dynamics of inflation. The second one is the loss function of the central bank, which describes the objectives of the monetary policy. The third component is the optimal monetary policy rule that shows how the central bank determines the optimum path for the nominal interest rate.

2.1 The structure of economy

In this subsection, there is a brief description of the log-linearized version of the New-Keynesian model, with sticky prices, analyzed by Clarida et al. (1999). According to this model, the evolution of an economy is represented by the following two-equation system:

$$y_t = E_t y_{t+1} - \varphi(i_t - E_t \pi_{t+1}) + u_t^d \tag{1}$$

$$\pi_t = \alpha E_t \pi_{t+1} + k y_t + u_t^s \tag{2}$$

where y_t is the output gap (this is, the difference between the actual and the potential output), π_t, is the inflation rate, $E_t y_{t+1}$ and $E_t \pi_{t+1}$ are the expected values of the output gap and inflation rate conditional on the available information at t, i_t is the nominal interest rate, u_t^d and u_t^s are, a demand shock and a cost shock respectively. The φ, k and α parameters, are positive constants.[2]

The IS curve, given by the equation (1), is a log-linearized version of Euler's equation for consumption, derived from the optimum decision of households on consumption and savings, after the imposition of the market's clearing condition. The expected value for the output gap shows that, since families prefer smoothing out consumption over time, the expected higher level of consumption leads to an increased current consumption, thereby increasing the present demand for the product.

On the other hand, the Phillips curve, given by equation (2), grasps the characteristic of overlapping nominal price, where firms have the probability α of maintaining the product price fixed, at any period of time (Calvo, 1983). Since the probability of α is supposedly constant and independent of the time elapsed since the last adjustment, the average duration in which the price

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[2] The aggregate behavioural equations (1) and (2) are explicitly derived from optimizing behaviour of firms and households, in an economy with nominal rigidities of prices and currency (Clarida et al., 1999)

remains fixed is 1/1-α. The discrete nature of price adjustment, resulting from this fact, encourages each firm to set a higher price when the higher is the expectation of future inflation.

Finally, shocks u_t^d and u_t^s comply with the following autoregressive processes:

$$u_t^d = \rho_{u^d} u_{t-1}^d + \hat{u}_t^d \tag{3}$$
$$u_t^s = \rho_{u^s} u_{t-1}^s + \hat{u}_t^s \tag{4}$$

where $0 \le \rho_{u^d}, \rho_{u^s} \le 1$, $\hat{u}_t^d$ and $\hat{u}_t^s$ are i.i.d random variables with zero mean and standard deviations σ_{μ^d} and σ_{μ^s}, respectively.

2.2 The central bank's loss function and the optimal monetary rule

Let us suppose that the monetary policy decisions are taken before the performance of the shocks u_t^d and u_t^s. Conditional on the information available at the end of the previous period, the monetary authority seeks to choose the current interest rate i_t and a sequence of future interest rates in order to minimize:

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

subjected to the structure of the economy, given by equations (1) and (2), where δ is the fixed discount factor. The loss function at time t, L_t, is given by:

$$L_t = \frac{1}{2}\left[\left(\pi_t - \pi^*\right) + \lambda y_t^2 + \mu_i \left(i_t - i^*\right)^2 + \mu_{\Delta i}\left(i_t - i_{t-1}\right)^2\right] \tag{6}$$

in which π^* is the inflation target, λ is the relative weight on the deviation of the output from the potential level, μ_i and $\mu_{\Delta i}$ are relative weights given in order to stabilize the interest rate around an implicit target, i^*, and of the interest rate at t-1, i_{t-1}.[3] The monetary authority is supposed to stabilize inflation around the inflation target, maintaining the output gap close to zero, and to stabilize the nominal interest rate around the target and that of the nominal interest rate at t-1.

[3] The interest rate smoothing is justified for several reason, such as: i) presence of uncertainties regarding the value of the data and coefficients of the macroeconomic model; ii) great changes in the interest rate might destabilize financial markets and foreign exchange; iii) constant variation in the short term interest rate, even if small, would cause great effect on aggregate demand and inflation. For a theoretical and empirical research on the interest rate smoothing of monetary policy, see Clarida et al. (1998), Sack (1998), Woodford (1999, 2003) and Sack and Wieland (2000).

To solve the optimization problem (5), it is assumed that monetary policy is discretionary.[4] This implies that the central bank takes the expectations of future variables as given and chooses the current interest rate for each period. Since there is no endogenous persistence in the inflation and output gap, the intertemporal optimization problem can be reduced to a succession of static optimization problems. Thus, taking the first order condition, we arrive at the following expression:

$$-\kappa\gamma E_{t-1}\left(\pi_t - \pi^*\right) - \lambda\gamma E_{t-1}(y_t) + \mu_i(i_t - i^*) + \mu_{\Delta i}(i_t - i_{t-1}) = 0 \qquad (7)$$

Solving into i_t, the monetary policy rule can be expressed as follows:

$$i_t = (1-\theta)\left[\beta_0 + \beta_1 E_{t-1}\left(\pi_t - \pi^*\right) + \beta_2 E_{t-1}(y_t)\right] + \theta i_{t-1} \qquad (8)$$

in which $\beta_0 = i^*$; $\beta_1 = \dfrac{\kappa\gamma}{\mu_i}$; $\beta_2 = \dfrac{\lambda\gamma}{\mu_i}$; $\theta = \dfrac{\mu_{\Delta i}}{\mu_i + \mu_{\Delta i}}$.

From equation (8) on, it can be observed that the optimal nominal interest rate for period t responds linearly to the deviations of the expected inflation from the inflation target and the output gap expected for period t. Regarding the smoothing parameter, θ, we can observe that: i) if $\mu_i > 0$ and $\mu_{\Delta i} > 0$, then $0 < \theta < 1$; ii) if $\mu_i = 0$ and $\mu_{\Delta i} > 0$, then $\theta = 1$; iii) if $\mu_{\Delta i} = 0$ and $\mu_i > 0$, then $\theta = 0$; iv) if $\mu_i = \mu_{\Delta i} = 0$, then θ will be indeterminate.

3 A monetary rule with time-varying parameters and a two-step procedure

With the aim of estimating the reduced reaction function form (8), an exogenous random shock to the interest rate, m_t, is included in that expression. It is presumed that this shock is *i.i.d* and can be interpreted as the pure random component of monetary policy. Besides, in order to capture changes in the policy conduction, it is considered that the reaction function parameters are time-varying and assume a dynamic random walk. This specification, proposed by Cooley and Prescott (1976) has been used in several papers and, is a way of considering Lucas' (1976) criticism regarding the

[4] Palma and Portugal (2011) found evidence in favour of a discretionary monetary policy in Brazil during the 2000 to 2010 period.

inadequacy of econometric models with constant parameters for policy evaluation.[5] Finally, the expected inflation and output gap values in (8) are substituted by their observed values. From these changes, we arrive at the following reaction function with time-varying parameters

$$i_t = \beta'_{0,t} + \beta'_{1,t}(\pi_t - \pi_t^*) + \beta'_{2,t} y_t + \theta_t i_{t-1} + e_t, \quad e_t \sim i.i.d.\mathrm{N}(0, \sigma_e^2) \tag{9}$$

$$\beta'_{i,t} = (1-\theta_t)\beta_{i,t}, \quad \mathrm{i} = 0,1,2 \tag{10}$$

$$\beta'_{i,t} = \beta'_{i,t-1} + \varepsilon_{i,t}, \quad \varepsilon_{i,t} \sim i.i.d.\mathrm{N}(0, \sigma_{\varepsilon,i}^2) \tag{11}$$

$$\theta_t = \theta_{t-1} + \varepsilon_{3,t}, \quad \varepsilon_{3,t} \sim i.i.d.\mathrm{N}(0, \sigma_{\varepsilon,3}^2) \tag{12}$$

in which $e_t = -\left[\beta'_{1,t}(\pi_t - E_{t-1}(\pi_t)) + \beta'_{2,t}(y_t - E_{t-1}(y_t))\right] + m_t$. Coefficients $\beta'_{1,t}$ and $\beta'_{2,t}$ ($\beta_{1,\mathrm{t}}$ and $\beta_{2,\mathrm{t}}$) measure the short-run (long-run) response and those of the Selic rate regarding inflation and the output gap.

As the forecasted errors of inflation and the output gap make up the term e_t, it is possible to observe that π_t and y_t are correlated with this error term. In that case, the estimation of (9)-(12) through the conventional Kalman filter via Maximum Likelihood cannot be performed because this procedure is derived under the assumption that the regressors and disturbances are not correlated.

In order to correct the endogeneity problem, instrumental variables will be used. In particular, the relationship between the endogenous regressors and their instruments will be given by:

$$\pi_t = z_t'\delta_1 + v_{1t}, \quad v_{1t} \sim N(0, \sigma_{v1}^2) \tag{13}$$

$$y_t = z_t'\delta_2 + v_{2t}, \quad v_{2t} \sim N(0, \sigma_{v2}^2) \tag{14}$$

in which z_t is the vector of the instruments. For simplicity's sake, it is assumed that the relationship among the endogenous regressors and their instruments, are constant.

3.1 A two-step Maximum Likelihood procedure

The two-step estimation procedure starts from the decomposition of π_t e y_t into two components: predicted components and prediction error components. Doing this, we have:

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[5] Examples of other works that suppose the parameters of the model follow a random walk are Cogley and Sargent (2001, 2005), Boivin (2006) and Kim and Nelson (2006).

$$\begin{bmatrix} \pi_t \\ y_t \end{bmatrix} = E\begin{bmatrix} \pi_t \\ y_t \end{bmatrix} | \psi_{t-1} + \begin{bmatrix} v_{1t} \\ v_{2t} \end{bmatrix} \quad (15)$$

$$\begin{bmatrix} v_{1t} \\ v_{2t} \end{bmatrix} = \Omega^{1/2}\begin{bmatrix} v_{1t}^* \\ v_{2t}^* \end{bmatrix}, \quad \begin{bmatrix} v_{1t}^* \\ v_{2t}^* \end{bmatrix} \sim i.i.d.N\left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix}\right) \quad (16)$$

where ψ_{t-1} is the available information in t-1 and Ω is the variance covariance matrix for a vector of prediction errors, $v_t = [v_{1t}\ v_{2t}]'$.

Taking a vector of 2x1 standardized prediction errors, $v_t^* = [v_{1t}^*\ v_{2t}^*]'$, we have the covariance structure between v_t^* and e_t:

$$\begin{bmatrix} v_t^* \\ e_t \end{bmatrix} \sim N\left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} I_2 & \rho\sigma_e \\ \rho'\sigma_e & \sigma_e^2 \end{bmatrix}\right) \quad (17)$$

where $\rho = [\rho_1\ \rho_2]'$is a constant correlation vector. As in Kim (2006), the Cholesky decomposition of the covariance matrix results in the following representation:

$$\begin{bmatrix} v_t^* \\ e_t \end{bmatrix} = \begin{bmatrix} I_2 & \rho\sigma_e \\ \rho'\sigma_e & \sqrt{(1-\rho'\rho)}\sigma_e \end{bmatrix}\begin{bmatrix} \varepsilon_t \\ \omega_t \end{bmatrix}, \quad \begin{bmatrix} \varepsilon_t \\ \omega_t \end{bmatrix} \sim i.i.d.N\left(\begin{bmatrix} 0_2 \\ 0 \end{bmatrix}, \begin{bmatrix} I_2 & 0_2 \\ 0_2 & 1 \end{bmatrix}\right) \quad (18)$$

where 0_2 is a 2x1 vector of zeros. From (18), then we have:

$$e_t = \rho_1\sigma_e v_{1t}^* + \rho_2\sigma_e v_{2t}^* + \omega_t^*, \quad \omega_t^* \sim N(0,(1-\rho_1^2-\rho_2^2)\sigma_e^2) \quad (19)$$

where ω_t^* is not correlated with either v_{1t}^* or v_{2t}^*. Equation (19) shows that e_t in equation (9) can be broken down into the following components: i) v_{1t}^* and v_{2t}^*, which are correlated with π_t and y_t; and ii) the ω_t^* component, that is not correlated with π_t and y_t. Substituting equation (19) in (9), results in:

$$i_t = \beta'_{0,t} + \beta'_{1,t}(\pi_t - \pi_t^*) + \beta'_{2,t}y_t + \theta_t i_{t-1} + \rho_1\sigma_e v_{1t}^* + \rho_2\sigma_e v_{2t}^* + \omega_t^* \quad (9')$$

In equation (9'), the new error term is not correlated with π_t, y_t, v_{1t}^* or v_{2t}^*.Therefore, the estimation procedure for the Maximum Likelihood (ML) is given in two steps:

Step 1: Estimate equations (13) and (14) through ML or Ordinary Least Squares (OLS) and obtain the standardized prediction errors, $\hat{v}_{1t}^*$ and $\hat{v}_{2t}^*$.

Step 2: Estimate through ML using the Kalman filter equation

$$i_t = \beta'_{0,t} + \beta'_{1,t}(\pi_t - \pi_t^*) + \beta'_{2,t} y_t + \theta_t i_{t-1} + \rho_1 \sigma_e \hat{v}_{1t}^* + \rho_2 \sigma_e \hat{v}_{2t}^* + \omega_t^* \qquad (9'')$$

together with equations (11) and (12).

As highlighted by Kim and Nelson (2006), the standardized prediction errors $\hat{v}_{1t}^*$ and $\hat{v}_{2t}^*$ are included in (9") as bias correction terms. This is similar to the two-step procedure proposed by Heckman (1976). In that case, the bias correction terms are inserted in order to capture possible changes in the degree of uncertainty associated with inflation and output gap, which are taken into account in the monetary policy rule.

3.1.1 The Augmented Kalman filter

The reaction function with time-varying parameters (9”) can be expressed as:

$$i_t = X_t' \beta_t + \rho_1 \sigma_e \hat{v}_{1t}^* + \rho_2 \sigma_e \hat{v}_{2t}^* + \omega_t^*, \quad \omega_t^* \sim N(0, (1-\rho_1^2 - \rho_2^2)\sigma_e^2) \qquad (20)$$

$$\beta_t' = \beta_{t-1}' + \varepsilon_t, \; \varepsilon_t \sim i.i.d.N(0, \Sigma_\varepsilon) \qquad (21)$$

where $X_t' = [1 \quad \tilde{\pi}_t \quad y_t \quad i_{t-1}]$, $\beta_t' = \left[\beta_{0,t}' \quad \beta_{1,t}' \quad \beta_{2,t}' \quad \theta_t\right]'$ and $\tilde{\pi}_t = \pi_t - \pi_t^*$ is the deviation of the inflation from its target.

For this model, the Kalman filter can be described using the following equations:

$$\beta_{t|t-1} = F\beta_{t-1|t-1}, \qquad (22)$$

$$P_{t|t-1} = FP_{t-1|t-1}F' + \Sigma_\varepsilon, \qquad (23)$$

$$\eta_{t|t-1} = i_t - X_t'\beta_{t-1|t-1} - \rho_1\sigma_e v_{1t}^* - \rho_2\sigma_e v_{2t}^*, \qquad (24)$$

$$H_{t|t-1} = X_t' P_{t|t-1} X_t + \sigma_{\omega*}^2, \qquad (25)$$

$$\beta_{t|t} = \beta_{t|t-1} + P_{t|t-1} X_t H_{t|t-1}^{-1} \eta_{t|t-1}, \qquad (26)$$

$$P_{t|t} = P_{t|t-1} - P_{t|t-1} X_t H_{t|t-1}^{-1} X_t' P_{t|t-1}. \qquad (27)$$

Although the Kalman filter provides the correct inference in β_t, the $P_{t|t-1}$ and $P_{t|t}$ variances are incorrect measures. To correct the endogeneity bias, the inference in β_t must be conditioned to the bias correction terms v_{1t}^* and v_{2t}^*. In this way, equation (27) provides the variance of β_t conditional on information up to time *t*, and on the bias correction terms. In contrast, the correct variance of β_t cannot be subjected to the bias correction terms. In order to purge the effects of these correction terms, the correct inferences of the conditional variances of β_t are obtained by the Augmented Kalman filter, where the following equations are inserted:

$$H^*_{t|t-1} = X_t' P_{t|t-1} X_t + \sigma_e^2, \tag{28}$$

$$P^*_{t|t} = P_{t|t-1} - P_{t|t-1} X_t H^{*-1}_{t|t-1} X_t' P_{t|t-1}, \tag{29}$$

$$P^*_{t+1|t} = F P^*_{t|t} F' + \Sigma_\varepsilon. \tag{30}$$

For a more accurate inference about β_t, the smooth values of these parameters, $\beta_{t|T}$, are estimated in which all the information available in the sample is used. According to Kim (2004) the smoothing filter is given by the following equations that are interacted for t=T-1, T-2,...,1:

$$\beta_{t|T} = \beta_{t|t} + P_{t|t} P^{-1}_{t+1|t} (\beta_{t+1|T} - \beta_{t+1|t}) \tag{31}$$

$$P^*_{t|T} = P^*_{t|t} + P_{t|t} P^{-1}_{t+1|t} (P^*_{t+1|T} - P^*_{t+1|t}) P^{-1\prime}_{t+1|t} P'_{t|t} \tag{32}$$

3.2 An alternative specification for the reaction function of the BCB

Following Minella et al. (2003), Minella and Souza-Sobrinho (2009) and Aragón and Portugal (2010), an alternative specification of the reaction function will also be estimated, which includes the deviation of the expected inflation from the inflation target. In this case, the monetary policy rule is expressed by:

$$i_t = \beta'_{0,t} + \beta'_{1,t}(\pi^e_{t,t+11} - \pi^*_t) + \beta'_{2,t} y_t + \theta_t i_{t-1} + e_t, \quad e_t \sim i.i.d N(0, \sigma_e^2) \tag{33}$$

$$\beta'_{i,t} = (1 - \theta_t)\beta_{i,t}, \quad \mathrm{i} = 0,1,2 \tag{34}$$

$$\beta'_{i,t} = \beta'_{i,t-1} + \varepsilon_{i,t}, \quad \varepsilon_{i,t} \sim i.i.d.N(0, \sigma^2_{\varepsilon,i}) \tag{35}$$

$$\theta_t = \theta_{t-1} + \varepsilon_{3,t}, \quad \varepsilon_{3,t} \sim i.i.d.\mathrm{N}(0, \sigma^2_{\varepsilon,3}) \tag{36}$$

where $\pi^e_{t,t+11}$ is the expected inflation, twelve months ahead, conditioned to the information available in *t*.

Since $\pi^e_{t,t+11}$ and y_t are potentially endogenous variables, the estimation procedure described above will be used in the following way:[6]

i) the regressions

$$\pi^e_{t,t+11} = z_t'\delta_1 + v_{1t}, \quad v_{1t} \sim N(0, \sigma^2_{v1}) \tag{37}$$

$$y_t = z_t'\delta_2 + v_{2t}, \quad v_{2t} \sim N(0, \sigma^2_{v2}) \tag{38}$$

[6] Regarding the determinants of the inflation expectations in Brazil, see Bevilaqua et al. (2007) and Carvalho and Minella (2012).

will be estimated through OLS or ML and those of standardized prediction errors $\hat{v}_{1t}^*$ and $\hat{v}_{2t}^*$ will be obtained;

ii) estimate the reaction function by ML through the Kalman filter

$$i_t = \beta'_{0,t} + \beta'_{1,t}\tilde{\pi}^e_{t,t+11} + \beta'_{2,t}y_t + \theta_t i_{t-1} + \rho_1\sigma_e\hat{v}_{1t}^* + \rho_2\sigma_e\hat{v}_{2t}^* + \omega_t^* \quad (39)$$

where $\tilde{\pi}^e_{t,t+11} = \pi^e_{t,t+11} - \pi_t^*$ is the deviation of the expected inflation from the inflation target.

4 Results

4.1 Description of the data and unit root tests

For the estimation of the BCB's reaction function specifications, monthly series were considered for the period between January 2000 and December 2011.[7] The series were obtained from the *sites* of the Institute of Applied Economic Research (IPEA) and the BCB.

The interest rate variable, i_t, is the annualized monthly accumulated Selic interest rate. This variable has been used as the major monetary policy instrument in the inflation targeting regime. Inflation is measured by the percentage variation accumulated during the last twelve months of the Broad Consumer Price Index (IPCA).[8] The inflation target series refers to the accumulated inflation target for the following 12 months. Since the National Monetary Council (CMN) sets inflation targets for calendar years, the data were interpolated.[9]

The expected inflation ($\pi^e_{t,t+1}$) concerns the median inflation forecast twelve months ahead (accumulated inflation between *t* e *t*+11) made by the market and collected by the BCB's Investor Relations and Special Studies Department (Gerin). For the period between January 2000 and

[7] Although the sample begins in January 2000, the observations used to estimate the reaction function (in step two) begin in November 2001. This is due to the use of the first 12 observations as initial values in the regressions, estimated in the first step, and of the following 10 observations to obtain the initial values of the regression coefficients in step two. This last procedure was suggested by Kim and Nelson (1999, 2006) to diminish the effect of arbitrary initial values of the β parameters on the value of the log-likelihood function.

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

[9] In the construction of the inflation targeting regime series, it was taken into account that the BCB pursued a set goal of 8.5% in 2003 and 5.5% in 2004, as well as a target of 5.1% in 2005. For details about the set goals and the announced goal for 2005, see Open Letters of 2003 and 2004 sent by the BCB to the Minister of Finance and the notes of the meeting of the Monetary Policy Committee (Copom) of September 2004.

October 2001, the BCB research does not present direct information regarding the expected inflation for the following twelve months but, it provides information about inflation expectations for the current year and the next. In this case, we follow Minella and Souza-Sobrinho (2009) approximating $\pi^{e}_{t,t+11}$, subtracting the inflation effective value, until the current month expectations for the current year and using the expectations for the following year proportionately to the number of remaining months

The output gap (y_t) is measured by the percentage difference between the index of industrial production seasonally adjusted (y_t) and the potential output (yp_t), which is $x_t = 100(y_t - yp_t)/yp_t$. Here arises an important problem because the potential product is an unobservable variable and therefore it must be estimated. Given this, it has become the trend of the output estimated by the Hodrick-Prescott (HP) filter, as a *proxy* for the potential output.

The instrument set includes a constant term, 1-3 lagged values of the Selic rate and of the exchange rate (ΔE_t), 1-3, 6, 9, 12 lagged values of the inflation (or expected inflation) and 2-4, 6, 9, 12 lagged values of the output gap.[10] Besides those variables, the time dummies d_{02M10} and d_{08M11} for 2002M10 and 2008M11:2009M1 were included in the regressions for the output gap (equations 14 and 38), and the dummy d_{02M11} for 2002M11 was included in the regressions for inflation and expected inflation (equations 13 and 37).[11]

Before proceeding with the estimations, the abovementioned variables were tested to see if they were stationary. To begin with, we investigated the integration order of the variables through six tests, namely: ADF (*Augmented Dickey-Fuller*); Phillips-Perron (PP); KPSS, proposed by Kwiatkowski et al. (1992); ERS, by Elliot et al. (1996); and tests MZ_{α}^{GLS} and MZ_{t}^{GLS}, suggested by Perron and Ng (1996) and Ng and Perron (2001). The null hypothesis of tests ADF, PP, ERS, MZ_{α}^{GLS} and MZ_{t}^{GLS} is that the series is non-stationary (unit root test) while in the KPSS tests, the null hypothesis is that the series is stationary.

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[10] The exchange rate is the percentage variation of the nominal exchange rate real/dollar (period average).
[11] These dummies were inserted to capture the strong current inflation increase and inflationary expectations at the end of 2002, the economic crisis of 2008 and an outlier (2002:10) in the output gap series. .

As indicated by Ng and Perron (2001), choosing the number of lags (*k*) was based on the Modified Akaike Information Criterion (MAIC) considering a maximum number of lags k_{max} = int(12(T/100)$^{1/4}$) = 13. A constant term (c) and a linear trend (t) were included as deterministic components for the cases in which those components were statistically significant.

Table 1: Unit root tests

Variable	Exogenous Regressors	ADF(k)	PP	KPSS	ERS(k)	MZ_{α}^{GLS}(k)	MZ_{t}^{GLS}(k)
i_t	c,t	-3.04$^{n.s}$(4)	-2.55$^{n.s}$	0.14***	4.53**(4)	-9.01$^{n.s}$(9)	-2.12$^{n.s}$(9)
Δi_t	-	-3.67*(0)	-3.92*	0.04$^{n.s}$	1.11*(0)	-23.3**(0)	-3.40**(0)
π_t	c	-1.75$^{n.s}$(13)	-2.27$^{n.s}$	0.48**	3.24***(13)	-6.43***(13)	-1.78***(13)
$\pi^{e}_{t,t+11}$	c	-3.09**(2)	-2.93**	0.40***	2.56**(2)	-10.5**(2)	-2.28**(2)
π^{*}_{t}	c	-2.87***(0)	-3.01**	0.11$^{n.s}$	4.84$^{n.s}$(0)	-6.57***(0)	-1.77***(0)
y_t	-	-3.51*(0)	-3.73*	0.03$^{n.s}$	1.77*(0)	-15.4*(0)	-2.77*(0)
ΔE^{*}_{t}	-	-4.49*(3)	-7.75*	0.15$^{n.s}$	0.78*(3)	-6.51***(13)	-1.75***(13)

Note: * Significant at 1%. ** Significant at 5%. *** Significant at 10%.$^{n.s}$ Non-significant.

The tests in Table 1, generally show that the root unit hypothesis can be rejected in the inflation series, expected inflation, target inflation, output gap and exchange rate. For the Selic rate, the results show that this variable is non-stationary.

Since the non-rejection of the null hypothesis of the root unit in the Selic rate may be due to the existence of a structural break in the trend function, two procedures were undertaken.[12] First, the Exp-W_{FS} statistic proposed by Perron and Yabu (2009) was used, in order to test the null hypothesis of non-structural break, in the trend function of the Selic rate against the alternative hypothesis of a break in the intercept and slope of the trend function at an unknown date.[13] The value calculated of that statistic (8.07) implies the rejection of the no break hypothesis at a significance level of 1%. In the face of this, two unit root tests with structural break were performed. Following Carrion-i-Silvestre et al. (2009), statistics MZ_{α}^{GLS} and MZ_{t}^{GLS} were used to test the unit root null hypotheses allowing a structural break in the unknown date tendency function under both hypothesis, null and

[12] See, for example, Perron (1989).

[13] Perron and Yabu (2009) introduce tests for structural breaks in the tendency function that do not need, *a priori* knowledge, if the noise component of the series is stationary or presents a root unit. These authors also show that, in the case in which the structural break is unknown, the functional Exp-W_{FS} of the Wald test produces a test with nearly identical limit distributions for the case of a noisy component I(0) or I(1). Due to this, the test procedures, with nearly the same size, can be obtained for both cases.

alternative. The values obtained for MZ_{α}^{GLS} (-29.2) and MZ_{t}^{GLS} (-3.79) allow us to reject the unit root hypothesis in the Selic rate at 5% significance.

4.2 Estimation of the reaction function with time-varying parameters

The first step to estimate the reaction function of the BCB consisted in obtaining the estimates of the standardized prediction errors, $\hat{v}_{1t}^{*}$ and $\hat{v}_{2t}^{*}$. For this purpose, equations (13), (14), (37) and (38), that relate the endogenous regressors with the instruments, were estimated through ML. As preliminary specification tests showed the presence of autoregressive conditional heteroskedasticity, it was considered that the errors of equations (13) and (37) follow a GARCH(1,1) and GARCH(2,1) process, respectively. It is also important to mention that the *F*-statistics for the estimated regressions in this first stage was always above the value of 10 indicated by Staiger and Stock (1997) as a rule-of-thumb threshold above which the weak instrument problem is not observed.

Table 2 shows the estimated parameters for the reaction function of the monetary policy (9") with and without the bias correction terms. The estimates for the standard errors $\sigma_{\varepsilon,i}$, i=0,1,2, are statistically significant, suggesting that there is a temporary variation in the β coefficients of the monetary policy rule. This evidence is corroborated by the likelihood ratio test (LR) calculated to the null hypothesis of constant parameters (H_0: $\sigma_{\varepsilon,0} = \sigma_{\varepsilon,1} = \sigma_{\varepsilon,2} = \sigma_{\varepsilon,3} = 0$).[14] For the specification with bias correction, the value and *p*-value of the LR statistics were 173.96 and 0.0000, respectively, indicating rejection of the null hypothesis at a 1% significance level. Since the LR test for the parameters stability is conservative[15], the results found here show that the BCB reaction to inflation and to the output gap have changed along time.

Regarding the endogeneity problem of the regressors in the reaction function, it can be observed that only the estimated coefficient for the correction term bias of the output gap, ρ_2, was

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[14] The log-likelihood value for the model with constant parameters and bias correction terms was -93.88

[15] See Kim and Nelson (1999, 2006).

significant. Further still, the value of the LR statistic (13.65) to test the null hypothesis of no endogeneity (H_0: $\rho_1 = \rho_2 = 0$) indicates the rejection of that hypothesis at a 1% significance level. These results show that to ignore possible endogeneity problems may result in serious biases in the time-varying coefficients of the monetary policy rule.

In order to know if the models are adequately specified, Table 2 also shows the Ljung-Box (*LB*) test for serial autocorrelation of the standardized prediction errors and for the squared standardized prediction errors, and the *H* statistic to test the null hypothesis that the standardized prediction errors are homoscedastic.[16] The results of these tests show that the standardized prediction errors are not serially correlated and show a constant variance. Besides, the null hypothesis that there is no conditional regressive heteroskedasticity (ARCH effect) in these prediction errors is not rejected.

Table 2: Estimates of the reaction function parameters (9'')

Parameters	Model with bias correction terms		Model without bias correction terms	
	Estimate	Standard deviation	Estimate	Standard deviation
$\sigma_{\varepsilon,0}$	0.0973	0.0368	0.1379	0.0232
$\sigma_{\varepsilon,1}$	0.0811	0.0103	0.0882	0.0114
$\sigma_{\varepsilon,2}$	0.0187	0.0074	0.0196	0.0096
$\sigma_{\varepsilon,3}$	7.43e-6	5.81e-5	0.0008	0.0017
σ_e	0.0946	0.0406	2.73e-5	0.0002
ρ_1	-0.1080	0.1575	-	
ρ_2	-0.8605	0.2120	-	
Specification tests				
$LB_1(24)$	21.991	(0.341)	28.020	(0.109)
$LB_2(24)$	15.672	(0.737)	10.987	(0.947)
$H(41)$	1.0173	(0.478)	1.0362	(0.455)
ln(L)	-6.8994		-13.7234	

Note: $LB_1(24)$ refers to the Ljung-Box statistic for serial autocorrelation of the standardized prediction errors up to order 24. $LB_2(24)$ refers to the Ljung-Box statistic for serial autocorrelation of the squared standardized prediction errors up to order 24. $H(41)$ refers to statistic *H* for testing the homoskedasticity of the standardized prediction errors. Values between brackets refer to the *p*-value.

The behavior of the reaction function coefficients with bias correction terms are presented below. In Figure 1, are the $\beta_{0t|T}$ trajectories and those of the persistence coefficient $\theta_{t|T}$, together with the confidence bands of ±1.0 standard deviation. As equation (8) shows, the $\beta_{0t|T}$ coefficient can be

[16] Regarding the *H statistic*, see Commandeur and Koopman (2007).

interpreted as the implicit target for the interest rate (i^*). It is possible to observe that, during most part of the period between 2003:1 and 2005:9, the estimates for the Selic rate remained above 14% per year. In contrast, during the period between 2006 and 2011, the estimated goal varied between 6.62% and 12.44%. The reduction in β_0 seems to be consistent with the higher stability of the Brazilian economy after 2003 and with the current global economic crisis since 2008, which favoured the BCB in pursuing lower targets both for inflation and Selic rates. In relation to the smoothing coefficient of the interest rate, the results reveal a relative stability of this parameter throughout time. Between November 2001 and December 2011, this coefficient fell only from 0.924 to 0.909.

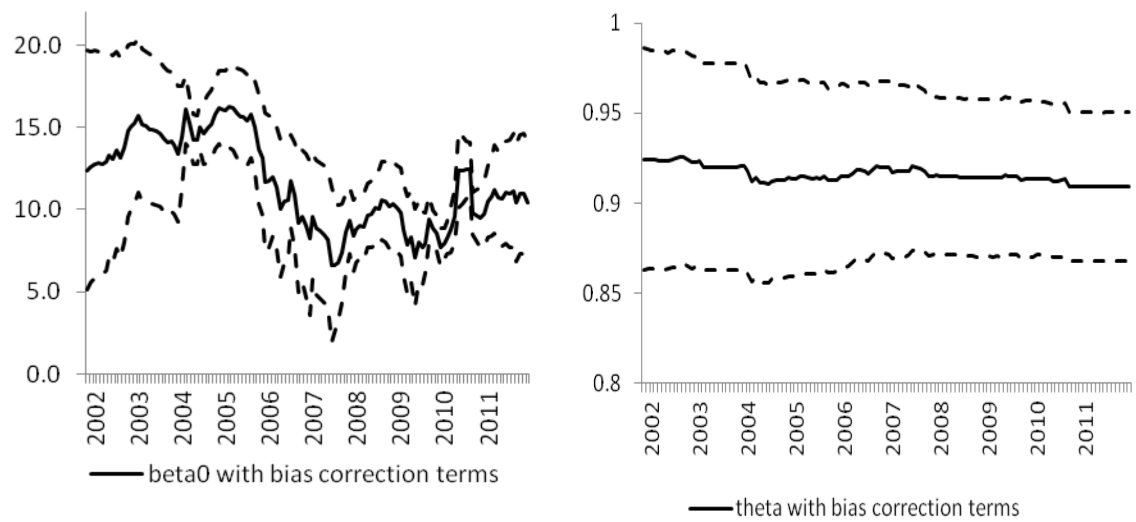

Figure 1: Evolution of coefficients $\beta_{0t|T}$ and $\theta_{t|T}$ (dashed lines indicate ±1 standard deviation)

Figure 2 shows the evolution of coefficient $\beta_{1t|T}$, which measures the long-run response of the Selic rate to the deviations of inflation from the target. The results indicate that that response presented a high oscillation during the period, varying between -3.5 and 5.5. It can also be found that, in approximately 61% of the analyzed period, the interest rate rule did not meet the Taylor principle because the value of this coefficient was less than 1 (blue line in the graph). This is in line with the evidence obtained by Bueno (2005) and Lima et al. (2007).

When comparing the behaviour of $\beta_{1t|T}$ with that of the deviation of inflation from the target, it can be verified that, in general, the BCB has risen (decreased) its reply in increase (reductions) periods in this deviation (see Fig. 2). However, two exceptions may be noted. In the first semester of 2003, the β_1 value decreased, while the inflation remained distanced from its target. This is observed again starting from March 2011, when the inflation gap rose and reached levels verified in 2005, while the response of the Selic rate to inflation was reduced.

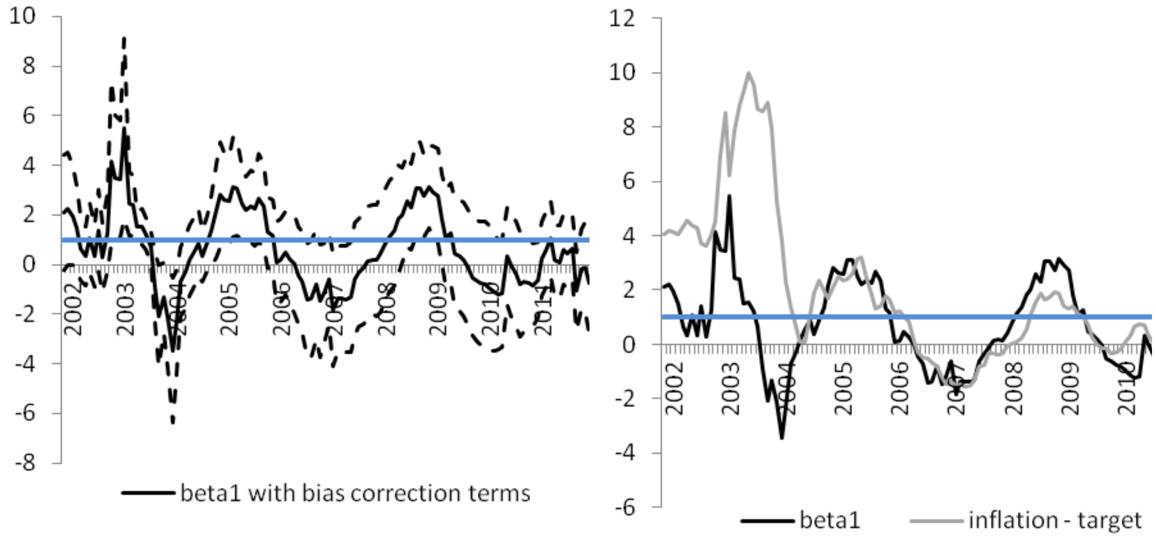

Figure 2: Evolution of the $\beta_{1t|T}$ coefficient (dashed lines indicate ±1 standard deviation) and that of the deviation of inflation from the target.

The response of the Selic rate to the output gap ($\beta_{2t|T}$) is shown in Figure 3. To begin with, it can be observed that this coefficient remained high between the fourth trimester of 2002 and the first semester of 2003, and presented more stability from 2004 until mid-2008. Although we cannot identify a clear relationship between $\beta_{2t|T}$ and the output gap, the estimates suggest that since the 2008-2009 economic crises, the BCB has increased the Selic rate response to the real activity.

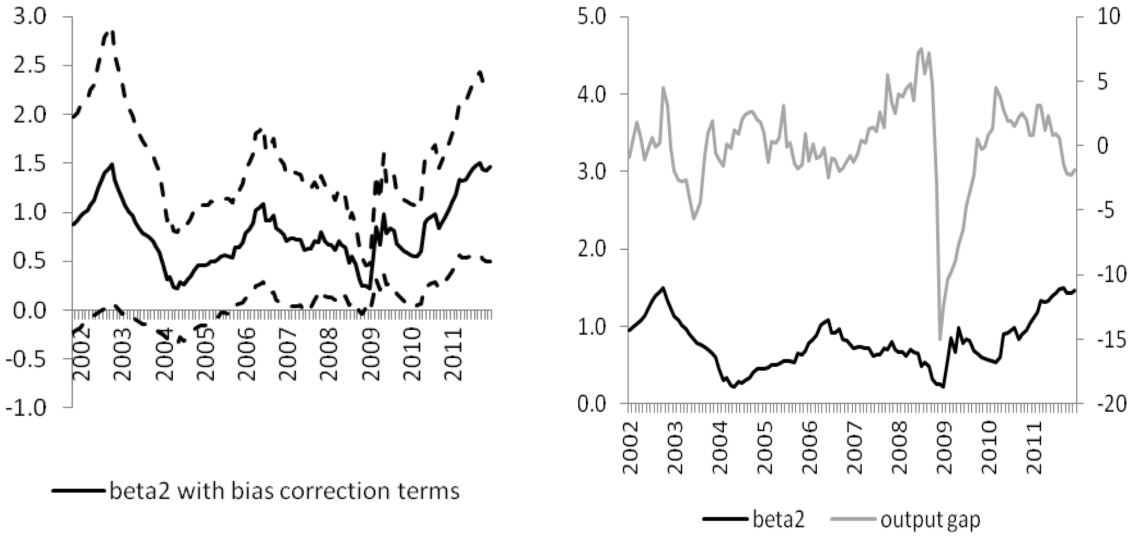

Figure 3: Evolution of the $\beta_{2t|T}$ coefficient (dashed lines indicate ±1 standard deviation) and the output gap.

The estimates of the reaction function parameters (39) are presented in Table 3. As in the previous model, the LR statistic (102,63) shows that the null hypothesis of the constant parameters is rejected at a 1% significance level.[17] Additionally, the LR test of the model with bias correction, against the model without correction, leads to the rejection of the null hypothesis of exogeneity of the expected inflation and output gap in the reaction function of the monetary policy.

Table 3: Parameter estimates of the reaction function parameters (39)

Parameters	Model with bias correction terms		Model without bias correction terms	
	Estimate	Standard Deviation	Estimate	Standard Deviation
$\sigma_{\varepsilon,0}$	0.0002	0.0009	0.0616	0.0698
$\sigma_{\varepsilon,1}$	0.1299	0.0373	0.1342	0.0378
$\sigma_{\varepsilon,2}$	2.13e-6	0.0003	4.70e-7	9.71e-6
$\sigma_{\varepsilon,3}$	0.0155	0.0019	0.0162	0.0024
σ_e	0.0619	0.0328	0.0002	0.0021
ρ_1	-0.5050	0.3594	-	
ρ_2	-0.8550	0.3937	-	
Specification Tests				
$LB_1(24)$	19.552	(0.488)	21.342	(0.377)
$LB_2(24)$	23.263	(0.276)	11.850	(0.921)
$1/H(41)$	1.4678	(0.112)	1.2767	(0.219)
ln(L)	-16.6871		-21.5890	

Note: $LB_1(24)$ refers to the Ljung-Box statistic for the serial autocorrelation of standardized prediction errors up to order 24. $LB_2(24)$ refers to the Ljung-Box statistic for serial autocorrelation of squared standardized prediction errors up to order 24. 1/H(41) refers to statistic 1/H to test the homoskedasticity of standardized prediction errors. Values between brackets refer to the *p*-value.

[17] In this case, the specification with constant parameters and bias correction terms presented a log-likelihood equal to -68.

Figures 4-6 show the trajectories of the estimated coefficients for the reaction function (39). Once again, the behaviour of $\beta_{0t|T}$ reveals a downward trend in the implicit target for the interest rate after 2003. In addition, Figure 4 shows the Selic rate smoothing, $\theta_{t|T}$, presented a small reduction, leaving 0.86 in 2001:11 to 0.78 in 2011:12.

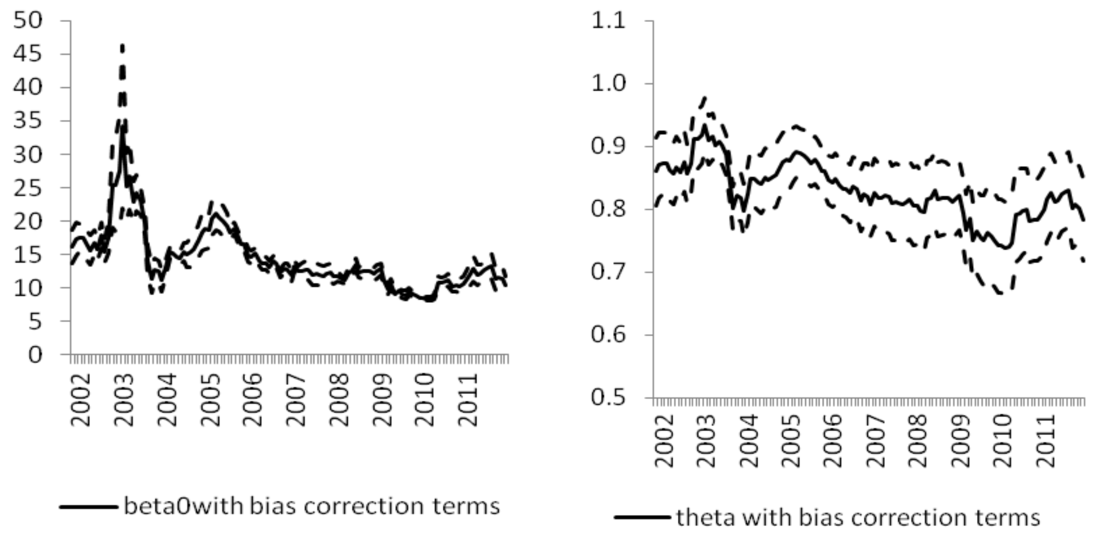

Figure 4: Evolution of coefficients $\beta_{0t|T}$ and $\theta_{t|T}$ (dashed lines indicate standard deviation)

The evolution of the long-run response of the Selic rate to the deviations of the expected inflation from the inflation target can be seen in Figure 5. Right from the start we can observe that this response satisfies Taylor principle in most part of the analyzed period. However, two exceptions to this behaviour can be highlighted. The first one concerns the passivity of the monetary policy in the months from March to September 2002, the period preceding the presidential elections of that year. The second exception is the period from 2010:9-2011:12, which has two unique characteristics: i) it is a period in which the value of β_1 diminished in spite of the inflation expectations having increased in relation to the inflation target; and ii) it is the only period in which the Selic rate response to the expected inflation reached negative values. [18]

When compared to estimates $\beta_{1t|T}$ for the reaction function (9”), shown in Figure 2, it is noteworthy that the BCB has responded more strongly to the expected inflation than to the current

[18] It is worth while pointing out that the confidence interval does not allow asserting that β_1 was significantly lower than zero during this period.

inflation. This procedure is consistent with a forward-looking policy maker and indicates that the BCB has been concerned, mainly, in anchoring inflation expectations to the inflation target set by the CMN.

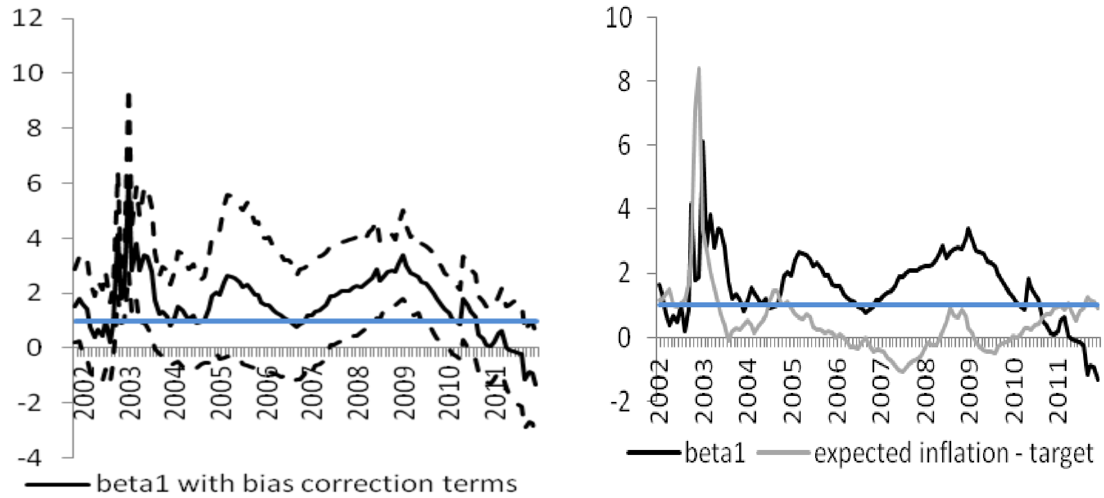

Figure 5: Evolution of the $\beta_{1t|T}$ coefficient (dashed lines indicate ±1 standard deviation) and of the expected inflation deviation from the target.

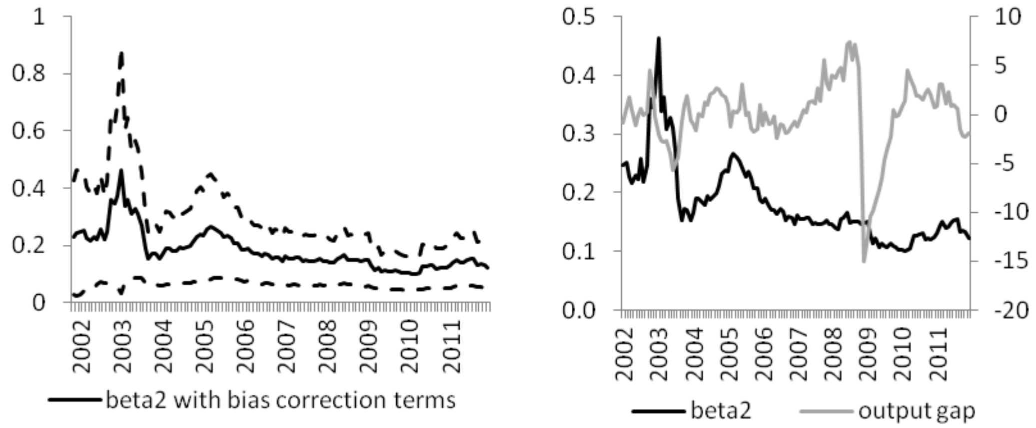

Figure 6: Evolution of the $\beta_{2t|T}$ coefficient (dashed lines indicate ±1 standard deviation) and of the output gap (y_t).

Finally, Figure 6 brings the response of the Selic rate to the output gap ($\beta_{2t|T}$).It can be observed that this coefficient has presented the most oscillation between 2002 and 2006. Starting in 2007, that coefficient remains relatively stable, varying between 0.10 and 0.17. Differently to the results presented in figure 3, we do not observe here a clear increase of that response starting from the 2008-2009 economic crisis.

5 Conclusion

In this paper, we have estimated a forward-looking reaction function with time-varying parameters, to identify possible changes in the conduction of the Brazilian monetary policy during the 2001-2011 period. In order to solve the endogeneneity problem of the policy rule regressors, a two-step procedure was used, similar to that of Heckman (1976). This methodology enables the consistent estimation of the hyper-parameters and the correct inference of the variances of the model's coefficients. Due to this, it was possible to analyze the dynamic behaviour of the BCB, in view of some macroeconomic variables such as inflation and output gap.

Before proceeding with the estimations, two LR tests were carried out. First, the validity of the null hypothesis of constant parameters was verified. The result showed that the coefficients of the BCB policy rule have changed along time. Regarding the endogeneity problem, the LR test rejected the hypothesis that inflation and the output gap are exogenous variables. Therefore, ignoring the endogeneity problems of these variables can result in a serious coefficient estimation bias.

The results obtained showed important changes in the BCB monetary policy rule coefficients. The implicit target for the Selic rate was reduced throughout the period. This probably resulted from the increased stability of the Brazilian economy after 2003 and was favoured by the recent global crisis. Regarding the response of interest rates to inflation, there was a considerable variation in time, albeit with a declining trend. The empirical evidence also indicates that: i) in general, the greater the inflation deviation (observed or expected) in relation to the target, the greater the policy response to that variable; ii) the BCB has responded more strongly to the expected inflation than to the observed inflation, reflecting in this way the forward-looking behaviour of this monetary authority; iii) since mid-2010, the response to inflation has been lower than 1, thus not satisfying the Taylor principle.

The response of the policy to the output gap differed between the reaction function specifications. When the observed inflation was inserted in the monetary rule, a relative stability of this response was noticed between 2003 and 2008, and an increase since the 2009 economic crisis. As to the specification of the reaction function, including the expected inflation, this rate remained stable after 2003.

For future research, this work may be advanced as follows: i) perform estimations with time-varying parameters for the reaction function specifications that are non-linear, due to the asymmetric preferences of the Central Bank (see, for example, Aragon and Portugal, 2010); ii) consider that the relationship between the endogenous regressors and its instruments are time-varying.

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