

Monetary policy in Costa Rica: an assessment based on the neutral real interest rate

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Abstract

In this study we assess the monetary policy stance in Costa Rica during the 2009-2018 period using an indicator of the real policy rate gap. We obtain estimates of the real neutral interest rate by using six methodologies, whose empirical consistency is evaluated in order to decide whether they are used in the final estimation. The updated value for the real neutral interest rate is 1.54%. The policy rate gap indicator shows appropriate empirical properties, among them a negative lead correlation with the output gap and core inflation. This suggests that the policy rate is successfully influencing the marginal cost of liquidity for financial intermediaries. Our analysis suggests that monetary policy in Costa Rica has responded mainly to inflation movements not related to temporary shocks, and that some policy adjustments could have been swifter.

Key words: monetary policy, inflation, unobservable variables.

JEL codes: E12, E31, E52.

1. INTRODUCTION

In 1898, the Swedish economist Knut Wicksell introduced the concept of the natural (or neutral) interest rate as “a certain rate of interest on loans which is neutral in respect to commodity prices, and tends neither to raise nor to lower them.” (Wicksell, 1938,

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pp.106). Wicksell argued that as long as the interest rate was lower than the natural rate of return on capital, there would be an incentive to borrow in order to accumulate capital. This accumulation process would eventually lead to an increase in the general price level due to the increase in aggregate demand. Thus, for Wicksell price stability would be achieved only if, all else constant, permanent discrepancies could be avoided between the current interest rate and the natural rate. Consequently, a discussion about the level of the interest rate would only make sense when it is compared to the level of the natural rate. However, Wicksell acknowledged that the natural rate is neither observable nor constant, since it would depend on the evolution of the factors that influence the return of capital.

The most relevant contribution to consolidate the concept of the natural interest rate in modern macroeconomics was made by Woodford (2003), who demonstrated that it is possible to derive a natural rate conceptually equivalent to that of Wicksell as a function of fundamental variables (consumer preferences, shocks of productivity) within the framework of a neo-Keynesian model, and that such rate is the one that would prevail in the absence of nominal rigidities.

Therefore, Woodford formulated the theoretical fundamentals that support the strategy followed by many central banks to conduct economic policy of price stabilization, by demonstrating that it is feasible to perform monetary policy through a rule for the interest rate without explicitly taking into account the money supply. This is particularly relevant for central banks that follow an inflation targeting scheme.

The neutral real interest rate (NRIR) is a fundamental reference to characterize, in real time and ex post, the monetary policy stance: contractive when the policy rate is higher than the NRIR, expansive when it is lower. Therefore, it is very useful for a central bank to have reliable estimates of the gap between the monetary policy rate and the NRIR when facing the monetary policy decision-making process. This is a challenge because the NRIR is an unobservable variable.

The Central Bank of Costa Rica (BCCR) has made several estimates of the NRIR for Costa Rica. The first exercise, by Muñoz and Tenorio (2007), used data for the period 1991-2006, under a crawling peg exchange rate regime. The NRIR was estimated using four methodologies: a semi-structural model proposed by Laubach and Williams (2003), the uncovered parity condition of interest, the Hodrick-Prescott filter to obtain a long-term trend, and an ad-hoc

approximation corresponding to the average of the effective real interest rate under a period of stable inflation.

Segura and Vindas (2012) used information for the period 2001-2011, which includes two exchange rate regimes (a crawling peg and a crawling band) and in addition to the methodologies used by Muñoz and Tenorio, they included a VAR estimation proposed by Brzoza-Brzezina (2003, 2006). Muñoz y Rodríguez (2016) used data for the period 2009-2015 under a more flexible exchange rate regime, along with the methodologies used in previous studies, their estimations incorporate Taylor rule state space models.

International organizations like the IMF (2016) and the OECD (2016) also estimated the NRIR as part of their evaluations of the Costa Rican economy. The IMF relied on a neo-Keynesian semi-structural model, a general equilibrium model, a monetary model and a Taylor rule augmented by expectations. This last method is also the one used by the OECD.

A common feature of previous studies for Costa Rica is their emphasis on the estimation itself. The present investigation has the objective of approximating an interest rate gap with coherent empirical properties and then analyzing its relationship with the state of the Costa Rican economy, which would allow us to make conclusions about the stance of the BCCR monetary policy. For the estimation of the NRIR we add two more methods to those used in Muñoz and Rodríguez (2016).

This paper is organized as follows: the methodology description is presented in section 2. The main results and the evaluation of the policy stand are discussed in section 3, finally the main conclusions are presented in section 4.

2. DATA AND METHODOLOGY

2.1 Monetary policy rate

In 2005, the BCCR began a process of modernization of its monetary policy in order to improve the compliance of the objective established by its Organic Law of maintaining the internal stability of the national currency. This mandate is interpreted as the achievement of low and stable inflation, in line with that of the main trading partners

of the country. This process led to the adoption of an inflation targeting monetary policy framework in January 2018.

An important decision in this process was the introduction of the monetary policy rate (MPR) as an instrument of monetary control, in June 2011¹. The greater variability of the exchange rate resulting from the adoption of the crawling band exchange rate regime at the end of 2006 and of the managed float at the beginning of 2015 has allowed an increasing independence for the use of the MPR, after decades of a crawling peg scheme that restricted the possibility of using the interest rate as a monetary policy instrument².

The MPR is defined as “...interest rate target of the Central Bank of Costa Rica. This indicator corresponds to the interest rate the Central Bank of Costa Rica uses as a reference in guiding the cost of one-day operations in the Integrated Liquidity Market into a corridor determined by the interest rates of the permanent facilities of credit and deposit in that market.”³

By changing this rate the BCCR aims to influence the marginal cost of liquidity for the financial intermediaries to channel the desired stance of monetary policy to the rest of interest rates in the financial system.

Before having this formal definition of the indicator, all empirical research requiring a monetary policy rate lacked a coherent monetary policy rate series long enough to allow quantitative analysis. For example, Muñoz and Tenorio (2007) and Segura and Vindas (2012) used the Deposit Basic Rate as an indicator of BCCR’s policy rate.

To overcome this issue, Castro and Chaverri (2013) defines an indicator that reflects the monetary policy stance the BCCR during the period January 1999 - May 2011. The data series resulting from linking the monetary policy rate indicator of Castro and Chaverri (2013) with the data of the MPR as of June 2011 is shown in Figure 1. This is the variable used as the rate of monetary policy throughout the present investigation.

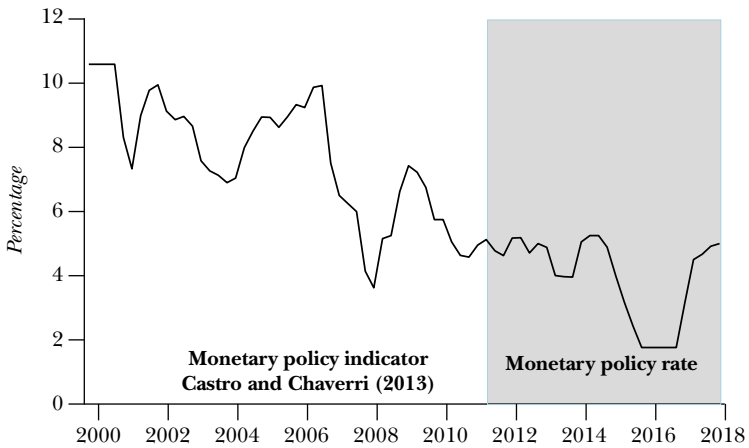
¹ Since 2005, the legal reserve requirement is at the maximum level allowed by the Organic Law.

² The increasing exchange rate volatility associated to the adoption of more flexible Exchange regimes was documented in BCCR (2018).

³ Monetary Policy Regulations, Title IV, Numeral 2, Literal D.

Figure 1

COSTA RICA: MONETARY POLICY



Source: Own elaboration with information from Castro and Chaverri (2013) and BCCR.

Additionally, Table 1 presents the data used in this study, as well as the corresponding source. The estimation period varies according to each methodology, but for the analysis of the NRIR the period 2009Q1-2018Q4 is considered.

2.2 Estimation methods

The NRIR is an unobservable variable whose level can change depending on the macroeconomic conditions, therefore a generalized practice is to use a set of methodologies that provide the authorities with a range of estimates. This research starts from the methodologies implemented by Muñoz and Rodríguez (2016) and incorporates two additional ones: a Bayesian SVAR with changing coefficients and a linear local trend model.

2.2.1 Bayesian SVAR with changing coefficients

This method follows the approach used by Lubik and Matthes (2015) of estimating a system of structural autoregressive vectors (SVAR)

Table 1

VARIABLES AND SOURCES		
<i>Variable</i>	<i>Data used</i>	<i>Source</i>
R_t	Nominal monetary policy rate	Central Bank of Costa Rica
R_t^{lp}	Weighted Treasury bonds rate, 5 years or more, secondary market	Superintendency of Securities (SUGEVAL)
π_t	12-month change in Consumer Price Index	National Institute for Statistics and Censuses (INEC)
π_t^M	BCCR target for 12-month inflation rate	Central Bank of Costa Rica
π_{t+n}^e	12-month change in Consumer Price Index at t+n	National Institute for Statistics and Censuses (INEC)
q_t	Tipo de cambio efectivo real multilateral con ponderadores móviles, logaritmo natural	Central Bank of Costa Rica
ρ_t	EMBI for Costa Rica	Bloomberg
α_t	Spread for 2020 bonds, Government of Costa Rica	Bloomberg
α	Average of α_t for estimation period	
y	Gross Domestic Product, chained volume at previous year's prices	Central Bank of Costa Rica
y^b	Percent difference of real GDP with respect to its potential level (Hodrick-Prescott, $\lambda = 1800$)	Own estimation

Source: own elaboration.

with changing coefficients (Time-varying parameter VAR, TVP-VAR) with real GDP growth rate, CPI inflation and the monetary policy rate as endogenous variables. The SVAR corresponds to that proposed by Primiceri (2005), where the variability of the system comes from the changing coefficients and stochastic volatility in the covariance matrix of the shocks that enter the model. Primiceri argues that the changing coefficients allow to capture possible nonlinearities as well as temporal variation in the lag structure of the model. Lubik and Matthes argue that this makes the method particularly apt to capture both the secular changes in the NRIR and those associated to the business cycle.

The model proposed by Primiceri (2005) is as follows

$$1 \quad Y_t = c_t + B_{1,t}Y_{t-1} + \dots + B_{k,t}Y_{t-k} + u_t \quad t = 1, \dots, T$$

where the variance covariance matrix Ω_t of the shocks u_t given by

$$2 \quad A_t \Omega_t A_t' = \Sigma_t \Sigma_t', \quad A_t = \begin{bmatrix} 1 & 0 & \dots & 0 \\ \alpha_{21,t} & 1 & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ \alpha_{n1,t} & \dots & \alpha_{nn-1,t} & 1 \end{bmatrix} \quad \Sigma_t = \begin{bmatrix} \sigma_{1,t} & 0 & \dots & 0 \\ 0 & \sigma_{2,t} & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \dots & 0 & \sigma_{n,t} \end{bmatrix}$$

Therefore, the system can be rewritten as

$$3 \quad \begin{aligned} Y_t &= X_t' B_t + A_t^{-1} \Sigma_t \varepsilon_t, & V(\varepsilon_t) &= I_n \\ X_t' &= I_n \otimes (1, y_{t-1}', \dots, y_{t-k}') \end{aligned}$$

Where the right hand side coefficients in are stacked in the vector B_t . In our case, $Y_t = (y_t, \pi_t, R_t)$. The dynamics of the time varying coefficients vector is specified as follows:

$$4 \quad \begin{aligned} B_t &= B_{t-1} + v_t \\ \alpha_t &= \alpha_{t-1} + \varsigma_t \\ \log \sigma_t &= \log \sigma_{t-1} + \eta_t \end{aligned}$$

In this model the innovations are jointly normally distributed with the following assumptions on the variance covariance matrix:

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$$V = \text{Var} \begin{pmatrix} \varepsilon_t \\ v_t \\ S_t \\ \eta_t \end{pmatrix} = \begin{bmatrix} I_n & 0 & 0 & 0 \\ 0 & Q & 0 & 0 \\ 0 & 0 & S & 0 \\ 0 & 0 & 0 & W \end{bmatrix}$$

where Q, S, W are positive definite matrices.

The model is estimated by Bayesian methods. In particular, the Gibbs sampler is used for the numerical evaluation of the posterior distribution of all parameters. The estimate of the NRIR corresponds to the average of the samples for the posterior density of the constant parameter in the equation for the interest rate.

For estimation, the a priori distributions of Primiceri (2005) were used:

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$$\begin{aligned} B_0 &\sim N(\hat{B}_{MCO}, 4 \cdot V(\hat{B}_{MCO})) \\ A_0 &\sim N(\hat{A}_{MCO}, 4 \cdot V(\hat{A}_{MCO})) \\ \log \sigma_0 &\sim N(\log \hat{\sigma}_{MCO}, I_n) \\ Q &\sim IW(k_Q^2 \cdot 40 \cdot V(\hat{B}_{MCO}), 40) \\ W &\sim IW(k_W^2 \cdot 4 \cdot I_n, 4) \\ S_1 &\sim IW(k_S^2 \cdot 2 \cdot V(\hat{A}_{1,MCO}), 2) \\ S_2 &\sim IW(k_S^2 \cdot 3 \cdot V(\hat{A}_{2,MCO}), 3) \end{aligned}$$

The model was estimated with quarterly data, with a calibration sample comprising 2002Q1 – 2008Q4, and an estimation method comprising 2009Q1 – 2018Q4.

2.2.2 Local linear trend model

The basis for this method is the decomposition formalized by the structural time series model by Harvey (1989), to express the real interest rate r_t as the sum of trend, cycle, seasonality and irregular components:

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$$r_t = \mu_t + \psi_t + \gamma_t + \varepsilon_t$$

where μ_t is the trend, ψ_t is the cycle y γ_t is the seasonal component. All components are stochastic and their perturbations are not

correlated. The real rate is computed from the monetary policy rate and the lead of the inflation rate.

The trend component μ_t evolves following a local linear trend model

$$\begin{aligned} \mu_t &= \mu_{t-1} + \beta_{t-1} + \eta_t \\ \beta_t &= \beta_{t-1} + \varsigma_t \end{aligned} \tag{8}$$

where η_t y ς_t are uncorrelated white-noise perturbations, with zero means and variances σ_η^2 y σ_ς^2 .

The ψ_t component is modeled as a cyclical function of time with frequency λ_c , that can be represented recursively as:

$$\begin{pmatrix} \psi_t \\ \psi_t^* \end{pmatrix} = \rho \begin{pmatrix} \cos \lambda_c & \text{sen} \lambda_c \\ -\text{sen} \lambda_c & \cos \lambda_c \end{pmatrix} \begin{pmatrix} \psi_{t-1} \\ \psi_{t-1}^* \end{pmatrix} + \begin{pmatrix} k_t \\ k_t^* \end{pmatrix} \tag{9}$$

where k_t and k_t^* are uncorrelated and have common variance σ_k^2 , and the model is stationary if $|\rho| < 1$.⁴

In this study, seasonality γ_t is introduced through a group of binary seasonal variables, whose effect sums to zero for the entire year:

$$\sum_{j=0}^{s-1} \gamma_{t-j} = \omega_t \tag{10}$$

where $s=4$ and ω_t is a zero-mean perturbation with variance σ_ω^2 .

The model is estimated in its state-space representation with the Kalman filter using monthly data for the period January 2000–June 2018. In this case the NRIR estimate is the smoothed estimate of μ_t in the signal equation 8. For the analysis, we take the quarter average of these values.

2.2.3 *Semistructural macroeconomic model with unobservable components*

This well-known approach pioneered by Laubach and Williams (2003)⁵ approximates the NRIR for the US economy through a parsimonious

⁴ The variables with a star result from the properties of the trigonometric functions, but have no interpretation. See Harvey (1989).

⁵ See also Laubach and Williams (2015) and Holston, Laubach and Williams (2016).

state-space model with new Keynesian theoretical underpinnings, in which changes in the interest rate affect consumption and investment decisions. The two signal equations are: an IS curve to describe equilibrium in the goods and services market (11); and a Phillips curve to explain the evolution of inflation (12). Furthermore, it is assumed that potential output grows at a rate that evolves following a random walk, like in equation (13), and that the NRIR is determined by potential output growth, like in equation (14). The complete system is:

$$11 \quad (y_t - y_t^p) = \sum_{s=1}^S \alpha_s^y (y_{t-s} - y_{t-s}^p) + \sum_{v=1}^V \alpha_v^r (r_{t-v} - r_{t-v}^n) + \chi_{1,t}' \alpha + \varepsilon_t^y$$

$$12 \quad (\pi_t - \pi_t^M) = \sum_{p=1}^P \beta_p^\pi (\pi_{t-p} - \pi_{t-p}^M) + \sum_{q=1}^Q \beta_q^y (y_{t-q} - y_{t-q}^p) + \chi_{2,t}' \beta + \varepsilon_t^\pi$$

$$13 \quad y_t^p = y_{t-1}^p + \dot{y}_{t-1}^p + \varepsilon_t^{y^p}$$

$$14 \quad r_t^n = c \dot{y}_t^p + z_t$$

where:

- y_t natural logarithm of GDP
- y_t^p natural logarithm of potential output
- r_t real monetary policy rate
- π_t^M inflation target
- $\chi_{1,t}'$ other variables explaining the output gap
- $\chi_{2,t}'$ other variables explaining
- $\varepsilon_t^{y,\pi}$ white-noise process with zero mean and constant variance
- \dot{y}_t^p potential output growth
- $\varepsilon_t^{y^p}$ zero-mean perturbation with constant variance
- z other factors explaining the neutral real rate

The other factors explaining the NRIR, z_t , are modeled as AR(1) from the estimated OLS errors of equation (14):

$$z_t = \delta z_{t-1} + \varepsilon_t^z$$

Estimation is done with the Kalman filter, with quarterly data for the 2009Q1– 2018Q2 period. Initial values for the parameters and for the state variables come from the OLS estimation of the signal equations. Initial values of unobservable variables (potential output and real neutral rate) were obtained from applying the Hodrick-Prescott filter to the GDP and real interest rate series.

2.2.4 *Implicit common stochastic trend*

Basdevant, Björkstén and Karagedikli (2004) proposed a model based on the assumption that the yield curve can be informative about the monetary policy stance. They state that there is a stochastic common trend in the nominal long-run and short-run interest rates, which is expressed in state-space form in the following way:

$$\begin{aligned} R_t^{\phi} &= r_t^n + \pi_{t+12}^e + \varepsilon_t^1 \\ R_t^{\psi} &= r_t^n + \pi_{t+12}^e + \alpha_t + \varepsilon_t^2 \\ r_t^n &= r_{t-1}^n + \vartheta_t^1 \\ \alpha_t &= \mu_0 + \mu_1 \alpha_{t-1} + \vartheta_t^2 \end{aligned}$$

Where the first two are signal equations and the rest are transition equations, with

- R^{ϕ} short-run nominal interest rate
- R^{ψ} long-run nominal interest rate
- π^e 12-month inflation expectations
- α_t risk premium

The system indicates that the short-term nominal interest rate equals the NRIR plus the 12-month inflation expectations and a stochastic perturbation, while the long-run nominal interest rate equals the short-run nominal rate plus a risk premium and a stochastic

perturbation. The transition equations assume a random walk for the NRIR and a stationary AR(1) for the risk premium. Perturbations are assumed independent and identically distributed with zero mean and constant variance. The model is estimated with the Kalman filter for the 2009Q1-2018Q2⁶ period.

2.2.5 Dynamic Taylor rules

Magud and Tsounta (2012) estimate the NRIR using two versions of the Taylor rule. In the first one, the nominal neutral interest rate comes from a formulation in which the nominal policy rate depends on the deviations of inflation from the central bank target, and of output with respect to its potential level, so that when both gaps are zero the short-run interest rate equals the nominal neutral rate. The model is expressed in state-space form with state equations for the short-run nominal interest rate (R_t^{sp}) and two transition equations: one for the nominal neutral rate (R_t^n) and the other for its variation rate g_t :

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$$\begin{aligned} R_t^{sp} &= R_t^n + \beta(\pi_t - \pi_t^M) + \theta y_t^b + \varepsilon_t^1 \\ R_t^n &= R_{t-1}^n + g_{t-1} \\ g_t &= g_{t-1} + \vartheta_t^1 \end{aligned}$$

where the transition process is a random walk for g_t . All perturbations are assumed independently and identically distributed with zero mean and constant variance. The model is estimated using the Kalman filter, with monthly data for the period January 2009–February 2018. The NRIR is obtained from the estimate for R_t^n , then averaged by quarter.

The second specification is an expectations-augmented Taylor rule, where the NRIR is estimated using a model that now includes signal equations both for the short-term and long-term nominal interest rates, with the same transition dynamics as in the previous model:

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$$\begin{aligned} R_t^{sp} &= r_t^n + \pi_{t+12}^e + \beta(\pi_t - \pi_t^M) + \theta y_t^b + \varepsilon_t^1 \\ R_t^{lp} &= r_t^n + \pi_{t+12}^e + \alpha + \varepsilon_t^2 \\ r_t^n &= r_{t-1}^n + g_{t-1} \\ g_t &= g_{t-1} + \vartheta_t^1 \end{aligned}$$

⁶ This approach is also used by Magud and Tsounta (2012).

Note that the nominal neutral interest rate now is given by the NRIR (r_t^n) plus the 12-month inflation expectation (π_{t+12}^e) .

2.2.6 Structural VAR with long-run restrictions

Brzoza-Brzezina (2002) suggests estimating the NRIR using a structural VAR with long-run restrictions à la Blanchard and Quah (1989). The real interest rate is defined as the sum of the NRIR and the interest rate gap:

$$20 \quad r_t = r_t^n + r_t^b$$

Besides, it is assumed that the neutral real interest rate and the interest rate gap follow stationary AR processes given by

$$21 \quad \begin{aligned} r_t^n &= \Phi_1(L)r_{t-1}^n + u_{1t} = \Xi_1(L)u_{1t} \\ r_t^b &= \Phi_2(L)r_{t-1}^b + u_{2t} = \Xi_2(L)u_{2t} \end{aligned}$$

where $\Xi_1(L)$ and $\Xi_2(L)$ are lag polynomials such that $\Xi_1(L) = (I - \Phi(L)L)^{-1}$. Thus, the interest rate can be expressed in terms of the structural perturbations u_{1t} and u_{2t} :

$$22 \quad r_t = \Xi_1(L)u_{1t} + \Xi_2(L)u_{2t}$$

Brzoza-Brzezina assumes that changes in inflation are a fraction Ψ of the deviation for the interest rate with respect to its neutral level:

$$23 \quad \Delta \pi_t = \Psi(r_t^b) = \Psi[\Xi_2(L)]u_{2t}$$

With this, the change in inflation and in the interest rate can be expressed from the structural perturbations:

$$24 \quad \begin{bmatrix} \Delta \pi_t \\ r_t \end{bmatrix} = \begin{bmatrix} S_{11}(L) & S_{12}(L) \\ S_{21}(L) & S_{22}(L) \end{bmatrix} \begin{bmatrix} u_{1t} \\ u_{2t} \end{bmatrix}$$

where $S_{ij}(L)$ are lag polynomials.

It is necessary to recover the structural perturbations u_t . For that, an unrestricted VAR is estimated

$$24 \quad \begin{bmatrix} \Delta \pi_t \\ r_t \end{bmatrix} = \begin{bmatrix} A_{11}(L) & A_{12}(L) \\ A_{21}(L) & A_{22}(L) \end{bmatrix} \begin{bmatrix} \Delta \pi_{t-1} \\ r_{t-1} \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix}$$

whose moving-average representation is

$$25 \quad \begin{bmatrix} \Delta \pi_t \\ r_t \end{bmatrix} = \begin{bmatrix} C_{11}(L) & C_{12}(L) \\ C_{21}(L) & C_{22}(L) \end{bmatrix} \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix}$$

In which reduced-form perturbations ε_{it} are a function of the structural perturbations:

$$26 \quad \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix} = \begin{bmatrix} s_{11}(L) & s_{12}(L) \\ s_{21}(L) & s_{22}(L) \end{bmatrix} \begin{bmatrix} u_{1t} \\ u_{2t} \end{bmatrix}$$

If the coefficients $S_{ij}(0)$ were known it would be possible to recover the structural perturbation from the residuals ε_t from the unrestricted VAR. To that end, Brzoza-Brzezina imposes the following restrictions:

- i) The variances of the structural perturbations are all equal to 1.
- ii) A long-run restriction is imposed so that $S_{11}(1)=0$ in the original system, which implies that the perturbation u_{1t} does not affect $\Delta\pi_t$.
- iii) Besides $S_{12}(0)=0$, so that the interest rate gap has no contemporaneous effect on inflation, that is, monetary policy operates with a lag.

From the estimated variance-covariance matrix for perturbations ε_t and restrictions (i)-(iii) it is possible to obtain the other coefficients $S_{ij}(0)$ according to:

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$$\begin{aligned}
 s_{11}(0) &= \pm \sqrt{\text{var}(\varepsilon_{1t})} \\
 s_{21}(0) &= \pm \left[\frac{C_{11}(1)}{C_{12}(1)} \right] \sqrt{\text{var}(\varepsilon_{1t})} \\
 s_{22}(0) &= \sqrt{-2 \frac{s_{21}(0)}{s_{11}(0)} \text{cov}(\varepsilon_{1t}, \varepsilon_{2t}) + s_{21}^2(0) + \text{var}(\varepsilon_{2t})}
 \end{aligned}$$

Finally, the NRIR can be computed as the result solely of structural perturbations:

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$$r_t^n = S_{21}(L) u_{1t}$$

Estimation is performed with monthly data for the period comprising January 2009–June 2018. Monthly estimates of r_t^n are averaged for analysis.

3. RESULTS

3.1 Estimates of the interest rate gap

The aim of this section is to assess the estimates of the interest rate gap obtained by applying the methods described in Section 2, in order to select the most suitable to compute a single gap indicator to characterize the BCCR's monetary policy stance during the period of analysis (Figure 11 in the Appendix shows the estimates of the neutral real interest rate). Figure 2 shows the estimates, computed as the percentage point differential in the real monetary policy rate and every NRIR estimate⁷. Most estimates show a similar behavior, with periods of expansionary or contractionary monetary policy that are relatively coincident.

The most notable differences occur in the estimates of the Taylor rules and the stochastic common trend. In particular, the gap estimated with the expectations-augmented Taylor rule remains

⁷ A one-year lead of inflation was used to compute the real interest rate.

negative during almost all the period considered, and the gaps resulting from the simple Taylor rule and the stochastic common trend show comparatively low values, stemming from NRIR estimates very close to the effective real rate.

In order to select the gap estimates, three requisites have to be met. First, when the real interest rate is higher than the NRIR, the monetary policy stance is contractionary, so that if that positive gap persists it is reasonable to expect a reduction in aggregate demand and then in the output gap and inflation. Thus, positive (negative) rate gaps should precede reductions (increases) in the output and inflation. Furthermore, even though monetary policy is forward-looking, it is also likely that policymakers respond to the contemporaneous behavior of inflation. Thus, it is also reasonable to assess whether the rate gap is correlated with it.

On the other hand, the evolution of each interest rate gap is compared with that of a Financial Condition Index (FCI) for Costa Rica obtained from the combination of 33 financial indicators using principal components. These indicators include credit and monetary aggregates, prices, interest rates and margins (including the monetary policy rate), ratios and indicators of the financial system, and factors linked with the state of the global economy. For more details see Álvarez (2016). Increases in the FCI point to less restrictive financial conditions, while reductions in it indicate the opposite case. Since the monetary policy stance should be transmitted to the financial sector, it should be expected that increases in the interest rate gap precede or coincide with reductions in the FCI.

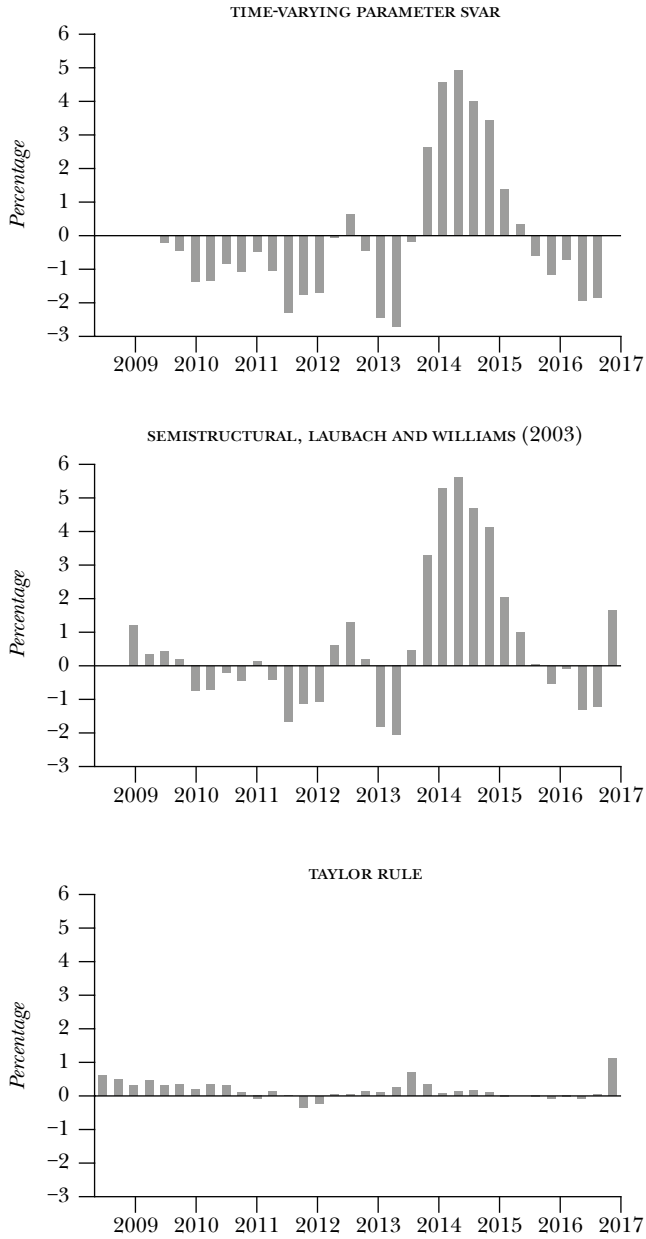
Ideally, an interest rate gap should meet these criteria, so for computation of the final indicator estimates those that do not are dropped. To verify compliance, we compute correlation coefficients between the output gap, core inflation and the FCI with several lags and leads of each interest rate gap⁸.

Most of the correlations with the output gap (Figure 3) show the expected pattern, that is, the highest correlation within the policy

⁸ The output gap is the one used in the macroeconomic modelling by the BCCR, which corresponds to an average of an estimate of potential output using a production function (Álvarez, 2018) and an estimate using a Hodrick-Prescott filter ($\lambda = 1800$). Details on trimmed-mean measure of core inflation can be found in Esquivel, Rodríguez and Vásquez (2011).

Figure 2

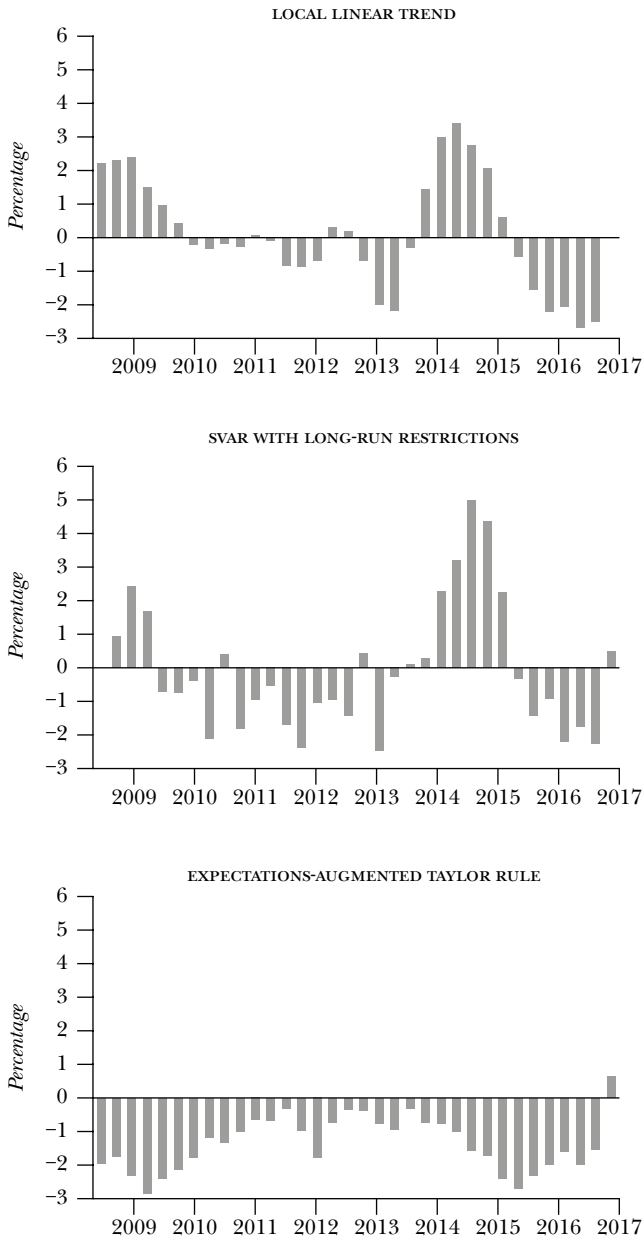
REAL RATE GAP ESTIMATES 2009Q1–2018Q2



Source: Own elaboration.

Figure 2 (cont.)

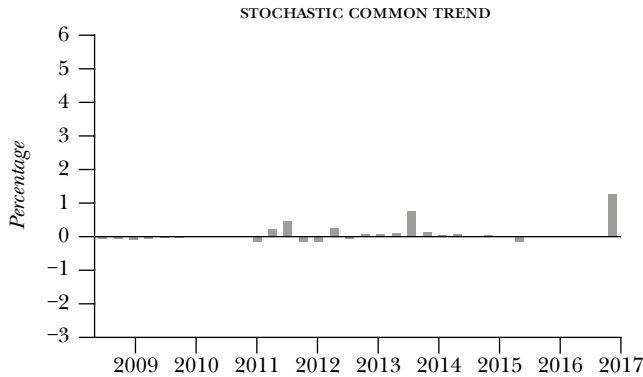
REAL RATE GAP ESTIMATES 2009Q1–2018Q2



Source: Own elaboration.

Figure 2 (cont.)

REAL RATE GAP ESTIMATES 2009Q1–2018Q2



Source: Own elaboration.

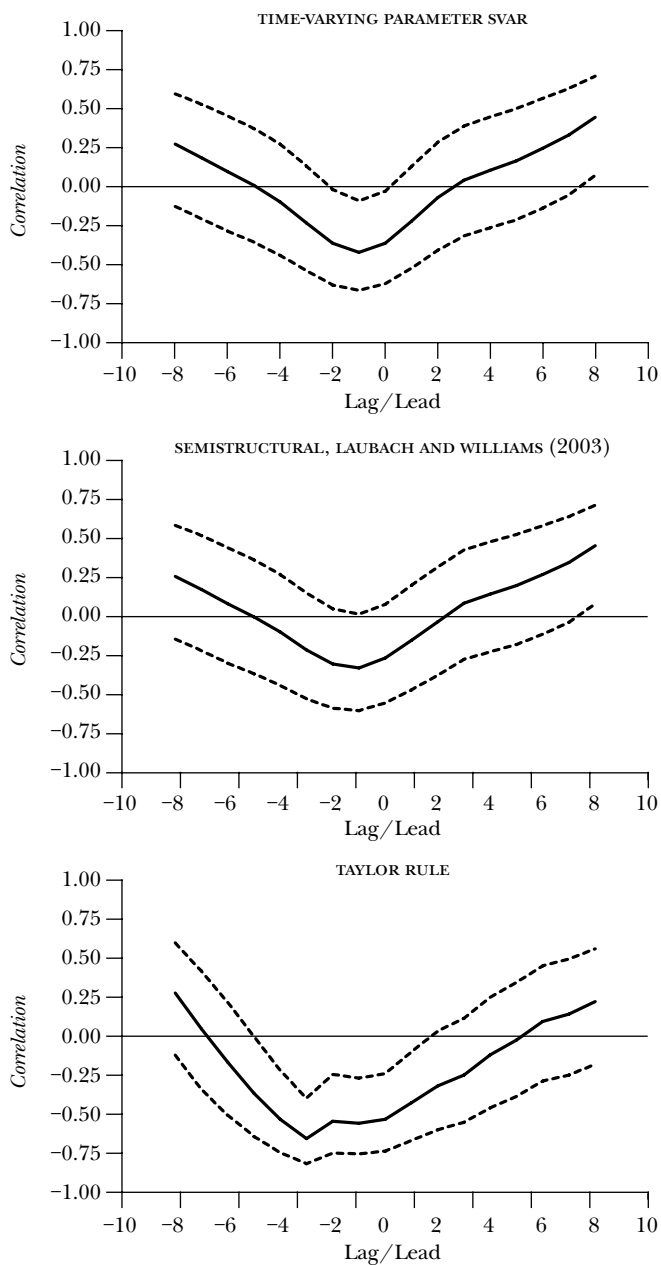
horizon is negative and precedes the output gap. This is the case for both SVAR models, the semi-structural model by Laubach and Williams (2003), the local linear trend model and the simple Taylor rule model, where in most cases the highest correlation of the output gap occurs with the interest rate gap of 1 to 3 quarters before.

On the other hand, in the case of the rate gap associated with the expectations-augmented Taylor rule, the highest correlation is positive and indicates that the output gap precedes the rate gap, while in the case of the stochastic trend model, correlations are relatively low and mostly positive.

Correlations with the trimmed-mean measure of core inflation are shown in Figure 4. The rate gaps of most models show the expected pattern: high negative correlations between core inflation and the rate gap about 4 quarters earlier (which are indicative of a delay in the effect of monetary policy on inflation) and positive contemporaneous correlations (which suggest a reaction of monetary policy to the current behavior of inflation). This pattern is not present for the estimates based on the Taylor rules or in the stochastic common trend.

Figure 3

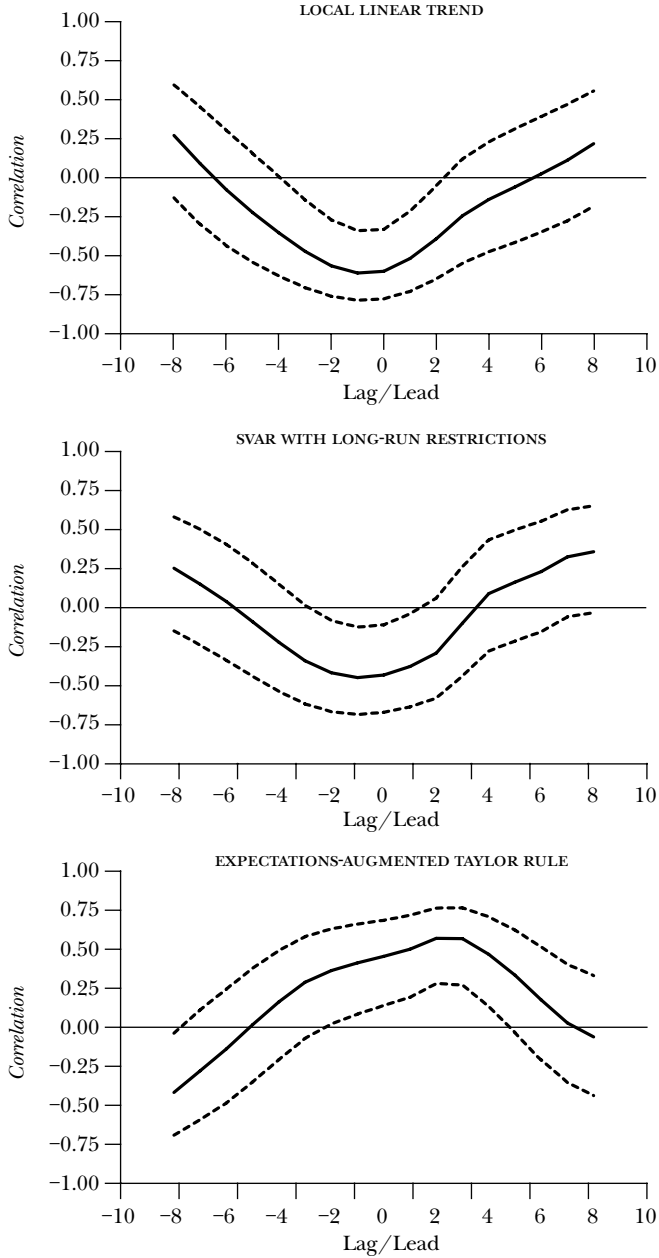
CROSS-CORRELATIONS OF RATE GAPS WITH OUTPUT GAP



Note: Confidence intervals computed with Fisher's z -transform, 5% significance.
Source: Own elaboration.

Figure 3 (cont.)

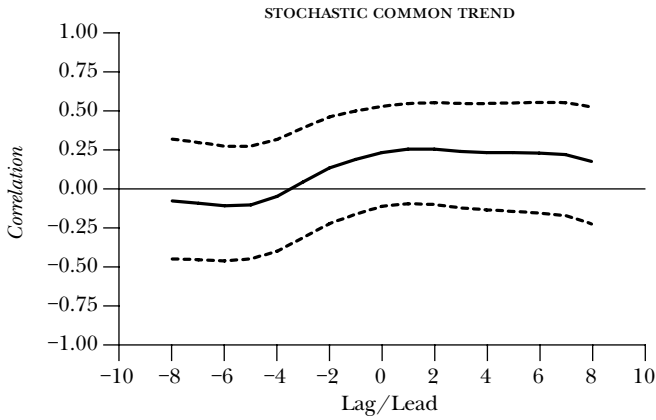
CROSS-CORRELATIONS OF RATE GAPS WITH OUTPUT GAP



Note: Confidence intervals computed with Fisher's z-transform, 5% significance.
Source: Own elaboration.

Figure 3 (cont.)

CROSS-CORRELATIONS OF RATE GAPS WITH OUTPUT GAP

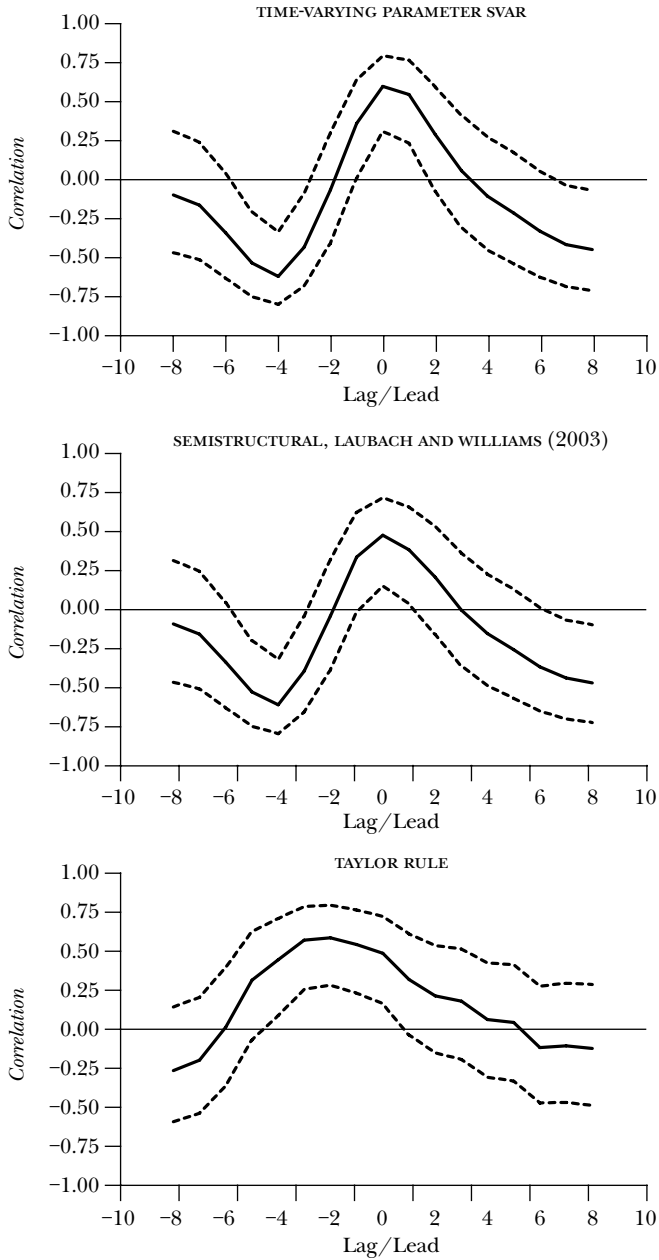


Note: Confidence intervals computed with Fisher's z-transform, 5% significance.
Source: Own elaboration.

Finally, the interest rate gap resulting from both SVAR models, from the Laubach and Williams (2003) model and from the local linear trend model show correlations with the FCI that exhibit the expected pattern: the highest correlations are positive and contemporaneous or indicative of a slight lead in the correlation of the interest rate gap and the FCI (Figure 5).

Figure 4

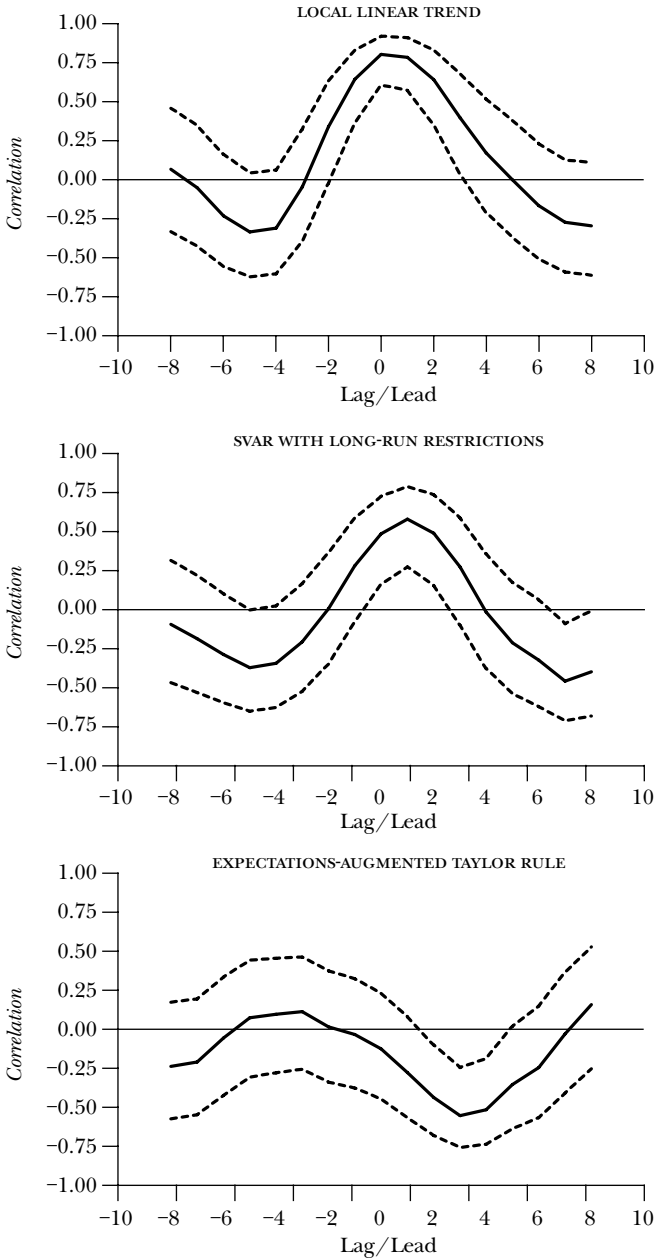
CROSS-CORRELATIONS OF RATE GAPS WITH TRIMMED-MEAN INFLATION RATE



Note: Confidence intervals computed with Fisher's z-transform, 5% significance.
Source: Own elaboration.

Figure 4 (cont.)

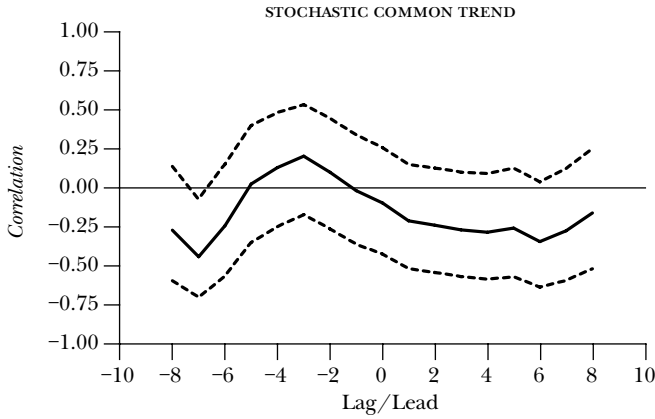
CROSS-CORRELATIONS OF RATE GAPS WITH TRIMMED-MEAN INFLATION RATE



Note: Confidence intervals computed with Fisher's z-transform, 5% significance.
Source: Own elaboration.

Figure 4 (cont.)

CROSS-CORRELATIONS OF RATE GAPS WITH TRIMMED-MEAN INFLATION RATE



Note: Confidence intervals computed with Fisher's z-transform, 5% significance.
Source: Own elaboration.

The analysis in this section suggests that estimates from the Taylor rules and the stochastic common trend do not have desirable empirical properties. Thus, the final NRIR indicator, which is shown in Figure 6 is computed as the simple average of the rest of estimates. During most of the period of study, the NRIR has fluctuated in the 0%-2% range, and its value at the end of the sample is 1.54%, which sits in the reference range of recent studies like Muñoz and Rodríguez (2016) and OECD (2016).

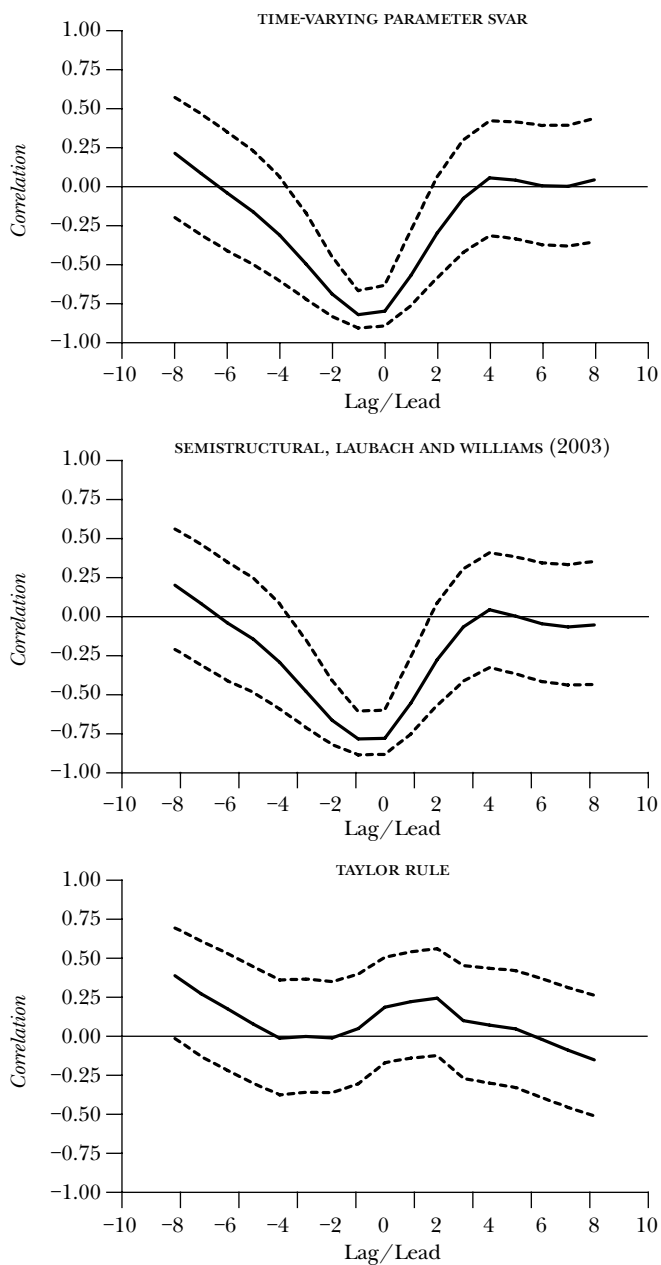
From the NRIR indicator an interest rate gap was computed, shown in Figure 7. This indicator shows a negative correlation with the output gap 1 quarter later, a negative and contemporaneous correlation with the FCI, and sizable correlations both with future core inflation (negative sign) and contemporaneous core inflation (positive sign). See Figure 13 in the Appendix⁹.

A noteworthy characteristic which can be inferred for all the period of analysis is that policymakers seem to have responded mostly to movements of inflation not associated with temporary shocks.

⁹ Besides, unit root tests suggest that the interest rate gap is stationary. See table 3 in the Appendix.

Figure 5

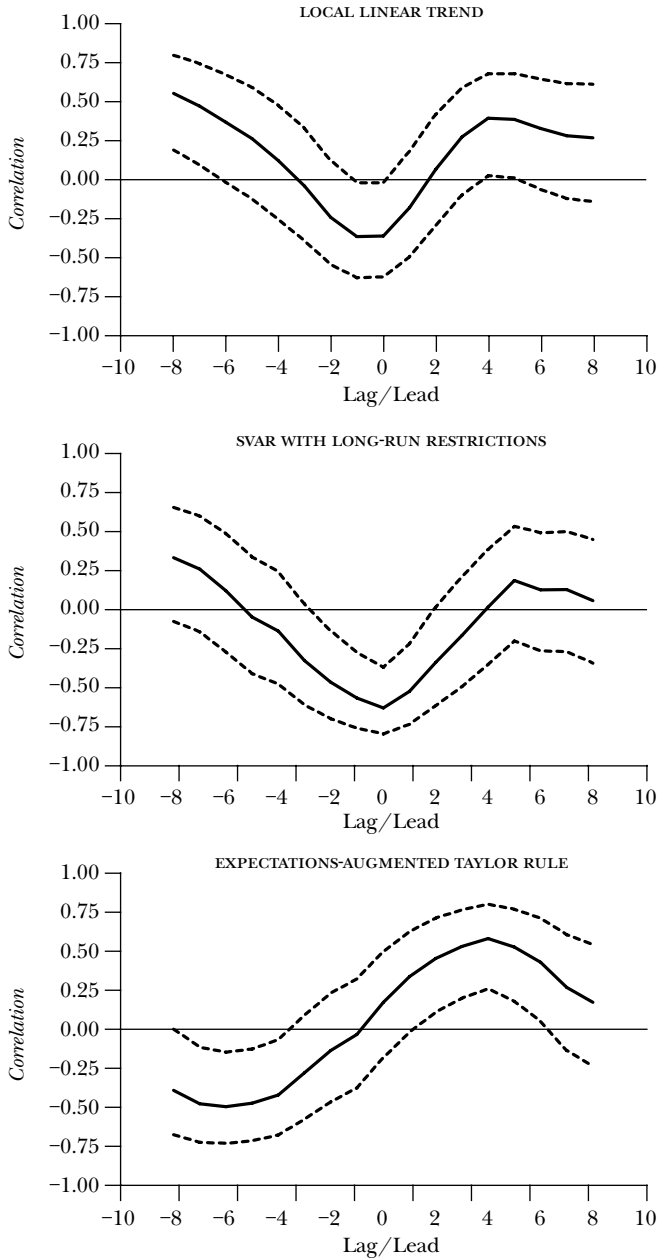
CROSS-CORRELATIONS OF RATE GAPS WITH FINANCIAL CONDITIONS INDEX



Note: Confidence intervals computed with Fisher's z-transform, 5% significance.
Source: Own elaboration.

Figure 5 (cont.)

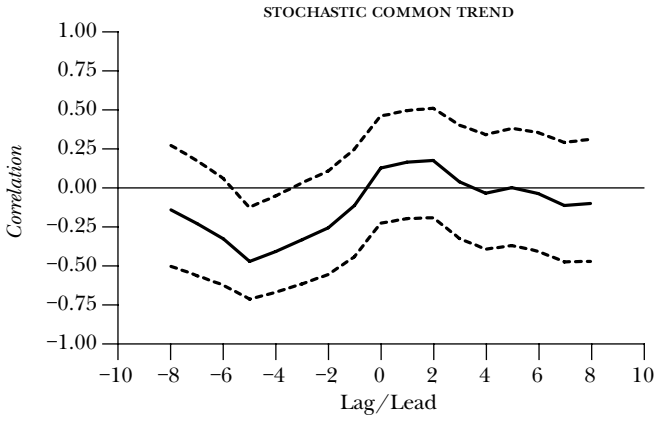
CROSS-CORRELATIONS OF RATE GAPS WITH FINANCIAL CONDITIONS INDEX



Note: Confidence intervals computed with Fisher's z-transform, 5% significance.
Source: Own elaboration.

Figure 5 (cont.)

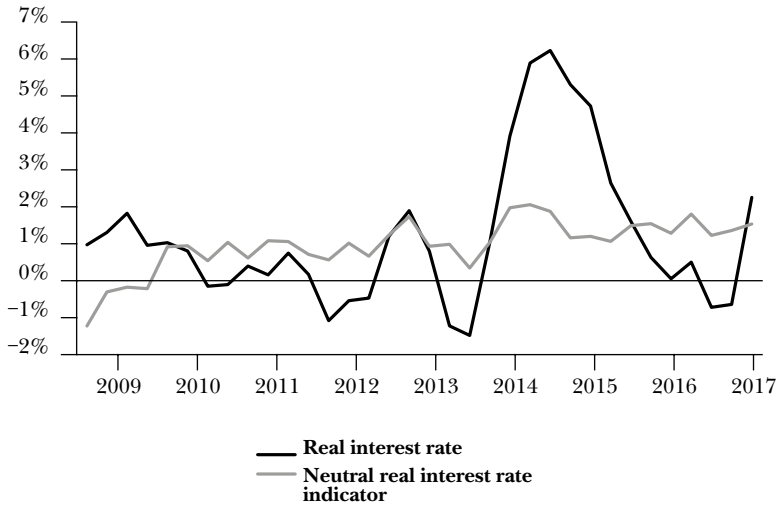
CROSS-CORRELATIONS OF RATE GAPS WITH FINANCIAL CONDITIONS INDEX



Note: Confidence intervals computed with Fisher's z-transform, 5% significance.
Source: Own elaboration.

Figure 6

NEUTRAL REAL INTEREST RATE INDICATOR FOR COSTA RICA



Source: Own elaboration.

To see this, first note that the interest rate gap shows its highest (positive) contemporaneous correlation with core inflation rather than CPI inflation (Figure 8). Furthermore, there seems not to be a sizable correlation between the interest rate gap with movements of the CPI inflation outside the target range when core inflation has been within its bounds (Figure 12 in the Appendix).

3.2 Monetary policy in Costa Rica 2009-2018

In order to assess whether the indicator of interest rate gap coherently reflects the monetary policy stance of Costa Rican policymakers, in this section a brief review of the evolution and characteristics of the monetary policy between 2009 and 2018 is presented. This period includes several policy decisions aimed at strengthening the ability of BCCR to fulfill its objective of maintaining a low and stable inflation, taken as part of a modernization process for monetary policy started in 2005.

The beginning of the period of analysis coincides with a time of expansionary monetary policy in several regions of the world, especially in the USA and the Euro area, which was a response to the financial crisis of 2008. This global excess of liquidity generated capital inflows to emerging economies like Costa Rica. As a consequence, Costa Rican economic agents had access to international financing at historically low interest rates. Besides, the need to finance the fiscal deficit was reflected in an upwards pressure on local interest rates, which in turn generated an increase in the Costa Rica premium¹⁰.

Additionally, between 2012 and 2015, the Costa Rican government issued the equivalent of 8% of its 2014 GDP in foreign debt bonds. At the same time, domestic and foreign-denominated credit to the private sector grew at a rate of 14.2% between 2009 and 2015 (see Figure 9).

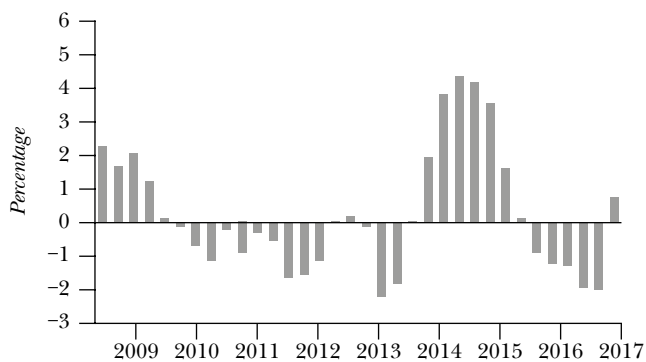
This capital inflow, under a crawling band regime¹¹, forced the BCCR to intervene to defend the lower limit of the band and allowed it to pursue several programs of reserve accumulation aimed at strengthening the economy in the face of external shocks. At the same time, the BCCR had to sterilize its intervention through open-market operations to control the risk the monetary expansion posed for future inflation. Liquidity

¹⁰ All of this with a low or null foreign-exchange risk, since the exchange rate stood at the lower limit of the crawling band for a long period.

¹¹ After more than 20 years of a crawling peg regime, in October 2006 Costa Rica adopted a crawling band regime which would last until February 2015, when a managed float was introduced.

Figure 7

REAL INTEREST RATE GAP FOR COSTA RICA

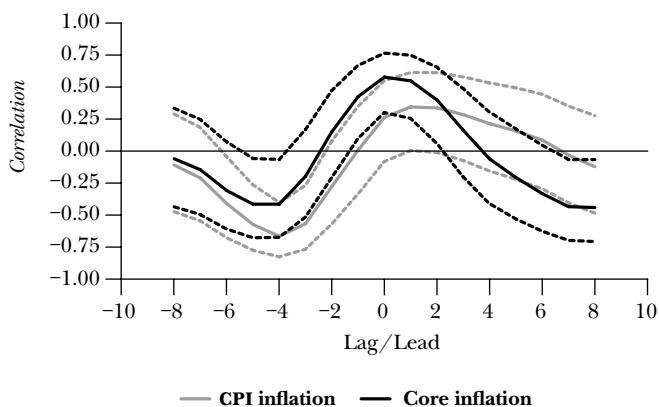


Source: Own elaboration.

Figure 8

**CROSS CORRELATION OF REAL INTEREST RATE GAP
WITH INFLATION VARIABLES**

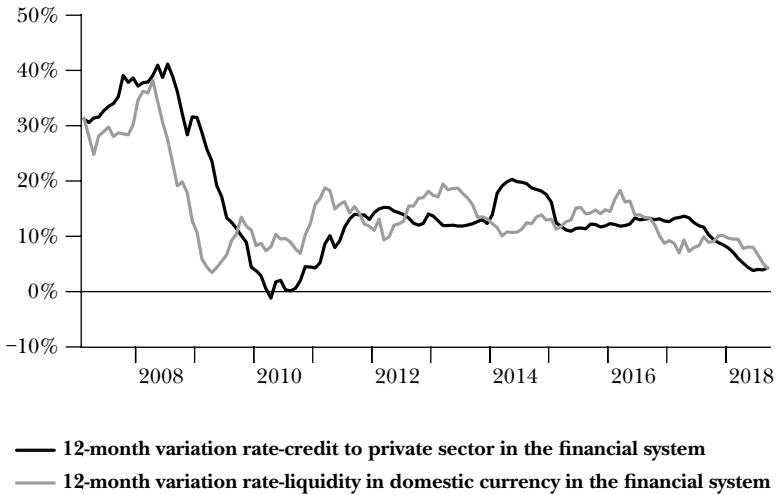
Trimmed-mean Inflation and CPI Inflation



Note: Confidence intervals computed with Fisher's z-transform, 5% significance.
Source: Own elaboration.

Figure 9

**CHANGE IN CREDIT TO THE PRIVATE SECTOR
AND IN LIQUIDITY IN DOMESTIC CURRENCY**



in domestic currency grew at an average rate of 20%, year-over-year, during the 2007-2009 period (see Figure 9).

During the first quarter of 2014 Costa Rica experienced a foreign exchange shock that increased the exchange rate 13.3% in 7 weeks. This event generated an increase in the inflation and exchange rate variation expectations. As a response to that shock, to avoid that inflation rose above the upper limit of the target range, the BCCR adjusted its policy rate from 3.75% to 5.25%. This contractionary policy stance was reflected during 2014 in the interest rate gap as positive values in Figure 10. In spite of this, inflation was outside the target range between July and December, returning to the target range in January 2015.

In 2014-2015 the fall in the international price of several commodities, especially oil, resulted in a positive terms-of-trade shock which lowered imported inflation and eventually contributed to a decrease in overall inflation. The policy response from BCCR comprised

communication with the population to explain the transitory nature of the shock, along with gradual decreases in the monetary policy rate. During eleven months, between February 2015 and January 2016, eight adjustments to the rate were applied amounting to 350 b.p., leading this indicator to 1.75%. Coincidentally, from 2015Q2 until 2017Q3 inflation was below the target range, even showing negative values. In spite of this, the interest rate gap suggests that a contractionary monetary policy stance was in place until the end of 2015, which leads to question whether the cuts in the monetary policy rate could have been done less gradually and sooner.

The monetary policy rate was held at 1.75% for 15 months, from January 2016 until April 2017. During that time, international rate began to rise, which translated into a lower, even negative, premium for saving in domestic currency. This, along with a high and persistent fiscal deficit, brought incentives for economic agents to increase their relative preference for saving in dollars and at the same time to de-dollarize their debts.

In consequence, the BCCR faced pressures in the FX market¹², which in turn translated into expectations of higher exchange rate variation that could pass on to inflation expectations and ultimately to domestic prices. In order to restore the premium for saving in domestic currency, it was decided to implement a gradual increase in the monetary policy rate, up to 300 b.p., between April and November 2017. However, the interest rate gap for this period still suggests a loose monetary policy stance during the first and second quarters of 2017 (Figure 10).

4. CONCLUSIONS

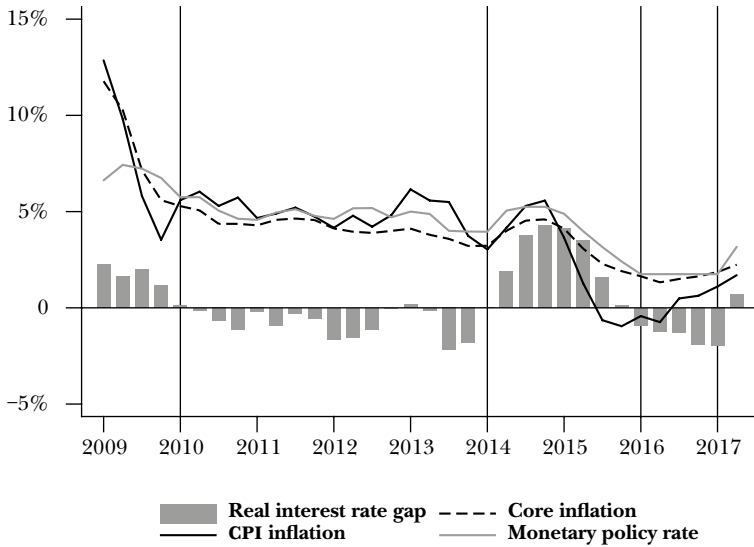
This study assesses the monetary policy stance for Costa Rica during the period 2009-2018 using an indicator of the real interest rate gap. This indicator is computed from a neutral real interest rate based on four estimates for which empirical coherence is evaluated. The value of the neutral real interest rate for Costa Rica is estimated at 1.54%, in line with previous estimates.

The interest rate gap shows expected properties: negative correlation with future output gap, negative contemporaneous correlation

¹² Several increases of the exchange rate were identified that were not coherent with the long-run trajectory set by its fundamentals.

Figure 10

**REAL INTEREST RATE GAP, INFLATION AND POLICY RATE
2009Q1-2018Q2**



Source: Own elaboration.

with the Financial Conditions Index, and sizable negative correlation both with future and current core inflation. This suggests that the use of the monetary policy rate has been successful in influencing the marginal cost of liquidity for financial intermediaries. This is a primordial factor for the correct functioning of an inflation targeting regime. However, a formal evaluation of the forecasting and modeling properties of these NRIR estimates was not the goal of this study.

The analysis of the monetary policy stand in Costa Rica suggests that policy has responded mostly to movements in inflation not related to temporary shocks. Besides, there have been cases in which it can be argued that adjustments in the policy stance could have been swifter.

Finally, the new NRIR and interest rate gap indicators are a valuable input in the ongoing process of improvement of the macroeconomic modelling at the Central Bank of Costa Rica.

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APPENDIX

Table 2

NRIR ESTIMATES FOR COSTA RICA 2007-2016					
	<i>Muñoz and Tenorio (2007)</i>	<i>Segura and Vindas (2012)</i>	<i>IMF (2016)</i>	<i>OECD^{a/} (2016)</i>	<i>Muñoz and Rodríguez (2016)</i>
Period of study	<i>2001- 2006</i>	<i>2001-2011</i>	<i>Several periods</i>	<i>2008-2015</i>	<i>Several periods</i>
Periodicity	<i>Quarterly</i>	<i>Monthly</i>	<i>Quarterly</i>	<i>Monthly</i>	<i>Monthly</i>
<i>Method</i>					
Ad-hoc (observed)	2,8%	2,0%			
Hodrick-Prescott filter	2,6%	1,9%			
Semistructural model	3,1%	2,2%	1,4%		0,6%
Structural VAR		1,3%			
Interest rate parity	3,1%	n.d.	1,6%		
Dynamic Taylor rule					1,2%
Expectations-augmented Taylor rule			2,4%	1,6%	3,0%
General equilibrium model			2,6%		
Monetary model			1,8%		
Stochastic common trend					1,2%
<i>Average</i>	2,9%	1,9%	1,9%	1,6%	1,5%

a/ Nominal rate is 4,6% in the original source, the real rate is presented after subtracting the inflation target.

Source: Own elaboration.

Table 3

UNIT ROOT TESTS FOR REAL INTEREST RATE GAP SERIES

	<i>Statistic</i>	<i>Probability</i>	<i>Critical value (1%)</i>
<i>Augmented Dickey-Fuller</i>			
No constant ^{1/}	0.0000	0.0000	
Constant ^{1/}	-3.0757	0.0386	
<i>Phillips-Perron</i>			
No constant ^{2/}	-2.1985	0.0289	
Constant ^{2/}	-2.1620	0.2232	
<i>Elliott-Rothenberg-Stock (DF-GLS)</i>			
Constante ^{1/}	-2.9384		-2.6392
<i>Kwiatkowski-Phillips-Schmidt-Shin (KPSS)</i>			
Constant ^{3/}	0.0816		0.7390

1/ One lag, automatic selection (Schwarz information criterion with a maximum of 8 lags considered)

2/ Bandwidth = 1 (Newey-West, Bartlett kernel).

3/ Bandwidth=4 (Newey-West, Bartlett kernel).

Source: Own elaboration.

Figure 11

NEUTRAL REAL INTEREST RATE ESTIMATES

2009I-2018II

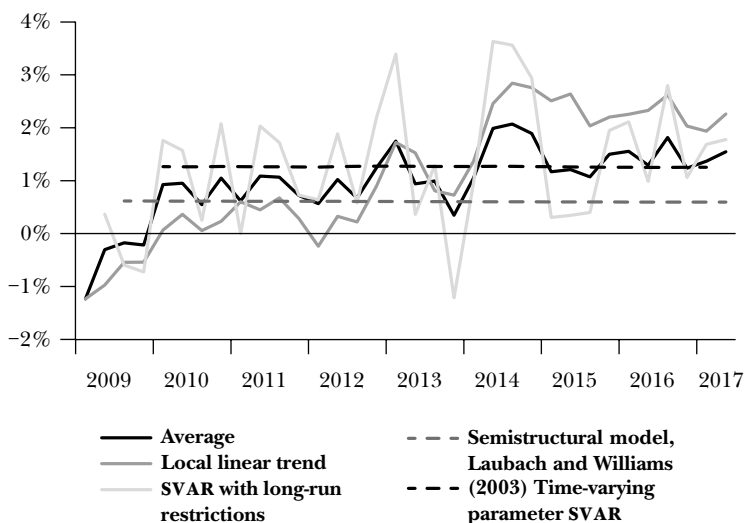


Figure 12

DEVIATIONS FROM TARGET OF CPI INFLATION
WHEN CORE INFLATION IS ON TARGET

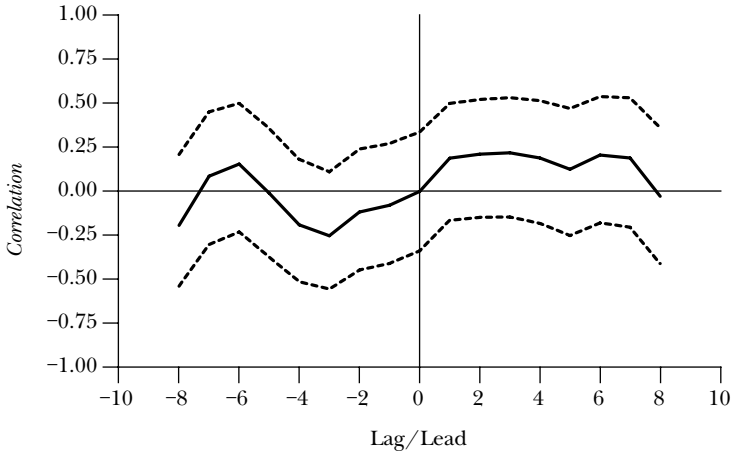
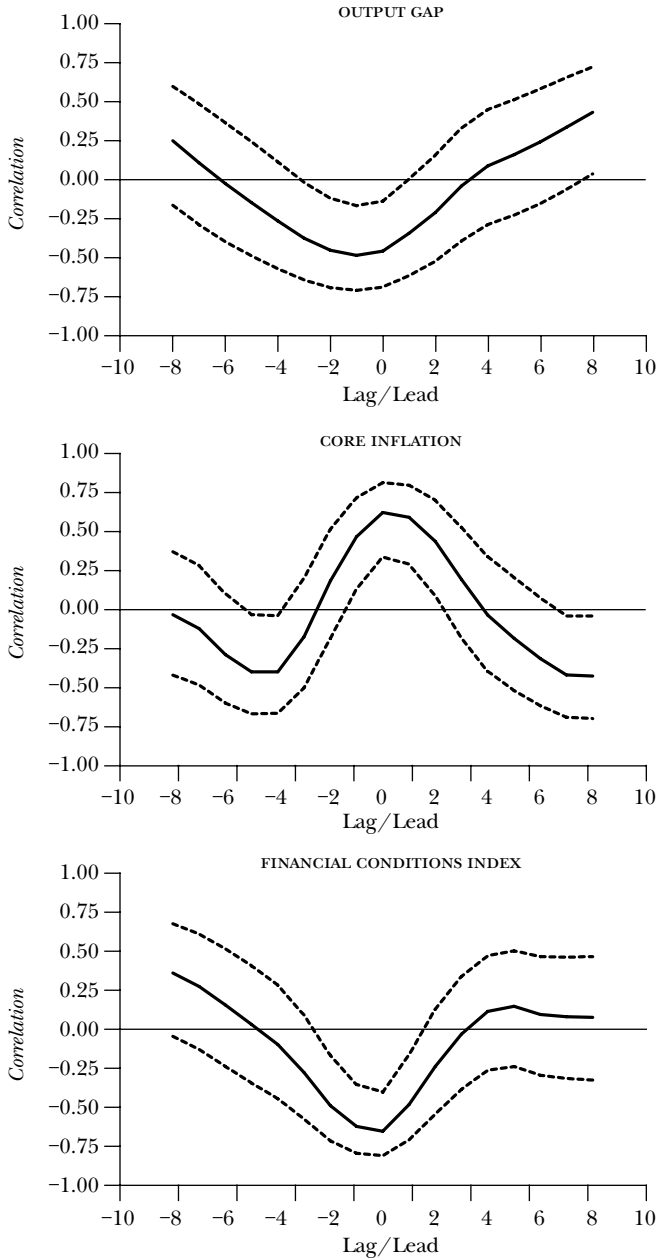


Figure 13

CROSS-CORRELATIONS OF REAL INTEREST RATE GAP WITH SEVERAL VARIABLES



Note: Confidence intervals computed with Fisher's z-transform, 5% significance.
Source: Own elaboration.