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INTERNATIONAL INTEREST RATE COMOVEMENTS AND MONETARY POLICY IN BRAZIL

COMOVIMENTOS DE TAXAS DE JUROS INTERNACIONAIS E POLÍTICA MONETÁRIA NO BRASIL°

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Abstract

This article shows that international spillovers or coordination in central bank interest rate setting are significant components of the Brazilian Taylor rule. For this, the short-term interest rate comovements of 28 countries and Euribor were estimated through a dynamic factor model. A nonlinear reaction function for the Central Bank of Brazil that includes global and regional factors evidenced that monetary policy in Brazil is influenced by the policy of other broad groups of countries and not just the United States and the Eurozone.

Keywords: comovements, dynamic factor model, Taylor rules, monetary policy.
JEL codes: E52, E58, C38.

Resumo

Este artigo mostra que spillovers internacionais ou a coordenação no ajuste das de taxas de juros dos banco centrais são componentes significativos da regra de Taylor brasileira. Para tal, são estimados os comovimentos das taxas de juro de curto prazo

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de 28 países e da Euribor através de um modelo de fatores dinâmicos. Uma função de reação não linear para o Banco Central do Brasil que inclui fatores globais e regionais fornece evidências de que a política monetária no Brasil é influenciada pela política de outros grupos amplos de países e não apenas dos EUA e da Zona do Euro.

Palavras-chave: comovimentos, modelo dinâmico de fatores, regras de Taylor, política monetária
Classificação *JEL*: E52, E58, C38.

INTRODUCTION

Central banks usually implement their monetary policy by managing short-term interest rates. Changes in the monetary policy of some economies, particularly the advanced ones, may influence that of other countries (Taylor, 2014), mainly in the case of emerging nations such as Brazil. The resulting monetary policy comovements are the subject of studies on interdependence, coordination, spillover, and/or contagion of interest rates between countries.

According to Caceres et al. (2016), in small open economies, some financial variables tend to move together with those that prevail abroad. There is evidence that emerging nations that manage their exchange rate fluctuations are more subject to foreign monetary policy influence (Edwards, 2015; Obstfeld, 2015; Aizenman et al., 2016; Rohit & Dash, 2019).

Some papers estimated the interest rate comovements in a group of countries and identified the factors that contribute to the coordination or interdependence of monetary policies between economies (Lindenberg & Westermann, 2012; Arouri et al., 2013; Chatterjee, 2016). Other studies focused on the transmission mechanisms of monetary policy to other nations, considering the effects of monetary adjustments from developed countries to emerging ones (Canova, 2005; Gray, 2013; Takáts & Vela, 2014; Edwards, 2015; Potjagailo, 2017; Anaya et al., 2017; Ratti & Vespignani, 2019).

In this study, we decomposed the common fluctuations among the interest rates of 28 countries and the Eurozone (Euribor), observed between 1996 and 2015[1], into a common global factor and common regional factors. For this, we used a dynamic factor model with multiple levels, as proposed by Kose et al. (2003, 2008).

At this point, there is a similarity to the work of Chatterjee (2016), who applied a dynamic factor model to examine monetary policy movements among the G5 countries. The author used the residuals of the Taylor rules adjusted to the nations to extract a factor based on the five countries in the sample and a second factor considering only Germany and the United Kingdom.

The present work analyzed the interest rate, and not unanticipated monetary policy shocks, in a larger group of countries covering the main continents. This allowed a more comprehensive analysis of the possible coordinations and interde-

[1] During data collection, we chose this period to allow for a broader set of countries used to estimate the global comovement of interest rates.

pendencies between monetary policies, given that central banks observe interest rates in other countries and perhaps not their unexpected movements.

After estimating the factors, we examined how common fluctuations affect the implementation of monetary policy in Brazil, a developing economy with strong exchange rate management, an increasing participation in international markets, and a monetary policy based on inflation targeting since 1999. These are the reasons why this economy was chosen for this analysis. In this step, we used the methodology proposed by Bai and Perron (1998, 2003) to estimate a nonlinear reaction function for the Central Bank of Brazil (BCB, as per its initials in Portuguese), whose parameters may vary due to the existence of (multiple) structural breaks in the implementation of the Brazilian monetary policy.

The results show that the factors capture the main economic events (financial crises) and explain on average 28% of the total variance of interest rates in all the economies. The nonlinear Taylor rule for Brazil evidences that common factors play an important role in the implementation of monetary policy by the BCB.

As a robustness analysis, we estimated other models with several indicators of global economic activity and a global inflation variable, to control for central bank reactions to common shocks to economic activity and the price level. In general, the results regarding the dates of the breaks and the statistical significance of the coefficients did not change substantially.

In other words, there may be interdependence between central bank decisions, even in periods without global shocks, and considering this possibility may provide better specification reaction functions and predictions about the interest rate.

The main contributions of this study are twofold. The first is the estimation of the commonalities in the implementation of monetary policy through a database comprising a large number of countries. This allows analyzing interdependence more broadly and thus observing that this phenomenon can occur between neighbouring economies or distant ones, and even among unexpected groups of countries. Second, to date there are no studies that include these measures of common monetary policy fluctuations in Taylor rules.

Besides this introduction, the paper is organized into four other sections. Section I contains a brief literature review. Section II describes the econometric method and the database. Section III presents the results. Lastly, our final considerations are provided.

I. COMOVEMENTS AND INTERDEPENDENCE OF MONETARY POLICY: EMPIRICAL EVIDENCE

The comovement of economic variables is a widely studied topic, either by empirical applications or by theoretical explanations. Interdependence, contagion, and spillover effects are some similar terms related to this subject.

Forbes and Rigobon (2001) defined contagion as the propagation of market disturbances from one country to another, as can be observed in some variables such as exchange rates and stock prices. There is a distinction between contagion and interdependence. The former specifically applies to financial crises, while the latter is characteristic of periods of economic stability (Forbes & Rigobon, 2001, 2002). According to Pesaran and Pick (2007), spillover effects are examples of interdependent movements.

For his part, Canova (2005) used a vector autoregressive (VAR) model and found substantial effects of US monetary policy on eight Latin American countries. According to the author, the contractionary monetary policy shocks in the United States induce a significant and instantaneous increase in interest rates in the Latin American countries in the sample.

Furthermore, Jannsen and Klein (2011) reported that monetary policy shocks in the Eurozone cause significant and proportional effects on interest rates and output in five Western European countries that are not part of the Eurozone. In addition, Potjagailo (2017) used a larger set of European countries and showed the existence of spillover effects of monetary policy shocks from the Eurozone to the 14 European countries, not included in that area.

Through tests for common serial correlation, cointegration, and codependence, Lindenberg and Westermann (2012) studied the dynamic behavior of short- and long-term interest rates for the G7 countries, using quarterly data from 1975 to 2010. Among their results, there is limited empirical evidence on cyclical comovements and a strong co-dependence among long-term interest rates of Italy, France, and the United Kingdom for the subsample during the period before the formation of the Eurozone (1979.01 – 1998.04).

Morever, using an international business cycle model and a sample of the six largest industrialized economies[2] between 1960 and 2006, Henriksen et al. (2013)

[2] Australia, Canada, Germany, Japan, the United Kingdom, and the United States.

found that fluctuations in aggregate price levels and nominal interest rates are more synchronized than those in output.

For their part, Arouri et al. (2013) investigated the synchronization and interdependence among short-term interest rates of France, the United Kingdom, and the United States during the period from 2005 to 2009. Using bi-directional feedback measures proposed by Geweke (1982) and smooth transition error correction models (STECM), the authors evidenced contemporary interdependence and an increase in the synchronization of the monetary policies of these countries in the period analyzed.

On the other hand, Chatterjee (2016) applied a dynamic factor model to examine monetary policy comovements among the G5 countries. The author used the residuals of the Taylor rules adjusted to the countries to extract two latent factors. The first one (global) considers all the nations in the sample, and the second one (European), only Germany and the United Kingdom. According to the author, the global factor captures the main economic events that occurred during the period analyzed and accounts for an average of 24% of the variation in monetary policy not explained by Taylor rules.

In addition, Takáts and Vela (2014) affirmed that the central banks of emerging economies adjust their interest rates according to changes in those of advanced countries. The results of Anaya et al. (2017) indicate that short-term interest rates of emerging economies tend to decline in response to shocks from unconventional expansionary monetary policy of the United States. In other words, innovation in US monetary policy strongly influences the implementation of monetary policy in the emerging countries analyzed. Several other authors reported similar results using different methodologies (Miranda-Agrippino & Rey, 2020; Passari & Rey, 2015; Rey, 2015, 2016; Aizenman et al., 2016; Ratti & Vespignani, 2019).

Lastly, Rohit and Dash (2019) used autoregressive vectors and the spillover index of Diebold and Yilmaz (2009) in a sample of five advanced and eight emerging economies. Their findings show that developing countries with a less flexible exchange rate are more subject to foreign monetary policy influences.

II. METHODOLOGY AND DATA

a. The Dynamic Factor Model with Multiple Levels

To decompose the fluctuations of the short-term interest rates of a sample of 28 countries plus the interest rate of the Eurozone into common movements associated with global and regional fluctuations and specific idiosyncratic movements of each country, we used a dynamic factor model and Bayesian techniques.

According to Stock and Watson (2010), there are three theoretical advantages or motivations for using Bayesian methods to estimate dynamic factor models. First, the method to compute the posterior distribution of the parameters and the factors can be more stable than those that maximize the likelihood function, mainly when there are many unknown parameters. Second, Markov Chain Monte Carlo (MCMC) methods can calculate the posterior distribution in models with latent variables that are not normally distributed or are non-linear, where it is extremely difficult to compute the likelihood directly. And third, it is possible to impose prior information on the parameters and factors.

In this respect, Otrok and Whiteman (1998)[3] developed a Bayesian approach to the dynamic factor model with only one factor. Following Tanner and Wong (1987), this approach is used to construct artificial observations or an unobservable indicator (latent factor) via data augmentation methods. Samples from the posterior distribution of the dynamic factors and relevant parameters are obtained by MCMC methods.

Moreover, Kose et al. (2003, 2008) proposed an extension of the Bayesian procedure of Otrok and Whiteman (1998). They used dynamic factor models with multiple levels to analyze international business cycles. In these models, geographic or economic characteristics (Kose et al., 2012) are used to identify economic oscillations, giving economic meaning to the estimated factors[4].

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3 They constructed a coincident indicator to measure economic activity based on four variables for the state of Iowa.

4 Crucini, Kose and Otrok (2011), Neely and Rapach (2011), and Moench, Ng and Potter (2013) are examples of other studies that have used this method.

In general, factor models (static or dynamic) assume that the variables under consideration are stationary. For some countries[5], the null hypothesis of a unit root is not rejected, and we took the first difference of this variable for all economies.

Following the approach proposed by Kose et al. (2003) and Neely and Rapach (2011), we considered that the variation of the interest rate of the th country in period follows a structure of common factors, expressed as:

$$\Delta r_{i,t} = \lambda_i^G f_t^G + \lambda_i^R f_{j,t}^R + \varepsilon_{i,t} \tag{1}$$

In Equation (1), the evolution of the observed variable $\Delta r_{i,t}$ depends on two common latent factors, f_t^G and $f_{j,t}^R$, plus an idiosyncratic component $\varepsilon_{i,t}$ with $E\left(\varepsilon_{i,t}, \varepsilon_{k,t-s}\right) = 0$ for $i \neq k$. The global factor f_t^G denotes the common fluctuations between the variations in the $N = 29$ short-term nominal interest rates considered. The regional factor $f_{j,t}^R$ represents the common fluctuations for each group of countries that belong to one of the j=1,2,…,5 pre-specified regions, categorized according to the common economic and geographic characteristics of the countries that form each region.

The term $\varepsilon_{i,t}$ is the idiosyncratic or specific component of the i −th observed unit, or the error term of the model, and captures purely national influences on the variations of nominal interest rates, or measurement errors of the dependent variables. The parameters λ are the factor loadings and represent the importance or weight of each factor in explaining the variation of the dependent variable.

It is assumed that f_t^G and $f_{j,t}^R$, are governed by an autoregressive process of order q:

$$f_t^G = \rho_1^G f_{t-1}^G + \rho_2^G f_{t-2}^G + \cdots + \rho_q^G f_{t-q}^G + u_t^G \tag{2}$$

$$f_{j,t}^R = \rho_{j,1}^R f_{j,t}^R + \rho_{j,2}^R f_{j,t}^R + \cdots + \rho_{j,q}^R f_{j,t-q}^R + u_{j,t}^R \tag{3}$$

The error terms in Equations (2) and (3) have normal distributions, $u_t^G \sim N(0, \sigma_G^2)$ and $u_{j,t}^R \sim N\left(0, \sigma_{j,R}^2\right)$, and are free of serial correlation, i.e., $E(u_t^G u_{t\ s}^G) = E\left(u_{j\ t}^R u_{j\ t\ \ s}^R\right) = 0$ for $\neq 0$. The error term in Equation (1) also has a normal distribution and follows an $AR(p)$ process:

[5] The tests proposed by Dickey and Fuller (1979, 1981) and Phillips and Perron (1988) do not reject the hypothesis of the presence of a unit root in the interest rate series of some countries. These results are available upon request.

$$\varepsilon_{i,t} = \rho_{i,1}\varepsilon_{i,t-1} + \rho_{i,2}\varepsilon_{i,t-2} + \cdots + \rho_{i,p}\varepsilon_{i,t-p} + u_{i,t}^{\varepsilon} \tag{4}$$

where:

$$E\left(u_{i,t}^{\varepsilon} u_{i,t-s}^{\varepsilon}\right) = \begin{cases} \sigma_i^2, \text{ if } i = k \text{ and } s = 0 \\ 0, \text{ otherwise} \end{cases} \tag{5}$$

The errors in Equations (2), (3), and (4) are not contemporaneously autocorrelated in any of their lags and leads, i.e., $E\left(u_{i,t}^{\varepsilon} u_{t-s}^{G}\right) = E\left(u_{i,t}^{\varepsilon} u_{j,t-s}^{R}\right) = E\left(u_{t}^{G} u_{j,t-s}^{R}\right) = 0$ for all i, j e s. Therefore, the global factor, the regional ones, and the idiosyncratic error terms of the model are orthogonal.

According to Kose et al. (2003, 2008), there are two identification problems: neither the signs nor the scales of factors and their loadings are separately identified. As a result, one of the factor loadings is constrained to be positive for each of f_t^G and $f_{j,t}^R$.

Following this strategy, we chose the most representative country of each region for this normalization[6]. Hence, the global factor loading for the United States and the regional factor loadings for the United States (North America), Brazil (Latin America), Eurozone (Europe), Japan (Asia), and Australia (Oceania) were restricted to being positive. The scales were identified by the strategy proposed by Sargent and Sims (1977) and Stock and Watson (1989, 1993), assuming that each variance of the error terms of the factors σ_G^2 and $\sigma_{j,R}^2$ for $j = 1,2,\cdots,5$ is equal to 1.

We simulated the posterior distribution for the parameters and factors of the model by successive draws from a series of conditional distributions through the MCMC procedure. The posterior distributions of the latent factors and the parameters of the model were obtained based on 100 000 MCMC replications, after a burn-in of 10 000 replications[7]. The order of the autoregressive processes[8] in Equations (2) – (4) is $p = q = 2$. For the prior distributions of the factor loadings, we assumed a multivariate normal distribution, given by:

$$\left(\lambda_i^G \ \lambda_{j\,i}^R\right) \sim N(0 \ I_2) \tag{6}$$

6 Similar results are obtained by changing the countries in the identification scheme considered.

7 Gibbs sampling requires choosing the initial values of the parameters to be estimated. To eliminate the dependence regarding this starting value, it is common to discard R_0 values (known as burn-in draws) by applying the Gibbs sampling method. In the MCMC procedure, the signs described previously are normalized after discarding the draws of the factor loadings that do not satisfy the constraints. In practice, inadmissible factor loadings are rarely drawn after the burn-in replications.

8 Similar results are obtained for different orders of p and q.

$$\left(\rho_{i,1}, \rho_{i,2}, \cdots, \rho_{i,p}\right)' \sim N\left(0; diag(1 \quad 0.5 \quad \cdots \quad 0.5^{p-1})\right) \tag{7}$$

$$\left(\rho_1^G, \rho_2^G, \cdots, \rho_q^G\right)' \sim N\left(0; diag(1 \quad 0.5 \quad \cdots \quad 0\ 5^{p-1})\right) \tag{8}$$

$$\left(\rho_{j,1}^R, \rho_{j,2}^R, \cdots, \rho_{j,q}^R\right)' \sim N\left(0; diag(1 \quad 0.5 \quad \cdots \quad 0.5^{p-1})\right) \tag{9}$$

The prior distributions in Equations (6) – (9) are similar to those employed by Kose et al. (2003, 2008) and Neely and Rapach (2011). The prior distribution for the variances of the error terms of Equation (1) follows an inverse gamma distribution $\sigma_i \sim IG(6; 0.001)$. The orthogonality between the factors and the error terms allows the following decomposition of the variance:

$$var\left(\Delta r_{i,t}\right) = \left(\lambda_i^G\right)^2 var(f_t^G) + \left(\lambda_{j,i}^R\right)^2 var\left(f_{j,t}^R\right) + var\left(\varepsilon_{i,t}\right) \tag{10}$$

The contribution of the global factor to the total variability is given by:

$$\theta_i^G = \frac{\left(\lambda_i^G\right)^2 var(f_t^G)}{var\left(\Delta r_{i,t}\right)} \tag{11}$$

Analogously, the relative contributions of the regional factors and the idiosyncratic error terms to the variance of the observed variable are defined, respectively, by the following expressions:

$$\theta_i^R = \frac{\left(\lambda_{j,i}^R\right)^2 var\left(f_{j,t}^R\right)}{var\left(\Delta r_{i,t}\right)} \tag{12}$$

$$\theta_i^C = \frac{var(\varepsilon_{i,t})}{var(\Delta r_{i,t})} \tag{13}$$

The decomposition in Equations (11) – (13) are functions of the model parameters and data, and are calculated in each iteration of the MCMC algorithm, which extracts from the respective posterior distributions the statistics necessary for every replication for each country. A high dispersion in the posterior distributions indicates uncertainty regarding their magnitudes[9].

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[9] By construction of the model, the factors are not autocorrelated. However, Kose et al. (2003) argued that the samples taken at each iteration of the MCMC algorithm will be autocorrelated purely due to sampling errors. To assure that the sum of the proportions of the variance explained by the global

b. The Taylor Rule with Factors

The empirical literature shows that central banks, mainly from emerging countries, react to the monetary policy decisions of other economies, especially large developed ones[10]. According to Calvo and Reinhart (2002) and Edwards (2015), in the presence of nearly perfect capital mobility, an increase in global interest rates, for example, caused by monetary policy adjustments of the US Federal Reserve, will result in an incipient external deficit, depreciation of the domestic currency, and monetary policy adjustments to reestablish equilibrium. However, if there is a fear of floating, local central banks will be likely to tighten their monetary policy as a way to avoid weakening the currency.

To check whether common fluctuations among short-term interest rates affect the adjustment of interest rates in Brazil, we included these variables as regressors in a Taylor rule. For this, we estimated the factors again, excluding Brazil from the group of countries, for the period from 2002 to 2015. The common factors assessed are used as proxies for the common global and regional fluctuations in monetary policy.

Based on Clarida et al. (1998), it is necessary to add the second lag of the dependent variable as an explanatory variable in the Taylor rule to control for possible problems of serial correlation, as well as to represent a greater smoothing of the central bank in the adjustment of the interest rates, as shown in the equation below:

$$i_t = \rho_1 i_{t-1} + \rho_2 i_{t-2}(1-\rho_1-\rho_2)\left[\alpha_0 + \alpha_1\left(E_t\pi_{t+j} - \pi^*_{t+j}\right) + \alpha_2 y_{t-1} + \alpha_3 \Delta e_{t-1} + \alpha_4 f_t^G + \sum_{i=1}^{5} \alpha_{4+i} f_{i,t}^R\right] + \varepsilon_t \qquad (14)$$

and regional factors and the idiosyncratic component for each country of the sample is 1, we adopted the strategy of Kose et al. (2003), which consists of orthogonalization of the factors in order, using the order global factor–regional factor–national factor, to compute the decompositions of the variance for each replication. Since the correlations between the raw factors are small, the order of orthogonalization has little impact on the results. All the results remain qualitatively the same under alternative orderings, and the quantitative differences are small.

10 Taylor (2007), Gray (2013), and Edwards (2015) included foreign interest rates in monetary policy rules to verify possible spillover effects and reported statistically significant coefficients of these estimated relations. Modenesi et al. (2013) calculated Taylor rules for Brazil between 2000 and 2010 including Libor as a proxy for the international interest rate in the original BCB reaction function, finding a positive and statistically significant response of the local interest rate to variations in Libor in the estimated reaction functions.

In Equation (14), the dependent variable i_t is the Brazilian benchmark rate, the monthly overnight SELIC[11] rate (expressed as % per year), $E_t\pi_{t+j}$ is a proxy for inflation expectations[12], π^*_{t+j} is the inflation target established by the Monetary Policy Committee (COPOM, as per its initials in Portuguese), y_{t-1} is an estimate of the output gap, obtained by applying the Hodric-Prescott filter to the seasonally adjusted industrial output index for Brazil in the period $t-1$, Δe_{t-1} is the 12-month variation of the nominal exchange rate with the dollar (R$/US$) in $t-1$, and f_t^G and $f_{i,t}^R$, for $i = 1,2,\cdots,5$, are estimates of the global and regional fluctuations[13].

The measure of the deviation of inflation expectations from the target is the one suggested by Minella et al. (2003), given by the weighted average between the deviations of expected inflation based on inflation targets for the years t and $t+1$:

$$D_j = \frac{12-j}{12}\left(E_j\pi_t - \pi_t^*\right) + \frac{j}{12}\left(E_j\pi_{t+1} - \pi_{t+1}^*\right) \tag{15}$$

Where j indexes the month and t, the year. Substituting Equation (15) in (14), we obtain:(15)

$$\begin{aligned} i_t = \rho_1 i_{t-1} + \rho_2 i_{t-2} + \gamma_0 + \gamma_1 D_j + \gamma_2 y_{t-1} + \gamma_3 \Delta e_{t-1} + \gamma_4 f_t^G \\ + \sum_{i=1}^{5} \gamma_{4+i} f_{i,t}^R + \varepsilon_t \end{aligned} \tag{16}$$

Where $\gamma_i = (1-\rho_1-\rho_2)\alpha_i$ for $i = 1,2,\cdots,9$. The coefficients in Equation (16) are the short-term parameters (Aizenman et al., 2011). The long-term coefficients α_i of the Taylor rule can be calculated by $\alpha_i = \frac{\gamma_i}{(1-\rho_1-\rho_2)}$.

c. Data Base

We used the money market interest rates of 28 countries plus the Euribor[14]. Short-term interest rates (money market) capture not only the temporal dynamics and

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11 The Selic rate, or "over Selic", is the Brazilian federal funds rate. Precisely, Selic rate is the weighted average interest rate of the overnight interbank operations - collateralized by federal government securities - carried out at the Special System for Settlement and Custody (Selic).

12 In this analysis, we used the average expectation for the next 12 months (% p/y) – obtained from the BCB (Focus survey).

13 The tests of Dickey and Fuller (1979, 1981) and Phillips and Perron (1988) reject the null hypothesis of the presence of a unit root in the variables of (14).These results are available upon request.

14 The inclusion of the Euribor reflects the behavior of the European Central Bank in adjusting its inter-

the magnitude of changes in interest rates adjusted by central banks (Crucini et al., 2011) but also market expectations for this policy instrument (Arouri et al., 2013).

Data were collected for the period from January 1996 to December 2015[15] from the International Financial Statistics of the International Monetary Fund (IFS – IMF). We chose this period to allow for a broader set of countries used to estimate global interest rate movements. The descriptive statistics presented in Table 1 show great heterogeneity in the average and standard deviations of the interest rates of the analyzed economies.

Table 1. Descriptive Statistics: Short-Term Nominal Interest Rates (% p/y)

Region	Country	Mean	Max	Min	Std. Dev.	Obs.
North America	Canada	2.70	5.8	0.24	1.668	240
	United States	2.57	6.54	0.07	2.317	240
Latin America	Argentina	11.49	91.19	1.2	11.838	240
	Bolivia	6.24	25.14	0.4	5.309	240
	Brazil	16.66	43.25	7.11	7.344	240
	Mexico	11.16	42.93	3.29	8.986	240
	Venezuela	9.17	49.2	0.1	8.917	240
	Uruguay	15.74	119.45	0.69	19.677	240
Europe	Denmark	2.56	6.11	-0.69	1.784	240
	Eurozone	2.37	5.81	-0.43	1.758	240
	Finland	2.42	5.113	-0.13	1.535	240
	Iceland	8.71	37.76	3.75	4.433	240
	Ireland	2.64	6.55	-0.19	1.98	240
	Poland	8.78	25.55	1.5	7.057	240

est rates. Shaw et al. (2016) stressed the importance of the Euribor in the transmission of monetary policy and for the good functioning of the European money market.

15 Updated data are not available for the same group of countries used in this research, including information on the economies that were significant factors in our results.

	Romania	26.39	217.8	0.3	32.058	240
	Russia	11.51	139.7	1	16.116	240
	Spain	2.58	9.03	-0.12	2.062	240
	Sweden	2.71	8.76	-0.52	1.696	240
	Switzerland	0.82	3.5	-2	1.087	240
	United Kingdom	3.53	7.5	0.35	2.442	240
Asia	China (Hong Kong)	2.45	17.75	0.06	2.6	240
	China (Macau)	2.70	12.25	0.09	2.691	240
	Indonesia	12.49	81.01	3.76	14.279	240
	Japan	0.16	0.521	0	0.182	240
	South Korea	5.07	25.63	1.48	4.114	240
	Singapore	1.56	7.75	0.02	1.529	240
	Thailand	3.83	23.87	0.96	4.412	240
Oceania	Australia	4.79	7.519	2	1.384	240
	New Zealand	5.245	10	2.281	2.227	240

Source: Prepared by the authors based on data obtained from the IFS – IMF.

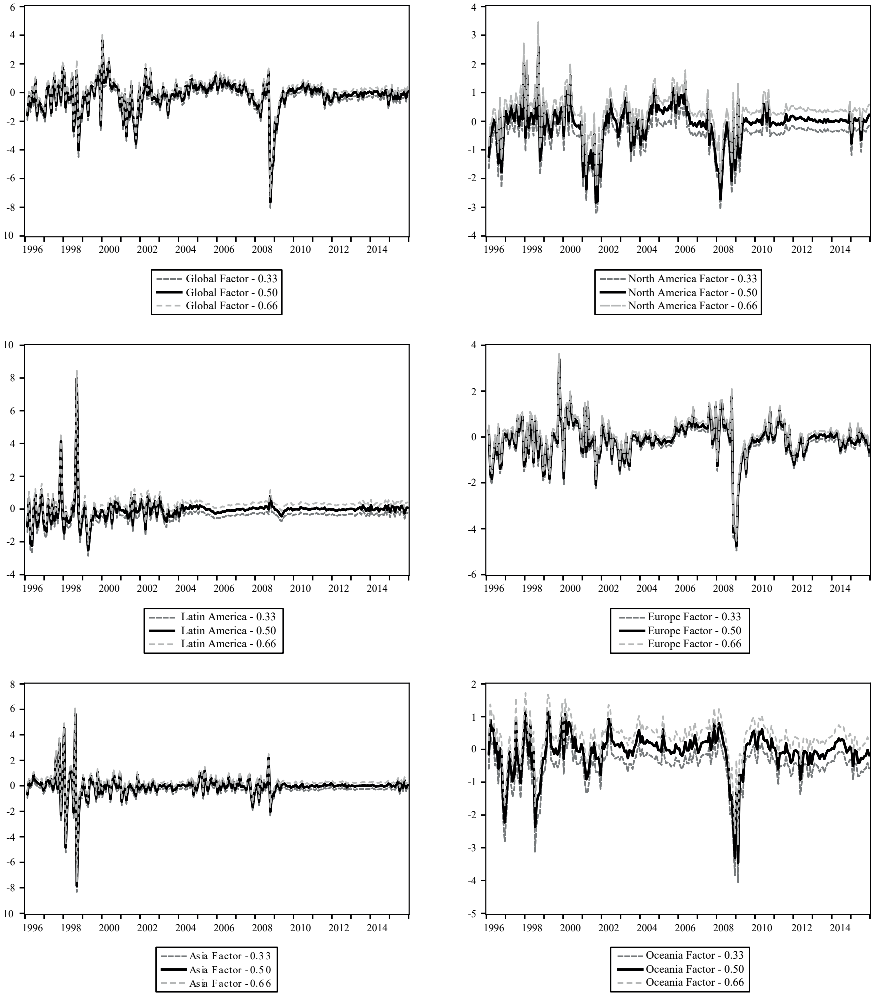
III. RESULTS

a. Results of the Dynamic Factor Model

Figure 1 shows the mean and the 0.33 and 0.66 quantiles for the posterior distributions of the global and regional factors, in the period from 1996.01 to 2015.12. The precision of the estimation is evidenced by the small distance between the mean and the quantiles[16].

[16] As a robustness analysis, the results are similar with different rankings in regional groupings and lags of p and q in Equations (1) to (4).

Figure 1. Global and Regional Factors, 1996.01 – 2015.12



Note: The solid lines denote the means, and the dotted lines represent the 0.33 and 0.66 quantiles of the posterior distribution of the global and regional factors.

The global factor captures most of the economic events that occurred during the two decades considered here. The peaks and troughs of the trajectory of the global factor coincide with the main economic and financial crises of this period. We highlight the currency crisis of Western Asian countries in 1997-1998, the

Russian debt crisis in 1998-1999, the collapse and ensuing bailout of the mega-fund manager Long-Term Capital Management in 1998, the dot-com stock market crash in 2000-2001, the Argentine crisis of 2001-2002, and the global financial crisis of 2008.

The trajectory of the global factor reveals relative instability in the period from the second half of the 1990s to the early 2000s, followed by a relatively stable period until the abrupt crash of 2008 during the global financial crisis that erupted that year.

The dynamics of the regional factors also evidences some economic events that had a direct impact on the path of interest rates in the countries of each regional group. For North America, there are peaks in the Russian financial crisis, the dot-com stock market collapse, and the 2008 financial crisis.

The peaks and troughs in the trajectory of the regional factor for Latin America are consistent with the Asian currency crisis (1997-1998), the Russian debt crisis (1998-1999), and the Brazilian currency crisis (1998-1999). This latent variable is unstable in the period between the second half of the 1990s and the first years of the next decade, reflecting the effects of the Argentine crisis (2001-2002). This is followed by a stable path with a small peak in 2008, in line with the global financial crisis.

The figure of the European regional factor coincide with the Russian, global, and European Union debt crises. For the Asian factor, it is possible to observe greater volatility between 1997 and 1999, in agreement with the Asian currency crisis of 1997-1998. After a period of relative stability of this factor, the 2008-2009 crash coincides with the global economic crisis of 2008.

Table 2 shows the results of the variance decomposition. For each one, the means and the 0.33 and 0.66 quantiles of their estimated posterior distributions are presented. In this analysis, the total period was divided into two sub-periods: 1996 – 2006 and 2007 – 2015. The first one covers the main financial crises that occurred in the period analyzed, while the second is characterized by the stability of the global economy until the world financial crisis of 2008 and the one of the European Union in 2011.

On average, the proportions and increase in the second sub-period, from 10.3% and 16.1% to 22.9% and 19.4%, respectively. The average contributions of the global and regional factors are higher in the second sub-period (42.3%) than in the first one (26.4%).

Table 2. Variance Decompositions for the sub-periods (in %)

|		Global Factor		Regional Factor		Idiosyncratic Component	
Region	Country	1996-2006	2007-2015	1996-2006	2007-2015	1996-2006	2007-2015
N. America	United States	33.69	32.72	16.09	6.12	50.22	61.16
	Canada	31.99	40.02	19.94	40.61	48.07	19.37
Latin America	Brazil	4.38	1.26	20.18	17.15	75.44	81.59
	Argentina	0.46	1.07	1.20	10.01	98.34	88.92
	Bolivia	0.54	0.18	2.76	13.11	96.70	86.71
	Mexico	12.55	1.57	28.50	18.62	58.95	79.82
	Uruguay	5.82	1.09	4.82	2.66	89.36	96.24
	Venezuela	0.47	4.56	1.49	6.63	98.04	88.82
Europe	Eurozone	27.35	69.51	62.31	7.29	10.34	23.20
	Denmark	18.55	5.19	20.62	37.69	60.83	57.12
	Finland	18.20	34.82	72.47	64.50	9.33	0.68
	Iceland	0.94	2.91	0.28	14.72	98.78	82.37
	Ireland	25.16	32.07	7.81	63.03	67.03	4.90
	Poland	1.87	5.07	0.32	3.68	97.81	91.25
	Romania	0.91	4.93	0.25	15.37	98.83	79.70
	Russia	15.90	0.17	4.18	0.58	79.93	99.25
	Spain	17.05	45.30	14.53	17.84	68.42	36.86
	Sweden	11.96	42.73	13.57	26.23	74.47	31.03
	Switzerland	2.56	12.16	1.20	8.67	96.25	79.17
	United Kingdom	1.99	76.48	0.40	1.97	97.61	21.55
Asia	Japan	5.08	24.87	5.51	4.58	89.41	70.54
	China (Hong Kong)	7.47	19.49	28.76	52.15	63.77	28.36

	China (Macau)	6.38	45.38	50.30	26.61	43.32	28.01
	Indonesia	0.98	0.14	2.61	0.51	96.41	99.35
	South Korea	4.37	50.60	9.85	11.05	85.78	38.35
	Singapore	9.82	8.27	35.20	20.36	54.97	71.36
	Thailand	1.12	18.63	2.33	6.63	96.55	74.73
Oceania	Australia	14.57	47.68	18.60	28.62	66.83	23.70
	New Zealand	15.02	33.69	20.55	34.64	64.44	31.68
Means	All	10.25	22.85	16.09	19.37	73.66	57.79
	North America	32.84	36.37	18.01	23.36	49.15	40.26
	Latin America	4.04	1.62	9.83	11.36	86.14	87.02
	Europe	11.87	27.61	16.49	21.80	71.64	50.59
	Asia	5.03	23.91	19.22	17.41	75.74	58.67
	Oceania	14.79	40.68	19.57	31.63	65.63	27.69

Source: Elaboration by the authors.

b. Results of the Taylor Rule with Common Fluctuations

According to Yüksel et al. (2013), monetary policy rules depend on the policymaker's behavior to the structure of the economy and its objectives, which are subject to change over time. We tested the null hypothesis of linearity in Equation (16) against the alternative of a model with structural breaks using the Bai and Perron (1998, 2003) tests[17]. The results of the UD max and WD max tests reject the null hypothesis of linearity, while the sequential test suggests three breaks[18]. The dates of these estimated breaks delimit four regimes during the period analyzed. Table 3 shows the results of the estimation process considering these different regimes.

[17] Aragón and Medeiros (2013) also used the structural break tests developed by Bai and Perron (1998, 2003) to check for possible changes in the implementation of monetary policy. Two of the three specifications applied by them indicate a change in the third quarter of 2003 (2003.07 and 2003.09). The estimates of the reaction functions among the regimes delineated by the date of the estimated break reveal that the BCB reacted more strongly to the deviations of inflation from the target and the output gap after 2003.

[18] These results are available upon request.

During the first regime (2002.01 to 2004.06), the coefficient of inflation expectations is positive and statistically significant. The other parameters in this regime are not statistically significant. In the second one (2004.07 and 2009.02), besides inflation expectations, the coefficient of the regional factor for Latin America, excluding Brazil, is negative and statistically significant. During the third regime (2009.03 to 2012.08), the output gap parameter is positive and statistically significant, while the deviation of inflation from the target is not significant. The regional factor for Oceania has positive and significant effects. In the last regime of the sample, from 2009.04 to 2015.12, the significant parameters belong to the exchange rate variation (negative sign[19]) and the global and regional factors, except the North American one.

Table 3. BCB Reaction functions – short-term coefficients between the estimated regimes

Dependent variable: SELIC rates (accumulated % per year) -				
	2002.03–2004.06	2004.07–2009.02	2009.03–2012.08	2012.09–2015.12
i_{t-1}	1.303***	1.428***	1.459***	1.460***
	(0.172)	(0.116)	(0.150)	(0.115)
i_{t-2}	-0.471***	-0.449***	-0.487***	-0.422***
	(0.167)	(0.115)	(0.158)	(0.120)
Dj	0.132**	0.192***	-0.102	0.053
	(0.058)	(0.054)	(0.089)	(0.061)
y_{t-1}	-0.025	-0.006	0.047***	-0.010
	(0.061)	(0.009)	(0.017)	(0.011)
Δe_{t-1}	0.010	-0.004	0.005	-0.006***
	(0.010)	(0.003)	(0.004)	(0.002)
Global Factor	-0.291	0.004	-0.117	0.336**
	(0.182)	(0.015)	(0.141)	(0.136)
Asia Factor	0.039	-0.018	0.231	-0.428*

19 According to Obstfeld and Rogoff (1995), stronger exchange rates tend to have negative effects on output, and central banks can compensate for this effect by reducing interest rates. As proposed by Ball (1999), a stronger exchange rate contracts aggregate demand by making foreign goods cheaper and domestic ones more expensive, reducing net exports. Interest rate cuts attenuate this contraction.

	(0.202)	(0.015)	(0.193)	(0.234)
Europe Factor	-0.256	0.001	0.127	0.544***
	(0.201)	(0.017)	(0.069)	(0.120)
Latin America Factor	0.072	-0.251***	-0.007	0.262***
	(0.062)	(0.063)	(0.119)	(0.101)
North America Factor	0.111	-0.015	-0.040	0.097
	(0.186)	(0.025)	(0.079)	(0.081)
Oceania Factor	0.180	0.004	0.114**	-0.168***
	(0.260)	(0.023)	(0.057)	(0.059)
Constant	2.666***	0.208	0.361	-0.276***
	(0.816)	(0.154)	(0.302)	(0.118)

Note: 1) Standard errors in parentheses. 2) * significant at 10%; ** significant at 5%; *** significant at 1%.

The regimes in Table 3 coincide with changes in the presidency of the BCB[20]. Estimates of the reaction function parameters between these periods suggest important changes in the implementation of the Brazilian monetary policy.

In the first and second regimes, empirical evidence shows that the BCB reacted more strongly to deviations in inflation expectations from the target, where the SELIC response to Dj increased from 0.132 to 0.192. There is a negative correlation between the interest rate of Brazil (SELIC) and that of other Latin American countries in the second regime. During this period, there is a gradual reduction in the former, while the latter remain without significant variations. Vartanian (2010) showed that in this period there are no signs of macroeconomic convergence among the Mercosur countries, where the effect of exchange rate and monetary policy shocks on some macroeconomic variables are distinct and asymmetric between these countries, even in periods of economic instability.

[20] During this period, the BCB was in charge of three presidents: Armínio Fraga Neto (1999.03 – 2003.01), Henrique de Campos Meirelles (2003.01 – 2011.01), and Alexandre Antonio Tombini (2011.01 – 2016.06).

In the third regime, the BCB reaction is more sensitive to variations in the output gap than to deviations from inflation expectations, which appears with a statistically insignificant coefficient. This result corroborates those of Aragón and Medeiros (2013, 2015), Barbosa et al. (2016) and Cortes and Paiva (2017), showing that there was a change in the implementation of monetary policy compared to the previous period, with the monetary authority acting in a more discretionary way.

In the last regime, the empirical evidence shows a new change in the implementation of Brazilian monetary policy, suggesting a BCB reaction to external monetary policy shocks, mainly after the global financial crisis of 2010. In addition to the response to changes in the exchange rate, there is a positive reaction from the SELIC to movements in global and regional monetary policy.

Gali and Gertler (2010), Taylor (2013) and Edwards (2015) observed that central banks may take into account the monetary policy decisions of other countries when adjusting their interest rates. Indeed, these studies highlight the influence of monetary policy shocks in developed countries, such as the United States, on the monetary policies of emerging markets, such as Brazil.

Moreover, Rey (2015) noted the existence of a global financial cycle, influenced by the monetary policy decisions of the Federal Reserve. Furthermore, the spillover effects of US monetary policy shocks and decisions affect emerging and developed economies and may lead to greater coordination between countries' monetary policies. In fact, Chatterjee (2016) identified a common factor related to monetary policy movements in a group of developed nations.

In general, except in the first regime, the common fluctuations of the external monetary policy, represented by the global and/or regional factors, are important variables in specifying the Taylor rule for Brazil in the overall period studied. The third regime contains the 2008 and 2010 crises that generate strong comovements due to the coordination and spillover effects of monetary policies.

c. Robustness Analysis: BCB Reaction Function considering fluctuations in global economic activity

The reaction of the monetary authority may take into account fluctuations in the economic cycle and inflation rates at the global level (Ratti & Vespignani, 2019). To control the effects of the international business cycle on the BCB monetary policy decisions, the global economic activity index[21], proposed by Kilian (2009), $igrea_t$, was included in the BCB reaction function. The new test results from Bai and Perron (1998, 2003) indicate the existence of three breaks, in 2004.06, 2010.05, and 2013.06. These figures are shown in Table 4.

The break dates are close to the results in Table 3 and are within their confidence intervals. Furthermore, the inclusion of the variable $igrea_{t-1}$ does not significantly change the previous estimates. The exceptions are the coefficients associated with the variables y_{t-1}, which is not significant in the third regime; Δe_{t-1}, significant in the third regime and not significant in the fourth; and the factor associated with the countries of Oceania, which is not significant in the third regime. The $igrea_{t-1}$ has a statistically significant coefficient, indicating a possible BCB reaction to fluctuations in the level of global economic activity, after the financial crises of 2008 and 2010.

Table 4. BCB Reaction Functions – Short-Term Coefficients between the Estimated Regimes

Dependent variable: SELIC rates (accumulated % per year) -				
	2002.03–2004.06	2004.06–2010.04	2010.05–2013.05	2013.06–2015.12
i_{t-1}	1.176***	1.375***	1.259***	1.187***
	(0.195)	(0.114)	(0.182)	(0.155)
i_{t-2}	-0.353*	-0.388***	-0.359***	-0.204
	(0.188)	(0.114)	(0.178)	(0.148)
Dj	0.118**	0.206***	-0.138	0.019
	(0.059)	(0.050)	(0.133)	(0.059)
y_{t-1}	-0.016	-0.004	0.023	-0.023
	(0.062)	(0.010)	(0.014)	(0.017)
Δe_{t-1}	0.023	-0.001	-0.010**	-0.003
	(0.014)	(0.002)	(0.005)	(0.003)

[21] Other measures of global economic activity were used: the cycle of global industrial production (not considering the United States) and that of the United States, extracted from the Hodric-Prescott filter. The results are similar to those provided in Table 4 and are available upon request.

Global Factor	-0.233	0.006	-0.183	0.303**
	(0.185)	(0.016)	(0.194)	(0.153)
Asia Factor	0.064	-0.012	0.244	-0.411*
	(0.202)	(0.016)	(0.320)	(0.232)
Europe Factor	-0.230	-0.001	0.024	0.556***
	(0.201)	(0.019)	(0.072)	(0.153)
Latin America Factor	0.064	-0.200***	0.139	0.259*
	(0.062)	(0.054)	(0.259)	(0.133)
North America Factor	0.080	-0.008	-0.149	0.044
	(0.186)	(0.024)	(0.128)	(0.080)
Oceania Factor	0.247	-0.001	0.004	-0.173***
	(0.263)	(0.023)	(0.062)	(0.064)
$igrea_{t-1}$	0.003	0.001	0.002**	-0.0013
	(0.002)	(0.0006)	(0.0009)	(0.0009)
Constant	2.638***	0.027	1.156***	0.340
	(0.869)	(0.121)	(0.330)	(0.239300)

Note: 1) Standard errors in parentheses. 2) * significant at 10%; ** significant at 5%; *** significant at 1%.

CONCLUSION

In this study, we decomposed the common movements of nominal short-term interest rates in 28 countries and the Eurozone (Euribor) between 1995 and 2015, into global and regional common factors and idiosyncratic movements specific to each country. For this, we used a Bayesian dynamic factor model with multiple levels.

The estimated factors capture the main economic events (financial crises) during the period analyzed. The decomposition of the variance shows that common factors explain on average 28% of the total variance in the interest rates of all the economies in the sample. In particular, for countries such as Canada, Finland, Ireland, China (Macau) and the Eurozone (Euribor), these percentages were 52%, 83%, 52%, 59.4%, and 76.2%, respectively.

The cases with the largest participation of factors in the decomposition of variance outside the crisis period are empirical evidence of monetary policy interdependencies. There are also more intense movements in common factors in periods of financial crisis, indicating monetary policy spillover or contagion effects (Forbes & Rigobon, 2001, 2002).

The extended Taylor rule for Brazil suggests that common factors play an important role in the implementation of monetary policy by the BCB. This was especially observed between 2012.9 and 2015.12, a sub-period delimited by the structural break test. In particular, the correlation of the common global and regional interest rate fluctuations is more relevant for the implementation of monetary policy in Brazil as of 2009.

As a robustness analysis, other models were estimated with several indicators of global economic activity (the reported result used the Kilian (2009) index) and a global inflation variable. In general, the results regarding the dates of breaks and the statistical significance of the coefficients did not change substantially. The findings also suggest that, after the crises of 2008 and 2010, there is a positive response from the BCB to variations in global economic activity.

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