**Área 4: Macroeconomia, Economia Monetária e Finanças**

**Reviewing monetary policy inertia and its effects: the fractional integration approach for an emerging economy**

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**Resumo**

Este artigo analisou a inércia de política monetária de uma economia emergente relevante com base na abordagem de integração fracionária (Geweke and Porter-Hudak, 1983; Baillie et al. 1996), de modo que evitaram-se estimativas espúrias potenciais para tal inércia que podem ser encontradas quando são adotados testes convencionais de raiz unitária, os quais são usualmente aplicados na literatura relacionada. Os resultados confirmaram a hipótese de uma elevada inércia de política monetária (Clarida et al., 1999), mesmo após o controle para memória-longa nas séries temporais. Ademais, foi identificado um efeito positivo sobre as expectativas inflacionárias seguindo-se um aumento do componente inercial. Isto significa que, sob metas de inflação, o Banco Central deve ajustar ambas a taxa básica de juros e a sua trajetória de inércia a fim de cumprir a meta de inflação no tempo.

**Abstract**

This paper analyzed the monetary policy inertia of a relevant emerging economy based on the fractional integration approach (Geweke and Porter-Hudak, 1983; Baillie et al. 1996), so that the work avoided the potential spurious estimates for such an inertia that can be found by adopting conventional unit root tests, which were usually applied in the related literature. The results confirmed the hypothesis of a high monetary policy inertia (Clarida et al., 1999), even after controlling for long-memory in time series. Furthermore, it was identified a positive effect on inflationary expectations following an increase of the inertia component. It means that, under an inflation targeting regime, the Central Bank should adjust both the basic interest rate and its inertia path in order to accomplish the inflation target over time.

**Keywords:** monetary policy inertia; Taylor rule; fractional integration; Brazil.

**JEL:** C14; E52; E58.

**1. Introduction**

Taylor (1993) suggested that fluctuations in the observed conditions, such as in the inflation rate and/or in the level of economic activity, stimulate changes at the level of the operational goals of the Central Banks (CBs), particularly concerning the short-term interest rate – Sack and Wieland (1999) and Woodford (2003). In turn, it is widely accepted that the explanatory power of an estimation for basic interest rates increases when policy inertia is taken into account. Thus, the coefficient of monetary policy inertia, as captured by the lagged interest rate, indicates that CBs adjust the policy instrument slowly and sparingly in response to macroeconomic fluctuations.

Regarding the international empirical literature, estimates show a statistically significant coefficient of inertia between 0.7 and 0.8, which is considered a high level (Clarida et al., 1999). Several studies argue that the high gradualism of the policy is a deliberate consequence of Central Banks` strategy in order to avoid volatility in interest rates and its impacts on the financial market, to deal with uncertainties regarding the future effects of changes in interest rates and to serve as an adequate mechanism to manage expectations – Sack and Wieland (1999) and Bernanke (2004). In contrast, studies such as those by Rudebusch and Svensson (1999) and Rudebusch (2006) argued that the inertial movement found in the literature represents a very slow, late behaviour of the authorities in response to macroeconomic evolution and that less timid movements would be more effective in stabilising both output and inflation.

The first main goal of our study was to test for a hypothesis of a particular statistical problem in estimates of monetary policy inertia: the existence of *long-memory* in data and its consequence in terms of an overestimation of such results. The statistical error of disregarding long-memory is that, for instance, when a time series is assumed as stationary by the conventional unit root tests, but it actually presents long-memory, the resulting auto-regressive inference commonly overestimates coefficients. One way to deal appropriately with long-memory in data is by the *fractional integration approach* (Geweke and Porter-Hudak, 1983; Baillie et al. 1996; Reisen et al., 2001; Reisen et al., 2003; Wei, 2006). Thus, our initial contribution to the associated literature was to estimate monetary policy inertia in Brazil, but only after controlling for such a potential problem. The particular case of Brazil is an important one as some authors have previously identified expressive values of its monetary policy inertia (Silva et al. 2015; Moreira, 2015). However, these studies did not deal with potential long-memory in data.

The second aim, in turn, was to identify possible inflationary effects of changes in monetary policy inertia. In order to make it possible, the *Kalman Filter* initially applied to extract a time behavior of the inertia coefficient for a specific estimated Taylor rule. In sequence, the paper tested for the hypothesis of a deterioration (an increase) of expected inflation as a result of a higher level of monetary policy smoothing, based on Generalized Method of Moments (GMM) regressions.

The remainder of this article is structured as follows: in the next section are presented some relevant issues in the theoretical and empirical literature; section 3, in turn, shows the methodological strategy, while sections 4 and 5 address the dataset and discuss the empirical findings from our estimates, respectively. At last, concluding remarks are presented and the list of references.

**2. Literature review: theoretical and empirical issues**

Since the seminal paper of Taylor (1993), many studies have developed and added new theoretical and empirical aspects into the literature about monetary policy rules, from the Taylor’s non-inertial policy rule to the inertial rule in Clarida et. al. (1999), including the focus on forward-looking issues as a better strategy for performing monetary policy.

According to Rudebusch (2006), there exists a consensus about the inertial nature of the basic interest rate adjustments over time. In other words, when central banks decide to move their policy rates (e.g. to react to output and/or inflation changes), such an adjustment is made in a smoothed path rather than in an abrupt one (Clarida *et al*., 1999; Sack and Wieland, 1999; Woodford, 2003).

From an empirical perspective, this gradualist monetary policy can be expressed by the following Taylor rule, especially due to its inertial component as the main aspect:

*it = it-1 + (1 – )[iLP + c1(Et t+j – t\* + c2(t-j – t\* + c3yt-1 + c4et-1 – et-2)] + t* (1)

Where it stands for the nominal basic interest rate in period t, iLP is the neutral interest rate, yt-1 is the output gap in period t-1, (Et t+j – t\* stands for the expected inflation deviation (from the target) over the future horizon t+j; et-1 – et-2) represents the nominal exchange rate change from t-2 to t-1 and, at last, t is a white noise shock. In turn, c1, c2, c3 and c4 are positive parameters and, most importantly,  stands for the inertial monetary policy component, which varies from 0 to 1. So, while  = 0 represents a limit case for a non-inertial monetary policy,  = 1 cover the extreme case of a fully inertial monetary policy.

Indeed, international empirical studies (including for the US economy) have pointed out that the inertial component normally assumes values around 0.8 (Clarida et al. 1999), which means a strong inertial path. Particularly for the Brazilian case, the evidence is not different. From Minela et al. (2002) and Moreira (2015) to Barbosa et al. (2016), among others, monetary policy rules applied to Brazilian data have reported inertial coefficients around 0.8, and even 0.9, thereby indicating an expressive smoothing behavior of the Selic rate (Brazil’s nominal basic interest rate).

If, on the one hand, the aforementioned literature has a consensus about a gradualist nature of the monetary policy, on the other hand, such works do not have the same convergent view regarding the motives for that inertia, as well as with respect to its true source. Among the potential reasons for an *intrinsic* smoothing monetary policy – i.e. an inertial policy rate resulting from a Central Bank’s deliberate strategy – are a) concerns with financial stability; b) management of public’s expectations and; c) a response to structural uncertainty.

One of the main arguments raised in the literature to justify gradualism in monetary policy is associated with the elimination of the volatility of the interest rate and the financial market, i.e., the objective is to generate interest rates less volatile than those suggested by a non-inertial policy rule. There is a fear that interest rate shock or abrupt movements in interest rates cause crises in financial markets (Sack and Willand (1999), Clarida *et al.* (1999), Woodford (2003), Bernanke (2004)). The rationality for this argument is the assumption that as the *ρ* coefficient increases, the variance of the level of the short-term interest rate decreases, providing greater financial stability given that a number of financial assets, such as exchange, fixed-income securities and shares, are correlated with movements in short-term interest rates.

According to Sack and Wieland (1999) and Bernanke (2004), there is a second justification for gradualism: using interest rate adjustments as a tool for expectation management. In an environment of partial adjustment, which is represented by a sequence of small adjustments of the interest rate in the same direction, an increase in the basic interest rate not only increases the current interest rate but also signals to the market that the rate should continue to rise for some time, thereby contributing to stability by forming expectations for future short-term rates. Thus, by leading market participants to anticipate that changes in the policy rate will be followed by other changes in the same direction, policy gradualism can increase the ability of the CB to affect the long-term interest rate, thereby influencing the behaviour of the economy with higher efficiency and lower volatility of the current basic rate.

The third justification that motivates gradualism in monetary policy is related to structural uncertainty. The rationale for this argument comes from the principle that uncertainty generates caution and that caution suggests a gradual adjustment in the policy rate (Bernanke, 2004). Sack and Wieland (1999) argued that one of the main assumptions made by the empirical literature regarding macroeconomic models and policy rules is the full knowledge of the economic structure. In reality, the parameters of the models are imprecisely estimated and typically vary unpredictably over time. An environment of uncertainty regarding the sensitivity of the parameters generates a trade-off in the policy-making process because aggressive adjustments in the policy rate, on average, approximate the output to its desired level but create greater uncertainty in the results. Moreover, policy makers do not know precisely the correct structure of the limiting model. Thus, it is more appropriate to moderately implement monetary policy due to the presence of strong parametric uncertainty (Brainard, 1967).

In turn, and most importantly for our objectives in this paper, there exists an alternative vision about the source of the aforementioned strong inertial component reported in empirical studies. According to this extrinsic view, estimation or statistical problems could be considered as causing spurious and high estimates for the  inertial component. In this work is explored a hypothesis associated with problems found in the integration order identification.

From a statistical perspective, the related literature is mostly based on typical Dickey-Fuller tests to identify unit root in data. In turn, one of the problems of such tests is that it can easily overestimate (or underestimate) the integration order of a specific time series (Diebold and Rudebush, 1991; Hasler and Wolters, 1994; Lee and Amsler, 1997). That is, the conventional unit root tests, such as Augmented Dickey-Fuller, Phillips-Perron and KPSS tests, among others, do not deal with the *long-memory* in data.

The statistical error of disregarding long-memory is that, for instance, when a time series is assumed as stationary by the conventional tests, but it actually presents long-memory, the resulting auto-regressive inference commonly overestimates coefficients. Thus, one way to deal appropriately with long-memory in data is by the fractional integration approach (Geweke and Porter-Hudak, 1983; Baillie et al. 1996; Reisen et al., 2001; Reisen et al., 2003; Wei, 2006). This approach allows for transforming a long-memory time series into a short-memory one, so that it can be avoided spurious estimates in autoregressive models.

**3. Methodological strategy**

The analysis of the Brazilian basic interest rate (SELIC) was divided into two steps. First, it was applied an univariate analysis, based on the method of Geweke and Porter-Hudak (GPH) (Geweke and Porter-Hudak, 1983), as well as semiparametric and parametric methods, especially ARFIMA models. Such an analysis allows for identifying the existence of long-memory in time series (Reisen, 1994; Reisen et al.2001; Wei, 2006).

Second, after testing for the fractionally integrated process of data and filtering eventual long-memory behavior of all variables, a multivariate analysis based on the estimation of *Generalized Method of Moments* (GMM) was applied, in order to controlling for the potential factors determining SELIC over time. The GMM method was applied for purposes of robustness of the estimates, as it is regarded as a method able to overcome potential problems from heteroskedasticity, autocorrelation and endogeneity (Hansen 1982). In turn, the instrumental variables set was based on the i-period lag for the explanatory time series, which thus satisfy the hypothesis of exogeneity of the instruments (Johnston 1984). Finally, to assure the appropriate specification of the instrumental variables, an analysis of *overidentification* was performed by the J-test (Gragg 1983; Hansen 1982).

It is also worth noting that before applying the multivariate analysis, it was applied the filter, where representing each times series, to extract the short-memory time series. In this case, to filter the series, the fractional integration parameter () considered a bandwidth equal to (see Reisen et al., 1994 and section 5). The correct of each time series was defined based on the GPH method. In contrast, the most part of the empirical studies regarding interest rates in Brazil applied typical Dickey-Fuller tests, which have low power against the fractional alternative (Diebold and Rudebush, 1991; Hasler and Wolters, 1994). Besides, Lee and Amsler (1997) found that the KPSS test cannot consistently distinguish between a process with nonstationary long memory and a unit root process (). In this perspective, the idea of this paper of using times series based on the fractional integration approach, as well as filtering possible long-memory process in data, can be regarded as an innovative trend in applied macroeconomics.

Furthermore, the paper estimated a time-varying monetary policy inertia path to infer on potential counter or pro-cyclical effects on inflationary expectations in Brazil. It was defined initially a state-space model to apply the *Kalman filter*, thus following the suggestion of Harvey (1989), and based on the extracted inertial coefficient time path, the effects on inflationary expectations was investigated.

**4. Dataset**

The data used in the present study covered the period from January 2004 to December 2017 on a monthly basis. The choice for the sample period considered the stabilisation of the inflation target at 4.5% per year for the *Broad Consumer Prices Index* (i.e. the official index of the Brazilian inflation targeting regime) over the most part of the time sample. Thus, this period can regard as one of stability or consolidation of such a regime in Brazil. The time series were the following ones and their graphical behaviour and descriptive statistics are presented in Figure 1 and Table 1, respectively:

*SELICt =* the basic interest rate of the Brazilian economy (effective *Selic* rate) (Source: Central Bank of Brazil);

*Pt* = the inflation rate accumulated in the last 12 months (Source: Ipeadata);

*EXP\_Pt* = the expected inflation rate for 12 months ahead (Source: Central Bank of Brazil);

*Yt =* Brazil’s output, measured by the index of economic activity of the BCB (ICB-Br) (Source: Central Bank of Brazil);

*NEXCHt* = the nominal exchange rate: Real/U.S. Dollar (R$/US$) (Source: Central Bank of Brazil);

REXCH = real effective exchange rate index (Source: Ipeadata);

DEBT = Net public debt/GPD ratio (%) (Source: Central Bank of Brazil).

**Figure 1: Graphical behaviour**



Source: drawn from the survey data.

**Table 1: Descriptive statistics**

|  |  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- | --- |
|  | **SELIC** | **P** | **EXP\_P** | **Y** | **NEXCH** | **REXCH** | **DEBT** |
| **Mean** | 12.32 | 5.75 | 5.21 | 131.23 | 2.37 | 70.36 | 48.78 |
| **Median** | 11.88 | 5.75 | 5.37 | 135.13 | 2.21 | 67.33 | 47.38 |
| **Maximum** | 19.75 | 10.22 | 7.29 | 148.78 | 4.04 | 102.86 | 67.00 |
| **Minimum** | 7.00 | 2.43 | 3.40 | 103.52 | 1.56 | 54.63 | 41.45 |
| **Std. Dev.** | 3.17 | 1.68 | 0.91 | 12.66 | 0.63 | 10.96 | 5.82 |

Source: drawn from the survey data.

**5. Results**

*5.1. Univariate analysis of the SELIC series*

According to Brockwell and Davis (2002), a fractionally integrated ARMA processes, or an ARFIMA, with , satisfy difference equations as follows:

, (2)

where and are polynomials of degrees and , respectively, satisfying and , for all such that ; is the backward shift operator; and is a white noise sequence with mean 0 and variance . In this case, the autocorrelation (ACF), , at lag , has the following property:

, as . (3)

This meaning that for an ARFIMA process converges to zero as at a much slower rate than for an ARMA process. Thus, fractionally integrated ARMA processes are said to have “long memory”. Otherwise, stationary processes whose ACF converges to zero rapidly, such as ARMA processes, are said to have “short memory”.

In general, based on the value of , the ARFIMA process is characterized as follows: i) if , the process is considered to have short memory; if , the process is stationary with long memory and mean reversion; if , the series is said to be nonstationary with long memory, but still shows mean reverting; and, if , the series is nonstationary and does not present mean reversion[[1]](#footnote-1).

In this context, the performance of traditional unit root tests in the presence of long memory has received much attention in recent years (Reisen et al., 2003). To Diebold e Rudebush (1991) and Hasler e Wolters (1994), for example, the Dickey-Fuller tests present low power against fractional alternative. Furthermore, Lee and Amsler (1997) demonstrated that the KPSS test could not consistently distinguish between a process with nonstationary long memory and a unit root process ().

Besides, according to Wei (2006), an ARFIMA process presents several characteristics that are similar to those of a nonstationary process in finite samples. For instance, the ACF of an ARFIMA model decays very slowly, that is, similar to the sample ACF of a nonstationary time series (ARIMA). In addition, the ARFIMA process and the nonstationary ARIMA present periodograms[[2]](#footnote-2) that diverge at the zero frequency. These similarities often lead to model misspecification. For example, a stationary ARFIMA model could be misspecified as a nonstationary ARIMA model. This overdifferencing may lead to some undesirable effects on parameter estimation and forecasting.

Figure 2 shows the ACF for the SELIC variable, considering 50 lags. As the sample autocorrelations decay slowly, the series seems to be nonstationary and possibly characterized by long memory behavior.

**Figure 2: Sample autocorrelation function for the SELIC**



Source: drawn from the survey data.

In order to verify if the SELIC series presents a unit root, the tests of Augmented Dickey-Fuller (ADF), Phillips-Perron (PP) and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) were used. The tests showed that the SELIC is nonstationary in level. Then, the first difference could result in an overdifferencing of the series, when compared with a fractionally integrated process. It is important do say that the ACF for the SELIC series in first difference showed positive and negative signs, indicating, in this case, which the series would be overdifferentiated (Morettin, 2008).

**Table 2: Unit root test for the variable SELIC in level**

|  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- |
| Variable | ADF | K | PP | K | KPSS | K |
| SELIC | -2.339 ns | 2 | -1.9362ns | 9 | 0,2532\*\*\* | 10 |

Source: drawn from the survey data.

Note: 1) \*\*\* Significant at 1%, ns Non-significant at 10%; 2) K = number of lags; 3) Tests performed with constant and/or trend. However, tests without constant and/or without constant and trend showed similar results.

To lead with this problem, this paper tested the presence of a fractionally integrated process in the SELIC by means of the Geweke and Porter-Hudak (GPH) method (Geweke and Porter-Hudak, 1983). Table 3 presents the results of the semiparametric estimation of for different bandwidths, , with equal to the number of observations and . According to results, regardless of the choice of , the estimates indicate a statistically significant behavior of long memory in the SELIC, namely: i) stationary for e ; and, nonstationary with mean reversion for e .

**Table 3: Estimated GPH for different bandwidths**

|  |  |  |  |
| --- | --- | --- | --- |
|  |  | GPH () | Standard-Error |
| 0.4 | 8 | 0.4621 | 0.2453 |
| 0.5 | 13 | 0.4694 | 0.2426 |
| 0.6 | 22 | 0.7397 | 0.0940 |
| 0.7 | 36 | 0.8974 | 0.0609 |

Source: drawn from the survey data.

To Reisen et al. (2001), the semiparametric estimators may exhibit bias in the estimator when the autoregressive or moving average parameters are nonzero. Then, Tables 4 and 5 present the best ARFIMA models estimated by semiparametric and parametric methods, respectively. To choose the best-estimated models was used the Akaike Information Criteria (AIC). For semiparametric estimators, when a model with autoregressive and moving average components was estimated, firstly the SELIC was filtered with , considering (see, Reisen, 1994). The Box-Pierce and Ljung-Box tests revealed that residuals of the estimated models were uncorrelated.

As can be seen in Tables 3, 4 and 5, the values ranged between 0.4621 and 0.8974. For information, the smallest AIC value occurred for the ARFIMA semiparametric model. In all cases, the value stayed below one, i.e., smaller than the unit root case. In short, the univariate results suggest that the SELIC series has long-memory, but with mean reversion. It means that if the analysis were based only on the typical unit root tests (Table 2), the SELIC series would be overdifferenced.

**Table 4: Estimated semiparametric models**

|  |  |  |  |
| --- | --- | --- | --- |
| Models | AR and/or MA estimates | Standard-Error | AIC |
| ARFIMA |  | 0.0194 | 140.95 |
| ARFIMA |  | 0.0753  0.0758 | 132.92 |
| ARFIMA |  | 0.0772  0.0948  0.0782 | 133.47 |
| ARFIMA |  | 0.0151  0.0622 | 135.64 |
| ARFIMA |  | 0.1621  0.1581  0.1569 | 134.56 |

Source: drawn from the survey data.

**Table 5: Estimated parametric models**

|  |  |  |  |
| --- | --- | --- | --- |
| Models | AR and/or MA estimates | Standard-Error | AIC |
| ARFIMA |  | 0.0224 | 213.14 |
| ARFIMA |  | 0.2831  0.1595 | 179.79 |
| ARFIMA | ns | 0.0250  0.0872 | 196,38 |
| ARFIMA | ns  ns | 0.3768  0.3787  0.3764 | 156.37 |
| ARFIMA |  | 0.0250  0.0958  0.0838 | 171.56 |

Source: drawn from the survey data.

Note: ns Insignificant coefficient.

*5.2. Estimated GPH for all the times series*

The results of the semiparametric estimates for () are shown in Table 6 (including SELIC). According to the estimates, the time series adopted for the estimation in the next sections can be classified as follows: i) stationary with long memory and mean reversion (SELIC, P and EXP\_P); ii) nonstationary with long memory, but with mean reversion (DLSP and REXCH); and, nonstationary (NEXCH and Y). Therefore, the filter was adopted, where representing each times series, to extract the short-memory time series. In the estimation of we considered (see Reisen et al., 1994).

**Table 6: Estimated GPH () for the time series**

|  |  |  |
| --- | --- | --- |
| Variable | GPH () | Std-Error |
| SELIC | 0.4694 | 0.2426 |
| P | 0.4454 | 0.1558 |
| EXP\_P | 0.4830 | 0.1202 |
| DLSP | 0.7498 | 0.1581 |
| NEXCH | 1.0660 | 0.1945 |
| Y | 1.0211 | 0.1226 |
| REXCH | 0.9954 | 0.2100 |

Source: drawn from the survey data.

*5.3. Identifying monetary policy inertia*

After filtering the time series by means of the fractional parameters () estimated by GPH, the GMM regressions were performed as a way to identify either the monetary policy inertia in Brazil or the factors which determine basic interest rate movements over time. For robustness purposes, regressions were estimated either with the nominal exchange rate (NEXCH) or with the real exchange rate (REXCH). It was also set the public debt/GDP ratio as a control variable to allowing fiscal effects on monetary policy decisions. The general regression model can be defined as:

*SELICt = SELICt-1 + 1 EXP\_P t-1 + 2 P t-1 + 3 Y t-1 + 4 EXCH t-1 + 5 DEBT t-1 + et* (4)

Regardless of using nominal or exchange rate, some robust evidence were obtained. First, the hypothesis of a high monetary policy inertia over the sample were confirmed, even after controlling the long-memory and the nonstationary in data. The estimates for ranges from 0.86 to 0.99 and with statistical significance at 1% (Table 7). In such a case, the results are convergent with those found in Moreira (2015) and Barbosa et al. (2016). Second, a backward-looking pattern in Brazil’s monetary policy was identified, an absence of statistically significant coefficients regarding expected inflation (EXP\_P) over all regressions was found. Instead, lagged observed inflation (P) presented statistical significance in two regressions (Eq. 2 and Eq. 3).

Furthermore, monetary policy in Brazil seems to ignore exchange rate movements. The related coefficients (4) did not present statistical significance. It is noteworthy that the public debt/GDP ratio presented a negative and statistically significant signal in all regressions in which we set such a component. This evidence demonstrates that monetary policy decisions were favorable to fiscal authorities, thereby lowering the cost of financial public spending in moments of an increase of the public debt/GDP ratio. Such a passive monetary policy can be understood as a case for fiscal dominance (Woodford, 2003).

In turn, all the estimated regressions satisfied the null hypothesis of *overidentification* andappropriate specification of the instrumental variables, such as showed by the J-test in each regression.

**Table 7: GMM estimates for SELIC t**

|  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- |
| **Explanatory**  **Variables** | **Equations** | | | | | |
| **Eq.1** | **Eq.2** | **Eq.3** | **Eq.4** | **Eq.5** | **Eq.6** |
| **SELIC** t-1 | 0.972\*\*\* | 0.865\*\*\* | 0.868\*\*\* | 0.936\*\*\* | 0.951\*\*\* | 0.994\*\*\* |
|  | (0.107) | (0.054) | (0.053) | (0.175) | (0.080) | (0.065) |
|  | [9.021] | [15.864] | [16.374] | [5.334] | [11.826] | [15.092] |
| **EXP\_P** t-1 | 0.076 | 0.121 | - | 0.209 | 0.312 | - |
|  | (0.182) | (0.104) | - | (0.265) | (0.192) | - |
|  | [0.419] | [1.161] | - | [0.790] | [1.623] | - |
| **P** t-1 | 0.144 | 0.114\* | 0.177\*\* | 0.133 | - | 0.055 |
|  | (0.087) | (0.063) | (0.071) | (0.132) | - | (0.073) |
|  | [1.642] | [1.788] | [2.491] | [1.004] | - | [0.757] |
| **Y** t-1 | 0.324\*\* | -0.020 | 0.069 | 0.212 | - | - |
|  | (0.151) | (0.070) | (0.065) | (0.141) | - | - |
|  | [2.145] | [-0.287] | [1.054] | [1.502] | - | - |
| **NEXCH** t-1 | 1.980 | - | 0.869 | - | - | - |
|  | (1.890) | - | (1.061) | - | - | - |
|  | [1.048] | - | [0.819] | - | - | - |
| **REXCH** t-1 | - | - | - | 0.117 | 0.104 | 0.015 |
|  | - | - | - | (0.127) | (0.087) | (0.045) |
|  | - | - | - | [0.918] | [1.201] | [0.349] |
| **DEBT** t-1 | -0.155\*\* | - | - | -0.151\*\* | -0.079\*\* | -0.051\* |
|  | (0.068) | - | - | (0.073) | (0.039) | (0.029) |
|  | [-2.269] | - | - | [-2.074] | [-2.025] | [-1.727] |
| **Adj. R2** | 0.760 | 0.888 | 0.873 | 0.782 | 0.808 | 0.879 |
| **J stat.** | 3.970 | 11.956 | 9.880 | 4.500 | 6.349 | 8.909 |
| **Prob. (J)** | 0.680 | 0.153 | 0.273 | 0.609 | 0.608 | 0.349 |
| **Instrument rank** | 12 | 12 | 12 | 12 | 12 | 12 |

Obs.: ( ) for standard error and [ ] for t-stat. (\*) confidence level at 10%, (\*\*) 5% and (\*\*\*) 1%. Instrument specification: exp\_p(-2), exp\_p(-3), exp(-4), exp\_p(-5), p(-2), p(-3), p(-4), p(-5), y(-2), y(-3), y(-4), y(-5).

Source: drawn from the survey data.

*5.4. Extracting monetary policy inertia over time*

In order to analyze the behavior of  over time the *Kalman Filter* was applied on the regression with the best fit to data (Eq. 2). However*,* according to Harvey (1989), it is firstly necessary to define the state-space model to apply the Kalman filter. Define the observations equation and the system equation as:

Jt = z’tt + dt +vt (5)

t = Ttt-1 + t 

Where Jt stands for the observed variables, which is related with a m x 1 vector, t, known as state vector (t = 1, …, T); zt is a N x m matrix, dt represents a N x 1 vector and vt is a N x 1 vector of disturbances with zero mean and covariance Gt. In turn, although the elements of t are not observable, they are determined by a Markov process of first order, while Tt is a m x m matrix and t is a b x 1 vector of uncorrelated residuals with zero mean and covariance matrix Qt. Therefore, the state-space specification assumes the following form (Equations 9 to 14):

Jt = SELICt = SELICt-1 + 1 EXP\_P t-1 + 2 P t-1 + 3 Y t-1 (7)

z’t = [SELICt-1 + 1 EXP\_P t-1 + 2 P t-1 + 3 Y t-1] (8)



(10)

(11)

(12)

After solving this state-space model and extracting the time behavior of t by the *Kalman filter*, the resulting trajectory can be seen in Figure 3. It is possible to observe that since the beginning of the sample there has been a decrease of the monetary policy inertia in Brazil, even though such values fluctuate around an expressively high inertia component.

**Figure 3: Brazil`s monetary policy inertia (t) from Jan/2004 to Dec/2017**



Source: drawn from the survey data.

*5.5. Testing for pro-cyclical effects of monetary policy inertia*

One important way to enhance the analysis is by asking: what do occur with Brazil’s inflationary expectations as a response to an increase in monetary policy inertia? The assumption of this paper is that an extremely high coefficient of inertia can be followed by a pro-cyclical monetary policy, i.e., by reducing the real interest rate in response to rising inflation pressures on the economy, which would result in persistent inflationary deviations around the inflation target and so an increase in inflationary expectations (Silva et al. 2015; Rudebusch and Svensson, 1999; Rudebusch, 2006).

In order to test for such a hypothesis an expected Phillips curve was estimated in the form:

*EXP\_P t = 1 EXP\_P t-1 + 2 P t-1 + 3 Y t-1 + 4 NEXCH t-1 + 5  t-1 + ut*(13)

While 1, 2, 3 and 4 are associated with control variables, the main goal in this step was to verify the estimated signal for 5, assuming the premise for 5 > 0, that is, a type of deterioration of inflationary expectations following higher levels of monetary policy inertia. Table 8 presents the estimates from three regressions based on Equation (15). Most importantly, for all tested specifications the t obtained positive and statistically significant values, thus confirming the theoretical assumption and converging with some of the previous studies on the subject (e.g. Moreira, 2015; Silva, et al. 2015). All the regressions, in turn, satisfied the exogeneity condition of instruments as the J-stat resulted in a validation of the null hypothesis.

**Table 8: GMM estimates for EXP\_P t**

|  |  |  |  |
| --- | --- | --- | --- |
| **Explanatory Variables** | **Eq.1** | **Eq.2** | **Eq.3** |
| **EXP\_P** t-1 | 0.745\*\*\* | - | 0.753\*\*\* |
|  | (0.103) | - | (0.087) |
|  | [7.230] | - | [8.598] |
| **P** t-1 | 0.034 | 0.276\*\*\* | 0.032 |
|  | (0.041) | (0.045) | (0.037) |
|  | [0.845] | [6.132] | [0.868] |
| **Y** t-1 | - | 0.104\* | - |
|  | - | (0.071) | - |
|  | - | [1.449] | - |
| **NEXCH** t-1 | - | - | 0.529 |
|  | - | - | (0.591) |
|  | - | - | [0.895] |
| ****t-1 | 0.077\*\* | 0.235\*\*\* | 0.077\*\*\* |
|  | (0.034) | (0.039) | (0.029) |
|  | [2.265] | [5.885] | [2.638] |
| **Adj. R2** | 0.527 | 0.202 | 0.509 |
| **J stat.** | 6.228 | 13.549 | 6.187 |
| **Prob. (J)** | 0.716 | 0.139 | 0.626 |
| **Instrument rank** | 12 | 12 | 12 |

Obs.: ( ) for standard error and [ ] for t-stat. (\*) confidence level at 10%, (\*\*) 5% and (\*\*\*) 1%. Instrument specification: exp\_p(-2), exp\_p(-3), exp(-4), exp\_p(-5), p(-2), p(-3), p(-4), p(-5), y(-2), y(-3), y(-4), y(-5).

Source: drawn from the survey data.

**6. Concluding remarks**

This work brought some contributions to the literature on monetary policy inertia. Firstly, by analyzing time series of a relevant emerging economy over the period from Jan/2004 to Dec/2017, the results confirmed the existence of a significantly high monetary policy inertia in Brazil, even after controlling the long-memory in data. In order words, despite of avoiding spurious estimates due to conventional unit root tests, which were so usually applied in previous studies, the GMM results corroborating the remarkably high values of the inertia component in empirical Taylor rules (Clarida et al. 1999). In doing so, this paper gave more strength to the intrinsic view about inertial monetary policy rules (Rudebusch, 2006), that is, such an inertia should not be regarded as resulting from spurious estimates due to statistical problems.

Second, the *Kalman filter* approach was applied to extract a time-varying monetary policy inertia as a way to identify potential effects of the latter on inflationary expectations. The underlying assumption of a pro-cyclical and inefficient monetary policy in a context of an increasing inertia component was robustly confirmed. Therefore, it is possible to assume that the higher (the lower) the level of monetary policy inertia the higher (the lower) the inflationary expectations, at least in Brazil and based on the time sample adopted. It means that although several arguments are consensually pointed out in favor of monetary policy smoothing, monetary authorities should closely manage it and avoid extreme persistence in basic interest rates in order to tame inflationary expectations.

An important way to testing the hypothesis and methodology in a more robust manner is by applying it for other economies, either based on times series or panel data methods. To develop theoretical models (resembling new Keynesian DSGE models) that aim at identifying an optimal time path of the monetary policy inertia is also a relevant step in this direction.

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**References**

Baillie, R.T., Chung, C.-F. and Tieslau, M.A. (1996). “Analysing inflation by the fractionally integrated Arfima-Garch model”, *Journal of Applied Econometrics*, Vol. 11, pp. 23-40.

Barbosa, F.H.; Camelo, F.D. and João, I.C. (2016). “A Taxa de juros natural e a Regra de Taylor no Brasil: 2003-2015”, *Revista Brasileira de Economia*, Vol. 70, pp. 399-417.

Bernanke, B. (2004). *Gradualism.* *Remarks at an economics luncheon*, Seattle, Washington.

Brainard, W. (1967), “Uncertainty and the effectiveness of policy”, *American Economic Review*, Vol. 57, pp. 411-425.

Brockwell, P. J. and Davis, R. A. (2001). *Introduction to time series and forecasting*. 2nd ed. Springer, New York.

Clarida, R., Galí, J. and Gertler, M. (1999), “The science of monetary policy: a new keynesian perspective”, *Journal of Economic Literature*, Vol. 37, pp. 1661-1717.

Diebold, F.X. and Rudebush, G. D. (1991) “On the power of Dickey–Fuller tests against fractional alternatives”, *Economics Letters*, Vol. 35, pp. 155-160.

Geweke, J. and Porter-Hudak, S. (1993) “The estimation and application of long memory time series models”, *Journal of Time Series Analysis*, Vol. 4, pp. 221-238.

Harvey, A.C. (1989). *Forecasting, structural time series models and the Kalman Filter*, Cambridge Univ. Press, Cambridge.

Hassler, U. and Wolters, J. (1994). “On the power of Dickey-Fuller tests against fractional alternatives”. *Economic Letters*, Vol. 45, pp. 1-5.

Lee, H.S. and Amsler, C. (1997) “Consistency of the KPSS unit root test against fractionally integrated alternative”, *Economics letters*, Vol. 55, pp. 151-160.

Minella, A., de Freitas, P.S., Goldfajn, I., and Muinhos, M.K. (2002), “Inflation targeting in Brazil: lessons and challenges”, working paper (Series 53), Central Bank of Brazil.

Moreira, R.R. (2015), “Reviewing Taylor rules for Brazil: was there a turning-point?”, *Journal of Economics and Political Economy*, Vol. 2, pp. 276-289.

Reisen, V.A. (1994) “Estimation of the fractional difference parameter in the ARIMA(p, d, q) model using the smoothed periodograma”, *Journal of Time Series Analysis*, Vol. 15, pp. 335-350.

Reisen, V.A., Abraham, B. and Lopes, S.R.C. (2001). “Estimation of parameters in ARFIMA processes: a simulation study”, *Communications in Statistics Simulation and Computation*, Vol. 4, pp. 787-803.

Reisen, V.A., Cribari-Neto, F. and Jensen, M.J. (2003). “Long memory inflationary dynamics: the case of Brazil”. *Studies in Nonlinear Dynamics & Econometrics*, Vol. 7, pp. 1-16.

Rudebusch, G. D. (2006). “Monetary policy inertia: fact or fiction?”, *International Journal of Central Banking*, Vol. 2, pp. 85-135.

Rudebusch, G. and Svensson, L. (1999). “Policy rules for inflation targeting”, in Taylor, J. (Ed.), *Monetary Policy* *Rules*, University of Chicago Press.

Sack, B. and Wieland, V. (1999). “Interest-rate smoothing and optimal monetary policy: a review of recent empirical evidence”, *Journal of Economics and Business*, Vol. 52, pp. 205-228.

Silva, M.J., Brasil, G.H. and Moreira, R.R. (2015), “Dynamic relations of the inertia of monetary policy: application to the Brazilian case by a Kalman approach”, *International Journal of Monetary Economics and Finance*, Vol. 9, pp. 1-24.

Taylor, J.B. (1993). “Discretion versus Policy Rules in Practice”. *Carnegie Rochester Conference Series on Public Policy*, Vol. 39, pp. 195-214.

Wei, W. (2006). *Time series analysis: univariate and multivariate methods*, 2nd ed., Pearson Education, Nova York.

Woodford, M. (2003). *Interest and prices: foundations of a theory of monetary policy*, Princeton University Press.

1. The intermediate memory of an ARFIMA model is defined when is in [-0,5; 0). [↑](#footnote-ref-1)
2. The periodogram function describes how the total variability of the series is partitioned over the various components relating to each of the Fourier frequencies. [↑](#footnote-ref-2)