TAYLOR RULE IN BRAZIL[1]

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Abstract

In this paper, we use the Taylor Rule to characterize empirically the Brazilian monetary policy before and after its major and succesful stabilization plan, Real Plan, launched in 1994. Specifically, we show how the inflation coefficient has changed after the stabilization plan was carried out. This is a natural experiment to test theories surrounding the Taylor Rule in which monetary instability is characterized by an inflation coefficient less than one, whereas monetary stability will have a greater than one coefficient (see Woodford's (2003)). Very suprisingly the paper shows that the inflation coefficient has remained less than one even after the stabilization. Our results are quite robust with respect to different samples, lags of variables, proxies for GDP, proxies for potential GDP and even with respect to econometric methods (see Bueno (2005a, 2005b)). The implications are very important both theoretically and empirically. First, it shows some gap in theory that deserves further investigation. Second, it suggests that the inflation targeting regime has been uneffective in Brazil confirming a feeling largerly spread among Brazilians.

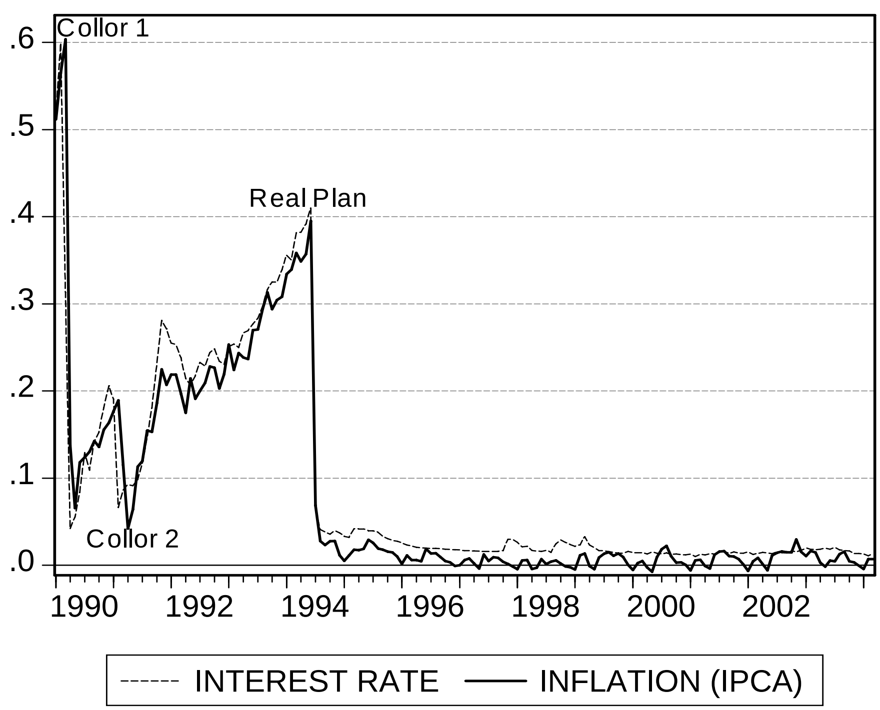
Key Words: Taylor Rule, inflation targeting, price stability

JEL: E52, E58, C32, C51

1 INTRODUCTION

Brazil is one of the greatest economies in the world and lived a high level of inflation for several years, mainly between 1980 and 1994. In July of 1994 the Real Plan was launched. Certainly it was one of the most well succeeded price stabilization plan ever made on the Earth. By looking at figure 1, one can see the high inflation level before the middle of 1994 (around 43% a month), and a more stable inflation starting in 1995, although Brazil still has an annual inflation around 7%. The graph shows that Collor 1 and Collor 2 Plans have failed, since the high inflation level returned shortly after they were implemented.

Figure 1: INTEREST X INFLATION - BRAZILIAN MONTHLY DATA



In this paper, we take the Taylor Rule to characterize empirically the monetary policy before and after the Real Plan. Specifically, we show how the inflation coefficient has changed after the Real Plan was carried out in 1994. This is a natural experiment to test theories surrounding the Taylor Rule in which monetary instability is characterized by an inflation coefficient less than one, whereas monetary stability will have a greater than one coefficient (see Woodford's (2003)).

The rationale for this principle is simple: if inflation grows, the nominal interest rate must grow even faster in order to increase the real interest rate, so as to push consumption and investment down, and therefore to refrain inflation from growing (see Clarida, Galí and Gertler (2000)).

Because the Real Plan has been a success since it was implemented, we would expect the inflation coefficient to be less than one before it had been launched and it would be greater than one after that. However, very suprisingly the results indicate that the inflation coefficient has stayed put around 0.6 even after the monetary stabilization. Various robustness checks were carried out and nothing changed.

After Taylor's (1993) influential paper, many empirical researches about Central Bank's rules have been done. However, they concentrate on large and closed economies like US (see Clarida, Galí and Gertler (2000) and Orphanides (2004) to cite two among many studies)[1]. Only a few researchers have been worried about the Taylor Rule in developing economies, particularly in Brazil. Bueno (2005a) has estimated the Taylor Rule by a multivariate Kalman Filtering model. Bueno (2005b) compares the Taylor Rule in Brazil with EUA using a Markov Switching Regime model to estimate them. Both works provide the same evidence of this paper by finding a less than one inflation coefficient for Brazil. Salgado, *et alli* (2001) estimate the Taylor Rule using a Threshold Autorregressive Model (TAR), but it is difficult to achieve a conclusion because they consider the interest rate as non-stationary.

It is beyond the scope of this paper to offer a formal explanation for this phenomenon, but we provide some free insights in Section 4. Notwithstanding, the implications are very important both theoretically and empirically. First, it shows some gap in theory that deserves further investigation. Second, it suggests that the inflation targeting regime has been uneffective in Brazil confirming a feeling largerly spread among Brazilians.

It would not be a surprise if we could relate this phenomenon to the inflationary memory still living even after 10 years the Real Plan has been put in place. In the 1960s Brazilian policy makers figured out a mechanism against the corrosive effects of inflation, the price indexation, which corrected wages, taxes and other contracts according to previous inflation. With the oil crises of 73 and 79 and with growing public deficits, such an indexation turned out to create a

[1] See Taylor (1999) for a survey on all these studies.

state of high, and far from steady, inflation. For instance in 1989, the monthly inflation reached over 84%. Between 1986 and 1994, several plans tried to knock the inflation down, however all failed but the Real Plan, perhaps because it involved some type of fiscal control.

This work challenges convetional theoretical results. A major monetary stabilization plas was implemented in an important and large country. Price stability was achieved, but the empirical Taylor Rule does not appear to characterize properly the stable period. Pointing out such a fact and asking which pieces are lacking in the model are the main contributions that we provide.

The remainder of this work is organized as follows. Section 2 develops the econometric modeling. Section 3 reports the results obtained by estimating the Taylor Rule. Section 4 analyses the consequences of the results in terms of policy recommendations. The last section concludes.

2 THE MODEL

We follow Clarida, Galí and Gertler (2000) to set up the model. Suppose the Central Bank defines a target rate given by:

$$i_t^* = i^* + g_\pi \left[E_t \left(\pi_{t,k} \right) - \pi^* \right] + g_x E_t \left(x_{t,q} \right), \tag{1}$$

where

$\pi_{t,k}$ is the inflation rate in log terms between periods t and $t+k$;

π^* is the target for inflation;

$x_{t,q}$ is the output gap between t and $t+q$;

E_t is the expectation taken with respect to the information available at t[2];

i^* is the desired nominal rate when both inflation and output are at their target levels.

This rule may be obtained by a macroeconomic model, where the Central Bank maximizes a quadratic loss function in deviations of inflation and output from their respective targets. Taylor (1993) proposes a rule with lagged inflation and output rather than their expected future values.

[2]We assume that $x_{t,0}$ and $\pi_{t,0}$ are fully observable at t. This is because we are using mainly monthly data, which are arguably observable currently.

Eventually, we may collapse into his model, since the rule proposed in equation 1 nests the Taylor rule as a special case.

2.1 BRAZILIAN INTERVENTIONS

We propose to expand that rule in order to consider some important aspects of Brazil. First, during the period at which we are looking there were several stabilization plans. They were designed to smooth price growth and, consenquently, the interest rates[3]. Moreover two very important structural changes were put in place: first the Real Plan in 1994, which caused the change from a high inflation level to a low price growth. Second, the change from fixed to flexible exchange rates in 1999. Such a richness of events serves as a natural laboratory to test theories supposedly designed to work out in different environments and with any particular institutional characteristics as it is our current case. We consider the following dummies during our data range:

Intervention	**Period**	**Measure**
Collor 1	Apr/90 to Aug/90	Abduction of financial assets
Collor 2	Mar/91 to Jun/91	Collor 1 + price control
Real Plan	Jul/94 on	New currency
Exchange	Feb/99 on	exchange fluctuation

The three first plans followed others which had failed to stabilize the price level[4]. In the first two plans, policy makers dried the liquidity of the market by limiting the amount of cash that people could draw out of their own banking accounts. But they failed when the government started to be sued and was forced by courts to release the money. The third plan was preceded of a fiscal adjustment. It constituted of total indexation of the economy to the new unit of value (URV). Once all contracts were indexed and the relative prices so adjusted, the old currency was extinguished and Real replaced URV. Since then the inflation has been less than 1% a month on average. In February, 1999, exchange rate in Brazil became fully flexible.

[3] Later on, we provide more details on this.

[4] The other plans were: Cruzado, from March/86 to Oct./86; Bresser, from July/87 to Sept./87; Summer, from Feb./89 to Apr./89. All those plans had a heterodox character of pegging prices someway. In general, they failed because there was no fiscal control and government always started to issue money again.

2.2 INTERNATIONAL RESERVES AND EXCHANGE RATE

Salgado, *et alli* (2001), following others, argue that the interest rate were also an instrument to control changes in the international reserves. Therefore interest rates reacted to variation of reserves, whose movements reflected the perception of the agents with respect to which level the exchange rate should be while it was pegged. Central Bank of Brazil's reports support this view. Therefore we add variation of reserves and exchange rate and posit the following rule:

$$i_t^* = i^* + g_\pi \left[E_t\left(\pi_{t,k}\right) - \pi^*\right] + g_x E_t\left(x_{t,q}\right) + X_t' \beta, \tag{2}$$

where

X_t stands for a vector with the additional variables like dummies, reserves and exchange[5];

β is a vector of parameters corresponding to these variables.

One may find a theoretical justification for the inclusion of proxies for exchange rate or reserves in Walsh (2003) or Taylor (1999).

2.3 SMOOTHING INTEREST RATE

There are some evidence that the policy reaction function given by equation 1 is too restrictive. Thus Clarida, Galí and Gertler (2000) assume that there is a Central Bank's tendency to smooth variations in the interest rates. There are means of technically justify the presence of the lagged interest rate in the rule (see Woodford, 1999). We simply point out that its presence may improve the stabilization performance of the rule. Thus the *actual* interest rate, i_t, is:

$$i_t = g_i i_{t-1} + \left(1 - g_i\right) i_t^* + \upsilon_t, \tag{3}$$

where

$g_i \in [0, 1]$ indicates the degree of smoothing of the interest rate changes;

υ_t is a zero mean exogenous shock on the interest rate.

Notice that this shock allows for a bit of reality, since the Central Bank does not have a

[5]Both reserves and exchange are in log.

perfect control over the interest rates, as equation 2 posits.

Combining the partial adjustment equation 3 with the target model 2, we find the policy reaction function:

$$i_t = g_i i_{t-1} + (1 - g_i)\left(\mu + g_\pi \pi_{t,k} + g_x x_{t,q} + X_t' \beta\right) + \varepsilon_t, \tag{4}$$

where

$\mu = i^* - g_\pi \pi^*$;

$\varepsilon_t = -(1 - g_i)\left\{g_\pi\left[\pi_{t,k} - E_t(\pi_{t,k})\right] + g_x\left[x_{t,q} - E_t(x_{t,q})\right]\right\} + \upsilon_t$.

In order to collapse into the Taylor's model, we assume $k = q = 0$. The following table maps each variable that we add with its respective coefficient.[6]

Variable	**Coefficient**
Collor 1	β_{C1}
Collor 2	β_{C2}
Real Plan	β_r
Exchange	β_e
Variation in reserves	$\beta_{\Delta R}$
Exchange Variation	$\beta_{\Delta e}$

3 RESULTS

3.1 DATA

In the appendix there is a complete description and some basic statistics of the data that we use in this work. We take the following quarterly data: SELIC Index for interest rate, Price Consumer index - IPCA for inflation and GDP calculated by Applied Research Economics Institute (IPEA). For monthly data, we take the same inflation and interest rate measure, but we use the monthly GDP calculated by the Brazilian Central Bank. It is important to notice that the IPCA is the official inflation rate of the government and it is used to base its monetary policy and inflation targeting.

[6] $\Delta x_t = x_t - x_{t-1}$ is the difference operator;

3.1.1 MONTHLY DATA

Since GDP may be mismeasured, or measured with some degree of delay, and also to check the robustness of the results, we also employ other two leading indicators of economic activity, hopefully better measured: consumption of electrical power in GWh and an index of industrial production, IND. The correlations between these variables are in the following table:

CORRELATION

	GDP	**GWh**	**IND**
GDP	1	0.453	0.379
GWh		1	0.852
IND			1

We have estimated the model by nonlinear least squares with Newey-West robust covariance matrix (see Davidson and Mackinnon, 1993) in order to correct for heteroskedasticity. The results are in the following table, where we have used a quadratic trend as a proxy for potential ouptut. We have chosen GWh as the proxy for ouptut; however, nothing changes qualitatively had we used either monthly GDP or the Production Index (see appendix). Also, if we had used lagged inflation and output gap or linear trend for potential ouptut, the results would not change at all.[7]

We have also tested the model with different samples. In the second column it is the entire sample. In the third it is the pre-Real Plan, when we had a very high inflation. The fourth column is the post-Real Plan sample. And the last column shows the results with flexible exchange rates.

[7] In the appendix, we repeat the regressions considering a closed economy. Nothing changes qualitatively either.

NLS - Newey-West Robust Covariance Matrix - Monthly Data - GWh

Period	90:02-03:12	90:02-94:06	95:01-03:12	00:01-03:12
μ	0.060^{*} (0.022)	0.016 (0.011)	0.018^{*} (0.004)	0.014^{*} (0.001)
g_π	0.663^{*} (0.115)	0.557^{*} (0.067)	0.597^{***} (0.353)	0.003 (0.091)
g_x	-0.063 (0.049)	-0.171^{**} (0.070)	0.004 (0.016)	-0.001 (0.014)
g_i	-0.001 (0.109)	-0.456^{***} (0.236)	0.877^{*} (0.043)	0.775^{*} (0.110)
β_{C1}	-0.071^{*} (0.022)	-0.036^{**} (0.016)		
β_{C2}	-0.036^{**} (0.016)	-0.009 (0.009)		
β_r	-0.042^{**} (0.021)			
β_e	-0.008^{**} (0.003)		-0.007^{**} (0.004)	
$\beta_{\Delta R}$	0.068^{**} (0.032)	0.165^{*} (0.037)	-0.109^{***} (0.061)	-0.008 (0.016)
$\beta_{\Delta e}$	0.140^{***} (0.078)	0.403^{*} (0.069)	0.049 (0.040)	0.000 (0.012)
Adjusted R^2	0.955	0.931	0.882	0.600
Log-likelihood	385.47	118.53	496.74	245.69
Ljung-Box (36)	40.55	19.62	37.02	69.52^{*}
MSE ($\times 10^3$)[8]	1.882		1.169	2.066

* significant at 1%; **significant at 5%; *** significant at 10%
Standard errors are reported in parentheses.

In general, the signs of the coefficients are not always in line with what is expected. Interest rate coefficient is negative before the Real Plan, for instance, although it is non significant.

Since the exchange rate was fixed before 1999, reserves reflected what was happening to exchange terms. Afterwards, the flexible exchange rate was adopted. A depreciation of exchange (R$/US$) implies an increase in inflation and calls for an increase in interest rate. An increase in the reserves comes from a devaluation of the domestic currency pushing out the exportations. Suprisingly, these variables become insignificant in the more recent period.

The analysis shows that the importance of g_i increases after the Real Plan, which is expected. In fact, in all developed economies, it is hardly seen g_i out of the range $[0.6; 0.9]$. This is a reason why it is weird to observe a non-significant g_i when running a regression with the entire sample or less than zero before the Real Plan. It is auspicious finding it significant after the Real Plan,

[8]We left out of the sample the first quarter of 2004. Then we used our estimated results to forecast future interest, and the remained sample as a benchmark testing. Thus, MSE is the quadratic mean square error of the forecasts.

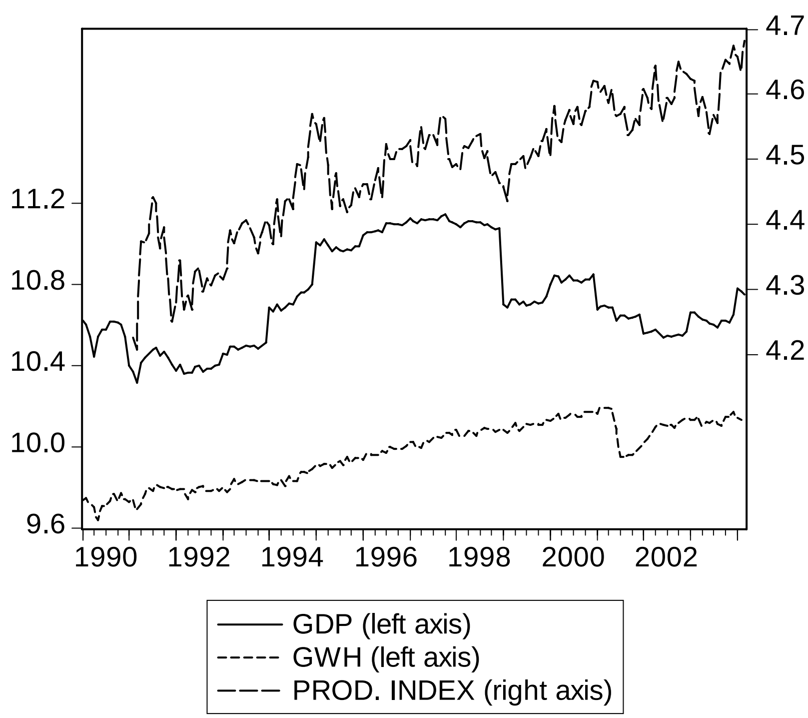
clearly reflecting the monetary stabilization.

One can see that g_π does not increase after the Real Plan and, even worse, it becomes non-significant after the adoption of the inflation-targeting regime (although the results in this case are not fully trustful because of the small number of observations.) Finding $g_\pi < 1$ was not expected at all and may reflect some features that the model was not able to take into account. Moreover, it would suggest us that the price evolution is unstable, although introspection and observation appear to indicate the opposite. The presence of variations in reserves and exchange rate seem not to drive down the inflation coefficient, because we obtain statistically equal coefficients without these variables and because of our robustness checks.

Output gap appears not to be important for determining the interest rate, since is non significant in almost every subsample. But this might explain the regularity observed in the inflation parameter around 0.7 in almost every output proxy that we use.

The robustness of the results are more striking if we consider figure 2, because it depicts series with very different volatilities, but yielding similar qualitative results.

Figure 2: OUTPUT - BRAZILIAN MONTHLY DATA (log)



3.2 QUARTERLY DATA

In order to test the robustness of the results regarding data frequency, we repeated the previous estimations using quarterly data. Since the GDP is better measured in this case, we did not use any other proxy. However, we have varied the output gap proxy (quadratic and linear detrending for potencial output), but the conclusions do not change too much. The following table shows the results for quadratic trend. In the appendix, we present an analysis with linear trend. There are three subsamples: entire sample, pre-Real Plan and post-Real plan.

NLS - Newey-West Robust Covariance Matrix - Quarterly Data - GDP

Period	91:1-03:4	91:1-03:4	95:1-03:4
μ	0.076^{*} (0.016)	0.085^{*} (0.016)	0.054^{*} (0.013)
g_π	0.899^{*} (0.049)	0.972^{*} (0.020)	0.628^{***} (0.338)
g_x	-0.226 (0.205)	-0.296 (0.189)	-0.123 (0.476)
g_i	0.004 (0.010)	-0.013 (0.013)	0.687^{*} (0.090)
β_r	-0.024 (0.016)	-0.033^{**} (0.016)	
β_e	-0.028^{*} (0.006)	-0.027^{*} (0.005)	-0.025^{**} (0.012)
$\beta_{\Delta R}$	0.025 (0.026)		-0.056^{**} (0.027)
$\beta_{\Delta e}$	0.083 (0.055)		0.006 (0.030)
Adjusted R^2	0.996	0.995	0.801
Log-likelihood	125.27	121.86	120.39
Ljung-Box (24)	26.82	16.50	35.55^{**}

* significant at 1%; **significant at 5%; *** significant at 10%
Standard errors are reported in parentheses.

First, let us look at the first regression. It shows g_π less than 1. Variations in reserves and exchange rate seem not to be significant with this data frequency. The coefficient for lagged interest rate is non significant, which may cause some surprise, although it complies with monthly data.

We have tested the jointly hypothesis that $g_\pi = 1, \beta_{\Delta R} = \beta_{\Delta e} = 0$. We reject the null at 10% level[9] using the $F - test$. In fact, in separated tests, Wald rejects $g_\pi = 1$ at 5%, but not

[9] $F = \frac{(RSSR-USSR)/\#restrictions}{USSRd/\#(observations-coefficients)} = \frac{(0.025576-0.021957)/3}{0.021957/43} = 2.3625$, where $RSSR$ is the restricted sum-of-

$\beta_{\Delta R} = \beta_{\Delta e} = 0$. This indicates that g_π must be less than 1. Thus we run the second regression, where we obtain again a $g_\pi < 1$.

In order to check for robusteness of the results, we run the model after the Real Plan. Given the small number of observations, the results are not very trustful, however they point in the same direction as monthly data, that is, g_π much less than one and g_i highly significant around 0.6. Therefore, our evidence with monthly data are reinforced.

4 MONETARY POLICY: EXPLANATIONS AND IMPLICATIONS

Since some kind of theory is lacking to explain the results, we provide insights, subject to be confirmed later. Significative part of investment in Brazil is financed by a public bank - BNDES. This bank charges a lower interest rate, because it is intended to stimulate investments and raise local industries. In order to get a feeling about this, the BNDES' budget is larger than Interamerican Development Bank - IDB. As a result, monetary policy has its effects mitigated by BNDES, because investors do face a lower interest rate than the targeted one by Central Bank.

Another very common explanation is that some prices are insensitive to interest rate. Many important prices that compose the consumers' budget are ruled over by government, like utilities' prices. Privatized firms must follow government agencies' instructions and generally have their contracts indexed to a price index. Therefore, even increasing the interest rate, these prices would not fall. Some people might suggest ruling out such prices from the price index. However, the target inflation includes such prices and people do bear them, controlled or not.

Most of the public deficit is financed by public bonds. Because of the high inflationary period that Brazil lived, most of them are indexed by price indexes. Therefore, an increase in the interest rate has the effect of increasing the public debt and private wealth. Since most of the debt is hold by banks, it seems to have a positive income effect higher than the substitution effect. Then the net effect is to increase the supply of credit, mitigating the expected monetary

squares residuals and $USSR$ is the unrestricted sum-of-squares residuals. (See Davidson and MacKinnon, 1993, ch. 6).

policy effect.

Still another explanation is that Brazilian consumers are very impatient. They would be so because of the high inflation they faced with, in such a way that it was more advantageous to consume sooner in order not to loose purchasing power. They would prefer consuming more today than saving, although they would have to pay a high interest rate. Because the borrowing real interest rate is already very high, an increase in the interest rate would have a marginal effect over the credit and therefore almost no effects on consumption. In other words, the demand for credit is in a very inelastic part of the curve making unfeasible the monetary policy.

All these considerations help extracting some major implications from our findings: The inflation-targeting regime in Brazil appears to have very low power to control inflation, whatever the theoretical explanation for that behavior be. Indeed while we are writing these notes (June, 2005) it is a common sense in media and among people that Central Bank of Brazil has pushed up interest rate too much. Notwithstanding the inflation has not fallen. Thus, if there exists a limit for the effectiveness of such a policy, then Brazil must have achieved it.

Our results indicate a puzzle from the theoretical point of view. Brazil has some monetary stability, applied a stabilization plan, however the inflation coefficient did not change, or, at least, did not increase enough to be greater than one. Then, there must exist some theoretical gap to be bridged.

Perhaps, the public deficit should be considered, since in Brazil expenditures plus interest payments are greater than revenues. This happens mainly because the government budget presents some rigidities due to the close tie between revenues and expenditures with education, health and social security. Thus in order to make policy effective, it would bem necessary further decrease in government expediture.

5 CONCLUSION

In this paper, we have used the Taylor Rule to analyse Brazilian monetary policy and its effects on the inflation. We have found that, despite the adoption of a inflation-targeting regime and despite observing a period with high inflation followed by one stable, the inflation coefficient of the Taylor Rule never was greater than one. This is an unexpected result because it suggests the existence of a monetary instability when it is clear that Brazil has prices under control. Our results are quite robust with respect to different samples, lags of variables, proxies for GDP, proxies for potential GDP, data frequency and even with respect to econometric methods (see Bueno (2005a, 2005b)).

Our findings are fundamental for two main reasons: First, Brazil, the biggest country in Latin America, experienced a long time with a high inflation level and succesfully implemented a heterodox plan to control it. This is a major contribution for economic policy as a whole. As a matter of fact, that plan eventually put inflation at civilized levels. Therefore it is a natural experiment to test theories and study their effects upon the economy. This fact is even more important if we recall that Brazil has a GDP of over US$ 400 billion/year and 200 million people.

The second reason is theoretical. They indicate an unexpected result. Therefore, it is necessary to look for a model which could explain the behavior that we have found. That is, because of the high inflation levels before the stabilization, an important empirical evidence is challenging theories and policy recommendations for stabilization.

One might question whether our results are not due to mismeasurement. We argue that this is not the case. The only variable that might be measured with error is the output gap. However, we do not find qualitative changes when we vary output, output gap measure, data frequency or econometric method. In fact all these measures are very distinct and by poiting in the same direction they reinforce our conclusions.

Although it is not the focus of this paper, we have provided some free insights about possible explanations for the phenomenon found here. It is clear the need of further investigation regarding the fundamentals of the Taylor Rule. In Brazil, empirical evidence matches the spread feeling among people about the ineffectiveness of the monetary policy with inflation targeting regimes.

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APPENDIX A: DATA DESCRIPTION

The data used in this work is described now.

BRAZIL - DATA DESCRIPTION

Series	GDP	IPCA	SELIC	Cons. Energy	PINDEX	Reserves	Exchange
	q	π	i	GWh	IND	$\mathrm{Re}\,s$	d
Source	BCB	IPEA	IPEA	Eletrobrás	IBGE	BCB	BCB
Seas. Adj.	YES	YES	NA	YES	YES	NA	NA
Freq.	Q/M	M	M	M	M	M	M
Units	B 2003 Real	1990 = 100	%	GWh	2002 = 100	US$ B	R$/US$
Range	1991-2003	1947-2003	1974-2004	1979-2004	1991-2004	1970-2005	1990-2004

BCB	Central Bank of Brazil	NA	Not applicable
IPCA	Consumer Price Index	PINDEX	Production Index
IPEA	Applied Research Economics Institute	IBGE	Brazilian Institute of Statistics
SELIC	Effective Federal Funds Rate		

Some basic statistics about these variables are presented in the following table.

BASIC STATISTICS - BRAZIL

Quarterly	π	q	i
Mean	0.20	26.53	0.24
Std. Dev.	0.31	0.10	0.33
Skewness	1.39	-0.54	1.48
Kurtosis	3.36	2.03	3.70
# **Obs.**	52	52	52

We proceed the Phillips-Perron unit root test with trend and intercept, unless otherwise noticed:

UNIT ROOT TESTS - Quarterly - BRAZIL

Variable	**Levels**	**First Differences**
q	-2.70	$-6.16^{*,b}$
π	-2.25	$-5.21^{*,b}$
i	-2.61	$-6.96^{*,b}$

* - Significant at 1% level

b - Without trend

Although we do not report results, we carried out Johansen's cointegration test and whatever is the hypothesis regarding trend and intercept, linear or quadratic, there are always at least one cointegrating vector.[10]

For the monthly data, we have.

BASIC STATISTICS

Monthly	π	i	q	GWh	IND
Mean	0.08	0.09	10.73	9.96	4.49
Std. Dev.	0.12	0.12	0.24	0.15	0.10
Skewness	1.81	1.66	0.31	−0.23	−0.53
Kurtosis	6.15	5.12	1.85	1.67	2.74
# **Obs.**	171	171	171	171	158

We proceed the Phillips-Perron unit root test with trend and intercept, unless otherwise noticed:

UNIT ROOT TESTS - Monthly - BRAZIL

Variable	**Levels**	**First Differences**
q	-1.36	-12.59^{b,*}
π	-4.30*	–
i	-4.36*	–
GWh	-2.79	-14.28^{b,*}
IND	-5.69*	-

* - Significant at 1% level

b - Without trend

[10]The tests here may be misleading, because of the structural breaks (see Cati, Garcia and Perron, 1988), but we abstract from that on the basis that inflation and interest rate must be cointegrated and hence there must exist a cointegrating vector between them.

APPENDIX B: CLOSED ECONOMY

NLS - Newey-West Robust Covariance Matrix - Monthly Data - GWh

Period	90:02-03:12	90:02-94:06	95:01-03:12	00:01-03:12
μ	0.078* (0.028)	0.073* (0.023)	0.019* (0.003)	0.014* (0.001)
g_π	0.737* (0.124)	0.743* (0.101)	0.410*** (0.247)	0.020 (0.083)
g_x	−0.067 (0.054)	−0.373** (0.164)	0.002 (0.013)	−0.002 (0.013)
g_i	0.027 (0.099)	−0.004 (0.237)	0.833* (0.053)	0.765* (0.107)
β_{C1}	−0.079* (0.022)	−0.071* (0.020)		
β_{C2}	−0.047** (0.019)	−0.043** (0.018)		
β_r	−0.059** (0.027)			
β_e	−0.008* (0.003)		−0.008** (0.003)	
Adjusted R^2	0.947	0.813	0.861	0.568
Log-likelihood	371.99	91.07	486.68	245.51
Ljung-Box (36)	49.49***	23.35	28.65	72.93*
MSE ($\times 10^3$)	5.211		1.781	2.432

* significant at 1%; ** significant at 5%; *** significant at 10%

Standard errors are reported in parentheses.

APPENDIX C: OTHER PROXIES FOR OUTPUT

NLS - Newey-West Robust Covariance Matrix - Monthly Data - GDP

Period	90:02-03:12	90:02-94:06	95:01-03:12	00:01-03:12
μ	0.054*** (0.022)	0.003 (0.007)	0.019* (0.004)	0.015* (0.001)
g_π	0.685* (0.103)	0.575* (0.062)	0.536*** (0.298)	0.025 (0.104)
g_x	−0.028 (0.019)	−0.055* (0.015)	−0.008 (0.019)	0.016 (0.024)
g_i	−0.001 (0.109)	−0.497*** (0.254)	0.869* (0.047)	0.792* (0.105)
β_{C1}	−0.061* (0.022)	−0.015 (0.015)		
β_{C2}	−0.033** (0.015)	−0.002 (0.008)		
β_r	−0.033*** (0.019)			
β_e	−0.014** (0.006)		−0.008** (0.004)	
$\beta_{\Delta R}$	0.067** (0.032)	0.148* (0.035)	−0.102*** (0.058)	−0.009 (0.017)
$\beta_{\Delta e}$	0.130*** (0.077)	0.418* (0.067)	−0.049 (0.038)	0.002 (0.014)
Adjusted R^2	0.955	0.935	0.882	0.555
Log-likelihood	385.55	119.99	496.78	245.92
Ljung-Box (36)	34.95	22.29	36.77	73.05*
MSE ($\times 10^3$)	9.344		1.147	1.917

* significant at 1%; **significant at 5%; *** significant at 10%

Standard errors are reported in parentheses.

NLS - Newey-West Robust Covariance Matrix - Monthly Data - IND

Period	90:02-03:12	90:02-94:06	95:01-03:12	00:01-03:12
μ	0.026 (0.019)	0.003 (0.015)	0.018* (0.005)	0.014* (0.002)
g_π	0.873* (0.059)	0.164 (0.102)	0.636 (0.420)	0.028 (0.118)
g_x	0.046 (0.063)	0.002 (0.073)	0.034 (0.085)	0.031 (0.047)
g_i	0.137 (0.130)	0.387** (0.160)	0.886* (0.051)	0.820* (0.117)
β_{C2}	−0.015 (0.014)	0.004 (0.015)		
β_r	−0.012 (0.017)			
β_e	−0.006*** (0.003)		−0.007** (0.004)	
$\beta_{\Delta R}$	0.031 (0.022)	0.011 (0.061)	−0.116 (0.070)	−0.009 (0.018)
$\beta_{\Delta e}$	0.102 (0.072)	0.917* (0.113)	−0.051 (0.042)	0.001 (0.015)
Adjusted R^2	0.981	0.980	0.883	0.556
Log-likelihood	432.10	126.25	496.89	245.96
Ljung-Box (36)	35.14	27.87	37.14	68.84*
MSE ($\times 10^3$)	2.592		1.039	1.972

* significant at 1%; ** significant at 5%; *** significant at 10%
Standard errors are reported in parentheses.

APPENDIX D: LINEAR TREND

NLS - Newey-West Robust Covariance Matrix - Linear Trend Quarterly Data

Period	91:1-03:4	91:1-03:4	95:1-03:4
μ	0.064* (0.013)	0.064* (0.013)	0.061* (0.016)
g_π	0.901* (0.052)	0.903* (0.051)	0.616*** (0.364)
g_x	−0.293*** (0.147)	−0.348** (0.159)	−0.295 (0.309)
g_i	0.003 (0.010)	−0.005 (0.009)	0.677* (0.077)
β_r	−0.006 (0.014)	−0.005 (0.014)	
β_e	−0.037* (0.008)	−0.038* (0.008)	−0.034** (0.016)
$\beta_{\Delta R}$	0.020 (0.023)		−0.059** (0.027)
$\beta_{\Delta e}$	0.092*** (0.053)	0.091*** (0.052)	0.016 (0.028)
Adjusted R^2	0.996	0.996	0.807
Log-likelihood	125.83	125.26	120.87
Ljung-Box (24)	25.61	22.53	32.42*

* significant at 1%; ** significant at 5%; *** significant at 10%
Standard errors are reported in parentheses.