THE TAYLOR RULE UNDER INQUIRY: Hidden states[1]

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Abstract

This work empirically evaluates the Taylor rule for the US and Brazil using Markov-Switching Regimes. I find that the inflation parameter of the US Taylor rule is less than one in many periods, contrasting heavily with Clarida, Galí and Gertler (2000), and the same happens with Brazilian data. When the inflation parameter is greater than one, it encompasses periods that these authors considered they should be less than one. Brazil is used for comparative purposes because it experienced a high level inflation until 1994 and then a major stabilization plan reduced the growth in prices to civilized levels. Thus, it is a natural laboratory to test theories designed to work in any environment. The findings point to a theoretical gap that deserves further investigation and show that monetary policy in Brazil has been ineffective, which is coherent with the general attitude of population in relation to this measure.

Key Words: Markov-Switching Regimes, Hidden States, Taylor Rule

JEL: E52, C32, C51

1 INTRODUCTION

Taylor (1993) describes US monetary policy using a simple linear function linking nominal interest, i_t, to inflation, π_t, and to output gap, x_t. Since then, a number of authors have established the economic fundamentals that support such an empirical finding, known as the Taylor Rule, roughly defined as follows: $i_t = g_\pi \pi_t + g_x x_t$. Taylor (1999) contains several empirical studies regarding it, and Woodford (2003), among others, was able to derive optimal rules that are closer versions of Taylor's.

Clarida, Galí and Gertler (2000) show that the inflation parameter of the Taylor Rule varies with different samples. In particular, the inflation parameter is greater than one after Volcker became chairman of the FED, and it is less than one before that event. Their results depend crucially on their choice regarding when to break the series. They make such a break to characterize a shift in the monetary policy. However, by varying the date the results may be modified, as well as being always arbitrary due to the exogenous determination of this factor.

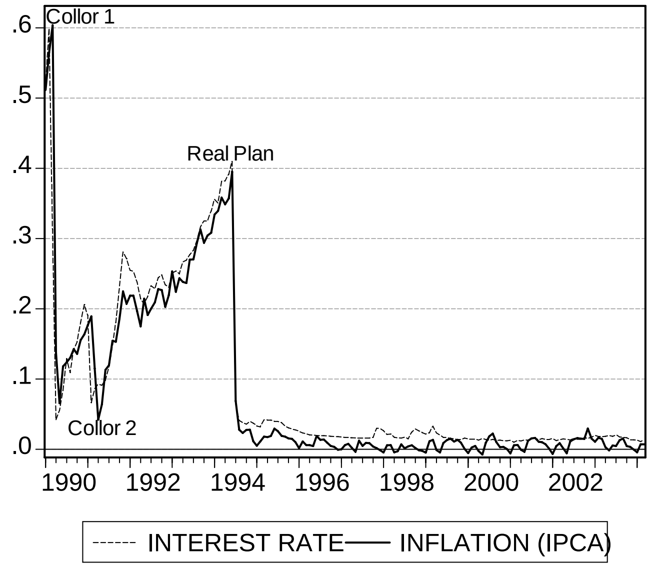
In this paper I use the Markov Switching Regime framework to estimate the regimes that the monetary policy may have experienced. The procedure is appropriate because the choice of the regimes is endogenous to the econometric model, mitigating the criticisms placed on Clarida, Gali and Gertler. We can then check the reliability of their assumptions and estimates. It is also interesting because one may access variations in regimes and verify how policy makers have changed their behavior, given that the econometrician does not directly observe such changes.

We apply the same procedure for Brazil for several reasons. The first is that the countrysuffered a very pronnounced latent hyperinflation period. However, in 1994 a monetary stabilization plan, the so-called Real Plan, was launched and brought inflation down to civilized levels. This successful plan clearly divided the regimes in Brazil and may be easily observed in figure 1.

Before the Real Plan, others were implemented, but failed, and I label them as Collor 1 and Collor 2 in the figure. Therefore, the procedure should divide clearly both periods, and this is indeed what happens as I demonstrate later on (see figure 5).

However, the same is not true for the US, at least not in the terms proposed by Clarida, Galí and Gertler. Therefore, I suggest that their conclusions were driven by the choice made regarding the date break and offer contrasting results and conclusions. In particular, I argue

Figure 1: Interest x Inflation - Brazilian Monthly Data



that there were a few periods in which extreme values have influenced the inflation parameter, forcing it above one. To support this claim, I rely on Bueno's (2005a) data set, who finds very similar conclusions to Clarida, Galí and Gertler by using a Kalman Filter and nonlinear least squares estimation procedures.

Notwithstanding, it is a fact that the US is a mature economy with no price indeterminacy. Therefore it is fair to ask why I arrive at such conclusions. That gives us the second reason to put Brazil in this paper.

After the Real Plan, inflation fell dramatically and has since fluctuated around 7% per annum. No one would claim that Brazil currently experiences high inflation levela or has uncontrolled prices, which would signify price indeterminacy. However, I show in this study that even after the Real Plan, the inflation parameter in Brazil has not increased and exceeded one (which means theoretically price determinacy) as would expected after such a plan. This astonishing finding supports Bueno (2005a, 2005b), which varies econometric methods, data frequency, proxies for output and output gap and mantains these conclusions. Therefore, by contrasting Brazil with the US, I am able to argue that there is some unexplained gap in theory which deserves more attention from subsequent researchers.

The remainder of this work is organized as follows: Section 2 discusses the econometric Taylor rule model that I use in the estimations; Section 3 reports the results using Markov-Switching regimes for the US; and Section 4 applies the same model to Brazilian data. The last section serves a conclusive point to this study.

2 THE TAYLOR RULE AND MARKOV SWITCHING REGIMES

I follow Clarida, Galí and Gertler (2000) to formulate the model. First, suppose the Central Bank defines a target rate given by:

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

where

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

$\pi_{s_t}^*$ 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[1];

$i_{s_t}^*$ is the desired nominal rate when both inflation and output are at their target levels;

$s_t \in \{1, 2, \ldots, M\}$ represents the prevailing regime at time t.

If one posits that the Central Bank maximizes a quadratic loss function in deviations of inflation and output from their respective targets, one may obtain this rule. Equation 1 nests the Taylor rule as a special case, when it is assumed that $k = q = -1$.

The process or rule is time-invariant conditional on an unobserved state or regime. The usual assumption for Markov-Switching regimes is that the unobserved realizations of the regime $s_t \in \{1, 2, \ldots, M\}$ follow a discrete time, discrete state Markov stochastic process, defined by the

[1]I assume that $x_{t,0}$ and $\pi_{t,0}$ are fully observable at t.

transition probabilities:

$$p_{ij} = \Pr\left(s_{t+1} = j \mid s_t = i\right), \sum_{j=1}^{M} p_{ij} = 1, \forall i, j \in \{1, 2, \ldots, M\},$$

where Pr stands for probability.

It is assumed that s_t follows an irreducible ergodic M state Markov process with the transition matrix P_M, where M obviously indicates the number of regimes. We refer to Hamilton (1994) and Krolzig (1997) for details regarding the statistical properties of the model. I use the expectation maximization algorithm to estimate the unobserved states and calculate the probabilities of being in each state.

2.1 SMOOTHING THE INTEREST RATE

Again, inspired by Clarida, Galí and Gertler (2000), I also assume that there is a Central Bank tendency to smooth variations in the interest rates, and therefore add lagged interest rates in the rule. Their presence may improve the stabilization performance of the rule. Thus the *actual* interest rate, i_t, is:

$$i_t = g_{i,s_t} i_{t-1} + \left(1 - g_{i,s_t}\right) i_t^* + \upsilon_t, \tag{2}$$

where

$g_{i,s_t} \in [0, 1]$ indicates the degree of smoothing of the interest rate changes;

$\upsilon_t \sim NID\left(0, \sigma_{s_t}^2\right)$ is a zero mean exogenous shock on the interest rate, whose variance also changes through time.

Notice that this shock allows for more realistic conditions, since the Central Bank does not have perfect control over the interest rate, as equation 1 posits.

Combining the partial adjustment equation 2 with the target model 1, we find the policy reaction function:

$$i_t = a_{s_t} + g_{i,s_t} i_{t-1} + \left(1 - g_{i,s_t}\right)\left(g_{\pi,s_t} E_t\left(\pi_{t,k}\right) + g_{x,s_t} E_t\left(x_{t,q}\right)\right) + \upsilon_t, \tag{3}$$

where

$a_{s_t} = (1 - g_{i,s_t}) \left(i^*_{s_t} + g_{\pi,s_t}\pi^*_{s_t}\right)$.[2]

It is noteworthy that the Markov-Switching Regimes method allows for non-linearities in the model because there are sudden changes, which are discrete. Indeed, this family of models is very flexible, making it extremely attractive. For example, one might keep the intercept and/or the variance of each regime fixed. Or one could vary them, and keep the other parameters fixed. What we must bear in mind, undoubtedly, is that the more flexible they are, the more parameters for estimation, which may bring out a number of practical issues. This is connected with the definition of the number of regimes that we should choose, estimation time and other concerns. For the purposes of this paper, I will assume that $k = q = 0$, and let every parameter free, in order to obtain the most flexible model.

3 EMPIRICAL RESULTS - USA

3.1 DATA

In the appendix there is a complete description and some basic statistics of the data that I use in this study.

For the US we use the Effective Federal Funds Rate as interest rate, the output gap calculated by the Congressional Budget Office (see Arnold, 2004), and the Personal Consumption Expenditure Index is used for inflation. I briefly use the data set of Orphanides (forthcoming) as proxy for expected inflation and output.[3]

3.2 ESTIMATES

It is not very well agreed on how one should define the number of regimes and which model to use. I estimate up to five regimes to access how the parameters change and what the shape of the regimes assumes, given the most flexible model at my disposal. Sims and Zha (2004) estimate a system with more than 8 regimes, but I think this is excessive.

[2] This parameter is not identified at all. However, I assume that g_{i,s_t}, $i^*_{s_t}$, g_{π,s_t} and $\pi^*_{s_t}$ do not vary in such a way that $a_{s_t} = \mu$, where μ is some constant. That means that by finding a_{s_t} varying over time, then there must exist a time-varying coefficient.

[3] I am grateful to Athanasios Orphanides, for kindly making his data set available.

Parameter/Regime	**R1**	**R2**
a	2.29×10^{-3}	3.37×10^{-3}
	(2.75×10^{-3})	(0.002)
g_i	0.901^{*}	0.908^{*}
	(0.027)	(0.087)
g_π	1.416^{*}	1.700
	(0.271)	(1.553)
g_x	0.266^{*}	0.381
	(0.069)	(0.559)
Std. Deviation ($\times 10^3$)	1.019	4.045
Duration	41.02	14.01
Prob.	0.746	0.254
Log-Likelihood	884.76	
Schwarz Criterion	-9.70	Volcker

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

Table 1: US - GDP - QUARTERLY DATA - TWO REGIMES

In this section, I analyze the cases with 2 and 4 regimes, which give us the same general conclusions had we used all models. In the appendix, the remaining models are detailed. I recall that we have used the CBO proxy for calculating the output gap. The estimates with 2 regimes are reported in table 1.

We observe first that both intercepts are statistically insignificant, and the coefficients of the lagged interest rate are very similar. This shows no time-varying intercept, confirming what Bueno (2005a) has encountered. Surprisingly, in the second regime, both the inflation and output gap parameters are statistically insignificant, but in the first regime they have reasonable values.

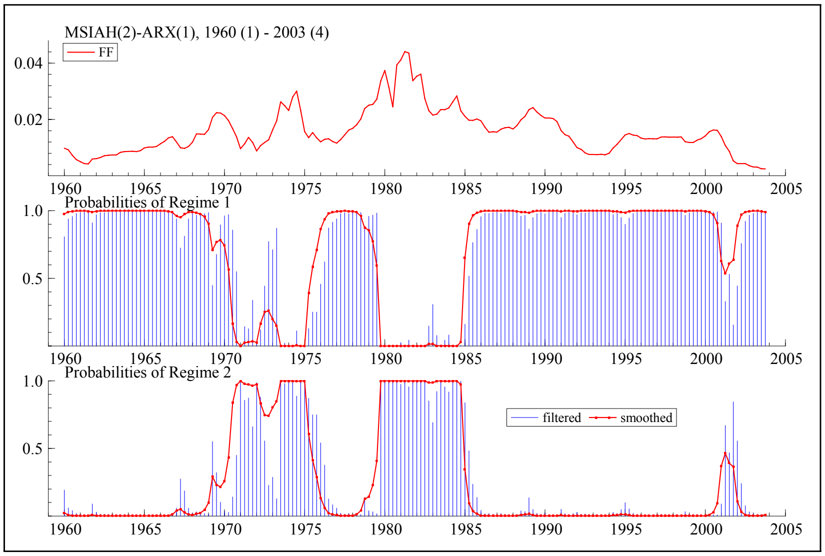
What is very interesting to observe is how these regimes are spread over time by looking at figure 2.

We observe a very well characterized period after Greenspan, in which the inflation coefficient is superior to 1. However, this same behavior characterizes the 1960's and the second half of the 1970's, diverging from Clarida, Galí and Gertler's (2000) assumption. This clearly suggests that the sample break proposed by the aforementioned is not that appropriate.

The other regime clearly characterizes the Vietnam War and the Volker-era, where the coefficients are not significant. Thus, two things must be noticed from now on. The first is that the last regime, which we are labelling Volcker-era, is always very well characterized,[4] regardless the number of regimes that we impose. Accordingly, when I increase the number of regimes, the first

[4]The Volcker's regime is pointed in the bottom of each table.

Figure 2: TWO REGIMES PROBABILITIES - US



regime is split into others. The second observation is that I always encounter non-significant coefficients for the Volcker-era, as the following table in this section and the tables in the appendix confirm. Looking at figure 3, we can observe how well the model fits the data.

If I increase the number of regimes to 4, I find weak evidence of time-varying intercept, because, in general, they are very low and statistically insignificant. Although the inflation parameter is significant in the intermediary regimes, it is greater than one only for the second regime, which has the shortest duration and lowest probability.[5] As a matter of fact, considering what other empirical studies have reported, such as Orphanides (forthcoming), it is absolutely unexpected to encounter many periods since the 1960's in which the inflation coefficient is non-significant or less than one.[6] These results can be easily confirmed in table 2.

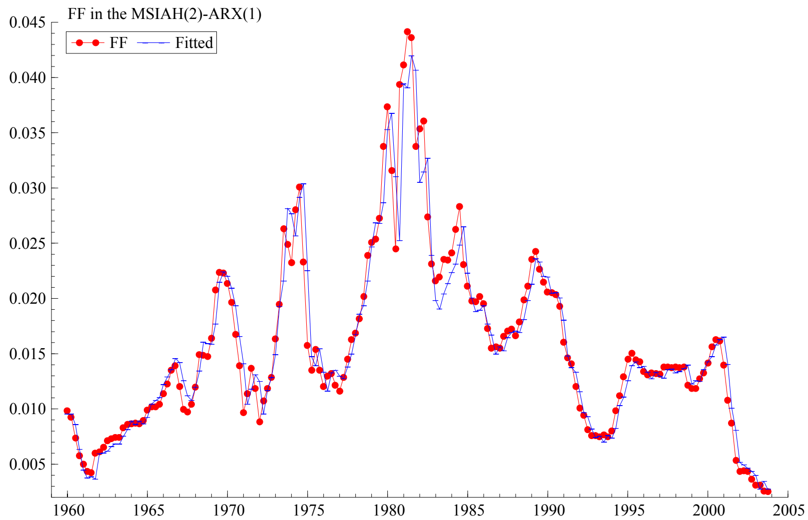
The output gap parameters have reasonable values and are significant in general. They reveal different intensities, but always some effort to reduce output gap variability in compliance with the Kalman Filter estimates made by Bueno (2005a).

Figure 4 shows the temporal distribution of the regimes. Recall that in Regime 1, the inflation parameter is non-significant and negative, and in regime 3 it is less than one; therefore, if one

[5]This same pattern is found with 5 regimes.

[6]With 3 regimes, this coefficient is greater than one with half probability. However we find it during most of the 1970's, unexpectedly

Figure 3: MODEL FIT WITH 2 REGIMES - US



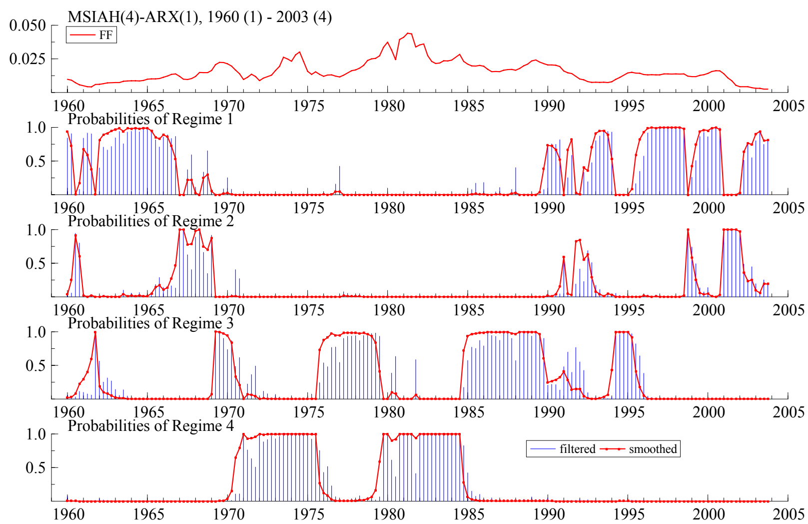
Parameter/Regime	**R1**	R2	**R3**	**R4**
a	2.00×10^{-3} $(2.00{\times}10^{-3})$	8.00×10^{-3} $(8.00{\times}10^{-3})$	0.002^{*} (0.001)	3.00×10^{-3} (0.002)
g_i	0.949^{*} (0.021)	0.735^{*} (0.076)	0.741^{*} (0.037)	0.924^{*} (0.090)
g_π	-0.259 (0.567)	1.243^{*} (0.372)	0.488^{*} (0.126)	1.417 (1.492)
g_x	0.400^{**} (0158)	0.128^{*} (0.042)	0.351^{*} (0.044)	0.556 (0.884)
Std. Deviation ($\times 10^3$)	0.396	0.852	0.839	4.092
Duration	6.29	3.35	8.82	17.84
Prob.	0.335	0.137	0.292	0.236
Log-Likelihood	923.19			
Schwarz Criterion	-9.55			Volcker

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

Table 2: US - GDP - QUARTERLY DATA - FOUR REGIMES

takes into account that Greenspan has been leading the FED throughout most of these regimes, one may be surprised. We also notice that it overlaps large portions of the findings of Sims and Zha (2004, p. 22). In particular, my regimes 1, 3 and 4 correspond roughly to their regimes 1, 2 and 3, respectively .

Figure 4: FOUR REGIMES PROBABILITIES - US



Generally speaking, the more regimes we allow for, then either the inflation coefficient greater than one is statistically non-significant, or less than one or, if significant and greater than one, it demonstrates the shortest duration and lowest probability. Therefore, we argue that what leads to a greater than one parameter are the few extreme observations that conceivably bias the results toward a higher absolute value. This is a common problem discussed in the literature and may be found, for instance, in the first chapter of Davidson and MacKinnon (1993). The results also seem to be in disagreement with Orphanides (forthcoming), but clearly show that there were regime switches in the US monetary policy, as documented by Sims and Zha (2004).

The findings here put under inquiry the results obtained by both Clarida, Galí and Gertler (2000) and Orphanides (forthcoming), because they diverge from them. In particular, I estimated the Markov Switching Model with the same data set of Orphanides (forthcoming), assuming $q = 0$ and $k = 1$. Orphanides' study is interesting because he has collected expected inflation reported in the Greenbook, where the FED reveals its expectations about inflation and potential output.

Hence, we can use directly observed expectations to estimate the model.

The tables are located in the appendix, but the analysis is discussed here. With two regimes, the qualitative findings are identical to mine; with 3 regimes, the results start to deviate from reasonable. For example, I find $g_\pi < 0$ in the third regime, or very high, although non-significant in the second regime. With 4 regimes, I find $g_i > 1$, and g_π either large, but non-significant, or less than zero. In other words, with three and four regimes, Orphanides' data demonstrates a $g_\pi > 1$ and significant in only one of the regimes. The shape of the pictures vary with respect to mine, however they overall clearity of the last regime as representing the Volcker-era.

The general conclusion is that the theories supporting the Taylor Rule need to be further developed to absorb the empirical results reported here. Anyone would disagree with the claim that the US has indeterminate monetary dynamics. Thus, the puzzle is not in the data set, since Bueno (2005a) has obtained similar results as Clarida, Galí and Gertler (2000) using two other estimation methods and the same sample break. Rather, this seeming incompatible finding, that is the inflation coefficient less than one for most time of the sample, should refer to some theoretical gap.

The results are even more striking if we observe that after the stabilization plan in Brazil, the inflation parameter did not increase to one or above, as I will discuss in the next section.

4 EMPIRICAL RESULTS - BRAZIL

4.1 BRAZILIAN INTERVENTIONS

I suggest the expansion of the section 2 rule in order to take into account some important aspects of Brazil. First, there were several stabilization plans during the period under analysis. They should smooth price growth and, thus, the interest rates as I discuss in depth later on. There were also two important structural changes in Brazil: first the Real Plan in 1994, which caused the change from a high inflation level to one of low price growth; second, the exchange rate modification from fixed to flexible rates in 1999. This envrionment is natural for experiments to test theories assumed to work in distinct economic settings and with particular institutional features, as is the case currently. Therefore, I consider the following dummies within my 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

Collor 1, Collor 2 and Real plans followed others aimed to stabilize the price level and failed.[7] Policy makers dried the liquidity of the market freezing the amount of cash that people could draw out of their own banking accounts in the first two plans. However, government started to be sued and was forced by courts to release the money, making them to fail. A fiscal adjustment preceded the third plan, which constituted a 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.

4.2 INTERNATIONAL RESERVES AND EXCHANGE RATE

Some, like Salgado, *et alli* (2001) among others, claim that the interest rate was also an instrument to control changes over international reserves, such that interest rates responded to reserve variations, which reflected the perception of the agents concerning which level the exchange rate should be while it was fixed. Central Bank of Brazil reports appear to support this opinion. Therefore, I add variation of reserves and exchange rate and assume the following rule:

$$i_t^* = i_{s_t}^* + g_{\pi,s_t}\left[E_t\left(\pi_{t,k}\right) - \pi^*\right] + g_{x,s_t} E_t\left(x_{t,q}\right) + X_t' \delta_{s_t}, \tag{4}$$

where

X_t stands for a vector with the additional variables like dummies, reserves and exchange[8];

[7] 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.

[8] Both reserves and exchange are in log.

Variable	Coefficient
Collor 1	β_{C1,s_t}
Collor 2	β_{C2,s_t}
Real Plan	β_{r,s_t}
Exchange	β_{e,s_t}
Variation in reserves	$\beta_{\Delta R,s_t}$
Exchange Variation	$\beta_{\Delta e,s_t}$

Table 3: CORRESPONDENCE BETWEEN COEFFICIENTS AND VARIABLES

δ_{s_t} is a vector of parameters corresponding to these variables.

Walsh (2003) or Taylor (1999) may provide a theoretical justification for the inclusion of such variables.

Thus, combining the partial adjustment equation 2 with target model 4, I find the policy reaction function:

$$i_t = a_{s_t} + g_{i,s_t} i_{t-1} + (1 - g_{i,s_t}) \left(g_{\pi,s_t} E_t (\pi_{t,k}) + g_{x,s_t} E_t (x_{t,q}) \right) + X_t' \beta_{s_t} + \upsilon_t, \tag{5}$$

where

$a_{s_t} = (1 - g_{i,s_t}) \left(i^*_{s_t} + g_{\pi,s_t} \pi^*_{s_t} \right)$;

$\beta_{s_t} = (1 - g_{i,s_t}) \delta_{s_t}$[9].

In order to collapse into the Taylor's model, I assume $k = q = 0$. Table 3 maps each variable that I add with its respective coefficient.[10]

4.3 DATA

For Brazil, I used the following quarterly data: SELIC Index for interest rates, Personal Consumption Expenditure - IPCA for inflation and GDP. For monthly data, the same inflation and interest rate measures are assumed, but I use the monthly GDP calculated by the Brazilian Central Bank. As alternative proxies for this measure, I employ the consumption of Energy in GWh and an Industrial production Index.

[9] Since the focus of this paper is not on the δ parameters, it is unnecessary to indentify them.

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

4.4 QUARTERLY DATA

I perform the same exercise with Markov-Switching for Brazil and report the results in this section. The only model that I was able to estimate with quarterly data had 2 regimes. The other models often did not converge, or presented overflow problems because of the restricted number of observations available. Thus, later on I employ monthly data to check the robustness of the results here. The output gap is obtained by quadratic detrending.

By looking at figure 5, one can perceive that the method effectively divides the periods pre- and post-Real, which is highly expected from the method. The first regime represents the post-Real Plan era, and the second, the pre-Real era. The transition between periods is nicely depicted by the smoothed line. Moreover, from the figure, one can conclude with the maintenance of monetary policy over subsequent governments: from 1995 to 2002 with President Fernando H. Cardoso, and from 2003 on, with President Luís I. Lula da Silva. This, of course, is in line with the opinion of Brazilian people.

Figure 5: TWO REGIMES PROBABILITIES - QUARTERLY - GDP - BRAZIL

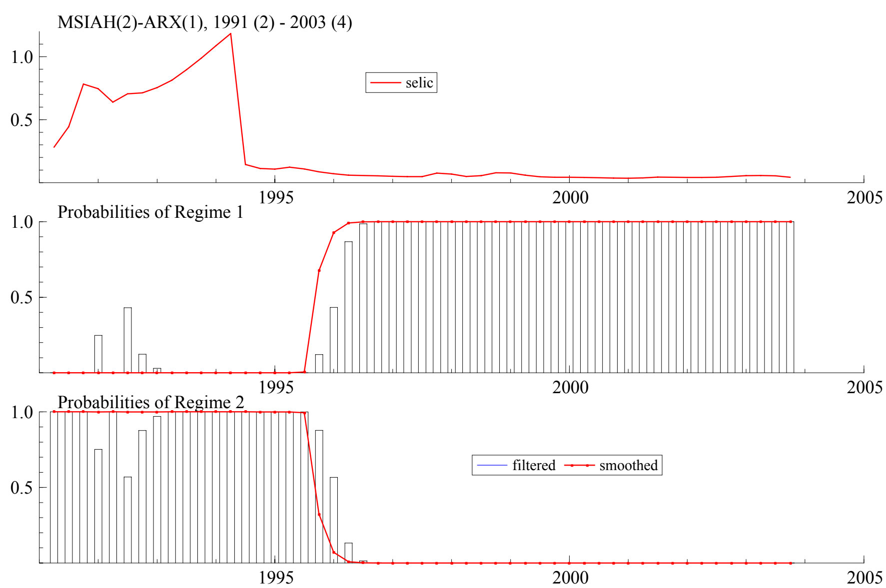


Table 4 shows the estimates.[11] It is a surprise to observe the inflation parameter to be non-

[11]The diganostic test of residuals showed no remaining autocorrelation.

Parameter/Regime	**R1**	**R2**
a	0.398 (6.55)	0.029 (0.018)
g_i	0.356* (0.123)	0.084* (0.026)
g_π	0.158 (0.131)	0.513* (0.098)
g_x	−0.484* (0.145)	0.110 (0.243)
$\beta_{\Delta e}$	0.007 (0.012)	0.486* (0.083)
$\beta_{\Delta R}$	−0.010 (0.009)	0.043*** (0.024)
β_r	−0.359 (6.558)	0.029*** (0.017)
β_e	−0.012* (0.004)	−0.017 (14.70)
Std. Deviation ($\times 10^3$)	6.80	15.86
Duration	∞	19.35
Prob.	1.00	0.000
Log-Likelihood	162.82	
Schwarz Criterion	−4.84	

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

Table 4: BRAZIL - GDP - QUARTERLY DATA - TWO REGIMES

significant and less than one after the Real Plan, while before it was less than one (as theory predicts), due to the presence of a sharp inflationary period.

Because of the small number of observations, I repeat the exercise with monthly data.

4.5 MONTHLY DATA

I estimate the Markov-Switching Regimes model for Brazil using 2, 3 and 4 regimes. I use three proxies for output, but always quadratic detrending for obtaining the output gap: GDP, GWh and an Industrial Production index. In this section, I report the results with GWh. Tables with other proxies are shown in the appendix.

Table 5 shows some differences compared to the quarterly data; for instance, $g_i < 0$ and g_x is distinct between tables. However, the inflation parameter seems to be very close of each other.

From the table presented, we observe that the intercept is time-varying, and the reaction to output gap is insignificant in the post-Real Plan period.

Figure 6 shows a very similar graph compared to quarterly data. In regime 1, the spikes before the Real Plan refer to the Collor and Real plans. In short, there is no surprise in the

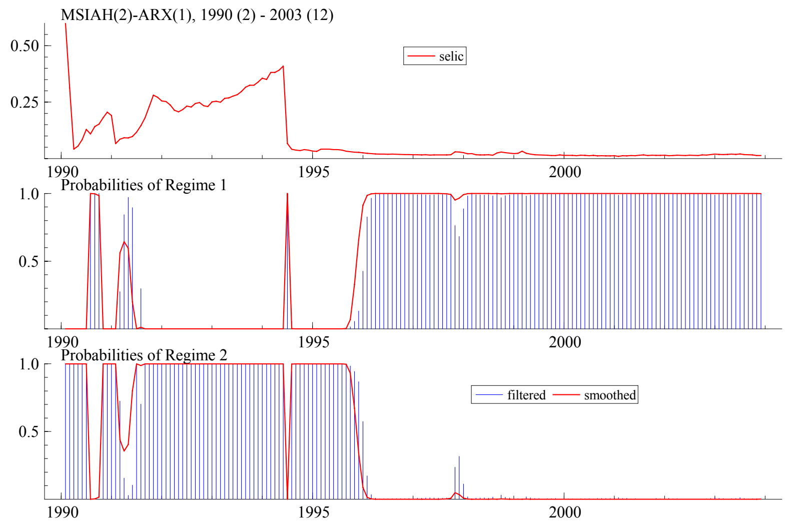
Parameter/Regime	**R1**	**R2**
a	0.137^{*} (0.007)	0.029^{*} (0.011)
g_i	0.129^{*} (0.011)	-0.399^{*} (0.101)
g_π	-0.037 (0.056)	0.498^{*} (0.049)
g_x	-0.004 (0.008)	-0.201^{*} (0.063)
β_{C1}	-0.040^{*} (0.004)	-0.072^{*} (0.013)
β_{C2}	-0.054^{*} (0.006)	-0.020 (0.019)
β_r	-0.004^{*} (0.001)	-0.007 (0.245)
β_e	-0.119^{*} (0.007)	-1.00×10^{-3} (0.012)
$\beta_{\Delta e}$	-0.008 (0.005)	0.631^{*} (0.060)
$\beta_{\Delta R}$	-0.002 (0.004)	0.175^{*} (0.040)
Std. Deviation ($\times 10^3$)	3.350	22.200
Duration	36.66	17.37
Prob.	0.679	0.321
Log-Likelihood	569.53	
Schwarz Criterion	-6.085	

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

Table 5: BRAZIL - GWh - MONTHLY DATA - TWO REGIMES

results, based on the findings with quarterly data.

Figure 6: REGIME PROBABILITIES - MONTHLY DATA - GWh - BRAZIL



Then I impose one more regime and estimate the parameters. Regime 2 comes out to represent the spikes stressed in figure 6, the crises in the second half of the 90's, and the transition between regimes 1 and 3. (See figure 7.)

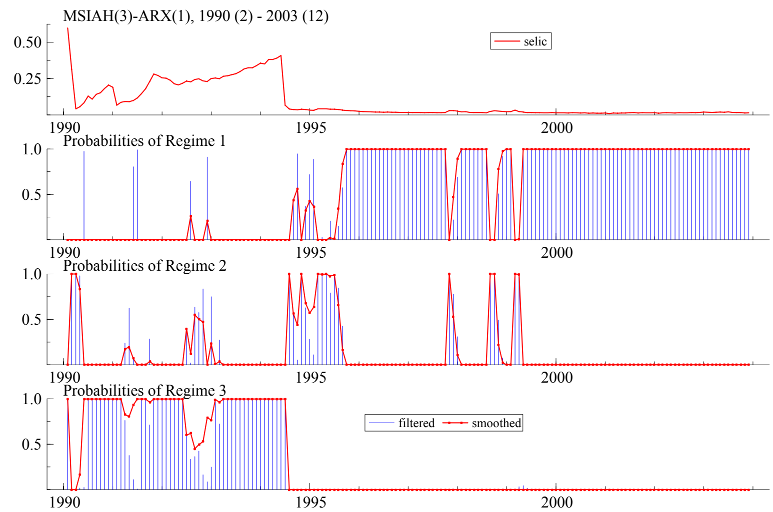
Exactly as in Bueno (2005a,b), which uses other estimation methods[12], table 6 shows that the lagged interest rate parameter becomes important after the Real Plan. At the same time, the inflation parameter gradually loses importance. In addition, again a time-varying intercept surfaces in these findings.

Imposing 4 regimes in the model, we perceive that regime 1 in previous models is once again divided. A close look at figure 8 reveals that its regimes 3 and 4 are similar to those found with the model with 3 regimes. Regime 2 represents the transition period in 2003, when President Lula took over the presidency, the transition from a fixed to flexible exchange rate regime in 1999, and other spikes not markedly characterized.

Table 7 provides no qualitative difference from previous analyses with 2 and 3 regimes. We do not annoy the reader with more tedious words to stress the point that the results demonstrate the

[12] Observe that this conclusion is not true at all to the US, because by varying econometric methods, the results change.

Figure 7: THREE REGIMES PROBABILITIES - MONTHLY DATA - GWh - BRAZIL

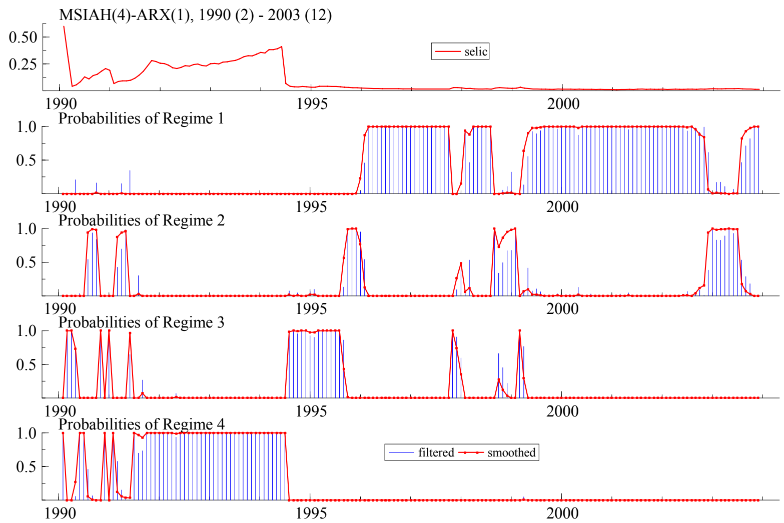


Parameter/Regime	R1	R2	R3
a	0.035^{**} (0.014)	0.194^{*} (0.008)	-0.017^{***} (0.009)
g_i	0.811^{*} (0.052)	-0.011^{*} (0.023)	0.397^{*} (0.127)
g_π	0.127 (0.121)	0.277^{*} (0.018)	0.531^{*} (0.104)
g_x	-0.019 (0.014)	-0.088^{*} (0.030)	0.098 (0.128)
β_{C1}	-0.001 (0.125)	-0.161^{*} (0.006)	0.012 (0.012)
β_{C2}	-0.013 (0.068)	-0.111^{*} (0.010)	0.015 (0.011)
β_r	-0.001 (4.00×10^{-3})	-3.00×10^{-3} (0.006)	0.107 (7.533)
β_e	-0.032^{*} (0.013)	-0.160^{*} (0.008)	-0.082^{***} (0.044)
$\beta_{\Delta e}$	-0.001 (0.002)	0.013 (0.028)	0.399^{*} (0.051)
$\beta_{\Delta R}$	-0.003 (0.002)	0.004 (0.007)	0.052 (0.040)
Std. Deviation ($\times10^3$)	1.306	2.674	15.334
Duration	21.07	2.93	16.54
Prob.	0.676	0.129	0.195
Log-Likelihood	698.56		
Schwarz Criterion	-7.17		

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

Table 6: BRAZIL - GWh - MONTHLY DATA - THREE REGIMES

Figure 8: REGIME PROBABILITIES - MONTHLY DATA - GWh - BRAZIL



ineffectiveness of the Taylor Rule in Brazil and that the results are not theoretically supported.

5 CONCLUSIONS

The results for the US turned out to be different during variations in the sample of the series. Any choice about where to start or to end a sample is arbitrary, and could distort the results that one obtains. Therefore, I let the data itself choose different regimes and estimate the parameters of the Taylor rule using Markov-Switching Regimes method, a completely new empirical approach, which yields conclusions that contrast directly with those of Clarida, Galí and Gertler (2000) and with Orphanides (forthcoming). Again, I did not find any evidence of a time-varying intercept for the US. The main (and surprising) finding is that the more regimes we allow for, then either the inflation greater parameter than one is statistically non-significant, or less than one or, if significant and greater than one, it has the shortest duration and the lowest probability. When it is statistically significant and greater than one, this happens in unexpected periods like the 1960's. These findings refute Clarida, Galí and Gertler's (2000) assumption about the break in the series. As a consequence, the greater than one inflation parameter seems to be driven by the existence of a few extreme observations in short periods of time that happen with low probability.

Parameter/Regime	R1	R2	R3	R4
a	0.065^{*} (0.011)	0.100^{*} (0.014)	0.140^{*} (0.008)	0.008 (0.007)
g_i	0.600^{*} (0.054)	0.389^{*} (0.074)	0.137^{*} (0.023)	0.433^{*} (0.122)
g_π	0.050 (0.046)	0.011 (0.063)	0.370^{*} (0.014)	0.189 (0.160)
g_x	-0.010^{***} (0.006)	-0.026 (0.030)	-0.052^{**} (0.025)	0.263 (0.160)
β_{C1}	-0.036 (3.593)	-0.041^{*} (0.002)	-0.111^{*} (0.006)	0.006 (0.012)
β_{C2}	-0.025 (0.015)	-0.042^{*} (0.006)	-0.074^{*} (0.007)	0.014 (0.113)
β_r	-0.001^{*} (3.00×10^{-3})	-0.005^{*} (0.001)	-0.001 (0.006)	0.104 (0.268)
β_e	-0.058^{*} (0.011)	-0.084^{*} (0.012)	-0.107^{*} (0.008)	-0.078^{***} (0.040)
$\beta_{\Delta e}$	0.001 (0.003)	-0.010^{*} (0.003)	0.007 (0.028)	0.562^{*} (0.050)
$\beta_{\Delta R}$	-0.003 (0.002)	-0.010^{**} (0.004)	-0.006 (0.008)	0.016 (0.039)
Std. Deviation ($\times10^3$)	1.010	1.252	2.268	11.675
Duration	23.27	4.11	3.05	8.58
Prob.	0.544	0.145	0.121	0.190
Log-Likelihood	711.12			
Schwarz Criterion	-6.80			

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

Table 7: BRAZIL - GWh - MONTHLY DATA - FOUR REGIMES

Then if we accept that most periods in the US economy existed an inflation parameter inferior to one, this raises a theoretical question about the dynamic stability of the economy. On the other hand, the reactions to output gap are almost always significant, revealing a permanent concern of policy makers about reducing output gap variability, regardless of the econometric model used, although with different intensities over time.

With quarterly data for Brazil, I estimated a Markov-Switching model with two regimes. I confirmed all the results of Bueno (2005a, b) and observed a continuation of the monetary policy after President Lula da Silva's executive succession. With monthly data I was able to conclude the existence of a time-varying intercept in Brazil (no matter the number of regimes that we impose), and three very well pronounced regimes: high inflation regime (pre-Real Plan), transitory regime during 1994/1995 when the Real Plan was implemented, and subsequent consistent monetary policy just altered over a few months due to international crises or domestic exchange rate surges[13]. We could also observe that the reaction to inflation is, in general, less than one or insignificant, and it seems to approximate one in regimes where high inflation was predominant. By contrast, the coefficient for the lagged interest rate becomes high and significant in the post-Real plan regimes, in compliance with the results using the Kalman Filter and nonlinear least squares (see Bueno, 2005a, b).

Overall, two conclusions emerge. Theoretically, there is space for further research into coherent explanations regarding the results obtained here. Practically, one can conclude that the Taylor Rule may be ineffective to conduct monetary policy. This is particularly important in Brazil, because there there is a widespread feeling among people that the conduction of monetary policy does not work at all. The Central Bank has increased the interest rate for several months between September/2004 and May/2005 without success in terms of reducing the inflation.

We point out that our results are in line with Sims and Zha (2004) who find different regimes for US monetary policy, and hence different coefficients. Thus other studies on monetary rules must take into account this data particularity.

[13] I believe that the same pattern would emerge with quarterly data had we enough observations for that.

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Series	Pot. GDP	GDP	PCE	EFFR
	q^n	q	P_t	i
Source	CBO	BEA	BEA	BG
Seas. Adj.	NA	YES	YES	NA
Freq.	Q	Q	Q	M
Units	B 2000 US	B 2000 US	2000 = 100	%
Range	1949-2014	1947-2003	1947-2003	1954-2004

BEA	U.S. Department of Commerce: Bureau of Economic Analysis
BG	Board of Governors of the Federal Reserve System
CBO	U.S. Congress: Congressional Budget Office
PCE	Personal Consumption Expenditures: Chain-type Price Index
EFFR	Effective Federal Funds Rate
NA	Not applicable

Table 8: DATA DESCRIPTION - USA

	q^n	q	i	π
Mean	6.99	6.97	0.015	0.009
Std. Dev.	0.53	0.55	0.008	0.007
Skewness	-0.16	-0.14	1.105	0.974
Kurtosis	1.82	1.85	4.414	4.094
# Obs.	220	228	198	227

Table 9: BASIC STATISTICS - US

APPENDIX A: DATA DESCRIPTION

UNITED STATES

This section is designed to define the data source and to give the basic statistics of the data I am using. Income and output are always in constant prices. The data-source that we use for the US is summarized in table 8.

Some basic statistics about these variables are presented in table 9. All variables are in log terms.

I proceed the Phillips-Perron unit root test (see Phillips and Perron, 1988) with trend and intercept, unless otherwise noticed.[14] See table 10.

[14]For simplicity, I do not follow Dickey and Pantula's (1987) procedure here.

Variable	Levels	1.st Differences
q	-2.49	$-10.65^{*,b}$
i	-2.34^{b}	$-10.91^{*,c}$
π	$-4.80^{*,b}$	

$*$ - Significant at 1% level

b - Without trend; c - Without trend and intercept

Table 10: UNIT ROOT TESTS - US

Series	GDP	IPCA	SELIC	Cons. Energy	PINDEX	Reserves	Exchange
	q	π	i	GWh	IND	$\operatorname{Re}s$	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		

Table 11: Data Description

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

Table 12: BASIC STATISTICS - BRAZIL

I also ran the Johansen's cointegrating test among q, π, and i (see Harris, 1995.) I rejected the hypothesis of no cointegration using both trace and max-eigenvalues statistics, with intercept and no trend, but I do not report the results.

The Johansen's test allows me to use the variables at their individual levels. But I could simply assume that interest rate is stationary as in Clarida, Galí e Gertler (2000), because of the empirical plausibility of this assumption, as well as the low power of the unit root tests. Moreover, stationarity is also a property found in many theoretical models.

APPENDIX B: BRAZILIAN DATA DESCRIPTION

BRAZIL

The data used in this work is described in table 11.

Some basic statistics about these variables are presented in table 12.

I proceed the Phillips-Perron unit root test in table 13 with trend and intercept, unless otherwise noticed.

Although I do not report results, we carried out Johansen's cointegration test and whatever the hypothesis regarding trend and intercept, linear or quadratic, there is always at least one cointegrating vector.[15]

[15] 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.

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

Table 13: UNIT ROOT TESTS - Quarterly - BRAZIL

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

Table 14: BASIC STATISTICS - BRAZIL

Basic statistics for monthly data are in table 14.

I proceed the Phillips-Perron unit root test with trend and intercept, unless otherwise noted in table 15

APPENDIX C: MARKOV-SWITCHING - US

3 REGIMES - US

5 REGIMES - US

APPENDIX D: ORPHANIDES DATA

Orphanides data is annualized. That's why the standard deviation seems higher.

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

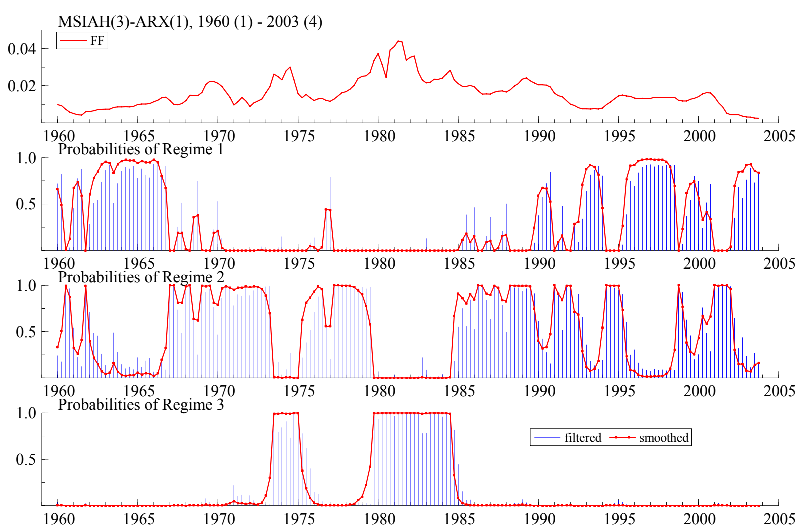
Table 15: UNIT ROOT TESTS - Monthly - BRAZIL

Parameter/Regime	**R1**	**R2**	R3
a	1.00×10^{-3}	-4.00×10^{-3}	0.005
	(2.00×10^{-3})	(5.00×10^{-3})	(0.005)
g_i	0.959*	0.849*	0.839*
	(0.022)	(0.038)	(0.131)
g_π	-0.364	1.350*	0.081
	(0.649)	(0.308)	(1.001)
g_x	0.436**	0.259*	0.369
	(0.212)	(0.076)	(0.419)
Std. Deviation ($\times 10^3$)	0.384	1.399	4.546
Duration	5.68	7.53	13.42
Prob.	0.319	0.515	0.166
Log-Likelihood	903.58		
Schwarz Criterion	-9.65		Volcker

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

Table 16: US - GDP - QUARTERLY DATA - THREE REGIMES

Figure 9: THREE REGIME PROBABILITIES - US

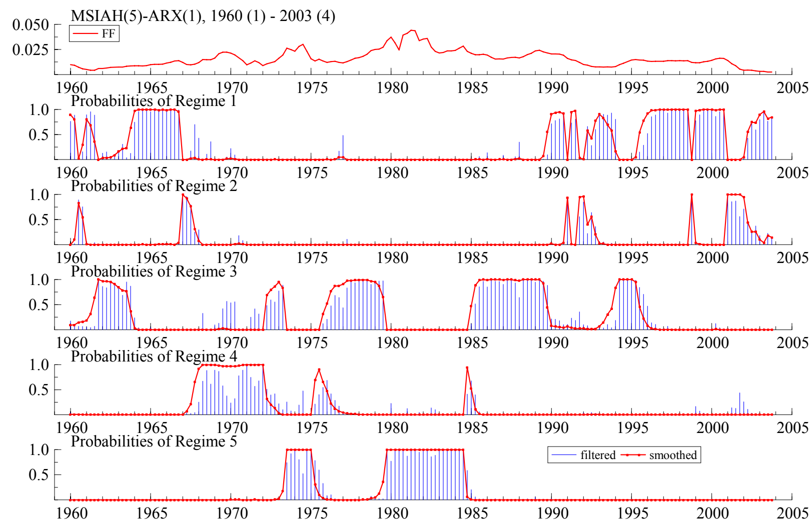


Parameter/Regime	R1	R2	R3	R4	R5
a	0.00×10^{-3} $(2.00{\times}10^{-3})$	-7.00×10^{-3} $(5.00{\times}10^{-3})$	$9.00 \times 10^{-3**}$ $(3.00{\times}10^{-3})$	-0.003 (0.003)	0.007 (0.004)
g_i	0.954* (0.019)	0.768* (0.053)	0.846* (0.029)	0.819* (0.080)	0.825* (0.128)
$\frac{g_\pi}{1-g_i}$	-0.050 (0.503)	1.003* (0.362)	0.735* (0.151)	2.124 (1.572)	-0.183 (0.925)
$\frac{g_x}{1-g_i}$	0.470** (0196)	0.076** (0.038)	0.397* (0.008)	0.313*** (0.172)	0.395 (0.405)
Std. Deviation ($\times 10^3$)	0.383	0.583	0.804	1.569	4.553
Duration	6.20	2.45	9.88	7.49	14.00
Prob.	0.318	0.094	0.301	0.130	0.158
Log-Likelihood	928.97				
Schwarz Criterion	-9.24				Volcker

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

Table 17: US - GDP - QUARTERLY DATA - FIVE REGIMES

Figure 10: FIVE REGIMES PROBABILITIES - US

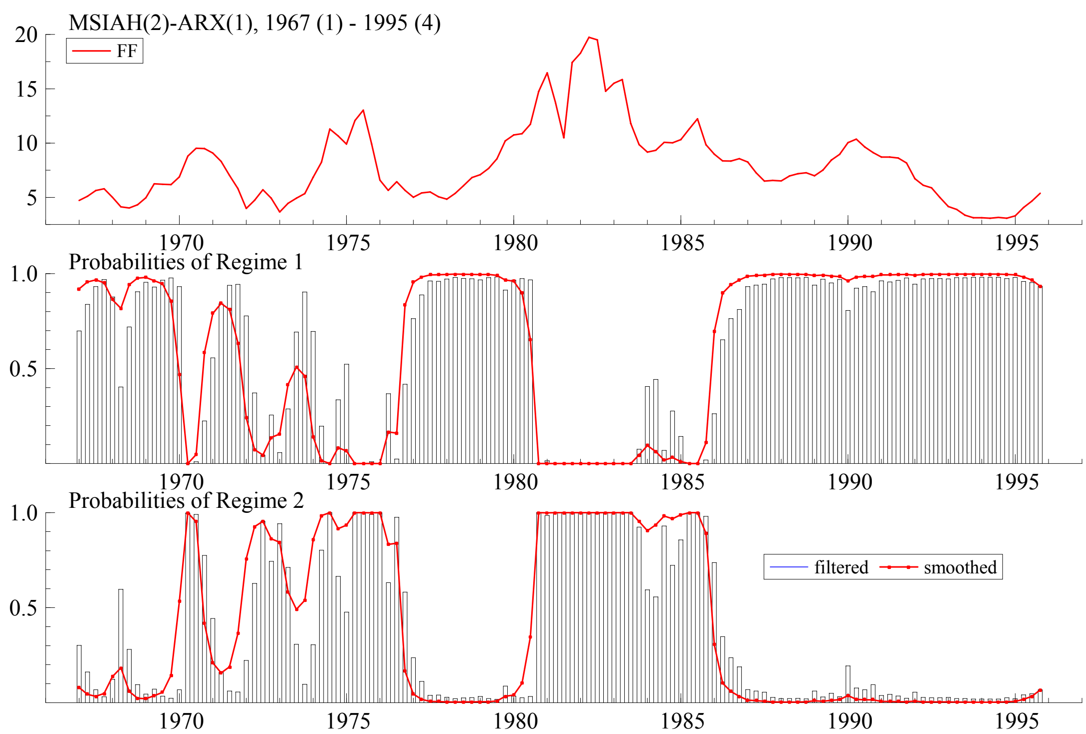


Parameter/Regime	R1	R2
a	-0.194 (0.213)	0.244 (0.860)
g_i	0.930* (0.032)	0.860* (0.081)
$\frac{g_\pi}{1-g_i}$	3.214** (1.393)	2.050 (1.519)
$\frac{g_x}{1-g_i}$	1.785*** (0.939)	0.251 (0.701)
Std. Deviation	0.480	1.933
Duration	17.24	10.07
Prob.	0.631	0.369
Log-Likelihood	-156.32	
Schwarz Criterion	3.19	

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

Table 18: US - ORPHANIDES GDP - QUARTERLY DATA - TWO REGIMES - EXPECTED INFLATION (k = 1, q = 0)

Figure 11: TWO REGIMES PROBABILITIES - ORPHANIDES DATA SET - (k = 1, q = 0)

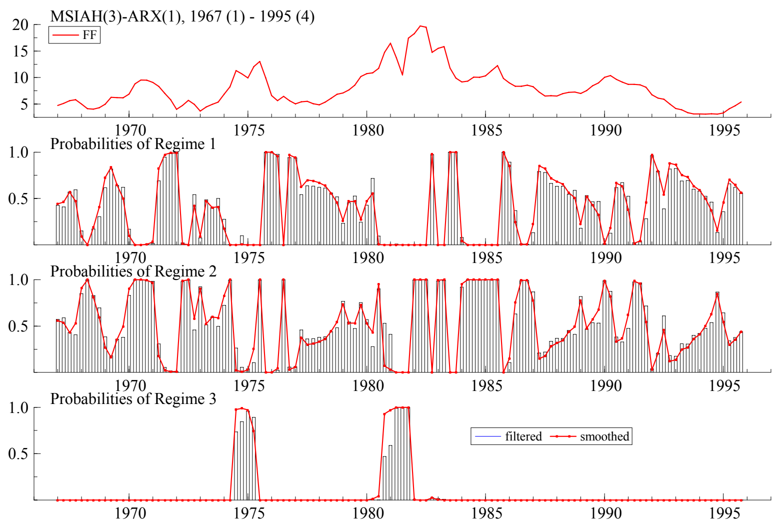


Parameter/Regime	**R1**	**R2**	**R3**
a	0.317 (0.283)	-0.249 (0.370)	37.750* (0.624)
g_i	0.792* (0.027)	0.993* (0.039)	-0.321 (0.302)
g_π	1.581* (0.368)	15.53 (84.172)	-1.384** (0.619)
g_x	0.958* (0.220)	-4.972 (22.522)	0.608* (0.153)
Std. Deviation	0.409	0.624	1.644
Duration	2.72	3.10	4.070
Prob.	0.367	0.678	0.239
Log-Likelihood	-138.49		
Schwarz Criterion	3.25		

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

Table 19: US - ORPHANIDES GDP - QUARTERLY DATA - THREE REGIMES - EXPECTED INFLATION (k = 1, q = 0)

Figure 12: THREE REGIMES PROBABILITIES - ORPHANIDES DATA SET - EXPECTED INFLATION (k = 1, q = 0)

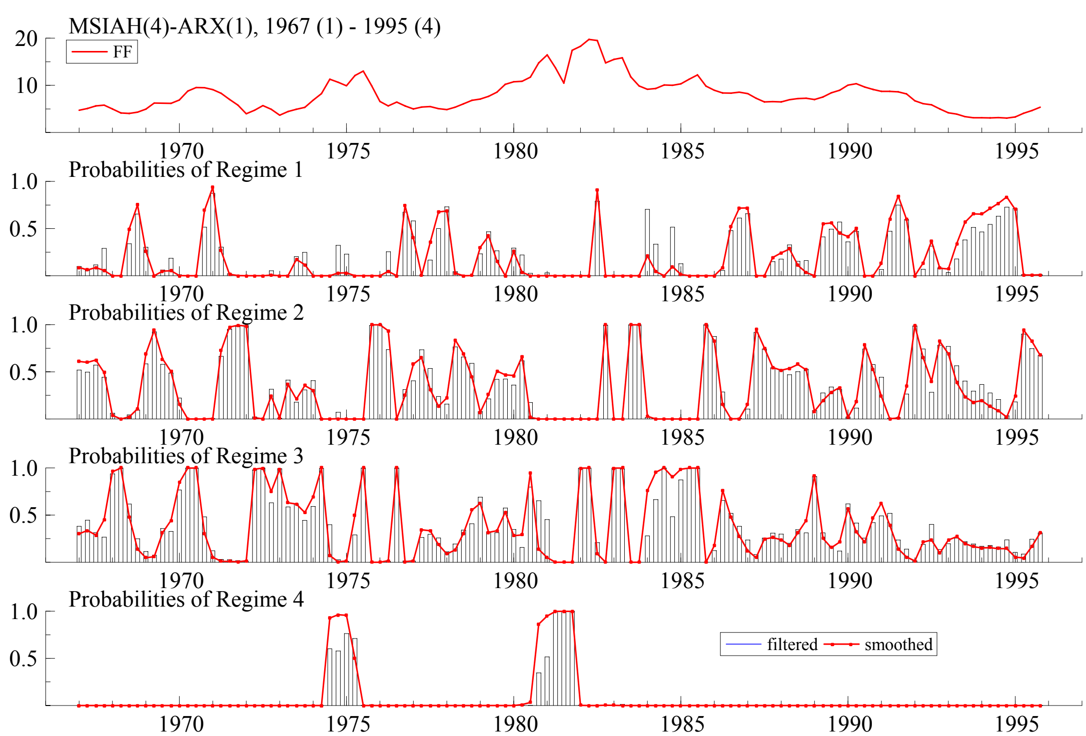


Parameter/Regime	R1	R2	R3	R4
a	-0.447^{**}	0.426	-0.378	38.241^{*}
	(0.211)	(0.305)	(0.624)	(11.222)
g_i	1.053^{*}	0.777^{*}	0.999^{*}	-0.342^{*}
	(0.023)	(0.032)	(0.052)	(0.326)
g_π	-2.218	1.571^{*}	85.083	-1.373^{***}
	(1.763)	(0.380)	(3,694.004)	(0.744)
g_x	-2.659^{*}	0.926^{*}	-56.667	0.613^{*}
	(0.979)	(0.234)	(2,404.580)	(0.163)
Std. Deviation	0.228	0.390	0.655	1.648
Duration	2.03	2.32	2.45	4.04
Prob.	0.199	0.347	0.385	0.070
Log-Likelihood	-131.98			
Schwarz Criterion	3.59			

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

Table 20: US - ORPHANIDES GDP - QUARTERLY DATA - FOUR REGIMES - EXPECTED INFLATION (k = 1, q = 0)

Figure 13: FOUR REGIME PROBABILITIES - ORPHANIDES DATA SET - EXPECTED INFLATION (k = 1, q = 0)



Parameter/Regime	R1	R2
a	0.140^{*} (0.007)	0.005 (0.013)
g_i	0.125^{*} (0.011)	-0.455^{*} (0.104)
g_π	-0.040 (0.054)	0.542^{*} (0.047)
g_x	-0.008^{***} (0.005)	-0.061^{*} (0.017)
β_{C1}	-0.040^{*} (0.040)	-0.036^{**} (0.017)
β_{C2}	-0.057^{*} (0.006)	-0.022 (0.033)
β_r	-0.006^{*} (0.001)	-0.035 (0.216)
β_e	-0.121^{*} (0.007)	0.038^{**} (0.015)
$\beta_{\Delta e}$	-0.011^{**} (0.005)	0.659^{*} (0.060)
$\beta_{\Delta R}$	-0.001 (0.004)	0.172^{*} (0.040)
Std. Deviation ($\times 10^3$)	3.284	21.969
Duration	31.12	15.49
Prob.	0.675	0.325
Log-Likelihood	571.59	
Schwarz Criterion	-6.11	

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

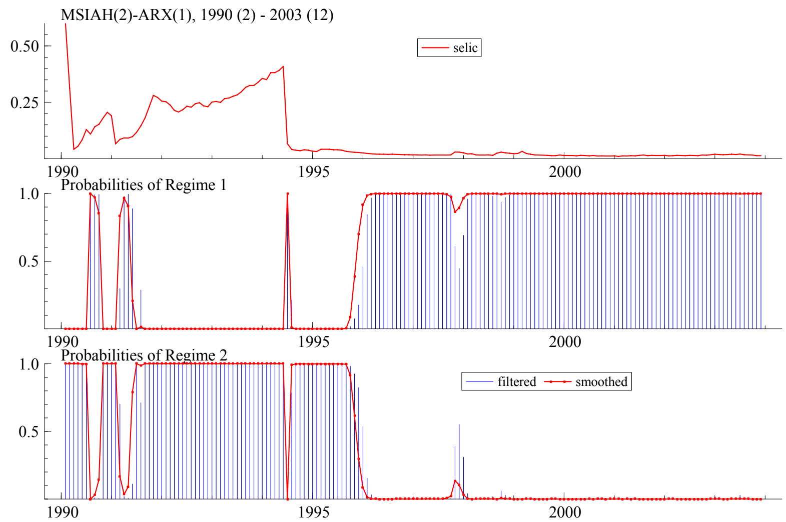
Table 21: BRAZIL - Gdp - MONTHLY DATA - TWO REGIMES

APPENDIX E: MARKOV-SWITCHING - BRAZIL

I present only the results with GDP because they are identical to using the Indutrial Production Index - IND.

4 REGIMES - BRAZIL

Figure 14: TWO REGIME PROBABILITIES - MONTHLY DATA - GDP - BRAZIL



Parameter/Regime	**R1**	**R2**	**R3**	**R4**
a	0.141^{*} (0.000)	0.090^{*} (0.007)	0.175^{*} (0.008)	-4.00×10^{-3} (0.007)
g_i	-0.020^{*} (1.00×10^{-3})	0.799^{*} (0.035)	-0.104 (0.092)	0.397^{*} (0.090)
g_π	-0.086^{*} (0.69×10^{-3})	0.093 (0.091)	0.333^{*} (0.051)	0.746^{*} (0.072)
g_x	0.005^{*} (0.16×10^{-3})	0.006 (0.010)	0.028^{*} (0.008)	0.014 (0.024)
β_{C1}	-0.037^{*} (0.000)	0.111^{*} (0.006)	-0.136^{*} (0.014)	-2.00×10^{-3} (0.011)
β_{C2}	-0.034^{*} (0.000)	0.113^{*} (0.004)	-0.087^{*} (0.000)	-0.010 (0.176)
β_r	-0.013^{*} (0.000)	-1.00×10^{-3} (0.001)	-0.011 (0.033)	0.124 (0.132)
β_e	-0.110^{*} (0.000)	0.093^{*} (0.007)	-0.141^{*} (0.007)	-0.117^{*} (0.033)
$\beta_{\Delta e}$	0.104^{*} (0.000)	-0.001 (0.002)	-0.095^{*} (0.028)	0.222^{*} (0.050)
$\beta_{\Delta R}$	0.014^{*} (0.000)	-0.002 (0.002)	-0.010 (0.010)	0.019 (0.028)
Std. Deviation ($\times10^{3}$)	0.006	1.267	2.751	9.691
Duration	1.81	16.02	5.62	14.30
Prob.	0.073	0.643	0.080	0.204
Log-Likelihood	784.36			
Schwarz Criterion	-7.68			

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

Table 22: BRAZIL - GDP - MONTHLY DATA - FOUR REGIMES

Figure 15: FOUR REGIMES PROBABILITIES - MONTHLY DATA GDP - BRAZIL