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The Recent Brazilian Disinflation Process And Costs*

Alexandre A. Tombini†

Sergio A. Lago Alves‡

Abstract

This Working Paper should not be reported as representing the views of the Banco Central do Brasil. The views expressed in the paper are those of the authors and do not necessarily reflect those of the Banco Central do Brasil.

This work revisits the recent disinflation process in Brazil and finds that solely the agents' perception that a policy rupture could occur is capable of triggering a change in the way firms and households used to behave in their pricing and consuming decisions. This change was captured by structural breaks in the parameters of a generalized hybrid Phillips curve, following the 2002 inflation shock. The paper also shows that such parameter changes led to an increase in the disinflation cost evidenced by a free market inflation gain that would have been observed should the coefficients on the Phillips curve have not changed. The paper finds that, maintaining the occurred paths for interest rates, output gap, nominal exchange rates, administered price inflation and exogenous shocks, the free market inflation would have been significantly lower in the absence of such structural break in the underlying inflation process, since mid 2002.

Keywords: Disinflation costs, inflation dynamics, inflation persistence, exchange rate pass-through, inflation expectations, inflation targets, Kalman filter.

JEL Codes: C22, C61, E31, E37, E52, E58

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1 Introduction

In the months preceding the 2002 presidential election, uncertainties surrounding the macroeconomic framework to be implemented by the upcoming administration and a heightened risk aversion in global markets, produced an unprecedented shock, making it harder for the Central Bank of Brazil to pursue the pre-determined inflation targets in the years that followed.

As shown in Alves, Areosa and Tombini (2005), time-varying estimates of a new Keynesian Phillips curve indicated an increase in the inflation inertia, with respect to both the free market and monitored items price inflations; an increase in the exchange rate pass-through coefficient; and a reduction in the future inflation expectations term. Such a structural change was accompanied by a strong sovereign risk premium shock that almost halved the nominal exchange rate within the short period from May to October 2002. As a consequence, a large inflation pressure emerged. Its effects lasted much longer than in previous episodes, due to the increase in inflation persistence that followed the 2002 shock. The Brazilian monetary authority reacted promptly, but found it even harder to disinflate having to face a flattened Phillips curve.

As it will be shown later on, such a structural change could not be strictly viewed as an instance in which the traditional Lucas' critique was in operation. From the outset of the new administration, monetary and fiscal policies continued to be implemented in a sound and consistent manner. Nonetheless, the agents' perception on the probability of a possible event in which such a rupture could occur, triggered a change in the way firms and consumers behaved in their pricing and consuming decisions. Hence, one of the important outcomes of the paper is the empirical finding that generalizes the well-stated results highlighted by Lucas. We now understand that the uncertainty regarding the anticipated perception of a possible policy change suffices to trigger changes in reduced-form Phillips curve parameters, irrespective of whether the event is confirmed or not. Incidentally these changes imposed higher disinflation costs to Brazil.

This paper presents a methodology to estimate a disinflation cost measure and applies it to the recent Brazilian case. The disinflation cost is presented in a simple way: the inflation gain that would have been observed should the coefficients on the Phillips curve have not changed.

This work is divided as follows: Section 2 presents some background about the context in which the uncertainty regarding the new administration was formed; Section 3 presents and updates one of the time-varying estimations shown in Alves, Areosa and Tombini (2005) and determines the disinflation cost measured in terms of inflation gains; Section 4 provides a few conclusions.

2 Background

2.1 Inflation

In July 1999, following the floating of the real earlier that year, Brazil adopted inflation targeting as the monetary policy framework. The Brazilian IPCA (broad consumer price index) was chosen as the inflation target index. Thus, we will focus our assessment on the behavior of this particular index, as in Alves, Areosa and Tombini[1] (2005).

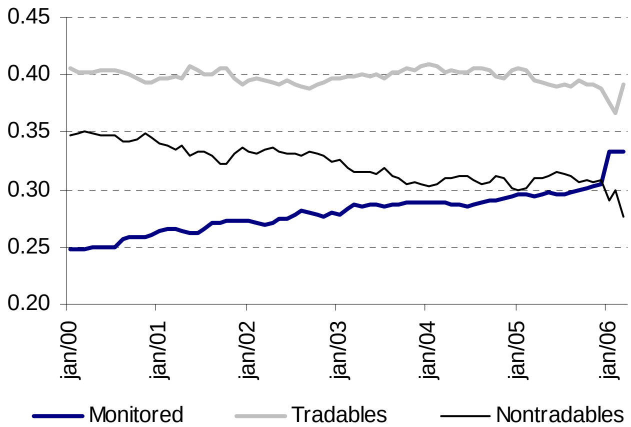

Figure 1: Weights in IPCA

It is a traditional approach in Brazil to disaggregate IPCA in its two main components: free market prices and monitored prices. The latter consist of government-administered prices and utility prices defined by contractual clauses as in the case of

[1]We address the readers to this paper in order to obtain more details about Brazilian inflation stylized facts.

telephone and power[2]. Freely determined prices, the ones directly affected by monetary policy, are broken down into tradables (only consumption tradable goods), and non-tradables. Figure 1 shows the relative weights of monitored, tradables and non-tradables prices in IPCA from 2000 on.

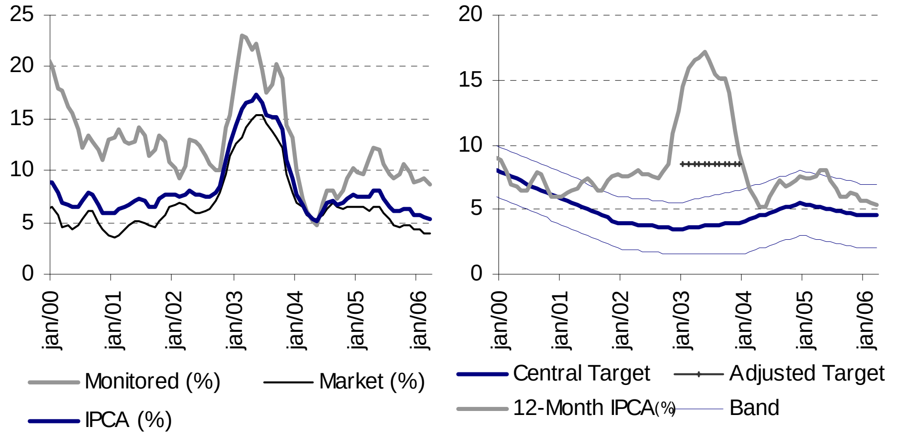

Figure 2: Inflation Targets and 12-Month Inflation Rates

In Brazil, monitored prices have systematically increased more than free market prices during the last decade - Figure 2 depicts the inflation targets, with tolerance bands, and the occurred inflation dynamics since 2000. With the floating of the real in early 1999, monitored prices moved far ahead of IPCA. Therefore, since monitored items have no close substitutes and are essentials, their shares in the consumption bundle have monotonically increased, reaching almost 30 percent in 2005 (Figure 1). In January 2006, the monitored basket definition was changed in order to add medicinal items, whose prices adjustments were regulated in 2003/2004, and exclude fuel alcohol, since its price adjustments criteria were deregulated in the past few years. As a consequence, monitored price weights changed since January 2006.

The last stylized fact is to a large extent explained by the long-lived pass-through of imported inflation - foreign inflation added to exchange rate depreciation - to monitored prices[3]. A significant portion of these prices is affected by international oil prices or is contractually adjusted by the general price index[4] (IGP-M), mostly

[2] A thorough analysis on monitored prices in Brazil is found in Figueiredo and Ferreira (2002).
[3] We will return to this issue in the next subsection.
[4] The IGP-M is released by Getulio Vargas Foundation (FGV), and is comprised of 60 percent

comprised by wholesale and tradable prices of consumption and intermediate items. The evolution of the 12-month foreign inflation[5] in domestic currency can be seen in Figure 3.

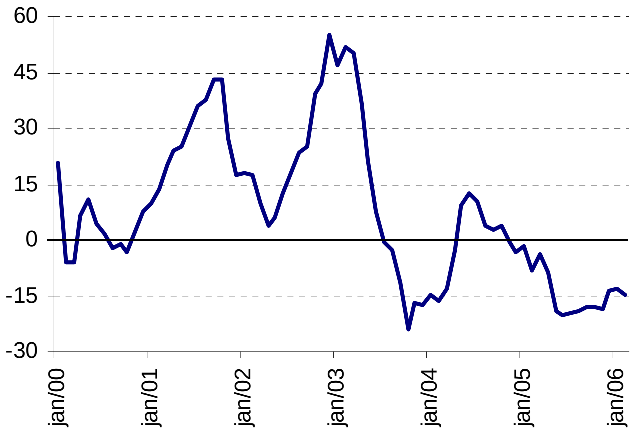

Figure 3: 12-Month Imported Foreign Inflation (%)

2.2 Policy

The previously mentioned combination of the domestically driven shock together with the increased risk aversion in mid 2002 unsettled domestic and foreign investors, decreasing the demand for "real" denominated claims. The rollover of domestic debt became quite challenging and Brazil experienced a sudden stop in capital flows, affecting both inflation and inflation expectations, through the sharp depreciation of the "real".

Against this unfavorable global and domestic backdrop, achievement of the stated inflation targets turned out to be extremely difficult, with IPCA[6] inflation reaching 12.5 percent in 2002, far above the 3.5 percent central inflation target[7].

The challenge in early 2003, with the inauguration of the new administration, was how to regain control over inflation and inflation expectations. The first move was to confirm the commitment of the new government with fiscal consolidation and

of the wholesale price index (IPA-M), 30 percent of consumer price index (IPC-M) and 10 percent of construction costs (INCC-M).

[5] We considered the US export prices (all commodities) as the foreign prices. The term 12-Month imported foreign inflation relates to depreciation added to foreign inflation.

[6] IPCA is the Brazilian broad consumer price index, used to gauge the inflation targets.

[7] In addition to the central inflation target, there was a 2-percentage points tolerance interval.

with the monetary policy framework, combining the inflation target and floating exchange rate.

The target for the primary surplus to GDP ratio was immediately raised to 4.25 percent of GDP. As to the monetary policy strategy, the Central Bank of Brazil continued to raise interest rates (Selic) early in 2003 to reverse the inflationary shock, as depicted in Figure 4.

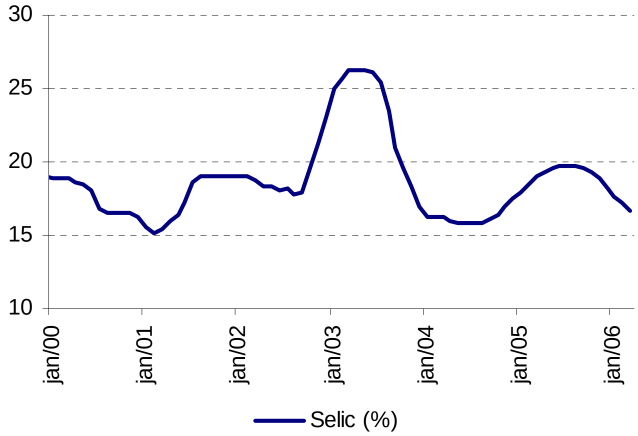

Figure 4: Monetary Policy

However, since the inflation target was missed by a very large margin in 2002, the original multiyear framework - which established a 4.0 percent inflation target for 2003 and 3.75 percent thereafter - was not a credible anchor anymore[8]. Their achievement in such a short time would have required a sharp economic contraction. Hence, in an open letter to the Minister of Finance, the Governor of the Central Bank of Brazil proposed an adjusted target trajectory that would be credible and allow the Central Bank of Brazil to regain control over inflation and inflation expectations, while smoothing out the economic cost of disinflation[9].

[8] In Alves and Areosa (2005), the authors derive a New-Keynesian Phillips curve (NKPC) incorporating indexation not only to past inflation, but also to inflation targets, generalizing the Woodford (2003) hybrid curve. One of their major findings is that the target ability of anchoring inflation - which measures the credibility of a short-run achievement of the inflation targets - has strongly decreased from mid 2002 to the end of 2003, when the target ability of anchoring inflation has been restored to the levels prevailing in 2000.

[9] The above mentioned work of Alves and Areosa (2005) shows that an "adjusted target" is what a central bank should optimally pursue in order to minimize a welfare-based loss function. Indeed, the welfare-based "adjusted target" can be defined as $\lambda \cdot \pi_t^{target} + (1-\lambda) \cdot \pi_{t-1}$, where λ is a parameter that measures the target ability of anchoring inflation. Note that the lower such a parameter is, the more the monetary authority should "adjust" the target with lagged inflation rates in order to maximize the welfare criterion.

Therefore, an adjusted inflation target for 2003 was announced: 8.5 percent for 2003 while the official inflation target for 2004 was set to 5.5 percent, with a tolerance band of 2.5 percentage points - see Figure 2. The flexibility to cope with the large size of the 2002 shock succeeded, with the Central Bank of Brazil being successful in regaining control over inflation expectations and delivering the new disinflation path.

3 The disinflation cost

We updated the general shaped hybrid Phillips curve approach shown in Alves, Areosa and Tombini (2005) to assess changes in the inflation process. To estimate time-varying coefficients we applied the Kalman filter method[10]. This section is highly based on that work.

In our final time-varying approach, we used an instrumentalyzed future inflation expectation term, considering the following state space model[11], in which all variables were seasonally adjusted:

$$\pi_t^{free} = \alpha_{1_t} \cdot \pi_{t-1}^{free} + \alpha_{2_t} \cdot \pi_{t-1}^{monit} + \alpha_{3_t} \cdot E_t \pi_{t+1}^{free} + \alpha_{4_t} \cdot \left(\Delta e_{t-1} + \pi_{t-1}^{f} \right) + \alpha_{5_t} \cdot x_{t-5} + \varepsilon_t \quad (1)$$

With:

$$\begin{gathered}
\alpha_{1_t} = \alpha_{1_{t-1}} + \xi_{1_t} \qquad ; \qquad \alpha_{2_t} = \alpha_{2_{t-1}} + \xi_{2_t} \\
\alpha_{4_t} = \alpha_{4_{t-1}} + \xi_{3_t} \qquad ; \qquad \alpha_{5_t} = \alpha_{5_{t-1}} + \xi_{4_t} \\
\xi_{i_t} \sim N\left(0, \sigma^2\right) \quad i.i.d. \quad \forall i \in \{1, 2, 3, 4\} \\
\alpha_{3_t} = \left(1 - \alpha_{1_t} - \alpha_{2_t} - \alpha_{4_t}\right) \\
\varepsilon_t = \alpha_6 \cdot \varepsilon_{t-1} + \xi_{\varepsilon_t} \\
\xi_{\varepsilon_t} \sim N\left(0, \sigma_\varepsilon^2\right)
\end{gathered}$$

[10] We considered monthly data. Inflation rates are not annualized, so some of the estimated coefficients (e.g., pass-through and output gap) are expected to be near to one third of the usually estimated ones in quarterly frequency. The inflation expectations term was instrumentally determined in the first step of a 2SLS estimation.

[11] We used the Brazilian IPCA (broad consumer price index). Issues concerning the chosen lags are detailed in Alves, Areosa and Tombini (2005). Using cross-correlograms exercises, the authors find that free-market inflation is mostly correlated to the 5-month lagged output gap. Regarding the verticality constraint, alternative models in which such a constraint was not imposed rejected the null of non-verticality. Indeed, the time-varying estimatives for the sum of inflation coefficients and pass-through ranged around 1. Regarding the time-varying coefficients specification, modeling them as describing random walk paths within the limited sample is very standard in the literature for allowing any possible level breaks or trend patterns to be captured. Regarding the AR(1) assumption on the error term, it is important in order to correct any problems coming from possible error serial correlation.

Where π_t^{free} is the market inflation, π_t^{monit} is the monitored price inflation, e_t is the nominal exchange rate, π_t^f is the foreign inflation[12], x_t is the output gap and ε_t is the error term. All of them are logarithmized.

Table 3 summarizes the estimates[13] for the error term auto-correlation and for the standard deviations[14]. Note that the autoregressive coefficient of the error term is very low and non-significant, indicating that our estimatives would not be biased should we estimate them without the AR(1) assumption on the error term.

Table 1: Estimates

Parameter	$Coef$	$St.Dev.$	z	p
α_6	-4.2E-06	7.5E-06	-0.56	0.58
σ	5.6E-03	2.1E-03	2.61	0.01
σ_ε	2.2E-03	1.2E-04	19.03	0.00

In the next subsection, we assess the time-varying estimates for the coefficients.

3.1 Time-varying coefficients

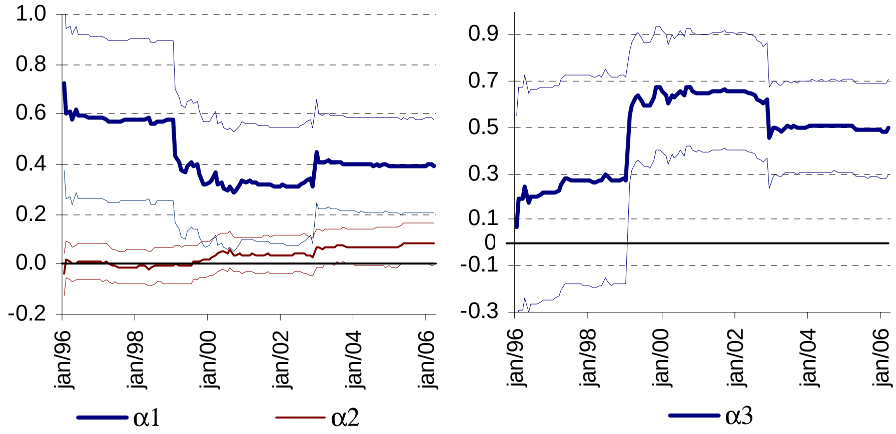

Figure 5: Persistence and Inflation Expectation Coefficients

Figures 5 and 6 show our time-varying estimates for the hybrid curve, within 2-standard deviation confidence intervals. Note that they differ a little from the

[12] We considered the US export prices (all commodities) as the foreign prices.

[13] We used the method Maximum likelihood (Marquardt) with the following sample: Mar 1995 to Mar 2006. Log likelihood: 572.90 .

[14] Once the time-varying coefficients were modeled as random walks, interpretations regarding σ^2 should be made with care, for the total variance must be calculated adding σ^2 to the variances implied by their time-varying paths within the limited sample.

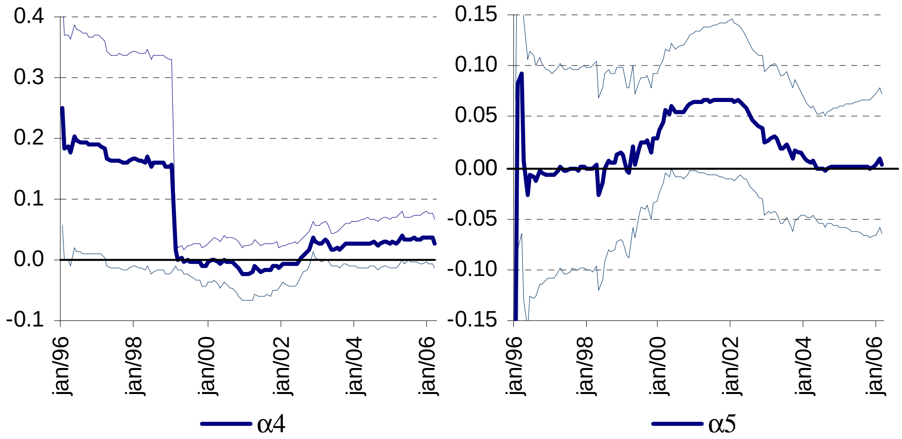

Figure 6: Foreign Inflation Pass-Through and Output Gap Coefficients

ones obtained in Alves, Areosa and Tombini (2005). Such a slight difference is probably due to different assumptions regarding the instrument list when obtaining the inflation expectation term. However, our estimates have the same pattern shown in that paper.

The main outcomes are the following:

- Inflation persistence from free market prices have slightly and persistently increased at the end of 2002, despite the sharp reduction observed from mid 1999 to 2002[15].

- Inflation persistence from monitored prices have been increasing, showing two persistent jumps: at the beginning of 1999 and at the end of 2002;

- The expectation term has slightly lost its importance at the end of 2002, regardless the fact that its coefficient is much higher than it was prior to 1999.

- The pass-through coefficient has persistently increased since mid 2002[16];

- The Phillips curve became flatter from mid 2002 on[17].

[15]Such a reduction is in line with Minella et al (2003) and Alves (2001).

[16]The finding is probably due to the increase in the Brazilian foreign trade that occurred since 2002. For more results on the Brazilian pass-through coefficient, see for example Bogdanski et al (2000), Goldfajn and Werlang (2000), Muinhos (2001), Belaisch (2003), Muinhos and Alves (2003), Minella et al (2003), Correa and Minella (2005) and Albuquerque and Portugal (2005).

[17]Alves, Areosa and Tombini (2005) found evidence that output gap coefficient rises when the economy is overheated, indicating that the Phillips curve may have a convex shape in relation to the output gap term. This evidence means that as the product level rises above its natural level,

3.2 Assessing the disinflation cost

In mid 2002, the Central Bank of Brazil had to react promptly to fight the inflationary effects caused by the sharp exchange rate depreciation as shown in Figure 3. It found it harder however, to fight inflation as a consequence of the structural changes in the underlying inflation process. This structural change raised the persistence of the inflationary shock and lowered the output gap coefficient. Therefore, monetary policy had to be tighter than before, increasing the cost of disinflation from mid 2002 on.

Note that the structural changes in the underlying inflation process cannot be strictly viewed as an instance in which the traditional Lucas' critique was in operation. Indeed, from the outset of the new administration, monetary and fiscal policies continued to be implemented in a sound and consistent manner. To the contrary, monetary policy continued to be conducted in a way to ensure the achievement of the inflation targets and fiscal policy became tighter to restore credibility and to achieve faster consolidation.

However, the agents' perception on the probability of a possible event in which a policy rupture could occur triggered a change in the way firms and households used to behave in their pricing and consuming decisions. Thus, such uncertainty, even though not confirmed ex-post, sufficed to cause the previously assessed changes in the aggregated parameters of the considered Phillips curve. Such changes surely imposed higher disinflation costs to the Brazilian economy.

A straightforward methodology is presented to estimate the disinflation cost and applied to the recent Brazilian disinflation process. We measure the inflation gain that would have been observed should the coefficients on the Phillips curve have not changed, maintaining the occurred paths for interest rates, output gap, nominal exchange rates, monitored inflation and exogenous shocks. So we simulate the probable free market inflation path that would have occurred in such environment. The difference between the occurred path and the simulated free market inflation

production restrictions induce upward pressures on inflation with an increasing magnitude, at the margin. Therefore, monetary policy may find it easier to reduce inflation if the economy is not in a recession, for the sacrifice ratio will be lower. Laxton et al (1998) and Clark et al (1995) found such evidence using US data. For a review on the evolution of views on this issue, see Clark and Laxton (1997).

rates will be our measure of the disinflation cost.

For forecasting purposes, it is usual to assume that $E_t\varepsilon_{t+j} = 0$ for any period $t+j$ in the future, so model consistent simulations can be run assuming $E_t\pi_{t+1}^{free} = \pi_{t+1}^{free}$ for any period $t+j$ in the future.

But now, due to the presence of occurred exogenous shocks, our simulation with rational expectations had to be made with care, for the correct relation was $\pi_{\tau+1}^{free} = E_\tau\pi_{\tau+1}^{free} + \bar{\varepsilon}_{\tau+1}$ for any past period τ in the sample. Since our time-varying estimations were carried out considering an instrumentalyzed inflation expectation term that could differ from the actual latent expectation term, the econometric residuals were not the correct series to be used as $\bar{\varepsilon}_\tau$. Indeed, if we assume that the time-varying parameters are pre-defined with a one period lag, note that we can lead (1) in one period and apply the iterating expectations property in order to obtain the following expression:

$$\pi_{\tau+1}^{free} = E_\tau\pi_{\tau+1}^{free} + \left[\varepsilon_{\tau+1} + \alpha_{3_{\tau+1}} \cdot \left(E_{\tau+1}\pi_{\tau+2}^{free} - E_\tau\pi_{\tau+2}^{free}\right)\right] \qquad (2)$$

Denoting the term inside the brackets by $\bar{\varepsilon}_{\tau+1}$, we realize that $\bar{\varepsilon}_\tau$ can depart from ε_τ depending on the magnitude of $\alpha_{3_\tau} \cdot \left(E_\tau\pi_{\tau+1}^{free} - E_{\tau-1}\pi_{\tau+1}^{free}\right)$.

Hence, in order to obtain a more accurate measure for $\bar{\varepsilon}_\tau$, we considered an iterative procedure described in the Appendix. With the new estimates for $\bar{\varepsilon}_\tau$, we simulated the new path for the free market inflation in a counter-factual approach, maintaining the occurred paths for interests rates, output gap, nominal exchange rates, monitored inflation and exogenous shocks $\bar{\varepsilon}_\tau$. In such simulations, we found the new rational expectation equilibrium path for the free market inflation from July 2002 on, considering that the parameters in (1) were fixed in the previous average levels occurred from January 1999 to June 2002.

Figure 7 compares the occurred seasonally adjusted monthly free market inflation path with the one that would have occurred should the coefficients have not changed. Note that the simulated inflation rates are much lower than the occurred ones until July 2004. This is due to the fact that the exchange rate began a long-lasting nominal appreciation trend. Since the new pass-through coefficients are higher than the previous ones, such higher inflation rates were expected to occur in the simulations.

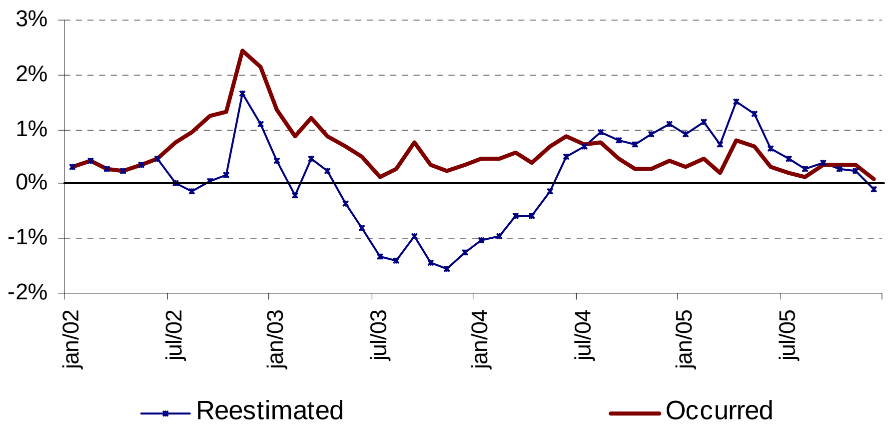

Figure 7: Counter Factual Simulation

Finally, we determined the accumulated disinflation cost as about 20 percentage points from July 2002 to December 2005, which means that the free market inflation could be, on average, 6 percentage points lower[18], in an annual basis, if the parameter had not changed after mid 2002, maintaining the occurred paths for the previously mentioned variables.

4 Concluding remarks

Using a Kalman filter approach, we updated some of the estimates of Alves, Areosa and Tombini (2005) and estimated time-varying parameters for a generalized hybrid Phillips curve. We found that many of the coefficients have moved to different levels on at least two occasions. The first one was just after the change from the "crawling peg" regime to the floating, in 1999. Then, a change in the parameters was expected, since the economic policy framework was significantly altered. The most remarkable change occurred in the pass-through coefficient, which fell significantly.

The second break occurred from mid 2002 on and was not related to a policy change. Instead, it was related to the agents' perception on the probability of a possible event in which a policy rupture could occur. Such uncertainty triggered a change in the way firms and households used to behave in their pricing and consuming decisions, causing the previously assessed changes in reduced-form Phillips

[18] In order to determine such an average, we considered the 42 months in the simulated sample and computed $(1-0.20)^{(12/42)} - 1 \approx -0.06$.

curve parameters.

Such changes also imposed higher disinflation costs to the Brazilian economy. The Central Bank of Brazil found it harder to restore inflation to their targets and keep prices under control afterwards. Therefore, we presented a counter factual method to estimate the disinflation cost and estimated it, in accumulated inflation terms, at 20 percentage points from July 2002 to December 2005. If the parameters had not changed from mid 2002 on, free market inflation could have been about, on average, 6 percentage points lower, in an annual basis, given the occurred path for interests rates, output gap, nominal exchange rates, monitored inflation and exogenous shocks.

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A Appendix

In order to obtain a more accurate measure for $\bar{\varepsilon}_\tau$, we considered an iterative procedure in which the first step was made assuming that a first prior for the expectation term was the occurred one. Hence we could determine a first estimate for $E_\tau \pi_{\tau+1}^{free}$ as follows:

$$\begin{aligned}\left(E_\tau \pi_{\tau+1}^{free}\right)_{Step\ 1} &= \alpha_{1_{\tau+1}} \cdot \pi_\tau^{free} + \alpha_{2_{\tau+1}} \cdot \pi_\tau^{monit} + \alpha_{3_{\tau+1}} \cdot \pi_{\tau+2}^{free} + \\ &+ \alpha_{4_{\tau+1}} \cdot \left(\Delta e_\tau + \pi_\tau^f\right) + \alpha_{5_{\tau+1}} \cdot x_{\tau-4}\end{aligned}$$

Hence, we obtained the following first step estimate for $\bar{\varepsilon}_{\tau+1}$:

$$\left(\bar{\varepsilon}_{\tau+1}\right)_{Step\ 1} = \pi_{\tau+1}^{free} - \left(E_\tau \pi_{\tau+1}^{free}\right)_{Step\ 1}$$

In the second step, we improved the estimate for $E_\tau \pi_{\tau+1}^{free}$ as follows:

$$\begin{aligned}\left(E_\tau \pi_{\tau+1}^{free}\right)_{Step\ 2} &= \alpha_{1_{\tau+1}} \cdot \pi_\tau^{free} + \alpha_{2_{\tau+1}} \cdot \pi_\tau^{monit} + \alpha_{3_{\tau+1}} \cdot \left(E_\tau \pi_{\tau+2}^{free}\right)_{Step\ 1} + \\ &+ \alpha_{4_{\tau+1}} \cdot \left(\Delta e_\tau + \pi_\tau^f\right) + \alpha_{5_{\tau+1}} \cdot x_{\tau-4}\end{aligned}$$

Therefore, the second estimate for $\bar{\varepsilon}_{\tau+1}$ was:

$$\left(\bar{\varepsilon}_{\tau+1}\right)_{Step\ 2} = \pi_{\tau+1}^{free} - \left(E_\tau \pi_{\tau+1}^{free}\right)_{Step\ 2}$$

We then repeated the following iteration, namelly j, until $\left(\bar{\varepsilon}_{\tau+1}\right)_{Step\ j}$ have converged to $\left(\bar{\varepsilon}_{\tau+1}\right)_{Step\ j-1}$, e.g. until a tiny tolerance criterion had been achieved:

$$
\begin{aligned}
\left(E_\tau \pi_{\tau+1}^{free}\right)_{Step\ j} &= \alpha_{1_{\tau+1}} \cdot \pi_\tau^{free} + \alpha_{2_{\tau+1}} \cdot \pi_\tau^{monit} + \alpha_{3_{\tau+1}} \cdot \left(E_\tau \pi_{\tau+2}^{free}\right)_{Step\ j-1} + \\
&\quad + \alpha_{4_{\tau+1}} \cdot \left(\Delta e_\tau + \pi_\tau^f\right) + \alpha_{5_{\tau+1}} \cdot x_{\tau-4} \qquad (3)
\end{aligned}
$$

And:

$$
(\bar{\varepsilon}_{\tau+1})_{Step\ j} = \pi_{\tau+1}^{free} - \left(E_\tau \pi_{\tau+1}^{free}\right)_{Step\ j-1} \qquad (4)
$$

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