Non-linear mechanisms of the exchange rate pass-through: A Phillips curve model with threshold for Brazil *

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Abstract

This paper investigates the presence of non-linear mechanisms of the pass-through from exchange rate to inflation in Brazil. In particular, it estimates a Phillips curve with a threshold for the pass-through. The paper examines whether the short-run magnitude of the pass-through is affected by the business cycle, direction and magnitude of the exchange rate change, and volatility of the exchange rate. For that purpose, three variables are tested as thresholds: i) output gap, ii) exchange rate change, and iii) exchange rate volatility. The results indicate that the short-run pass-through is higher when the economy is booming, when the exchange rate depreciates above some threshold and when the exchange rate volatility is lower. These results have important implications for monetary policy and are possibly related to pricing-to-market behavior, menu costs of price adjustment and uncertainty about the degree of persistence in exchange rate movements.

Keywords: Exchange Rate Pass-Through, Threshold, Inflation, Non-linearity, Brazil
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1. Introduction

The non-linearity of the Phillips curve has been an important topic in the recent research on monetary policy. First, IS and Phillips curves are the benchmark equations in the New Keynesian approach to monetary policy—largely used in policy-oriented research in central banks. Second, the slope of the Phillips curve has different implications for stabilization policies. For example, the cost of disinflation depends on the shape and steepness of the Phillips curve.

Although the majority of the initial studies worked with a linear specification of the Phillips curve, recent research has investigated theoretically and empirically the possibility of existence of non-linearities.[1] Three major ideas come out about the shape of the Phillips curve. The first, pointed out by Laxton, Rose and Tambakis (1998), Tambakis (1998), Scharling (1999) and Bean (2000), considers that the Phillips curve is convex. Accordingly, the sensitivity of inflation to the economic activity increases with the level of the latter. Therefore, the output cost of disinflation is lower when the economy is booming. In contrast, Stiglitz (1997) and Eisner (1997) claim that the Phillips curve is concave on the grounds that non-competitive firms (i.e., in monopolistically competitive or oligopolistic markets) would be more reluctant in raising prices than in lowering them to obtain and/or maintain market share. Finally, Filardo (1998) considers that the Phillips curve is neither entirely convex nor concave, but a combination of both (a concave-convex curve). The disinflation cost depends on the point where the economy stands.

This paper investigates the presence of non-linearities in the magnitude of the pass-through from exchange rate changes to prices. The understanding and quantification of the impact of the pass-through is important for the conduct of monetary policy, especially in inflation targeting countries such as Brazil, which was hit by strong exchange rate shocks. In fact, the possibility of the presence of a non-linear pass-through in Brazil has motivated an important empirical investigation, such as in Bogdansky, Tombini and Werlang (2000), Goldfajn and Werlang (2000), and Carneiro, Monteiro and Wu (2002). The sources of non-linearity considered are usually the nominal and real exchange rates, degree of openness and output level.

Following this discussion, this paper aims to investigate the possibility of a non-linear pass-through in Brazil using threshold models. It estimates a Phillips curve with three different variables as

[1] See, for instance, Chadha, Masson and Meredith (1992), Laxton, Meredith and Rose (1995), Stiglitz (1997), Filardo (1998), Dupasquier and Ricketts (1998), Schaling (1999), Nobay and Pell (2000), Aguiar and Martins (2002), and Clements and Sensier (2002).

threshold: i) output gap; ii) nominal exchange rate change; and iii) exchange rate volatility. The first question is whether economic activity affects the magnitude of the pass-through. The second one is whether the pass-through is symmetric with respect to the direction of the exchange rate change—whether appreciations or depreciations have symmetric effects on prices—and to the magnitude of the exchange rate change.

The estimations in this paper indicate that the short-run pass-through is higher when the economy is booming, when the exchange rate depreciates above a threshold value and when exchange rate volatility is lower. These results affect monetary policy and are possibly related to pricing-to-market behavior, menu costs of changing prices, and uncertainty as to the degree of persistence of exchange rate changes.

The article is organized as follows. Section 2 presents the theoretical underpinnings and develops a model used to give support to the presence of a non-linear pass-through related to the direction and magnitude of the exchange rate change. The following section sets forth the methodology of threshold models with the presence of endogenous variables. Section 4 presents the specification of the Phillips curve with threshold and the estimation results. The last section concludes the text.

2. Theoretical basis

The literature on exchange rate pass-through reports several sources of non-linearity, indicating that the degree of pass-through can be related to some macroeconomic variables. In this paper, we examine three of them: the level of economic activity, the exchange rate change and its volatility.

The business cycle as a source of non-linearity is well known in the literature. For Brazil, some studies—including Goldfajn and Werlang (2000), Carneiro, Monteiro and Wu (2002), and Muinhos (2001)—investigate the possibility that business cycle affects the degree of pass-through. The hypothesis is that the lower the aggregate demand, lower is the room for pricing adjustments. These papers test that hypothesis using cross terms between some measure of economic activity and the exchange rate term in the Phillips curve.

The literature also indicates the possibility that the degree of pass-through depends on the direction or magnitude of exchange rate change. Empirical works, such as Mann (1986), Goldberg (1995), Gil-Pareja (2000), Mahdavi (2002) and Olivei (2002), have found asymmetric pass-through related to the direction of exchange rate change. On the other hand, studies such as Ohno (1989) and

Pollard and Coughlin (2004) indicate the presence of asymmetry associated with the magnitude of exchange rate change.

We present a model that can deal with these possibilities. It follows Blonigen and Hayes (1999), Gil-Pareja (2003), Pollard and Coughlin (2004) and the pricing-to-market literature, which argues that an exporter firm discriminates prices across countries to where it sells its products as the exchange rate changes. The exporter may adjust its markup for a specific country to accommodate changes in the exchange rate.

We assume that the economy of the "domestic" country is comprised by three sectors, each one with several industries. The industries in sector M have a production technology that uses imported inputs. This sector produces internally the products $x_{M,i}$, where the subscript *M* indicates the sector and *i* represents the industry type (i=1,...,n). Sector D is comprised by domestic industries that use only domestic inputs in the production of goods $x_{D,j}$, where *j* designates the industry type (j=1,...,k). Sector F, in turn, is formed by subsidiaries of foreign firms, which produce abroad (say, in one country, called "external") the products $x_{F,z}$, where *z* is the industry type (z=1,...,h), and sell them in the domestic country. We consider that firms have some market power, but each industry is comprised by several firms, implying that there is struggle for market share. Subsidiaries of foreign firms also face competition of substitute domestic goods. Sectors M, D and F comprise a proportion of the economy equals to θ_M, θ_D and $(1-\theta_M-\theta_D)$, respectively.

Because some goods have close substitutes and, thus, are weakly separable of the others in the consumer's utility function, demand for good $x_{M,i}$ is $x_{M,i} = f(p^{M,i}, p^s, y)$, where $p^{M,i}$ is the price of good $x_{M,i}$, p^s is a vector with the prices of the substitute goods of $x_{M,i}$ and y is the output level of the domestic economy. Similarly, demand for good $x_{D,j}$ of sector D is given by $x_{D,j} = f(p^{D,j}, p^s, y)$, and that of sector F by $x_{F,z} = f(p_E^{F,z}, p^s, y)$, where $p_E^{F,z}$ is the domestic price of imported goods.

Regarding the price decision of firms in sector M, which use imported inputs, the price of input $w_{M,i}$ depends on the exchange rate, denoted by e, measured as the price of foreign currency in domestic currency units. In addition, the production cost of each firm depends on its production level, $x^*_{M,i}$, where the superscript * denotes an individual firm. Thus, the production cost of the firm is given by $c(x^*_{M,i}, w_{M,i})$. Assuming that costs are homogeneous of degree one with respect to input prices, we can write $c(x^*_{M,i}, w_{M,i}) = w_{M,i}(e)\phi(x^*_{M,i})$, where $\phi(x^*_{M,i})$ is a function that depends only on the quantity produced.

Each firm takes the price of its competitors as given and maximizes its profits:

$$\max_{p^{M,i}} L = p^{M,i}.x^*_{M,i} - w_{M,i}(e)\phi(x^*_{M,i}). \qquad (1)$$

Maximizing this equation with respect to the price generates the first-order condition:

$$\frac{\partial L}{\partial p^{M,i}} = x^*_{m,i} + p^{M,i}\frac{\partial x_{M,i}}{\partial p^{M,i}} - w_{m,i}(e)\phi(x^*_{M,i})'\frac{\partial x_{M,i}}{\partial p^{M,i}} = 0,$$

which can be rewritten as

$$\frac{\partial x_{M,i}}{\partial p^{M,i}}\left[p^{M,i}\left(1-\frac{1}{\varepsilon^M}\right) - w_{M,i}(e)\phi(x^*_{M,i})'\right] = 0, \qquad (2)$$

where $\varepsilon^M = -\frac{\partial x_{M,i}}{\partial p^{M,i}}\frac{p^{M,i}}{x_{M,i}}$ is the price-elasticity of demand.

Using $\upsilon^{M,i}$ for the markup over the marginal cost of the firm in sector M, where $\upsilon^{M,i} = \frac{1}{1-1/\varepsilon^M}$, we can rewrite equation (2) as:

$$\frac{\partial x_{M,i}}{\partial p^{M,i}}\left[\frac{p^{M,i}}{\upsilon^{M,i}} - w_{M,i}(e)\phi(x^*_{M,i})'\right] = 0,$$

implying

$$p^{M,i} = w_{M,i}(e)\phi(x^*_{M,i})'\upsilon^{M,i}. \qquad (3)$$

The solution of the maximization problem results in the usual solution, where the price of each firm in the market is determined by a specific markup, $\upsilon^{M,i}$, over the marginal cost, $w_{M,i}(e)\phi(x^*_{M,i})'$. Note that those costs are affected by exchange rate changes.

Doing the same for sector D, we find a similar expression. The difference is that marginal costs of those firms are not affected by the exchange rate:

$$p^{D,j} = w_{D,j}\phi(x^*_{D,j})'\upsilon^{D,j}. \qquad (4)$$

In sector F, firm's decisions are made in its headquarters, located in the foreign country. We assume that foreign firms produce in the foreign country, using only inputs produced there and their final products are imported by their subsidiaries in the domestic economy. Production costs denominated in foreign currency depend on the quantity produced, $c_F(x^*_{F,z}) = \phi_F(x^*_{F,z})$. Since profit maximization decision is made in the foreign country, the price is determined in the currency of the foreign country, $p^{F,z}$. Therefore, firm's problem is given by:

$$\max_{p^{F^*,z}} L = p^{F,z}.x^*_{F,z} - \phi_F(x^*_{F,z}), \qquad (5)$$

whose first-order condition indicates that the price that maximizes profit is determined by a markup over the marginal cost, all expressed in foreign currency:

$$p^{F,z} = \phi_F(x^*_{F,z})' \upsilon^{F,z}. \qquad (6)$$

However, this good is sold by its subsidiary in the domestic country, and its price should be converted into local currency, multiplying the price by the exchange rate. Thus, the price in local currency in the domestic country is:

$$p_E^{F,z} = p^{F,z} e$$

The price index in the economy is determined by a weighted average of the prices in the three sectors, $P = p_M\theta_M + p_D\theta_D + p_F(1-\theta_M-\theta_D)$.

This model illustrates the reasons why an exchange rate appreciation could lead to a higher or lower pass-through than in the case of an exchange rate depreciation. We summarize below the possible different situations.

We consider first sector F, which imports final products through domestic subsidiaries. In the case of an exchange rate depreciation, foreign firms have three options: i) reduce their markup to keep stable the price in local currency (absence of pass-through); ii) keep their markup, increasing the price charged in the domestic country to reflect completely the exchange rate change (complete pass-through), which may imply a market share reduction; or iii) a combination of the previous two possibilities (partial pass-through).

Therefore, pricing-to-market theory delivers an explanation for a partial pass-through. When subsidiaries firms are trying to build up or keep their market shares, a local currency depreciation results in a lower pass-through than that when there is an appreciation. When exchange rate depreciates, firms that produce externally can offset the potential price increase lowering theirs markup and keeping the prices charged by their subsidiaries. The extension of this effect on the price level in the economy depends on ε^F, the price-elasticity of demand for these firms' goods, and $(1-\theta_M-\theta_D)$, the share of sector F in the economy.[2]

In contrast, if the firms that produce abroad face a restriction on their production capacity, an exchange rate appreciation can result in a lower pass-through than a depreciation can. The restriction capacity limits the fall of domestic price that the appreciation could generate. As before, the effect on the economy price level depends on the degree of openness of the economy.

An exchange rate depreciation can have a higher effect also because of the behavior of the domestic firms that use imported inputs. A depreciation implies a higher cost, which may result in losses or reduction in their markup. To avoid those losses, firms tend to change prices more quickly. In the case of appreciation, firms' profits would increase, which could result in a longer period to adjust prices downwards.

This model can also deal with the presence of asymmetry in the pass-through related to the magnitude of exchange rate changes if we assume the presence of menu costs. If price changes are costly, a small change in the currency value can be accommodated within the markup margin. Consequently, menu costs increase the possibility that firms will adjust price only if the exchange rate change surpass some threshold. Therefore, the presence of menu costs can result in a asymmetric pass-through related to small and large exchange rate changes.

In addition, firms assess the degree of persistence of exchange rate changes. Variations that are considered permanent have a quicker pass-through. On the other hand, firms tend to postpone their pass-through decisions when exchange rate changes are considered transitory.

3. Threshold models with endogenous variables

A natural way of modeling economic series using non-linear models is to determine different states of nature or regimes and allow different dynamic behavior for the variables conditional on the regime prevailing in each moment (Franses and van Dijk, 2000). This means that some properties of the time series, such as mean and autocorrelation, may change across regimes.

One way of doing this is using threshold models, where the sample is divided into classes based on the value of an observed variable—whether it surpasses or not some threshold. This kind of model—Threshold Autoregressive (TAR) Model—was initially proposed by Tong (1978) and Tong

[2] That is, the degree of openness of the economy.

and Lim (1980) and spread in the recent applied economic literature. When the threshold is not known—as usual in practice—it needs to be estimated. The simplest model is the SETAR (Self-Exciting Threshold Autoregressive Model), where the threshold is given by a lagged term of the dependent variable—y_{t-d}, where d>0. An AR(1) model of two regimes and d=1 can be rewritten as:

$$y_t = \begin{cases} \phi_0^1 + \phi_1^1 y_{t-1} + \varepsilon_t & if \quad y_{t-1} \langle \tau \\ \phi_0^2 + \phi_1^2 y_{t-1} + \varepsilon_t & if \quad y_{t-1} \geq \tau \end{cases}, \tag{7}$$

where τ is the threshold value and ε_t is an i.i.d. white noise sequence conditional on the history of the series, denoted by $\Omega_{t-1} = \{y_{t-1},..., y_{t-p-1}, y_{t-p}\}$, with zero mean and variance σ^2. Alternatively, this model can be expressed as:

$$y_t = (\phi_0^1 + \phi_1^1 y_{t-1})[1 - I(y_{t-1} \langle \tau)] + (\phi_0^2 + \phi_1^2 y_{t-1}) I(y_{t-1} \geq \tau) + \varepsilon_t, \tag{8}$$

where I(.) is an indicator function that takes a value equal to either one or zero, depending on the regime at time *t*.

For models such as SETAR and others of the TAR type with exogenous regressors, there is a well-developed theory of inference and estimation[3]. In the case of models with endogenous variables, in turn, the theory is still working in progress. Caner and Hansen (2004) develop an estimator and an inference theory for this kind of model, with the restriction that the threshold variable must be exogenous.

A model of this type is as follows. Let $\{y_t, z_t, x_t\}_{t=1}^n$ be the information set, where y_t is unidimensional, z_t is a M-dimension vector (regressors), and x_t is a K-dimension vector (instruments), with K≥M. The threshold variable $q_t = q(x_t)$ can be an element or a function of the vector x_t. In a general way, the structural equation can be written as:

$$\begin{cases} y_t = \theta_1^{'} z_t + \zeta_t & q_t \langle \tau \\ y_t = \theta_2^{'} z_t + \zeta_t & q_t \geq \tau \end{cases}, \tag{9}$$

or in a more compact way,

$$y_t = \theta_1^{'} z_t [1 - I(q_t \langle \tau)] + \theta_2^{'} z_t I(q_t \geq \tau) + \zeta_t, \tag{10}$$

where $\tau \in$ T and T is the set of the possible threshold values.

[3] See, for example, Chan (1993), Hansen (1996, 1999, 2000) and Caner (2002).

As mentioned earlier, in this formulation the error term is correlated with z_t—at least one variable in vector z_t is endogenous—and thus equation (10) cannot be estimated by ordinary least squares because parameters would be biased and inconsistent.

The method proposed by Caner and Hansen (2004) is based on the estimation of a reduced form equation for the endogenous variables as a function of instrument variables, that is, a model with the conditional mean of the endogenous variables as a function of exogenous variables. The estimated values are plugged into structural equation (10) and the threshold value is estimated by minimizing the sum of the squared residuals. The parameters of the structural equation are estimated in a third step, when the sample is divided according to the estimated threshold. The estimation is conducted using the two-stage least square method (2SLS) or the generalized method of moments (GMM).

Therefore, the first stage (conditional expectations model of z_t) is given by:

$$z_t = f(x_t, \beta) + u_t, \qquad (11)$$

$$E(u_t \mid x_t) = 0 \qquad (12)$$

where β is a vector with parameters, u_t is the error term and $f(.,.)$ is a function. In particular, this function can also be conditioned on the threshold value (which can be equal or different from that in the structural equation), becoming:

$$f(x_t, \beta) = (\beta_1^{'} x_t)[1 - I(q_t \langle \tau)] + (\beta_2^{'} x_t)[I(q_t \geq \tau)] \qquad (13)$$

The parameters β in equation (11) can be obtained by OLS, for each $\tau \in T$, as:

$$\hat{\beta}_1(\tau) = \left(\sum_{t=1}^{n} x_t x_t^{'} [I(q_t \langle \tau)] \right)^{-1} \sum_{t=1}^{n} x_t z_t^{'} [I(q_t \langle \tau)],$$

$$\hat{\beta}_2(\tau) = \left(\sum_{t=1}^{n} x_t x_t^{'} [I(q_t \geq \tau)] \right)^{-1} \sum_{t=1}^{n} x_t z_t^{'} [I(q_t \geq \tau)].$$

Using parameters $\hat{\beta}$, we can obtain the values $\hat{z}_t$ that will replace z_t in the structural equation. Doing it recursively for every $\tau \in T$, the threshold value in the structural equation can be chosen by the minimization of the sum of the squared residuals, using a grid search. For every τ, let Y, Z_L and Z_G denote the vector y_t and the matrices $z_t[I(q_t \langle \tau)]$ e $z_t[I(q_t \geq \tau)]$, respectively. Thus, the threshold estimator is obtained from:

$$\hat{\tau} = \arg\min_{\tau \in T} S_n(\tau), \qquad (14)$$

where $S_n(\tau)$ is the sum of the squared residuals in the regression of Y on $\hat{Z}_L$ and $\hat{Z}_G$. The set of threshold values in (14) should be such that each regime has sufficient observations to generate reliable parameter estimation. According to Franses and van Dijk (1999), a safe choice is at least 15% of the sample.

Given the estimated threshold value $\hat{\tau}$, the sample is divided into subsamples, and parameters of the equation (10) can be estimated by 2SLS as:

$$\hat{\theta}_1 = \left[\hat{Z}_L^{'} \hat{X}_L \left(\hat{X}_L^{'} \hat{X}_L \right)^{-1} \hat{X}_L^{'} \hat{Z}_L \right]^{-1} \left[\hat{Z}_L^{'} \hat{X}_L \left(\hat{X}_L^{'} \hat{X}_L \right)^{-1} \hat{X}_L^{'} Y \right] \qquad (15)$$

$$\hat{\theta}_2 = \left[\hat{Z}_G^{'} \hat{X}_G \left(\hat{X}_G^{'} \hat{X}_G \right)^{-1} \hat{X}_G^{'} \hat{Z}_G \right]^{-1} \left[\hat{Z}_G^{'} \hat{X}_G \left(\hat{X}_G^{'} \hat{X}_G \right)^{-1} \hat{X}_G^{'} Y \right], \qquad (16)$$

where $\hat{Z}_L$, $\hat{Z}_G$, $\hat{X}_L$ e $\hat{X}_G$ stand for the matrices with observations $z_t[I(q_t \langle \hat{\tau})]$, $z_t[I(q_t \geq \hat{\tau})]$, $x_t[I(q_t \langle \hat{\tau})]$ and $x_t[I(q_t \geq \hat{\tau})]$, respectively.

Caner and Hansen (2004) show that those estimators are consistent, although not necessarily efficient. Their applicability is conditioned on the exogeneity of the threshold variable.

4. Phillips curve model

Aiming to test the possibility of the presence of non-linear mechanisms in the pass-through from exchange rate to inflation, we estimate some Phillips curve models for Brazil combined with the methodology of regime switching described in the previous section.

The specification of a Phillips curve, which relates inflation to a measure of real disequilibrium (output gap), inflation expectations, past inflation, exchange rate change and external inflation, with a threshold variable, can be formulated as:

$$\begin{cases} \pi_t^L = \alpha_1^1 E_t \pi_{t+1} + \alpha_2^1 \pi_{t-1} + (1 - \alpha_1^1 - \alpha_2^1)^1 (\Delta e_{t-1} + \pi_{t-1}^*) + \alpha_4^1 h_{t-1} + \varepsilon_t & \text{if } q_t \langle \tau \\ \pi_t^L = \alpha_1^2 E_t \pi_{t+1} + \alpha_2^2 \pi_{t-1} + (1 - \alpha_1^2 - \alpha_2^2)^2 (\Delta e_{t-1} + \pi_{t-1}^*) + \alpha_4^2 h_{t-1} + \varepsilon_t & \text{if } q_t \geq \tau \end{cases},$$

where π_t^L is the free IPCA inflation (headline inflation measured by the Broad National Consumer Price Index, but excluding administered prices), π_t is the headline IPCA inflation, π_t^* is a measure of external inflation (PPI in the US), h_t is output gap[4], e_t is the logarithm of the nominal exchange rate, $E_t(.)$ is the expectations operator conditional on the information available at *t*, Δ is the difference operator ($\Delta e_{t-1} = e_{t-1} - e_{t-2}$), q_t is the threshold variable, $\tau \in$ T and T is the set of possible values for q_t.

The dependent variable is the free price inflation instead of the headline inflation because administered prices have a different price dynamics, partially obeying contract rules. Note that the estimated pass-through refers to the pass-through from the previous quarter to the current inflation, that is, it captures only the short-run effect of exchange rate changes.

To enable the joint estimation, the previous equations becomes:

$$\begin{aligned}\pi_t^L = &\left(\alpha_1^1 E_t\pi_{t+1} + \alpha_2^1\pi_{t-1} + (1-\alpha_1^1-\alpha_2^1)^1(\Delta e_{t-1}+\pi_{t-1}^*) + \alpha_4^1 h_{t-1}\right)\left(1 - I[q_t \langle \tau\right) + \\ &\left(\alpha_1^2 E_t\pi_{t+1} + \alpha_2^2\pi_{t-1} + (1-\alpha_1^2-\alpha_2^2)^2(\Delta e_{t-1}+\pi_{t-1}^*) + \alpha_4^2 h_{t-1}\right)\left(I[q_t \geq \tau\right) + \varepsilon_t\end{aligned}.$$

Based on the theoretical indications for a non-linear exchange rate pass-through, we evaluate three threshold variables: i) business cycle, measured by output gap; ii) magnitude of the nominal exchange rate change; and iii) a measure of exchange rate volatility. We use quarterly data from 1995:1 through 2004:4, estimating using 2SLS, with instrument variables for the inflation expectations term.

The first estimated specification has output gap as the threshold variable. In this model, all parameters, except for that of the output gap, are subject to regime switching. The estimation results are the following (the *t* statistics are in parentheses):

$$\underset{}{\pi_t^L} = \underset{(7.88)}{0.72E_t\pi_{t+1}} + \underset{(3.28)}{0.29\pi_{t-1}} \underset{(-0.41)}{- 0.01(\Delta e_{t-1}+\pi_{t-1}^*)} + \underset{(2.33)}{0.15h_{t-1}} \quad if\ h_{t-1} \langle -1.89\% \qquad R^2 = 0.38$$

$$\pi_t^L = \underset{(2.91)}{0.62E_t\pi_{t+1}} + \underset{(1.31)}{0.29\pi_{t-1}} + \underset{(2.48)}{0.09(\Delta e_{t-1}+\pi_{t-1}^*)} + \underset{(2.33)}{0.15h_{t-1}} \quad if\ h_{t-1} \geq -1.89\%.$$

According to that estimation, there is a non-linearity in the pass-through term related to the business cycle: the exchange rate pass-through is not statistically different from zero in the regime where the economy is well-below its capacity, whereas it is around 9% when the economic activity is

[4] The output gap used in the estimation was generated using a production function model. See, for example, the box "Methodologies for estimating the potential output" in the December 2003 Inflation Report of the Central Bank of Brazil and Muinhos and Alves (2003) for a description of the methodology.

higher. The exchange rate pass-through is significantly higher when output is above some threshold[5], estimated in 1.98% below the potential output. One limitation of this result is its implication that exchange rate appreciations have a higher pass-through when the economy is booming than when output gap is below the threshold.

The other estimated parameters are in line with those found in the literature using models without a threshold variable. Except for the backward-looking term, all the coefficients are statistically significant at 5% and their point estimates are not very different across regimes. In addition, the estimated value for the threshold splits the sample into two approximately equal parts—18 observations when $h_{t-1} < -1.89$, and 22 when $h_{t-1} \geq -1.89$. This means that, although the sample size is not large, none of the regimes was estimated with an extremely low number of observations.[6] Actually, we have tried several specifications, using different instruments for the expectations term[7], and the results were very robust.

The second estimated model considers the nominal exchange rate change as the threshold variable. Similarly to the previous model, all parameters are allowed to vary with the regime change, except for the output gap parameter, kept constant in both regimes. The estimation results are the following[8]:

$$\begin{aligned}
\pi^L_t &= \underset{(3.85)}{0.64}E_t\pi_{t+1} + \underset{(2.18)}{0.34}\pi_{t-1} + \underset{(0.85)}{0.02}(\Delta e_{t-1} + \pi^*_{t-1}) + \underset{(2.67)}{0.22}h_{t-1} \qquad if\ \Delta e_{t-1} \langle\ 1.97\% \qquad R^2 = 0.35 \\
\pi^L_t &= \underset{(4.55)}{0.63}E_t\pi_{t+1} + \underset{(1.41)}{0.27}\pi_{t-1} + \underset{(3.05)}{0.10}(\Delta e_{t-1} + \pi^*_{t-1}) + \underset{(2.67)}{0.22}h_{t-1} \qquad if\ \Delta e_{t-1} \geq 1.97\% .
\end{aligned}$$

Those results indicate that the short-run effect of the exchange rate change on inflation is asymmetric. In the case of large exchange rate depreciations, the estimated pass-through for the following quarter is around 10%, whereas appreciations or small depreciations do not have a

5 This result is in line with those in Goldfajn and Werlang (2000), which estimate a panel data for 71 countries and find that appreciations have a higher pass-through to prices when the economy is booming, and Carneiro, Monteiro and Wu (2002), which estimate a backward-looking Phillips curve for Brazil with the pass-through coefficient as a function of unemployment rate and real exchange rate level.

6 Including the period previous to the launch of the Real Plan is not recommendable because the inflation dynamics in a high inflation regime is very different from the current one, which would distort the estimation.

7 We have used the usual procedure of selecting as instruments the exogenous variables at several lags.

8 In that specification, we have used for the backward-looking inflation and output gap terms the average of their values at t-1 and t-2, that is, $\pi_{t-1} = \left(\pi^A_{t-1} + \pi^A_{t-2}\right)/2$ and $h_{t-1} = \left(h^A_{t-1} + h^A_{t-2}\right)/2$, where the superscript A means actual values. That specification has generated better fitting.

statistically significant effect. Therefore, the pass-through is greater when depreciations are equal to or larger than 2%. Although we have not found statistical significance that an appreciation in the previous quarter would affect current inflation, we should not infer that appreciations are not transmitted to prices. This transmission can take place with more lags than in the case of depreciation.

Since the estimated threshold is not zero, a slightly positive value (close to 2%) suggests the presence of menu costs, where small exchange rate changes are not promptly transmitted to prices. In this case, firms tend to postpone their decisions, adjusting their markup in the short-run, and adjust prices only when the exchange rate change is perceived as permanent. If the exchange rate change surpasses some limit, even if the change is temporary, the costs of not adjusting prices are high, leading firms to change prices more rapidly.

As before, the estimated parameters are robust with respect the instruments used and are statistically significant at 5% (except the one on lagged inflation when $\Delta e_{t-1} \geq 1.97\%$). Besides, the number of observations in each regime was balanced (16 observations in the large depreciation regime and 24 in the other) and the estimated values for the coefficients are close to those found in the literature. Note that in both estimations the forward-looking inflation term is greater than the backward-looking component.

One limitation of the estimated models, in particular of the second model, is that it does not make any distinction between the pre-1999 period, when the exchange rate was mostly managed, and the following period of floating rate. In fact, when we include a step dummy into the exchange rate term, the results deteriorate significantly, in terms of both data fitting and signs and statistical significance of the parameters. This result may be related to the increase in the number of parameters to be estimated when we include a dummy variable, reducing the degree of freedom of the estimation. Furthermore, the restriction of the coefficients adding to one becomes more stringent.

Because of those limitations, we have estimated a third model, using the exchange rate volatility as the threshold variable. That estimation would tend to resolve, at least partially, the problem of the separation of the exchange rate regimes (before and after January 1999) because the threshold estimation tend to classify the observations of the managed system period into the low volatility regime. In principle, that estimated regime could also contain observations from when the exchange rate was relatively stable during the floating period.

In addition, that estimation considering exchange rate volatility as the threshold variable aims to capture the inflationary effects in two different situations: i) when agents perceive the exchange rate change as transitory; and ii) when they perceive it as permanent. Our hypothesis is that the probability that agents consider the change as permanent is higher in periods of low exchange rate volatility and lower in periods of greater volatility.[9] Thus, in exchange rate instability periods, agents tend to consider changes as temporary, whereas in periods of greater stability, they would consider them as permanent. Thus, we would expect a lower pass-through in the first case when compared to the latter.

We have used the standard deviation of daily changes in the exchange rate within each quarter as the measure of volatility. The estimation results were the following:

$$\pi_t^L = \underset{(0.71)}{0.36}E_t\pi_{t+1} + \underset{(0.21)}{0.04}\pi_{t-1} + \underset{(0.94)}{0.60}(\Delta e_{t-1} + \pi_{t-1}^*) + \underset{(2.79)}{0.20}h_{t-1} \qquad if \ \sigma_{e,t-1} \langle \ 0.07\% \qquad R^2 = 0.34$$

$$\pi_t^L = \underset{(6.64)}{0.66}E_t\pi_{t+1} + \underset{(2.55)}{0.30}\pi_{t-1} + \underset{(2.34)}{0.04}(\Delta e_{t-1} + \pi_{t-1}^*) + \underset{(2.79)}{0.20}h_{t-1} \qquad if \ \sigma_{e,t-1} \geq 0.07\%.$$

In terms of magnitude, the point estimates indicate a greater pass-through in the low volatility period than in the high volatility one (60% and 4%, respectively). However, the estimated pass-through is not statistically significant in the low volatility regime, although it is significant in the other regime. Additionally, those parameter values are close to those found in the literature for the periods of managing and floating.[10] The resulting sample division assigned most of the observations of the managed system in the low volatility regime. The observations corresponding to values below the threshold comprise the 1995:4–1998:2 period.

Figure 1 presents the quarterly free price inflation—measured in the left axis—, the exchange rate change (regarding the previous period to ease the visualization, because of the lagged effect of one quarter on inflation)—measured in the right axis—and a line that indicates the threshold value (-1.89%) estimated in the model with the threshold determined by the output gap—separating the periods when the output gap is above and below the threshold. We point out some periods, described in Table 1, in which the estimated model can explain, at least partially, the relation between exchange rate and inflation. Table 1 records the corresponding values, besides including the headline inflation (measured by IPCA).

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[9] Albuquerque and Portugal (2005a), for example, explore the relationship between exchange rate volatility and inflation in Brazil, using a bivariate GARCH model.

[10] Muinhos and Alves (2003), por instance, have found a coefficient reduction from 51% to 6% after the change in the exchange rate regime, and Albuquerque and Portugal (2005b), using a Kalman filter model, have estimated parameters values around 42% and 4%, respectively.

Figure 1 – Free price inflation, exchange rate change and output gap

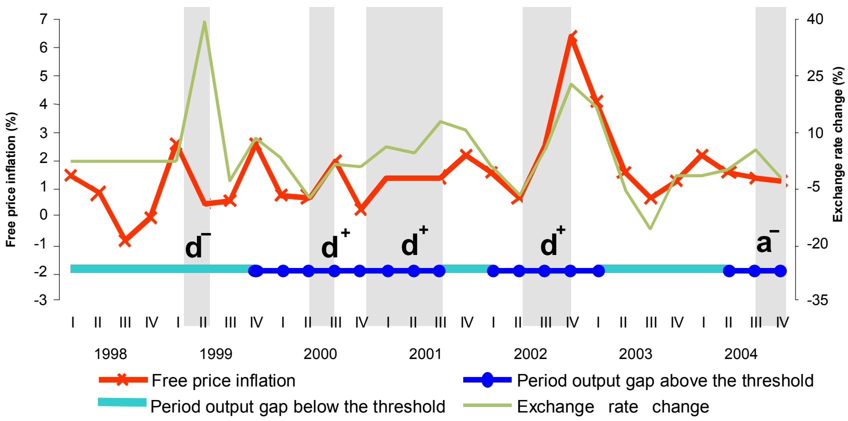

Table 1 – Inflation and exchange rate change in selected periods

Period	Characteristic	Headline inflation in t-1	Headline inflation in t	Free price inflation in t-1	Free price inflation in t	Exchange rate change in t-1
1999: II	d^{-}	2.88	1.05	2.59	0.49	39.33
2000:III	d^{+}	0.66	3.18	0.71	1.93	1.64
2001:II	d^{+}	1.42	1.52	1.42	1.40	4.39
2001:III	d^{+}	1.52	2.33	1.40	1.41	12.69
2002:III	d^{+}	1.44	2.58	0.62	2.56	4.87
2002:IV	d^{+}	2.58	6.56	2.56	6.34	22.30
2004:IV	a^{+}	1.94	2.00	1.35	1.31	-2.30
2005:I	a^{+}	2.00	1.79	1.31	1.72	-6.80

Where: d^{+} means depreciation with a booming economy

d^{-} means depreciation with recession

a^{+} means appreciation with a booming economy

a^{-} means appreciation with recession

Comment: The exchange rate change is calculated based on the quarterly exchange rate average

In the second quarter of 1999 (immediately after the floating), for instance, in spite of the 39% exchange rate depreciation in the previous quarter, free price inflation was only 0.49% and the headline inflation stood at 1,05%, both below the previous quarter values. In that period, output gap was below the estimated threshold (economic slowdown), which implies, according to the model, a low pass-through to inflation. The depreciations in the third quarter of 2000 and during 2001, in turn, were followed by higher increase in the inflation rate. In that period, output was higher than the estimated threshold.

In mid-2002, when the economy was booming again, a strong depreciation was accompanied by a great inflation rise. In the last quarter, for instance, when the depreciation in the previous quarter reached 22%, the free price inflation went from 2.56% to 6.34%, and the headline inflation from 2.58% to 6.56%.

On the other hand, although inflation had fallen along 2005, it did not follow so strongly the exchange rate change in the last quarter of 2004 and first quarter of 2005. One possible explanation lies on the asymmetry of the short-run pass-through with respect to appreciation and depreciation, put in evidence by the model with the threshold given by the exchange rate change. Moreover, the model estimates the short-run pass-through, that is, the effect on current inflation of the exchange rate change of the previous quarter. In 2005, the appreciation ended up affecting prices, but with lags greater than one quarter. Possibly initial movements of appreciation were not perceived immediately as having longer duration, postponing the effect on prices.

5. Conclusions

This paper explored the possibility of the presence of non-linearity in the mechanisms of pass-through from the exchange rate to inflation in Brazil. We have estimated models for the Phillips curve combined with the methodology of threshold models. In those models, the parameter values depend on whether the economy is in one or another regime, which are determined endogenously by means of an observed variable.

The choice of variables used was based on the possible sources of non-linearity of the pass-through reported in the literature. In particular, we have examined three sources: i) business cycle; ii) exchange rate change; and iii) exchange rate volatility.

The estimations indicate the presence of non-linear mechanisms in the short-run pass-through in Brazil. The short-run pass-through is higher when the economy is booming, when the exchange rate depreciates above some threshold, and when the exchange rate volatility is lower. Those results have important implications for monetary policy and are possibly related to a pricing-to-market behavior, menu costs to change prices, and uncertainty about the degree of persistence of exchange rate changes.

Nevertheless, future research is warranted to unveil with more details the exchange rate pass-through mechanism. In particular, working in a more disaggregated level for pricing formation—for example, disaggregating according to industrial sectors—can allow a better comprehension of the extension and reasons behind those non-linear mechanisms.

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