PASS-THROUGH FROM EXCHANGE RATE TO PRICES IN BRAZIL

AN ANALYSIS USING TIME-VARYING PARAMETERS FOR THE 1980 – 2002 PERIOD[†]

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Abstract

The aim of the present paper is to analyze the pass-through from exchange rate to inflation in Brazil from 1980 to 2002. Initially, we developed a model of a profit-maximizing firm based on the pricing-to-market approach presented by FEENSTRA and KENDAL (1997). In order to adapt the model to the Brazilian reality, we considered the following aspects: (i) the firm sells its product both in the domestic market – where it has some pricing power – and in the foreign market – where it is a price-taker; (ii) costs are a function of the exchange rate; (iii) the degree of openness is included in the demand equation. Results show that the Kalman Filter yields better results than linear models with time-invariant parameters and that the inflationary environment and the exchange rate regime perceived by the agents affect the degree of pass-through. We can observe a reduction in the pass-through to consumer price indices (IPCA and IGP-DI) after the implementation of the Real plan, and a more intense reduction after the adoption of the floating exchange rate regime in 1999. These results are in line with other estimates presented in the literature. The pass-through to wholesale prices, however, is relatively constant and its levels are close to one throughout the period. This also seems to be a consistent result if we consider a small (price-taking) economy in the foreign market.

Key words: exchange rate, pass-through, Kalman filter
JEL Classification: E31, F41

I – Introduction

There have been several studies about the effects of exchange rate movements on the economy, especially on prices. However, most of these studies are concerned with developed economies, which behave differently from the Brazilian economy. Moreover, few of these studies focus on Brazil, and many of them fail to use a longer study period that includes the years prior to the price stabilization brought about by the Real Plan. Furthermore, new econometric techniques developed in recent years have made it possible to analyze the relationship between exchange rate and prices, although such alternatives remain underexplored even in industrialized countries.

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The effects of exchange rate movements on prices in different economic scenarios are of paramount importance in order for us to evaluate whether they depend upon the macroeconomic environment, as this information is relevant for monetary policy decisions. Evidence suggests that such relation exists; an example is the different pass-through behavior of developed and emerging economies. According to CALVO and REINHART (2000), emerging economies showed a pass-through from exchange rate to inflation about four times higher than that of developed economies, and the variance of inflation compared to exchange rate variation was 43% for emerging economies and 13% for developed ones. The authors conclude that there is a lower tolerance to exchange rate fluctuations in emerging economies.

The impact of an exchange rate devaluation on prices is both direct, through an increase in import prices, and indirect, through the effects on aggregate demand. In the first case, the increase results from the share of imports in the price index as well as from the rise in input costs. On top of that, devaluation also places pressure for nominal wages to rise, due to the change in real wage. In the second case, the effects on aggregate demand are due to (i) changes in the relationship between foreign and domestic prices, (ii) the effects on interest rates, since the foreign capital movements are affected, and (iii) the wealth effect, since there may possibly be a relevant number of firms that hold foreign exchange positions. The change in the expenditure structure (between domestic and imported goods) will be greater the higher the price-elasticity of exports and imports, and the degree of openness of the economy (LOSCHIAVO and IGLESIAS, 2002).

AMITRANO, GRAUWE and TULLIO (1997) describe the following three stages in the pass-through of exchange rate devaluation to domestic inflation:

1) Pass-through to import prices: since there is a second-order effect on profit, which increases the average revenue and decreases the quantity demanded, the increase in profit depends on the demand elasticity. As the prices constitute a mark-up over costs, exporters might not increase them, especially if there are menu costs as well as expectations that devaluation is temporary;

2) Pass-through from import prices to domestic prices: the degree of pass-through depends on the characteristics of the economy: the more open an economy is (i.e., stronger presence of imported goods), the higher the impact of the increase in import prices over the domestic prices.

3) price behavior after devaluation: price adjustment leads to changes in nominal wages. The degree of price adjustment depends on whether the economy is in a recession or on whether there is a restrictive fiscal policy, so as to avoid the price-wage spiral.

The studies on the exchange rate pass-through originate from the investigation of the validity of the purchasing power parity (PPP) theory. After the devaluation of the US dollar in the 1970s, US price levels did not increase as much as the exchange rate, seemingly casting some doubt on the validity of the PPP theory. Many studies were carried out[3] to test the PPP, but the conclusion is that the parity is valid in the long run but not in the short run, that is, the pass-through from exchange rates to prices is incomplete.

Another finding is that the volatility of PPP deviations could have remained stable over time. According to ROGOFF(1996), the reasons for such deviations should not be restricted to institutional factors that are specific to the 20th century. KLEIN (1990) reminds us that the difference in the price effects between the US dollar devaluation in 1977-81 and after 1985 and its appreciation in 1982-85 offers the following empirical evidence: the pass-through is unstable and its change over time is a result of the structure of the economy. EINCHENGREEN (2002) highlights that the pass-through is not independent of the monetary regime. If the commitment to inflation control is serious and if monetary policy decisions are clear, the agents will reckon the validity of a temporary exchange rate shock by the monetary authority as very unlikely, taking longer to adjust their prices in response to a change in the exchange rate. Therefore, if the pass-through is high, the short-term effect of a change in the exchange rate will be stronger on inflation than on the product, due to the reluctance to adopt a tighter monetary policy. FRANKEL (1978) found evidence of PPP in hyperinflations, which was already expected due to the predominance of monetary shocks in such situations. However, the tests rejected the parity for more stable monetary environments. All of the studies conducted reached the following conclusions: (i) real exchange rates converge to PPP in the very long term at too low a speed of convergence, and (ii) short-term deviations from PPP are high and volatile (ROGOFF, 1996).

Based on these results, the economic theory attempted to explain such deviations. The following explanations arose: the role of nontradeables in the economy (ROGOFF, 1996), the existence of sticky prices that may influence relative prices (DORNBUSCH, 1976), adjustment costs (DIXIT, 1989, KRUGMAN, 1988) and the existence of pricing-to-market (KRUGMAN, 1986). For applications of these theories, see KIMBROUGH (1983), FISHER (1989), GOLDBERG and KNETTER (1996), VIAENE and VRIES (1992), PARSLEY (1995), ANDERSEN (1997), FEENSTRA and KENDAL (1997), BORENZSTEIN and de GREGORIO (1999), SMITH (1999), BETTS and DEVEREUX (2000), GOLDFJAN and WERLANG (2000), LIEDERMAN and BAR-OR (2000), OBSTFELD and ROGOFF (2000), TAYLOR (2000). Among the studies about Brazil, we highlight those carried out by FIORENCIO and MOREIRA (1999), BOGDANSKI, TOMBINI and WERLANG (2000), MUINHOS (2001), CARNEIRO, MONTEIRO and WU (2002), FIGUEIREDO and FERREIRA (2002), BELAISCH (2003), MUINHOS and ALVES (2003) and MINELLA, FREITAS, GOLDFAJN and MUINHOS (2003).

[3] For a good literature review on pass-through and PPP tests, see GOLDBERG and KNETTER, (1996) and KLAASSEN (1999), respectively.

II – The theoretical model

We developed a model of a domestic firm that may choose between selling its production in the domestic or in the foreign market, or in both. The price for period *t* is set in *t-1* by maximizing the expected profit. Models like this can be found in most pass-through studies that use the pricing-to-market approach. The difference here is that those models consider that the firm sells only to the foreign market and, hence, its decision is concerned with foreign prices. The model developed here considers that the firm is a perfect competitor (therefore a price taker) in the foreign market, but has some market power domestically. Hence, the firm can choose domestic prices, given its domestic and foreign demand, foreign prices, and costs involved.

Our model is based on FEENSTRA and KENDAL (1997), with some pertinent changes : (i) as previously mentioned, the decision concerns domestic prices; (ii) we consider the presence of imported inputs, implying that costs are a function of the exchange rate; (iii) we include the degree of openness in the demand function. The following equations define the model.

The total revenue of the firm in the domestic market is given by:

$$RT^{dom} = p.x^{dom}(p^{dom}, p^{imp}, y, ope)$$

The total revenue resulting from exports and expressed in domestic currency is:

$$RT^{\exp} = s.p^{\exp}.x^{ext}(p^{\exp}, p^{*}, y^{*})$$

where p and p^{exp} are the prices charged by the firm in the domestic and foreign market, x^{dom} and x^{ext} are the domestic and foreign demand, p^* is the price of competitors in the foreign market, p^{imp} is the price of imports competing with the firm's product domestically, and y and y^* are the domestic and foreign income, respectively. The nominal exchange rate, expressed in domestic currency units per foreign currency, is given by s. The variable *ope* represents the degree of openness, included here for its relevance in explaining inflation, as pointed out by several authors. The use of this variable is justified by the studies of TERRA (1998) and ROMER (1993,1998), and by the contagion of domestic price indices by the higher presence of import goods. The degree of openness is regarded as a proxy for the competition faced by domestic products, being therefore a relevant variable in the demand function.

According to FEENSTRA and KENDAL (1997), the firm sells z units of currency in the future market at price f_t in order to protect itself from the exchange rate risk. Thus, its profit (or loss) with the transaction is given by z (f_t – s_t). Exchange rate protection is also a firm's decision variable.

Hence, the firm's profit is given by:

$$\Pi_t = (p_t - c_t) * x_t^{dom}(p_t, p_t^{imp}, y_t, ope) + (s_t.p_t^{\exp} - c_t^{*}) * x_t^{\exp}(p_t^{\exp}, p_t^{*}, y_t^{*}) + z_t * (f_t - s_t)$$

The firm maximizes the expected utility of profits. Then, the problem of the firm is given by:

$$\underset{p_t, z_t}{Max} E_{t-1}\{U[(p_t - c_t) * x_t^{dom}(p_t, p_t^{imp}, y_t, ope) + (s_t.p_t^{\exp} - c_t^{*}) * x_t^{\exp}(p_t^{\exp}, p_t^{*}, y_t^{*}) + z_t(f_t - s_t)]\} \quad (1)$$

Using a second-order Taylor's expansion, we have[4]:

$$U(\Pi_t) \approx U(E_{t-1}[\Pi_t]) + U'(E_{t-1}[\Pi_t]) * (\Pi_t - (E_{t-1}[\Pi_t]) + \tfrac{1}{2} U''(E_{t-1}[\Pi_t]) * (\Pi_t - E_{t-1}[\Pi_t])^2$$

However, it is known that (Π_t -($E_{t-1}[\Pi_t]$)) = 0 and that (Π_t - $E_{t-1}[\Pi_t]$)^2 is the conditional variance of profits, herein referred to as $var_{t-1}(\Pi_t)$. Such considerations allow us to rewrite (1) as:

$$\underset{p_{t(1'},z_t}{Max} \{U(E_{t-1}[\Pi_t]) + \tfrac{1}{2} U''(E_{t-1}[\Pi_t]) * \mathrm{var}_{t-1}(\Pi_t)\} \qquad (1')$$

To calculate $E_{t-1}[\Pi_t]$, conditional mean of Π_t, consider $E_{t-1}p_t^{exp} = p_t^{exp}$ and $E_{t-1}x_t^{exp} = x_t^{exp}$ This supposition can be made if we consider that foreign contracts for sales in *t* are negotiated in *t-1*. So, we have:

$$E_{t-1}[\Pi] = p_t * E_{t-1}[x_t^{dom}(p_t, p_t^{imp}, y_t, ope)] - E_{t-1}(c_t) * x_t^{dom}(p_t, p_t^{imp}, y_t) + E_{t-1}(s_t) * p_t^{\exp} * x_t^{\exp}(p_t^{\exp}, p_t^*, y_t^*) + z_t.f_t - z_t.s_t -$$
$$- c_t * x_t^{\exp}(p_t^{\exp}, p_t^*, y_t^*)$$

Supposing $E_{t-1}(s_t) = e_t$ and by rearranging the terms above, we obtain:

$$E_{t-1}[\Pi_t] = E_{t-1}[x_t^{dom}(p_t, p_t^{imp}, y_t, ope)] * (p_t - E_{t-1}(c_t)) + x_t^{\exp}(p_t^{\exp}, p_t^*, y_t^*) * (e_t p^{\exp} - E_{t-1}(c_t^*)) +$$
$$z_t * (f_t - e_t) \qquad (2)$$

Calculating the conditional variance of profits and naming $E_{t-1}[(s_t-e_t)^2]$, which is the conditional variance of the exchange rate, as σ^2_s, we have:

$$\mathrm{var}_{t-1}(\Pi_t) = E_{t-1}\{[\Pi_t - E_{t-1}(\Pi_t)]^2\}$$
$$\mathrm{var}_{t-1}(\Pi_t) = \sigma_s^2 * [x_t^{\exp}(p_t^{\exp}, p_t^*, y_t^*) - z_t]^2 \qquad (3)$$

Using (2) and (3), (1') can be rewritten as:

$$\underset{p_t,z_t}{Max}\{U[E_{t-1}[x_t^{dom}(p_t, p_t^{imp}, y_t, ope)] * (p_t - E_{t-1}(c_t)) + x_t^{\exp}(p_t^{\exp}, p_t^*, y_t^*) * (e_t.p^{\exp} - E_{t-1}(c_t^*)) +$$
$$z * (f_t - e_t)] + \tfrac{1}{2} U''[E_{t-1}(\Pi_t)] * \sigma_s^2 * [x_t^{\exp}(p_t^{\exp}, p_t^*, y_t^*) - z_t]^2\} \qquad (1'')$$

Deriving the equation above in relation to z_t, a first-order condition is that:

$$U'(E_{t-1}(\Pi_t)) * (f_t - e_t) - U''(E_{t-1}(\Pi_t)) * \sigma_s^2 * x_t^{\exp}(p_t^{\exp}, p_t^*, y_t^*) * p_t^{\exp} - z_t^*) = 0$$

From where it follows that:

$$z_t^* = [-U'(E_{t-1}(\Pi_t)) * (f_t - e_t) / U''(E_{t-1}(\Pi_t)) * \sigma_s^2] + x_t^{\exp}(p_t^{\exp}, p_t^*, y_t^*) * p_t^{\exp}$$

However, $-U'(E_{t-1}(\Pi_t)) / U''(E_{t-1}(\Pi_t))$ is the inverse of the Arrow-Pratt absolute risk aversion coefficient (R_u). Hence,

$$z_t^* = ((f_t - e_t) / R_u * \sigma_s^2) + x_t^{\exp}(p_t^{\exp}, p_t^*, y_t^*) * p_t^{\exp} \qquad (4)$$

The optimal future contract has a term that represents the speculative purchase (or sale) of foreign currency, and a second term that corresponds to the contribution of foreign sales to the total revenue of the firm[5].

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[4] It is necessary to disregard the rest in the equation since, otherwise, it would be necessary to incorporate the term U'''(.) – third derivative of the utility function– about which the economic theory has no assumptions.

Using (2), (3) and (4), equation (1) can be rewritten as:

$$\underset{p_t, z_t}{Max}\{U[E_{t-1}[x_t^{dom}(p_t, p_t^{imp}, y_t, ope)]*(p_t - E_{t-1}(c_t)) + x_t^{\exp}(p_t^{\exp}, p_t^*, y_t^*)*(e_t p^{\exp} - E_{t-1}(c_t^*)) +$$
$$[(f_t - e_t)/R_u * \sigma_s^2) + x_t^{\exp}(p_t^{\exp}, p_t^*, y_t^*) * p_t^{\exp}]*(f_t - e_t)] + \tfrac{1}{2}U''[E_{t-1}(\Pi_t)]*\sigma_s^2 * (x_t^{\exp}(p^{\exp}, p_t^*, y_t^*) -$$
$$(f_t - e_t)/R_u * \sigma_s^2) + x_t^{\exp}(p_t^{\exp}, p_t^*, y_t^*) * p_t^{\exp}]^2 \qquad \mathbf{(1''')}$$

The first-order condition with respect to p_t is:

$$U'E_{t-1}(\Pi_t))*\{[\delta E_{t-1}(x_t^{dom}(p_t, p_t^{imp}, y_t, ope)) \big/ \partial p_t]*(p_t - E_{t-1}(c_t)) + E_{t-1}(x_t^{dom}(p_t, p_t^{imp}, y_t) - R_u^{-2} * [(f_t - e_t)^2 / \sigma_s^2] * \delta R_u \big/ \delta p_t\}$$
$$-\tfrac{1}{2}U''(E_{t-1}(\Pi_t))\sigma_s^2 * \{-2 * R_u^{-3} * [(f_t - e_t)^2 / \sigma_s^2] * \delta R_u \big/ \delta p_t\} = 0$$

From where we get p_t, the optimal price to be charged by the firm:

$$p_t = E_{t-1}(c)_t - E_{t-1}[\, x_t^{dom}(p_t, p_t^{imp}, y_t, ope)]/\, \eta_t \quad (5)$$

where η_t is the price-elasticity of demand $(\delta E_{t-1}(x_t^{dom}(p_t, p_t^{imp}, y_t, ope))/\delta p_t)$.

The next step is to transform the equation above into an equation that can be tested empirically. To do that, we need to make a few assumptions concerning the demand and cost functions. Let us consider the demand function presented in FEENSTRA and KENDAL (1997):

$$x_t^{dom}(p_t, p_t^{imp}, y_t) = (\alpha / p_t) - (\beta / p_t^{imp}) * y$$

This function has the requested properties, that is, it is decreasing in the domestic price and increasing in the price of the imported competitor and in the income. Besides, as pointed out by the authors, such a function allows the demand for the domestic product to be null. This will happen for domestic and imported price levels that are sufficiently high, for the whole market to be supplied. In such case, the local product will be demanded if $p_t < q_t(\alpha/\beta)$.As previously mentioned, a difference in relation to the original study is that here we will consider that variable y will not be regarded as income but as the deviation from the potential product instead[6].

Let us also consider that the domestic price of the imported good in an imperfect market depends not only on the actual import price but also on the presence of other foreign competitors in the same market. Thus, the higher the degree of openness, the less freedom the importer will have to pass elevated mark-ups on to the consumer. Therefore, we consider that some weight is placed on competition when setting the prices for the consumer. So p^{imp} is given by:

$$p^{imp} = (p^M)^{\phi}.ope^{-\vartheta}, \qquad \phi > 0, \vartheta > 0$$

where p^M is the price imports have when they arrive in Brazil and *ope* is the degree of openness of the economy. The demand function has the following form:

[5] This result is similar to the one presented by FEENSTRA and KENDAL (1997). The difference lies in the second term, which, in that work, is the total revenue the firm should obtain with external sales expressed in domestic currency.

[6] The product deviation from its natural level, as a proxy for idle capacity, is a relevant variable in pass-through and inflation studies. The idea is that during a recession (i.e., with high idle capacity), there is more difficulty in passing cost increases on to final prices

$$x_t^{dom}(p_t, p_t^{imp}, y_t, ope) = [(\alpha / p) - (\beta.(p^M)^{-\phi}.ope^{\vartheta})]y$$

The degree of openness in the function above also presents the required properties. Deriving the demand function in relation to the variable *ope*, we observe a negative sign: the higher the degree of openness (and hence, market competition), the smaller the demand for a certain product. Likewise, using the function above to derive *p* in relation to *ope*, the sign is also negative[7]. This sign is expected because, according to the literature, there is an inverse relationship between inflation and the degree of openness, whose reasons may be found, for instance, in TERRA (1998) and ROMER (1993, 1998). According to TERRA (1998), there is a negative relationship between inflation and the degree of openness in economies with high level of external indebtedness, since, if the major part of the debt belongs to the public sector, taxes will have to be increased. The less open an economy is, the higher the exchange rate devaluation required to produce trade surpluses, leading to an increase in the liabilities expressed in domestic currency and, hence, a greater need to obtain revenues through the inflationary tax. ROMER (1993, 1998) also establishes a negative relationship, but the cause lies in the implicit commitment of the monetary policy: the more closed an economy is, the greater the benefits of a surprise inflation will be[8].

Substituting the demand function above in equation (5) we have:

$$P^2 = E_{t-1}(c_t) * \left[\left(\alpha \big/ \beta \right) y_{t-1}^{-1} P_{t-1}^{M^{-\phi}} ope_{t-1}^{\vartheta} \right] \qquad (6)$$

Next, some assumptions should be made about the costs. Since there are imported inputs, let us consider that the costs are an increasing function of the exchange rate, assuming the form $c_t = As^{\theta}$. Let us also consider that the purchase of inputs to produce goods in *t* is made in *t-1*, therefore applying the exchange rate in effect at the time. Hence, $c_t = As_{t-1}^{\theta}$ [9].

Thus, the price equation is expressed as $P_t^2 = As_{t-1}^{\theta} \left(\alpha \big/ \beta \right) y_{t-1}^{-1} P_{t-1}^{M^{-\phi}} ope_{t-1}^{\vartheta}$. Applying the natural logarithm on both sides of the equation:

$$\ln p_t = (1/2)\ln(A\alpha / \beta) + 1/2\,\theta \ln(s_{t-1}) - 1/2\,\phi \ln(p_{t-1}^M) + 1/2 \ln y + \vartheta / 2 \ln(ope) \qquad \textbf{(6´)}$$

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[7] $\delta x(.) / \delta ope = - \vartheta\beta y(p^M)^{-\phi} ope^{\vartheta-1}] < 0$ e $\delta p(.)/\delta ope = -\alpha\vartheta\beta y(p^M)^{-\phi} ope^{\vartheta-1} / (x+\beta y(p^M)^{-\phi} ope^{\vartheta})^2 < 0$.

[8] However, one must be aware of the difference between the negative relationship between the degree of openness and inflation and the positive relationship between the degree of openness and the exchange rate pass-through, as recalled by GOLDFAJN and WERLANG(2000). The latter relationship is positive because a more open economy means a higher presence of imported goods in the price index. The higher the contribution of imported goods, the higher the increase in the price index whenever an exchange rate devaluation occurs.

[9] Other assumptions may be made in order to remove the expectation operator from the equation. One of them consists in adopting the assumption of FEENSTRA and KENDAL (1997). If costs follow a time process such as $\ln c_t = \ln c_{t-1} + \varepsilon_t$, where $\varepsilon_t = \varepsilon_{t-1} + v_t$ (v_t is a white noise), then $E_{t-1}(\ln c_t)$ is equal to $\ln c_{t-1}$ plus a residual term. However, the authors do not consider costs as a function of exchange rates, but we can reach the same conclusion if we assume such relationship and if we also consider that the exchange rate follows a random walk as the one described here. Considering that costs are negotiated in (t-1) to be paid in t with the exchange rate in effect at that period would add some algebric complexity to the solution, since we would have to consider the term $E_{t-1}(s_t)$ throughout the exercise. For simplification, we chose the first alternative presented here.

Generalizing,

$$\ln p_t = \mu + \alpha_1 \ln(s_{t-1}) + \alpha_2 \ln(y_{t-1}) + \alpha_3 \ln(ope_{t-1}) + \alpha_4 \ln(p^M_{t-1}) + \varepsilon_t \qquad (6'')$$

where $\gamma_0 = (½)\ln(A\alpha/\beta)$, $\alpha_1 = (½)\theta$, $\alpha_2 = ½$, $\alpha_3 = ½\vartheta$, $\alpha_4 = ½\,\phi$ and ε_t is a white noise error.

First, a linear model with constant parameters will be used to test equation 6´´, in order to check whether the parameters changed along the study period, especially parameter α_1. If there is any evidence of time instability, a specification with time-varying parameters will be tested through the Kalman Filter, an algorithm used to compute the optimal value of a state vector.

III – Data

We used a quarterly sample, from 1980 to 2002, and data were obtained from the websites of IPEA and of the Central Bank of Brazil[10]. To deseasonalize the series, we used the X-11 method. The following variables were used:

a) *igp_des*: Deseasonalized generalized price index _ internal availability (IGP-DI /FGV) ;
b) *ipa_des*: Deseasonalized wholesale price index (IPA/FGV),
c) *ipca_des*: Deseasonalized broad consumer price index (IPCA/IBGE);
d) *cambio*: Nominal exchange rate, selling values, in Brazilian Reais vis-à-vis US dollars, monthly average;
e) *gap*: Deviation of GDP from its potential level. The first step for its calculation consisted in deseasonalizing the GDP series provided by IBGE. After that, the trend of the series was extracted using the Hoddrick-Prescott filter. The difference between the observed value and the trend calculated by the Hoddrick-Prescott filter is the proxy for the GDP deviation from its potential level. The expected sign of this variable in the price level is positive: if GDP is below its level, a decrease in the product implies a decrease in its gap value, which becomes more negative. Thus, we expect an increase in recession to reduce prices. If the economy is strong, with GDP above its potential level, an increase in the GDP – which causes a rise in prices – increases the positive gap value;
f) *ope*: Represents the degree of openness of the economy. It is calculated as the ratio between the sum of exports and imports and the GDP. As previously mentioned, the response of prices to this variable has a negative sign;
g) *p_imp_des*: Refers to the deseasonalized import price index. A rise in import prices is expected to increase prices directly, due to the presence of imported goods in the price index, and indirectly, due to its presence in production costs.

[10] http://www.ipeadata.gov.br and http://www.bcb.gov.br , respectively.

IV– Empirical analysis

IV.1 – Stationarity and cointegration

The first step before working with the series is to check whether they are stationary. In this regard, Tables 1 and 2 show the results of stationarity tests in level and in first difference, respectively. The optimal number of lags for AD&F tests was based on Akaike information criteria.

Note that the *variable gap* is stationary, while the variables *openess, ipca_des, ipa_des* and *preço_imp_des* have unit roots, both on the *Augmented Dick & Fuller* test (AD&F) and on the Phillip-Perron test (PP). For the variable *cambio*, the ADF test described the series as stationary, whereas the PP test indicated the presence of unit roots[11]. For the IGP-DI, the ADF test shows the series as nonstationary – at a 5% significance level – both in level and in first difference. The PP test presented the series as being I(1). In Table 2, the variables *ope, e, p, pm, igp, ipa* express the variables *openess, cambio, ipca_des, preço_imp_des, igp_des* and *ipa_des*, respectively, in first difference.

TABLE 1 –Stationarity test for variables in levels

Variable	ADF test statistics	Phillip-Perron test statistics
cambio	-2.0976(3)*[b]	-1.0816[a]
gap	-2.8529(5)*[a]	-4.3333*[a]
ipca_des	-1.3636(2)[a]	-0.2156
preço_imp_des	-3.2960(0)	-0.7077[b]
igp_des	-1.7115(2)[b]	-0.2722
ipa_des	-1.7339(5)[b]	-0.2911
ope	-2.1726(8)	-1.4139[b]

* Null hypothesis of the presence of unit root rejected at 5%
[a] Test made without a trend term; [b]Test made without trend and intercept terms; figures between parentheses indicate the optimal number of lags for the test.

[11] The Dickey-Fuller GLS and Kwiatkowski-Phillips-Schimidt-Shin tests also identified "câmbio" as I(1).

TABLE 2 – ADF stationarity test for variables in first difference

Variable	ADF test statistics	Phillip-Perron test statistics
ope	-3.1675**(7)	-38.2634*
e	-2.4634(2)[a]	-3.9883*
p	-3.5620(0)*	-3.3988*
pm	-10.5994(0)*	-10.6048*
igp	-2.4602(2)[a]	-3.3673*
ipa	-2.6223(2)[a**]	-3.0297[a*]

*, ** Null hypothesis of the presence of unit roots rejected at 5% and at 10%, respectively
[a] Test made without a trend term; [b]Test made without trend and intercept terms; figures between parentheses indicate the optimal number of lags for the test.

TABLE 3 – Cointegration test - IPCA

Series: IPCA_DES OPENNESS CAMBIO PRECO_IMP_DES
Lags: (in first differences): 1 to 2

Trend assumption: deterministic linear trend

Unrestricted cointegration rank test

Hypothesized No. of CE(s)	Eigenvalue	Trace test	5% critical value	1% critical value
None	0.224714	32.09613	47.21	54.46
At most 1	0.072061	9.443602	29.68	35.65
At most 2	0.029431	2.787392	15.41	20.04
At most 3	0.001445	0.128677	3.76	6.65

*(**) rejection of the null hypothesis at 5%(1%); the trace test indicates no cointegration equation at 5% and at 1%

TABLE 4 – Cointegration test - IGP-DI

Séries: IGP_DES OPENNESS CAMBIO PRECO_IMP_DES
Lags: (in first differences): 1 to 2

Trend assumption: deterministic linear trend

Unrestricted cointegration rank test

Hypothesized No. of CE(s)	Eigenvalue	Trace test	5% critical value	1% critical value
None	0.248336	37.23425	47.21	54.46
At most 1	0.077971	11.82773	29.68	35.65
At most 2	0.049677	4.602857	15.41	20.04
At most 3	0.000764	0.067981	3.76	6.65

*(**) rejection of the null hypothesis at 5%(1%); the trace test indicates no cointegration equation at 5% and at 1%

TABLE 5 – Cointegration test – IPA-DI

Series: IPA_DES OPENNESS CAMBIO PRECO_IMP_DES
Lags: (in first differences): 1 to 2

Trend assumption: deterministic linear trend

Unrestricted cointegration rank test

Hypothesized No. of CE(s)	Eigenvalue	Trace test	5% critical value	1% critical value
None	0.270498	41.30977	47.21	54.46
At most 1	0.082633	13.23980	29.68	35.65
At most 2	0.058054	5.563713	15.41	20.04
At most 3	0.002702	0.240829	3.76	6.65

*(**) rejection of the null hypothesis at 5%(1%); the trace test indicates no cointegration equation at 5% and at 1%

With regard to cointegration tests, Tables 3 to 5 show that cointegration vectors are not present in any of the three cases considered (IPCA, IGP-DI, IPA). Therefore, we have to use the first difference of the variables.

IV.2 – LINEAR MODELS

First we tested the model using OLS. For all price indices, as shown in Appendix I, the models showed specification errors (Reset test), parameter and/or variance instability (CUSUM of squares tests), autocorrelation of residuals (for IGP and IPA) and Arch residuals (for IPA). Given these results and the previous knowledge about changes in the inflation pattern and exchange rate policy after 1994, we made two attempts to model these changes: to split the sample into two periods and to include dummy variables in the exchange rate coefficient.

The two subsamples refer to the pre- and post-Real periods, and are used to check whether there are significant changes in the parameters in these periods. The first subsample, covering the pre-Real plan, goes up to 1993, while the second one starts in 1995. The year 1994 was not included in any of the samples because we consider it as a transition period, where agents could predict the changes in the monetary policy. Inflation indices were still influenced by the high inflationary levels of the previous period. Hence, the pattern observed in 1994 may be neither characteristic of the pre-Real period nor of the post-Real one.

The inclusion of dummy variables to indicate the three major periods of Brazilian monetary policy concerning inflation and exchange rate aims at verifying whether such inclusion is enough to model the breaks suggested in the previous analysis. The model will therefore have the following form:

$$P = \mu + (\alpha_1 + \alpha_{14} d4 + \alpha_{15} d5) * e_{t-1} + \alpha_2 * gap_{t-1} + \alpha_3 * ope_{t-1} + \alpha_4 * PM_{t-1} + \varepsilon_t \qquad (7)$$

Dummy variables *d4* and *d5* represent the post-Real period with pegged and floating exchange rates, respectively, with unity values assigned to the periods t they intend to represent. Thus, the exchange rate coefficient for the pre-Real period is α_1, while for the 1994:III to 1998:IV period it is $\alpha_1+\alpha_4$,and $\alpha_1 + \alpha_5$ between 1999:I and 2002:IV.

Splitting the period into two subsamples is not enough to eliminate specification errors, parameter instability, and presence of autocorrelation (see Appendix I). The inclusion of dummy variables allowed correcting such problems and perceiving the change in the exchange rate coefficient after the Real Plan. However, we cannot reject the null hypothesis that $\alpha_1 + \alpha_{15}$ equals zero, which means that the pass-through from exchange rate to inflation after 1999 is null – a contradiction from the economic standpoint.

Parameter instability, significant differences between the variables for the post-Real sample, specification errors in the model pointed out by Reset tests may be an indicative sign that analyzing the period using models with time-invariant parameters is not the most adequate approach, especially when we consider the exchange rate pass-through, whose behavior changed after 1995.

The inclusion of dummy variables in the exchange rate coefficient – as in equation 7 – avoids some of the problems detected by the tests, especially residual autocorrelation and parameter and/or variance instability. However, if on the one hand the tests revealed some instability in the exchange rate coefficient, on the other hand, they yielded results which do not seem coherent since the pass-through from exchange rate to inflation is statistically null for the three indices. The correction of the instability and autocorrelation by means of dummy variables shows that the exchange rate coefficient should not be regarded as time-invariant. The incapability of such dummy variables to identify the changes that occurred after the floating exchange rate regime led us to test the initial model using the Kalman Filter.

IV.3 –The Kalman Filter Analysis

The Kalman Filter was applied in a general-to-specific process. First, we tested a model where all coefficients were stochastic. Afterwards, we restricted the number of stochastic coefficients based upon the statistical significance of the variance coefficient in the state equation and upon the information criteria. Thus, the variables whose state variance coefficient (parameters $\vartheta_{\mu,t}$ and $\vartheta_{\alpha it}$ in equations *8*) were not significant were considered as having time-invariant parameters. The advantage of such procedure is that if we consider that only the exchange rate coefficient is stochastic and that other coefficients vary over time, the results found for the exchange rate will incorporate the movements in those coefficients regarded as time-invariant.

Another decision refers to the space equation format, i.e., whether it is a random walk or an AR(1) process. In the first case, the effects of the stochastic coefficients are assumed to be permanent, whereas in the second case, the effects, although persistent, are regarded as temporary. Since exchange rate shocks are not permanent – once they are passed on to prices, in different degrees,– we adopted the AR(1) format for the space equation. If the estimated AR parameter is close to unity, a random walk formulation will be tested.

We tested the inclusion of dummy variables in the state equation of the exchange rate parameter in order to verify whether the exchange rate regime or the price dynamics affects that equation. Thus, three dummy variables were tested. The first of them – d_1 – assumed a value equal to zero for periods when there was a managed exchange rate system in Brazil, and equal to one for the other periods with officially floating exchange rates (March 1990 to February 1995 and January 1999 to December 2002). The second dummy variable – d_2 – differs from d_1 as it assumes a value equal to the unit in those periods with effectively and not only officially floating exchange rates. Therefore, d_2 refers to the period known as managed exchange rate system. "Managed" means that the monetary authorities interfere in the exchange rate market, but have no intention to maintain the exchange rate stable at a given level regarded as ideal by the government, as occurred between March 1990 and July 1994 (see ARAUJO and FILHO (2002)). Thus, d_2 assumes a unit value from July to September 1994 and from January 1999 on, and a zero value for the other periods. The purpose of such distinction is to verify whether the announced exchange rate regime is relevant to price setting or how exchange rates behave in practice. Finally, the third dummy variable – d_3 – aims at comparing the price dynamics in high-inflation periods with stable periods. Therefore, d_3 has a unit value for the pre-Real period, and a zero value for the post-Real period.

Chart 1 – Tested Dummies

Dummy	**Plan**	**Period with a unit value**
D_1	Officially floating exchange rates	1990:I – 1995:I; 1999:I – 2002:IV
D_2	Managed exchange rate system	1994:III ; 1999:I – 2002:IV
D_3	High inflation	1980:I – 1994:II
D_4	Real Plan with pegged exchange rates	1994:III – 1998:IV
D_5	Real Plan with floating exchange rates	1999:I – 2002:IV
D_{cruz}	Cruzado plan	1986:I – 1986:III
D_{bress}	Bresser plan	1988:III
D_{ver}	Summer plan	1989:I
D_{col1}	Collor 1 plan	1990:I – 1990: II
D_{col2}	Collor 2 plan	1991:I – 1991:II
D_{real}	Real plan	1994:III – 2002:IV

We also tested dummy variables related to the economic plans announced in the course of the study period, according to CATI, GARCIA and PERRON (1999). The variables have a unit value throughout the period in which economic plans were in effect, and a zero value for the other periods, except for the post-Real period, which was not included in the referred paper. Two dummies were assigned to the post-Real period: d_4, with a unit value between July 1994 and December 1998, and d_5 with a unit value after the exchange rate devaluation in 1999. The dummy variables are shown in Chart 1.

The initial model had therefore the following form[12]:

$$\begin{aligned}
&p = \mu_t + \alpha_{1,t} e_{t-1} + \alpha_{2,t} gap_{t-1} + \alpha_{3,t} ope_{t-1} + \alpha_{4,t} pm_{t-1} + \varepsilon_t \\
&\mu_t = c_1 + c_2 \mu_{t-1} + \vartheta_{\mu,t} \\
&\alpha_{1,t} = a_{11} + a_{12} \alpha_{1,t-1} + \sum_{i,j} a_{1j} d_i + \vartheta_{\alpha 1,t} \\
&\alpha_{2,t} = a_{21} + a_{22} \alpha_{2,t-1} + \vartheta_{\alpha 2,t} \\
&\alpha_{3,t} = a_{31} + a_{32} \alpha_{3,t-1} + \vartheta_{\alpha 3,t} \\
&\alpha_{4,t} = a_{41} + a_{42} \alpha_{4,t-1} + \vartheta_{\alpha 4,t}
\end{aligned} \qquad \textbf{(8)}$$

[12] In order to ensure a positive variance term, ε_t and $\vartheta_{\alpha 1,t}$ were defined as var[exp(ε_t)] and var[exp($\vartheta_{\alpha 1,t}$)], respectively. The choice for naming the variables through the text just as ε_t and $\vartheta_{\alpha 1,t}$ was adopted for simplification .

The final model, however, changed according to the inflation index used, as shown in the following sections.

IV.3.1 – IPCA

For IPCA, the model to be estimated was[13]:

$$p = \mu_t + \alpha_{1,t} e_{t-1} + \alpha_2 gap_{t-1} + \alpha_3 ope_{t-1} + \alpha_4 pm_{t-1} + \varepsilon_t$$

$$\mu_t = c_1 + c_2 \mu_{t-1} + \vartheta_{\mu,t}$$

$$\alpha_{1,t} = a_{11} + a_{12}\alpha_{1,t-1} + a_{13} d_2 + \vartheta_{\alpha 1,t}$$

In Table 6 - which shows the results obtained – we note that the coefficient of *gap* ($\alpha_{2,t}$) is significant and has the expected sign. The coefficient of the degree of openness ($\alpha_{3,t}$), although it contains the expected sign, is not significant. The coefficient $\alpha_{4,t}$, which refers to import prices, does not contain the expected positive sign, and is not significant.

Table 6 –Kalman Filter : IPCA Results

Variable	coefficient	Standard error	t-statistics	p-value
		Measure Equation		
α_2	1.636715	0.731031	2.238912	0.0252
α_3	-0.044387	0.089447	-0.496233	0.6197
$\alpha_{4,t}$	-0.001186	0.000802	-1.478920	0.1392
		State Equation - Intercept		
c_1	0.007883	0.013579	0.580584	0.5615
c_2	0.948579	0.049568	19.13699	0.0000
$\vartheta_{\mu,t}$	-5.954234	0.392567	-15.16744	0.0000
		State Equation – *Pass-through coefficient*		
a_{11}	0.489162	0.148228	3.300074	0.0010
a_{12}	0.005248	0.188764	0.027799	0.9778
$a_{1,3}$	-0.488220	0.242746	-2.011240	0.0443
$\vartheta_{\alpha 1,t}$	-2.208140	0.262412	-8.414797	0.0000
	Final State	Root MSE	z-statistics	p-value
$\mu_{T+1\|T}$	0.046336	0.065347	0.709071	0.4783
$\alpha_{1,T+1\|T}$	0.001166	0.331521	0.003517	0.9972
Log-likelihood	66.15722	Akaike Information Criteria		-1.225716
Hannan-Quinn Criteria	-1.102508	Schwartz Information Criteria		-0.920184

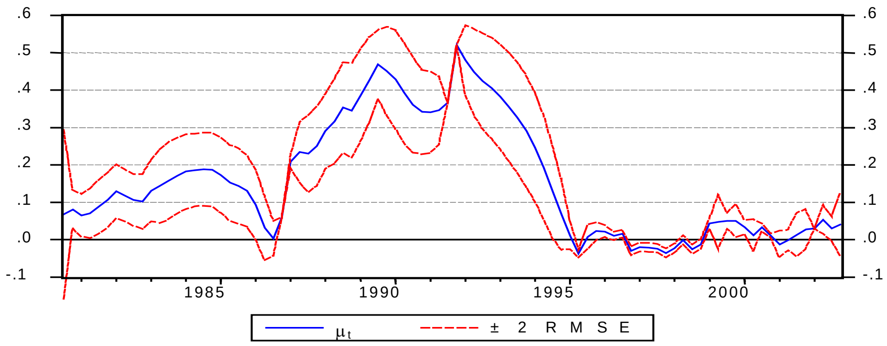
As far as variances are concerned, we note that the variances of the state equation for the intercept and the exchange rates ($\vartheta_{\mu,t}$ $\vartheta_{\alpha 1,t}$) have significant coefficients, which means that they are

[13] Appendix I shows some other models tested for IPCA, IGP-DI and IPA using the Kalman Filter.

effective. In other words, the coefficients are actually varying over time and, hence, the Kalman Filter captures changes in these coefficients that a model with constant parameters would not.

As for the intercept, we observe that c_1 is not significant whereas c_2 is. This means that, although the mean of the intercept is null, shocks are persistent on it. Such a result is expected in an inflation model if we consider that this variable captures the inflationary inertia of the period, since the Kalman filter with varying parameters on the constant is equivalent to estimating the stochastic trend of the series. So, we can consider that the constant, to some extent, represents the inflationary inertia. Graph 1 shows that this coefficient becomes not only smaller but also more stable after the implementation of the Real Plan, underscoring the idea of a remarkably low inflationary inertia after price stabilization.

Graph 1 – IPCA: Smoothed Estimate of μ_t

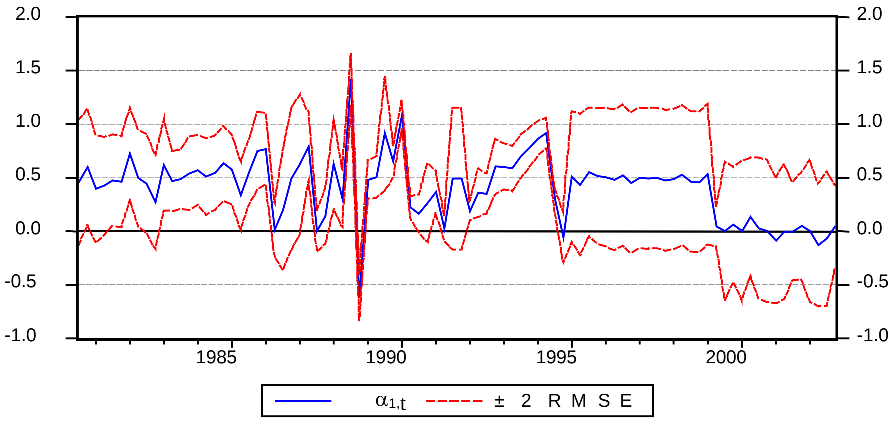


Concerning the behavior of the exchange rate coefficient, $\alpha_{1,t}$, a_{11} is significant, while a_{12} is not. In the former case, the exchange rate pass-through is important in explaining prices, regardless of the period. However, shocks on this coefficient do not propagate through time. In other words, the stochastic process resembles a white noise. The forecast for period t+1 is given by $E(\alpha_{1,t+1}) = a_{11} + a_{12}\alpha_{1,t-1} + a_{13}d_2$. Since a_{12} and a_{13} are indifferent from zero, the best forecast of the value for the exchange rate *pass-through* ($\alpha_{1,t+1}$) is the mean of the process, a_{11}. Hence, an increase of 1% in the exchange rate causes an average increase, in the period analyzed, of 0.49% in the inflation rate. Finally, the dummy variable d_2, is significant, implying that the intervention in the exchange rate market affects the pass-through dynamics.

By analyzing Graph 2, which shows the smoothed estimates of the pass-through coefficient, we can clearly identify three different periods in the behavior of α_1. These periods may be associated with three different moments of the Brazilian economy throughout the sample period.

The first period goes from 1980 until the implementation of the Real plan. A considerably high and volatile exchange rate pass–through characterizes this period, with peaks close to one, which illustrates the exchange rate/price spiral typical of high inflations. The mean pass-through for the period is 0.49 (see Appendix II), but there are moments of sharp reductions that may be associated with the different economic plans (1986:II, 1987:III, 1988:IV, 1990:II, 1991:II, 1992:I).

Graph 2 – IPCA: Smoothed Estimate of $\alpha_{1,t}$



The second period covers 1995 to 1998, where the mean drops to 0.42 (Appendix II), and so does its volatility, showing a more stable behavior over time. Finally, the third period starts in 1999, when the floating exchange rates were adopted and the coefficient remarkably decreased, yielding a mean value of around 0.4.

At first, the strong decrease in 1999 is not expected since, in large devaluations, a higher pass-through from exchange rates to prices is assumed. However, the Brazilian economic scenario at the time, with recession and extremely volatile exchange rates, may have favored a contrary behavior. In this scenario, price setters would not be able to increase their prices proportionately to the devaluation as they used to do before, due to the economic slowdown, which inhibits demand, and to the uncertainty about the future. If the exchange rates do not mantain that higher level, the costs to reverse the price increase (menu costs and reputation costs, for instance) could be much higher[14]. Furthermore, in times of pegged exchange rates, changes in exchange rates are considered to be permanent and, therefore, agents have an extra incentive to adjust their prices as soon as possible. However, in times of floating exchange rates, the resulting uncertainty and the presence of factors such as menu costs and hysteresis (see DIXIT, 1986), make agents "wait and see" until they can be sure that the (de)valuation is permanent and until they know the new exchange rate level.

IV.3.2 – IGP-DI

The IGP-DI model is quite similar to that used for IPCA, with time-varying coefficients for the exchange rate and intercept. Although d_2 was not significant, its inclusion yielded better results than the inclusions of other dummy variables (also nonsignificant) or than its absence. The three periods related to the behavior of the exchange rate coefficient are more noticeable than in the case of IPCA, and the decrease in 1999 is less intense as well, (Table 7 and graphs 3 and 4).

[14] For a detailed discussion, see DIXIT (1986).

Table 7 Kalman Filter : IGP-DI Results

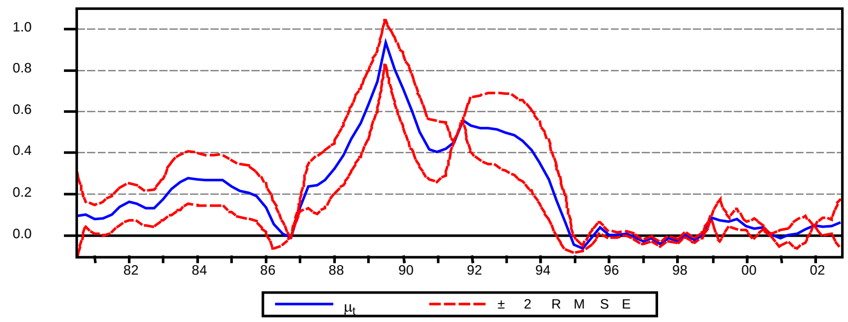
Variable	Coefficient	Standard error	t-statistics	p-value
		Measurement Equation		
α_2	1.9712	1.0770	1.8303	0.0672
α_3	0.0389	0.1638	0.2377	0.8121
α_4	0.0003	0.0015	0.1715	0.8638
		State Equation - Intercept		
c_1	-5.1799	0.3117	-16.6200	0.0000
c_2	0.0113	0.0200	0.5671	0.5707
$\vartheta_{\mu,t}$	0.9457	0.0414	22.8407	0.0000
		State Equation - Pass-through Coefficient		
a_{11}	0.3246	0.1807	1.79645	0.0724
a_{12}	-0.0075	0.1871	-0.0401	0.9680
$a_{1,3}$	-0.2702	0.2547	-1.0611	0.2887
$\vartheta_{\alpha 1,t}$	-2.0026	0.3888	-5.1514	0.0000
	Final State	Root MSE	z-statistics	p-value
$\mu_{T+1\|T}$	0.0690	0.0918	0.75120	0.4525
$\alpha_{1,T+1\|T}$	0.0536	0.3674	0.1458	0.8841

Log-likelihood	44.0753	Akaike Information Criteria	-0.7350
Hannan-Quinn Criteria	-0.6118	Schwartz Information Criteria	-0.4295

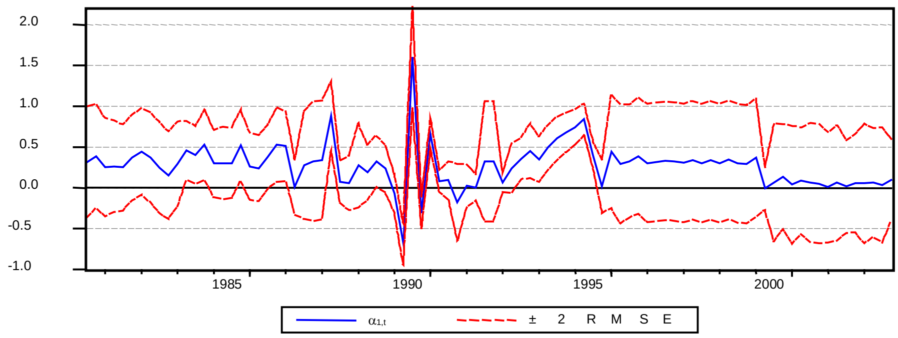
As for the significance of fixed parameters and their signs, import prices are still indifferent from zero. α_2 and α_3 are significant, although the former one does not present the expected sign.

With regard to the exchange rate coefficient, again, it has a white noise with drift. We may also note that most of the sharp reductions in the coefficients are the same ones found for IPCA (1986:II, 1987:III, 1989:I, 1989:IV, 1990:II to 1990:4, 1991:I, 1991:II, 1992:I, 1994:I). The coefficient – or the exchange-rate elasticity of prices - is 0.3246 for the whole period. Calculating the mean of the filtered estimates of $\alpha_{1,t}$ for the three periods, we have an elasticity of 0.33 for 1980:I to 1994:II – if we remove the above mentioned periods from the sample, this value goes to 0.40 – 0.27 from 1994:III to 1998:IV and 0.07 from 1999:I on (see Appendix II). Thus, the Real plan led to a decrease in the pass-through, but the change in the exchange rate system and the adoption of the inflation targeting regime in 1999 caused a sharper decrease in this coefficient.

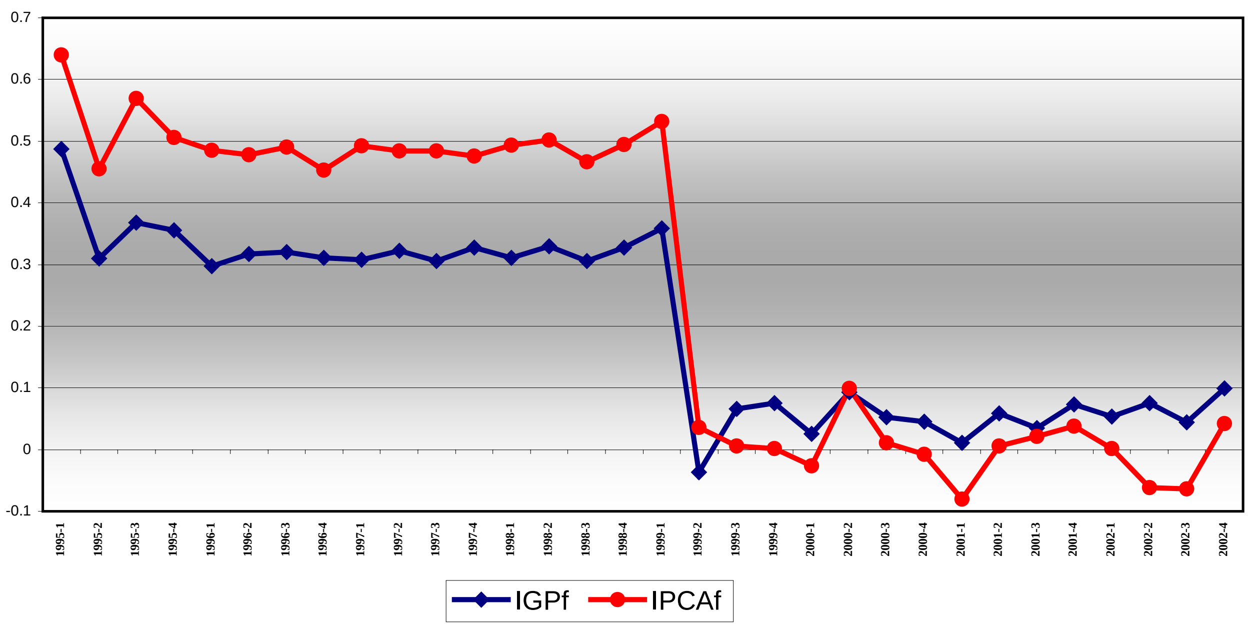
Graph 3 – IGP-DI: Smoothed Estimate of $\mu_{,t}$



Graph 4 – IGP-DI: Smoothed Estimate of $\alpha_{1,t}$



Graph 5 – Filtered Estimates of $\alpha_{1,t}$ for the Post-Real period: IGP-DI *vs.* IPCA



Graph 5 draws some attention to the comparison of both indexes through the filtered estimates of the exchange rate coefficients in both cases (IPCA and IGP-DI) after the Real plan. Until 1999, the exchange rate pass-through to IPCA was, on average, higher than to the IGP-DI, which justifies the selection of the latter one as the index used to realign contracts. After 1999, we have an opposite situation, when the IGP-DI – which consists mostly of wholesale prices - had a pass-through higher, on average, than that of IPCA.

IV.3.3 –IPA

The best model that the Kalman filter identified for the wholesale price index (IPA) had only the exchange-rate coefficient as time-varying. The model assumes the following form:

$$p = \mu_t + \alpha_{1,t} e_{t-1} + \alpha_{2,t} gap_{t-1} + \alpha_{3,t} ope_{t-1} + \alpha_{4,t} pm_{t-1} + \varepsilon_t$$

$$\alpha_{1,t} = a_{11} + a_{12}\alpha_{1,t-1} + \vartheta_{\alpha 1,t}$$

In case of IPA, the most satisfactory results were those where only the exchange rate coefficient varies over time. Again, $\alpha_{1,t}$ follows a white noise with drift, as shown in Table 8. This means that the best forecast for the pass-through from exchange rates to IPA is the mean of the process. We can also note, according to Graph 6, that the only change in the behavior of the coefficient is the peaks in the early 1990s and a smaller volatility after the Real plan. However, the mean of the coefficient was relatively stable: 0.90 from 1980:I to 1994:II, 0.86 from 1994:3 to 98:4 and 0.87 from 1999:I on (see Appendix II). If we exclude the moments in which there was a sharp decrease in the filtered coefficient (basically the same as with IPCA and IGP-DI: 1986:II, 1987:III, 1989:II, 1990:II, 1991:I, 1991:II, 1992:I, 1994:III, 1994:IV, 1999:II) – the mean goes to 0.93 between 1980:I and 1994:II, 0.89 for 1994:II/1998:IV and 0.88 after 1999.

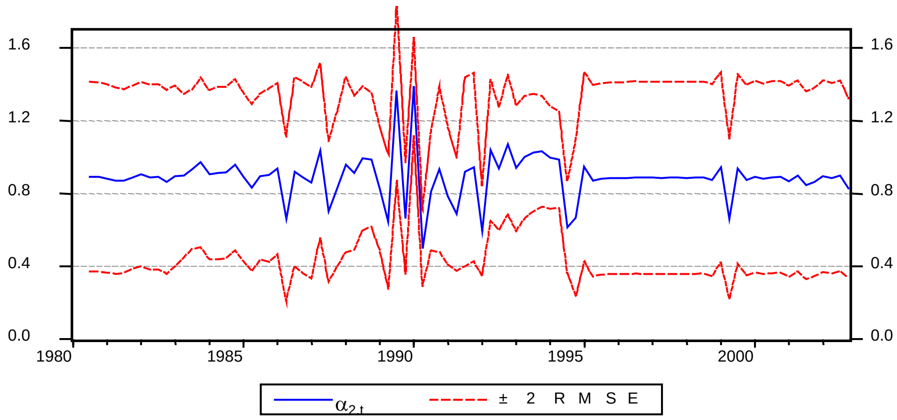
The values presented here are quite high compared to the ones found for the other two indices. As they are close to the unit, they also suggest a virtually complete pass-through from exchange rate to wholesale prices. This result seems to demonstrate what the economic theory had already predicted: wholesale prices are more strongly affected by exchange rate movements and, in the absence of nontradeables, the pass-through from exchange rate to inflation is almost complete. An explanation to this behavior can be found in the theoretical model developed in this paper. Since Brazil is a small economy, it is a price-taker in the foreign market and it is not able to affect international prices as predicted by the pricing-to-market models applied to developed economies15. A smaller pass-through to consumer prices (IPCA and IGP) reflects the absorption of the exchange rate by retailers, which can be explained – given the maximization model presented in this paper – by an attempt to avoid a reduction in demand that is not offset by a rise in prices. This does not necessarily mean losses for the agents, but only a change in their profit margins.

[15] Those models consider that, in face of an exchange rate devaluation in country B, the exporting firm in country A will reduce its exporting prices for B in order not to have a high reduction in sales in that country, given the weight of that market on its global demand. The result of such an action is that prices in B will rise less than proportionally to the exchange rate devaluation, resulting in an incomplete pass-through and evidence of the rejection of PPP.

Table 8- Kalman Filter : IPA Results

Variable	Coefficient	Standard error	t-statistics	p-value
		Measure equation		
μ_t	0.0471	0.0491	0.9590	0.3376
α_2	1.9392	1.031	1.8809	0.0600
α_3	-0.0937	0.2224	-0.4212	0.6736
α_4	0.0019	0.0048	0.4032	0.6868
		State Equation - Pass-through Coefficient		
a_{11}	1.0908	0.3694	2.9526	0.0032
a_{12}	-0.2263	0.3667	-0.6171	0.5372
$\vartheta_{\alpha1,t}$	-2.7168	0.9236	-2.9417	0.0033
	Final State	Root MSE	z-statistics	p-value
$\alpha_{1,\ T+1\|T}$	0.9029	0.2631	3.4323	0.0006
Log-likelihood	26.7479	Akaike Information Criteria		-0.4166
Hannan-Quinn Criteria	-0.3270	Schwartz Information Criteria		-0.1944

Graph 6 – IPA: Smoothed estimate of $\alpha_{2,t}$



V – Comparison of Results

According to the results obtained herein, there is a decrease in the pass-through from the exchange rate to IPCA and IGP-DI after the stabilization of the Real, and a sharper one after the shift in the exchange rate regime in 1999. Before 1999, the effects of an exchange rate shock in period *t* on IGP-DI would be complete after approximately four quarters (considering the absence of further shocks). After 1999, only 32% of the shock would have been absorbed by the index in an equal period. For IPCA, between 1980 and 1998, the pass-through of the shock would be complete in two quarters before 1999, but after that year it would represent only 7%. The exception is IPA-DI, which keeps an almost complete pass-through in the third quarter.

The exchange rate pass-through behavior found in the present paper is in line with other estimates reported in the literature. MINELLA, FREITAS, GOLDFAJN and MUINHOS (2003) analyzed the post-Real period and also found a change in the exchange rate pass-through to prices after 1999 (considering the 12-month exchange rate variation with one lag). However, the magnitude of the change is different, depending on the approach adopted: the Central Bank's structural model, the Phillips curve, or a VAR model. Nonetheless, the graph of the recursive estimation of the coefficients found in the Phillips curve for IPCA is very similar to Graph 2 presented in section IV.3.1 of this paper.

The results for IPCA in the post-Real period are also similar to the ones presented by MUINHOS and ALVES (2003) for free prices – which correspond to approximately 70% of IPCA – by applying a non-linear Phillips curve. The authors found an exchange rate pass-through of 0.51 between 1995:I and 1998:IV and of 0.06 from 1999:I on. The values observed for the period after 1999 are also similar to the ones presented by CARNEIRO, MONTEIRO and WU (2002), who found a quarterly exchange rate pass-through between 1999 and 2002 of 6.4% on average.

Our results for IPCA are also close to the ones presented by BELAISCH (2003), who observed a 6% exchange rate pass-through in Brazil in a 3-month period through a VAR model. However, the results are considerably different for the other indices (27% for the IGP and 34% for the IPA). Despite this difference, the relation between the magnitudes of the coefficients is the same: the pass-through to IPA is larger than to IGP, which is larger than to IPCA.

GODLJAN and WERLANG (2000) found a six-month accumulated pass-through of about 21% for European economies and 38% for emerging ones between 1980 and 1998. HAUSSMAN, PANIZZA and STEIN (1999) also encountered different pass-through values among the analyzed countries. For instance, the USA, the UK and Japan have an average 12-month accumulated pass-through of 3%; Germany, Canada and Norway, 7%; Switzerland, Greece, Israel and Korea, 16%; Australia and Peru, 21%, and Mexico, Paraguay and Poland, over 50%, among others.

MUINHOS (2001) uses a sample with quarterly data from 1980 to 2000, different estimations of the Phillips curve, with and without an expectation term and with linear and non-linear specifications (the latter of which contains cross-terms), also including a short sample relating to the period after 1995. The results of the linear specification point towards a pass-through coefficient of 0.10 in the small sample if the expectation term is not included and of 0.09 if this term is included. In the non-linear specification, the pass-through coefficients are 0.24 without the expectation term, 0.12 with the term, and 0.55 for the whole sample. However, the results do not indicate changes in the exchange rate pass-through after 1995 in the whole sample, or in both samples, after 1999. When the authors show the behavior of pass-through coefficients after 1998 for the small sample, there is a break in this coefficient after the floating exchange rate regime was adopted. The average coefficient for 1998 is, in this case, higher than 0.5 while it is of about 0.1 after 1999, a result that is in line with the ones presented in this paper. Nevertheless, such change is not identified in the whole sample (1980 to 2000) when the coefficient for 1998 is also around 0.1.

VI – Final Remarks

Some conclusions may be drawn in light of what was discussed in this paper. The first conclusion is that the model developed from a pricing-to-market approach is able to indicate changes that occurred in the exchange rate coefficient throughout the study period. Furthermore, amongst the tested formulations, non-linear models and time-variant coefficients are more suitable than OLS coefficients with time-invariant parameters, even when the sample is divided. The results obtained by MUINHOS (2001) who, as other previously mentioned authors, observed a decrease in the exchange rate pass-through after 1999 only when using the small sample, lends further support to the Kalman Filter to the detriment of time-invariant models, when such a long and complex period is analyzed.

In comparison with the results found by GOLDFJAN and WERLANG (2000) and by HAUSSMAN, PANIZZA and STEIN (1999), the data presented herein show that the pass-through from the exchange rate to prices in Brazil is not only smaller after the adoption of the floating exchange rate system in 1999, but also quite similar to the ones observed in stronger economies. Our paper shows that the macroeconomic environment affects the way prices will respond to exchange rate movements, as it is possible to identify three different patterns in the pass-through coefficient: the first one is characterized by a high inflation period, the second one concerns the period of low inflation and pegged exchange rates, and the third one refers to the period of stable prices and floating exchange rates. The presence of the dummy variable d_2 also suggests that the type of exchange rate regime observed by the agents – more than the one officially announced – also affects the response of prices to exchange rate movements.

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APPENDIX I – LINEAR MODEL RESULTS

A.I – Linear Models – Complete Sample (TABLES)

Table A.1 - IPCA

Coefficient	Estimate	Standard Error	t-statistics	p-value
μ	0.0508	0.0286	1.7784	0.0789
α_1	0.8398	0.0684	12.2763	0.0000
α_2	2.1079	0.7436	2.8349	0.0057
α_3	-0.3304	0.1884	-1.7534	0.0831
α_4	-0.0007	0.0037	-0.1965	0.8447

R^2	0.6433	Mean dependent var	0.2891
R^2 adjusted	0.6265	S.D. dependent var	0.3205
S.E. of regression	3.2603	AIC	-0.3690
LM test (1st order)	0.0065†	SIC	-0.2301
ARCH-LM test (1st order)	17.6433*	F-statistic	38.3271
Ramsey-Reset test (2nd order)	2.9960*	Prob(F-statistic)	0.0000

* significant at 5%;† for higher orders the presence of residual autocorrelation was also rejected

Table A.2 – IGP-DI

Coefficient	Estimate	Standard Error	t-statistics	p-value
μ	0.0644	0.0290	2.2220	0.0289
α_1	0.8083	0.0695	11.6346	0.0000
α_2	1.9492	0.7552	2.5812	0.0116
α_3	-0.1666	0.1914	-0.8703	0.3866
α_4	0.0008	0.0038	0.1975	0.8439

R^2	0.6152	Mean dependent var	0.2955
R^2 adjusted	0.5971	S.D. dependent var	0.3134
S.E. of regression	0.1989	AIC	-0.3380
LM test (1st order)	2.5224**	SIC	-0.1991
ARCH-LM test (1st order)	0.5579†	F-statistic	33.9716
Ramsey-Reset test (1st order)	3.4552*	Prob(F-statistic)	0.0000

* significant at 5%; ** significant at 10%;† for higher orders the presence of residual autocorrelation was also rejected

Table A.3 – IPA

Coefficient	Estimate	Standard Error	t-statistics	p-value
μ	0.0636	0.0292	2.1778	0.0322
α_1	0.8119	0.0699	11.6112	0.0000
α_2	1.9018	0.7600	2.5024	0.0143
α_3	-0.1830	0.1926	-0.9504	0.3446
α_4	0.0008	0.0038	0.2105	0.8338

R^2	0.6145	Mean dependent var	0.2955
R^2 adjusted	0.5964	S.D. dependent var	0.3151
S.E. of regression	0.2002	AIC	-0.3252
LM test (2st order)	3.2633*	SIC	-0.1863
ARCH-LM test (2nd order)	5.6300*	F-statistic	33.8757
Ramsey-Reset test (1st order)	3.5620*	Prob(F-statistic)	0.0000

* significant at 5%

A.II – Alternative Linear Models (TABLES)

Table A.4 – IPCA

Coefficient	Pre-Real	Post-Real	Model with dummy variables
μ	0.1415*	0.0173*	0.0554
	(0.0516)	(0.0032)	(0.0269)**
α_1	0.7262*	0.0568	0.8888
	(0.1077)	(0.0355)	(0.0655)*
α_{14}	-	-	-0.5680*
			(0.1669)
α_{15}	-	-	-0.8551**
			(0.3792)
α_2	2.4121*	0.4505*	2.17881*
	(0.9129)	(0.1683)	(0.0022)
α_3	-0.1325	-0.0125	-0.1591
	(0.2798)	(0.0265)	(0.1824)
α_4	-0.0012	0.0005	-0.0008
	(0.0054)	(0.0005)	(0.0035)
R^2	0.4888	0.2304	0.7006
R^2 adjusted	0.4471	0.1164	0.6789
S.E. of regression	2.3945	0.0147	0.1816
LM test (1st order)	0.0879†	13.5814*	0.7835†
ARCH-LM test (1st order)	7.8165*	6.5385*	23.0965*
Ramsey-Reset test (1st order)	2.9151***(a)	0.2304**	

*significant at 1%; **significant at 5%; ***significant at 10%; (a) test in second order; † for higher orders, residual autocorrelation was also rejected

Table A.5 – IGP-DI

Coefficient	Pre-Real	Post-Real	Model with dummy variables
μ	0.1808*	0.0222*	0.0707*
	(0.0526)	(0.0042)	(0.0282)
α_1	0.6341*	0.0876**	0.8428
	(0.1097)	(0.04551)	(0.0686)*
α_{14}	-	-	-0.4192*
			(0.1748)
α_{15}	-	-	-0.8658**
			(0.3971)
α_2	2.2269*	0.0186	1.9876
	(0.9301)	(0.21571)	(0.7234)
α_3	0.0843	0.0351	-0.0184
	(0.2851)	(0.0340)	0.0394)
α_4	0.0006	0.0010	-0.0003
	(0.0055)	(0.006)	(0.0008)
R^2	0.4210	0.2655	0.6564
R^2 adjusted	0.3737	0.1567	0.6316
S.E. of regression	0.2252	0.0189	0.1902
LM test (1st order)	4.2357(a) **	0.3868†	1.3723 **
ARCH-LM test (1st order)	0.0830†	0.6568†	1.4793**
Ramsey-Reset test (1st order)	2.3705(a)***	3.6850**(a)	0.6564

*significant at 1%; **significant at 5%; ***significant at 10%; (a) test in second order; † for higher orders, residual autocorrelation was also rejected

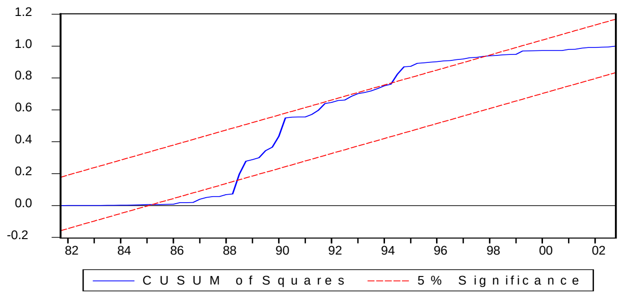
Table A.6– IPA-DI

Coefficient	Pre-Real	Post-Real	Model with dummy variables
μ	0.1721*	0.0252*	0.0688*
	(0.0536)	(0.0055)	(0.0284)
α_1	0.6562*	0.1023***	0.8483*
	(0.1112)	(0.0617)	(0.0692)
α_{14}	-	-	-0.4344*
			(0.1764)
α_{15}	-	-	-0.7943**
			(0.4007)
α_2	2.2346*	-0.3548	1.9477*
	(0.9468)	(0.2924)	(0.7299)
α_3	0.0242	0.0599	-0.0342
	(0.2902)	(0.0461)	(0.1928)
α_4	0.0006	0.0011	0.0007
	(0.0056)	(0.0009)	(0.0037)
R^2	0.4256	0.324	0.6541
R^2 adjusted	0.3787	0.2239	0.6291
S.E. of regression	0.2293	0.0256	0.1919
LM test (1st order)	3.9428**(a)	0.0117†	2.4383†
ARCH-LM test (1st order)	4.6245**(a)	3.4787***(a)	4.4998**
Ramsey-Reset test (1st order)	2.8311***	6.4948**(a)	

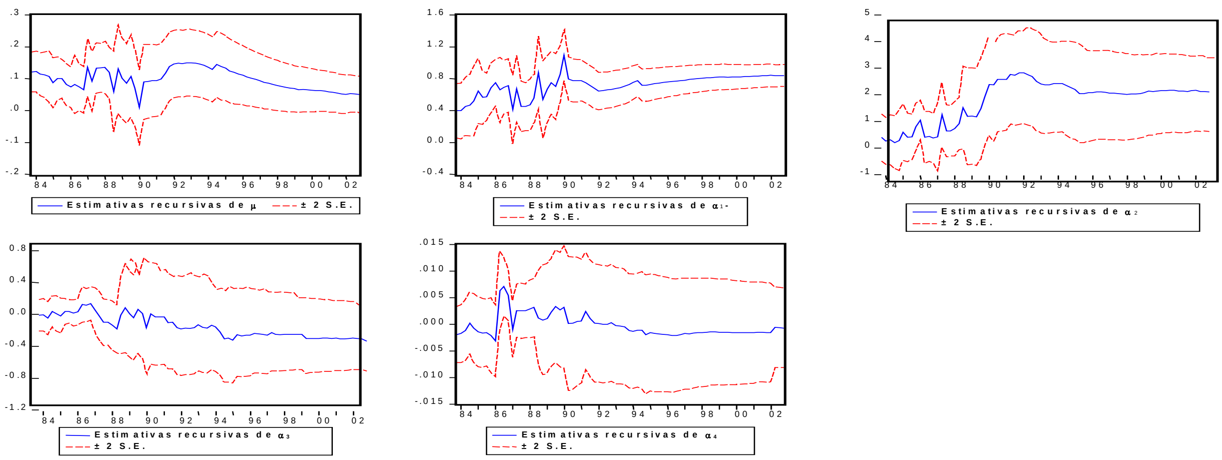
*significant at 1%; **significant at 5%; ***significant at 10%; (a) test in second order; † for higher orders, residual autocorrelation was also rejected

A.I – Linear Models (GRAPHS)

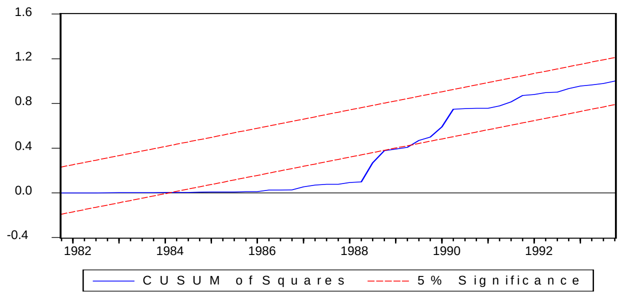
Graph A.1– IPCA – CUSUM of Squares test (complete period)



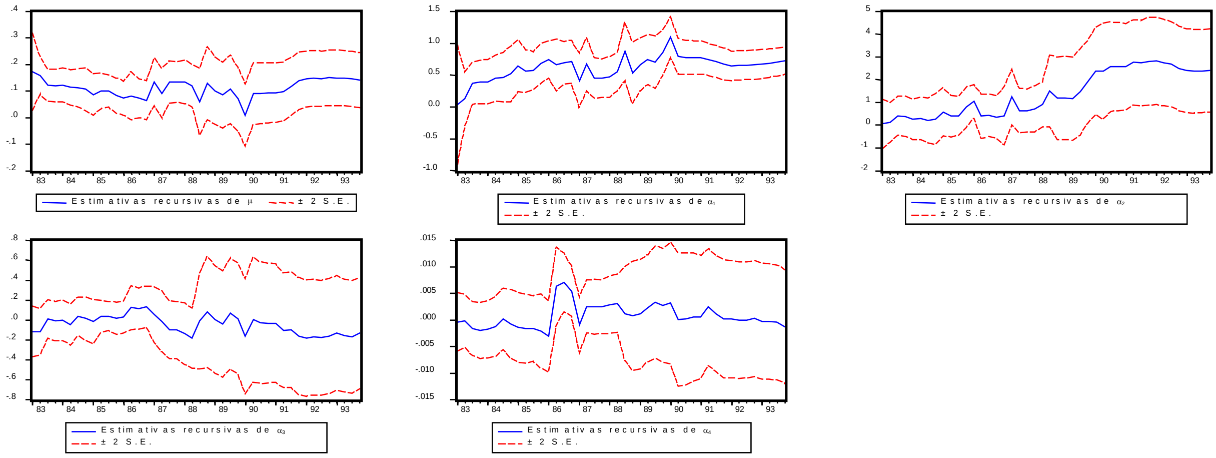
Graph A.2 – IPCA – Recursive Coefficients (complete period)



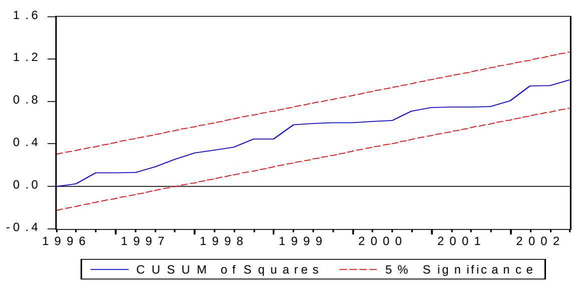
Graph A.3 – IPCA – CUSUM of Squares test (pre-Real period)



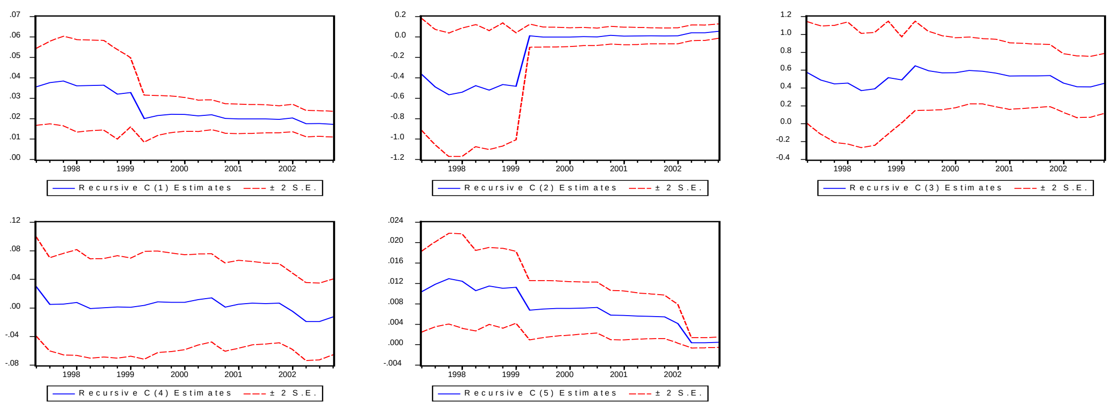
Graph A.4 – IPCA - Recursive Coefficients (pre-Real period)



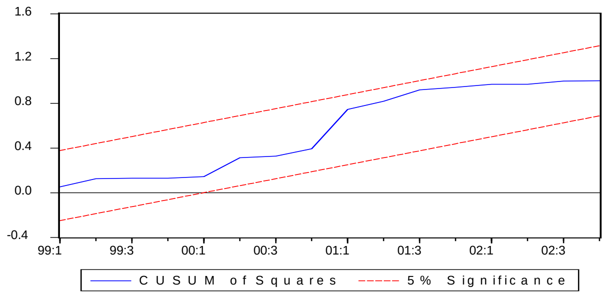
Graph A.5 – IPCA - CUSUM of Squares Test (post-Real period)



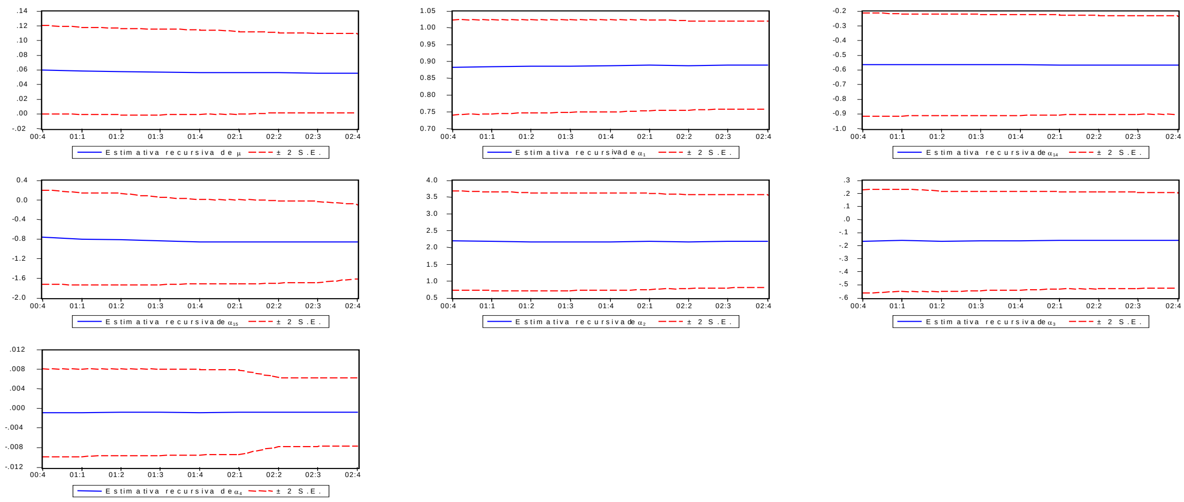
Graph A.6 – IPCA - Recursive Coefficients (post-Real period)



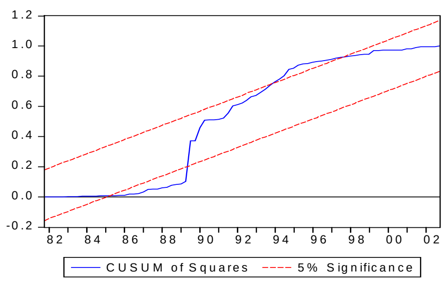
Graph A.7 – IPCA - CUSUM of SquaresTest - model with dummy variables



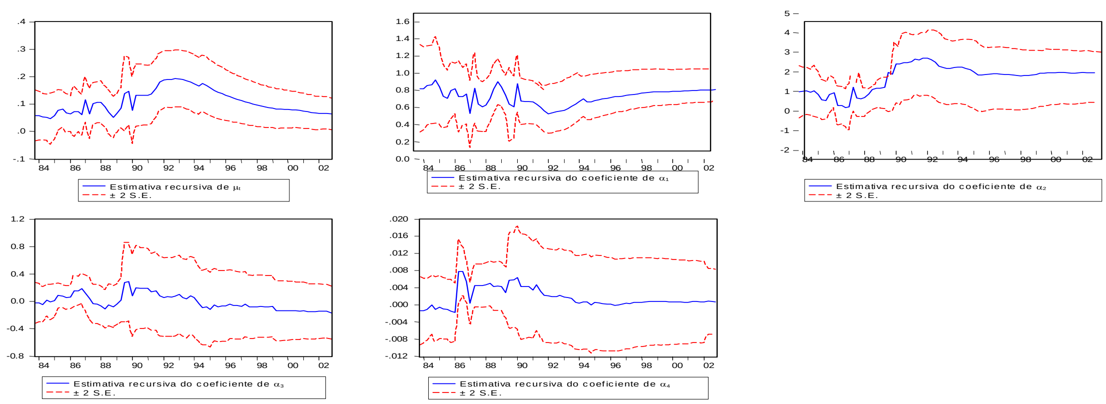
Graph A.8 – IPCA - Recursive Coefficients - model with dummy variables



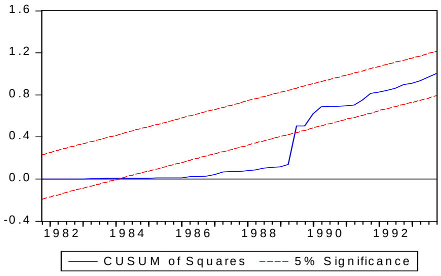
Graph A.9 – IGP - CUSUM of Squares test (complete period)



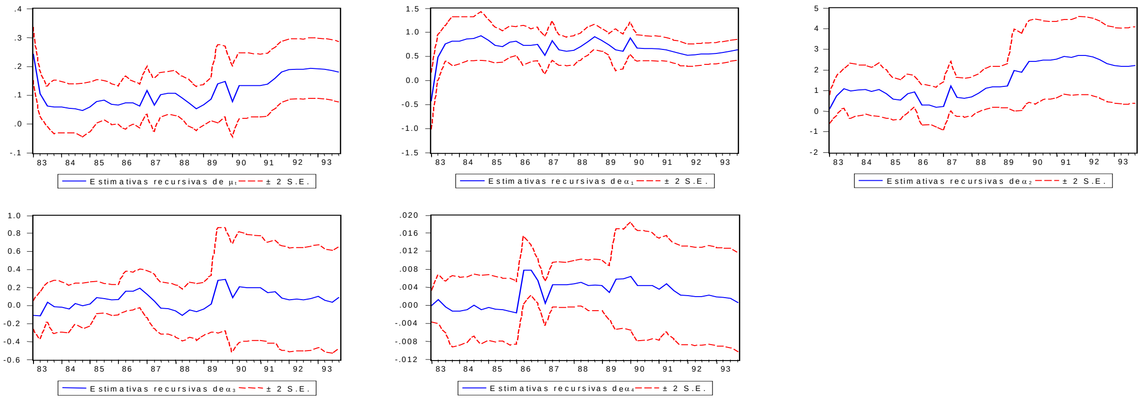
Graph A.10 – IGP - Recursive Coefficients (complete period)



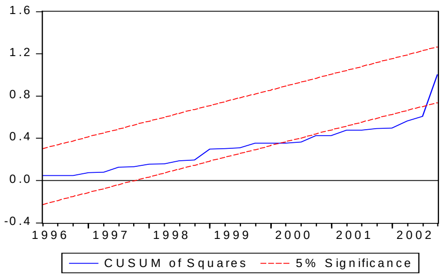
Graph A.11 – IGP - CUSUM of Squares test (pre-Real period)



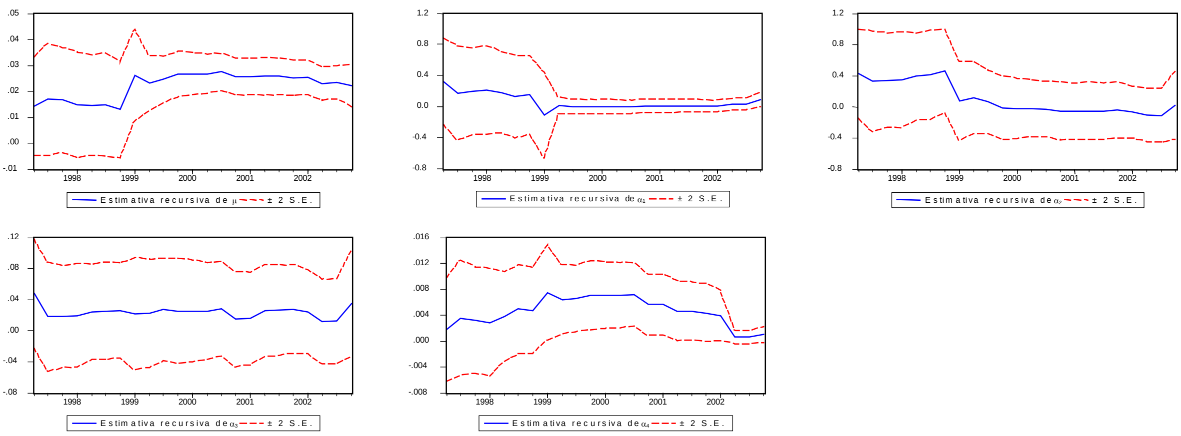
Graph A.12 – IGP - Recursive Coefficients (pre-Real period)



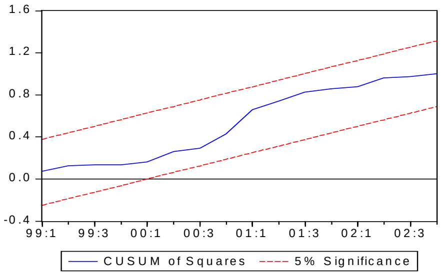
Graph A.13 – IGP - CUSUM of Squares Test (post-Real period)



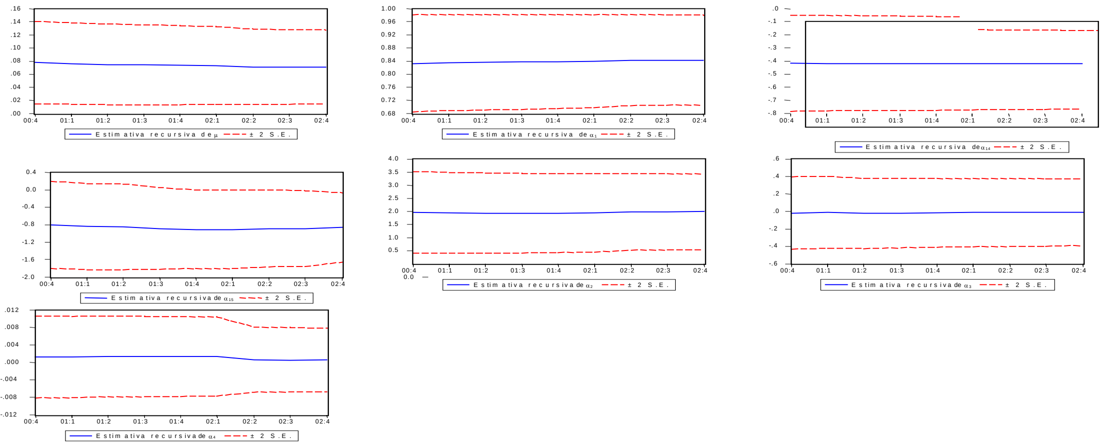
Graph A.14 – IGP Recursive Coefficients (post-Real period)



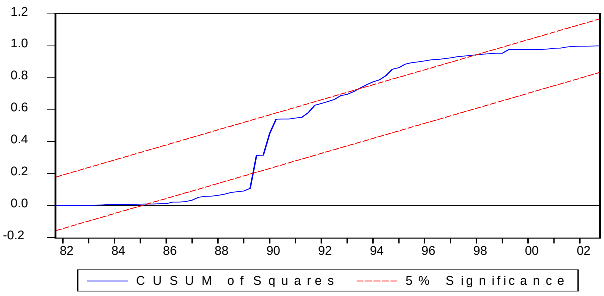
Graph A.15 – IGP - CUSUM of SquaresTest - model with dummy variables



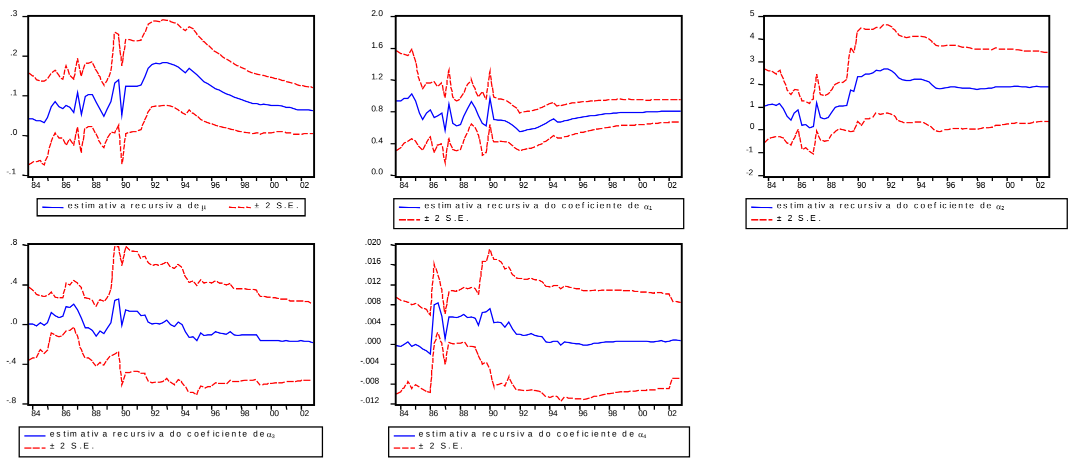
Graph A.16 – IGP - Recursive Coefficients - model with dummy variables



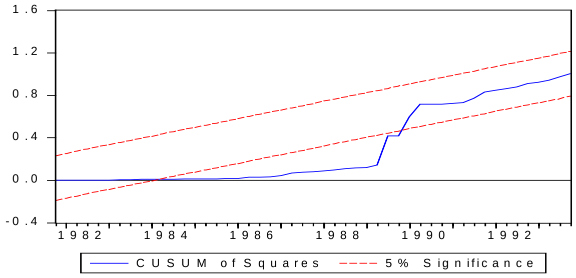
Graph A.17 – IPA - CUSUM of Squares test (complete period)



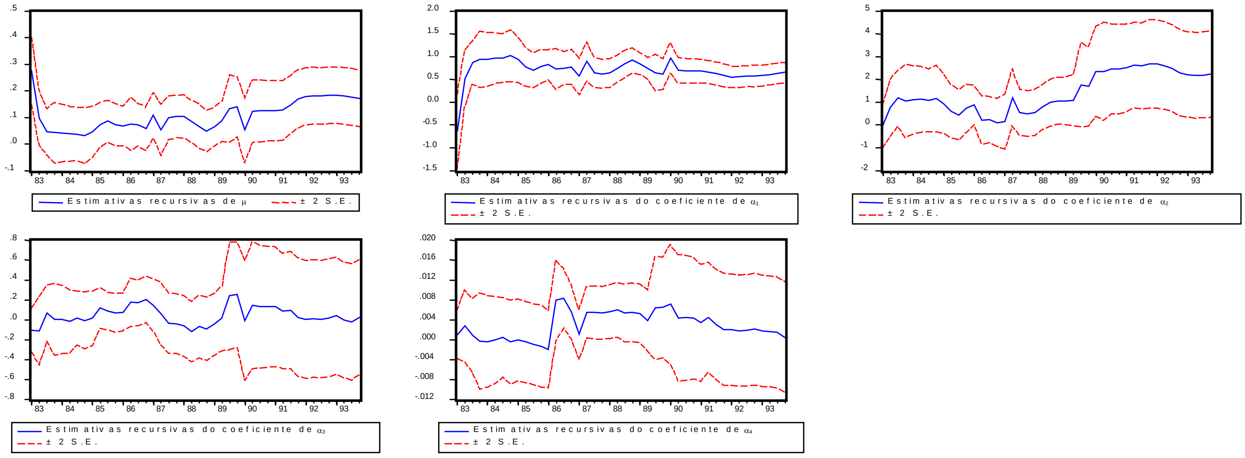
Graph A.18 – IPA - Recursive Coefficients (complete period)



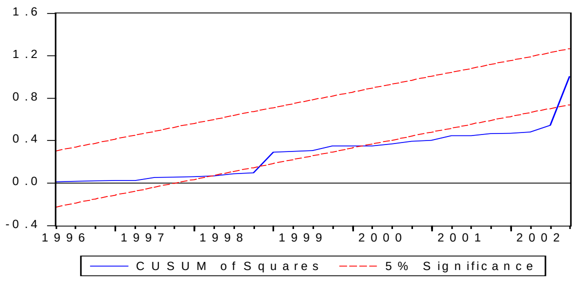
Graph A.19 – IPA - CUSUM of Squares test (pre-Real Period)



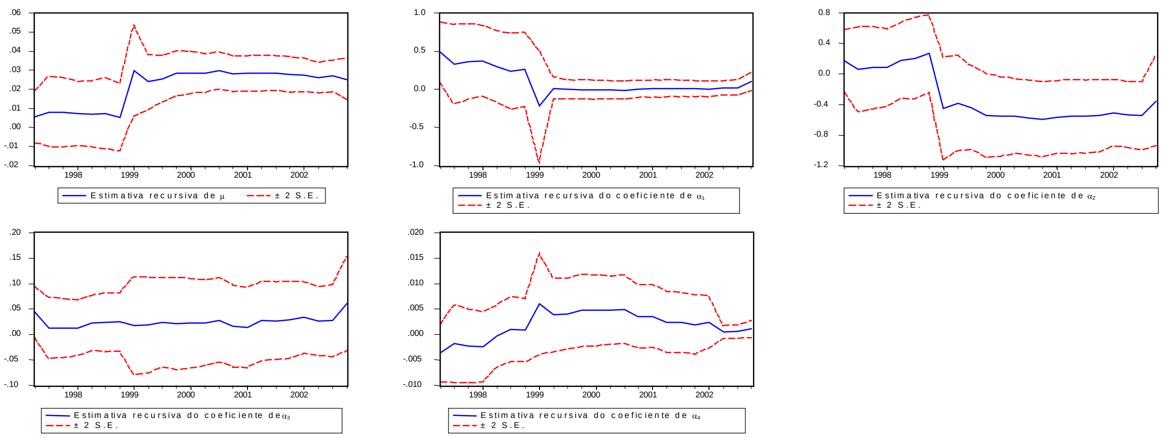
Graph A.20 – IPA - Recursive Coefficients (pre-Real period)



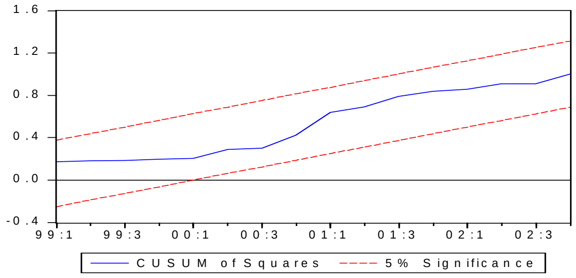
Graph A.22 – IPA - CUSUM of Squares Test (post-Real period)



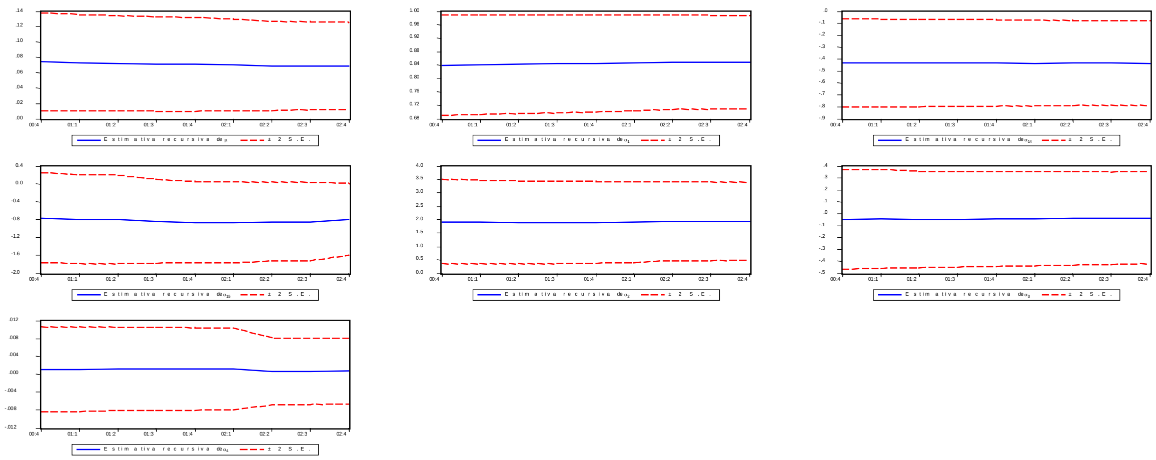
Graph A.23 - IPA –Recursive Coefficients (post-Real period)



Graph A.24 – IPA - CUSUM of SquaresTest - model with dummy variables



Graph A.25 – IPA - Recursive Coefficients - model with dummy variables



APPENDIX II – FILTERED COEFFICIENTS– VALUES AND MEANS

II.1 – Filtered Coefficients of $\alpha_{1,t}$ – IPCA

DATE	Coefficient Filtered Estimate	DATE	Coefficient Filtered Estimate	DATE	Coefficient Filtered Estimate
1980-1	0.0000	**1987-4**	0.3051	**1995-3**	0.5409
1980-2	0.4582	**1988-1**	0.7671	**1995-4**	0.4747
1980-3	0.4620	**1988-2**	0.4572	**1996-1**	0.4550
1980-4	0.5130	**1988-3**	1.4162	**1996-2**	0.4481
1981-1	0.3686	**1988-4**	-0.4444	**1996-3**	0.4612
1981-2	0.4391	**1989-1**	0.6641	**1996-4**	0.4225
1981-3	0.5050	**1989-2**	0.7343	**1997-1**	0.4624
1981-4	0.5096	**1989-3**	0.8732	**1997-2**	0.4543
1982-1	0.6443	**1989-4**	0.6297	**1997-3**	0.4541
1982-2	0.4166	**1990-1**	1.0293	**1997-4**	0.4454
1982-3	0.3926	**1990-2**	0.1772	**1998-1**	0.4643
1982-4	0.3397	**1990-3**	0.1075	**1998-2**	0.4717
1983-1	0.6340	**1990-4**	0.2846	**1998-3**	0.4367
1983-2	0.4925	**1991-1**	0.4021	**1998-4**	0.4646
1983-3	0.5250	**1991-2**	0.1042	**1999-1**	0.0819
1983-4	0.5661	**1991-3**	0.4583	**1999-2**	0.0197
1984-1	0.5688	**1991-4**	0.4620	**1999-3**	0.0378
1984-2	0.5024	**1992-1**	0.1729	**1999-4**	0.0218
1984-3	0.5192	**1992-2**	0.3231	**2000-1**	0.0034
1984-4	0.5487	**1992-3**	0.3244	**2000-2**	0.1217
1985-1	0.4631	**1992-4**	0.5412	**2000-3**	0.0393
1985-2	0.2699	**1993-1**	0.5287	**2000-4**	0.0208
1985-3	0.4645	**1993-2**	0.4926	**2001-1**	-0.0557
1985-4	0.5364	**1993-3**	0.5694	**2001-2**	0.0342
1986-1	0.4541	**1993-4**	0.6168	**2001-3**	0.0422
1986-2	-0.1555	**1994-1**	0.6633	**2001-4**	0.0649
1986-3	0.3079	**1994-2**	0.6673	**2002-1**	0.0301
1986-4	0.4595	**1994-3**	0.0581	**2002-2**	-0.0423
1987-1	0.6253	**1994-4**	-0.2927	**2002-3**	-0.0421
1987-2	0.7878	**1995-1**	0.5671	**2002-4**	0.0570
1987-3	0.0552	**1995-2**	0.4270		

Period	Mean
1980:1 / 1994:2	**0.4876**
1994:3 / 1998:4	**0.4213**
1999:1 / 2002:4	**0.0346**

II.2 – Filtered Coefficients of $\alpha_{1,t}$ – IGP

DATE	Coefficient Filtered Estimate	DATE	Coefficient Filtered Estimate	DATE	Coefficient Filtered Estimate
1980-1	0.0000	**1987-4**	0.2429	**1995-3**	0.3683
1980-2	0.3245	**1988-1**	0.5097	**1995-4**	0.3553
1980-3	0.3221	**1988-2**	0.4317	**1996-1**	0.2972
1980-4	0.3336	**1988-3**	0.6012	**1996-2**	0.3174
1981-1	0.2432	**1988-4**	0.5513	**1996-3**	0.3206
1981-2	0.2933	**1989-1**	0.3359	**1996-4**	0.3108
1981-3	0.3307	**1989-2**	0.0694	**1997-1**	0.3082
1981-4	0.4270	**1989-3**	1.3132	**1997-2**	0.3228
1982-1	0.4174	**1989-4**	-0.4109	**1997-3**	0.3058
1982-2	0.3100	**1990-1**	0.5020	**1997-4**	0.3279
1982-3	0.2353	**1990-2**	-0.0620	**1998-1**	0.3114
1982-4	0.2520	**1990-3**	-0.1438	**1998-2**	0.3298
1983-1	0.4193	**1990-4**	-0.1819	**1998-3**	0.3057
1983-2	0.5477	**1991-1**	0.0956	**1998-4**	0.3273
1983-3	0.4818	**1991-2**	0.1157	**1999-1**	0.3591
1983-4	0.5161	**1991-3**	0.3245	**1999-2**	-0.0370
1984-1	0.2967	**1991-4**	0.3221	**1999-3**	0.0658
1984-2	0.3107	**1992-1**	0.0712	**1999-4**	0.0754
1984-3	0.3191	**1992-2**	0.2772	**2000-1**	0.0256
1984-4	0.4376	**1992-3**	0.3870	**2000-2**	0.0929
1985-1	0.2023	**1992-4**	0.4595	**2000-3**	0.0525
1985-2	0.2105	**1993-1**	0.3525	**2000-4**	0.0456
1985-3	0.3234	**1993-2**	0.4620	**2001-1**	0.0115
1985-4	0.3559	**1993-3**	0.5155	**2001-2**	0.0584
1986-1	0.2405	**1993-4**	0.5212	**2001-3**	0.0347
1986-2	-0.1499	**1994-1**	0.5229	**2001-4**	0.0732
1986-3	0.2249	**1994-2**	0.5412	**2002-1**	0.0538
1986-4	0.3225	**1994-3**	0.0797	**2002-2**	0.0758
1987-1	0.4050	**1994-4**	-0.5037	**2002-3**	0.0447
1987-2	0.8969	**1995-1**	0.4877	**2002-4**	0.0996
1987-3	0.1352	**1995-2**	0.3094		

Period	Mean
1980:1 / 1994:2	**0.3283**
1994:3 / 1998:4	**0.2712**
1999:1 / 2002:4	**0.0707**

II.3 – Filtered Coefficients of $\alpha_{1,t}$ – IPA

DATE	Coefficient Filtered Estimate	DATE	Coefficient Filtered Estimate	DATE	Coefficient Filtered Estimate
1980-1	0.8895	**1987-4**	0.8348	**1995-3**	0.8796
1980-2	0.8895	**1988-1**	0.9677	**1995-4**	0.8856
1980-3	0.8947	**1988-2**	0.9313	**1996-1**	0.8857
1980-4	0.8908	**1988-3**	1.0106	**1996-2**	0.8862
1981-1	0.8783	**1988-4**	0.9829	**1996-3**	0.8889
1981-2	0.8655	**1989-1**	0.8009	**1996-4**	0.8863
1981-3	0.8688	**1989-2**	0.6956	**1997-1**	0.8867
1981-4	0.8939	**1989-3**	1.3409	**1997-2**	0.8854
1982-1	0.9078	**1989-4**	0.7006	**1997-3**	0.8868
1982-2	0.8897	**1990-1**	1.3730	**1997-4**	0.8864
1982-3	0.8871	**1990-2**	0.4938	**1998-1**	0.8857
1982-4	0.8664	**1990-3**	0.8160	**1998-2**	0.8874
1983-1	0.8993	**1990-4**	0.9189	**1998-3**	0.8857
1983-2	0.9068	**1991-1**	0.7605	**1998-4**	0.8866
1983-3	0.9495	**1991-2**	0.6860	**1999-1**	0.8934
1983-4	0.9780	**1991-3**	0.9356	**1999-2**	0.6588
1984-1	0.9097	**1991-4**	0.8791	**1999-3**	0.9353
1984-2	0.9202	**1992-1**	0.5981	**1999-4**	0.8755
1984-3	0.9310	**1992-2**	1.0506	**2000-1**	0.8920
1984-4	0.9626	**1992-3**	0.9564	**2000-2**	0.8817
1985-1	0.8816	**1992-4**	1.0820	**2000-3**	0.8913
1985-2	0.8319	**1993-1**	0.9528	**2000-4**	0.8882
1985-3	0.8974	**1993-2**	1.0176	**2001-1**	0.8697
1985-4	0.9112	**1993-3**	1.0407	**2001-2**	0.8887
1986-1	0.8956	**1993-4**	1.0437	**2001-3**	0.8383
1986-2	0.6583	**1994-1**	1.0076	**2001-4**	0.8656
1986-3	0.9227	**1994-2**	0.9711	**2002-1**	0.8949
1986-4	0.8820	**1994-3**	0.6012	**2002-2**	0.8863
1987-1	0.8942	**1994-4**	0.6700	**2002-3**	0.8860
1987-2	1.0067	**1995-1**	0.9480	**2002-4**	0.8304
1987-3	0.6892	**1995-2**	0.8685		

Period	Mean
1980:1 / 1994:2	**0.9034**
1994:3 / 1998:1	**0.8606**
1999:1 / 2002:4	**0.8673**