EXCHANGE RATE AND INFLATION: A CASE OF SULKINESS OF VOLATILITY

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Abstract

There is a paucity of methodologically sound studies directly addressing exchange rate and inflation volatilities, and the existing ones suggest a relation between these variables, although there is no consensus about its characteristics. Therefore, it is necessary to verify the effects of exchange rate volatility, especially under an inflation-targeting regime where the monetary authority must know, as precisely as possible, the factors that affect price behavior. This paper seeks to establish the relation between exchange rate and inflation volatilities by adopting a more sophisticated econometric methodology than those applied so far - a bivariate GARCH model, dealing directly with the effects of conditional volatilities, which has been largely unexplored by the literature. We find a semi-concave relation between exchange rate and inflation variances, differently from what was estimated for financial series and in line with the intuition obtained from other studies. The article innovates by (i) trying to establish a relation between exchange rate and inflation volatilities and its implication for the monetary policy, (ii) applying a multivariate GARCH model, using conditional variances to analyze the relation between those volatilities and (iii) showing that traditional tests performed with exogenously constructed volatility series are sensitive to the criteria chosen to construct such series and do not reveal relevant features of that relation.

Key-words: exchange rate, inflation, volatility, GARCH models

JEL Classification: E31; F41.

I – INTRODUCTION

Although the literature about the impact of exchange-rate volatility is not as extensive as the one available on exchange-rate pass-through, some authors highlight such relation. Whether the impacts are significant or not remains controversial. However, CAPORALE and PITTIS (1995), point out that

"Economic theory concerned with behavior under uncertainty suggests that agents´ decisions are based upon the conditional distribution of the relevant random variables. In the presence of risk aversion, not only the conditional mean but also the higher moments, in particular the conditional variance, will play a role. (p. 397)"

The study of the effects of exchange rate volatility on inflation becomes important since we must verify whether that volatility should be a worry for the Central Bank's actions when monetary

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policy decisions are at stake. This is so because higher inflation volatility means higher uncertainty, affecting inflation expectations, a crucial variable in decisions under an inflation-targeting regime.

It is possible to classify the literature in two groups: those authors to whom there is no connection between exchange rate and macroeconomic variables volatilities and those who states the opposite. According to the first group, exchange rate volatility is not important to macroeconomic variables, since empirical evidence shows a substantial increase in it during floating exchange rate regimes, while macroeconomic variables do not present a similar rise in their volatilities. The second group establishes that relation focusing on different variables, mainly in the effects over trade. Concerning prices, analysis are made under different frameworks, the most common of them being exposed further on this introduction.

Concerning the first group of authors, we highlight the works of KRUGMAN (1988), BAXTER and STOCKMAN (1988), FLOOD and ROSE (1995), OBSTFELD and ROGOFF (2000), ROGOFF (2001) and DUARTE and STOCKMAN (2002).

KRUGMAN (1988) justifies the disconnection of exchange rate fluctuations from real fluctuations by a circular logic: since the exchange rate is volatile, it disconnects from fundamentals and, by being disconnected, exchange rates may be more volatile. The first is caused by the lack of reactions of firms to an increased volatility, both for seeing it as temporary and for the pricing-to-market, which prevents firms from altering their prices in countries to which they export their products as much as expected. On the other hand, by being disconnected, exchange rates may be more volatile because, if exchange rate fluctuations were immediately followed by changes in product allocation, changes in the trade account would be such that they would lead to exchange rate devaluation. If this devaluation were immediately passed through to prices, the resulting high inflation differentials (and, therefore, differentials in real interest rates) would bring the exchange rate back to levels that are “closer to reality” and, hence, volatility would be reduced.

BAXTER and STOCKMAN (1988) develop a study with a 49-country sample in the post-war period (1960 to 1985) under different exchange rate regimes and conclude that some series of real exchange rate, trade and industrial production became more volatile after the Bretton Woods period. However, evidence would not attribute the changes to the exchange rate regime, but to exchange rate itself.

To FLOOD and ROSE (1995), who carried out a monthly study on nine industrialized countries from 1960 to 1991, if the volatilities of macroeconomic variables do not change with the exchange rate regime, it is because there is not a clear trade-off between the reduction in exchange rate volatility and macroeconomic stability. When it comes to inflation, the authors do not find a trade-off between its level or volatility and exchange rate volatility. The results would be corroborated by a sticky price model, in which price variations are a function of the output gap, expected inflation and the output is a function of exchange rates, domestic and external prices and of the *ex-ante* real interest rates.

OBSTFELD and ROGOFF (2000) attribute the exchange rate disconnect puzzle (high volatility of exchange rate disconnected, apparently, from fundamentals)[3] to a mix of trade costs, monopoly and pricing-to-market in the domestic market. The tradables markets are very segmented, having effects on the exchange rate in such a way that they control its variation, even though nontradables push it up. With pricing-to-market at retail level, consumers will be insulated from exchange rate effects until these have been incorporated by wholesale import prices and, only afterwards, passed through to consumers in a period longer than that suggested by the half-life of PPP. Price rigidity also contributes to that disconnection: with rigid prices and macroeconomic variables (e.g. consumption) insulated from exchange rates in the short run, the exchange rate adjustment has a minimal effect and, therefore, it has to be high to allow for the clearing in financial markets. Real effects would be, therefore, too delayed to be captured by econometric tests. As incomplete pass-through in the short run is also pointed out by DEVEREUX and ENGEL (2002), the exchange rate volatility is much higher than that resulting from shocks to other variables.

ROGOFF (2001) points out that, in the macroeconomic chaos of the 1970s, there was a perspective that the calmness in exchange rate markets would follow inflation stabilization since, even in the weak version of PPP, price instability is incompatible with exchange rate stability. The belief was strengthened by Dornbusch's overshooting model, where exchange rate has a disproportionate adjustment to monetary shocks in the short run. Therefore, monetary instability would lead to even higher exchange rate instability. However, the evidence in the following decades showed that the latter is, at most, part of the former, since the volatilities of the world's main currencies remained at high levels. This volatility has costs over exports and imports, in addition to hedging costs. However, one should not analyze the effects of volatility by looking at the exchange rate regime – since volatility is always higher under floating exchange rates – but rather ask if that volatility makes product, investment and consumption more volatile. Nonetheless, the author points out that the empirical issue is not solved, as the differences in effects may be due to microeconomic distortion of the models. Although exchange rate volatility is a disturbance, the author does not consider it strong enough to be an economic policy target.

DUARTE and STOCKMAN (2002) credit this fact to the presence of noise traders, so that the same variability in exchange rate does not affect real variables.

As for the authors who find some relation between exchange rate and macroeconomic variables volatilities, we highlighted those whose works deal, in some sense, with inflation volatility or whose conclusions may be related to it. Concerning specifically the impacts of exchange rate volatility on prices, the available literature is even scantier, with most articles focusing on the volatility of relative prices across countries (basically, developed countries) or pursuing different aims (for instance, investment decisions), in which the result concerning inflation is a by-product.

[3] According to the authors, the *Purchase Power Parity Puzzle* – that is, the non-validity of PPP in the short run – may be considered as a special case of the *Exchange Rate Disconnect Puzzle*.

Few studies refer to the effects of exchange rate volatility on domestic price inflation, either in mean or in variance.

However, even with a wide-ranging literature, without a defined framework or line of study, we may classify the articles that stablish a relation between exchange rate and price volatility in six broad lines, according to their focus: hysteresis, trade effects, source of shocks, impact on relative prices, role of welfare and impact on inflation.

Concerning the role played by hysteresis, we mention the work of DIXIT (1989). His conclusions are that trade flows and prices depend on the investment made on a future basis and, consequently, depend on both expectations and on higher moments of the distributions involved. Therefore, the pattern of changes is affected when the environment is altered (KRUGMAN, 1988).

As for the role of trade markets in the impacts over prices, we may mention the works of CALVO and REINHART (2000a and 2000b). In CALVO and REINHART (2000a), they argue that changes in commodity prices are a frequent source of disturbances in emerging economies, requiring an adjustment in exchange rates. If this adjustment occurs, then one can notice a similar degree of volatility in commodity prices. Hence, their value in domestic currency should be relatively stable (a devaluation would reduce commodity prices). However, if there is fear of floating, exchange rates do not adjust and commodity prices in local currency also fall. Their results show that those prices in local currency are more volatile than exchange rates, especially for emerging economies. Besides, in most cases, the correlation between those two variables is small and not significant, which means that exchange rate does not play its role as a shock absorber, probably a result of the fear of strong fluctuations.

As for CALVO and REINHART (2000b), in emerging markets, exchange rate volatility is harmful to exports and imports since they become less competitive due to difficulties in being hedged – since forward markets are illiquid or inexistent – and to a high pass-through. Therefore, exchange rate volatility will have effects on inflation, allowing for a contractionary devaluation in the absence of perfect capital mobility, contrary to the results of traditional macroeconomic models.

The source of shocks is recalled by HAUSSMANN, PANIZZA and STEIN (2001) and by BARKOULAS, BAUM and CAGLAVAN (2002). In the former, real shocks imply that exchange rate flexibility is an important factor for output stabilization. The authors also find a negative and significant correlation in their tests between pass-through and measures of volatility. The test is performed for 12 industrialized countries and 26 developing ones, and the authors construct different measures to analyze to what extent a country has, indeed, a floating exchange rate. Those measures are related to the behavior of flotation, the extent to which there is intervention by means of using reserves or interest rates to stabilize the exchange rate. The aim is to analyze if there is any interference of the monetary authority in the exchange rate market even if a country declares itself as a free floater.

For BARKOULAS, BAUM and CAGLAVAN (2002), an analysis considering the effects only of exchange rate level on trade is incomplete, since it does not lead to optimal behavior forecasts;

therefore, second moments need to be related. With an extraction model, they show that the direction and the optimizing decision about exports and imports depend on the source of the shock. Agents form their expectations about future exchange rates based on the available information and, as invoices are issued in foreign currency, they are exposed to the exchange rate risk. Therefore, the more accurate the information is, the better the forecasts will be, implying effects on trade flows not only in the level but also in volatility. The model is tested for developed and developing countries and considers three sources of shocks over exchange rates: microstructure, stochastic behavior of fundamentals and disturbance in future policy signs. Its development shows that the response of trade flow movements may increase or reduce in face of exchange rate volatility, according to its source and magnitude.

The relation between exchange rate and relative prices volatilities is present in WEI and PARSLEY (1995), ANDERSEN (1997), SMITH (1999), DEVEREUX and ENGEL (2003) and CHEN (2004). WEI and PARSLEY (1995) analyze a panel with 12 tradable sectors in 91 pairs of OECD countries (total of 14 countries) to study PPP deviations and find that increases in exchange rate volatility associated with increases in the variability of price differentials (consistent with the idea of sticky prices in the local currency and incomplete pass-through) and a more stable exchange rate promote a faster convergence to PPP. In PARSLEY and WEI (2000), the authors present a panel analysis of 27 tradables over 88 quarters in 96 cities in the USA and Japan, finding that exchange rate volatility has a positive and significant effect on price dispersion (measured as variability of relative prices) between countries and that when one adds this variable (among others), part of what is primarily attributed to the border effect is reduced.

ANDERSEN (1997) presents an intertemporal macroeconomic overlapping generations model for an open economy with nominal rigidity. There, the author also highlights the effects of nominal volatility on the deviation of real exchange rates (relative prices).

SMITH (1999) tests a regression of changes in the real domestic price of good *i* – defined as $P_i = ep_i^c / \pi$, where *e* is the nominal exchange rate, P_i^e is the external price of good *i* and π is inflation – against the exchange rate. According to the model developed, the value of the coefficient will tell whether the exchange rate variance will increase or reduce the volatility of real domestic prices. Results show that an increase in exchange rate volatility causes a drop in inflation volatility in about 31% of times. Another indicator proposed by the authors is the value of the equation's R^2, which shows how much of the variance in the real domestic price is explained by exchange rate movements.

DEVEREUX and ENGEL (2003) assert that free exchange rates allow for the adjustment of relative prices when goods prices are sluggish. ENGEL and ROGERS (2001), on their turn, study border effects on relative prices for a sample of 55 European countries from 1981 to 1997, finding that exchange rate volatility explains part of the deviations in those prices.

The literature reviewed by CHEN (2004) points to an increase in price stickiness in face of the uncertainty derived from exchange rate volatility (i.e., firms become more reluctant to adjust their prices due to the possibility of a later reversion in the exchange rate). Besides, volatility would explain much of PPP failure in cross-country analysis and reduces the speed of mean reversion or of adjustments towards PPP. By testing the persistence of relative prices (speed of convergence) the author finds a positive and significant coefficient for exchange rate volatility. Thus, the higher the former is, the stickier the prices are. In other words, higher exchange rate volatility means poorer response of inflation to exchange rate movements.

Welfare literature is recalled by GHOSH, GULDE, OSTRY and HOLGER (1997) and by SUTHERLAND (2002). The former points to inflation volatility as being, at least, as important as average inflation. The authors analyze 140 countries over a 30-year period and, afterwards, divide them into 9 groups according to their exchange rate regime. Then, they regress inflation volatility against the Central Bank's turnover rate, degree of openness, volatilities of output growth and of money supply and interest rates, plus dummies for fixed and intermediate exchange rate regimes. However, when they split the countries into groups, results show that inflation volatility is lower under the floating and intermediate exchange rate regime for countries with low inflation.

SUTHERLAND (2002) recalls that, in the welfare literature, exchange rate volatility has a direct effect on welfare when the pass-through is incomplete. Hence, in this case, monetary policy should take exchange rate volatility into consideration. The developed model shows that increasing or decreasing exchange rate volatility to obtain domestic price stability may be an optimum from the welfare point of view. Decision will depend on the model's parameters, which indicate whether prices are set according to the producer's currency or local currency and if domestic demand reacts to exchange rate changes. These patterns will tell whether the relation between exchange rate and inflation volatilities is positive or negative.

Among the works that deal directly with inflation volatility, we have BARONE-ADESI and YEUNG (1990), BLEANEY (1996), SEABRA (1996) and BLEANEY and FIELDING (2002). BARONE-ADESI and YEUNG (1990) use descriptive statistics of developed countries between 1961 and 1984 and simple OLS regressions between real and nominal exchange rates and output. Their results show the effects of volatilities on output and their positive correlations with average inflation.

BLEANEY (1996) also states that the present value of an investment project depends on the expected value of future demand, price level and relative prices. Therefore, uncertainty over relative prices affects investment decisions and, hence, output growth. By testing a sample with 41 developing countries between 1980 and 1990, the author finds a negative relation between real exchange rate instability and growth, and a strongly positive correlation between inflation and real exchange rate.

SEABRA (1996) uses a model of intertemporal optimization with asymmetric adjustment costs. As in DIXIT (1989) and KRUGMAN (1986), even if the exchange rate is at a level that makes investment profitable, when uncertainty comes, the firms will not have incentives to enter the market due to asymmetric costs (for entering or leaving the market). The critical value that leads a firm to invest is a function of uncertainty: if uncertainty is high, in case of an exchange rate devaluation (appreciation), but the exchange rate is at a level below (above) the critical level, the optimal decision will be to wait before making a movement (*wait-and-see* strategy). This attitude impacts not only on aggregate output but also on inflation: if firms expand or enter the market, there is an increase in aggregate supply and, consequently, prices will fall, while the opposite will lead to a price increase.

In face of the evidence of lower inflation under fixed exchange rates, BLEANEY and FIELDING (2002) develop a model to test whether controlled exchange rates mean lower inflation and higher variability of output and inflation. The model of the BARRO and GORDON (1983) type, where policymakers set inflation and output targets, is tested through a panel that included 80 developing countries from 1980 to 1989, leading to the conclusion that not only mean inflation but also output and inflation standard deviations vary with the exchange rate regime (a floating regime, for instance, presents higher inflation and lower output volatilities).

In light of what has been shown here so far, we can conclude that there is a paucity of methodologically sound studies directly addressing exchange rate and inflation volatilities. However, results regarding not only the effects on inflation but also on other variables suggest that this relation exists, although there is no consensus about its characteristics. Therefore, it is necessary to verify the effects of exchange rate volatility, especially under an inflation-targeting regime where the monetary authority must know, as precisely as possible, the factors that affect price behavior. Once inflation volatility may be considered as a measure of uncertainty, and since inflation expectations are important variables in monetary policy actions, understanding the mechanisms that may affect this uncertainty becomes relevant in order to take more precise decisions in face of such mechanisms.

Hence, this paper seeks to establish the relation between exchange rate and inflation volatilities by adopting a more sophisticated econometric methodology than those applied so far: instead of constructing exogenous volatility series (by using standard deviations or variance of subsamples or windows) we will apply a bivariate GARCH model, working with conditional volatility series. The purpose of this procedure, as will be clear later in this paper, is to adopt a measure that is not sensitive to individual selection criteria, besides a more suitable econometric technique. At this point, it is important to make clear that we are not referring to the overall effects of exchange rate volatility, which include effects on investment decisions, international trade, etc. Our aim is to verify whether this variable is relevant enough so that the Central Bank should monitor it in its decisions about price behavior. This is the economic contribution proposed by this paper, since studies of this kind are scarce, especially in Brazil.

The paper is divided into six sections, including this introduction. Section II introduces the theoretical model that led to the econometric tests, while section III presents the data. The results obtained by the use of traditional methods (i.e.: unconditional variance series) are presented in section IV. Section V shows the results of the bivariate GARCH model, while section VI concludes.

II – THE THEORETICAL MODEL

Our approach is based on BLEANEY and FIELDING (2002), with slight modifications. The government has a utility function Z, of the BARRO and GORDON (1983) type, to be maximized. Z is given by equation (1), which represents the case where the government of a country faces a trade-off between price stabilization and output growth above its equilibrium level.

$$Z = -0.5\pi^2 - 0.5b\,(y - y^* - k)^2 \qquad (1)$$

Where π is inflation, y is the output level and y* is potential output. The term b>0 is the relative weight that the government gives to output maximization instead of price stability and k>0 represents the inflationary bias of the government. The presence of *b* and *k* comes from the assumption that, in developing countries, governments tend to attribute a higher weight to output growth to the detriment of price stability.

The restriction imposed by the authors upon function Z consists of an expectations-augmented Phillips Curve, including the exchange rate regime. Here, we have the first difference to the model of BLEANEY and FIELDING (2002) since we will focus not on the real but on the nominal exchange rate. Our restriction will be a Phillips Curve for an open economy similar to that one in BOGDANSKI, TOMBINI and WERLANG (2000). Hence, our restriction is given by

$$\pi_t = a_0\,\boldsymbol{\pi_t^e} + a_1\,\pi_{t-1} + a_2\,\pi_{t-2} + a_3\,(y-y^*)_{t-1} + a_4\Delta(p_t^{ext} + s_t) + \varepsilon_t \qquad (2)$$

where p_t^{ext} is the foreign price level, s_t, the nominal exchange rate and π_t^e the inflation expectation between period t and period t+1.

Let us, then, also assume that the exchange rate follows a random walk. This assumption is present in almost all partial equilibrium studies, since the forecasts for the exchange rate path do not perform better than a random walk. Thus, we have

$$s_t = s_{t-1} + \eta_t \qquad \eta_t \sim N(0, \sigma^2_\eta) \qquad (3)$$

applying (2) and (3) to (1), and obtaining the first-order condition for the maximization of Z with respect to π, we have

$$\pi_t = K' + \beta\pi^e + a_1\beta\pi_{t-1} + a_2\beta\pi_{t-2} + a_4\beta\pi_t^{ext} + a_4\beta\eta_t + \beta\varepsilon_t \qquad (4)$$

where $\beta = (b)/(1+b)$ and $K' = \beta a_3$

Some assumption also must be made concerning the behavior of π_t^e. Let us, then, consider that inflation expectations are of the form:

$$\pi_t^e = \pi_{t-1} + \nu_t \qquad (5)$$

We opted for π_{t-1} instead of π_t in the equation above since inflation in period *t* is not known in that period.[4] Therefore, agents will consider the information available in period *t*, that is, π_{t-1}, when forming their expectations.

Thus, substituting (5) in (4) we get that

$$E(\pi) = K' + (1+a_1)\,\beta\pi_{t-1} + a_2\beta\pi_{t-2} + a_4\beta\pi_t^{ext} \qquad (6)$$

[4] In the Brazilian case, for instance, inflation of month *t* is published during the first two weeks *t+1*.

Being inflation variance given by

$Var(\pi) = \beta^2E(\nu_t)^2 + \beta^2 a_4^2\, E(\eta_t)^2 + \beta^2\, E(\varepsilon_t)^2$ (7)

But, from (8), we have that $E(\varepsilon_t)^2$ is the inflation variance. Hence,

$Var(\pi) = \mu_0 var(\nu_t) + \mu_1 var(\eta_t)$ (8)

where $\mu_\mathbf{0} = \beta^2/(1-\beta^2)$ and $\mu_\mathbf{1} = \mu_\mathbf{0} a_4$

Inflation variance is, therefore, a function of ν_t (the variance of the shock expected in *t* in relation to *t-1* inflation) and of η_t (variance of the exchange rate process).

From these results, we may test the relation between volatilities and our aim is to do that by using a multivariate GARCH model. However, due to the small sample available – since Brazil adopted a floating exchange rate regime only in January 1999 – the estimation of a multivariate GARCH model with three variables is not viable because of the large number of terms to be estimated in the process. Aside from that, the inflation expectation research published by the Central Bank of Brazil started only in April, 2000, reducing our sample even further. Therefore, we will assume that ν_t is constant and, hence, equation (8) becomes:

$Var(\pi) = \mu'_0 + \mu_1 var(\eta_t)$,

where $\mu'_0=\mu_0 var(\nu_t)$ is the new constant.

Although the assumption that ν_t is constant is strong, it is not unlikely. Table 1 shows the result of a regression of π_t^e against π_{t-1} and a constant. If our hypothesis that the shock expected to *t+1* in comparison with *t-1* is, on average, constant, then the residuals of this equation should be homoskedastic. The data for π_t^e refer to the average market expectations for IPCA[5] inflation in month *t+1* as in the last business day of month *t-1*. These data are published by the Investor Relations Group (Gerin) from the Central Bank of Brazil.[6] As we may see, we accept the null hypothesis of homoskedasticity, which allows us to assume that ν_t is constant.[7]

Besides, if we compute $\nu_t = \pi_t^e - \pi_{t-1}$, since these data are available, one can notice that almost the entire series is within the interval of one standard deviation from the mean, as shown in Graph 1. The longest period in which it was outside that interval was during the moment of political uncertainty caused by the results of the presidential election of October/November 2002[8] (December 2002 to April 2003) and the first months of the new government.

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[5] Index of consumer prices considered by the Central Bank in the inflation targeting

[6] http://www4.bcb.gov.br/?FOCUSERIES

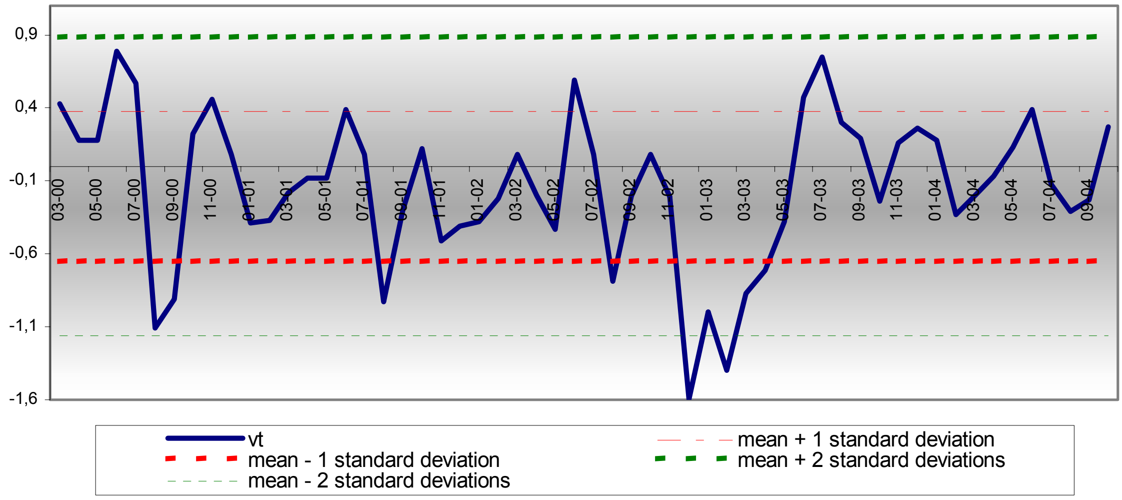
[7] Equivalent tests to π_t and s_t from equations 8, 9 and 10 accepted the alternative hypothesis of heteroskedasticity.

[8] In this election, the opposition had its left-wing candidate, Luis Inácio Lula da Silva, elected and there was a great uncertainty concerning the future of monetary policy. There was some fear that the fight against inflation would be abandoned and episodes that occurred on other occasions in Brazilian history - such as default, price freezing and capital controls - would happen again.

Table 1: Estimation of Equation $\pi^e_t = c + \pi_{t-1}$ (method: OLS)

Variable	Coefficient	Standard deviation	t-statistics	p-value
π_{t-1}	0.1593	0.0542	2.9385	0.0049
C	0.4463	0.0555	8.0351	0.0000
MA(1)	0.6576	0.1145	5.7452	0.0000
R^2	0.4584	Durbin-Watson		1.8755
adjusted R^2	0.4376	White Test for homoskedasticity: (p-value)		0.4354

Graph 1 – Evolution of $\nu_t = \pi^e - \pi_{t-1}$



III – Data

The data used in our estimations were the following:

a) Price Index: Extended Consumer Price Index (IPCA), consumer price index published by the Brazilian Institute of Geography and Statistics (IBGE),[9] December/1993=100 and considered by the Central Bank of Brazil as the reference index in the inflation targeting regime;

b) Exchange Rate: Exchange rate R$/US$, selling prices, monthly average ;

c) External Prices: *Producer price index* (PPI), published by the *Bureau of Labour Statistics*[10] (*commodities*, final goods).

d) GAP: output gap. It was computed by subtracting the industrial production series published by IBGE (used as a *proxy* for monthly GDP) from the trend obtained by the Hoddrick-Prescott filter.

All series were deseasonalized by the X-12 method and, afterwards, taken in logarithms (ln). Next, unit root tests were performed. All series, except for *gap* have unit roots, as shown in Table 2 and, therefore, they were taken in first differences. The series in first difference of *Price Index*, *Exchange Rate* and *External Prices* are henceforth referred to as *IPCA*, *E* and *PPI*, respectively.

[9] http://www.ibge.gov.br
[10] http://www.bls.gov/data/

Table 2 – Unit Root Test

Variable	ADF test statistics	Critical value at 5%	ADF test statistics –first difference of the variable	Critical value at 5%
Price Index	-2.170379	-3.478305	-3.904127	-3.478305
Exchange Rate	0.383569 [(a)]	-1.945745	-7.427513	-3.478305
External Prices	2.767025 [(a)]	-1.945745	-8.014687	-3.479367
GAP	-2.765390 [(b)]	-1.946072	-	-

Note: (a): test performed without trend or intercept; (b): test performed without trend (see ENDERS, 1995, ch. 4.7)

IV – TESTS WITH UNCONDITIONAL VOLATILITY

The first step to test the relation between exchange rate and inflation volatility was to apply the methods found in the literature to analyze the relations between the volatilities of exchange rate and other macroeconomic variables. In such cases, the most common measure of volatility is calculated by the standard deviation from the mean, either by splitting the series into small subsamples and computing the standard deviation of each one of them – what reduces the sample size – or by adopting rolling windows, maintaining the original sample size. Another common measure is the variance of such subsamples or windows.

ENGEL and ROGERS (2001) and BOWE and SALTVEDT (2004) measure volatility as the variance of the first difference of the series. LEVY-YEYATI and STURZENEGGER (2002), WEI and PARSLEY (1995), PARSLEY and WEI (2000), CHEN (2004), HAUSMANN, PANIZZA and STEIN (2001), BLEANEY and FIELDING (2002), BLEANEY (1996), BAXTER and STOCKMAN (1988) consider volatility as the standard deviation of the series in the period. All these authors, however, work with panel data and, thus, have one observation by country or one observation by country per year, resulting in quite a small sample for time series analysis. Furthermore, the small frequency of data, with annual observations may cause, in some cases, loss of relevant information, since in a one-year period many of the shocks and their effects on variances may have already been accommodated.

FLOOD and ROSE (1995) and BARONE-ADESI and YEUNG (1990) calculate volatility as the standard deviation of the first difference of exchange rates in subsamples. GHOSH et alli (1997) compute variance as the three-year-centered moving standard error of the inflation residual in an AR(1) equation (according to the authors, the results are similar to those obtained by using the standard deviation of inflation). CALVO and REINHART (2000) measure volatility by a frequency distribution of monthly percentage changes in the exchange rate, choosing the values 1% and 5% as thresholds. The higher the probability of staying within the interval, the smaller the volatility. CASTELLANOS (2004) considers the variance decomposition of a VAR to analyze the effects of shocks to a variable on the variance of another. CAPORALE and PITTIS (1995), in their turn, try to analyze the statistical properties of some economic variables under different exchange rate regimes by using ARCH models to generate the conditional variance.

BASTOURRE and CARRERA (2004) attribute the lack of macroeconomic studies about volatility to the lack of a pattern to define or to measure volatility. According to them, the use of *rolling windows*, instead of subsamples, has the advantage of reducing information loss (resultant from the reduced sample size). However, this procedure is also limited since it is not an easy task to determine the ideal number of observations in a window. In addition , the method to compute these series implies a high correlation, which may affect the quality of estimators. Furthermore, it is possible that the true relation between the volatilities of two different series is altered. For instance, once the exchange rate regime varies over time, a certain window may contain two different regimes.

In this paper, we opted for three different methods to calculate the unconditional volatility series. The first one is constructed by computing one standard deviation from the mean in rolling windows with 4, 6, 8 and 12 observations in each window (series are computed as the first difference of the natural logarithm of the variable on a monthly basis). The rolling window procedures were chosen to maintain the sample size. The second one considers the variances, instead of the standard deviation. Finally, we tested a VAR between the price index (IPCA) and the exchange rate (E) and analyzed the resulting variance decomposition.

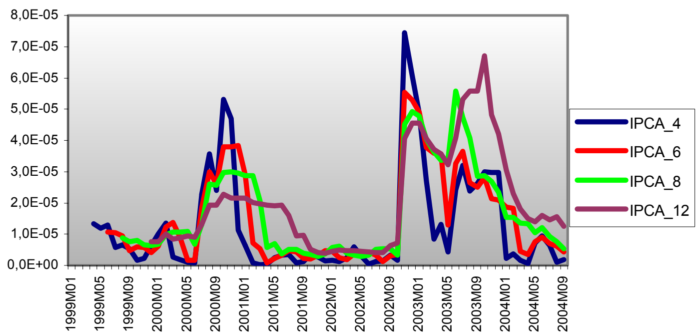
IV. 1 – Rolling Windows with standard deviations

The volatilities computed by the standard deviations are presented in Graphs 2 and 3, where *E_i* and *IPCA_i* are the volatilities of E and IPCA, respectively, within a window of size *i*. In this analysis, it is possible to note that the series are sensitive to the size of the window. As Table 3 shows, the unit root test for *IPCA _i* is also affected by window size: *IPCA_4* is stationary and so is *IPCA_6*, although we reject the presence of unit roots in the former at a level of significance of 10%. However, *IPCA_8* and *IPCA_12* have unit roots. Since *E_i* is always stationary, we computed the first differences of *IPCA8* and *IPCA12,* named *d_ IPCA_8* and *d_IPCA_12,* respectively.

The estimation results also are very sensitive to window size. The number of lags in each VAR was chosen by taking into consideration the information criteria, absence of residual autocorrelation (LM test), absence of correlation between variables, and parsimony. In all models the dummy variable d2002_M11 - which assumes the unity value for November 2002 – was included, since in all series there is a peak in that month, probably associated with the political crisis. Its inclusion allowed us to correct problems of residual autocorrelation or correlation between the variables found in the model. For similar reasons, the dummy variables d1999 in the four-month window and d2003_M10 in the 12-month window were included. The latter assumes the unity value for April and May 1999 (peak in E_4) while the former equals the unity value for October 2003 (peak in IPCA_12).

Tables 4 to 7 report the results of the four VARs estimated for different window sizes. In the four-month window, the lagged terms of a variable in its respective equation and the effect of inflation variance on exchange rate variance are considered to be statistically significant. With regard to the six-month window, there are significant cross-terms. However, the Wald test shows that the sum of the lagged coefficients of E_6 in the IPCA_6 equation is not statistically different from zero, and the same happens to the lagged coefficients of IPCA_6 in the E_6 equation. Only the dummy and first lag of a variable are significant in the equation. In the eight-month window, only E_8(-1) in the equation for E_8 is significant, while only the dummy is significant in the D_IPCA_8 equation. However, in this VAR, the correlation between IPCA_8 and E_8 equals –0.43, which may jeopardize the OLS estimation. Finally, the VAR between d_IPCA_12 and E_12 reports the coefficient of E_12(-1) as the only significant one in the E_12 equation. *E_12(-1), E_12(-6)* and *E_12(-7)* are significant in the d_IPCA_12 equation and, according to the Wald test, their sum is statistically different from zero at a 10% level.

Graph 2 – Variances of IPCA (standard deviations from the mean) - *Rolling Windows*



Graph 3 – Variances of E (standard deviations from the mean) - *Rolling Windows*

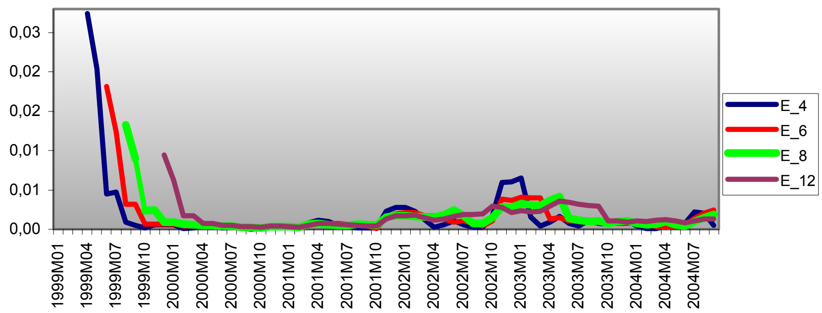


Table 3 – Unit Root Test – standard deviations

Variable	ADF test statistics	Critical Value at 5%	ADF test statistics – first difference of variables	Critical Value at 5%
IPCA_4	-3.469652	-3.480463	-	-
IPCA_6	-1.853977[a]	-1.946072	-	-
IPCA_8	-1.304640[a]	-1.946253	-7.001495	-3.486509
IPCA_12	-0.882262[a]	-1.94654	-5.987452	-3.49215
E_4	-11.95967	-3.481595	-	-
E_6	-10.85743 [b]	-2.908420	-	-
E_8	-9.505375[b]	-2.910019	-	-
E_12	-7.552068[b]	-2.913549	-	-

Note: (a): test performed without trend or intercept; (b): test performed without trend (see ENDERS, 1995, ch. 4.7)

Table 4 – VAR for four-month windows

Variables	E_4	IPCA_4	Variables	E_4	IPCA_4
E_4(-1)	0.279034	-5.65E-05	D2002_M11	0.005080	7.11E-05
	(0.04494)	(0.00040)		(0.00099)	(8.7E-06)
	[6.20854]	[-0.14234]		[5.13185]	[8.12319]
IPCA_4(-1)	21.86008	0.720580	D1999	0.012058	1.59E-06
	(7.73206)	(0.06834)		(0.00153)	(1.4E-05)
	[2.82720]	[10.5446]		[7.88981]	[0.11770]
C	0.000369	2.27E-06	R-squared	0.885547	0.738392
	(0.00016)	(1.4E-06)	Adj. R-squared	0.877917	0.720952
	[2.29918]	[1.59645]	F-statistic	116.0581	42.33774

Note: standard deviations between parenthesis; t-statistics between brackets.

Table 5 – VAR for six-month windows

Variables	E_6	IPCA_6	Variables	E_6	IPCA_6
E_6(-1)	0.902627	0.003413	IPCA_6(-1)	0.488179	0.824558
	(0.11178)	(0.00178)		(7.08504)	(0.11290)
	[8.07474]	[1.91570]		[0.06890]	[7.30320]
E_6(-2)	0.039786	-0.006402	IPCA_6(-2)	1.836084	0.145744
	(0.14826)	(0.00236)		(9.09015)	(0.14486)
	[0.26835]	[-2.70965]		[0.20199]	[1.00613]
E_6(-3)	0.023939	-0.000179	IPCA_6(-3)	-4.371941	-0.072066
	(0.15482)	(0.00247)		(8.76520)	(0.13968)
	[0.15463]	[-0.07236]		[-0.49878]	[-0.51594]
E_6(-4)	-0.134537	0.004324	IPCA_6(-4)	7.948727	-0.095348
	(0.11393)	(0.00182)		(8.64624)	(0.13778)
	[-1.18086]	[2.38140]		[0.91933]	[-0.69202]
E_6(-5)	-0.101762	0.000542	IPCA_6(-5)	-26.50704	-0.005847
	(0.08741)	(0.00139)		(7.82891)	(0.12476)
	[-1.16425]	[0.38908]		[-3.38579]	[-0.04687]
E_6(-6)	0.070599	-0.001031	IPCA_6(-6)	19.65405	-0.014069
	(0.04817)	(0.00077)		(6.15600)	(0.09810)
	[1.46568]	[-1.34296]		[3.19266]	[-0.14341]
C	0.000203	1.86E-06	D2002_M11	0.002768	4.59E-05
	(0.00011)	(1.8E-06)		(0.00043)	(6.8E-06)
	[1.82062]	[1.04619]		[6.50619]	[6.76973]
R-squared	0.881021	0.858823	Adj. R-squared	0.845868	0.817112
F-statistic	25.06256	20.58970			

Table 6 – VAR for eight-month windows

Variables	E_8	D_IPCA_8	Variables	E_8	D_IPCA_8
E_8(-1)	0.793006	0.001550	D_IPCA_8(-2)	5.380433	-0.043150
	(0.09864)	(0.00097)		(9.18843)	(0.09008)
	[8.03975]	[1.60262]		[0.58557]	[-0.47903]
E_8(-2)	0.009701	-0.000623	C	0.000203	-1.79E-06
	(0.06740)	(0.00066)		(0.00011)	(1.1E-06)
	[0.14393]	[-0.94255]		[1.82026]	[-1.63756]
D_IPCA_8(-1)	1.875515	0.098667	D2002_M11	0.001494	4.21E-05
	(9.21495)	(0.09034)		(0.00053)	(5.2E-06)
	[0.20353]	[1.09218]		[2.82229]	[8.11445]
R-squared	0.720969	0.578104	Adj. R-squared	0.694645	0.538302
F-statistic	27.38862	14.52466			

Table 7 – VAR for twelve-month windows

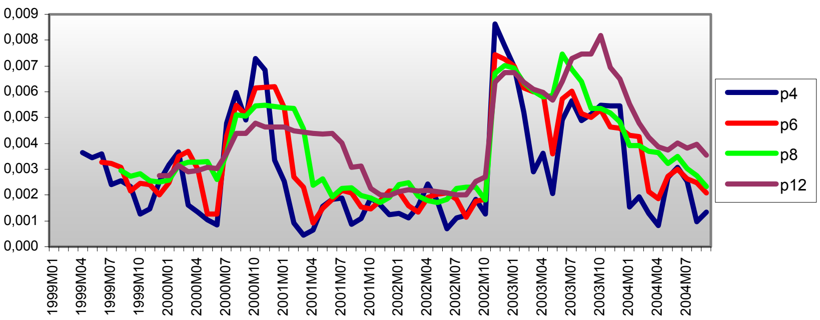
Variables	E_12	D_IPCA_12	Variables	E_12	D_IPCA_12	Variables	E_12	D_IPCA_12
E_12(-1)	1.126580	0.006572	E_12(-7)	-0.013448	0.001657	D_IPCA_12(-6)	6.952438	-0.085580
	(0.15476)	(0.00167)		(0.07273)	(0.00078)		(8.11413)	(0.08743)
	[7.27949]	[3.94088]		[-0.18491]	[2.11490]		[0.85683]	[-0.97882]
E_12(-2)	-0.086404	-0.004059	D_IPCA_12(-1)	-7.315069	0.154255	D_IPCA_12(-7)	-8.153833	0.016438
	(0.27024)	(0.00291)		(9.98464)	(0.10759)		(7.60833)	(0.08198)
	[-0.31973]	[-1.39396]		[-0.73263]	[1.43376]		[-1.07170]	[0.20050]
E_12(-3)	-0.151181	0.000891	D_IPCA_12(-2)	5.382736	0.106947	C	8.75E-05	-1.55E-06
	(0.27491)	(0.00296)		(9.84580)	(0.10609)		(9.6E-05)	(1.0E-06)
	[-0.54993]	[0.30091]		[0.54670]	[1.00807]		[0.90991]	[-1.49882]
E_12(-4)	0.197026	-0.002785	D_IPCA_12(-3)	-11.84970	-0.034154	D2002_M11	-0.000391	2.64E-05
	(0.26046)	(0.00281)		(10.5016)	(0.11316)		(0.00034)	(3.7E-06)
	[0.75645]	[-0.99235]		[-1.12837]	[-0.30182]		[-1.13746]	[7.12747]
E_12(-5)	-0.111431	0.002046	D_IPCA_12(-4)	-0.519314	-0.114882	D2003_M10	-0.001797	1.46E-05
	(0.23913)	(0.00258)		(10.1884)	(0.10978)		(0.00035)	(3.8E-06)
	[-0.46598]	[0.79386]		[-0.05097]	[-1.04645]		[-5.07601]	[3.82675]
E_12(-6)	0.021749	-0.004004	D_IPCA_12(-5)	14.21636	0.084784	R-squared	0.931680	0.854298
	(0.16543)	(0.00178)		(9.21351)	(0.09928)	Adj. R-squared	0.898556	0.783654
	[0.13147]	[-2.24615]		[1.54299]	[0.85401]	F-statistic	28.12652	12.09307

In sum, the relation between those two variables is sensitive to window size. According to the window size selected, we may accept or reject that the exchange rate variance affects inflation variance and the other way round, as well as accept or reject that lagged values of inflation variance will affect it.

IV.2 - Rolling Windows with variances

Once again, we have series that are very sensitive to window size, as shown in Graphs 5 and 6 (*pi* and *ei* are the volatility series for IPCA and E, respectively, computed as the variance of the sample inside the window). Concerning stationarity, the only difference from the standard deviation case is that the variance of IPCA in the six-month window is not stationary. Hence, we took the first difference of *p6*, *p8* and *p12*, and named them as *dp6*, *dp8* and *dp12*, respectively.

Graph 4 – Variances of IPCA (variances) - *Rolling Windows*



Graph 5 – Variances of E (variances) - *Rolling Windows*

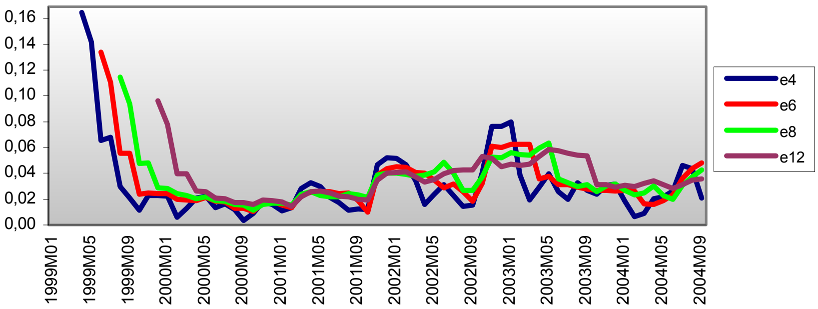


Table 8 – Unit Root Test – variances

Variable	ADF test statistics	Critical Value at 5%	ADF test statistics – first difference of variables	Critical Value at 5%
p4	-3.256578[b]	-2.906923	-	-
p6	-2.515552 [b]	-2.908420	-7.359868	-3.483970
p8	-0.845988 [a]	-1.946253	-7.339836	-3.486509
p12	-0.416680 [a]	-1.946654	-5.919263	-3.492149
e4	-6.571509	-3.481595	-	-
e6	-5.864106 [b]	-2.908420	-	-
e8	-5.436113 [b]	-2.910019	-	-
e12	-4.535080 [b]	-2.913549	-	-

Note: (a): test performed without trend or intercept; (b): test performed without trend (see ENDERS, 1995, ch. 4.7)

Table 9 – VAR for four-month windows

Variables	E4	p4	Variables	E4	p4
e4(-1)	0.652446	0.002503	C	0.008844	0.000748
	(0.06047)	(0.00665)		(0.00322)	(0.00035)
	[10.7898]	[0.37628]		[2.74707]	[2.11127]
p4(-1)	0.079097	0.700972	R-squared	0.664259	0.499806
	(0.83935)	(0.09233)	Adj. R-squared	0.653428	0.483671
	[0.09424]	[7.59224]	F-statistic	61.33301	30.97599

Note: Standard deviations in parentheses ( ) and t-statistics in square brackets [ ] .

Table 10 – VAR for six-month windows

Variables	E6	dp6	Variables	E6	dp6
e6(-1)	0.763029	0.008275	dp6(-1)	0.712408	0.022092
	(0.11265)	(0.01364)		(0.86253)	(0.10442)
	[6.77362]	[0.60681]		[0.82595]	[0.21157]
e6(-2)	0.010049	-0.012623	dp6(-2)	0.433394	-0.120232
	(0.09029)	(0.01093)		(0.85151)	(0.10308)
	[0.11130]	[-1.15486]		[0.50897]	[-1.16635]
C	0.006372	3.82E-05	D2002_M11	0.029878	0.005608
	(0.00240)	(0.00029)		(0.00771)	(0.00093)
	[2.66047]	[0.13188]		[3.87664]	[6.01086]
R-squared	0.725513	0.433191	Adj. R-squared	0.700559	0.381663
F-statistic	29.07471	8.406887			

Table 11 – VAR for eight-month windows

Variables	E8	dp8	Variables	E8	dp8
e8(-1)	0.899426	0.020323	dp8(-1)	0.066888	0.022146
	(0.11770)	(0.01505)		(1.05644)	(0.13511)
	[7.64180]	[1.35014]		[0.06331]	[0.16392]
e8(-2)	-0.059502	-0.013575	dp8(-2)	0.190518	-0.006954
	(0.09883)	(0.01264)		(1.05476)	(0.13489)
	[-0.60208]	[-1.07402]		[0.18063]	[-0.05155]
C	0.005153	-0.000214	R-squared	0.726663	0.033797
	(0.00249)	(0.00032)	Adj. R-squared	0.706416	-0.037773
	[2.06625]	[-0.66959]	F-statistic	35.88956	0.472226

Table 12 – VAR for twelve-month windows

Variables	E12	dp12	Variables	E12	dp12
e12(-1)	0.975153	0.041250	dp12(-1)	-0.904573	0.177430
	(0.10864)	(0.01242)		(1.15946)	(0.13250)
	[8.97595]	[3.32254]		[-0.78017]	[1.33909]
e12(-2)	-0.054296	-0.029020	dp12(-2)	-0.194250	0.177073
	(0.10135)	(0.01158)		(1.11872)	(0.12784)
	[-0.53574]	[-2.50567]		[-0.17364]	[1.38506]
C	0.002762	-0.000406	R-squared	0.834838	0.244949
	(0.00225)	(0.00026)	Adj. R-squared	0.821625	0.184545
	[1.22679]	[-1.57822]	F-statistic	63.18329	4.055166

Tables 9 to 12 show the results of the four estimated VARs. D2002_M11 was included for the six-month window case. For the four-month window VAR, only the lagged terms of each variable are significant and, differently from the previous case, the volatility of IPCA does not affect the exchange rate volatility. As for the six-month window, contrary to what was observed in the standard deviation case, the only significant terms are the dummy and the first lag of the exchange rate volatility in its own equation. In the eight-month window, we do not find the correlation problem we found before but, again, the only term that is significant is *e6(-1)* in the equation for the exchange rate variance. Finally, the VAR between *dp12* and *e12* indicates e12*(-1)* as the only significant variable in the equation for *e12* . In the equation for inflation variance, *e12(-1)* and *e12(-2)* are significant and the Wald test shows that their sum is statistically different from zero at a 10% level.

In sum, we notice that the results differ from the ones obtained when the standard deviation was used to compute the variance series concerning unit root tests, the number of lags in the VAR and the significance of some variances. None of the models showed that inflation volatility is affected by its lagged term, differently from what happens to exchange rate volatility. When it comes to cross-terms, we find that exchange rate volatility is significant in explaining inflation volatility in the 12-month windows. Hence, one can realize that results are sensitive not only to window size but also to the method chosen to compute volatility. In addition, since there are lagged effects in the case of exchange rate variance, we reinforce the adequacy of a GARCH-type model.

IV.3 – Variance decomposition in a VAR model

The last exercise performed in this section was to test a VAR between the price index and the exchange rate and to analyze variance decomposition. Since both series have unit roots, as shown in Table 1, we first tested for the presence of cointegration vectors, considering a linear trend in data (since IPCA is trend stationary). The number of lags in the cointegration test must be equivalent to the number of lags in the VAR between the variables. Hence, we tested a VAR between *price index* and *exchange rate* and found one lag by means of the information criteria. However, in that case, the LM test accepted the alternative hypothesis of residual autocorrelation, eliminated only by using four lags in the VAR. Therefore, our cointegration test was performed considering four lags. As shown in Table 13, the Trace and Eigenvalue tests do not accept the null hypothesis of presence of a cointegration vector.[11] For this reason, we will test a VAR between IPCA and E, i.e., the first differences of price index and exchange rates.

Table 13 – Cointegration test between exchange rate and consumer price index (ln of variables)

Trend assumption: Linear deterministic trend

Series: ln (exchange rates) and ln (consumer price index)

Lag interval (in first differences): 1 to 4

Unrestricted Cointegration Rank Test (Trace)

Number of cointegration vectors under Ho	***Eigenvalue***	***Trace statistic***	***Critical Value (5%)***	***p-value *****
None	0.104774	8.310901	15.49471	0.4328
At most one	0.018996	1.227468	3.841466	0.2679

The Trace test rejects the null hypothesis of the presence of at least one cointegration vector at 5%

** MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Number of cointegration vectors under Ho	***Eigenvalue***	***Trace statistic***	***Critical Value (5%)***	***p-value *****
None	0.104774	7.083432	14.26460	0.4793
At most one	0.018996	1.227468	3.841466	0.2679

The Eigenvalue test rejects the null hypothesis of the presence of at least one cointegration vector at 5%

** MacKinnon-Haug-Michelis (1999) p-values

The next step is to perform the Granger Causality test, since in variance decomposition factorization is made using the Cholesky method, where the ordering of variables may affect the result. The right ordering has the most exogenous variable first and, according to the test in Table 14, it means that E must precede IPCA. Table 15 shows the VAR results, where D2002_M11 was included once again, while Table 16 presents the variance decomposition.

The variance decomposition is shown in Table 16. We find that about 3% of the IPCA variance in t+1 may be explained by shocks in E in period t. There are increasing accumulated effects over time, and shocks in E explain around 42% of IPCA variance after 12 months. A shock in IPCA, in its turn, does not have an immediate effect on the variance of E, however it has lagged effects, although on a smaller scale.

[11] The conclusion is the same if we consider only one lag.

Table 14 –Granger Causality Test [12]

Null Hypothesis	Number of Obs.	F-statistic	p-value
E does not Granger-Cause IPCA	66	9.0601	5.0E-05
IPCA does not Granger-Cause E		1.4686	0.2323

Table 15 – VAR between E and IPCA

Variables	E	IPCA	Variables	E	IPCA
E(-1)	0.603612	0.035565	IPCA(-1)	0.918499	0.655619
	(0.12170)	(0.01111)		(1.27348)	(0.11626)
	[4.95968]	[3.20095]		[0.72125]	[5.63920]
E(-2)	-0.116778	-0.003658	IPCA(-2)	-2.703953	-0.205941
	(0.15010)	(0.01370)		(1.54984)	(0.14149)
	[-0.77801]	[-0.26692]		[-1.74467]	[-1.45551]
E(-3)	0.118308	0.022020	IPCA(-3)	2.459311	0.145171
	(0.12392)	(0.01131)		(1.54983)	(0.14149)
	[0.95469]	[1.94634]		[1.58683]	[1.02602]
E(-4)	0.082738	-0.005089	IPCA(-4)	-2.539942	0.062882
	(0.10334)	(0.00943)		(1.15037)	(0.10502)
	[0.80066]	[-0.53944]		[-2.20794]	[0.59875]
C	0.016518	0.001811	D2002_M11	-0.134947	0.014337
	(0.00836)	(0.00076)		(0.03438)	(0.00314)
	[1.97558]	[2.37300]		[-3.92558]	[4.56817]
R-squared	0.458232	0.716785	Adj. R-squared	0.369580	0.670441
F-statistic	5.168839	15.46653			

Note: Standard deviations in parenthesis ( ) and t-statistics in square brackets [ ] .

Table 16 – Variance Decomposition (Cholesky ordering: E IPCA)

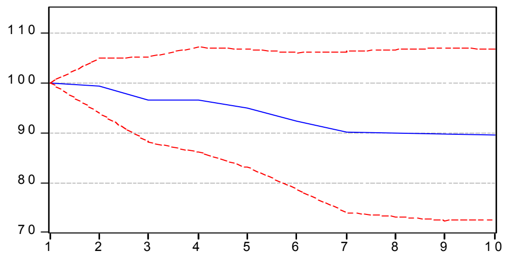
Variance decomposition of E:				***Variance decomposition of IPCA:***			
Period	Standard Error	E	IPCA	Period	Standard Error	E	IPCA
1	0.030575	100.0000	0.000000	1	0.002791	3.038439	96.96156
2	0.035574	99.49638	0.503623	2	0.003410	7.113075	92.88693
3	0.037020	98.21512	1.784882	3	0.003662	15.73568	84.26432
4	0.037345	98.23839	1.761613	4	0.003953	27.00305	72.99695
5	0.038343	97.46899	2.531007	5	0.004164	32.38073	67.61927
6	0.039049	95.48982	4.510177	6	0.004308	36.37634	63.62366
7	0.039491	93.57150	6.428498	7	0.004410	39.26637	60.73363
8	0.039649	92.88324	7.116760	8	0.004497	41.37337	58.62663
9	0.039819	92.27025	7.729753	9	0.004548	42.32425	57.67575
10	0.040014	91.74086	8.259143	10	0.004576	42.55194	57.44806
11	0.040189	91.44510	8.554900	11	0.004592	42.44773	57.55227
12	0.040340	91.39917	8.600828	12	0.004605	42.23988	57.76012

Graphs 6 to 9 show these decompositions over time, as well as the interval of ± 2 standard errors. We notice that shocks to the variables have positive effects on variances and that, apart from the impact of IPCA on E, they are different from zero. Therefore, we cannot rule out the

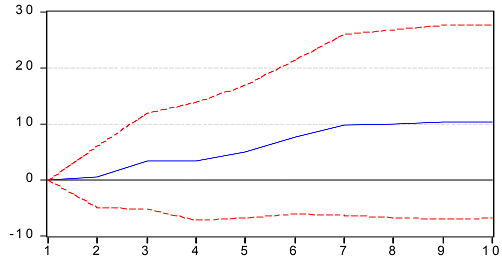
[12] It is important to include as many lags as possible in variable *x* that may be significant over variable *y*. We tested an equation with 13 lags in both variables and the highest significant lag of *x* over *y* was the third lag of E over IPCA. In the Granger Causality test the null hypothesis that IPCA Granger-Causes E is rejected both with 3 and with 13 lags.

hypothesis that shocks to the exchange rate – represented by η_t in equation (14) – might affect inflation variance.

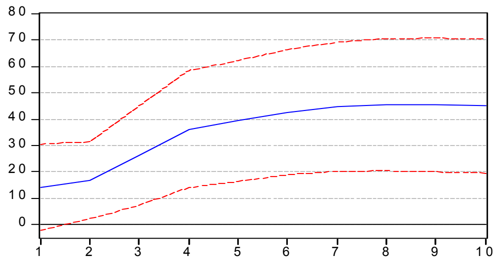
Graph 6 - Percent E variance due to E



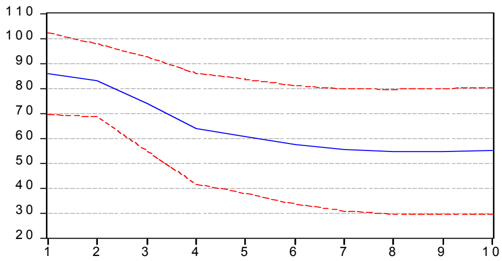
Graph 7 - Percent E variance due to IPCA



Graph 8 - Percent IPCA variance due to E



Graph 9 - Percent IPCA variance due to IPCA



Based on the results presented in this section, we may conclude that the traditional measures used to verify whether there is a relation between the volatilities of exchange rate and macroeconomic variables (standard deviations or variances in subsamples) yield results that are sensitive to the subsample size, leading us to accept or reject the significance of the relation according to the window size we are working with.

The variance decomposition, in its turn, indicates that shocks to the exchange rate affect inflation variance. Since volatility is also a measure of uncertainty, this result sounds more intuitive than some of those presented before: if the exchange rate affects inflation and has delayed effects (incomplete exchange rate pass-through in the short run), shocks to that variable will affect the uncertainty about future inflation. Besides, an adequate exchange rate model must consider the presence of conditional heteroskedasticity, as illustrated in Table 17. In this case, it is necessary to generate volatility series for both variables in the same way - hence, to consider conditional variance for both - and not simply compare the variance series obtained from a GARCH (p,q) model for the exchange rates with an exogenous measure of inflation volatility. Furthermore, we show that variance decomposition reports that shocks to the IPCA affect its variance, just as well as some of the results obtained in the rolling window procedure show us that IPCA volatility is affected by its past values, reinforcing the application of the test for a bivariate GARCH model with E and IPCA.

Table 17 - OLS Equation for E

Variable	Coefficient	Standard Error	t-statistic	p-value
C	0.0059	0.0092	0.6378	0.5258
AR(1)	0.4341	0.0964	4.5041	0.0000
R^2	0.2351	LM Test (1 lag) [(a)]		0.8708
Adjusted R^2	0.2235	ARCH-LM Test (1 lag)		28.6673 [(b)]

Note: (a) null hypothesis of absence of autocorrelation accepted also for higher number of lags; (b) null hypothesis of absence of ARCH residuals rejected at 1%.

V – TESTS WITH CONDITIONAL VARIANCE– BIVARIATE GARCH

Testing a GARCH model requires, first, some assumption about the mean equation and, bearing this in mind, we considered three different cases. The first one considers only lagged terms of each variable, the second one considers a Phillips Curve for the IPCA equation (according to equation 2 in Section II) and the lagged values for the exchange rate and, finally, the third case considers the Phillips Curve for the IPCA and a random walk for the exchange rate (equation 3 in Section II). The number of lags in the variables in the equations for IPCA and exchange rate was chosen by considering both the cross-correlograms and the OLS models.[13] With regard to variance specifications, we tested five different options: diagonal-Vec (BOLLERSLEV, ENGLE and WOOLDRIDGE, 1988), constant correlation (CCORR, from BOLLERSLEV, 1990), full parameterization (Vec), the BEKK restriction (ENGLE and KRONER, 1993) and the dynamic conditional correlation (DCC, from ENGLE, 2000). Only under the BEKK restriction convergence was achieved, and we consider some reasons for that further ahead in this section.

Table 18 shows the results found for each of the three assumptions about mean equations. For each assumption, different simulations were made changing the convergence criteria and the number of iterations. Therefore, it is possible that, for each assumption, we end up with more than one result achieving convergence. The choice between them was made based on two criteria. The first one was the calculation of the eigenvalues, ensuring that the condition of covariance stationarity was respected (see ENGLE and KRONER, 1993). The second one was the value of the likelihood function, being reported on table 18 the results with the highest values of the likelihood function among those respecting the covariance stationarity condition. The general form of mean, variance and covariance equations under the BEKK model is defined by:

Mean equations

$$IPCA = \delta_0 + \delta_1 IPCA_{t-1} + \delta_2 E_{t-1} + \delta_3 E_{t-2} + +\delta_4 GAP_{t-2} + \delta_5 PPI_{t-1} + \varepsilon_{1,t}$$

$$E = \gamma_0 + \gamma_1 E_{t-1} + \gamma_2 E_{t-2} + \varepsilon_{2,t}$$

Variance and Covariance equations[14]

[13] When they pointed to different number of lags, we chose the highest one.

[14] Variance and covariance equations are from ENGLE and KRONER (1993), equation 2.3, pages 5 and 6, without suppressing the GARCH terms.

$$h_{11} = c_{11} + a_{11}^2\varepsilon_{1,t-1}^2 + 2a_{11}a_{21}\varepsilon_{1,t-1}\varepsilon_{2,t-1} + a_{21}^2\varepsilon_{2,t-1}^2 + g_{11}^2 h_{11,t-1} + 2g_{11}g_{21}h_{12,t-1} + g_{21}^2 h_{22,t-1}$$
$$h_{22} = c_{22} + a_{12}^2\varepsilon_{1,t-1}^2 + 2a_{12}a_{22}\varepsilon_{1,t-1}\varepsilon_{2,t-1} + a_{22}^2\varepsilon_{2,t-1}^2 + g_{12}^2 h_{11,t-1} + 2g_{12}g_{22}h_{12,t-1} + g_{22}^2 h_{22,t-1}$$
$$h_{12} = c_{21} + a_{11}a_{12}\varepsilon_{1,t-1}^2 + (a_{12}a_{22} + a_{21}a_{12})\varepsilon_{1,t-1}\varepsilon_{2,t-1} + a_{21}a_{22}\varepsilon_{2,t-1}^2 + g_{12}g_{11}h_{11,t-1} + (g_{11}g_{22} + g_{12}g_{21})h_{12,t-1} + g_{21}g_{22}h_{22,t-1}$$

In order to make the analysis clearer, we renamed the coefficients above as:

$$c_{11} = \alpha_0;\ a_{11}^2 = \alpha_1;\ 2a_{11}a_{21} = \alpha_2;\ a_{21}^2 = \alpha_3;\ g_{11}^2 = \alpha_4;\ 2g_{11}g_{21} = \alpha_5;\ g_{21}^2 = \alpha_6$$
$$c_{22} = \beta_0;\ a_{12}^2 = \beta_1;\ 2a_{12}a_{22} = \beta_2;\ a_{22}^2 = \beta_3;\ g_{12}^2 = \beta_4;\ 2g_{12}g_{22} = \beta_5;\ g_{22}^2 = \beta_6$$
$$c_{21} = \mu_0;\ a_{11}a_{12} = \mu_1;\ a_{12}a_{22} + a_{21}a_{12} = \mu_2;\ a_{21}a_{22} = \mu_3;\ g_{12}g_{11} = \mu_4;\ g_{11}g_{22} + g_{12}g_{21} = \mu_5;\ g_{21}g_{22} = \mu_6$$

Hence, the variance and covariance equations can be rewritten as:

$$h_{11} = \alpha_0 + \alpha_1\varepsilon_{1,t-1}^2 + \alpha_2\varepsilon_{1,t-1}\varepsilon_{2,t-1} + \alpha_3\varepsilon_{2,t-1}^2 + \alpha_4 h_{11,t-1} + \alpha_5 h_{12,t-1} + \alpha_6 h_{22,t-1}$$
$$h_{22} = \beta_0 + \beta_1\varepsilon_{1,t-1}^2 + \beta_2\varepsilon_{1,t-1}\varepsilon_{2,t-1} + \beta_3\varepsilon_{2,t-1}^2 + \beta_4 h_{11,t-1} + \beta_5 h_{12,t-1} + \beta_6 h_{22,t-1}$$
$$h_{12} = \mu_0 + \mu_1\varepsilon_{1,t-1}^2 + \mu_2\varepsilon_{1,t-1}\varepsilon_{2,t-1} + \mu_3\varepsilon_{2,t-1}^2 + \mu_4 h_{11,t-1} + \mu_5 h_{12,t-1} + \mu_6 h_{22,t-1}$$

By analyzing Table 18, we notice that the results for the mean equations are quite similar, as well as the values in the variance equation for cases (1) and (2). Case (3) differs from the other two but, since that model has ARCH residuals for the equation of E and serial correlation of residuals for both mean equations, it may not be considered as a good model.

By considering the comparison between variance equations in cases (1) and (2), the differences lie in the signs of g_{12} and g_{22}, in the values of a_{11}, a_{22} and a_{12} and in the significance of coefficients μ_1, β_1 and β_2, that is, the impact of $\varepsilon^2_{1,t-1}$ on the conditional variance of E and in the covariance and impact of $\varepsilon^2_{1,t-1}\varepsilon^2_{2,t-1}$ on the conditional covariance of E.

It can be seen from Table 19 that the significance of μ_1, β_1 and β_2 is the only significant difference between both cases. The difference in the signs of g_{12} and g_{22} does not affect the final result because these coefficients are considered under three situations: (I) squared values; (ii) multiplied by each other, (iii) multiplied by coefficients that are statistically equal to zero. The differences in a_{11}, a_{22} and a_{12}, in their turn, fall within standard deviation boundaries, thus, they may not be considered significant. It is important to notice that for the inflation equation all cases provided the same signals and the same significance (i.e. if statistically equal to or different from zero). Therefore, our results for the response of IPCA to shocks in E are robust.

The Wald Test was performed to decide between cases (1) and (2). The unrestricted case – that is, case (2) – was preferred to the detriment of cases (1) and (3), as shown in Table 20 . Hence, we will consider case (2) as our results from now on.

By analyzing the results in the second column of Table 18, one can notice that the conditional variance of IPCA is affected (statistically significant) by shocks to the IPCA, E and shocks common to both. However, since α_1 and α_3 are square coefficients, we cannot determine whether the effects of IPCA and E shocks have a positive or negative sign. Lagged variances and covariances, however, do not play a significant role in explaining IPCA variance.

As for the conditional variance of E, it is affected by its lagged values and by lagged values of the conditional variance of IPCA – the latter of which goes undetected by almost all tests with unconditional variances – although we also cannot make assertions about the sign. Shocks common to both variables ($\varepsilon_{1,t-1}\varepsilon_{2,t-1}$) and covariance have a positive and significant sign. Graph 7 shows the estimated conditional variances over time. The correlation coefficient computed for the conditional variances of E and IPCA equals 0.9035.

Table 18 – Bivariate GARCH Results

	Variables	Case 1	Case 2	Case 3[a]
Function Value		548.5888337	558.7008734	548.0509755
Equation for IPCA	Constant	0.00282329*	0.00197323*	0.0011409*
		(0.0005596)	*(0.00050458)*	*(0.00051313)*
	$IPCA_{t-1}$	0.55715402*	0.57849083*	0.68315183*
		(0.07376155)	*(0.0609667)*	*(0.05967424)*
	E_{t-1}	-	0.03863824*	0.06380358*
			(0.00980401)	*(0.00779833)*
	E_{t-2}	-	-0.00073484	-0.00920641
			(0.00964137)	*(0.0073217)*
	GAP_{t-2}	-	0.01673604**	0.01552899**
			(0.00977043)	*(0.0095768)*
	PPI_{t-1}	-	0.09498593**	0.10285839*
			(0.05367548)	*(0.05499651)*
Equation for E	Constant	0.00669639	0.01229489*	0.01917512*
		(0.00424251)	*(0.0043067)*	*(0.00534423)*
	E_{t-1}	0.8094822*	0.60752556*	-
		(0.11426506)	*(0.12549928)*	-
	E_{t-2}	-0.22750191**	-0.16772685	-
		(0.13597143)	*(0.10694474)*	
Conditional variance of IPCA	α_0	0	0	0
	α_1	+	+	+
	α_2	+	+	+
	α_3	+	+	+
	α_4	0	0	0
	α_5	0	0	0
	α_6	0	0	0
Conditional variance of E	β_0	0	0	0
	β_1	+	0	+
	β_2	+	0	+
	β_3	+	+	+
	β_4	+	+	+
	β_5	+	+	+
	β_6	+	+	+
Covariance	μ_0	+	+	-
	μ_1	-	0	-
	μ_2	-	-	-
	μ_3	-	-	-
	μ_4	0	0	0
	μ_5	0	0	0
	μ_6	0	0	0

Notes: (a) case presents residual autocorrelation in both mean equations (LM test); residuals of ARCH-type in the exchange rate equation; standard deviations in parentheses; * and ** denote significance at 5% and 10%, respectively.

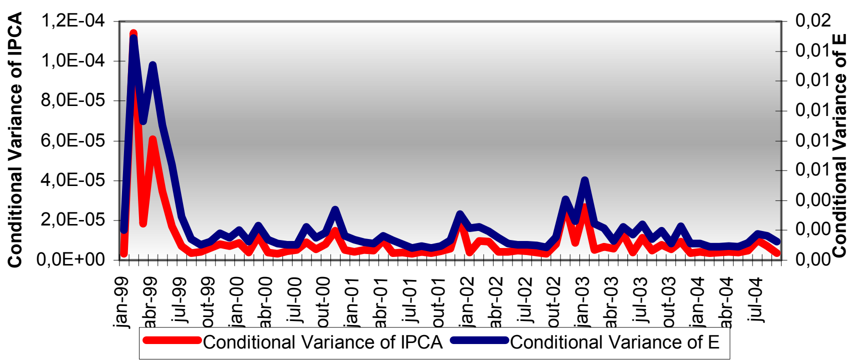
Table 19 - Estimated Parameters in Variance and Covariance Equations

g_{11}	-0.1041899	-0.0513807	0.03980506*
	(0.1348117)	*(0.11993625)*	*(0.14292026)*
g_{21}	0.00810385	0.01085598	-0.00743866*
	(0.01683404)	*(0.01418979)*	*(0.0156595)*
g_{12}	12.52650218*	-13.00132891*	13.12325566*
	(1.55221793)	*(1.68386503)*	*(1.98208477)*
g_{22}	0.41076103**	-0.43513845*	0.50120845*
	(0.23448881)	*(0.21248228)*	*(0.22752382)*
a_{11}	0.27491257*	0.40438372*	-0.44434782*
	(0.12289452)	*(0.13130784)*	*(0.17146475)*
a_{21}	0.0595035*	0.05305404*	-0.05393149*
	(0.0131604)	*(0.01178906)*	*(0.01283429)*
a_{12}	-3.42461647**	-5.71185332*	7.4208698
	(1.98123913)	*(2.10969805)*	*(2.41134849)*
a_{22}	-0.46199166*	-0.59176609*	0.78655762*
	(0.1406473)	*(0.14044079)*	*(0.16694871)*
c_{11}	-0.00000016	-0.00000008	0.00000006
	(0.007226)	*(0.00455714)*	*(0.00519676)*
c_{21}	0.00211834*	0.00175652*	-0.00189984*
	(0.00037541)	*(0.00033542)*	*(0.00031271)*
c_{22}	0.00366595	0.0038599	-0.00462995
	(0.00836578)	*(0.00918004)*	*(0.01001528)*
	(0.00836578)	*(0.00918004)*	*(0.01001528)*

Table 20 –Wald Test[15]

Cases tested	Observed χ^2_q statistic	**Null hypothesis:** Variables added in case (2) are not jointly significant
Case (1) *vs* Case (2)	20.22	Reject
Case (2) *vs.*Case (3)	21.30	Reject

Graph 7 – Conditional Variances - IPCA and E



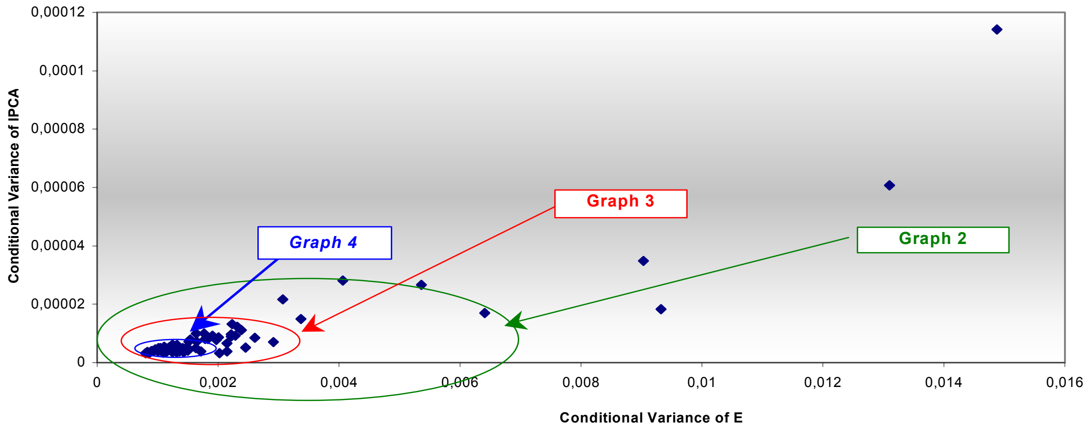
[15] Wald Test: $-2(l_r-l_u) \sim \chi^2_q$, where q is the number of added variables, l_r and l_u are the log-likelihood of the restricted and unrestricted cases, respectively. Under H_o, the added variables are not jointly significant.

At first, we considered that the lack of convergence for specifications other than the BEKK model would result from the small size of our sample. However, this may be questioned since the BEKK specification has more parameters than some of the other specifications tested. The negative sign of shocks in E over the conditional covariance and Graphs 8 to 12 suggest that the sign of shocks in E over the conditional variance of IPCA may not be the same all the time. If this is true, then we may have a reason for the non-convergence of specifications that, instead of working with squared terms (imposing the positivity of the matrix), try to find a sign for the relation.

Graphs 8 to 12 are dispersion graphs with the conditional variance of E on the horizontal axis and of IPCA on the vertical axis. Graph 1 plots the entire sample and one can clearly see four outliers in that graph, which correspond to the period between February and May 1999 (i.e. in the first months after the exchange rate crises which led to the change in the exchange rate regime and before the adoption of the inflation-targeting regime in June of that year). Hence, we excluded these observations and built Graph 2. Again, we have five outliers that were removed to construct Graph 3 (June 1999, November 2000, December 2001, December 2002 and January 2003). Graph 4, in its turn, was built using only the region with the highest concentration of observations (57% of the sample).[16]

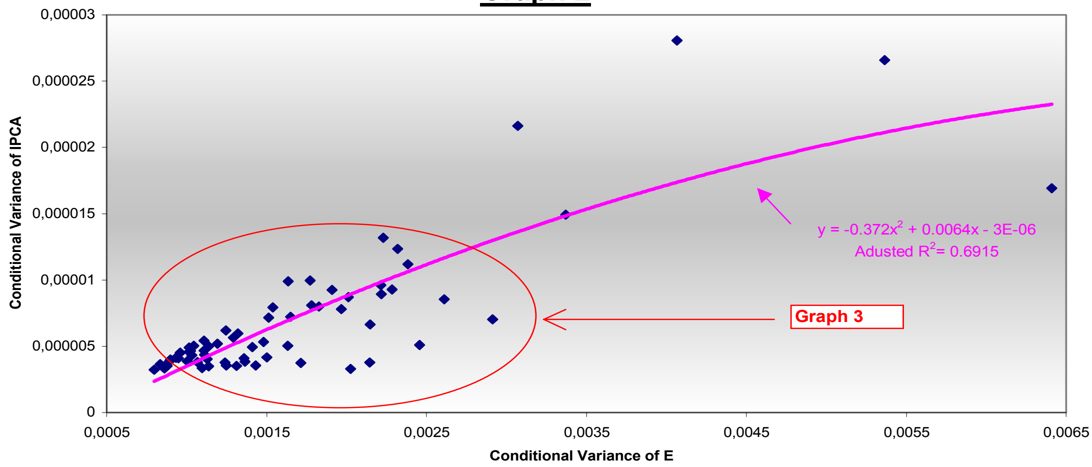
A trend was included in Graphs 2 to 4 and the adjusted R^2 of each trend equation was reported. The results show that if the sign of the relation between volatilities is not inverted by high values of exchange rate volatility, we can say that, at least, the response of inflation volatility to exchange rate volatility decreases as the latter rises. We can only say, for sure, that the relation between those two variables is not an increasing one; instead it has a (semi-) concave form, as opposed to the convex form observed in financial variables. If we have the *smile of volatility* in financial cases, Graphs 1 to 3 suggest that we have a case of *grumpiness of volatility* for the price index.

<u>Graph 1</u>

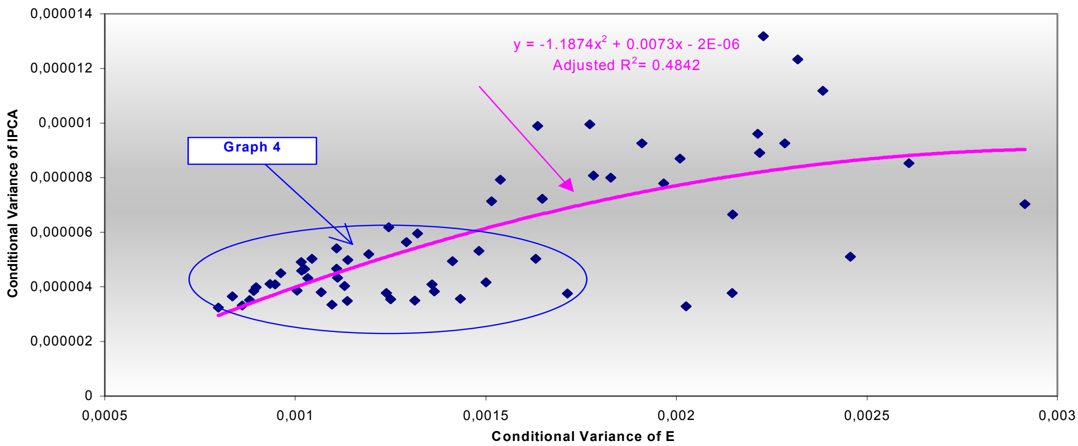


[16] The observations removed from Graph 4 are related to Graph 1: January to July 1999; November 1999 to January 2000; March, August, October and November 2000; April 2001; December 2001 to March 2002; October 2002 to March 2003; May, July, September and November 2003 and August 2004.

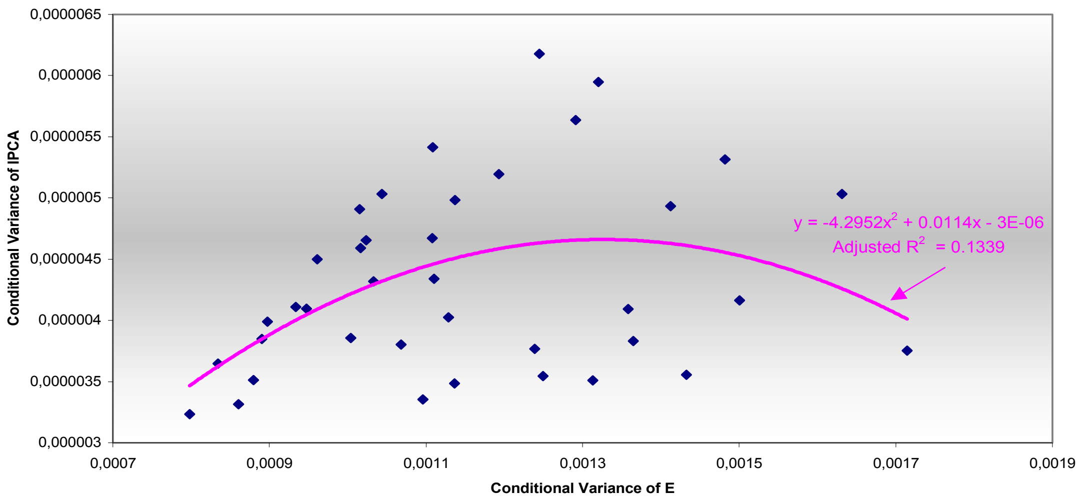
Graph 2



Graph 3



Graph 4



The results presented herein seem to be in line with DIXIT (1989) and SEABRA (1996) where, under uncertainty, the firm chooses to adopt a *wait-and-see* strategy. It also recalls us of the circular logic of KRUGMAN (1988): fluctuations are possible because they have small effects and they have small effects because they are possible: when volatility is very high, inflation response is reduced and effects are small. It also reminds us of the many results presented earlier in this paper, including those for other economic variables: LEVY-YEYATI and STURZENEGGER (2002), who found higher output volatility under fixed exchange rates; BARKOULAS, BAUM and CAGLAVAN (2002), who show that trade flow may increase or decrease in face of exchange rate volatility; ENGEL and ROGERS (2001), DEVEREUX and ENGEL (2003), WEI and PARSLEY (1995) and PARSLEY and WEI (2000), who find an inverse relation between exchange rate volatility and variations in relative prices between countries; SUTHERLAND (2002), who shows that the optimum relation between exchange rate volatility and prices – under a welfare approach – has different signs according to the parameters of the model, as in SMITH (1999), where the value of the estimated parameter will tell whether the exchange rate variance will increase or decrease the volatility of real domestic price; GHOSH, GULDE, OSTRY and HOLGER (1997), who find that inflation volatility is lower under floating and intermediate exchange rate regimes in countries with low inflation. Finally, this result corroborates the one found in ALBUQUERQUE and PORTUGAL (2004) where it is shown that the pass-through from exchange rates to consumer prices in Brazil decreased after the floating regime.

From our results, we conclude that maybe the exchange rate disconnect puzzle, as put by OBSTFELD and ROGOFF (2000), may be explained, at least when it concerns price variations. First, because there is a significant relation between exchange rate volatility and the volatility of a macroeconomic variable (i.e. prices). The gap is in the magnitude of this relation, as shown in Graph 7. In addition, since exchange rate effects are only part of the price formation process – the quarterly pass-through for Brazil is estimated between 0.04 and 0.06 – the proportion of deviations also should be smaller.

Besides, in the presence of nonlinearities, the puzzle could be justified by the following logic: in periods of high volatility, price setters do not have the same pattern of behavior due to higher adjustment costs (reputation costs, as in PARSLEY (1995) menu costs, etc), which inhibits the degree of price adjustment and, therefore, inflation volatility has a small amplitude. When the exchange rate volatility is not very high, inflation volatility answers more clearly. This explanation will be reinforced if the sign inversion present in Graph 4 is observed in further studies.

Concerning implications for monetary policy, results suggest that it is interesting to keep the stability of the national currency, since increases in exchange rate volatility imply an increase also in inflation volatility and, consequently, higher uncertainty, which may affect expectations about future inflation. However, in periods of crisis, when exchange rate volatility is very high, the monetary authority would not need to promote a strong intervention in exchange rate markets to protect the inflation target (thus incurring costs of intervention, such as loss of reserves and debt increases), since, in those moments, price setters do not answer with the same intensity as exchange rate movements.

VI – CONCLUSIONS

The analysis of results presented in sections IV.3 and IV.4 show that the use of unconditional variances leads us to results that are sensitive to the chosen measure of volatility, which is based on subjective criteria.

Moreover, the significance of lagged terms or positive effects of shocks in a variable over variances shows that it is appropriate to work with GARCH-type models also to analyze the effects on inflation variance.

By testing the model with a bivariate GARCH model, we find that there is a relation between exchange rate and inflation variances and that this relation is semi-concave, differently from what was estimated for financial series and in line with the intuition obtained from other studies. Besides, we deal directly with the effects of conditional volatilities, which have not been explored by the literature so far.

The caveats of this paper basically lie in the small sample available for Brazil, with the floating regime for exchange rates having started only in 1999. Because of that, we cannot establish with certainty whether the problems with convergence encountered were due to the sign instability or to the small period involved. Nonetheless, we tend not to rely too much in the small sample explanation, since three out of the other four restrictions tested – diagonal VEC, CCORR and DCC – have less parameters to be estimated. Furthermore, the small sample size prevented us from testing interesting variations of the problem, such as a GARCH-M formulation. Thus, a large sample is essential to the extensions we plan.

However, this article innovates by (i) applying a multivariate GARCH model, thus, considering conditional variances to analyze the relation between volatilities, (ii) trying to establish a relation between exchange rate and inflation volatilities and its implication for the monetary policy and (iii) showing that traditional tests performed with exogenously constructed volatility series are sensitive to the criteria chosen to construct such series and do not reveal relevant features of that relation.

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