Pass-through of exchange rate shocks in Brazil as a small open economy

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Abstract

An econometric model is used to analyse pass-through from the exchange rate to inflation, considering the monetary policy instrument variable, the target variable, the level of economic activity, the level of credit and the exchange rate. This system is exogenously affected by commodity prices, the level of external sector activity and the uncertainty perceived by international investors. Using vector error correction models, we find greater pass-through to administered prices than to free prices and a marked asymmetry characterized by stronger exchange rate pass-through when the domestic currency appreciates (deflationary effect) than when it depreciates (inflationary effect). Moreover, we note that the asymmetry in pass-through to domestic prices is due to the behaviour of free prices, since asymmetry in administered prices is not very significant.

Keywords

Monetary policy, foreign exchange rates, prices, inflation, measurement, mathematical analysis, Brazil

JEL classification

E31, E52, F31

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I. Introduction

The exchange rate is one of the main prices in the economy. Given a certain level of domestic and external prices, it is the exchange rate that regulates the international purchasing power of domestic currency-denominated rents and the purchasing power of foreign currency rents over goods denominated in domestic currency, so that the competitiveness of domestic industry, the cost of imported inputs and the ability of the domestic economy to absorb global consumer goods are determined by the exchange rate, at least in the short run. It is thus of interest to both academics and monetary policy practitioners to know how exchange rate movements affect the pass-through mechanisms whereby the base interest rate influences its target variable, namely the rate of inflation for final consumption goods (Brun-Aguerre, Fuertes and Phylaktis, 2012).

The literature defines exchange rate pass-through as the percentage change in domestic prices given a 1% change in the exchange rate, i.e., it is the exchange rate elasticity of domestic prices (Campa and Goldberg, 2005; Assis, Cerqueira and Feijó, 2019). However, it should be noted that in the vector autoregressive (VAR) model methodology and variants such as the vector error correction (VEC) and structural vector autoregressive (SVAR) model methodologies, exchange rate pass-through is measured by means of impulse response functions, which, by construction, measure the elasticity of domestic prices to exchange rate shocks not anticipated by the model, i.e., the reaction of domestic prices to exchange rate shocks as such and not the partial reaction of domestic prices to any change in the exchange rate as measured by other methodologies.

This paper uses different VEC specifications to measure exchange rate pass-through in the Brazilian economy over recent years (2003–2016).[1] The VAR model is very frequently used in the literature on exchange rate pass-through (Belaisch, 2003; Minella and others, 2003; Nogueira, 2007; Souza and Alves, 2010; Pimentel, Luporini and Modenesi, 2016),[2] but we find that little consideration is given to testing for cointegration between variables, since if one or more cointegrating relationships are detected, the system being examined should be represented by a VEC. Moreover, non-inclusion of the error correction term in the VAR in differences may lead to autocorrelations in the residuals (ultimately leading to unnecessary lags being added to deal with the problem) and to biased forecasts, since a variable that is important in describing the dynamics of the system is being omitted.

This paper differs from the rest of the relevant literature on the Brazilian economy in that it gives due weight to cointegration tests and handles pass-through from the exchange rate to prices with a model that treats external shocks as unmodelled (exogenous) variables, given that commodity prices, for example, are set in international markets and depend on variables other than those endogenous to the macroeconomic system that is important for monetary policy. This system is described by the policy instrument variable, the policy target variable, the nominal exchange rate, the level of credit and the level of domestic activity. Therefore, it does not seem reasonable to treat as endogenous a variable that depends on a number of factors external to the system. Moreover, we consider a wider range of shocks that may interfere with measurement of the domestic price response to exchange rate shocks: in addition to international commodity prices, we include as exogenous variables the level of external sector activity and international investors' perception of short-term uncertainty. We believe that we thus better distil genuinely unexpected exchange rate shocks and their impact on prices. We have based our choice of variables on a review of the Brazilian and international literature,[3] filling in gaps that we

[1] The era following the fixed exchange rate regime and the instability caused by the 2002 elections.

[2] These studies obtain their main results from vector autoregressive (VAR) models. The Brazilian literature, curiously, relegates to the background the possibility of cointegration between variables endogenous to its estimation models.

[3] Particular mention may be made of Belaisch (2003), Minella and others (2003), Nogueira (2007), Capistrán, Ibarra and Francia (2012), Burstein and Gopinath (2014), Donayre and Panovska (2016), Pimentel, Luporini and Modenesi (2016), Brun-Aguerre, Fuertes and Greenwood-Nimmo (2017) and Bejarano Aragón, Moura and Moura (2018).

consider significant, and on the Central Bank of Brazil's publications dealing with its own VAR and VEC models (Central Bank of Brazil, 2012b and 2013).

The paper shows that there is more pass-through to administered prices than to free prices. There is also evidence of asymmetric pass-through[4] in the full version of the extended consumer price index (IPCA) and in the free price category within the IPCA. This asymmetry is characterized by stronger pass-through of domestic currency appreciation shocks than depreciation shocks: in absolute terms, the long-run effect on prices of an appreciation shock is about twice as large as the effect caused by a depreciation shock. Moreover, it is shown that the pass-through of a domestic currency appreciation to free and administered prices is of similar magnitude, while there is much greater pass-through to administered prices than to free prices when the domestic currency depreciates. This suggests that, in general, market-regulated consumer prices are subject to a competition effect that counteracts the pass-through of positive cost shocks to the final consumer.

In sum, Brazilian monetary policy cannot ignore this important non-linearity (pass-through asymmetry) in the relationship of the exchange rate to domestic prices in Brazil, which deserves further investigation, as well as other possible non-linearities (Carneiro, Monteiro and Wu, 2004; Correa and Minella, 2006) in the variables that are important for monetary policy. The discovery of non-linearities as strong as the asymmetry shown in this paper reveals how urgent it is to review the use of purely linear models when approximating relationships of relevance to monetary policy.

Section II that follows reviews the empirical literature of recent years. Section III presents our own empirical analysis. Section IV, lastly, draws conclusions.

II. Exchange rate pass-through in Brazil over recent years

In an emerging economy like Brazil's, with a highly volatile exchange rate (Minella and others, 2003; Kaltenbrunner and Painceira, 2015) and its own foreign exchange market idiosyncrasies,[5] a proper understanding of the relationship between exchange rate movements and final consumer prices is essential for the design of an appropriate monetary policy that can help stabilize inflation (Minella and others, 2003; Mishkin, 2004; Nordstrom and others, 2009; Ghosh, Ostry and Chamon, 2016). The autoregressive vector methodology and its variants are very widely employed in the literature to measure exchange rate pass-through (Jacobson and others, 1999; Central Bank of Brazil, 2012b), and their use is interesting, as they provide an analysis with minimal researcher interference and with a priori application of restrictions derived from the theory (Central Bank of Brazil, 2012b). We shall now review a selection of papers from recent years in which this type of methodology has been used to study the situation in Brazil (see table 1).

[4] See section III.2.

[5] Ventura and Garcia (2012) present the sui generis relationship between the futures market and the spot market in Brazil, where they find evidence that the short-term futures market (transactions for foreign currency delivery within 30 days) is the locus of price formation, with prices transmitted to the spot market through arbitrage transactions.

Table 1
Summary of studies using the VAR, SVAR and VEC methods
to measure exchange rate pass-through without considering asymmetry

Authors	Model	Pass-through measuring method	Sample	Cointegration testing	Twelve-month pass-through to the IPCA *(Percentages)*
Belaisch (2003)	VAR	$\frac{\Delta P_{t,t+j}}{\Delta E_{t,t+j}} \times 100$	June 1999 to December 2002	No	17.00
Minella and others (2003)	VAR	$\frac{\Delta P_{t,t+j}}{\varepsilon_{Et}} \times 100$	September 1994 to December 2002	No[a]	17.90
	VAR		September 1994 to June 2002	No[a]	14.10
Nogueira (2007)	VAR	$\frac{\Delta P_{t,t+j}}{\varepsilon_{Et}} \times 100$	January 1995 to June 1999	No	131.00
	VAR		July 1999 to December 2004	No	11.00
Souza and Alves (2010)	VEC	$\frac{\Delta P_{t,t+j}}{\Delta E_{t,t+j}} \times 100$	January 1999 to December 2002	Yes	12.57
	SVAR		January 2003 to December 2009	Yes[b]	1.78
	VAR		January 2003 to December 2009	Yes	1.53

Source: Prepared by the authors, on the basis of A. Belaisch, "Exchange Rate Pass-Through in Brazil", *IMF Working Papers*, No. 2003/141, Washington, D.C., International Monetary Fund (IMF), 2003; A. Minella and others, "Inflation targeting in Brazil: constructing credibility under exchange rate volatility", *Journal of International Money and Finance*, vol. 22, No. 7, December 2003; R. P. Nogueira, "Inflation targeting and exchange rate pass-through", *Economia Aplicada*, vol. 11, No. 2, June 2007; R. G. Souza and A. F. Alves, "Relação entre câmbio e preços no Brasil: aspectos teóricos e evidências empíricas", thirty-eighth national meeting, Brazilian Association of Graduate Programmes in Economics (ANPEC), Salvador, 7–10 December 2010 [online] https://www.anpec.org.br/encontro2010/inscricao/arquivos/000-63cb7b0661b466d2e5760e4a82f422dc.pdf [accessed on 5 January 2023].

Note: VAR: vector autoregressive model; SVAR: structural vector autoregressive model; VEC: vector error correction model; IPCA: extended consumer price index.

[a] The method used by Belaisch (2003) and all those who follow her is the ratio of the changes in the price level and the exchange rate accumulated from t until $t+j$, given a shock at t. The other method considers only the relationship between the cumulative change in the price level and exchange rate shocks not anticipated by the model.

[b] No cointegrating relationship was identified for this sample period.

Belaisch (2003) measures the exchange rate pass-through ratio for $t=j$ after a given shock of $t=1$ as the ratio between the cumulative change in the price level and the cumulative change in the exchange rate, both up until $t=j$.

Belaisch (2003) estimates a VAR(2) for the period from July 1999 to December 2002, including the oil price index (IMF, 2022), the physical production index for industry (the old Monthly Industrial Survey-Physical Production (PIM-PF) prepared by the Brazilian Institute of Geography and Statistics (IBGE)), the exchange rate, and one of the following price indices, for each case: the general price index-domestic supply (IGP-DI), the wholesale price index (IPA), the IPCA, the tradable goods IPCA, the non-tradable goods IPCA, the administered prices IPCA and the free prices IPCA.[6] The variables were used in first difference, as the unit root tests indicated that they were all I(1).[7] The impulse response functions were calculated from the orthogonalized residuals using the Cholesky decomposition. The author finds that exchange rate pass-through to the IPA is much more persistent and faster than that to the IPCA. The IGP-DI also responds quickly, but the impact is smaller and shorter-lived than in the case of the IPA. Among consumer prices, tradable goods prices are the most sensitive to exchange rate shocks. Free prices show similar behaviour, but less marked. The impact on non-tradable goods prices is small, but persists for 12 months, which is consistent with the idea that these experience second-order shocks after the exchange rate shocks (the direct inflationary impact on tradable goods

[6] The first two (the IGP-DI and the IPA) are produced by the Brazilian Institute of Economics (FGV IBRE, 2022) and the other five (the IPCA and its subindices) by the Brazilian Institute of Geography and Statistics (IBGE, 2022).

[7] Integrated of order 1.

prices is transmitted with some lag to non-tradable goods prices).[8] Administered prices react faster than free prices, but the impact on the former is less persistent and loses statistical significance after one quarter. Belaisch (2003) identifies a 12-month pass-through of 53% for the IGP, 120% for the IPA, 17% for the full IPCA, 15% for both free and tradable goods prices, 5% for administered prices and 12% for non-tradable goods prices.

This greater impact of exchange rate shocks on the IPA than the IPCA challenges the notion that the strength of exchange rate pass-through diminishes along the production chain (McCarthy, 2000; Choudhri, Faruqee and Hakura, 2005), as identified by Capistrán, Ibarra and Francia (2012) in the Mexican case in their observation of pass-through to imported product prices, producer prices and consumer prices.

Minella and others (2003) estimate two specifications of the VAR for the period from September 1994 to December 2002. In both specifications, the authors include seasonally adjusted industrial production as a proxy for the level of activity, the emerging markets bond index plus (EMBI+)[9] as a way of controlling for the financial crises in the period (the Mexican, South-East Asian, Russian, Argentine and Brazilian crises), the exchange rate of the Brazilian real against the dollar and the monthly average of the effective Over/SELIC rate. In one, prices were measured by the IPCA, and in the other, the IPCA was broken down into administered and market (free) prices.

Unlike Belaisch (2003), Minella and others (2003) estimate a larger exchange rate pass-through for administered prices than for free ones, but the measurement method differs: these authors estimate the exchange rate pass-through as the ratio between the cumulative 12-month change in the price level and the value of the exchange rate shock in the first of those 12 months. Thus, they estimate exchange rate pass-through of 32.7% for administered prices and 17% for free prices. If the second half of 2002 is removed from the sample, exchange rate pass-through is 19.7% for administered prices and 7.8% for free prices (the drop in the expected values is considerable, but, taking a 95% confidence level, there is no statistical difference between the two estimates in the pass-through values for each price category). For the full IPCA, pass-through was 17.9% in the full sample and 14.1% in the reduced sample. When considering only the inflation targeting period (July 1999 to December 2002), exchange rate pass-through is lower: 20% for administered prices, 11.3% for free prices and 13.1% for the full IPCA. However, the values for administered and free prices are not statistically significant in the period of the inflation targeting regime, probably owing to the paucity of observations.

Nogueira (2007) analyses exchange rate pass-through and the "fear of floating"[10] in a number of developed economies (Canada, Sweden and the United Kingdom) and emerging ones (Brazil, Czechia, Mexico, the Republic of Korea and South Africa) that have adopted an inflation targeting regime. Following the methodologies used by Calvo and Reinhart (2002) and Ball and Reyes (2008), the author identifies a reduction in exchange rate intervention in all countries after the adoption of the inflation targeting regime. The question is whether exchange rate intervention is driven by a fear of inflation or a more generalized fear of floating. The monetary authority of a country adopting a targeting regime would have an incentive to try to influence the exchange rate if it identified high pass-through of the exchange rate to prices in the economy, so that fear of inflation may translate into a fear of floating, but the latter would be characterized by attention to exchange rate movements that was inconsistent with the monetary policy regime adopted in the economy (i.e., there would be exchange rate interventions beyond what was needed to meet policymakers' publicly stated objectives).

[8] Some non-tradable goods may have tradable goods as inputs, or there is simply a tendency for relative prices to readjust.

[9] Specifically, the authors used the EMBI for the period from September 1994 to December 1998 and the EMBI+ for the rest of the sample.

[10] The concept is taken from Calvo and Reinhart (2002) and refers to interventions in the foreign exchange market by the monetary authority of an economy that appear inconsistent with announced monetary policy objectives, even when the adoption of a floating exchange rate regime has been officially declared.

With these perspectives, Nogueira (2007) analyses macroeconomic data (international reserves, interest rates, exchange rates of national currencies per dollar unit, inflation rates, output growth rates) for developed economies between January 1985 and December 2004 and for emerging economies between January 1995 and December 2004 by analysing the variability of the data before and after the adoption of the inflation targeting regime and using the VAR method to measure exchange rate pass-through to prices. The author finds that exchange rate pass-through is lower in developed economies than in emerging ones and that it falls significantly in both after the adoption of the inflation targeting regime. Through variance decomposition, however, he finds that the exchange rate remains an important determinant of the inflation rate, even after the adoption of the targeting regime, which would seem to justify the reactions of most central banks in the economies analysed to exchange rate movements, taking the form of interventions that are intended to avoid exchange rate pass-through spilling over to consumer prices and are therefore consistent with the targeting regime, being induced by the fear of inflation. In the Brazilian case in particular, the author finds that, given a 1% exchange rate shock at $t = 1$, there is a 1.3% change in the general price level at $t = 12$ before the adoption of the inflation targeting regime. After the adoption of this regime, the cumulative 12-month impact on the general price level is only 0.11%. In other words, the cumulative 12-month pass-through ratio of exchange rate shocks decreases from 1.3% before the adoption of the targeting regime to 0.11% afterwards. Even so, it should not be lightly assumed that the adoption of the regime was the determining factor in this decline in exchange rate pass-through, since, except for the first half of 1999, the period before the targeting regime coincides with the exchange rate peg period, so that the sharp reduction in pass-through may be partly explained by the adoption of a floating exchange rate (Albuquerque and Portugal, 2005).[11]

An exception in the Brazilian literature as regards the concern with testing for cointegration between variables of interest is Souza and Alves (2010), but in our view these authors do not consider enough variables, do not address exogenous shocks as we do, do not deal with the issue of asymmetry and do not show whether the choice of lags for estimating the models eliminates possible autocorrelation of the residuals, which if present casts doubt on the estimates in the context of these models. Souza and Alves (2010) investigate exchange rate pass-through in the period from 1999 to 2009 and, identifying a structural break in the exchange rate in January 2003, split the analysis into two samples, finding evidence of a cointegrating vector between the IPCA, the exchange rate, the Institute for Applied Economic Research (IPEA) industrial production index and the oil price index,[12] also provided by IPEA, in the period from January 1999 to December 2002, with no cointegrating relationship between the variables in the rest of the sample. Thus, estimates are made using the VEC method for the first period and the VAR and SVAR methods for the second. The methodology for measuring exchange rate pass-through is the same as that used by Belaisch (2003). The authors find strong evidence of a reduction in exchange rate pass-through between the two periods, with an estimated 12-month pass-through of 12.57% in the period from January 1999 to December 2002 and an estimated 12-month pass-through of 1.53% using VAR and 1.78% using SVAR in the period from January 2003 to December 2009.

Pimentel, Luporini and Modenesi (2016) introduce the possibility of asymmetric exchange rate pass-through, i.e., impacts of different magnitudes on prices following an exchange rate appreciation or depreciation. For the period from January 1999 to November 2013, the authors estimate that the inflationary impact of an exchange rate depreciation is significantly larger than the deflationary impact of an exchange rate appreciation. They use the IPCA as a measure of the general price level; the monthly index of physical industrial production (PIM-PF) calculated by IBGE as a proxy for the level of activity stimulated by aggregate demand; the monthly average buying exchange rate between the Brazilian real and the dollar; and the commodity price index calculated by IPEA and the dollar-denominated

[11] Having access to more data on the inflation targeting regime, however, the Central Bank of Brazil (2011) reports a decline in exchange rate pass-through during the period this regime was in force.

[12] Endogenous model variables.

international oil price published by the International Monetary Fund (IMF) as proxies, used in different specifications, for cost behaviour (aggregate supply shocks). With different specifications, and without including asymmetry in exchange rate pass-through, the authors estimate for the floating exchange rate period up to May 2012 that, given an exchange rate shock at t, between 5.9% and 7.2% of the exchange rate change caused by that shock is passed through to prices in the 12-month period starting at t. When asymmetry is included, an unexpected 1% appreciation in the exchange rate at $t = 1$ leads to a 0.024% change in the price level at $t = 12$. This is surprising, because exchange rate appreciation would be expected to have a deflationary impact. In the case of a 1% exchange rate depreciation, the positive change in the price level at $t = 12$ is between 0.056% and 0.077%.

III. Brazil as a small open economy

1. Some methodological considerations

When variables of relevance to monetary policy in an economy are analysed, it is usual for them to have a unit root, i.e., these are variables that have a stochastic trend and are thus non-stationary. Examples generally include monetary aggregates, the general price level, the exchange rate, indirect indicators of the level of activity and sometimes even the base rate. For this reason, when estimating the interrelationships between these variables in a VAR, it is usual to work with the first differences of the series or with the first differences of the logarithms of the series.

When a set of integrated series of order 1 (I(1)) has a statistically stable long-run relationship, these variables are said to be cointegrated (Banerjee and others, 1993, chap. 5). Every cointegrating relationship is represented by an error correction model (Engle and Granger, 1992; Banerjee and others, 1993, chaps. 5 and 8). This name comes from the fact that the one-period lagged deviation of the long-run relationship enters the equation describing the short-run dynamics of at least one of the variables. That is, when z_t, I(0), with zero mathematical expectation, is a linear combination of n I(1) variables, the first difference at t of at least one of the n endogenous variables in the model is significantly affected by the deviation of the long-run relationship between the n variables at $t - 1$. Thus, a first-difference VAR model is only suitable for measuring exchange rate pass-through if we eliminate the possibility of there being an error correction term relating to a long-run relationship between the endogenous variables.

Moreover, we do not consider it reasonable to treat variables such as international commodity or oil prices, which are commonly used as proxies for inflationary cost or supply shocks, as endogenous to the economic system that is relevant to domestic monetary policy. This system is basically described by five main variables: the monetary policy instrument variable, the monetary policy target variable, the level of activity in the domestic economy, credit (elasticity of domestic liquidity)[13] and, of course, the exchange rate, which is the factor of interest for this paper. These variables coincide with those used by the Central Bank of Brazil in its vector autoregressive models and with the monetary policy transmission channels it considers (Central Bank of Brazil, 2012a, 2012b and 2013).

Furthermore, in order to correctly isolate the effects of exchange rate shocks unanticipated by the model on domestic prices, it was decided to include three variables exogenous to the endogenous domestic monetary policy system: an international commodity price index, an index representing the economic activity of the external sector, and a measure of international investors' perception of uncertainty, since at times of greater uncertainty in financial markets investors seek out safer assets, preferably

[13] In other words, the ability to expand liquidity coupled with the willingness to lend, given the banking system's reserves in cash and securities that are immediately convertible to goods and assets at face value.

denominated in currencies that rank higher in the hierarchy of the international monetary system,[14] which also leads to reactions in base interest rates to forestall very large and abrupt depreciations of the domestic currency; in other words, it is an important predictive variable for the system we are concerned with in this paper. Brazil is deemed here to be too small to affect these variables, i.e., they exogenously impact the system described by the five endogenous variables mentioned above, these being variables that are very much affected in the sphere of the domestic economy and, more specifically, that of domestic monetary policy.

Capistrán, Ibarra and Francia (2012) formulate the same consideration in the case of the Mexican economy with respect to variables in the United States economy that affect it exogenously. They use these variables, much as those in this article are used, as proxies for Mexican external sector variables, which are insignificantly affected by the Mexican economy. Given the size of the Brazilian economy and its participation in world trade and capital flows, it is reasonable to treat it as a small open economy in the sphere of analysis we are concerned with.

Lastly, our measure of exchange rate pass-through, as befits the methodology used, will start from the analysis of impulse response functions, which we will derive using the methodology of generalized impulse response functions, which are robust to the ordering of variables (Pesaran and Shin, 1998).

Let us take a VAR(p) with exogenous variables as follows:

$$Y_t = \sum_{i-1}^{p} \Gamma_i Y_{t-i} + \Upsilon D_t + \Phi X_t + \varepsilon_t \tag{1}$$

Taking the standard assumptions of stability of the VAR(p) process, no error autocorrelation (but without restricting the possibility of contemporaneous correlation between the components of ε_t), independence between the errors and the stochastic and deterministic regressors, and non-perfect multicollinearity between Y_t,..., Y_{t-p}, X_t and D_t, the stochastic process $\{Y_t\}_{t=-\infty}^{+\infty}$ can be represented as:

$$Y_t = \psi(L)D_t + \Xi(L)X_t + \Theta(L)\varepsilon_t \tag{2}$$

where $\psi(L)$, $\Xi(L)X_t$ and $\Theta(L)$ are infinite-order matrix polynomials.

Taking Ω_{t-1} as the known history of the economy up to $t-1$, we can write:

$$E(Y_{t+k} \mid \Omega_{t-1}) = \psi(L)D_{t+k} + \sum\nolimits_{i=1}^{\infty} \Xi_{i+k}X_{t-i} + E\left(\sum\nolimits_{h=0}^{k} \Xi_{k-h}X_{t+h} \mid \Omega_{t-1}\right) + \sum\nolimits_{i=1}^{\infty} \Theta_{k+i}\varepsilon_{t-i} \tag{3}$$

If we consider that the economy is affected at t by a vector $\delta' = (\delta_1,..., \delta_n)$ of shocks to the n variables of Y_t, we get:

$$E(Y_{t+k} \mid \varepsilon_t = \delta, \Omega_{t-1}) = \psi(L)D_{t+k} + \sum\nolimits_{i=1}^{\infty} \Xi_{i+k}X_t + E\left(\sum\nolimits_{h=0}^{k} \Xi_{k-h}X_{t+h} \mid \Omega_{t-1}\right) + \Theta_k\delta + \sum\nolimits_{i=1}^{\infty} \Theta_{k+i}\varepsilon_{t-i} \tag{4}$$

[14] Currencies with a wider currency area, such as the dollar, which is accepted worldwide, the euro, which is accepted throughout Europe and in countries that have substantial trade relations with Europe, and the Japanese yen, which is also accepted as a means of payment in transactions between residents and non-residents in a number of economies that trade with Japan. These are the main currencies in the currency hierarchy, whose ranking criterion is the international liquidity of the financial asset. The Brazilian real clearly stands lower in the international currency hierarchy than the currencies listed above.

With the shock occurring at the given t, we are interested in the time profile of the response of each variable in a system to an unexpected shock due to the dynamics of the system itself. The impulse response function of each variable is a measure of this profile. So given δ and Ω_{t-1}, we obtain a function that for each value of k returns the response Y_{t+k} to a shock at t. Pesaran and Shin (1998) call this function the generalized impulse response function, and we represent it as:

$$GI(K, \delta, \Omega_{t-1}) = E(Y_{t+k}|\varepsilon_t = \delta, \Omega_{t-1}) - E(Y_{t+k}|\Omega_{t-1}) = \Theta_k\delta \tag{5}$$

Considering that at t there is only one shock to variable j and supposing that the errors of the autoregressive form of Y_t follow a multivariate normal distribution,[15] $\varepsilon_t \sim N(0, \Sigma)$, we can write:

$$E\left(\varepsilon_t \mid \varepsilon_{jt} = \delta_j\right) = \left(\sigma_{1j}, \sigma_{2j}, \ldots, \sigma_{mj}\right)' \sigma_{jj}^{-1} \delta_j = \Sigma \varsigma_j \sigma_{jj}^{-1} \delta_j \tag{6}$$

where $(\sigma_{1j}, \sigma_{2j}, \ldots, \sigma_{mj})'$ is the vector of contemporaneous covariances between the errors of each equation of the system, σ_{jj} is the variance of the j-th element of ε_t and ς_j is a vector of zeros, except the number 1 in its j-th entry.

Scaling the shock δ_j to one standard deviation of ε_{jt}, Pesaran and Shin (1998) define the generalized impulse response function as:

$$GI_j\left(k\right) = \sigma_{jj}^{-1/2} \Theta_k \Sigma \varsigma_j \tag{7}$$

Let us now assume that the n endogenous variables of Y_t are I(1), but are cointegrated, so that we observe $Y_t \sim CI(1,1)$ (Engle and Granger, 1992). We then represent the model in its error-corrected form:

$$\Delta Y_t = \alpha\beta' Y_{t-1} + \sum_{i=1}^{p-1} C_i \Delta Y_{t-i} + BX_t + \alpha\beta' HD_t + \in_t \tag{8}$$

where D_t is a vector of deterministic components that may be present in the cointegrating relationship (so that $\alpha\beta' Y_{t-1}$ is a zero mean stationary process) and B, H and C_i, with $i = 1, \ldots, p-1$, are coefficient matrices.

Following Pesaran and Shin (1998), given that the first difference of Y_t is $I(0)$, let $C = I_n - \sum_{i=1}^{p-1} C_i$ and $\alpha\perp$ and $\beta\perp$, such that $\alpha'\alpha\perp = \beta'\beta\perp = 0$, as long as $\alpha\perp C\beta\perp$ is the full range, we get:

$$\Delta Y_t = \sum_{i=0}^{\infty} Z_i X_{t-i} + \sum_{i=0}^{\infty} A_i \alpha\beta' D_{t-i} + \sum_{i=0}^{\infty} K_i \in_{t-i} \tag{9}$$

where Z_i, A_i and K_i, for $i = 0, 1, \ldots$, coefficient matrices.

Let $\sum_\epsilon$ be the covariance matrix of ϵ_t and $\psi_k = \sum_{j=0}^{k} K_j$ with $\psi_0 = I_n$, then the generalized impulse response function for endogenous level variables, scaled to a unit shock, is:

$$GI_j\left(k\right) = \sigma_{jj}^{-1} \psi_k \sum\nolimits_\epsilon \zeta_j \tag{10}$$

[15] Pesaran and Shin (1998) point out in their second footnote that "when the distribution of the errors ε_t are non-normal, one could obtain the conditional expectations $E(\varepsilon_t|\varepsilon_{jt} = \delta_j)$ by stochastic simulations, or by resampling techniques if the distribution of errors is not known".

In a model using the natural logarithms of the nominal exchange rate and the level of consumer prices, the impulse response function of the price level logarithm in the event of a unit exchange rate shock provides exactly one measure of exchange rate pass-through, as it represents the elasticity of the price level in the event of an exchange rate shock.

As an extension of the original model, the effects on free and administered prices and the presence of asymmetry in exchange rate pass-through will also be separately analysed. The free and administered price index series were constructed by setting the July 2003 index value at 100 and accumulating the percentage changes for each price category. We shall now explain the methodology used to consider the possibility of asymmetry in exchange rate pass-through.

Given *LEXCHANGE+* and *LEXCHANGE-* defined as follows:

$$LEXCHANGE_t^{+} = \sum_{i=1}^{t} \omega^{+} \Delta e_t \,; \begin{cases} \omega^{+} = 1, & \text{if } \Delta e_t > 0 \\ \omega^{+} = 0, & \text{if } \Delta e_t \leq 0 \end{cases} \tag{11}$$

$$LEXCHANGE_t^{-} = \sum_{i=1}^{t} \omega^{-} \Delta e_t \,; \begin{cases} \omega^{-} = 0, & \text{if } \Delta e_t \geq 0 \\ \omega^{-} = 1, & \text{if } \Delta e_t < 0 \end{cases} \tag{12}$$

where Δe_t is the first difference of the natural logarithm of the exchange rate.

To carry out the empirical analysis presented in the next subsection, the following steps were taken: (i) selection of the monetary policy-related variables serving to model the system that concerns us, based on the literature and the models used by the Central Bank of Brazil; (ii) execution of unit root tests to determine the order of integration of each of the variables; (iii) determination of the number of lags for a VAR with variables in level; (iv) application of the cointegration test; (v) preliminary VEC and VAR estimations and observation of the generalized impulse response functions; (vi) for each VEC, testing for the significance of the coefficients of each variable in the cointegrating relationship and also weak exogeneity; and (vii) final specifications. To determine the number of lags, lastly, we used two methodologies that were equivalent in all the models presented in this paper: the maximum lag was set at eight and the Lagrange multiplier test for autocorrelation was used to find the smallest VAR with non-autocorrelated residuals. Following the Central Bank of Brazil (2012b), we used the Hannan-Quinn information criterion to determine the size of the VAR and added lags as necessary to eliminate autocorrelation in the residuals of the LM test. These methodologies were applied to the VAR in levels, so that the VECs have one lag less.[16]

2. Why is the exchange rate passed through asymmetrically?

Asymmetry in exchange rate pass-through means that aggregate prices react differently depending on the direction of the exchange rate shock. Asymmetric import prices in the domestic economy can be explained by the market power of firms exporting to it, which are able to maintain their foreign currency margins in the event of an exchange rate depreciation and to profit when there is an appreciation. Other possible explanations include constraints on the distribution networks of firms exporting to the domestic economy or export limitations in the countries from which the domestic economy imports, so that higher demand for imported goods resulting from the increased international purchasing power of the domestic currency in periods of appreciation puts upward pressure on the foreign currency price of imported goods, thereby reducing the impact of the exchange rate appreciation (Brun-Aguerre,

[16] To ensure parsimony in the size of the VAR, we only rejected the null hypothesis of no autocorrelation when the p-value was less than 0.01.

Fuertes and Greenwood-Nimmo, 2017). Another possible cause of pass-through asymmetry arises when a firm faces reasonable competition in the international market and fights for market share in foreign markets. In the event that an importing country's currency appreciates, the firm maintains its price in its own currency and allows full exchange rate pass-through to the price in the importing country's currency. In the event of a depreciation, the exporting firm, in order to maintain market share, absorbs some of the loss of purchasing power of the importer's currency by cutting the price in its own currency, thereby reducing the pass-through to the price paid by the importer (Pimentel, 2013, pp. 36–37; see Pimentel, 2013; Pimentel, Luporini and Modenesi, 2016; Brun-Aguerre, Fuertes and Greenwood-Nimmo, 2017; Assis, 2017, chap. 1; and Assis, Cerqueira and Feijó, 2019 for more detailed theoretical discussions on this topic).[17]

In the international literature, the issue has been investigated in studies such as those by Pollard and Coughlin (2003), who observe asymmetric short-term exchange rate pass-through in half the 30 industrial sectors analysed in the United States between 1978 and 2000, but find that no one type of asymmetry predominates; Bussière (2013), who in a study of the Group of Seven (G7) countries between 1980 and 2006 finds a tendency towards greater pass-through of local currency depreciations than appreciations in the short run; and Delatte and López-Villavicencio (2012), who study quarterly data between 1980 and 2009 for Japan, Germany, the United States and the United Kingdom and find greater pass-through of depreciations in the long run. In the Brazilian case, the literature on the asymmetry of exchange rate pass-through is still quite sparse. We highlight here the studies by Pimentel, Luporini and Modenesi (2016) and Assis (2017), which find asymmetries in opposite directions: the former authors find greater pass-through of depreciations than of appreciations, while Assis finds just the opposite, in addition to observing even stronger asymmetry than was found by Pimentel, Luporini and Modenesi (2016).

3. Empirical analysis

To achieve the objectives of this study by means of the methodology described, we will use the measures presented in table 2 for the theoretical variables that concern us. This table is complemented by the analysis of variables used in our model, included in tables 3, 4 and 5.

Table 2
Variables used

Variable	Measure	Abbreviation	Source
Real base rate	Annualized Over/SELIC interest rate (cumulative for the month) deflated by the percentage change in the monthly general price index (IGP-M) for the current month[a]	INTEREST	Central Bank of Brazil and Brazilian Institute of Economics (FGV IBRE)
Exchange rate	Real exchange rate/average free dollar for the month	EXCHANGE	Series 3697 of the Central Bank of Brazil Time Series Management System (SGS)
Cumulative positive percentage changes in the exchange rate	See section III.1	LEXCHANGE+	
Cumulative negative percentage changes in the exchange rate	See section III.1	LEXCHANGE-	
Activity level	Central Bank of Brazil economic activity index (IBC-Br)	IBC	Central Bank of Brazil
Consumer price index	Extended general consumer price index	IPCA	Brazilian Institute of Geography and Statistics (IBGE)

[17] This is an example of pricing to market, where an exporting firm faced with different market structures in the markets it sells to (one in the domestic market and one in international markets) price-discriminates between these markets, making it a more aggressive competitor in international markets (which leads to the exchange rate pass-through asymmetry described above) and something more akin to an oligopolist in the domestic market (Krugman, 1986).

Variable	Measure	Abbreviation	Source
Credit	M4/monetary base, the latter seasonally adjusted using the moving average method in a multiplicative model	CREDSA	Central Bank of Brazil
Free consumer price index	Cumulative changes in the IPCA for free prices, taking July 2003 as 100	FREE	Constructed from the percentage change series available from the Institute of Applied Economic Research (IPEA), Ipeadata [online] http://www.ipeadata.gov.br/Default.aspx, using the definitions of the Central Bank of Brazil
Administered consumer price index	Cumulative changes in the IPCA for administered prices, taking July 2003 as 100	ADMS	
Commodity prices	Brazil index of commodity prices (IC-BR) converted to dollars	IC	Central Bank of Brazil
External sector activity	United States total industrial production index (2012 average=100)	IND^{US}	Board of Governors of the Federal Reserve System, "Industrial Production: Total Index" [online] https://fred.stlouisfed.org/series/IPB50001N
International investors' perception of uncertainty	Chicago Board Options Exchange (CBOE) Volatility Index (VIX) (monthly average of daily closing values)	VIX	Chicago Board Options Exchange (CBOE), "CBOE Volatility Index: VIX" [online] https://fred.stlouisfed.org/series/VIXCLS[b]

Source: Prepared by the authors.

[a] Consistent with the practice of the Central Bank of Brazil, "Revisão dos modelos de vetores autorregressivos com fundamentação econômica – 2012", *Relatório de Inflação*, vol. 14, No. 3, September 2012.

[b] VIX measures are now available for emerging markets, including specifically for Brazil (this should perhaps be treated as an endogenous variable), but the observed data for these measures do not cover the entire sample period dealt with by this paper.

Table 3
Descriptive statistics for the endogenous variables

	DLOG(IPCA)	DLOG(IBC)	DLOG(EXCHANGE)	INTEREST	LOG(CREDSA)
Mean	0.004944	0.001829	0.000691	5.932555	2.850260
Median	0.004787	-0.003013	-0.005452	5.975032	2.820874
Maximum	0.013114	0.118665	0.188575	31.737390	3.199159
Minimum	-0.002103	-0.063896	-0.070322	-11.410070	2.627624
Standard deviation	0.002616	0.036181	0.036942	8.110612	0.157833

Source: Prepared by the authors.

Table 4
Descriptive statistics for the exogenous variables

	DLOG(IND^{US})	DLOG(IC)	LOG(VIX)
Mean	0.000976	0.003159	2.883016
Median	-0.001173	0.004653	2.815523
Maximum	0.046533	0.081658	4.137396
Minimum	-0.050636	-0.181264	2.381176
Standard deviation	0.017284	0.034099	0.352920

Source: Prepared by the authors.

Table 5
Endogenous variables in the extended models

	LEXCHANGE-	LEXCHANGE+	DLOG(FREE)	DLOG(ADMS)
Mean	2.016269	2.931270	0.005010	0.004829
Median	2.165229	2.884575	0.005000	0.003500
Maximum	2.967498	4.242634	0.012000	0.033600
Minimum	0.885990	2.094315	-0.003500	-0.011100
Standard deviation	0.599040	0.639886	0.002873	0.005353

Source: Prepared by the authors.

We considered that using the real interest rate in the way described above would shed more light on the effectiveness of monetary policy and that it was more in line with the previous result of Assis (2017), who, in exercises similar to those carried out in this paper, found that while the Central Bank of Brazil was more reactive to shocks affecting the inflation expectations expressed in the targeting regime operating in the country, a monetary policy shock was more powerful when it was able to raise the real interest rate at the same time, in line with the literature following the so-called "Taylor rule".

In addition, the credit proxy was the one used by the Central Bank of Brazil (2012b) in several of its models. However, given the strong seasonality of the monetary base in December, when it grows strongly because of banks' need for liquidity and increased demand for paper money from the public, it was necessary to strip out this effect from the monetary base series before including this series in the models, as the seasonal dummies did not adequately reflect the seasonal effect.[18] Thus, the proxy obtained reflects the relationship between the total money supply (on the broadest definition) and the trend, multiplied by a stochastic component, of the monetary base. This reflects the capacity of the banking system, given the behaviour over time of the money supply in its strict official sense, to expand the liquidity available in the economy, i.e., it is a measure of the elasticity of the Brazilian monetary system.[19]

Including commodity prices is necessary, not only because of their historical importance for the country, but also because they are very important for understanding a number of economic phenomena of the last decade and a half in Brazil (Bresser-Pereira, 2009; Finello and Feijó, 2017). The atypical appreciation of commodity prices, which peaked at the height of the subprime crisis (in the second half of 2008), and its resumption after the "eye of the storm" with a new peak in April 2011, led to appreciation of the currencies of a number of commodity exporters, including Brazil, which seems to be linked to what some researchers have called the reprimarization of the Brazilian export model,[20] in opposition to the description of Dutch disease in Bresser-Pereira (2009). Moreover, it led to an unprecedented build-up of international reserves. As a commodity price measure we use the Brazil commodity index (IC-BR), calculated in reais by the Central Bank of Brazil and converted to dollars. Including the commodity index also serves to control for cost shocks arising from the external sector that are not due to exchange rate depreciation, and is a measure of changes in nominal prices in the benchmark foreign currency.

In the area of monetary policy, it is also important to address the impacts caused by movements in the level of activity in the external sector, since this affects the current transactions balance, which is a variable essentially linked to the exchange rate, impacting domestic economic activity and tradable goods prices. Because the United States is one of Brazil's main trading partners, as well as the world's largest economy and issuer of the international reserve currency, we have selected the United States total industrial production index as a proxy for the level of external sector activity.[21]

Initially, we also considered introducing the federal funds rate as an international benchmark interest rate control, but the behaviour of the series[22] raised difficulties that were not compensated for by substantial control advantages, as the interest rate differential is largely explained by changes in the Brazilian interest rate, given that the effective federal funds rate is much more stable than the effective SELIC rate.

As a way to increase exogenous shocks to the Brazilian monetary policy system by introducing information not captured by the external sector activity level or by commodity prices, we included a

[18] To test whether the seasonal dummies used explained the effect, the credit variable was regressed without seasonally adjusting the base vis-à-vis the dummy variables, and the effect was found to be still very strong in the residuals.

[19] In this case, elasticity means the ability of liquidity in the economy to expand and contract in response to public demand for it.

[20] See Oreiro and Feijó (2010).

[21] Industrial production: total index (IPB50001N), index 2012=100, monthly, not seasonally adjusted (see [online] https://fred.stlouisfed.org/series/IPB50001N).

[22] This includes a break since the introduction of the "near-zero" nominal interest rate policy in the post-crisis period and long periods of stability interrupted by periods of increase or decrease. These phenomena can compromise the results of exercises like those conducted in this section.

measure of risk perception in international financial markets. Thus, the VIX index enters the model as a way of controlling for exogenous shocks to financial flows into the Brazilian economy. The VIX index measures the expected short-term volatility of a wide range of exchange-traded securities, which reflects the "mood" of international investors, undoubtedly an important factor in the conduct of domestic monetary policy.

Once the endogenous and exogenous variables had been chosen, we tested for the presence of unit roots in each of the series. The tests used were the augmented Dickey-Fuller (ADF) test, the Elliot, Rothenberg and Stock (ERS) optimal point test and two Ng and Perron (NgP) tests, MZa and MZt. For the ADF, an augmented Dickey-Fuller regression was first fitted for each series to choose the smallest number of lags with non-autocorrelated residuals. Where necessary, centred seasonal dummies were included,[23] as these are constructed in such a way that they do not shift the test distribution, so that the test value[24] can be compared with the critical values calculated in EViews. These dummies are so constructed that they take the value $(P\text{-}1)/P$ in the seasonal period of interest and $\text{-}1/P$ in the rest of the seasonal periods, with P being the number of seasonal periods. For example, if the data are monthly, the January dummy takes the value 11/12 every January and -1/12 in the other months. In the case of the NgP and ERS tests, the quadratic spectral kernel method and automatic Andrews bandwidth selection were used for the spectral estimation. The sample consisted of monthly data from August 2003 to August 2016 (157 observations). The tests were performed with constant only, and with constant and trend for the natural logarithms of the chosen series (except for the INTEREST variable) and constant for their first differences.[25] Table 6 summarizes the results of the tests. The values of interest for the tests of the level series are presented in the annex (see table A1.2).

Table 6
Summary of unit root test results

Series	ADF		NgP				ERS	
	D	D and T	D		D and T		D	D and T
			MZa	MZt	MZa	MZt		
INTEREST	I(0)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
LOG(EXCHANGE)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
LOG(IBC)	I(1)	I(1)	I(1)	I(0)	I(1)	I(1)	I(1)	I(1)
LOG(IPCA)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
LOG(CREDSA)	I(1)	I(1)	I(1)	I(1)	I(0)	I(0)	I(1)	I(1)
LOG(FREE)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
LOG(ADMS)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
LOG(IC)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
LOG(IND^{US})	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
LOG(CPI)[a]	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
LOG(VIX)	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)

Source: Prepared by the authors.
Note: ADF: augmented Dickey-Fuller test; NgP: Ng and Perron tests; ERS: Elliot, Rothenberg and Stock optimal point test; D: drift; T: trend.
[a] Logarithm of the United States Consumer Price Index for all urban consumers, prepared by the Bureau of Labor Statistics of the United States Department of Labor. This variable is used to estimate the real exchange rate adjusted by the variation of domestic and external prices. See [online] https://www.bls.gov/cpi/.

[23] When their inclusion considerably reduced the number of lags needed to eliminate autocorrelation in the residuals. See IHS Global Inc. (2015, p. 942) for a discussion of these dummy variables.

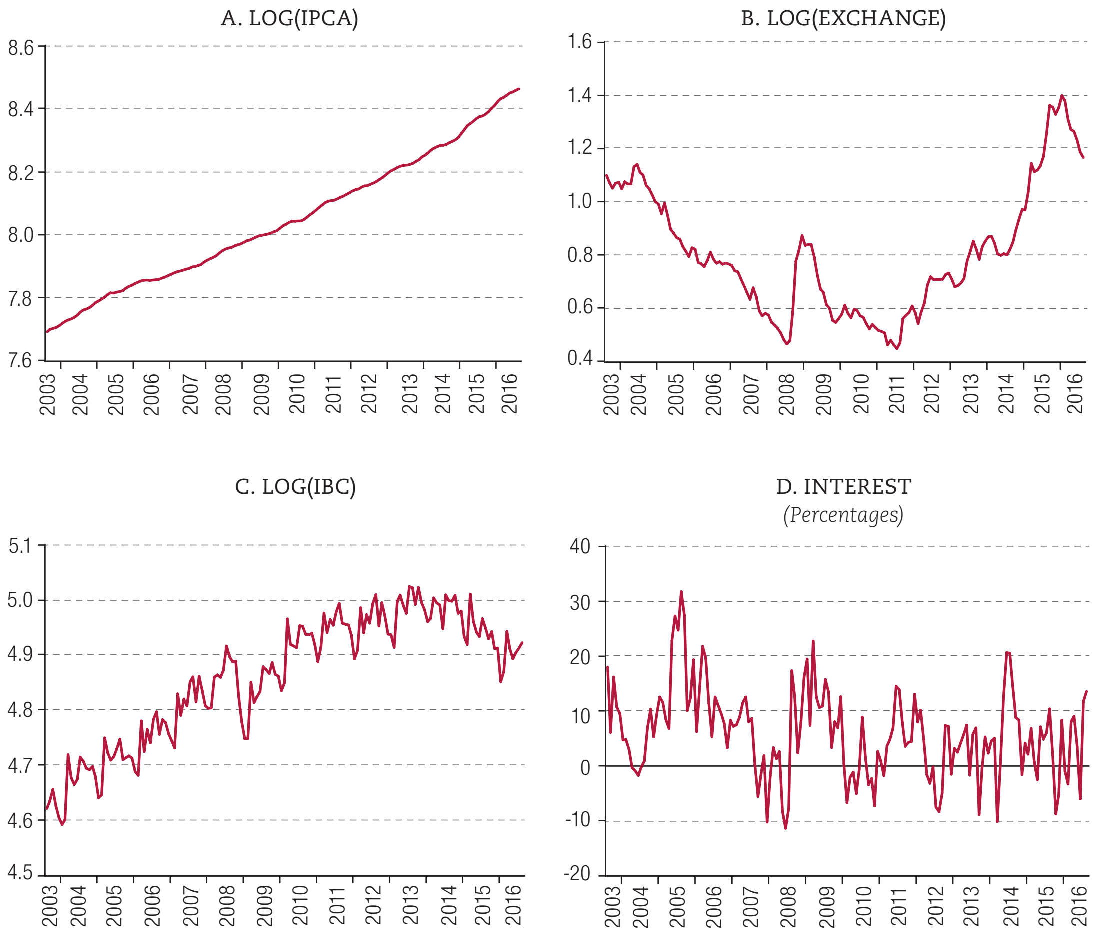
[24] The value corresponding to the t-statistic (which actually follows a variant of the Dickey-Fuller distribution) of the coefficient of the lagged variable in the ADF regression.

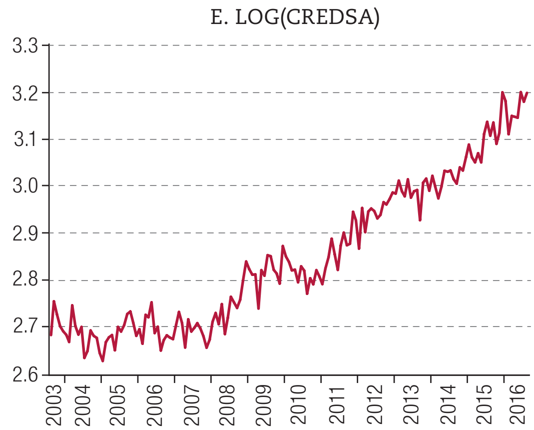
[25] The NgP and ERS optimal point tests do not allow specification to be carried out without a deterministic term. Since at most we consider that the original series had a linear deterministic trend, it is not reasonable to consider the presence of a trend in the first differences.

In the case of the VIX index, employing a union of rejections methodology (Harvey, Leybourne and Taylor, 2011), we reject the unit root for log(VIX), as tests with a constant reject the unit root null hypothesis and tests with a constant and trend do not reject the presence of unit root in the series.

After a series of tests including and excluding variables, we arrived at four models: the first with the full IPCA; the second with the IPCA decomposed into free and administered prices; the third with the full IPCA and the exchange rate decomposed to obtain the pass-through asymmetry; and the fourth with inflation and the exchange rate decomposed. In the first models, three lags were identified in the VAR in levels, and both cointegration tests (Johansen trace and maximum eigenvalue test) pointed to two cointegrating vectors. Their respective VECs were estimated with two lags. When pass-through asymmetry is included, the autocorrelation in the residuals increases considerably, requiring the VAR to be increased so as not to reject non-autocorrelation of the residuals by the Lagrange multiplier (LM) test; including the first difference of log(VIX) in a different way from log(VIX) as we did in the previous two models helped to maintain parsimony, and the VECs were estimated with six lags. With seven lags in the VAR in levels, both tests indicated a cointegrating vector (see figure 1). In all models, the first differences of the logarithms of IC and IND^{US} were taken (see figure 2).

Figure 1
Endogenous variables, 2003–2016



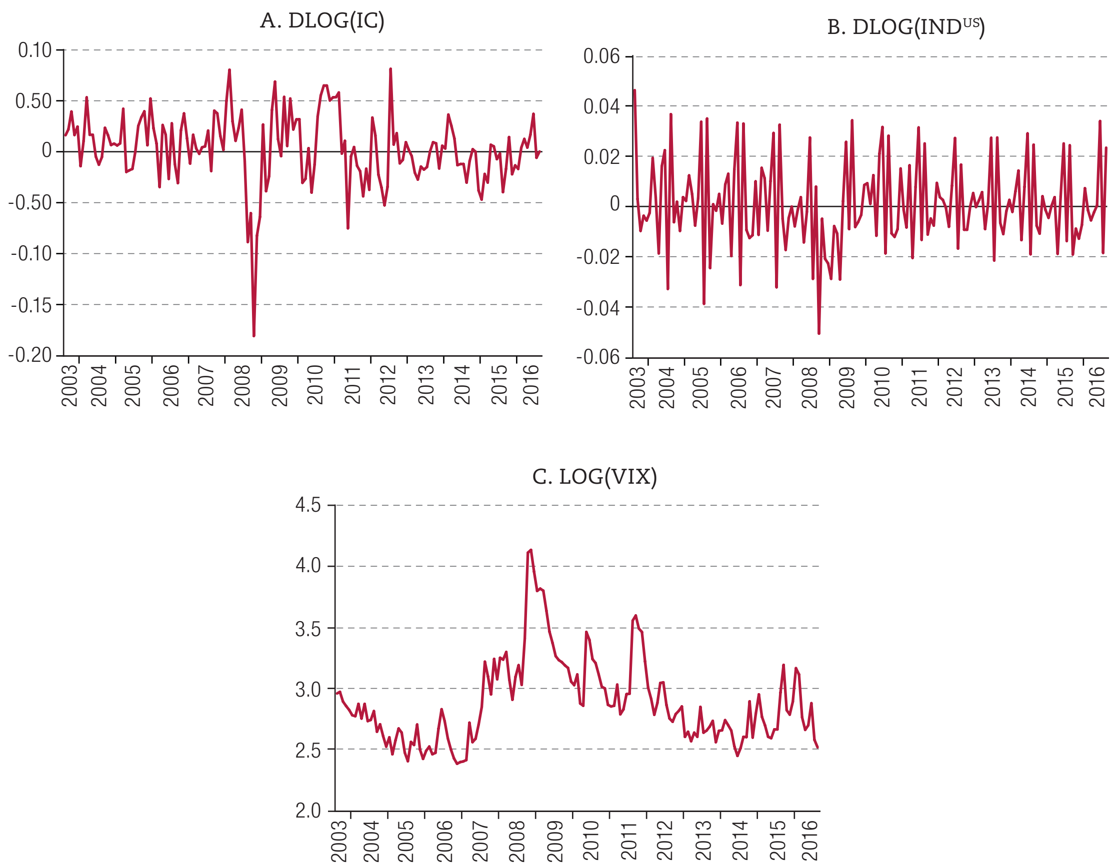


Source: Prepared by the authors.

Note: Measures for the variables: IPCA: extended consumer price index; EXCHANGE: real exchange rate/average free dollar for the month; IBC: Central Bank of Brazil economic activity index (IBC-Br); INTEREST: Annualized Over/SELIC interest rate (cumulative for the month) deflated by the percentage change in the monthly general price index (IGP-M) for the current month; CREDSA: M4/monetary base, the latter seasonally adjusted using the moving average method in a multiplicative model.

Figure 2

Unmodelled variables, 2003–2016



Source: Prepared by the authors.

Note: Measures for the variables: IC: Brazil index of commodity prices (IC-BR) converted to dollars; IND^{US}: United States total industrial production index (2012 average=100); VIX: Chicago Board Options Exchange (CBOE) Volatility Index (VIX) (monthly average of daily closing values).

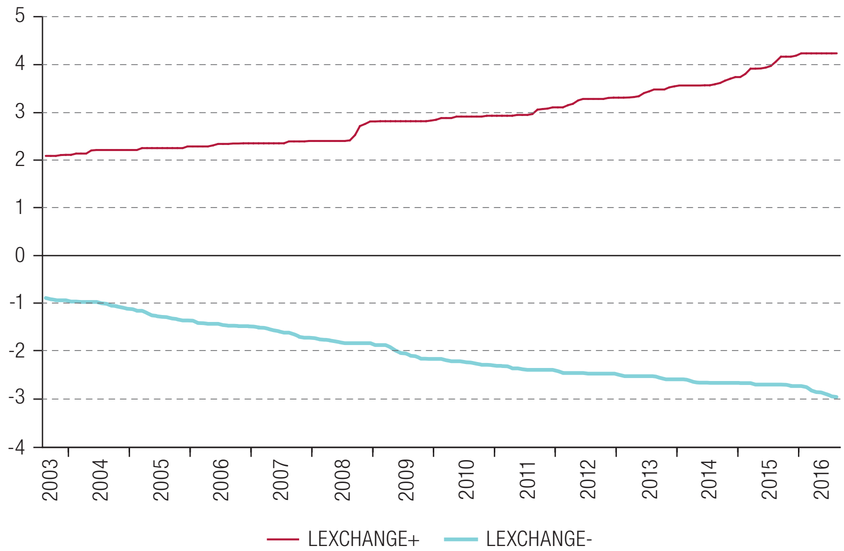
In the first model, the first cointegrating vector was subjected to the restriction that the log(IPCA) coefficient was 1 and the log(IBC) coefficient was 0, while in the second vector the opposite was done, so that both vectors were identified without further restrictions and weak exogeneity of the exchange rate cannot be rejected. In the model with free and administered prices not including pass-through asymmetry, the identification restrictions on the cointegrating vectors were 1 for the log(FREE) coefficient and 0 for the log(IBC) coefficient in the first vector, while in the second vector they were 1 for the log(IBC) coefficient and 0 for the log(ADMS) coefficient. Again, weak exogeneity of the exchange rate cannot be rejected.

In the model with asymmetry and the full IPCA, the cointegrating vector included neither LSNE (the cumulative negative percentage changes in the exchange rate) nor log(CREDSA). In the latter model, no restriction beyond the coefficient 1 for log(FREE) can be imposed on the cointegrating relationship, at the risk of the LM test rejecting non-autocorrelation in the residuals. This difference in the restrictions (and these could not be statistically rejected) between the asymmetric models could explain the inconsistency between the estimates showing that the pass-through of depreciations was larger for the two IPCA components than for the full IPCA.

The base of the full IPCA in the same period was changed to make the series comparable, as shown in figure 3.

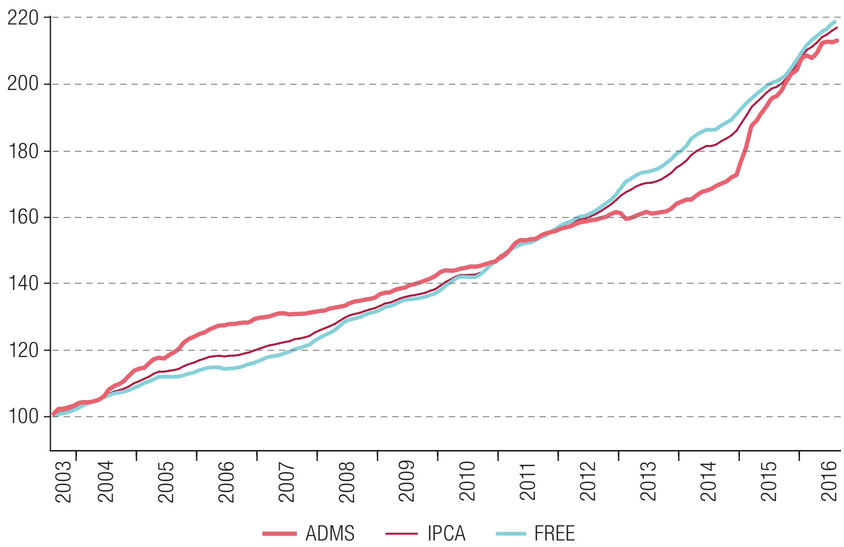
Figure 3
Additional endogenous variables, 2003–2016

A. EXCHANGE
(Cumulative percentage variations)



B. IPCA

(Index: 2003=100)



Source: Prepared by the authors.

Note: EXCHANGE: real exchange rate/average free dollar for the month; LEXCHANGE+: cumulative positive percentage changes in the exchange rate; LEXCHANGE-: cumulative negative percentage changes in the exchange rate; IPCA: extended general consumer price index; ADMS: cumulative changes in the IPCA for administered prices, taking July 2003 as 100; FREE: cumulative changes in the IPCA for free prices, taking July 2003 as 100.

Tables 7 and 8 show the exchange rate pass-through estimated by the models using the methodology described. We present the results of the 12-period price response, taking into account the exchange rate shock in the first period. The long-run effect is the one that remains unchanged in the price level after a certain period, i.e., it is the one that establishes the increase in prices, other things being equal, relative to what they would be if the shock had not occurred at $t = 1$, from a certain time after the shock and for an indeterminate period.

Table 7

Exchange rate pass-through

t	Symmetric IPCA	Symmetric decomposed IPCA		Asymmetric IPCA	
		FREE	ADMS	Appreciation	Depreciation
1	0.0061	-0.0024	0.0281	-0.0043	0.0082
2	0.0147	0.0038	0.0440	-0.0179	0.0135
3	0.0203	0.0119	0.0458	-0.0362	0.0142
4	0.0256	0.0198	0.0445	-0.0461	0.0175
5	0.0330	0.0275	0.0493	-0.0796	0.0404
6	0.0420	0.0360	0.0585	-0.0927	0.0458
7	0.0499	0.0429	0.0677	-0.1134	0.0492
8	0.0566	0.0480	0.0769	-0.1313	0.0548
9	0.0627	0.0523	0.0859	-0.1458	0.0633
10	0.0677	0.0554	0.0939	-0.1559	0.0692
11	0.0719	0.0577	0.1005	-0.1699	0.0777
12	0.0755	0.0595	0.1062	-0.1825	0.0802
Long run	0.0930	0.0623	0.1258	-0.2422	0.1095

Source: Prepared by the authors.

Table 8
Exchange rate pass-through, decomposed IPCA model with asymmetric pass-through

t	Appreciation		Depreciation	
	FREE	ADMS	FREE	ADMS
1	0.0140	-0.0410	-0.0012	0.0424
2	0.0156	-0.0779	0.0021	0.0624
3	0.0026	-0.0980	0.0047	0.0620
4	0.0001	-0.1214	0.0062	0.0765
5	-0.0317	-0.1599	0.0295	0.1109
6	-0.0462	-0.1598	0.0305	0.1382
7	-0.0561	-0.1754	0.0341	0.1477
8	-0.0707	-0.1926	0.0475	0.1391
9	-0.0843	-0.2029	0.0595	0.1413
10	-0.0975	-0.2030	0.0666	0.1521
11	-0.1209	-0.1965	0.0781	0.1589
12	-0.1368	-0.1930	0.0789	0.1657
Long run	-0.2337	-0.2161	0.1302	0.1955

Source: Prepared by the authors.

With our pass-through estimation methodology, tables 7 and 8 should be interpreted as follows: assuming that the model reflects the real dynamics of the variables in the economy, the value of a cell corresponds to the elasticity of the price level in that period relative to a given exchange rate shock at $t = 1$, so that if the value of the cell for $t = s$ is 0.1, this means that, if the exchange rate changes unexpectedly by $x\%$ at $t = 1$, the price change expected by the model relative to what prices would have been in the absence of the shock at $t = 1$ is (x*0.1)% at $t = s$. In other words, the impulse response function measures the change in prices relative to what they would have been if there had not been a shock (unpredicted by the model dynamics) at $t = 1$. Thus, for example, the first model tells us that, if the exchange rate undergoes a change unpredicted by the model dynamics of 1% in the current period, current prices will be 0.0061% higher and will accumulate a change of 0.0755% in 12 months relative to what they would have been in the absence of the shock, assuming all else remains constant.

The above results, like those of Minella and others (2003), indicate a much larger exchange rate pass-through in administered prices than in free prices. Free prices are more subject to market competition, unlike administered prices, which still include contractual prices indexed to price indices that are more sensitive to the exchange rate than the IPCA, such as the monthly general price index (IGP-M). The IGP-M is heavily influenced by the IPA, which is an index more sensitive to exchange rate movements (Belaisch, 2003; Albuquerque and Portugal, 2005), as well as being a kind of producer price index, associated therefore with prices at the top of the production chain (unlike the IPCA, which measures the prices of final goods to consumers), and it has been observed that pass-through decreases along the chain (McCarthy, 2000; Choudhri, Faruqee and Hakura, 2005; Capistrán, Ibarra and Francia, 2012). Furthermore, it has been shown, as in Assis (2017), that there is asymmetric pass-through in which the pass-through of domestic currency appreciation is much stronger than the pass-through of depreciation. This suggests that the exchange rate can be used to control inflation via appreciation. Thus, there is the possibility of offsetting inflationary shocks from other sources through appreciation of the domestic currency, either by attracting capital via a higher interest rate (a transmission mechanism whose power is underestimated when the type of pass-through asymmetry noted in the econometric exercises conducted in this paper is present) or by intervening directly in the currency spot or futures market. The latter seems to have been the preferred route for the Central Bank of Brazil (Garcia and Volpon, 2014; Rossi, 2015) and possibly for good reasons, as Ventura and Garcia (2012) point out.[26]

[26] The authors note that the exchange rate in Brazil, unlike the world's major money markets, is determined not in the spot market but in the very short-term futures market.

We also observe greater asymmetry in free prices and little asymmetry in administered prices, so that the marked asymmetry in pass-through from the exchange rate to the IPCA is due much more to market-regulated prices than to prices contracted for over the course of the year or controlled by the government.

IV. Conclusions

The demonstration of a marked asymmetry in exchange rate pass-through characterized by greater pass-through of domestic currency appreciation shocks indicates that interventions by the monetary authority to strengthen the exchange rate are quite important in preventing further deviation of the inflation rate from its target in the face of other inflationary shocks. Thus, rather than fearing the detrimental effects on inflation control that may result from exchange rate depreciation, there could be a case for reacting through interest rates or intervention in the currency spot or futures market (see Garcia and Volpon, 2014; Rossi, 2015; Assis, 2017, chap. 2). The reason is that domestic currency appreciation would serve to generate deflationary pressures in the face of other inflationary shocks that might occur or be expected, and not just to avoid the second-order effects of cost shocks. Thus, our results indicate that the power of the exchange rate to transmit monetary policy shocks for the purpose of holding down inflation has been underestimated.

In sum, there is strong evidence that exchange rate pass-through is non-linear (asymmetric), and the monetary authority should take this asymmetry into account in its models. It is further suggested that there needs to be more research into non-linear relationships between exchange rates and domestic prices and between the variables of interest to policymakers.

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Annex A1

Table A1.1
Lagrange multiplier tests for autocorrelation of residuals

Lags	Symmetric IPCA		Symmetric FREE and ADMS		Asymmetric IPCA		Asymmetric FREE and ADMS	
	Statistic	p-value	Statistic	p-value	Statistic	p-value	Statistic	p-value
1	39.0808	0.0362	47.2527	0.0993	56.7191	0.0153	59.7967	0.1388
2	20.3334	0.7291	36.0607	0.4658	35.3087	0.5013	42.1443	0.7452
3	22.2339	0.6222	28.7980	0.7976	41.6040	0.2399	52.8725	0.3270
4	30.0535	0.2223	43.9895	0.1693	46.9853	0.1040	52.4281	0.3425
5	20.6480	0.7121	36.2741	0.4559	52.0696	0.0405	59.2674	0.1495
6	27.1980	0.3461	38.6004	0.3529	51.1179	0.0489	53.7773	0.2965
7	35.2182	0.0843	37.8543	0.3847	53.5763	0.0299	50.5711	0.4113
8	19.5033	0.7724	31.2219	0.6952	25.2953	0.9088	28.2160	0.9925
9	22.6693	0.5969	47.4854	0.0954	27.8162	0.8337	42.8133	0.7208
10	25.1717	0.4528	49.9947	0.0605	30.8326	0.7127	47.6918	0.5262
11	19.3288	0.7811	33.7468	0.5762	34.1944	0.5547	50.2111	0.4252
12	30.0456	0.2226	52.5850	0.0366	26.6668	0.8712	51.7590	0.3667

Source: Prepared by the authors.

Table A1.2
Unit root test

	ADF		ERS		Ng-Perron			
					MZa		MZt	
	C	C and T	C	C and T	C	C and T	C	C and T
Critical value at 5% significance	-2.879	-3.438	3.144	5.651	-8.1	-17.3	-1.98	-2.91
Series								
INTEREST	-4.774	-4.672	1.320	2.385	-21.831	-45.101	-3.303	-4.697
LOG(EXCHANGE)	-1.337	-1.489	13.273	40.810	-2.022	-1.869	-0.986	-0.799
LOG(IBC)	-2.757	0.090	28.957	6.256	0.614	-17.576	-0.387	-2.481
LOG(IPCA)	1.873	0.079	4 537.436	60.310	2.126	-0.301	14.643	0.162
LOG(FREE)	1.535	-1.052	5 398.554	77.923	2.106	0.230	15.911	0.129
LOG(ADMS)	-2.046	0.190	657.620	27.254	2.110	-3.121	5.468	-1.129
LOG(CREDSA)	-2.050	-0.313	29.632	5.810	1.199	-19.633	0.645	-3.267
LOG(IC)	-2.589	-2.246	25.554	19.174	-1.012	-5.015	-0.619	-1.434
LOG(IND^{US})	-2.622	-2.605	8.356	14.357	-3.117	-6.191	-1.030	-1.731
LOG(VIX)	-2.904	-2.888	1.746	5.670	-16.201	-16.356	-2.789	-2.833

Source: Prepared by the authors.

Table A1.3
Cointegration tests, p-values

	Symmetric IPCA		Symmetric FREE and ADMS		Asymmetric IPCA		Asymmetric FREE and ADMS	
	Trace	Maximum eigenvalue	Trace	Maximum eigenvalue	Trace	Maximum eigenvalue	Trace	Maximum eigenvalue
0	0.0000	0.0001	0.0000	0.0000	0.0094	0.0218	0.0002	0.0000
1	0.0046	0.0072	0.0022	0.0049	0.1733	0.0903	0.2151	0.2598
2	0.1957	0.3816	0.1324	0.2258				

Source: Prepared by the authors.