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Exchange Rate Determination: An Application of a Monetary Model for Brazil

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Abstract The goal of this paper is to test a variant of the monetary exchange rate determination model, described by Obstfeld and Rogoff (1996), for the Brazilian economy in the recent period. The model starts with the Cagan (The Journal of Political Economy, 66(4):303–328, 1958) money demand, which is complemented by the hypotheses of purchase power parity (PPP) and uncovered interest parity (UIP). We used monthly data of exchange rate, GDP, interest rate for Brazil, and U.S. interest rate and inflation as proxies for international variables. We applied cointegration tests to identify a long run relationship among the variables. The estimated error correction model offers an exchange rate determination model in the short run. Due to potential endogeneity of some variables, GMM was applied to estimate a long-run model of exchange rate determination. The forecasting results of both estimatives were compared with a random walk approach. The results point to the existence of a long and short run equilibrium Real/dollar exchange rate using the structural model, which may be the achievement of this paper.

Keywords Exchange rate determination · Cointegration · Monetary model

JEL F21 · F30 · F17 · F47 · C53

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Introduction

Over the last few years, economic analyses of exchange rates experienced modifications which contributed both to the theory of exchange rate determination and to an empirical observation of exchange rate behavior. Development of econometrics and greater data availability stimulated the growth of empirical work about this subject (Sarno and Taylor 2002).

Despite growth of knowledge concerning exchange rate behavior, a great number of issues remain unsolved, mainly because of financial crises and the development of new currencies, such as the Real in Brazil. One of these issues is the inability to reject the null hypothesis that exchange rates would follow a random walk path. The work by Meese and Rogoff (1983) points to this evidence for main currencies of Europe and Japan. In other words, economists did not have much to say about exchange rate forecasts.

Since the establishment of monetary and inflation target policies, the debate over the determination of exchange rate has increased. An increase in data availability during the 1990's permitted greater statistical confidence, and new works based on the long term path rejected the random walk hypothesis.

For example, in Brazil, which was established as an inflation target regime, one of the strategies against inflation was the provision of an appreciated exchange rate, with interventions of the Central Bank. Among the instruments used were government papers indexed in foreign currency and others which increased Brazilian external and domestic debt. Fiscal imbalance became a critical issue during the Asian crisis in 1997 and the Russian crisis in 1998. At the same time, other instruments helped maintain the exchange rate value, such as foreign capital inflow tax cut and the publication of a normative list to increase the supply of foreign currencies.

The process of the domestic currency overvaluation lasted from 1995 to 1999. At the beginning of 1999, the Central Bank could not keep the pegged float regime and there was a huge depreciation of the exchange rate followed by a floating exchange rate. To avoid the exchange rate pressure turning into a new inflationary process, an inflation target regime for monetary policy was adopted based on a high official interest rate (SELIC). Since then, Real/dollar parity started to float and the curiosity about its determination rose. Despite that, there are few papers in the literature that evaluate the exchange rate behavior in Brazil, and these are mainly work done by officers from the Central Bank of Brazil.

The goal of this paper is to establish a long-run relationship among the Brazilian exchange rate and other key monetary policy variables, following the Cagan money demand model for an open economy, described in Obstfeld and Rogoff (1996). We applied cointegration analysis to identify a long and short run equilibrium relationship and to compare the forecasts results with the random walk approach. In addition, we estimated a Generalized Method of Moments (GMM) equation based on the model, which allowed us to control for endogenous variables.

Besides the introduction, this paper is organized into five sections. The following section presents a brief overview of exchange rate theories. The third describes the theoretical model. The next discusses the econometric procedure, including unit root and cointegration tests and GMM. Finally, in the last section we offer concluding remarks.

Theories of Exchange Rate Determination

International trade theory, in its elasticity approach (Bickerdike 1920; Robinson 1947; Metzler 1948), offered support to trade relations among countries through time. This literature works with marshallian partial equilibrium theory and affirms that there is a consensus about the basic determinants of the import and export demands: real exchange rate, real domestic income, and real demand of the rest of the world.

However, since there was an increase in international trade after the Second World War, some models added income determination to the balance of payments. The principal theoretical jump (Dornbusch 1980) was the integration of income and relative price. In the 1960’s, standard analysis was made by comparative statics in models of demand determination using the income and exchange rate by stipulating relative prices. In sequence, open economics macroeconomic models (e.g. Mundell 1963; Fleming 1962) brought organized structures to theory, which included assets and capital mobility in theses economies.

The development of asset markets and globalization of financial markets made the focus in the balance of payments turn to the capital account. The monetary model and the portfolio equilibrium model, of Walrasian tradition, surpassed the Mundell-Fleming’s model. They essentially criticized the lack of expectation and affirmed that asset markets had a leading role in open macroeconomics and a positive influence on the model.

The monetary approach, developed in the 1970’s, came as an answer to the increase in exchange markets liberalization in many countries. The exchange rate in this model is an asset in which interest rate is adjusted almost instantaneously to help balance the national currency international demand. This model goes in an opposite direction of the previous one, which accepted exchange rate determination to balance the trade flow. Monetarists believe that exchange rate floating can have a movement similar to prices in the asset market.

More recently, models based on general and partial equilibrium were developed, according to open macroeconomics with temporal optimization, which considers time and expectations in the decision making process. In Obstfeld and Rogoff (1996) we found an explanation of why the intertemporal analysis of the current account became common in the 1980’s. This model recognizes that private consumption and investment decisions result in an inter temporal calculation of the agents that take into account, for instance, future expectations on productivity and demand growth.

The benchmark on the debate was the issue raised by Meese and Rogoff (1983) who compared models of exchange rate econometric forecast with a simple random walk exchange rate model. The authors started with general specification and compared structured models with a random walk. They concluded that structured models performances were worse than the prediction based on a simple exchange rate random walk model.

Meese and Rogoff’s paper became a reference on short term exchange rate dynamics, not only for forecasting but also for models of exchange rate determination. This Meese and Rogoff puzzle is one of the six central enigmas in international macroeconomics and is a particular manifestation of the exchange rate dissociation problem. It alludes to the fact that empirical work has found a weak relationship among the exchange rate and lots of macroeconomics aggregates. Theory tells us they have considerable influence.

The international literature considering developing countries is quite poor. We found some papers in Brazil, most of them linked to the Central Bank of Brazil Working Papers. For example, Muinhos et al. (2001) test a model using the unemployment rate and the current account to reach a real long-term exchange rate. They compare their model with a random walk to the Brazilian exchange rate and conclude that the first tends to achieve results that are more realistic.

Muinhos et al. (2003) argue that the possibility of the exchange rate being a random walk is not the best hypothesis for the Brazilian case. They use a model with Uncovered Interest Parity (UIP) from 1999 to 2001 and verify that this captures the exchange rate behavior in Brazil better than the random walk. Finally, Moura and Lima (2007) test the adaptability and forecast power of some empirical models. They observe that those including variables that capture monetary policy (e.g., money supply and interest rate), risk (EMBI) and trade flows had better forecast power on the nominal exchange rate than a random walk.

The available data concerning the Real is fitting now, since we have monthly observations from 1995 until 2009. Those papers do not use cointegration analysis and vector error correction models (VECM) to compare it with a random walk. This is where our paper brings some innovation to the literature. It also brings exchange rate determination back to economic reasoning.

The Model

Obstfeld and Rogoff (1996) describe a discrete money demand model and apply to it the Keynesian money supply equation, with Purchase Power Parity (PPP) and UIP. Considering the equation

$$m_t - p_t = -\eta \mathrm{i}_{t+1} + \phi y_t \tag{1}$$

where: m_t is the log of nominal money at time t; pt is the log of price index at t; η is the semielastic demand for real balances in terms of expected inflation; it+1 is the nominal interest rate at t+1; and yt is the log of real GDP. From the UIP hypothesis, we find the interest rate differential between countries occurs according to the currency movement

$$\mathrm{i}_{t+1} = \mathrm{i}^*_{t+1} + E_t e_{t+1} - e_t \tag{2}$$

where: i^*_{t+1} is the interest rate on foreign-currency bonds (which we call international interest rate); and the differential $E_t e_{t+1} - e_t$ represents the difference between the expected value of the exchange rate at $t+1$ and t.

The hypothesis of uncovered parity with the agent's perfect prediction is due to the supposition of the inexistence of arbitrage. In the end, the purchase power parity is assumed, so, substituting i_{t+1} in and using $e = p/p^*$ in (1):

$$\left(m_t - \phi y_t + \eta \mathrm{i}^*_{t+1} - p^*_t\right) - e_t = -\eta(E_t e_{t+1} - e_t) \tag{3}$$

Exchange rate solution, with PPP and UIP at t is:

$$e_t = \frac{1}{1+\eta}\left(m_t - \phi y_t + \eta \mathrm{i}^*_{t+1} - p^*_t\right) + \frac{\eta}{\eta+1} e_{t+1} \tag{4}$$

In the next period (t+1):

$$e_{t+1} = \frac{1}{1+\eta}\left(m_{t+1} - \phi y_{t+1} + \eta \mathrm{i}^{*}_{t+2} - p^{*}_{t+1}\right) + \frac{\eta}{\eta+1} e_{t+2} \tag{5}$$

Substituting (4) in (3) we get a 2 period result:

$$\begin{aligned} e_t = {} & \frac{1}{1+\eta}\left(m_t - \phi y_t + \eta \mathrm{i}^{*}_{t+1} - p^{*}_{t}\right) + \frac{\eta}{\eta+1} \\ & \times \left(\frac{1}{1+\eta}\left(m_{t+1} - \phi y_{t+1} + \eta \mathrm{i}^{*}_{t+2} - p^{*}_{t+1}\right) + \frac{\eta}{\eta+1} e_{t+2}\right) \end{aligned} \tag{6}$$

By s interaction, we find the exchange rate equation in a stochastic process:

$$e_t = \frac{1}{1+\eta} \sum_{s=t}^{\infty} \left(\frac{\eta}{1+\eta}\right)^{s-t} E_t\left(m_s - \phi y_s + \eta \mathrm{i}^{*}_{s+1} - p^{*}_{s}\right) \tag{7}$$

Equation 7 shows a positive relation between money supply and exchange rate and a negative relation between the GDP and the exchange rate. This is justified by the idea that a product rise elevates the money demand and, as the latter is static because of monetary policy, the domestic prices go down to reach real balances and the domestic currency obtains a higher value.

Therefore, in this work we will verify Eq. 7 to Brazilian data since 1995. We suppose linearity in the parameters and exogeneity of international interest rate and international prices in order to approximate the exchange rate as a function of e_t (m, y, i^*,p):

$$e_t = \alpha m_t - \phi y_t + \eta i_t{}^* - \beta p_t{}^* + \varepsilon \tag{8}$$

With ε being the random error term.

Econometric Analysis

Data

Because the Real currency was instituted in 1994 and the exchange rate was fixed at that time, we chose to use monthly data from January 1995 until June 2009. We consider the U.S. federal interest rate (Fed Funds) as a proxy for international interest rate (i^*) and the U.S. producer price index (PPI), August 2005=100, as a proxy for international price variation (p^*). The nominal exchange rate (e) is the monthly average price of one dollar in Reals (R$/US$). As for money supply, we used the monetary basis M1, monthly average, which comprises instant liquidity liabilities (m). For the GDP (y) we used the monthly series of the gross domestic product in current values (Reals) calculated by the Central Bank of Brazil using the trimestral GDP research conducted by The Brazilian Institute of Geography and Statistics (IBGE). Both the GDP and the M1 series were seasonally adjusted using the X12 multiplicative method. Also, the GDP was deflated using the Consumers National Price Index (INPC/IBGE) starting in January 1995. We transformed all the variables in natural logarithms, as required in (8). The final sample showed 174 observations. We obtained the data series on the Central Bank of Brazil website.

Besides those variables, we included impulse dummies in some periods of the years 1998, 1999, 2001, 2002 and 200 to capture shocks in the exchange rate and one level dummy to measure the change in its policy administration, from pegged to floating.[1] Therefore, dummy3 covers the period when Brazil's exchange rate started to float from January 1999 until June 2009; its value is 1 (floating) or 0 (pegged). For impulse dummies we have 1 from December 1998 until March 1999 (dummy2); 1 in November and December 2001 (dummy5); 1 from August 2002 till February 2003 (dummy1); 1 from December 2008 until March 2009 ("dummy4"). These were used mainly to avoid graphically observed pick movement in the exchange rate.

Unit Root Tests

We begin the econometric analysis by testing for unit root. Both DF and augmented DF are criticized because of the great distortion of the test size and power. Modifications proposed by Elliott et al. (1996), and Ng and Perron (2001) overcame those problems with the development of the ADF test and the Phillips and Perron (1988) test. The modifications were: the use of GLS to detrend the data and the application of a modified Akaike information criteria to lag selection, as suggested by Ng and Perron (2001).

Table 1 summarizes unit root tests results of variables in level utilizing MADF GLS and Ng–Perron. For both test procedures, we chose automatic modified Akaike lag selection with a maximum of 14 lags. In the latter, we used the spectral autoregressive estimation without trend using GLS (AR GLS-detrended).

The series are not stationary at level, both used modified ADF-GLS and Ng–Perron with constant and trend. As the statistics were not significant at 10%, 5%, or 1% we cannot reject the null hypotheses of unit root. Additionally, the series are not stationary at level, and we proceed with the analysis of cointegrating vectors.

Cointegration

Cointegration tests are useful to investigate long-term relationships between two variables, as proposed by Engle and Granger (1987), and Johansen (1988).

Engle and Granger (1987) proposed a long-term equation estimation using Order of Least Squares (OLS) in which the residual should be stationary. First, we have to determine the integration order in each series. If the series are integrated of different order, one can conclude that they are not cointegrated. If results point that the series are I(1), the next step is to estimate a long-term relationship. In case they are cointegrated, an OLS regression can give us a consistent estimator of the vector cointegration parameters. Therefore, to determine cointegration in series, we must analyze the residual obtained from the OLS estimation of the series. If stationary, we conclude that a stationary linear combination exists and series are cointegrated.

[1] Ferreira and Tullio (2002) documented the history of the Brazilian exchange rate regime, which went from pegged to floating in 1999.

Table 1 Unit Root Test

Series	Model	Lag Numbers	MADF-GLS	MZt
e	C	2	−0.17	0.17
e	C,T	2	−0.82	−0.85
m	C	14	0.90	1.24
m	C,T	14	−0.98	−1.10
y	C	2	1.77	1.84
y	C,T	4	−1.53	−1,54
i^*	C	11	−1.02	−23.9[a]
i^*	C,T	3	−1.74	−1.98
p^*	C	5	0.78	0.69
p^*	C,T	5	−1.66	−1.77

Series in log. "C" indicates constant, "T" indicates trend. ([a]) Significant at the 5% level.

The error term contributes to long term equilibrium adjustment, so considering the series at first difference and adding this term one period lagged, we have an error correction model. This process also gives us a long term adjustment speed.

It is important to observe that most cointegrated models of economic literature focus on cases in which the series has a unique unit root. This is because traditional regressions are applied when the series are I(0). Actually, a small number of variables are integrated of orders greater than 1. Therefore, the authors mostly use the expression cointegration to show that the series are cointegrated of order 1.

With the optimal choice of two lags, we estimated an Autoregressive Distributive Lag Relationship (ADLR) for the Brazilian exchange rate using OLS. We included one dummy in the structural break after 1999 (dummy3), a dummy to capture the shock in the exchange rate during the presidential election in 2003 (dummy1), a dummy during the exchange rate fluctuation process (dummy2) and another for the months of November and December 2001[2] (dummy5). After normalizing, in equilibrium, we obtained the Engle–Granger estimated equation

$$e = 0.016m + 0.43y - 0.05\mathrm{i}^* - 1.12p^* + 0.64\,\mathrm{dummy1} + 0.77\,\mathrm{dummy2} + 0.78\,\mathrm{dummy3} - 0.95\,\mathrm{dummy5} \tag{9}$$

The estimated ADL using OLS permitted the residual extraction so we could apply it as a correction vector in short run. Once stationary, we confirmed that series are cointegrated as shown in Table 2. We emphasize that the unit root test applied to the residuals was the DF test and the critical values are those calculated by Engle and Yoo (1987).

In sequence, from general to specific, we estimated an OLS regression with 10 lags to find the short-term relationship using the Error Correction term (ECM) to evaluate the short run dynamic. The static solution is such:

$$\mathrm{e} = 0.012 - 0.13\mathrm{m} + 0.39\mathrm{y} - 0.16\mathrm{i}^* - 2.56\mathrm{p}^* - 1.18\mathrm{ECM}(-1) \tag{10}$$

[2] Dummy for the financial crisis period (December 2008–March 2009) was not significant.

Table 2 Results for Residuals Unit Root Test

Type of Residual	DF Statistic	Critical Value at 1%
Constant	−11.56[a]	−2,6
Trend and constant	−11,53[a]	−2,6

[a] There is no unit root, i.e., it is stationary

The short run analysis suggests a different effect than that proposed by the theoretical model. The idea that a greater money supply contributes to currency appreciation (minus sign), but in the long run this effect turns to depreciation. GDP growth contributed to the currency depreciation in both the long and short run. Both international interest rate and prices were found to have an appreciation effect on their increase. Due to the speed of adjustment, any deviation of the current state from its long run relationship is adjusted rapidly.

Generalized Method of Moments

Hamilton (1994) and Johnston and Dinardo (1997) affirm that Hansen (1982) GMM has been useful in estimating parameters in linear and non-linear models. The dissemination of this method has some advantages over others, such as the fact that GMM estimator does not need the process distribution path or normality. Its standard error is consistent even if the error is characterized by heteroscedasticity.

GMM estimation starts by equalizing the moments of the population origin (μ'k) to the sample ones (m'k). In the generalized method, we estimate one distribution parameter substituting the information of any population moment by a sample moment. Then, we choose the estimative parameter in a way that the theoretical relationship is mostly satisfied. The theoretical relationship is then substituted by the estimative sample in order to minimize the weighted distance between the estimated and theorical values.

For short-term estimations, we applied tests on residuals to verify the presence of autocorrelation and heteroscedasticity. In the Breusch–Godfrey correlation series, we succeeded in rejecting the null hypothesis of autocorrelation (F statistic 0.24, and a probability of 78%). In the Autoregressive Conditional Heteroscedasticity Test (ARCH), we did not reject the hypothesis of autocorrelation (F statistic 9.38, and a probability of 0). Also we could not reject the heteroscedasticity in residuals (F statistic of 6.6 and a probability of 0%). In regards to normality, the Jarque–Bera test rejects the presence of normality (F statistic 3, and a probability of 4%).

Because of the problems described and the possible endogeneity of GDP and M0, we applied a regression using GMM, with the lagged series as instruments. The estimators generated at the GMM are robust and do not require exact information about the probability distribution errors (Moura and Lima 2007).

We used 11 instruments[3] in nine parameters, so we had two over-identifying restrictions. As instruments, we used the lagged values money supply and GDP of the last two periods plus five dummies.

[3] List of instruments: $m(-1)$, $m(-2)$, $m(-3)$, $y(-1)$, $y(-2)$, $y(-3)$ dummy1, dummy2, dummy3, dummy4, dummy5.

The estimated equation is as follows:

$$e = \underset{(0.19)}{0.38m} \underset{(0.13)}{+0.36y} \underset{(0.09)}{-0.20i^*} \underset{(0.7)}{-2.33p^*} \underset{(2.68)}{+0.12\text{dummy1}} \underset{(-3.54)}{0.2\text{dummy2}} \underset{(5.25)}{+0.44\text{dummy3}} \underset{(-2.14)}{-0.5\text{dummy4}} \underset{(-3.69)}{-0.14\text{dummy5}} \tag{15}$$

In the instrument regression, we obtained an adjusted R statistic of 0.88 and a J statistic of 0.0009. The latter is important to evaluate the equation overidentification (Hansen 1982). In this case, J statistic multiplied by the number of observations generated a statistic to test the null hypothesis that overidentification is satisfied. The obtained value was 0.15; therefore, the conditions were satisfied.

The signs obtained were similar to the long run relationship obtained in (9). GMM regression emphasizes money supply and GDP positive effects over the exchange rate depreciation. International interest rate was negative and, again, the greatest coefficient was the one for international prices, which may be due to PPP—when international prices rise, domestic prices become cheaper, increasing the demand for local products and the currency appreciates.

Forecasting

Once we determined the variables as cointegrated, we used the same methodology as used in Cheung et al. (2005) to make a simple comparison in forecast performance between the structural models with a random walk. We used error correction specification of the theoretical model. It allowed for the long run interaction effect of the variables in generating forecast. Hence, we forecasted using the GMM specification.

To evaluate the forecasting accuracy of the different structural models, the ratio between the mean squared error (MSE) of the structural models and a driftless random walk dynamic method is used.[4] A value smaller (larger) than one indicates a better performance of the structural model (random walk). Table 3 summarizes the results, showing that the MSE results for three periods ahead are favorable to the structural model.

Using the MSE rate, the structural model had a better performance in forecasting, followed by the GMM estimation. However, the Theil coefficient of the GMM estimation was lower (0.07 average) than the structural model (0.5 average).

Concluding Remarks

The model proposed by Obstfeld and Rogoff (1996) provides a starting point for the idea that the exchange rate is a relative price between currencies. We tested it with the Brazilian currency and the results obtained were quite intriguing.

First, the cointegration analysis suggests that there is economic interdependency in the exchange rate determination for the Brazilian case. We observe, however, at both the long run and GMM estimation, GDP growth contributed to an exchange rate depreciation, as well as money demand. This might be explained by the fact that GDP growth, in the long run, can suggest a monetary or a consumption expansion,

[4] Due to the driftless random walk model proposed in Meese and Rogoff (1983).

Table 3 MSE Comparison

	1 Month	12 Months	18 Months
Structural model	0.05	0.05	0.04
GMM	0.10	0.10	0.11
Random walk	0.88	0.88	0.88

which amplifies the excess of local currency—at least inside the country. Concerning the VEC, the expected result for GDP is achieved only in the short term of the model. However, this is the atheoretic part, describing deviations from the equilibrium path in the one and other direction.

Second, concerning the interest rate, we did not obtain the expected sign suggested by theory regarding the international interest rate, which suggests a more detailed examination. One observation is that Brazil has one of the highest interest rates in the world, so the variance of other countries' interest rate may not affect the national interest rate. In addition, the non-existence in literature of empirical observations of the UIP may be causing this unclear observation in the Brazilian data.

Third, forecasting comparison suggests that macroeconomic variables influence the exchange rates in Brazil despite the random walk approach. We think that further research could be done with data from other developing countries, because prior studies consider data only from developed countries (Meese and Rogoff (1983) and Cheung et al. (2005)).

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Appendix

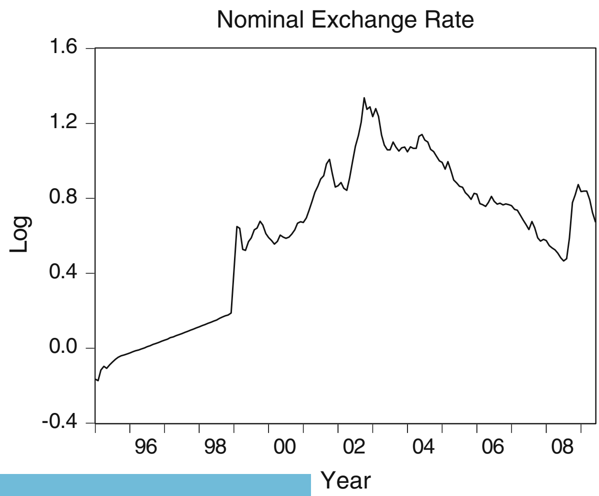

Fig. 1 Nominal exchange rate

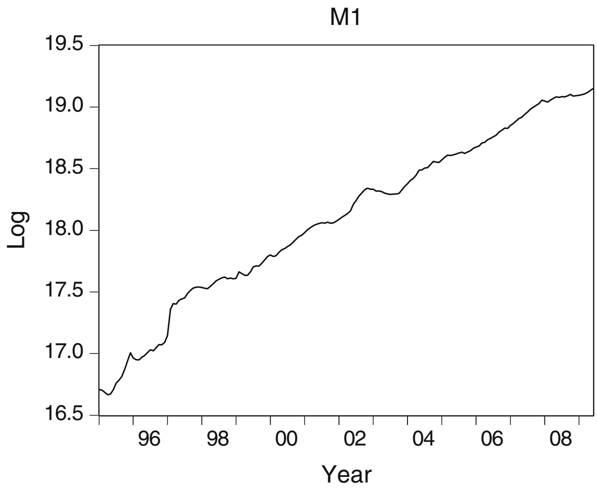

Fig. 2 M1

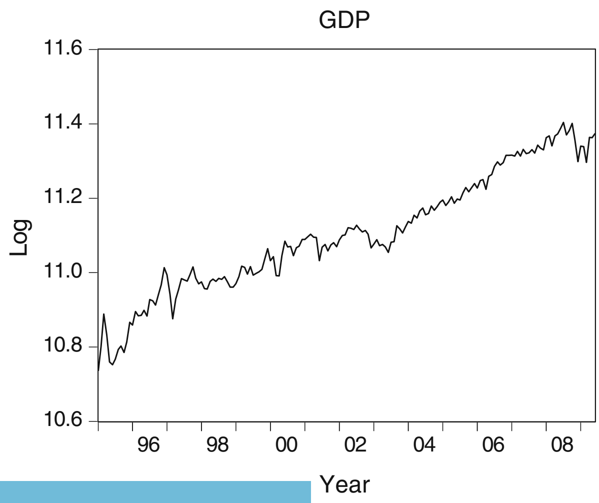

Fig. 3 GDP

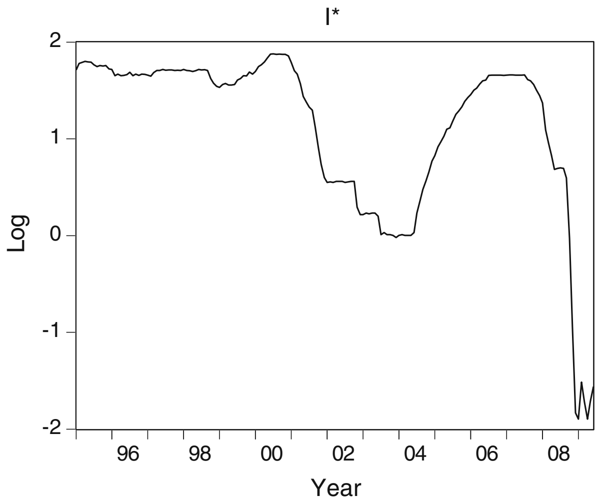

Fig. 4 I*

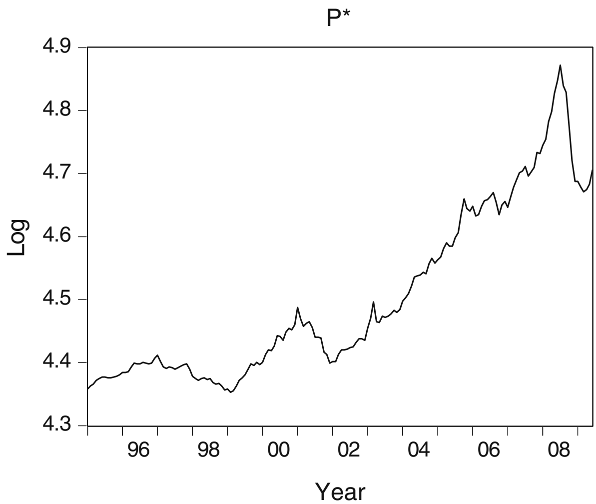

Fig. 5 P*

Source: Central Bank of Brazil.

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