IS THE BRAZILIAN REAL A COMMODITY CURRENCY? LARGE SAMPLE EMPIRICAL EVIDENCE

Filipe Monteiro de Castro Albert
University of Essex, Brazil
E-mail: filipe_albert@yahoo.com.br

Paolo Edoardo Coti-Zelati
Municipal University of São Caetano, Brazil
E-mail: coti_zelati@outlook.com

Davi Lucas Arruda de Araújo
Mackenzie Presbiteryan University, Brazil
E-mail: davi_lucas89@hotmail.com

Submission: 14/08/2013
Accept: 28/08/2013

ABSTRACT

Brazil is one of the world's largest base materials exporters, and this paper examines through large time series samples whether the Brazilian Real can be characterized as a commodity currency. The Real/US dollar real exchange rate and a real commodity prices index are found to be non-stationary and not co-integrated, while a risk premium appeared to have a large and statistically significant long term relationship with exchange rate movements. Combined first difference models showed that real exchange rate elasticity to risk premium is twice as large as to commodity prices, although both variables have considerable influence. Some specifications outperformed a random walk model with respect to root mean square forecast errors for many horizons, but the latter still better determined the exchange rate in longer terms.

Keywords: Exchange rates; Commodity Prices; Co-integration.

http://www.ijmp.jor.br ISSN: 2236-269X
v. 5, n. 1, October – January 2014.

1. INTRODUCTION

An affirmation like the one above might sound extremely alarmist today, when most developed and emerging economies follow similar and stable macroeconomic policies. Back in the 1970s, however, and for many years after that, it made perfect sense for a country with a currency crisis every three or four years like Brazil. It was kind of a mantra for those defending fixed exchange rate regimes, and not until the second half of the nineties, when big economic reforms came into place, it became outdated.

A floating exchange rate regime is today a powerful tool for current account imbalances adjustments and one of the most indispensable policies for Brazilian economic development. In that sense, understanding what factors relate to exchange rate movements and how forecasting could be improved is a very important issue in Brazil. Most local economist have turned to interest rate differentials or current account balances to try and explain how Brazilian Real's value is determined and this paper will look into international literature to find evidence, pros and cons of these traditional methods. More attention, however, will be drawn to Meese and Rogoff (1983) successful exhibit of how a simple random walk can outperform structural exchange rate determination models, and also to a fresh attempt of introducing commodity prices to explain currency movements done by Chen and Rogoff (2002). This last paper, in fact, is going to be of highest importance for the empirical part of this study.

As most developing economies and a few well developed ones, Brazil exports a great deal of base materials. These commodities are traded internationally and have their price settled in US dollars without any direct interference of a specific producer. Therefore, a surge in commodity prices should lead to an increase of Brazilian exports value in dollars, more foreign currency within the country and a possible appreciation of the country's exchange rate. If this actually happens, the Real could be characterized as a "commodity currency", as have been the Australian and Kiwi dollars in more than one occasion. Indeed, evidence for developed countries is pretty easy to find, but dealing with real exchange rates in emerging economies is a tougher job and empirical works are scarce. This is mostly due to lack of reliable data, since, as Brazil, most developing countries have a very unstable past and short series of data. Special care is then given to the quality of data used.

The key objective of this paper is not to examine all variables that interfere with the Brazilian real exchange rate, but only to show if real commodity prices have a significant influence over it. In order to do that, both monthly and weekly data from March 1999 to June 2010 were gathered so that econometric tests could have maximum validity. As in many other studies, the real exchange rate and real commodity prices were found to be non-stationary. After that, variables appeared to have a long run relationship only when Johansen's co-integration tests used very unusual specifications, casting doubts over their results.

Therefore, actual estimation of the real exchange rate continued with models in first differences that considered variables not to be co-integrated. They pointed to statistically significant coefficients of real commodity prices with theoretically expected signs, yet with far less influence than a risk premium measure. Out-of-sample forecasting performances of some specifications were then compared with a simple random walk, and results diverged from classic findings of Meese and Rogoff (1983). By the measure of mean root square error, both specifications outperformed random walk depending on the time forecasting horizon.

This paper is constructed as follows. Section 2 is a comprehensive review of theoretical approaches to exchange rate determination, added by a deep analysis of some insightful empirical works both classic and more recent. Section 3 is dedicated to data description and some graphical evidence of the relationship we are trying to check. This includes how series that were not readily available were constructed. Section 4 shows how unit root and co-integration tests were used to establish some properties for the data, evolving to the econometric specifications used to determine the relationship between the Brazilian Real/US dollar real exchange rate and real commodity prices. Section 5 compares out-of-sample forecasts from our models with results from a simple random walk. Section 6 concludes.

2. LITERATURE REVIEW

2.1 Early exchange rate models

There are many ways in which economists today can try to model exchange rates medium and long term behaviours. All of those are based, at least in some level, on a modern approach to purchasing power parity (PPP) defined in the early 1920's mainly by Swedish economist Gustav Cassel. The idea that in the long term equilibrium currencies (E) must reflect the ratio of domestic prices (P) to foreig prices (P*), or E=P/P*, is still broadly discussed and sometimes useful to compare currencies valuations, as in the famous Big Mac index calculated every year by The Economist magazine. Comparing prices of a homogeneous product that exists in many countries is a clever way to approximate how exchange rates deviate from long term equilibrium. Yet, also in this case basic criticism to PPP cannot be avoided, since non-tradable inputs, taxes and other interferences may apply. Defining an appropriate index to calculate PPP is the biggest problem, given that besides all interferences cited above, there are also significant differences in preferences between countries that make price indexes hard to compare.

In Frenkel (1979), evidence of a working PPP regime in the 1920's was found, but the same could not be said about the floating exchange rates period of the 1970's. Real shocks (oil price shocks, productivity growth, fiscal policies swings, etc…) were said to be the difference, giving extreme volatility to exchange rates and permanent deviations from PPP in the latter period. Also, there are many studies that have tried and failed to show any convergence of exchange rates to PPP on the long term. They all use different techniques and datasets, and are summarized in Rogoff (1996), where the author made popular the concept of a purchasing power parity puzzle. At that point, the Balassa-Samuleson effect of higher productivity in the tradable sector of a richer country was one of the possible factors leading to long term PPP deviations.

Understandings on how exchange rates behave evolved on the post-war era to a balance of payments flows approach. Current account balances were linked to currencies fluctuations, and how to minimize volatility using monetary and fiscal policies was a subject much debated in that time. Fixed exchange rates regimes became fashionable. During the 1960's, however, the famous Mundell-Fleming model of exchange rate determination was developed and enlightened common knowledge of monetary policy's role in economic stabilization. Today's comprehension that a central bank should be independent and therefore exchange rates should float is due to Mundell (1963) and Fleming (1962) findings. The incompatibility of free capital mobility, fixed exchange rates and independent monetary policy was proven, even though criticism to the whole theory not working with stock variables existed.

This theoretical background, that brought attention to government policies, led to a new field of research in the 1970's, where Frenkel (1976), Bilson (1978) an

Dornbusch (1976) took prominence. This was the monetary approach to exchange rate determination, and can be divided into the flexible price models, where the first two economists were responsible, and the sticky price models derived by the latter. Although only able to partially explain (and sometimes not even that) exchange rates movements, these were an important counterweight to the equally incomplete balance of payments flows method. They can be differentiated by the assumption in flexible price models that PPP holds for every period, allowing prices to change instantaneously. Frenkel (1976) and Bilson (1978) arrive to exchange rates influenced by money supply, real income and expectations. A more controversial finding is that rising interest rates differentials cause currency depreciation, contrary to the popular (but not completely verified) belief that high interest rates should attract capital inflows and cause exchange rates to appreciate. This is explained by differences in inflation expectations. A rise in the domestic interest rate is viewed as higher expected future inflation and adjustment in prices will lead to a less valued exchange rate.

Dornbusch (1976) also found interesting results after relaxing the assumption that PPP holds in the short term. Price adjustments occur in a slower pace in the goods markets than in the assets market, leading to an initial overshoot of exchange rates when monetary policy changes. In that sense, the sticky price model can relate to Mundell-Fleming's when it comes to a domestic currency appreciating after an increase in domestic interest rates.

All different monetary models of exchange rate determination, however, have been tested throughout the years with rather disappointing results. Some particular studies have found that exchange rates are co-integrated with monetary fundamentals, meaning there's a long term relationship between them. It can be said, though, that most use insufficiently long time series datasets for statistical inference and sometimes even consider data from periods with a fixed exchange rate regime or other misspecifications. Robust empirical evidence of this theory is scarce.

2.2 Real interest rates differentials models

From the monetary approach of exchange rate determination derived the very popular method of modelling real exchange rates with real interest rate differentials. This method uses the concept of uncovered interest rate parity (UIP), where in a risk-neutral environment real interest rates should converge to equilibrium where no arbitrage gains can be made. Persistent changes in real interest rates differentials are seen as of great influence in real exchange rate determination, although robust empirical evidence is again a big challenge.

Using Rosenberg (1996) notations, we can algebraically demonstrate UIP's intuition in a very simple way. Domestic assets returns (i) must equal foreign assets returns (i*) adjusted by expected changes in nominal exchange rates ($\dot{e}^e$), or i = i* + $\dot{e}^e$. Assuming that the parity is valid for n periods or maturities (short and long run) and rearranging, we will arrive at $\dot{e}^e_n = n(i - i^*)$. This can be expressed in real terms:

$$\dot{q}^e_n = n(r - r^*)$$

Where $\dot{q}^e_n$ is the expected change in the real exchange rate for n periods, and $(r - r^*)$ is the real interest rate differential. Working with a risk premium Ø, we can evolve to the covered interest rate parity equation (CIP):

$$\dot{q}^e_n = n(r - r^*) - n\text{Ø}$$

Note that rational expectations and perfect capital mobility are basic assumptions in this line of thought, perhaps partially explaining why it is so hard for it to be empirically proven.

Given this theoretical framework, many studies in the past two decades have tried to model and forecast real exchange rate behaviours. Mixed results have been obtained at best, with a considerable improvement for long term horizons, due to many different data and specification problems. A common problem and perhaps the most serious one are well described by McCallum (1994b). His main explanation for the UIP failure in the short run is that it does not consider a system of equations. In other words, modelling real exchange rates through uncovered interest rate parity ignores a simultaneity bias that is clear from all agents and central banks actions. A policy reaction function is developed, and other economists like Christensen (2000) made use of that more recently, although with equally unsatisfactory results even for the long term. Since it considers a central bank reaction to deviations of inflation target, a newer dataset that would include emerging countries (who adopted inflation targeting just more recently) and a larger sample for developed economies could probably help.

Meredith and Chinn (1998) follow a similar line of work and extend McCallum's macroeconomic model, using data from G7 countries. They recognize its "intrinsic dynamics" and attribute to the failure of UIP, especially in the short run, time-varyi risk premia and expectation errors. Only in longer horizons, when shocks to premia tend to fade, fundamentals appear to relate to interest and exchange rates. This relationship between risk premia and exchange rates will be explored in the empirical part of this paper.

Another paper with a good insight for the empirical part of this study is Edison and Pauls (1991). As they try to relate real exchange rates to real interest rate differentials, dynamic models (something that will be seen in other papers) slightly improve results after a series of test failures. Both nominal and real variables are found to be non-stationary and not co-integrated, something that is quite common in the literature and as we will see occurs with our dataset. The problem is that not even 100 observations are used by the authors, putting results under doubt.

With more observations and using data for a larger array of countries, MacDonald and Nagayasu (2000) also found non-stationary for both exchange rates and interest rates, along with very weak co-integration results for individual countries. When panel co-integration tests were made, however, strong evidence of a long-run relationship between fundamentals is seen. This is used as an argument for how deficiency of structural models may come from estimation methods rather than the data itself.

Also of interest for the empirical part of this paper is a short study about evidences of UIP working in the case of Brazil. Ferreira (2008) uses exchange rate expectations data and an Instrumental Variables (IV) model to try and reduce the negative simultaneity bias occurring in uncovered interest rate parity. Several misspecification and robustness problems can be drawn from the apparently acceptable results. First, using the Brazilian target interest rate with a US 3-month treasury bill is not ideal. Short term market rates could have been used for Brazil, or maybe longer maturity US bonds could have been used. Other problem arises from not including control variables in the model. It is known that the Real (BRL) fluctuation in the last decade was affected by some exogenous shocks that could have been controlled by including dummy variables or risk premium, for example. In the end, the biggest issue is that results look ok but no robustness checks are made. There are no tests for stationary or co-integration and these issues are not even mentioned, maybe because of the small sample used.

2.3 Seminal papers: Meese and Rogoff 1983/1988

Considering all questions discussed above, and without doubt being the best critique to fundamental exchange rate models made so far, there is a classic paper by Meese and Rogoff (1983) that shall be looked into more carefully. Many insights can be drawn from this very controversial study that shows random walk as a better exchange rates forecasting tool than many structural models.

The authors' idea is to compare out of sample root mean square forecast errors of different modelling strategies. Side by side there is a Frenkel-Bilson flexible price monetary model, a Dornbusch-Frankel sticky price variation and a current account augmented Hooper-Morton model. These structural models are compared to a random walk model, forward exchange rates and univariate and vector auto regressions (VAR). Statistics are calculated for the dollar/pound, dollar/mark, dollar/yen and trade weighted dollar relations, during the floating exchange rate period of the 1970's. Ordinary least squares (OLS), general least squares and Fair's (1970) instrumental variables techniques are used. A general specification for the structural models can be defined as:

$$s = a_0 + a_1(m - m^*) + a_2(y - y^*) + a_3(r_s - r_s^*) + a_4(\pi^e - \pi_*^e) + a_5TB + a_6TB^* + u$$

Where s is the real exchange rate logarithm, $(m - m^*)$ is money supply ratio logarithm, $(y - y^*)$ is the log of real income differential, $(r_s - r_s^*)$ is short-term interest rates differential and $(\pi^e - \pi_*^e)$ the expected long-run inflation differential. TB stands for trade balance and u is the disturbance term. Depending on the theoretical background constraints are added up, but all cases look up for $a_1 = 1$.

As already mentioned in this literature review, the monetary approach for exchange rates determination (including interest rate differential models) might suffer from endogeneity of explanatory variables. This is also indicated by Meese and Rogoff, and not just briefly. The VAR specification, where variables are not treated as exogenous *a priori*, yield results that support endogeneity possibilities. Even so, explanatory variables are said to be legitimate regressors in OLS and GLS, with endogeneity not precluding consistent estimation of structural parameters. In any case, instrumental variables (IV) technique can be used (and in fact it was) if the error term follows an autoregressive process. In the end, however, results from GLS where not worse than IV estimations.

Finally, root mean square errors (RMSE) were taken from one month, six months and twelve months horizon forecasts. Results from the random walk model were not inferior to all other six specifications, being significantly better than them in longer horizons. It is particularly intriguing that models with fundamentals were not able to improve random walk out of sample forecasts even so they were based on realized values of all explanatory variables. Between the three structural models, none could be defined as superior. Estimations in first difference did not help. Regarding the inability of forward exchange rates to beat random walk, the existence of a risk premium is once again brought up. Again, this will be treated in the empirical part of this paper, given Brazil’s unstable past. Still, strong assumptions like market efficiency and rational expectations are also to blame.

In sum, Meese and Rogoff establish many reasons for unsatisfactory results with structural models, mainly simultaneity bias, sampling error and misspecifications, although none could fully explain poor results. It is obvious, however, that the period considered for this study was not ideal, given all structural shocks. Also, expectations variables, still today notably difficult to model, were even less trustworthy back then. Nevertheless, the excellent structure and questions generated by the paper make it very important for exchange rate modelling studies still today, being a benchmark for almost all interesting works in the past two decades.

Meese and Rogoff (1988) is an extension of their previous work, with a larger dataset that allows for unit root and co-integration tests. As seen in most papers, real exchange rates and real interest rates are defined as non-stationary and there is no strong evidence of a stationary linear combination of the two (no co-integration). Authors realize that reasons for non-stationary are different between variables and suggest that there is an omitted variable in the relationship. Note, however, that although sample size is sufficient for these tests, the ones employed (Dickey-Fuller and Engle-Granger) can be considered outdated, with stronger versions available nowadays. Even so, more recent papers with modern tests (including this one) fail to find different results.

Rerunning regressions, now with an extra technique, the generalized method of moments (GMM), the random walk model continued to be unbeatable when considering root mean square errors. An improvement from other studies is that the theoretically anticipated sign of the coefficients can be seen. Failure to produce better results falls again on real structural shocks, with some consideration of speculative bubbles getting in the way.

A more recent extension of Meese and Rogoff's classic paper of the early 1980's is Cheung, Chinn and Pascual (2003), where exchange rate models of the nineties are put into proof by similar criteria. Main differences are the inclusion of a productivity differentials model and a composite specification, along with estimations in first-difference and error correction specification. Forecasting horizons are also distinct, with one, four and 20 quarters ahead. Even with these changes and a very long dataset, results are not much different from past ones. No structural model outperformed random walk using mean square error as criteria, and direction of change depended on the model/specification/currency bundle. This is actually a useful result, since it shows that what might work in one period, or for one currency, might not work for a different structure. From this author point of view, it is an argument against the improved results of panel data specifications. The idea of a unique real exchange rate determination model for every country/period combination sounds misleading.

2.4 Commodity currencies: introducing a new fundamental

Another fundamental that can be included in exchange rate determination models, but only in specific cases, are commodity prices. Differently from terms of trade, that can be defined to any country and have been extensively studied as a part of exchange rate models, commodity prices can only influence countries reasonably dependent on commodity exports and is a factor that is not often included in exchange rate studies. In fact, most empirical studies are only about developed commodity exporters such as Australia, Canada and New Zealand, leaving the relationship between commodity prices and exchange rate in developing countries to be explored. The interesting idea behind choosing commodity prices instead of terms of trade is that these prices are most of the times completely defined internationally, not reflecting any domestic fundamental. In that sense, their changes might represent real shocks, something that has been showed as not captured by traditional models.

A well-structured paper on this topic is Chen and Rogoff (2002), where they analyse if the three countries cited above have so-called commodity currencies. They all fit in the well-developed, small open economy criteria for a good empirical work, with globally integrated financial markets and floating exchange rate regimes. Australia and New Zealand are highly dependent on commodity exports, while Canada has a more diversified trade pattern but still holding large amounts of metals, wood and oil exports. Given that background, authors use variations of this simple linear trend regression of real exchange rates to real commodity prices indexes weighted by the export pattern of each country:

$$\ln(Real\ Exchange\ Rate)_t = \alpha + \beta * t + \gamma * ln(Real\ Commodity\ Price)_t + \varepsilon_t$$

They use the non-parametric GMM Newey-West approach to correct for the biased standard errors estimates, and manipulate the equation above with a Hodrick-Prescott filter and an AR(1) process for the residuals. This last inclusion brings Durbin-Watson statistics towards 2 and therefore end the significant positive serial correlation observed before. Coefficients have consistent estimates, but they are significant only for Australia and New Zealand.

A problem with these results is the assumption of stationary for all variables, since tests for unit roots cannot be taken with such small samples (fewer than 100 observations). As we have seen, there is vast empirical evidence of exchange rates being non-stationary, and although authors don't recognize this, a dynamic OLS model to estimate co-integrating relations is designed, also with positive yet likely misleading results. Again, a small sample is the obstacle for co-integration tests, and a model in first-difference is developed to account for non-stationary and non-co-integration. Coefficients are then significant and with the correct expected sign (as seen in other specifications), suggesting with greater robustness that changes in real commodity prices have great impact on real exchange rates. Commodity prices are considered a good alternative for the missing shocks seen on other structural models.

Main criticism can be made on the assumptions of stationary and co-integration, since there is major evidence otherwise. Most results can be invalid, and interpretations would not be that strong. It is not explained why samples used were so small, given the availability of extended monthly datasets. Also, periods with structural breaks were considered and the Hansen test used to account for that, which relies on asymptotic properties, is not valid. Still, the paper is very comprehensive, since it considers endogeneity problems (and uses a GMM IV specification to account for that), measures the persistence of shocks, and tries productivity differentials as one influence over real exchange rates (yet with le explanatory power than commodity prices). The simple approach to the relationship between real exchange rates and real commodity prices will be followed for the case of Brazil in the empirical part of this paper, with greater attention to stationary and co-integration features.

First, however, it is also interesting to look at evidences for developing countries. Cashin, Céspedes and Sahay (2003) used data from 58 different commodity dependent countries to establish a long run relationship from their base materials exports prices and real exchange rates. Authors had a large data set (from 1980 to 2002) and use Gregory-Hansen co-integration test to allow for structural shifts, since working with unstable developing economies. One third of all countries analysed seem to have co-integrated real exchange rates and real commodity prices, with statistically significant estimates for the real commodity prices elasticity of real exchange rates. Good improvements from previous works are the indication from weak exogeneity tests of real exchange rate adjustment to long-run equilibrium, and extremely fast half-lives of adjustment (only ten months). Criticism can be made to the dataset used, which included very underdeveloped countries, with illiquid, fixed or pegged exchange rates. Also, no control variables or dummies to account for macroeconomic instabilities were included, what could have improved results.

3. DATA DESCRIPTION AND GRAPHICAL EVIDENCE

A major concern in this paper was to work with data from reliable sources and for a period without greater disturbances, knowing the unstable past of Brazil's macroeconomic environment. Therefore, only observations from March 1999 to June 2010 were collected, as in order to minimize excessive volatility or structural breaks in the series. Using older information would cause disruptions, since Brazil only adopted a free floating exchange rate regime and an informally independent, inflation targeting monetary policy, in the beginning of 1999. Before that, and especially before the Real was adopted in 1994, the country would suffer from chronic hyperinflation, frequent currency crisis, capital flights and other severe economic imbalances, turning the development of appropriate econometric specifications a true nightmare.

Also, it was seen as crucial for the success of following econometric procedures to have large enough samples, so tests that have asymptotic properties

ISSN: 2236-269X
v. 5, n. 1, October – January 2014.

could have valid results. In that sense, both monthly and weekly datasets were constructed, the latter having an excellent size of 584 observations. Both datasets used the exact same time series, except for Brazil's inflation (P), where the main index was only available on a monthly basis. The seasonally adjusted IPCA (national consumer price index), calculated since 1979 by IBGE (Brazilian Official Geography and Statistics Institute) with surveys on the 11 greater metropolitan areas in the country, was used for monthly regressions. For weekly models, the chosen index was the IPC calculated by an economics research private institute, FIPE, since 1973. It uses surveys taken in São Paulo, Brazil's largest metropolitan area, and compare moving average prices seen in the last four weeks with a period before that, always weekly updated. They both have high correlation and come from equally renowned institutes.

Note that a private weekly inflation measure is even older than the official survey, a heritage of many years of rampant inflation and lack of statistical scrutiny in government offices. Although prices are well controlled in Brazil for almost two decades now, many other indexes are still calculated by different institutes. The two indexes are used in this paper to deflate exchange rate and international commodity prices series.

For the same purpose, the seasonally adjusted urban all items consumer price index (CPI-U) calculated by the United States Department of Labour's Bureau of Labour Statistics was used. This measure of inflation is produced only monthly, inducing us to use a cubic spline interpolation[1] method to arrive at weekly data. It was taken as foreign inflation (P*) since only the relationship between the Brazilian Real and the US dollar is considered.

That takes us to the nominal exchange rate (E) used in this study, accounted by the Brazilian Central Bank; the commercial mid spot rate of Brazilian Real per US dollar (R$/US$). Data was gathered through Bloomberg. To turn it into the real exchange rate (RER) used in regressions, basic purchasing power parity (PPP) theory led to the formula $RER = E(\frac{P^*}{P})$. Also through Bloomberg, JP Morgan's Emerging Market Bond Index (EMBI) for Brazil was collected to participate as a control variable. It considers the spreads between Brazilian foreign debt bonds and a theoretically riskless US Treasury bond, working as a measure of risk premium. As in every series used in this paper, index base date was March 1999 = 100. It can be seen in figure 1 how the EMBI captures some abrupt moves from real exchange rate, especially during the unstable electoral period of 2002/03, when left wing presidential candidate, Luis Inácio Lula da Silva, was about to win his first mandate. Risk premium rose to unprecedented levels in a few months, just to fall back after fears of an extreme change in the country's orthodox economic policies didn't materialize.

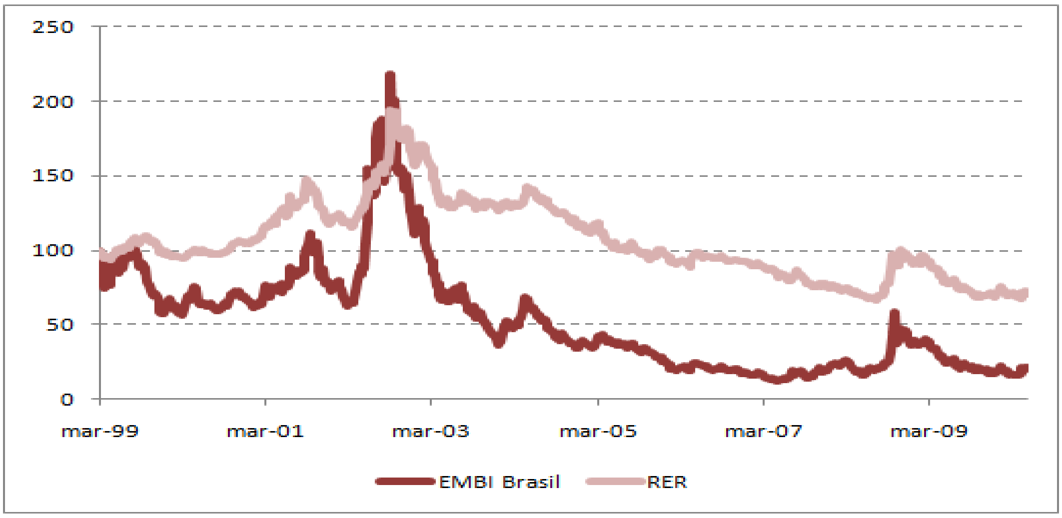

Figure 1. Real/US dollar exchange rate. March 1999 to June 2010

Finally, a commodity prices index suitable for Brazil's pattern of trade had to be created. It required prices from different trade boards in the world, always aiming for the most liquid spot or future commodity contracts traded. Also, weights used reflected the average participation of each commodity on all Brazilian base materials exports during the period between March 1999 and June 2010. In the end, almost 95% of all commodity exports were counted, and the index composition can be summarized as in Table 1. Exports of Petroleum and its products were not considered (as in other significant studies), given their outstanding price volatility. Anyhow, oil has only been significant for Brazilian exports more recently, when large deep water reserves were discovered. In fact, Brazil used to be a petroleum and fuels small net importer for many years.

ISSN: 2236-269X
v. 5, n. 1, October – January 2014.

Table 1. Composition of Non-fuels Commodity Price Index

March 1999 – June 2010		
Product	**Weight**	**Source**
Iron Ore	19,3%	SFCJ
Soybeans	14,8%	CBOT
Sugar	10,8%	NYMEX
Soybean Meal	8,7%	CBOT
Poultry	8,0%	IMF
Coffee	7,2%	NYMEX
Cattle Feeder	7,0%	CME
Hardwood Pulp	6,7%	IMF
Orange Juice	4,4%	NYMEX
Aluminium	3,5%	LME
Soybean Oil	3,3%	CBOT
Lean Hog	2,1%	CME
Ethanol	1,7%	CEPEA
Corn	1,6%	CBOT
Cotton	0,8%	NYMEX

Note: SFCJ (Sinter-feed Carajás), CBOT (Chicago Board of Trade), NYMEX (New York Mercantile Exchange), IMF (International Monetary Fund), CME (Chicago Mercantile Exchange), LME (London Metals Exchange), CEPEA (Centro de Estudos Avançados em Economia Aplicada). World Market Prices (USD)

It is also interesting to see that base materials excluding fuels responded for roughly 40% of Brazilian exports between 1999 and 2008 on average, according to data from the World Trade Organization (WTO). It is not nearly as much as in other Latin American countries, but it is twice as much as in Canada and around what Australia exports, as exemplified on figure 2. That shows how Brazil’s trade pattern is diversified, yet dependent on commodity prices in some level. We should expect then results for Brazil comparable to Chen and Rogoff’s findings for Australia.

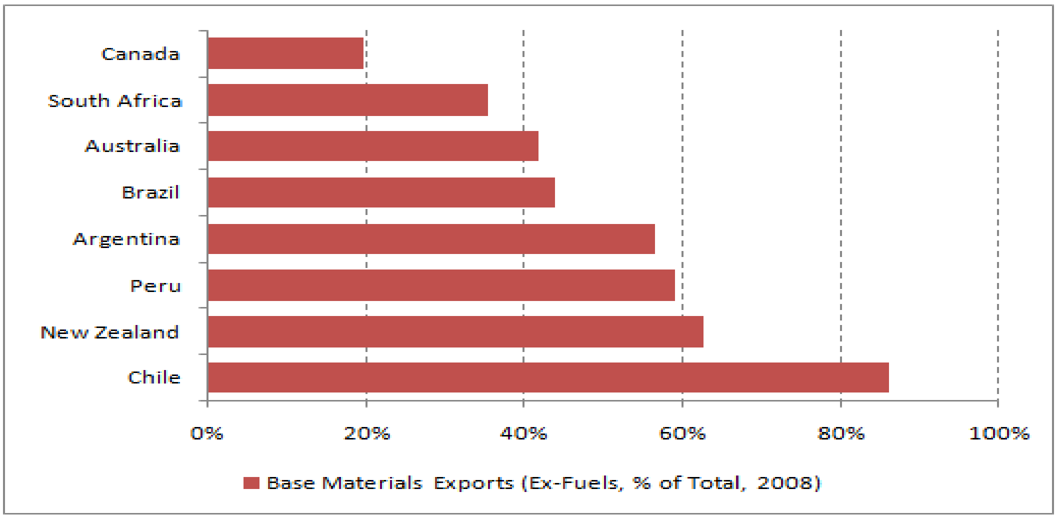

Figure 2. Base Materials Exports excluding Fuels (% of total, 2008). Selected Countries

It is graphically evident in figure 3, however, that Brazil has its own special features that add volatility to the real exchange rate, so that we needed more information than only commodity prices (as Chen and Rogoff used for Australian regressions) to explain currency movements. As said before, the country has a quite unstable past, and control variables are needed. A dummy variable, for example, was used to try and capture 2002/03 pre electoral swings, but with much less success than risk premium as it will be described later in this paper.

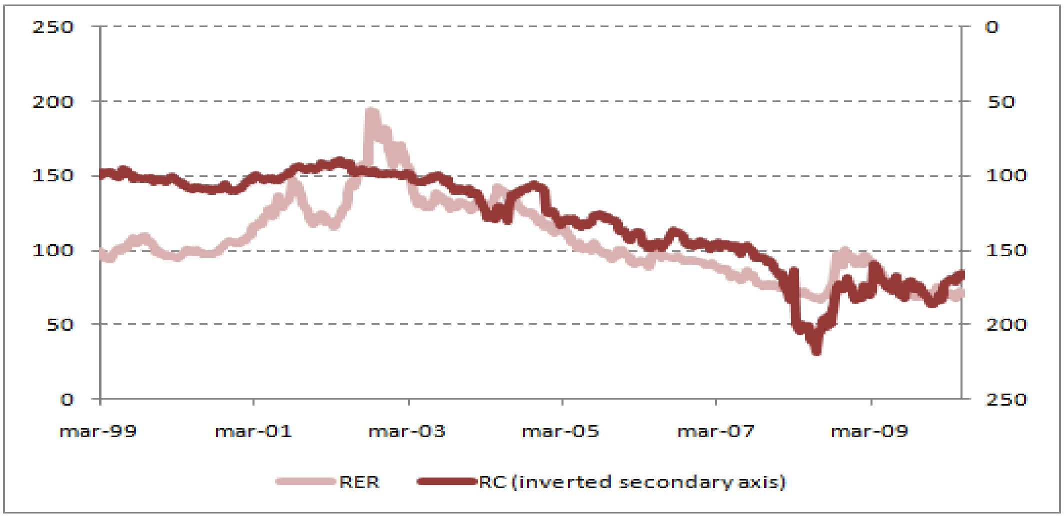

Figure 3. Real/US dollar exchange rate, Real Commodity Price Index, March 1999 to June 2010

4. EMPIRICAL ANALYSIS

As described before, the idea behind this paper and its econometric procedures is not to validate (or not validate) any kind of theoretical approach to exchange rates determination, nor it is to present the most efficient model in forecasting Brazil's Real movements. It is simply to determine whether or not commodity prices influence the Brazilian currency, and perhaps measure the size of its impact. A starting point was the simplest model defined by Chen and Rogoff:

$$\ln(Real\ Exchange\ Rate)_t = \alpha + \beta * t + \gamma * ln(Real\ Commodity\ Price)_t + \varepsilon_t$$

Before evolving from that, however, important steps had to be taken for results not to be misinterpreted. First of all, with a large enough sample of 584 observations, Augmented Dickey-Fuller (ADP) and Phillips-Perron (PP) unit root tests were ran, producing similar results that pointed to non-stationary of both the log of Real/US dollar real exchange rate and the log of the real commodity price index calculated for

Brazil. Insignificant differences between tests were already expected, since the dataset was constructed in order not to have structural breaks, something that could have been captured by the more complex PP test.

Large sample testing confirmed results obtained using monthly data that reduced the number of observations to 136. Table 2 summarizes these results; there are in line with previous findings for different exchange rates. Note that with the smaller sample, t-statistics tended to be more negative and closer to critical values. Even so, results were far from leading us to wrongly reject the null hypothesis of a unit root.

Table 2. Augmented Dickey-Fuller and Phillips-Perron unit root tests, in levels with intercept and trend

Levels	ln (Real Exchange Rate)		ln (Real Commodity Prices)		ln (EMBI)	
t-Stat (ADF)	**- 1.8225**	**- 2.0659**	**- 2.3782**	**- 2.5173**	**- 2.0923**	**- 2.0738**
Adj. t-Stat (PP)	**- 1.8605**	**- 2.2093**	**- 2.3758**	**- 2.6223**	**- 2.0328**	**- 2.2238**
Coefficient	- 0.0091	- 0.0493	- 0.0184	- 0.0869	- 0.0152	- 0.0648
Std. Error	[0.0050]	[0.0238]	[0.0077]	[0.0345]	[0.0073]	[0.0312]
Observations	584	136	584	136	584	136

Note: Critical values for the weekly data are -3.9738, -3.4175 and -3.1311 for 1, 5 and 10 percent (%) levels of significance. For monthly data, critical values are -4.0279, -3.4437 and -3.1466 for 1, 5 and 10 percent (%) levels of significance.

Many possible I(0)/deterministic trends specifications become automatically invalid after results like the ones above, including ones with detrending methods or autoregressive process for the residuals that were used in other papers to correct for clear serial autocorrelation. Just as an exercise, however, a model with linear trend and AR(1) process for the residuals was taken into account, with apparently reasonable, yet clearly deceiving, results. At this point, including the log of the risk premium variable, which is also non-stationary (look at Table B for results), did bring some improvements for coefficients and parameters, yet not changing the validity of results. Results for these specifications are then not reported.

Given all that, the next step is to run co-integration tests that can say whether or not stationary linear combinations between variables, and therefore long term relationships between them, exist. This is made easier since all three variables are integrated of the same order (order one or I(1)), as seen after unit root tests in first differences (Table 3).

Table 3. Augmented Dickey-Fuller and Phillips-Perron unit root tests, in first differences with intercept

Levels	In (Real Exchange Rate)		In (Real Commodity Prices)		In (EMBI)	
t-Stat (ADF)	**- 26.1974**	**- 6.5058**	**- 23.7237**	**- 10.6576**	**- 25.6931**	**- 11.166**
Adj. t-Stat (PP)	**- 26.1261**	**- 11.6189**	**- 23.7215**	**-10.6576**	**- 25.6897**	**- 11.166**
Coefficient	- 1.0834	- 1.0031	- 0.9851	- 0.9295	- 1.0627	- 0.9886
Std. Error	[0.0413]	[0.0869]	[0.0415]	[0.0872]	[0.0413]	[0.0867]
Observations	584	136	584	136	584	136

Note: Critical values for the weekly data are -3.4413, -2.8662 and -2.5693 for 1, 5 and 10 percent (%) levels of significance. For monthly data, critical values are -3.4800, -2.8832 and -2.5784 for 1, 5 and 10 percent (%) levels of significance.

Johansen's test is the most convenient and up-to-date tool, and results are found to be negative for co-integration between real exchange rate and real commodity prices in many specifications. Allowing for a linear deterministic trend and including an intercept, a long term relationship between these two variables was only found with more than 72 lags for weekly data and 24 lags for monthly data, and considering a 10% confidence level. Both trace and maximum eigenvalue tests pointed for the same result using monthly observations and only trace test suggested co-integration for weekly data, results that cannot be considered as anywhere near robust.

A large number of lags can be used in financial markets high frequency variables modelling, but this is not the exact case. The inclusion of the risk premium did not help the results for the group. In fact, it was even trickier to find anything near co-integration when all three variables were considered. Results for the co-integration tests with only the two basic variables and weekly observations can be summarized as in Table 4.

Table 4. Johansen's Co-integration test, linear deterministic trend, Intercept and 72 lags

	Trace	**Max-Eigenvalue**
Statistic	14.1597	11.8672
Critical Value (0.1)	13.4287	12.2965
Prob	**0.0787**	**0.1157**
Observations	511 after adjustments	
	In (Real Exchange Rate)	**In (Real Commodity Prices)**
Normalized Coint. Coefficients	1.000	1.1122
Std. Error	-	[0.1599]
t-Stat	**-**	**6.9555**
	D[In (Real Exchange Rate)]	**D[In(Real Commodity Prices]**
Adjustment Coefficients	- 0.0303	- 0.0015
Std. Error	[0.0106]	[0.0090]
t-Stat	**- 2.8584**	**- 0.1666**

If these results were to be considered, interpretation would be that a 1% change in real prices of Brazilian exported commodities cause a 1,11% impact in the Real/US dollar real exchange rate in the long term. The theoretical relationship between variables is negative (higher commodity prices lower the exchange rate), so the positive sign of the normalized coefficient is as expected, since it is always the inverse of the actual coefficient (β). Adjustment coefficients (alphas) give us the short term corrections, and the negative and more statistically significant one shows the most endogenous variable. In this example, real exchange rates have a greater contribution in reducing short term deviations from long run equilibrium, as seen by the t-Stat closer to the MacKinnon-Haug-Michelis (1999) p-values. Again, this would be a result theoretically anticipated.

The biggest problem is that the specification used for these tests are completely relaxed, and results are not robust enough to produce strong evidence of co-integration between the real exchange rate and commodity prices. Even so, a vector error correction model (VECM) was produced with incredibly long and confusing (probably misleading) results. Since there is great doubt about the co-integrating relation between variables, these modelling results are not reported.

Interesting enough, yet not surprisingly, is the fact that the real exchange rate proved to have a long run relationship with the risk premium variable. Co-integration tests for only these two variables were positive (both trace and maximum-eigenvalue), allowing a linear deterministic trend, intercept and only 2 lags (with standard 5% confidence interval).

During the whole period considered in this paper (a quite volatile one we must say), it is clear that the Brazilian currency followed most of the movements in the measure of risk. Perhaps disregarding the pre-electoral period of 2002/2003 would worsen these results. Anyway, results of the co-integration tests between the real exchange rate and Brazilian EMBI are demonstrated in Table 5.

Note the normalized co-integrating coefficient with the correct negative sign, which is the inverse of the actual β that shows risk premium and exchange rate's theoretical direct relationship. Also, the real exchange rate is again the more endogenous variable, adjusting in the short run. The traditional specification with few lags and robust results attest the variables long term interdependence (yet normality of the residuals could not be verified). This would lead us to develop a vector err

ISSN: 2236-269X
v. 5, n. 1, October – January 2014.

correction model (VECM), but that would not be within this paper's scope. Our goal continues to be testing for the commodity currency profile of the Brazilian Real. Unfavourable results seen earlier, however, direct us to a last resource modelling specification.

Table 5. Johansen's Co-integration test, linear deterministic trend, intercept and 2 lags

	Trace	**Max-Eigenvalue**
Statistic	18.5079	14.9883
Critical Value (0.05)	15.4947	14.2646
Prob	**0.0170**	**0.0384**
Observations	581 after adjustments	
	In (Real Exchange Rate)	**In (Real Commodity Prices)**
Normalized Coint. Coefficients	1.000	- 0.6258
Std. Error	-	[0.1009]
t-Stat	**-**	**- 6.2021**
	D[In (Real Exchange Rate)]	**D[In(Real Commodity Prices]**
Adjustment Coefficients	- 0.0088	- 0.006
Std. Error	[0.0036]	[0.0109]
t-Stat	**- 2.4444**	**- 0.5504**

Following common procedure when variables are found to be non-stationary and not co-integrated, a model in first difference that avoids spurious regressions was developed. Only variables used so far were included, but all possible influences in exchange rates discussed in Section 2, especially interest rate differentials, could have taken part in the experiment. The basic first difference regression augmented by the risk premium measure can be written as:

$$\Delta\ln(\text{Real Exchange Rate})_t = \alpha + \beta_1 * \Delta\ln(\text{Real Commodity Price})_t + \beta_2 * \Delta\ln(\text{EMBI})_t + \varepsilon_t$$

Both explanatory variables have statistically significant coefficients at the 5% confidence level, with $\beta_2$being almost twice as large as β_1. This is in line with co-integration tests results that pointed to a closer relationship between the real exchange rate and risk premium. Also, including lags for the EMBI (t-3, t-4) slightly improved some coefficients, while no commodity prices lag proved to be statistically significant. Similarly beneficial was to include a dummy variable for "stress periods", namely when real exchange rate varied more than 10% in one week. That occurred three times. Once right before Brazilian elections, in September 2002, and also on the first two weeks of October 2008, in the beginning of the most recent financial markets crash and consequent credit crunch with persistent recession in developed countries.

Results for the base first difference model and improved variations can be seen in Table 6. Regressing the real exchange rate against solely the real commodity prices index resulted in larger and statistically significant coefficient, but with an extremely low explanatory power (adjusted R^2). All coefficient signs are correct and as expected for every specification.

Table 6. Dependent Variable: Log of R$/US$ Real Exchange Rate Sample. Period: March 1999 – June 2010

	I(1) / Non-Cointegration			
	Basic 1^{st} differencing	**Augmented 1^{st} differencing**	**Augmented 1^{st} differencing with lags**	**1^{st} differencing with dummy**
ln(Real Commodity Prices)	- 0.2048 [0.0510]	- 0.1279 [0.0379]	- 0.1305 [0.0379]	- 0.1107 [0.0367]
ln(EMBI)	- -	0.2195 [0.0100]	0.2214 [0.0100]	0.2145 [0.0097]
$ln(EMBI)_{t-3}$	- -	- -	0.0210 [0.0100]	- -
$ln(EMBI)_{t-4}$	- -	- -	0.0236 [0.0101]	- -
Dummy	- -	- -	- -	0.0541 [0.0080]
Durbin-Watson Stat	2.1843	2.1290	2.1424	2.1176
Adj. R^2	0.0253	0.4645	0.4699	0.5046
Observations	583	583	579	583

Note: Standard Errors in parentheses.

It is clear from results that real commodity prices fluctuations then have a quite not negligible influence over real exchange rates, although overshadowed by the larger impact of the EMBI when this variable is considered. Log-linear models in first differences denote elasticity in their coefficients and in the augmented one with lags, total real exchange rate elasticity to the risk premium is more than double the one to real commodity prices.

The main question that arises from these regressions is the validity of including a co-integrating variable (EMBI) in the first differencing specifications. In Chen and Rogoff (2002) this approach is considered not appropriate if a long term relation between variables is observed, but there are no further explanations on the subject. Also, there is no third (not co-integrating) variable in their models, which makes things substantially different. Theoretical proof of this affirmation could not be found by this author, nor could similar analysis could be found in the literature review. However, perhaps it is wiser to consider the specifications that included the risk premium variable very carefully, especially when gauging forecasting performance on Section 5.

5. FORECASTING PROPERTIES

Following the results presented in the last section, we advance to a simple comparison of out-of-sample forecasting performance of main specifications. As in Meese and Rogoff (1983), the idea is to measure the root mean square forecast errors of fundamental models (in this case with commodity prices) and compare to a random walk (in this case with a trend) error. They were calculated according to the authors cited formulation:

$$RMSE = \left\{ \sum_{s=0}^{N_{k-1}} [F(t+s+k) - A(t+s+k)]^2 / N_k \right\}^{1/2}$$

Where k are the forecast horizons, s the modelling choice, F(t) the forecast value, A(t) the actual value already known and N_k the total number of forecasts in the projection period. Basic real exchange rate versus real commodity prices regression and the one including the risk premium variable were chosen, and four different forecasting horizons (1 month, 6 months, 12 months and 5 years) were selected to measure how models behave in the long run. Results in Table 7 show a good performance of the pure commodity prices specification in the short run, which fades away over time. Random walk predictions are better only in the longer horizon, contradicting Meese and Rogoff findings of predominance of the random walk for every term. Modelling with the risk premium variable give us low and balanced root mean square errors. Note that all out-of-sample forecasts use actual values of explanatory variables.

Table 7. Root mean square forescat errors*

Horizon	Random Walk (with a Trend)	Basic 1st differencing	Augmented 1st differencing with lags
1 month	3.69	2.98	3.06
6 month	2.98	2.93	2.16
12 months	10.64	16.92	6.02
5 years	9.27	13.71	13.02

Note: *Approximately in percentage terms

These interest findings show how the Brazilian exchange rate cannot be easily modelled, and all results undoubtfully suffer from omitted variables problems. This could have been corrected with an extended research and perhaps the use of monthly data only (not every possible variable is available in weekly terms).

6. CONCLUSION

It is clear from the diversity of exchange rate determination models developed through the years that establishing what variables or fundamentals influence currencies valuations is not an easy task. This is especially true when dealing with developing economies such as Brazil. In this paper we were able to at least determine some important features for the Real/US dollar exchange rate, which kept us away from extremely misleading econometric results.

As described in other papers for different currencies and other related variables, the Brazilian real exchange rate and real commodity prices were found to be non-stationary. Also, no strong signs of co-integration between these variables could be seen after many tests specifications. A long term relationship with a risk premium variable, however, was easily spotted. These mixed results led us to different modelling approaches, which produced interesting yet not quite as expected results.

An array of basic first difference models was best considered, trying to avoid any severe misspecification. Their consistent results allow us to say that, in the case of Brazil, during the last eleven years the real exchange rate elasticity to risk premium was more than twice as large as the one to real commodity prices, although both variables had reasonably large influences. Compared with a traditional random walk model, forecasting performances of these models are strikingly good. Contrary to classic findings, structural first difference models outperformed random walk when considering root mean square forecast errors in some different horizons. Random walk proved to be a better forecaster only in the longer period.

Many improvements can be made to all specifications defined in this paper. There are clearly omitted variables that could have produced better results, but due to either lack of time or availability of data could not be used. Also, the question regarding the validity of using a co-integrating variable on first differences models is left open.

More research is necessary if one is to accomplish the difficult task of defining the most efficient exchange rate determination model for Brazil. From this work, however, it is possible to say that commodity prices have a moderate influence on the country's currency, although characterizing the Real as a commodity currency might be extreme.

REFERENCE

BILSON, J. (1978). The Monetary Approach to the Exchange Rate: Some Empirical Evidence. **International Monetary Fund (IMF)** Staff Reports n. 25, p. 48-75.

CHEUNG, Y., CHINN, M., PASCUAL, A. (2003). **Empirical Exchange Rate Models of the Nineties: Are Any Fit to Survive?** Santa Cruz Center for International Economics.

CHEN, Y. and ROGOFF, K. (2002) Commodity Currencies and Empirical Exchange Rate Puzzles. **International Monetary Fund (IMF)** Working Paper 02/27. Staff Reports 76.

DORNBUSCH, R., (1976), Expectations and Exchange Rates Dynamics, **The Journal of Political Economy**. v. 84, n. 6, p. 1161-1176.

FRENKEL, J. (1979). Purchasing Power Parity: Doctrinal Aspects from the 1920s. **National Bureau of Economic Research**. Paper n. 4.

EDISON, H., PAULS, D. (1991). A Re-Assessment of the Relationship Between Real Exchange Rates and Real Interest Rates, Board of Governors of the Federal Reserve System. **International Finance Discussion**. Paper n. 408.

FERREIRA, A. (2008). The Simultaneity Bias of Uncovered Interest Rate Parity: Evidence for Brazil, **University of São Paulo**, FAERP

FLEMING, M. (1962). Domestic Financial Policies Under Fixed and Under Floating Exchange Rates. **International Monetary Fund (IMF)** Staff Papers

FRENKEL, J., 1976, A Monetary Approach to the Exchange Rate: Doctrinal Aspects and Empirical Evidence, **The Scandinavian Journal of Economics**, v. 78, n. 2, p. 200-224.

MACDONALD, R., NAGAYASU, J. (2000). The Long Run Relationship Between Real Exchange Rates and Real Interest Rate Differentials: A Panel Study. **International Monetary Fund (IMF)** Staff Papers, v. 47, n. 1.

MCCALLUM, B.T., (1994b). Monetary Policy and the Term Structure of Interest Rates, **NBER Working Paper**, 4938.

MEESE, R., ROGOFF, K. (1983), Empirical Exchange Rate Models of the Seventies: Do They Fit Out of Sample? **Journal of International Economics**. n. 14, p. 3-24

MEESE, R., ROGOFF, K. (1988) Was it Real? The Exchange Rate-Interest Differential Relation Over the Modern Floating-Rate Period. **The Journal of Finance**, v. 43, n. 4, p. 933-948.

MEREDITH, G., CHINN, M. (1998). Long-Horizon Uncovered Interest Rate Parity, **NBER Working Paper**, 6797.

MUNDELL, R. (1963). Inflation and Real Interest. **The Journal of Political Economy**. n. 71, p. 280-283.

ROGOFF, K. (1996). The Purchasing Power Parity. **Journal of Economic Literature**. n. 34, p. 647-688.

ROSENBERG, M. (1996). **Currency Forecasting: A Guide to Fundamental and Technical Models of Exchange Rate Determination**. McGraw-Hill, 2nd edition.