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The Simultaneity Bias of the Uncovered interest Rate Parity: Evidence of Brazil

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The Simultaneity Bias of the Uncovered Interest Rate Parity: Evidence for Brazil

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Abstract

We test *ex ante* uncovered interest parity (UIP) for Brazil using survey data of exchange rate expectations from the Brazilian Central Bank. Using data from 2001M11 until 2007M12 and Ordinary Least Squares, we found that the estimated UIP parameter is smaller than one, which is a common finding. We then develop a model that explains how a negative bias can arise due to the simultaneous actions between the Central Bank and speculators. Our results, using Instrumental Variables, show that the bias can be reduced, and lend support to *ex ante* UIP. The reduced form, dynamically complete model provides the best fit for expected exchange rate changes, as it aims to represent the data generation process of the observed data, in contrast to the single structural equation.

Keywords: Uncovered Interest Rate Parity, Simultaneity.

JEL Classification codes: E43, E52, E58, F31

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Introduction

Although support for uncovered interest rate parity (UIP, hereafter) has been growing, this hypothesis still causes some embarrassment from the empirical point of view[2]. There are competing explanations for the failure of short-run UIP (for example, risk, Peso problems, improper econometric techniques etc) but none seem to be widely accepted and there is no consensus on the subject. Hence, there is an open field for investigation and space to work towards some sort of consensual explanation.

This paper builds on the work of McCallum (1994) who puts forward a model that recognizes the simultaneous action of agents (globalized speculators) and Central Bankers in determining equilibrium interest and exchange rates. His model assumes interest rate smoothing and reaction against exchange rate changes, and implies that the failure of the hypothesis can be policy driven.

The endogeneity issue has already been investigated. Many authors have recognized the potential of this explanation in solving the UIP problem, for example, Meredith & Chinn (1998) and Favero & Giavazzi (2004) (to cite just a few) and for exchange rates, in particular, one can see Engel & West (2005). Kugler (2000) was a pioneer in seeing the main implications of McCallum (1994)'s model for ordinary least squares (OLS) estimations. He applied the model to analyze the term structure of interest rates and derived the asymptotic bias using McCallum (1994)'s policy reaction function. On the other hand, Christensen (2000) tested the policy reaction function of the McCallum (1994) model for the US, Germany and Japan but did not find supporting evidence regarding the size of the parameters needed to generate the negative bias.

Our contribution is to test *ex ante* UIP for Brazil while taking into account the endogeneity problem. We also develop a simple macroeconomic model that does not hinge on the assumption of "leaning against the wind" and shows that reaction against prices can be enough to generate the bias on UIP. Furthermore, we show the associated asymptotic bias and provide a hint as to the size of the structural parameters needed to generate a negative bias. Our results show that the Instrumental Variable (IV) estimation of *ex ante* UIP reduces the Ordinary Least Squares (OLS) bias. There is evidence supporting *ex ante* UIP and that the dynamically complete model, which better represents the observed data (when variables are in equilibrium), produces the best fit.

The rest of the paper is organized as follows. We first present the model and discuss both the asymptotic bias and the fully dynamic model for expected exchange rate changes. The penultimate section is dedicated to the empirical findings and the final to conclusions.

Endogeneity

We conclude from McCallum (1994)'s article that the empirical failure of short-run UIP is due to researchers overlooking the fact that this hypothesis (concerning equilibrium in the assets market) belongs to a system of equations. Hence, a regression of UIP using OLS

[2]See, for example, Isard (2006) and Chinn (2006).

produces estimated parameters that cannot have a structural interpretation and could also be subject to simultaneity bias. A shortcoming of his model is that it contained only two equations, the policy function and UIP itself. Another complication is that monetary authorities react to exchange rate changes but not to deviations of inflation from its target. In order to overcome these limitations, our paper considers a Taylor rule type function (under a strict inflation target) as well as other equilibrium relationships, such as the modeling of the demand and supply side of the economy.

This section aims to illustrate how a negative bias can arise from the OLS regression. Our objective is to obtain a closed-form analytical solution for the reduced form model along the lines of McCallum (1994) and Engel & West (2005), for instance, but without resorting to the explicit inclusion of exchange rates in the policy function. A possible justification is the non supportive result presented by Christensen (2000) on "leaning against the wind". The model presented here describes a simplified open economy, as opposed to using more detailed model specifications - see for instance, the interesting works of Meredith & Ma (2002) and Alexius (2002). A complex structure would require numerical solutions and simulations, which is an avenue of investigation that we chose not to follow.

As can be seen below, the first equation of the system stands for the UIP relationship under imperfect capital mobility, while the remaining three equations represent the monetary policy reaction function, the Phillips curve and the IS relationship, respectively. As can also be inferred, they result from the subtraction of the foreign equation from the domestic counterpart, assuming that parameters are analogous in both economies:

$$s_t = s^e_{t+1} - (i_t - i^*_t) + \xi_t, \tag{1}$$

$$i_t - i^*_t = \rho(i_{t-1} - i^*_{t-1}) + (1-\rho)[i^n_t - i^{*n}_t + \lambda[\pi_t - \pi^*_t - (\pi^T - \pi^{*T})], \tag{2}$$

$$\pi_t - \pi^*_t = \pi^e_{t+1} - \pi^{e*}_{t+1} + \eta_1(h_t - h^*_t) + e^s_t, \tag{3}$$

$$h_t - h^*_t = -\eta_2[i_{t-1} - i^*_{t-1} - (\pi_t - \pi^*_t)] + e^d_t, \tag{4}$$

where s_t is the natural logarithm of the nominal exchange rate (defined as the domestic price of the foreign currency); i_t is the nominal interest rate paid on a one-period bond. The superscript e denotes expected values and the asterisk denotes an exogenous determined foreign variable or the foreign economy; ξ_t represents all other variables that explain differences in nominal returns. In fact, we will think of ξ_t as a risk term, as is often done in the literature. For simplification, we start with the assumption that ξ_t is white-noise. The variable π_t stands for the inflation between $t-1$ and t, while π^e_{t+1} is the inflation forecast made at t for the following period and π^T is the inflation target for $t+1$ known at t; π^T is constant and equal to zero by hypothesis, i.e. $\pi^T = \pi^{*T} = 0$. The variable i^n_t is the neutral interest rate, i.e. $i^n_t = r_t + \pi^T$. The letter r_t represents the equilibrium real interest rate, which is determined by two real factors: the marginal product of capital of a larger foreign economy and risk premium. The time subscript in r is explained by the time-varying risk, which implies a time-varying, neutral, real rate. The log of the output gap is represented by h_t. The error terms e_t^s and e_t^d stand for supply and demand shocks, respectively, and are both random variables. The other letters are parameters: ρ is the smoothing term, $0 < \rho < 1$; λ measures the extent to which money authorities react to deviations of inflation from target, and $\lambda > 1$; η_1 and η_2, both positive quantities, measure the sensitivity of the actual inflation differential to the output gap and the sensitivity of the output gap to the lagged real interest rate, respectively.

We complete the model by assuming a process for inflation expectations

$$\pi_{t+1}^e - \pi_{t+1}^{e*} = \phi \Delta s_t + (1-\phi)(\pi^T - \pi^{*T}), \tag{5}$$

where ϕ shows the extent to which the expected inflation differential is anchored in relative purchasing power parity and $(1-\phi)$ on the inflation target differential[3], and $0 < \phi < 1$. Also note that we can write

$$i_t^n - i_t^{*n} = r_t + \pi^T - (r_t^* + \pi^{*T}).$$

Hence, the process for the nominal natural interest rate differential is simply given by the real interest rate differential which we express as

$$i_t^n - i_t^{*n} = r_t - r_t^* = \xi_t + \mu_t, \tag{6}$$

where μ_t is the forecast error of exchange rate depreciation. The meaning of (6) is that, in the absence of shocks and in initial equilibrium, the monetary authority will set the nominal interest rate at a level that will not induce flows of capital. The rule also prescribes adjusting i_t to shocks in risk.

As UIP is often tested using $nid_t = \Delta s_{t+1}^e + \xi_t$ where Δ stands for the first difference and $nid_t = i_t - i_t^*$, we have to show that nid_t and ξ_t are correlated. We start by substituting (5) and (4) into (3) which gives

$$\pi_t - \pi_t^* = \phi \Delta s_t + \eta_1\{-\eta_2[i_{t-1} - i_{t-1}^* - (\pi_t - \pi_t^*)] + e_t^d\} + e_t^s. \tag{7}$$

Substituting (6) and (7) into (2) and solving the resulting expression for the nid_t, we can write

$$nid_t = \alpha_0 nid_{t-1} + \alpha_1 \Delta s_t + \alpha_2 \xi_t + e_t, \tag{8}$$

where

$$\alpha_0 = \frac{[\lambda(1-\rho) + \rho]\eta_1\eta_2 - \rho}{\eta_1\eta_2 - 1},$$

$$\alpha_1 = \frac{\lambda\phi(\rho - 1)}{\eta_1\eta_2 - 1},$$

[3] One could think of ϕ as measuring some sort of expectational pass-through mechanism and $1-\phi$ the degree of credibility of the Central Bank.

$$\alpha_2 = 1 - \rho,$$

and,

$$e_t = \frac{(\rho - 1)[\lambda(e_t^s + \eta_1 e^d) + (1 - \eta_1\eta_2)\mu_t]}{\eta_1\eta_2 - 1}.$$

Observe that the variable nid_{t-1} is predetermined and, because ξ_t, e_t^s, e^d and μ_t are all i.i.d., e_t is also exogenous and i.i.d. In order to obtain the reduced form, we have to take into consideration rational expectations UIP. Substituting the process for the nid_t in (8) into equation (1) and solving for expected exchange rate changes gives

$$\Delta S_{t+1}^e = \alpha_0 nid_{t-1} + \alpha_1 \Delta s_t + (\alpha_2 - 1)\xi_t + e_t. \tag{9}$$

Then we postulate a bubble-free linear solution using the relevant state variables, as below

$$\Delta s_t = \gamma_0 nid_{t-1} + \gamma_1 \xi_t + \gamma_2 e_t. \tag{10}$$

In order to solve for Δs_t, we use the method of undetermined coefficients. After abandoning a non-stationary root ($\gamma_0 = 1$), one reaches the following solution for the nid_t.

$$nid_t = \frac{\lambda\phi(\rho - 1)}{\rho(\lambda\phi - 1) + [\lambda + \rho(1 - \lambda)]\eta_1\eta_2 - \lambda\phi}\xi_t \tag{11}$$

In summary, the conclusion is that nominal interest rate and the variable ξ_t are correlated, which will render OLS estimators biased and inconsistent.

Asymptotic Bias

Now, if you wish to estimate equation (1) by ordinary least squares (OLS)

$$\Delta S_{t+1}^e = \beta_0 + \beta_1 nid_t + \epsilon_t \tag{12}$$

where $\epsilon_t = -\xi_t$. The asymptotic value of β_1 will be

$$\text{plim}(\hat{\beta}_1) = \beta_1 + \frac{\text{Cov}(nid_t, \epsilon_t)}{\text{Var}(nid_t)}. \tag{13}$$

where plim is the probability limit when the sample size grows to infinity. Hence, Bias $=$ $\text{Cov}(nid_t, \epsilon_t)/\text{Var}(nid_t)$. As $\text{Var}(nid_t) > 0$, the sign of the bias will depend on how $\text{Cov}(nid_t, \xi_t)$ differs from zero. As $\epsilon_t = -\xi_t$, the bias will be negative only if $\text{Cov}(nid_t, \xi_t) > 0$. As we assumed that nid_{t-1} and ξ_t are uncorrelated, we can write

$$\text{Cov}(nid_t, -\xi_t) = -\mathbb{E}(nid_t \xi_t), \tag{14}$$

and, hence, we need to find $\mathbb{E}(nid_t \xi_t)$, as shown below

$$\mathbb{E}(nid_t \xi_t) = \mathbb{E}\left\{ \frac{\lambda\phi(\rho - 1)}{\rho(\lambda\phi - 1) + [\lambda + \rho(1-\lambda)]\eta_1\eta_2 - \lambda\phi} \xi_t^2 \right\}.$$

For reasonable parameter values, the negative covariance in the UIP structural equation will arise, leading to statistical bias.

Since the β_1 of the "population" is equal to one, the estimated parameter can be negatively biased according to the simple model above. This suggests, for instance, that IV estimation is more appropriate for UIP tests, provided one has the proper instruments. However, as the structural equation will not represent the observed data for the expected exchange rate change, we will first derive its reduced form, dynamically complete model.

The dynamically complete model

Substituting the process for nid_t from (11) into (1) and solving for expected exchange rate changes gives

$$\Delta s^e_{t+1} = (\kappa - 1)\xi_t, \tag{15}$$

where

$$\kappa = \frac{\lambda\phi(\rho - 1)}{\rho(\lambda\phi - 1) + [\lambda + \rho(1-\lambda)]\eta_1\eta_2 - \lambda\phi}.$$

Solving (15) for ξ_t and writing the result with one lag results in

$$\xi_{t-1} = \frac{1}{\kappa - 1}\Delta s^e_t. \tag{16}$$

We then consider that the variable representing risk is serially correlated, which is a feature of our data and also a frequent assumption of the literature. We introduce some dynamics into the reduced form by assuming serial correlation of the AR(1) type

$$\xi_t = \theta_0 + \theta_1\xi_{t-1} + \zeta_t, \tag{17}$$

where ζ_t is white noise. Substituting (16) into (17) , generates

$$\xi_t = \theta_0 + \frac{\theta_1}{\kappa - 1}\Delta s^e_t + \zeta_t, \tag{18}$$

and we finally substitute (18) into (15)

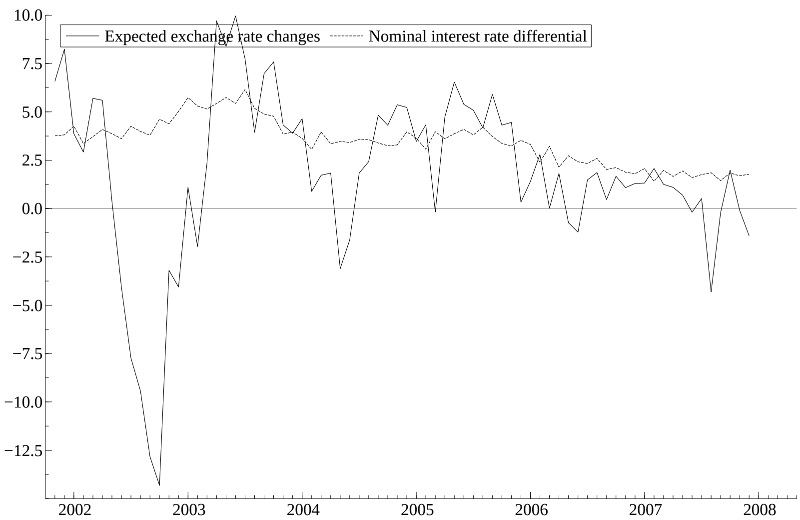
$$\Delta s^e_{t+1} = \theta_0(\kappa - 1) + \theta_1\Delta s^e_t + \frac{1}{\kappa - 1}\zeta_t. \tag{19}$$

As will be shown, our tests reveal that serial correlation was eliminated when estimations were made using the fully dynamic model in (19).

Empirical Results

Data from exchange rate surveys was obtained from the Brazilian Central Bank, while data on interest rates (the annual Selic for Brazil and the three-month Treasury Bill for the US) was taken from the Institute of Applied Economic Research (Ipea). The Selic was transformed into a three-month rate and the expected change in the exchange rate (consistent with the interest rate of month t) was calculated as the average of the daily forecasts during t for $t+3$ minus the spot rate at t.

We initially present Graph 1, for which the series of ΔS^e_{t+1} and nid_t are plotted. A feature of the data that stands out is the large drop in expected exchange rate changes from 2002 and 2003. The sharp depreciation of the Brazilian Real in that period was largely caused by a hike in default risk due to uncertainty regarding the presidential election. The nominal interest rate differential, however, remained relatively stable.



Graph 1: Monthly changes in ΔS^e_{t+1} and nid_t

We first estimated equation (12) using OLS and found a parameter close to 1. In fact, the 99% confidence interval contains 1 which is already a surprising result favoring *ex ante* UIP (see Table 1). However, the point estimate is below 1, which could be due to a negative bias of the type presented in our theoretical model.

We then estimated the model by two stage least squares using the second lag of the nid_t as an instrument[4]. Results presented in Table 2 show that the parameter is larger than when using OLS and that the point estimate is slightly closer to the one predicted by UIP. This lends support to the conclusion that the negative bias of UIP is reduced by purging the endogenous component from the nid_t.

[4] Using the first lag generates a point estimate of 1.05 for the nid_t. We opted for the second lag because of the possible correlation between nid_{t-1} and the error term.

Table 1: **OLS Estimation of Equation (12)**

The dependent variable is ΔS^e_{t+1}

	Coefficient	Std. Error	t-prob
Constant	-0.009	0.016	0.548
nid	0.805	0.440	0.072
F(1,72): 3,334 [0.072]	n=74	R^2: 0.044	

Diagnostic Tests

AR 1-5 test:	F (5,67) =	20.755	[0.000]
ARCH 1-5 test:	F(1,72) =	3.334	[0.072]
Normality:	$\chi^2(2)$ =	40.816	[0.000]
Heteroscedasticity:	F(2,69) =	3.0538	[0.054]

Table 2: **IV Estimation of Equation (12)**

The dependent variable is ΔS^e_{t+1}

	Coefficient	Std. Error	t-prob
Constant	-0.022	0.017	0.205
nid	1.124	0.481	0.022
n=72			

Diagnostic Tests

AR 1-5 test:	F (5,65) =	20.931	[0.000]
ARCH 1-5 test:	F(5,60) =	10.952	[0.000]
Normality:	$\chi^2(2)$ =	60.935	[0.000]
Heteroscedasticity:	F(2,67) =	2.715	[0.073]

The earlier two models' lack of predictive power is due to the fact that the reduced form model better describes the nature of the observed data. The estimated error of the *ex ante* UIP presented serial correlation, as can be seen in the diagnostic tests in Table 1. We detected serial correlation of the first order - the autoregressive parameter is $\hat{\theta}_1 = 0.768$ with a t-probability of 0 and the constant is not significant - which justifies testing equation (19). Results presented in Table 3 show that the estimated model does not have problems of serial correlation, normality or functional form misspecification. The remaining problems are related to heteroscedasticity, both unconditional and time conditional. We decided not to deal with the problem of heteroscedasticity, because the reduced form parameters are a combination of the structural parameters, making inference senseless.

Table 3: **OLS Estimation of Equation (19)**

The dependent variable is ΔS^e_{t+1}			
	Coefficient	Std. Error	t-prob
Constant	0.002	0.003	0.455
ΔS^e_t	0.777	0.073	0.000
F(1,70): 113.2 [0.000]	n=72	R^2: 0.62	
Diagnostic Tests			
AR 1-5 test:	F (5,65) =	1.717	[0.143]
ARCH 1-5 test:	F(5,60) =	5.232	[0.005]
Normality:	$\chi^2(2)$ =	1.961	[0.375]
Heteroscedasticity:	F(2,67) =	19.961	[0.000]
RESET:	F(1,69) =	0.000	[0.990]

Conclusion

We showed that the simultaneity bias holds even when monetary authorities react to price changes, which complements the work of McCallum (1994), Kugler (2000) and implies that Christensen (2000)'s results are not conclusive evidence against the simultaneity bias hypothesis. Using IV techniques and data from the Brazilian Central Bank on exchange rate expectations, we also found results supporting *ex ante* UIP.

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