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

Alex Luiz Ferreira,
(University of São Paulo, Department of Economics, FEARP, Brazil[1])

Abstract

In this paper we test *ex ante* uncovered interest parity (UIP) using survey data of exchange rate expectations from 2001M11 until 2007M12, available at the Brazilian Central Bank. Using Ordinary Least Squares, we found that the estimated UIP parameter is smaller than one, which is a common finding of the literature. We then develop a model that explains how a negative bias can arise from the simultaneous actions between the Central Bank (through a policy reaction function) and speculators (through UIP). Our results using Instrumental Variables techniques show that the bias can be reduced, and lend support to UIP. The reduced form dynamically complete model provides the best fit for expected exchange rate changes, as it is supposed 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

[1]Address for correspondence: Av. Bandeirantes, 3900. CEP 14040-900, Ribeirão Preto, São Paulo, Brazil. Phone:+ 55 16 36023908, e-mail: alexferreira@usp.br.

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Introduction

Although support for uncovered interest rate parity (UIP, hereafter) has been growing, this hypothesis still brings some embarrassment from the empirical point view[2]. It predicts that a high (low) domestic interest rate in comparison to another country is associated with a depreciating (appreciating) exchange rate. However, several authors have found point estimates of the UIP parameter of the wrong sign implying that interest rates rise when agents expect a nominal exchange rate appreciation, which is counterintuitive. The widespread theoretical use combined with a mix of both supporting and undermining empirical evidence is puzzling and largely provides the motivation for the present paper. In this sense, as pointed out by Flood & Rose (2001), it is always easy to justify another look at the UIP.

There are competing explanations for the failure of short-run UIP (for example, risk, Peso problems, irrational agents, improper econometric techniques etc) but none seems to be widely acceptable and there is not a consensus on the subject. Hence, there is an open field for investigation and space to build on the direction of some sort of consensual explanation.

This paper develops on the work of McCallum (1994*b*) who put forward a model that recognizes the simultaneous action of agents (globalized speculators) and Central Bankers in determining equilibrium interest and exchange rates. The model assumes a function with interest rate smoothing and reaction against exchange rate changes, implying a policy driven failure of the hypothesis.

The endogeneity issue has already been investigated. Many authors have recognized the potential of this explanation for the UIP problem, for example, Meredith & Chinn (1998), Kirikos (2002) and Favero & Giavazzi (2004) (to cite a few) and to exchange rates, in particular, one can see Engel & West (2005). Kugler (2000) pioneered in seen the main implications of McCallum (1994*b*)'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*b*)'s policy reaction function. The reaction function in his model is characterized by the slow change in interest rates and resistance against exchange rate changes. One of his conclusions is that interest rate smoothing and leaning against the wind (reaction against exchange rate changes) lead to a negative relationship of the spot exchange rate change and the lagged forward premium and, thus, a rejection of the UIP. On the other hand, Christensen (2000) tested the policy reaction function of the McCallum (1994*b*) model for a sample of developed countries (the US, Germany and Japan) but did not find evidence supporting the significance of the parameters needed to generate the negative bias.

We aim to contribute to this literature by testing *ex ante* UIP for Brazil and taking into account the endogeneity problem. We develop a simple macroeconomic model that does not hinge on the assumption of "leaning against the wind" and show that reaction against prices can be enough to generate the bias on UIP[3]. Furthermore, we show the

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

[3]As a matter of fact, McCallum (1994*a*) had already recognized that a more general reaction function

associated asymptotic bias and provide a hint on the size of the structural parameters needed to generate a negative bias. We also performed Ordinary Least Squares (OLS) and Instrumental Variable (IV) estimation of *ex ante* UIP. Results show that the IV estimation reduces the bias; there is evidence supporting *ex ante* UIP and that the dynamically complete model, which better represents the observed data (i.e., when variables are in equilibrium), produces the best fit.

The rest of the paper is organized as follows. We first present a model with UIP and a policy reaction function that takes into account the fact that Central Banks react to deviations of expected inflation from target and instrument smoothing. Two subsections discuss both the asymptotic bias and the reduced form dynamically complete model for expected exchange rate changes. The penultimate section is dedicated to the empirical part and the final to conclusions.

Endogeneity

We conclude from McCallum (1994*b*)'s article that the empirical failure of short-run UIP is due to the non-recognition that this hypothesis (concerning equilibrium in the assets market), in fact, belongs to a system of equations. Hence, a regression of UIP using OLS would produce estimated parameters that cannot have a structural interpretation and could also be subjected 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 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*b*) and Engel & West (2005), for instance, but without recurring to the explicit inclusion of exchange rates in the policy function [a possible justification is the not supportive result presented by Christensen (2000) on "leaning against the wind"]. The model presented here describes a simplified open economy, as opposed to 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 other last 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:

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

- including reaction against prices - would be theoretically plausible and could be empirically stronger.

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

$$\pi_t - \pi_t^* = \pi_{t+1}^e - \pi_{t+1}^{e*} + \eta_1(h_t - h_t^*) + e_t^s, \tag{3}$$

$$h_t - h_t^* = -\eta_2[i_{t-1} - i_{t-1}^* - (\pi_t - \pi_t^*)] + e_t^d, \tag{4}$$

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 it is often done in the international finance literature (this way of thinking will also provide an intuitive interpretation of the model results). For simplification, we start with the assumption that ξ_t is identically and independently distributed. The variable π_t stands for the inflation between $t-1$ and t, while π_{t+1}^e is the one period ahead inflation forecast (made at t) 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_t^n is the neutral interest rate, i.e. $i_t^n = r_t + \pi^T$. The letter r_t represents the equilibrium real interest rate, which is determined by 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 - we will come this point later on this analysis. The log of the output gap is represented by h_t. The error terms e_t^s and e_t^d stand for a supply and demand shock, respectively, and are both random variables. The other letters are parameters. For example, ρ is the smoothing parameter, $0 < \rho < 1$; 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[4], 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

[4] 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.

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

where μ_t is the forecast error of exchange rate depreciation. The former assumption is reasonable if we consider rational expectations UIP under imperfect capital mobility and the process for the expected inflation differential determining the behavior of the real interest rate differential such as the one given by (5). 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. This is more easily seen if one considers $\rho = 0$, $\pi_{t+1}^e = \pi_{t+1}^{*e}$, $h_t = h_t^* = 0$ and no shocks. In this case, if the monetary authority sets $nid_t = \xi_t$, there will not be any capital inflows or outflows, i.e. the UIP condition will be satisfied without any expected change in exchange rates. So when the instrument used for targeting inflation is the interest rate, the rule (in this model) 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

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

which solved for $\pi_t - \pi_t^*$ results

$$\pi_t - \pi_t^* = \frac{\eta_1\eta_2 nid_t + \phi\Delta s_t + \eta_1 e^d + e_t^s}{1 - \eta_1\eta_2}. \tag{8}$$

We then substitute (8) and (6) into (2)

$$nid_t = \rho nid_{t-1} + (1-\rho)\left(\xi_t + \mu_t + \lambda\frac{\eta_1\eta_2 nid_t + \phi\Delta s_t + \eta_1 e^d + e_t^s}{1 - \eta_1\eta_2}\right), \tag{9}$$

and we can solve the expression for the nid_t by writing

$$nid_t = \alpha_0 nid_{t-1} + \alpha_1\Delta s_t + \alpha_2\xi_t + e_t, \tag{10}$$

where

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

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

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

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 (10) into equation (1) and solving for expected exchange rate changes gives

$$\Delta S^e_{t+1} = \alpha_0 nid_{t-1} + \alpha_1 \Delta s_t + (\alpha_2 - 1)\xi_t + e_t. \tag{11}$$

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

$$\Delta s_t = \gamma_0 nid_{t-1} + \gamma_1 \xi_t + \gamma_2 e_t. \tag{12}$$

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[5]

$$nid_t = \frac{\lambda\phi(\rho - 1)}{\rho(\lambda\phi - 1) + [\lambda + \rho(1-\lambda)]\eta_1\eta_2 - \lambda\phi}\xi_t \tag{13}$$

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

Asymptotic Bias

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

$$\Delta S^e_{t+1} = \beta_0 + \beta_1 nid_t + \epsilon_t \tag{14}$$

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

$$\text{plim}(\hat{\beta}_1) = \beta_1 + \frac{\text{Cov}(nid_t, \epsilon_t)}{\text{Var}(nid_t)}. \tag{15}$$

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

$$\text{Cov}(nid_t, -\xi_t) = -\mathbb{E}(nid_t\xi_t), \tag{16}$$

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 positive correlation is reflected into a negative covariance in the UIP structural equation leading to the statistical bias. Suppose, for instance, that $\mathbb{E}(\xi_t^2) = \sigma_\xi = 1$, $\lambda = 1.5$, $\rho = 0.6$, $\phi = 0.5$, $\eta_1 = 0.2$ and $\eta_2 = 0.2$, which

[5] Detailed results can be obtained with the author upon request.

are parameter values that replicate the typical negative bias found in the literature. Then $\mathrm{Cov}(nid_t, \epsilon_t) = -0.35$, and we can finally write

$$\mathrm{Bias} = \frac{\rho(\lambda\phi - 1) + [\lambda + \rho(1-\lambda)]\eta_1\eta_2 - \lambda\phi}{\lambda\phi(\rho - 1)} \approx -3.$$

The estimator of β_1, named $\hat{\beta}_1$, will be

$$\mathrm{plim}(\hat{\beta}_1) \approx \beta_1 - 3, \tag{17}$$

where plim is the limit of the probability when the sample size grows to infinity. Since the β_1 of the "population" is equal to one, the estimated parameter will be biased towards -2, according to the simple model above and for the given parameter values. This suggests, for instance, that instrumental variables techniques are more appropriate for UIP tests. The reason is that they purge the endogenous component from the *nid*, provided one can have proper instruments. In the empirical part of the paper, we present two stage least squares estimations of *ex ante* UIP. 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, which will also be tested.

The dynamically complete model for the expected exchange rate

Substituting the process for nid_t from (13) into (1) and solving for expected exchange rate changes gives

$$\Delta s^e_{t+1} = (\kappa - 1)\xi_t, \tag{18}$$

where

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

Solving (18) for ξ_t and writing the result with one lag results

$$\xi_{t-1} = \frac{1}{\kappa - 1}\Delta s^e_t. \tag{19}$$

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[6]. We introduce some dynamics in the reduced form by assuming serial correlation of the AR(1) type

$$\xi_t = \theta_0 + \theta_1\xi_{t-1} + \zeta_t, \tag{20}$$

where ζ_t is white noise. Substituting (19) into (20) , generates

[6]As mentioned earlier, our earlier assumption that ξ_t was independently distributed was intended to simplify the analysis for the calculation of the bias.

$$\xi_t = \theta_0 + \frac{\theta_1}{\kappa - 1}\Delta s_t^e + \zeta_t, \tag{21}$$

and we finally substitute (21) into (18)

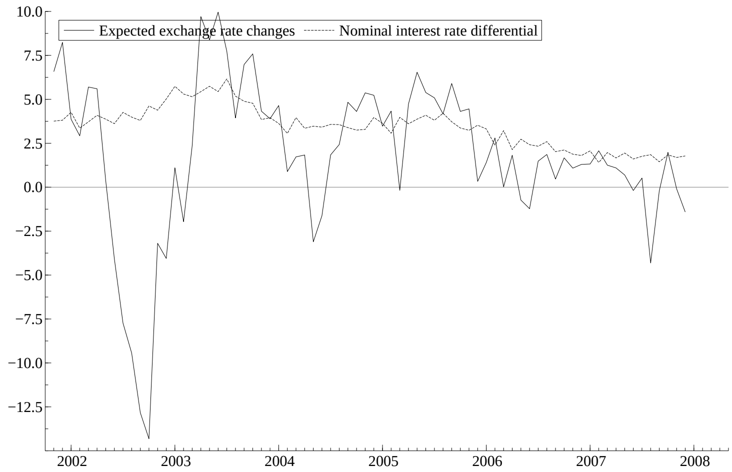
$$\Delta s_{t+1}^e = \theta_0(\kappa - 1) + \theta_1 \Delta s_t^e + \frac{1}{\kappa - 1}\zeta_t. \tag{22}$$

The equation above represents the dynamically complete model of the reduced form. As it will be discussed below, our tests showed that there was no more serial correlation after estimating the model in (22).

Empirical Results

We were able to test for *ex ante* UIP because of data availability on expected changes in exchange rates. The use of expected exchange rate changes as the dependent variable also avoided problems regarding the time series properties of the expectational error. Data from exchange rate surveys was obtained from the internet site of the Brazilian Central Bank (known as the Focus data - Focus is the agency that collects expectations from the market). On the other hand, data on interest rates was taken from the Institute of Applied Economic Research (Ipea) and corresponds to the Brazilian Selic and the North-American Treasury Bill with three-month maturity. The Selic was transformed into a three-month rate while the expected exchange rate change consistent with the interest rate of month t, for instance, 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_{t+1}^e 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 at that period, largely caused by a hike in default risk due to uncertainty in the presidential elections, possibly provoked an overshooting of the currency that was reflected into a future foreseen appreciation. However, the nominal interest rate differential remained relatively stable (in fact, it increased) which might reflect the option of money authorities in reacting positively to the increase in risk.



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

We first estimate equation (14) using OLS and found a parameter close to 1. In fact, the 99% confidence interval contains 1 which is already a surprising result favoring 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.

Furthermore, we tested *ex ante* UIP using instrumental variables techniques, i.e. two stage least squares using lags of the dependent variable as instruments. Results presented in Table 2 show that the parameter is larger than the OLS one 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.

Table 1: **OLS Estimation of Equation (14)**

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]

The lack of predictive power of the earlier two models is due to the fact that the reduced form model better describes the nature of the observed data. However, we tested another model with the lagged dependent variable in the specification and without the

Table 2: **Instrumental Variable Estimation of Equation (14)**

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]

† We employed two stage least squares, using the second lag of the nid_t as instrument.
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.

nid. The reason is that 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 $\hat{\beta}_1 = 0.768$ with a t-probability of 0 and the constant is not significant. This finding was taken into consideration by modeling the reduced form with dynamics according to equation (22). Results are shown in Table 3. The estimated model does not have problems of serial correlation, normality and functional form misspecification. The remaining problems are related with heteroscedasticity, unconditional and time conditional. As heteroscedasticity would be problematic to inference and because the reduced form parameters are a combination of the structural parameters (so inference would not be interesting), we decided not to deal with this problem.

Table 3: **OLS Estimation of Equation (22)**

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

In this paper we found support for *ex ante* UIP and built a model that seems to represent the data generation process of the observed data. Our model shows that a bias on UIP estimations can arise when there is a simultaneous relation between agents and Central Banks. Our model does not rely on the hypothesis of leaning against the wind but it is based on a monetary policy reaction function with smoothing. We showed that an expectational pass-through and a time varying neutral real interest rate can explain the need of monetary authorities to react to UIP shocks and thus we were able to unveil the correlation between these shocks and nominal interest rates.

The main contributions to the international finance literature are: *i*) showing that the simultaneity bias holds even when monetary authorities react to price changes, which complements the work of McCallum (1994*b*), Kugler (2000) and implies that Christensen (2000) results are not concluding evidence against the simultaneity bias hypothesis; *ii*) testing *ex ante* UIP using data on exchange rate expectations of the Brazilian Central Bank and instrumental variables techniques while finding supportive results.

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