QUASI PURCHASING POWER PARITY: STRUCTURAL CHANGE IN THE MEXICAN PESO/US DOLLAR REAL EXCHANGE RATE

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Resumen: Analizamos si el tipo de cambio real peso/dolar se revierte a un valor de equilibrio de largo plazo, y si este valor es único. Utilizamos un método para verificar estacionariedad que permite un número desconocido de cambios estructurales en el nivel de la serie. Al utilizar datos anuales (1925-1994), nuestros resultados proveen evidencia en favor de la *cuasi* paridad del poder adquisitivo. En particular, encontramos que el tipo de cambio real peso/dolar ha fluctuado estacionariamente alrededor de un nivel de largo plazo durante 70 años, perturbado por una serie de eventos, domésticos y externos, durante o alrededor de 1981.

Abstract: This paper analyzes whether the real exchange-rate of the Mexican peso/US dollar revert to a long-run equilibrium value, and whether this value is unique. We use a method for testing stationarity, that allows for an unknown number of structural breaks in the level of the series. Using a long span of annual data covering the period 1925-1994, our results provide evidence favoring long-run Quasi-Purchasing Power Parity. In particular, we find that the real peso/dollar exchange rate has fluctuated stationarily around a 70 year long-run level, perturbed by a series of events, both domestic and external, in or around 1981

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1. Introduction

The issue of whether real exchange rates, RER, revert to a long-run equilibrium value has been a widely researched area in international finance during the last decade. Mean reversion in this context implies that relative prices -valued in a common currency- tend to converge over long spans of data, thus supporting the doctrine of Purchasing Power Parity, PPP).[1] This parity doctrine is central to many theoretical models of exchange rate determination.

It is common practice in the literature to apply unit root tests to investigate whether the RER reverts to its (equilibrium) long-run mean. Following the influential paper of Perron (1989), there are a number of studies showing the relevance of allowing mean shifts in modelling the long-run behavior of RERs. See, for instance, Corbae and Ouliaris (1991), Perron and Vogelsang (1992), Culver and Papell (1995), and Baum, Barkoulas, and Caglayan (1999). In this literature, the number of structural breaks allowed in the deterministic trend function is fixed *a priory*, based mainly on visual inspection of the data. For many of the real exchange rates series analyzed in the above papers, it is not unambiguous how many significant structural breaks have occurred within the sample. Hegwood and Papell (1998) argue that rejection of a unit root in real exchange rate data only implies that PPP holds in the absence of structural breaks. This means that PPP requires reversion to a *constant* mean. In their empirical investigation (which includes several long annual periods of real exchange rates), they use a two step procedure. After they establish that the RERs are stationary using Augmented Dickey-Fuller, ADF, tests, they apply a sequential test for structural breaks, developed by Bai and Perron (1998a), to find that there are indeed multiple structural breaks in most of the RERs analyzed.[2] These findings led them to conclude that the series revert to an occasionally changing mean, and called this phenomenon Quasi-PPP.

In this paper, we test the stationarity of the Mexican peso/US dollar RER allowing for an unknown (endogenously determined) number of structural breaks in the level of the series. Although the application of a 'standard' ADF test would indicate rejection of a unit

[1] It is difficult to expect PPP to be valid in the short-run, due to trade barriers, transaction costs, foreign exchange market interventions, etc. These factors affect the basic assumption of perfect intercountry commodity arbitrage.

[2] This implies, however, that none of the identified breaks were sufficiently strong so as to induce unit root behaviour in the series.

root, as in the case of the series analyzed by Hegwood and Papell (1998), there is strong evidence of a major change in the long-run behaviour of the Mexico RER, starting around the beginning of the 1980s. This has not been taken into account in previous studies concerning the peso/dollar RER. Mexico's internal and external economic environment was particularly interesting during those years. In 1979, the government adopted a model based on oil exports, following the oil field discoveries of 1978 and the 150% increase in oil prices the following year. However, this oil-based strategy ended with a decrease in oil prices in 1981, leaving the country with an enormous external debt, which had been contracted to develop the oil industry. As documented in Aspe (1993), the chronology of the financial crises begins with the worsening of Mexico's terms of trade around the middle of 1981, mainly as a result of the decline in oil prices. Then, in 1982 increases in international interest rates accelerated capital outflows. The macroeconomic adjustment of 1982 implied a 500% nominal devaluation of the peso (from 25 to 150 pesos per dollar), while the inflation rate rose from 29% to nearly 100%. By 1982 the RER had depreciated 272% with respect to the previous year.

We utilize a long span of data for the peso/US dollar RER, covering the period 1925-1994.[3] The evidence on the stationarity of the RER between Mexico and the US is mixed thus far. Avalos and Hernández (1995), do not find evidence against a unit root over the period 1961-1994 using both annual and quarterly data. In Mejía and González (1996), the unit root hypothesis is marginally rejected using an Augmented Dickey-Fuller test and annual data over the longer period 1940-1994; similar conclusions are reached in Galindo (1995). In these papers there is no allowance for structural breaks in the data. Our results indicate that the peso/US dollar RER is better modeled as a stochastically stationary AR process around a long-run level perturbed by a single structural break, in 1981. Along the lines of Hegwood and Papell (1998), this implies that Quasi-PPP holds.

The next section presents the econometric methodology, based on the procedures and methods in Bai (1997b), Bai and Perron (1998a, 1998b), and Noriega and Ramírez-Zamora (1999). Section 3 presents

[3] The data source is Alzati (1997), who constructs a series of RERs for the period 1895-1994. He argues, however, that some data points along the period (1910-1920) could be extremely distorted by effects of the Mexican Revolution. In 1925 the central bank (Bank of Mexico) was established, and with it the generation of official statistics. We chose 1994 as the final year due to the potential break occurring in 1995 (following the peso devaluation in late 1994), leaving very little data points afterwards to identify it.

and discusses results for the Mexico/US RER. Finally, section 4 summarizes final comments.

2. Econometric Methodology

The procedure for testing for the presence of a unit root with an unknown number of structural breaks in the deterministic trend function, based on the Unit Root Rejection Stopping Rule, URR-SR, works as follows (for details see Noriega and Ramírez-Zamora, 1999). Denoting by Y_t the logarithm of the observed real exchange rate series, we first estimate (by OLS) the following Mean Stationary MS and Difference Stationary, DS models, respectively:

$$\Delta Y_t = \mu + \sum_{i=1}^{m} \theta_i DU_{it} + \alpha Y_{t-1} + \sum_{i=1}^{k} a_i \Delta Y_{t-i} + \varepsilon_t, \qquad (1)$$

$$\Delta Y_t = \sum_{i=1}^{k} a_i \Delta Y_{t-i} + \varepsilon_t, \qquad (2)$$

for $t = 1, 2, ...T$, where T is the sample size, ε_t is an iid process, and DU_{it} is a dummy variable allowing changes in the mean's level, that is, $DU_{it} = \mathbf{1}(t > T_{b_i})$, where $\mathbf{1}(\cdot)$ is the indicator function and T_{b_i} is the unknown date of the i^{th} break. In the MS model (1), Quasi-PPP holds whenever $-2 < \alpha < 0$, in which case Y_t fluctuates stationarily around a deterministic level μ, (possibly) perturbed by m level shifts. Under the DS specification (2), $\alpha = 0$ (the null hypothesis), and the real exchange rate behaves like a random walk, implying that PPP does not hold. In determining the autoregressive order k for each model, we use the k – max criterion, as in Noriega and Ramírez-Zamora (1999) and Perron (1997). In order to discriminate between these two models, we simulate the distribution of the t-statistic for the null hypothesis of a unit root ($\alpha = 0$ in (1)), called $\widehat{\tau}$, under the hypotheses that the true models are the MS model (1) and the DS model (2), both estimated from the data.[4] We call these empirical densities $f_{MS_m}(\widehat{\tau})$, $(m = 0, 1, 2, ...)$ and $f_{DS}(\widehat{\tau})$, respectively.

[4] We use 10,000 replications for each model. A similar approach is used by Kuo and Mikkola (1999), who use bootstraped critical values, based on stationary and non-stationary ARIMA models fitted to the US/UK real exchange rate series. However, they do not consider the case of structural breaks in the trend function.

For determining the location of breaks, the criterion we use chooses, among all possible combinations of m break dates, the one which yields the smallest residual sum of squares (called min RSS) from (1). This is done for all values of $k \leq k\max$. As in Bai and Perron (1998b), we utilize a dynamic programming algorithm to obtain global minimizers of the RSS.[5] Note that this criterion implies simultaneous determination of m breaks via a global search.

In order to determine the number of breaks, we equip the above procedure with the URR-SR, which indicates the termination of the search. Under the URR-SR, we proceed sequentially: after we estimate equation (1) with $m = 0$, the relevance of both the null (a unit root) and alternative (a MS model with $m = 0$) hypotheses are analyzed in terms of the position where the sample estimate of the t-statistic for testing a unit root ($\widehat{\tau}_{sample}$) lies relative to the empirical densities of $\widehat{\tau}$ under the estimated MS model (1) and DS model (2). If as a result it is concluded that the null hypothesis can not be rejected, or that it is not possible to discriminate between hypotheses, then we allow the procedure to search and locate one structural break in the level of the series, and the relevance of both the null of a unit root and the alternative of a MS model with a single structural break is analyzed. This process continues until the null hypothesis is rejected and the alternative hypothesis most supported by the data is found. After the search finishes, we suggest analyzing the results from allowing one additional break. That is, comparing the relevance of both the null and alternative hypotheses under two different trend specifications. As can be seen, this is a sequential procedure which globally searches for an increasing number of structural breaks.[6]

3. Results and Discussion

We first present results obtained from the application of the URR-SR. This results are then compared to those obtained from the application

[5] With thanks to Pierre Perron for providing us with his GAUSS code, which was adapted for this study.

[6] Some authors have used versions of this rule in empirical applications (for the case of models allowing for up to two breaks in the trend function): Clemente, Montañés and Reyes (1998), Ohara (1999), Mehl (2000), Aggarwal, Montañés and Ponz (2000). Arestis and Biefang-Frisancho Mariscal (1999) conclude that "...unit root tests that do not account sufficiently for the presence of structural breaks are misspecified and suggest excessive persistence" (p. 155).

of the Parameter Constancy Stopping Rule (PC-SR, based on Bai, 1997b).

The empirical results are presented in table 1. The first column indicates the number of breaks allowed in the trend function under the alternative hypothesis, m. Column 2 reports the value of the estimated value of k (starting from an upper value of $k \max = 10$). Column 3 reports the estimated break dates under the min RSS criterion. Columns 4-6 report, respectively, the Akaike Information Criterion, AIC, the sample estimate of the t-statistic for testing a unit root ($\widehat{\tau}_{sample}$), and the standard error of regression. The last two columns report the rejection probabilities of difference stationary and mean stationary models for the real exchange rate data, using exact critical values based on the Monte Carlo distributions of the Dickey-Fuller type $t-$statistic. These values indicate the position where the sample estimate of the t-statistic for testing a unit root ($\widehat{\tau}_{sample}$) lies relative to those distributions. To draw exact inference on the unit root hypothesis through $\widehat{\tau}_{sample}$, we calculate, under each density, the probability mass to the left of $\widehat{\tau}_{sample}$, denoted $\Pr[\widehat{\tau} \leq \widehat{\tau}_{sample} \mid f_{DS}(\widehat{\tau})]$, and $\Pr[\widehat{\tau} \leq \widehat{\tau}_{s}ample \mid f_{MS}(\widehat{\tau})]$, respectively.

From the reported probabilities based on $\widehat{\tau}_{sample} = -3.78$ (with $m = 0$), we can conclude that it is very unlikely that this estimated value of the t-statistic for testing a unit root in the RER could have been generated by a DS model. On the other hand, the probability associated with the MS model (69.5%) indicates that this specification is much more plausible.

The disproportionate changes observed in the nominal exchange rate, and the relative price indices in the early 80s, led us to apply the procedure for testing the null of a unit root against the alternative of stationary fluctuations around a level perturbed by one structural break. As reported in the second row of table 1, the min RSS criterion selects 1981 as the break date, with $\widehat{k} = 5$. The corresponding t-statistic for testing a unit root ($\widehat{\tau}_{sample}$) is -6.43, and the p -values in the last two columns show a clear rejection of the DS model in favor of the MS model with a single structural break in the level of the series. In fact, the probability under the MS model with one structural break lies nearly in the middle of the empirical distribution (0.48), suggesting that this specification is even more plausible than the MS one without a structural break. Additionally, both the AIC and the standard error of the regression indicate a better fit for the model allowing for a single structural break.[7]

[7] It should be noted, however, that this break was not strong enough to induce

Table 2 reports results of the application of the parameter constancy stopping rule.[8] Over the entire sample (1925-1994) a significant break is identified in 1981 for the peso/US dollar RER. Upon dividing the sample into two subsamples separated by this break, no additional significant breaks are found by the procedure. The table also shows a not-very-tight 95% confidence interval for the break date. Note that the break date is the same as the one obtained under the 'unit root rejection' stopping rule.

Hence, from the results of applying the URR-SR, we can conclude that the peso/US dollar RER is better modeled as a stochastically stationary AR process around a long-run level perturbed by a single structural break, implying that Quasi-PPP holds. Since we are able to reject the unit root hypothesis for our data, the restrictive dynamic structure of the adjustment process relating nominal exchange rates and relative price indices implied in unit root tests, as discussed in Steigerwald (1996), is not binding in our case. The estimated break date under this procedure, 1981, is confirmed using the parameter constancy stopping rule.

4. Conclusions

This paper has shown that the Mexican peso/US dollar real exchange-rate does revert to a long-run equilibrium value. Our results show, however, that this value underwent an upward level shift during 1981. According to some authors, this date coincides with the worsening of Mexico's terms of trade, mainly as a result of the decline in oil prices. The macroeconomic adjustment of 1982 implied a 500% nominal devaluation of the peso, which translated into a 272% depreciation of the real exchange rate, with respect to the previous year. Our results provide evidence favoring long-run Quasi-Purchasing Power Parity, and imply that it is possible to separate a stationary cycle for the real exchange rate from a long-run deterministic level. In particular, the peso/US dollar RER has fluctuated stationarily around a 70 year

unit root behaviour in the data.

[8] In table 2, the trimming parameter, π, is selected such that $k+3 \leq T_{b_1} \leq T-3$, that is, $\pi \times T_s = 3$, where T_s represents either the sample size, or the size of a subsample (see Andrews, 1993). For example, for the full sample of the real exchange rate in the table, 1925-1994, we have 70 observations, and $\pi \times 70 = 3$ implies $\pi = 0.043$. Tests were also carried out for the case $\pi \times T_s = 6$. We obtained the same qualitative results.

long-run level, perturbed by a series of events, both domestic and external, in or around 1981.

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Table 1

Broken Trend Models for Real Exchange Rate Mexico/US 1925-1994

$$\Delta Y_t = \mu + \sum_{i=1}^{m} \theta_i DU_{it} + \alpha Y_{t-1} + \sum_{i=1}^{k} a_i \Delta Y_{t-i} + \varepsilon_t$$

m	$\hat{k}$	$\widehat{T}_b$	AIC	$\widehat{\tau}_{sample}$	$\hat{\sigma}_\varepsilon$	$\Pr[\widehat{\tau} \leq \widehat{\tau}_{sample} \mid f_{DS}(\widehat{\tau})]$	$\Pr[\widehat{\tau} \leq \widehat{\tau}_{sample} \mid f_{MS}(\widehat{\tau})]$
0	0	...	-0.3041	-3.78	0.2048	0.0098	0.6953
1	5	1981	-0.4917	-6.43	0.1785	0.0007	0.4809

Table 2

Test Statistics, Break Points and Confidence Intervals

				Critical Values			*Confidence Interval*
Sample	$Sup - Wald$	$\widehat{T}_b$	π	*90%*	*95%*	*99%*	*95%*
1925-1994	11.30	1981	.042	9.84	11.53	16.73	(1978-1984)
1925-1981	5.26	1976	.0526	12.17	14.54	18.91	
1982-1994	7.75	1987	.2307	11.07	14.60	27.85	