FUNDAMENTALS OF REAL EXCHANGE RATE DETERMINATION: WHAT ROLE IN THE PESO CRISIS?

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Resumen: Se investiga el papel desempeñado por factores estructurales en el comportamiento histórico del tipo de cambio real en México. Se demuestra que determinantes "fundamentales" (que se omitieron en investigaciones anteriores) son cointegrados con el tipo de cambio nominal y los precios relativos, mientras una prueba de la versión sencilla de PPA (sin los factores fundamentales) falla. La ecuación de cointegración indica una subvaluación severa a finales de los 1980s y una sobrevaluación ligera antes de la devaluación de diciembre de 1994. Los factores fundamentales no son capaces de explicar la magnitud del choque que recibió México en aquel momento.

Abstract: We examine the role of structural factors in the Mexican real exchange rate experience since 1970, particularly in the crisis of December, 1994. We find that "fundamental" determinants of the real exchange–omitted from previous research–are cointegrated with nominal exchange rates and relative prices, while tests of PPP alone fail. The cointegrating equation indicates a severe undervaluation during the 1980s and only modest overvaluation in the period immediately preceding the devaluation in December, 1994. Nothing in the fundamentals can account for magnitude of the blow Mexico suffered at that time.

Clasificación JEL: F31

Palabras clave: tipo de cambio real, paridad de poder adquisitivo, crisis cambiaria, cointegración, Real Exchange Rates, Purchasing Power Parity, Exchange Rate Crisis, Cointegration.

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1. Introduction

This paper revisits the remarkable Mexican real exchange rate experience since 1970. Our aim is to understand the role of structural factors that have largely been omitted from prior research on the subject.[1] We also analyze the extent to which changes in these fundamentals may have played a role in the crisis that began in December, 1994.

While the empirical focus here is the recent Mexican experience, the economic problem under study generalizes broadly. Cycles of gradual overvaluation and sudden correction continue to plague the world's independent currencies, and the speculative attacks themselves frequently do lasting harm to the real economies of the countries affected.[2]

The problem is compounded by great uncertainty about when a currency is overvalued and when it is not. It seems at least plausible that less uncertainty with regard to the fundamental determinants of equilibrium nominal and real exchange rates might reduce the severity of speculative attacks, but what reasonable guidance might economists give policy makers on this score? Simple versions of purchasing power parity (PPP) theory have long since been thought dubious, both on theoretical and empirical grounds. But while more

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[1] Time series tests of Mexican real exchange rate behavior that ignore structural determinants (though they may allow for "structural breaks") include: Noriega and Medina (2003), Galindo (1995), Ghoshroy-Saha and Van Den Berg (1996), Mejía and González (1996). Ávalos and Hernández (1995) reject a stand-alone purchasing power parity theory of real exchange rates on the basis of unit root and cointegration tests, but find an important role for structural factors on the basis of standard regression techniques. While these latter results are plausible–and in fact anticipate some of what we report here–their statistical inference is plagued by the problem of spurious regression, arising from the use of non-stationary regressors. Nevertheless, their principal finding –that real exchange rate variability is negatively related to GDP growth– is valuable as well as statistically valid.

[2] Debt deflation among domestic borrowers (frequently banks) leaves bankruptcies that are not easily liquidated, as well as a political stalemate between foreign creditors (on the one hand) eager to see the losses socialized and paid off and domestic taxpayers (on the other hand) loathe to assume the privately contracted debts of their countrymen. Public sector debts raise the specter of further monetization of deficits, inflation, and devaluation. International creditors, whether parties to the original collapse or not, are understandably reluctant to lend, facing the likelihood that any credit they extend will immediately go to paying off creditors at the front of the queue, while they themselves are consigned to the rear. Dornbusch, *et al.* (1995) provides a good discussion.

nuanced theories of real exchange rate determination are now common even in undergraduate textbooks,[3] empirical estimations most often proceed as if this weren't the case.[4]

2. PPP, Yet Again

There are now decades of empirical work attempting to measure the degree to which international arbitrage in goods markets anchors the domestic price level to the nominal exchange rate. In the balance hang questions of whether PPP is a valid closure rule in theoretical models –and if so, over what time horizon– and more quotidian questions of whether a specific nominal exchange rate is in fact over- or undervalued at a given moment in time.

The specification of PPP theory for the purposes of empirical estimation has nearly always involved some variant on the following equation:

$$s_t = p_t - p_t^* + u_t \tag{1}$$

In equation 1 (all in logs) the nominal exchange rate equates to the relative price of domestic output plus an error term. Advances in time series econometrics in the 1970s showed that statistical tests of the theory using data in levels would find spurious correlations, since the nominal exchange rate, domestic price levels, and foreign price levels are very unlikely to be stationary variables.[5] Tests of PPP in the 1980s frequently used OLS on first differences of the data. While this slightly alters the theory under scrutiny,[6] it will at least generate consistent parameter estimates when the differenced data are in fact

[3] See for example Krugman and Obstfeld (2003), Chapter 15.

[4] The last decade has witnessed a great deal of effort directed towards raising the power of statistical tests sufficiently to reject the null hypothesis of a unit root in the real exchange rate data, in order to rescue some semblance of standard PPP theory. These efforts include assembling ever longer time series, bringing the power of panel estimation to bear, and allowing for unexplained structural breaks. Relevant surveys of this literature include Froot and Rogoff (1996), and Edwards and Savastano (1999).

[5] Granger and Newbold (1974).

[6] Here relative price levels need not *equate* with the nominal exchange rate (due to differences in preferences and measurement, to distortionary tariffs and transport costs, etc.) but they should move together over time. This is *relative* PPP, as opposed to *absolute* PPP.

stationary. Further advances in time series econometrics in the 1980s showed that if there is a long-run equilibrium relationship between variables, they should be cointegrated, which is to say that some linear combination of these nonstationary variables should be stationary.[7] For more than a decade, now, efforts to test PPP have focused on rejecting the unit root in the real exchange rate series directly; or, barring that, on finding cointegration among the variables in (1), allowing, possibly, for sluggish adjustment to shocks.

When the data fail to support a version of PPP according to either of these methods –which they frequently do– it has nearly always occasioned a call for stronger statistical tests to uproot the anomaly (see footnote 4). At the risk of belaboring the obvious, however: insufficient statistical power is not the only possible explanation for failure to find mean-reversion in the real exchange rate data. It could as well be that the omission of other relevant variables implies reversion, albeit to a moving target.[8]

Unfortunately, the parsimonious specification of equation (1) ignores all of the structural determinants of *real* exchange rates that the theoretical literature identifies. PPP is grounded on the microeconomics of arbitrage (the famous "Law of One Price"), but it has been well understood at least since the 1960s that the presence of nontraded goods –for which no international arbitrage exists– can lead to systematic movements in real exchange rates inconsistent with PPP (Balassa, 1964; Samuelson, 1964).[9] It follows that economic phenomena that impact the relative price of nontraded goods must be included in models of real exchange rate determination.

Theoretical models and empirical tests have established the important determinants of nontraded goods prices.[10] The best understood and most frequently cited in the literature include: i) differential rates of productivity growth (Balassa , 1964; Samuelson, 1964);

[7] Engle and Granger (1987).

[8] Hegwood and Papell (1998) carry out tests of "quasi" PPP, with reversion to a mean that changes as a result of structural breaks. They find that such a specification results in much faster mean reversion than the literature typically finds. The structural breaks themselves, however, remain unexplained.

[9] Since non-traded goods play an important role in consumer price indices (and a lesser –but still positive– role in wholesale price indices), factors which inflate domestic non-tradeable goods prices will also inflate the overall index relative to the foreign index.

[10] Edwards (1999) and Froot and Rogoff (1996) provide comprehensive literature reviews.

ii) changes in the pattern of demand or supply, due to changes in resource endowments, preferences, or levels of government expenditure (Kravis and Lipsey, 1983; Bhagwati, 1984);[11] and iii) changes in the terms of trade (Marion, 1984), *inter alia*).

If these structural determinants are sufficiently stable over time, their omission will not complicate the life of the econometrician.[12] If, however, they vary importantly over the sample, the parsimonious specification above will be flawed. Failure to reject a unit root in real exchange rates, or to find cointegration among relative prices and the nominal exchange rate could reflect precisely the fact that these variables evolve in consort with the structural variables that determine the real exchange rate over longer horizons.

3. Modeling Bilateral US-Mexico Real Exchange Rates

A unit root test performed on the Mexican real exchange rate for the period 1969-2000 fails to reject a unit root under a wide variety of lag specifications.[13] Failure to reject the unit root is the point of departure for the empirics of this paper. A generalized theory of PPP should allow for the possibility that structural variables impact real exchange rates. As in Dibooglu (1996) we posit such a long-run equilibrium relationship between the variables in the following equation:[14]

[11] Within the latter category one might reasonably consider changes in relative levels of government spending, comprising, as these do, a change in the composition of demand. Government spending is generally thought to fall more heavily on non-traded goods, thus tilting aggregate demand in that direction. As these are inelastically supplied (unlike traded goods), the relative price must rise in consequence of this shift in demand. An elaboration of this point appears in Froot and Rogoff (1996).

[12] They may, nevertheless, complicate the life of the policy maker: if such a model were employed out-of-sample, during a period when the structural variables actually began to vary, the model's guidance would be misleading.

[13] Data sources are described below. The sample was chosen to maximize the data available for estimation of equation (2). The nearest rejection stems from the Ng and Perron (2001) m-test which finds a p-value of 11%.

[14] We focus consciously on long-run structural determinants of real exchange rates, to the exclusion of variables reflecting current and capital account rigidities and imperfections, and strictly monetary phenomena. Our justification for this is that changes in these variables should ultimately be reflected in the relative price indices themselves, at least in the long-term. For an example of a modeling strat-

$$\beta_0 + \beta_1 s_t + \beta_2 (p_t - p_t^*) + \beta_3 (pr_t - pr_t^*)$$
$$+\beta_4 (g_t - g_t^*) + \beta_5 (rpo_t) = 0 \qquad (2)$$

With all variables in logs, and the asterisk denoting "foreign" (in this case, US), we posit a long-run equilibrium relationship between the nominal peso-dollar exchange rate (S_t) and various determinants including the relative price of domestic (Mexican) output ($p_t - p_t^*$), the relative productivity of the domestic tradeable goods sector ($pr_t - pr_t^*$), the relative share of government consumption to Gross Domestic Product (GDP) at home *vs.* abroad ($g_t - g_t^*$), and the real world price of oil (rpo_t). In the empirical implementation, the left hand side of equation to should equate to a disturbance term that is itself stationary with mean zero. The data are annual, from 1969-2000.[15]

Data considerations merit special attention. Cointegration studies suffer an inherent tension: while longer series are strictly better, insofar as they help reveal the common stochastic trends among variables, longer series also increase the likelihood of spanning a regime change or structural break. The presence of structural breaks, in turn, makes it more difficult to detect common stochastic trends among the variables.[16] In the present case, which is limited by the availability of adequate historical productivity measures for Mexico, we do not test

egy that finds a significant predictive role for these intermediate-term variables see Ávalos y Hernández (1995).

[15] The specific sources of the data are as follows: for relative price data we use the US CPI and the Mexican PPC (also a consumer price index). The bilateral exchange rate data comes from the Banco de Mexico. The relative government spending shares come from the Penn World Tables. US productivity data is the Commerce Department's output per man-hour in the manufacturing sector. A complete and consistent series for Mexican productivity is not available. We constructed this series in the following manner: for the 1960s and 1970s we used the data series "employment in the industry of transformation", interpolating missing years. For 1981-1988 we used the series "employment in the manufacturing sector". For 1989-2000 we updated the employment series with the available "growth rate in paid employment in the manufacturing sector" (Instituto Nacional de Estadística, Geografía e Informática, INEGI). Finally, we used INEGI data on sectoral output for each year, divided by the appropriate employment figure. The oil data come from the US Department of Energy, Energy Information Administration: http://www.eia.doe.gov/emeu/aer/txt/ptb0519.html.

[16] Zietz (2000) presents Monte Carlo evidence that structural breaks increase the incidence of type 2 (false negative) errors, and that the Engle-Granger technique is particularly susceptible.

explicitly for structural changes. In one sense, the question is moot, since the presence of such changes should only make it more difficult to find cointegration, which in fact we do find decisively. In another sense, the spirit of this analysis is to consider the effects of the inclusion of structural variables in a real exchange rate test. The behavior of an individual series, which may appear in isolation to have breaks, might alternatively be explained by shocks buffeting a variable with which it is cointegrated, as opposed to being driven by some completely unexplained *deus ex machina* that we call a structural break.

Turning to cointegration tests, we pre-test for the order of integration by applying a battery of tests to the variables in (2). The conclusions robustly indicate –both in terms of statistical significance and with respect to inclusion/exclusion of constants and trends– that the series are I(1).[17]

To test for cointegration between the set $\mathbf{x}$ of variables including the nominal exchange rate, relative price levels, relative productivities, relative government spending, and the world real price of oil, we construct the following error correction model:[18]

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[17] We began with the method suggested in Dickey and Pantula (1987), testing recursively $H_0 : (s)$ unit roots *vs.* $H_1 : (s - 1)$ unit roots, with a stopping condition when H_0 cannot be rejected. We began, arbitrarily, with 4 unit roots (tested *vs.* 3), and worked our way down. The number of lags included in the ADF estimation in each case was determined by the general-to-specific method (Hall, 1994), beginning with $k_{max} = 5$. Under these conditions, the tests robustly indicated I(1) processes, with the sole exception of relative prices, for which the hypothesis I(2) *vs.* I(1) could not be rejected at the 10% level of confidence. Resorting to a forward and reverse Dickey-Fuller test (Leybourne, 1995), we reject the null of I(2) *vs.* I(1) at the 10% level with a constant. Finally, we tested this variable with the method proposed in Ng and Perron (2001), which corrects for the size distortion in the presence of negative moving average errors. Again, we determined lag length according to Hall's method. The 4th lag was found to be significant at 10%. When the Ng and Perron *M*-tests were run on 4 lags, a unit root in first differences was rejected at the 5% level by all of the tests. When the same tests were run on 3 lags, a unit root in first differences was rejected at the 10% level. The cumulative evidence of these several tests leads us to conclude that relative prices, like the other variables in the model, are I(1).

[18] The number of lags –two, in this case– was determined by the likelihood ratio test, comparing various orders of VARs. The Akaike information criterion concurred with the LR test. The Schwarz and Hannan-Quinn information criteria, along with the "final prediction error" criterion, all indicated a lag length of 3. See table 2 for comparative results.

$$\Delta\mathbf{x_t} = \pi\mathbf{x_{t-1}} + \sum_{\mathrm{i}=1}^{2}\pi_{\mathrm{i}}\Delta\mathrm{x}_{\mathrm{t-i}} + \varepsilon_{\mathbf{t}} \tag{3}$$

where

$$\pi = \alpha\beta^{'} \tag{4}$$

Here the speed of adjustment parameter (α) multiplies the deviation from long term equilibrium (since $\beta^{'}\mathbf{x_t} = \mathbf{u_t}$, where β is the parameter vector and $\mathbf{u}_t$ is the stationary error term from the equilibrium relation posited in equation 2).

The rank (r) of π then corresponds to the number of cointegrating vectors. The λ -trace statistic tests $H_0 : r \leq n$ against $H_1 : r > n$. The λ-max (maximal eignvalue) statistic tests the $H_0 : r = n$ against the alternative $H_1 : r = n + 1$. Results are reported in table 1. Parameter values are reported in table 2 and table 3.

The tests strongly reject the absence of a cointegrating relationship in these data. The trace statistics reject $r = 0$ (in favor of $r > 0$) at the 1% confidence level, while the maximal eigenvalue statistics reject $r = 0$ (in favor of $r = 1$), also at 1%. In fact the trace statistics rejects $r = 1$ (in favor of $r > 1$) at 5% also. The maximal eigenvalue statistics cannot reject $r = 1$ in favor of $r = 2$.[19]

Table 1
Test Statistics for Cointegration

Hypothesized No. of CE (s)	*Eigenvalue*	*Trace Statistic*	*Max-Eigen Statistic*
None	0.811	104.58**	48.33**
At most 1	0.603	56.24**	26.82
At most 2	0.472	29.42	18.56

Note: Sample(adjusted): 1972 2000. Included observations: 29 after adjusting endpoints. *(**) denotes rejection of the hypothesis at the 5%(1%) level.

19 It is commonly the case that such ambiguous results emerge from the two tests. Given the difficulty of interpreting the meaning of a second cointegrating vector in this context, and the ambiguity of the results, in what follows we limit ourselves to consideration of a single cointegrating vector.

Table 2
Estimated Cointegration Vector $\beta = [\beta_1 \beta_2 \beta_3 \beta_4 \beta_5]$
(β_1 normalized to 1; standard errors in parentheses)

	s	*relp*	*rprod*	*relg*	*rpo*
Model 1 2nd Order VECM	1.00	-0.965 (0.00675)	2.388 (0.70088)	0.053 (0.01009)	-0.298 (0.11086)
Model 2 3rd Order VECM	1.00	-0.961 (0.00228)	2.582 (0.23230)	0.048 (0.00371)	-0.261 (0.03558)
Model 3 2nd Order VECM with weakly endogenous *s*, *rprod*, *relg*, and *rpo*.	1.00	-0.962	2.757	0.0592	-0.359

Table 3
Speed of Adjustment Vector (standard errors in parentheses)

	α_1 (Δ *s*)	α_2 (Δ *relp*)	α_3 (Δ *rprod*)	α_4 (Δ *relg*)	α_5 (Δ *rpo*)
Model 1	-0.398 (0.74640)	0.434 (0.30087)	0.033 (0.11460)	1.599 (5.29317)	-0.116 (0.71831)
Model 3	0	0.577 (0.0679)	0	0	0

Table 2 shows that the parameters of the cointegrating vector are highly significant and have the expected sign in all cases but one.[20] Equilibrium relative prices must move with the nominal exchange rate, *ceteris paribus*, in the long run. Higher domestic relative productivity is associated with an appreciation of the nominal exchange rate. Higher relative government spending likewise associates with a revalued nominal exchange rate. Increases in the real world price of oil are associated with devalued nominal and real exchange rates.[21]

Table 3 reports the speed of adjustment parameter estimates. The signs of these coefficients are the intuitive ones for the variables that are most plausibly adjusting, namely the nominal exchange rate and relative prices. The significance levels are low, however, with only the term driving the changes in relative prices coming anywhere near standard significance levels. This raises the possibility that some of the variables are *weakly exogenous* (Johansen, 1992) with respect to the parameters of the cointegrating vector, and that the dynamic adjustment might be correspondingly less complex. Table 4 reports likelihood ratio tests for weak exogeneity of each of the variables (i.e.

Table 4

Tests of Weak Exogeneity

	α_1=0 (Δ *s*)	α_2=0 (Δ *relp*)	α_3=0 (Δ *rprod*)	α_4=0 (Δ *relg*)	α_5=0 (Δ *rpo*)
LR test	0.318	2.158	0.125	0.148	0.036
	(0.572)	(0.141764)	(0.723)	(0.700)	(0.848)

Note: LR distributes as χ^2 on (1) d.f., *p*-value in parentheses.

[20] Because of the ambiguity in the optimal lag structure (see previous footnote) a 3rd order VECM was also estimated. The cointegrating vector was again found to be highly significant, and the parameter estimates were found to have the same signs (and similar magnitudes) as in the 2nd order model. The cointegrating vectors from both models are reported in table 2. All other results reported here are from the 2nd order VECM.

[21] While the expected effect of changes in real oil prices is not obvious, the result here seems counterintuitive: the bilateral peso-dollar exchange rate rises (read: "peso devaluation") when oil prices rise, despite the fact that Mexico is a net exporter and the US a net importer of oil. One plausible explanation is that oil prices have a depressive effect on the US economy, causing a fall in demand for Mexico's non-oil exports.

we test the parameter restriction that $\alpha_i = 0$ for each variable i). Weak exogeneity cannot be rejected.

Whereas the adjustment parameter on relative prices is close to standard significance levels while the other adjustment parameters are not, and whereas it is likely that relative prices will do the vast majority of the adjusting (particularly as the sample spans long periods of fixed nominal exchange rates and/or dirty floating, with far fewer spells of truly free-floating rates), we re-estimate the model imposing the restriction that the other four speed of adjustment parameters are all equal to zero. Table 2 shows that this restriction changes the parameters of the estimated cointegrating vector very little, while table 3 shows that the significance level of the speed-of-adjustment parameter on relative prices has risen very significantly.

We use the Johansen procedure to test various restrictions on the cointegrating vector obtained above (and reported as model 1 in tables 2 and 3). First, we test the hypothesis that relative PPP alone better fits the data. This amounts to the following parameter restrictions:

$$[\beta_3 = \beta_4 = \beta_5 = 0] \tag{5}$$

Note that this can be equivalently viewed as a test of whether the variables we include to augment PPP theory jointly add information explaining systematic deviations from PPP. Results are reported in table 5.

Table 5

Tests of Cointegration Parameter Restrictions

<table>
<tr><td colspan="4">Restriction: $\beta_3 = 0, \beta_4 = 0, \beta_5 = 0$
Description: cointegration parameters on *relg, rprod*, and *rpo* are zero</td></tr>
<tr><td>Restricted Log-likehood</td><td>LR Statistic</td><td>Degrees of Freedom</td><td>Probability</td></tr>
<tr><td>95.94650</td><td>16.95887</td><td>3</td><td>0.000721</td></tr>
<tr><td colspan="4">Restriction: $\beta_1 = -\beta_2$
Description: proportionality of *s, relp*</td></tr>
<tr><td>Restricted Log-likehood</td><td>LR Statistic</td><td>Degrees of Freedom</td><td>Probability</td></tr>
<tr><td>97.50199</td><td>13.84789</td><td>1</td><td>0.000198</td></tr>
</table>

The LR statistic distributes as chi-squared (on 3 d.f., for the present case of 1 cointegrating vector). We calculate the value at 16.95, for a significance level of .0007 (.07%). In other words, we have strong evidence that relative productivities, government spending, and the real price of oil are appropriate variables to include in order to solve the puzzle of systematic deviations from real exchange rate stationarity.

It would perhaps be preferable to perform tests directly on the real exchange rate, rather than on the more cumbersome separation of the nominal exchange rate from relative prices which we have reported thus far. Cheung and Lai (1993) point out that measurement issues relating to the calculation of price indices may imply that proportionality between the nominal exchange rate and prices does not obtain, even if PPP does hold. Proportionality ($\beta_2 = -\beta_1$) should thus be tested as a restriction on the cointegrating vector. If it holds, cointegration tests may be carried out directly on the real exchange rate, reducing the number of parameters to estimate. Unfortunately table 5 also indicates that we can strongly reject proportionality, and thus may *not* proceed with direct estimation of a cointegrating vector using the real exchange rate.

Since we cannot impose proportionality or exclude any of the variables from the original specification, we proceed to estimate the vector error correction model, VECM, of equation (3), imposing 1 cointegrating vector, using all of the variables originally discussed, and including 2 lags of changes. Table 6 reports the results of a Choleski decomposition, based on this VECM, for a four year forecasting horizon.[22]

The fundamental insight from this exercise is the importance of relative government spending, which explains over a quarter of the forecast error variance of the nominal exchange rate and relative price levels.[23] This is consistent with Dibooglu (1996), who also finds that this variable overwhelms the other determinants.[24]

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[22] The decomposition was accomplished using *E-views* econometric software, version 5.0.

[23] When the order of variables is reversed in this procedure, we find only one significant inconsistency: the proportion of the variance explained by *relp* and by *s* are virtually reversed, i.e. *s* becomes insignificantly small, while *relp* assumes all of the explanatory power. The forecast error variance attributed to *rpo, relg*, and *rprod* do not change importantly under a reverse ordering.

[24] Since Dibooglu (1996) was not able to reject proportionality, the variance decomposition reported there is with respect to the real exchange rate.

Table 6
Choleski Variance Decomposition

	Innovation in				
	s	*relp*	*rprod*	*relg*	*rpo*
s	71.74	0.08	1.19	26.38	0.59
relp	69.02	0.77	0.92	27.75	1.52
rprod	10.08	4.67	74.69	9.51	1.02
relg	5.16	9.94	21.72	59.65	3.51
rpo	12.14	4.95	45.40	17.12	20.35

4. The Mexico Crisis, Revisited

An important discussion began in and about Mexico as early as 1993, regarding the sustainability of the fixed exchange rate. Dornbusch and Werner (1994) estimated 20% overvaluation, and issued a controversial call for a "one-time" devaluation. In the ensuing political turmoil of 1994, an opportunity to devalue without losing face was passed up. An insistent Finance Ministry led by Pedro Aspe argued publicly that real exchange rate appreciation was consistent with underlying fundamentals (such as improvements in productivity resulting from Salinas's microeconomic reforms), and worried privately that a devaluation would destroy hard-won credibility (Aspe, 1993, chapter 1).

Figures 6, 7, and 8 report respectively the actual path of the (log) real exchange rate, the long-run equilibrium value of the real exchange rate (at each moment in time) that we derive from the cointegrating relationship reported above, and the difference between the two, which can be interpreted as the percent over-valuation of the real exchange rate.

Figure 6 vividly captures the turmoil of the 1980s, as Mexico found itself with a classic transfer problem when international capital flows abruptly stopped. According to Figure 8 the extent of undervaluation reached over 60 percent at the height of the debt crisis. Figure 6 also illustrates the rapid appreciation through the late 80s and early 90s, as the *Pacto de Solidaridad* stabilized the wage-price spiral and the Brady Plan began to ease some of the external pressures. One clear finding that emerges is that the real devaluation of the early-to-mid 1980s finds no explanation in movements of "fundamentals". Figure 7 suggests that fundamentals called for a 9% revaluation between 1980 and 1986, and a 38% real devaluation between 1986 and 1990. In fact (as figure 6 shows) the actual real exchange rate *de*valued some 65% between 1980 and 1986, and *re*valued some 32% between 1986 and 1990.

The subsequent period seems to indicate a larger role for fundamentals, at least until the 1995 collapse. The real exchange rate appreciation associated with this period of stabilization was widely noted, occasionally with dread. The analysis here, based on fundamental determinants of real exchange rates, finds a magnitude of overvaluation in 1993 of 23%, which is almost exactly the midpoint between the estimates in Warner (1996: 25%) and Dornbusch and Werner (1994: 20%). Furthermore, we find that this overvaluation falls by about half in 1994.

What of the movements of other determinants of real exchange rates? From 1990 to 1995, the relative share of government spending in Mexico recovered (after falling precipitously in the mid-1980s; see figure 1), which, *ceteris paribus*, could account for a 4% appreciation of the equilibrium nominal exchange rate between 1990 and 1994. Meanwhile the relative productivity (figure 2) of Mexican firms does improve from 1988 to 1990, explaining some 9% of the increase in the equilibrium real exchange rate during this period. This variable is essentially stagnant between 1990 and 1994, however, and thus is not a likely candidate for shedding light on the revaluation question of that time period. The real price of petroleum, meanwhile, falls some 36% during the 1990-1994 period (figure 3), which according to the model appreciates the equilibrium real exchange rate some 14%.

In summary, the real exchange rate appreciated 26% between 1990 and 1994 (pre-crash), while the rate we estimate to be the equilibrium rate appreciated 21%, with most of the explanation due to changes in government spending relative to GDP, and to oil prices. The real misalignment by the end of 1994 was just 13%, based on fundamentals. Clearly the exchange rate collapse was out of all proportion to the actual misalignment.

Figure 1
Share of Federal Government Consumption to GDP in Mexico Relative to US (in logs)

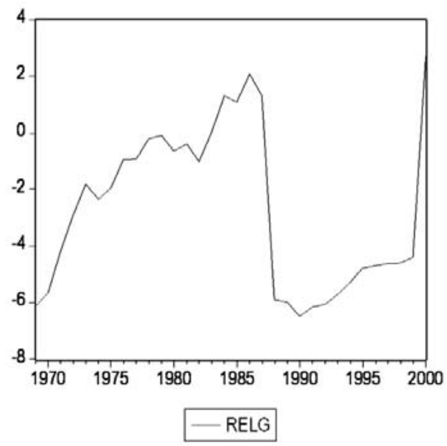


Figure 2
Relative Mexican Productivity (in logs)

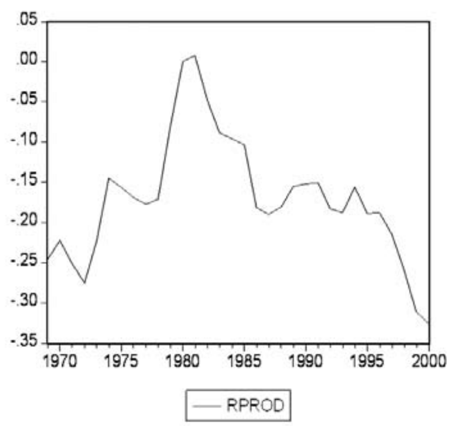


Figure 3
Real $ Price of Petroleum (in logs)

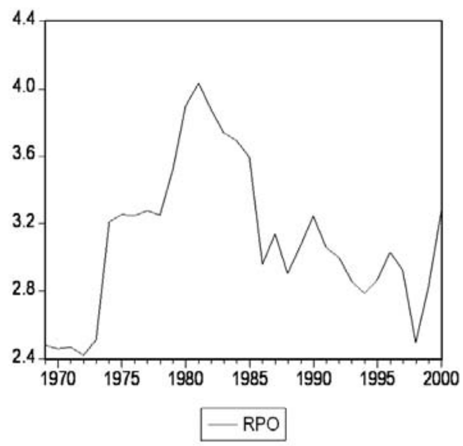


Figure 4
Ratio of Mexican to US CPI (in logs)

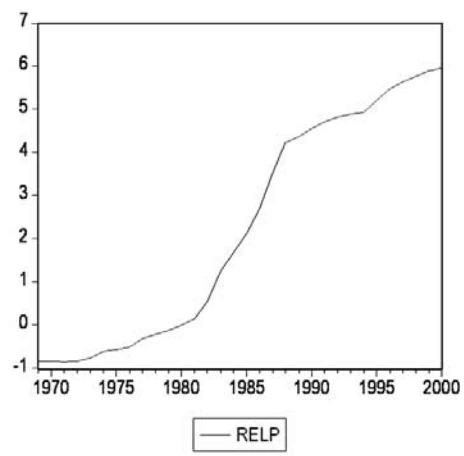


Figure 5
Nominal Exchange Rate, Pesos/ \$ (in logs)

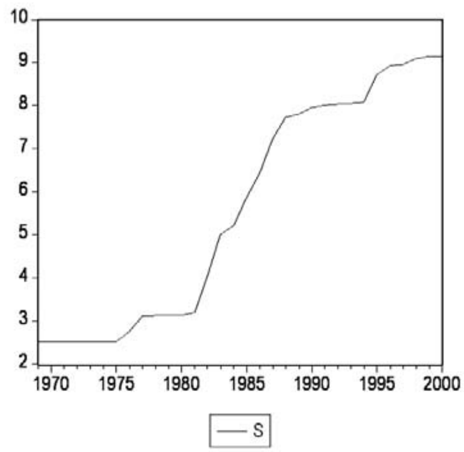


Figure 6
Historical Real Exchange Rate (log)

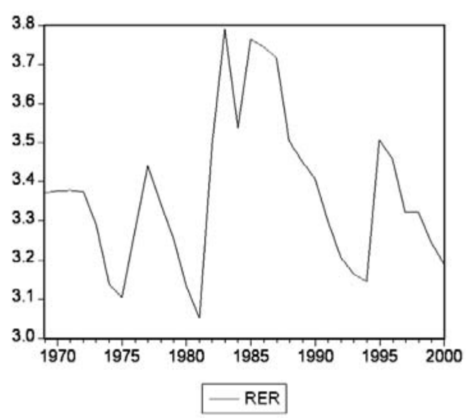


Figure 7
Estimated Equilibrium Long-Run Real Exchange Rate

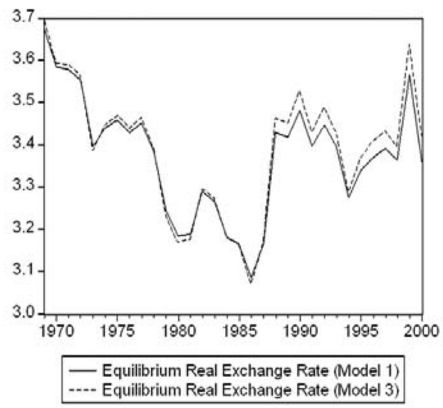
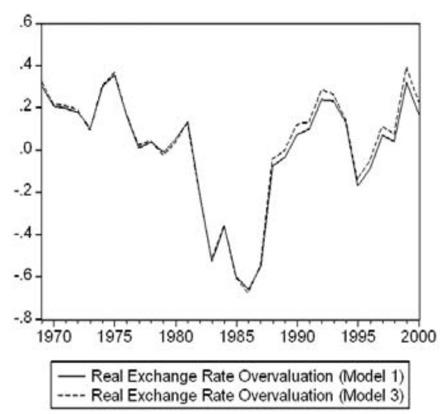


Figure 8
Estimated Real Exchange Rate Overvaluation



5. Conclusion

We find that "fundamental" determinants of real exchange rates as identified by the theoretical literature are in fact cointegrated with nominal exchange rates and relative prices, for the bilateral case of Mexico and the US, while tests of PPP alone fail to reject a unit root in the real exchange rate. The clear implication is that any attempt to gauge the appropriate level of the nominal exchange rate must consider more than just the relative price levels.

When we use the cointegrating equation to derive the path of the "equilibrium" real exchange rate we find a rather severe undervaluation during the debt crisis of the 1980s and a modest overvaluation in the period immediately preceding the catastrophic devaluation in December, 1994. The analysis suggests that nothing in the fundamentals of long-run real exchange rate determination can account for the blow Mexico suffered at that time.

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