Random Walks and Half-Lives in Chilean and Mexican Peso Real Exchange Rates: 1980 – 2003.

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Abstract: Several papers have shown that high-inflation contributes to mean reversion in real exchange rates. This paper studies the Chilean peso (CLP) and Mexican peso (MXN) real exchange rates over 1980-2003. Three datasets are used: two with quarterly and monthly bilateral data (against the U.S. dollar) with consumer and producer price indices; and another with monthly real effective rate exchange rates (REER). Unit root tests do not reject the root in levels for both currencies. Half-lives, however, contrast markedly: at 5 years or infinity for the Chilean peso and between 1 and 3 years for the Mexican peso. The CLP thus clearly behaves like a random walk, while the MXN implies mean reversion patterns faster than the 3 to 5 years "consensus" proposed by Rogoff (1996) for industrial countries. We present evidence that the sharp MXN depreciations and Mexico's relatively higher inflation record may have lead to these contrasting results.

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I. Introduction

At high rates of inflation, nominal shocks dominate real shocks. Evidence on the German hyperinflationary experience of the 1920s in Frenkel (1978) has led researchers to search for this stylized fact in postwar high inflation countries. Typical examples are McNown and Wallace (1989) for 4 countries over monthly 1976-1986 data, Mahdavi and Zhou (1994) for 13 countries over quarterly data at varying periods, and Choudhry (1999) for 4 Eastern European countries over monthly data at varying periods from 1991 to 1997. While the estimated coefficients of cointegrating regressions between exchange rates and relative prices are much closer to the PPP-based value of unity in high inflation countries, Salehizadeh and Taylor (1999) find large deviations from the theoretical values in a large sample of 27 developing economies. Findings such as these have led some to argue that "empirical evidence from countries that have experienced varying inflation rates is, however, more ambiguous" [Bleaney et al. (1999), p.839].

A very influential study by Rogoff (1996) refers to the "remarkable consensus" of 3-5 year half-lives of deviations from PPP in long-horizon data for currencies of industrial countries. Bayesian evidence in Kilian and Zha (2002) has found, however, considerable uncertainty about the half-life, with an average of 90% posterior error bands covering values between 2 and 15 years. The half-life, defined as the number of periods required for a unit shock to dissipate by one half, is a good *measure of persistence*. A conceptual issue is how to reconcile the high short-term volatility of real exchange rates with slow convergence to (long-run) PPP. Nonlinear mean reversion proposed by Michael et al. (1997) and the structural change hypothesis suggested by Cheung and Lai (1998) can be offered as explanations for this fact.

Further evidence with new econometric tools to handle persistence in real exchange rates may lead to new findings. This is because the null hypothesis of the second stage of PPP research, as surveyed by Froot and Rogoff (1995), is simply the random walk hypothesis. The no-rejection of the null implies non-stationarity, while the alternative hypothesis comprehends the stationary autoregressive process. Any half-life figure must be based on the estimate of the autoregressive parameter, which may be imprecise. Murray and Papell (2002) and Cashin and McDermott (2003) apply to real exchange rates the adjustment to the autoregressive parameter proposed by Andrews (1993) in the AR (1) case and by Andrews and Chen (1994) to the more general AR (p) case. These methods are called *median unbiased* (MU) estimations of AR parameters in DF or ADF-type regressions and essentially deal with the problem of downward bias in the AR parameter. More specifically, "if the least squares estimate equals .8, say, one does not use .8 as the estimate of α, but rather, one uses the value of α that yields the least squares estimator to have a median of .8" (Andrews, 1993, p. 140).

While abstracting from nonlinear mean reversion, this paper employs the MU methods for adjustment of the autoregressive process for quarterly data from 1980:1 to 2003:4 and monthly data from 1980:1 to 2003:12 of two important Latin American currencies against the U.S. dollar (USD): the Chilean peso (CLP) and Mexican peso (MXN). The only other study we know of that addresses half-lives for Latin American countries finds that "estimates for developing countries appear much more dispersed than those for industrial countries. Most of the half-lives for developing countries are less than 3 years" (Cheung and Lai, 2000, p. 388). Such result was presented for blocks of countries and not for individual countries.

We are specifically interested in comparing two real Latin American currencies that behave very differently and, for this reason, may have different degrees of persistence. While performing unit roots testing is common practice in many studies on the real exchange rate in Latin America, no study has focused on the connection between these tests and the persistence as measured by the half-life. Typical recent examples that have not discussed half-lives include: the battery of unit root tests for Argentinean macroeconomic variables (including the real exchange rate) conducted by Carrera et al. (2003) and the analysis for real exchange rates in the very long run for Brazil by Mollick (1999).

Similarly, the measurement of price indices requires explanation. Although the consumer price index (CPI) is the most widely used index in PPP tests, its calculation involves large amounts of non-tradables and different baskets of goods across countries. Engle (1999), for example, finds that relative prices of non-traded goods account for almost none of the movements in U.S. real exchange rates. Kim (1990) does not reject the random walk hypothesis for CPI-based U.S. real exchange rates; yet he obtains cointegration for series calculated with the wholesale price index (WPI). Others find evidence (in 14 out of 27 countries) in favor of PPP as a cointegration concept under CPI, although the coefficients do not correspond to the unitary values implied by PPP, which led them to conclude about a "no discernible pattern" [Salehizadeh and Taylor (1999), p. 192].

Given the conflicting evidence on the appropriate price index, this paper employs three different datasets. Two are with bilateral data (against the U.S. dollar) under both CPI and PPI indices. The other collects data from both CLP and MXN with respect to countries that represent large part of the trade against Chile and Mexico, constructed with PPI series. This is a sort of the real effective rate exchange rate (REER), advocated by the IMF as a better measure of real competitiveness of a currency than bilateral rates. Analyzing effective rather than bilateral real exchange rates captures the international competitiveness of a country against its major trade partners. It also helps avoid biases related to the choice of base country in bilateral real exchange rate analyses. This has been refereed to as "the U.S. dollar problem" that might carry substantial bias in Latin American countries. Nevertheless, bilateral rates against the USD are more often used in practice, and one is tempted to wish to know if the results change across different real exchange rate definitions.

We estimate ARMA models and find substantial moving average (MA) terms for the CLP real exchange rates process, while an augmented AR model fits the MXN nicely. Five unit root tests, including those proposed by Ng and Perron (2001) that are able to deal with near unit roots in the AR case or large MA terms, uniformly coincide in suggesting integrated of order one [hereafter I(1)] processes for both real exchange rates. These results are obtained for both quarterly and monthly datasets.

The evidence on half-lives, however, contrasts remarkably: at 5 years or infinity for the Chilean peso and between 1 and 3 years for the Mexican peso. The CLP thus clearly behaves like a random walk, while the MXN implies a mean reversion pattern faster than the 3 to 5 years “consensus” by Rogoff (1996) on industrial countries. In addition to MU estimation, the half-life calculation is herein calculated in two ways, including the correction factor proposed by Rossi (2005), who found that the lower bound of the confidence interval is around 4 to 8 quarters for most currencies, while the upper bounds are infinity for all currencies.

We conjecture that the several abrupt depreciations of MXN and Mexico’s relatively higher inflation record may have lead to these contrasting results. We also discuss in this paper features of CLP that make it smoother in the sample, possibly due to “real exchange rate” targeting by the Chilean Central Bank.

This paper contains four more sections. Section II presents the data employed and section III reviews the empirical models to be implemented. Section IV summarizes our main findings and section V concludes the paper and indicates extensions for further work.

II. The Data and Institutional Setting

The bilateral dataset is taken from the IMF's *International Financial Statistics* (http://ifs.apdi.net/imf/logon.aspx) and covers the period 1980:1 to 2003:4 for quarterly data and 1980:1 to 2003:12 for monthly data, except for the Mexican peso real exchange rate, which covers the period 1981:1 to 2003:12 for the monthly data constructed with the PPI. All other real exchange rates (CLP under CPI and PPI and MXN under CPI) run from 1980:1 to 2003:12. We employ logs on all series as in the standard equation ($q_t \equiv s_t - p_t + p^*_t$) and obtain the log real exchange rate series in figure 1 for CLP constructed with PPI and with CPI, respectively. Figure 2 contains the same figure for MXN. For both price indices, the overall graph representations look similar. Yet there seems to be some discrepancy in the estimates, particularly for the real CLP, in which there are noticeable gaps in the early and late 1980s as well as in more recent years. [1]

[Figures 1 and 2 here]

The multilateral dataset comes from Argentina's Center for International Economics (CEI: Centro de Economía Internacional; http://www.cei.gov.ar/home.htm) of the Secretary of Foreign Affairs and covers monthly data from 1980:1 to 2003:12. The multilateral (effective) real exchange rate (REER) is adjusted by the producer price indices. The destination weights vary over the years and are calculated every year, representing about 71% of total Chilean and Mexican trade. The countries included for the CLP are: Argentina, Brazil, Canada, United States, Mexico, Peru, Germany, Spain, France, Italy, Belgium, Netherlands, U.K., Sweden, and Japan.

[1] Using data from 1980 to 2003, the 1970s period of high inflation in Chile is absent in our empirical evidence. The study by Fischer et al. (2002) on inflationary episodes in market economies displays a remarkable difference under high-inflation. In Chile, from October 1971 to May 1977 (68 months), the cumulative inflation rate was 127,958%. In Mexico, from December 1985 to August 1988 (33 months), the cumulative inflation rate was 724% and from February 1982 to July 1983 (18 months), the cumulative inflation rate was 180%.

The countries for the MXN REER calculation are: Argentina, Brazil, Canada, Chile, United States, Peru, Germany, Spain, France, Italy, Belgium, Netherlands, U.K., Switzerland, and Japan.

Since the REERs resemble the behavior of the pair of series just displayed under bilateral real rates, we omit both multilateral graphs. Analyzing effective rather than bilateral real exchange rates measures the international competitiveness of a country against its major trade partners and addresses robustness, thereby avoiding biases due to the "U.S. dollar problem".[2]

The general pattern of the graphs is that the CLP appears smoother, while MXN displays pronounced peaks with the 1982, 1987, and 1994 depreciations. The latter one, in particular, represented a drastic move into the floating rate scheme after so many variations of pegged systems in the 1980s and early 1990s. The Chilean experience is perhaps less known and ended up less abruptly. Gallego and Loayza (2002) refer to the "golden period for growth in Chile" as remarkable and their growth accounting exercise showed that the large increase in the growth rate after 1985 was due mostly to an expansion of total factor productivity. After 13 years of sustained high growth rates, Chile experienced a slowdown in 1998. Figure 1 shows that the real exchange rate appreciated until 1997, which is consistent with a Balassa-Samuelson effect of productivity in the tradable sector implying a stronger real currency. Morandé and Tapia (2002) argue that growth in demand far outpaced output, forcing stricter monetary policy and higher interest rates. The Central Bank of Chile, reluctant to abandon the exchange rate band, pursued several "second-best" options between 1990 and 1997 aiming at accommodating a stronger

[2] See Ellis (2001) for several issues on the calculation of REERs that deal with which currencies to include, the appropriate weighting scheme and the price measures to use. She emphasizes that the choice of the index depends on the issue being investigated and that there is no single "right" measure. Her message is that since real exchange rates are intended to capture competitiveness, "it would be preferable to deflate the nominal exchange rates with some measure of *producer* prices or costs rather than consumer prices" (Ellis, 2001, p. 12).

peso.[3] After the Russian crisis of 1998 and the devaluation of the Brazilian real in January of 1999, the floating rate regime was adopted by Chile in September of 1999.

Figures 3 and 4 graph Chilean and Mexican quarterly inflation rates measured by the CPI (the ones for PPI are very similar) against the logarithm of their respective real exchange rates. In Chile the real exchange rate follows a weaker pattern with inflation changes than in Mexico. In both countries the early 1980s is subject to relatively higher rates of inflation, while recent data suggests price stabilization. Inspection of figures 3 and 4 leads to the conclusion that Mexico had higher inflation rates than Chile over the period. Also, the abnormally high quarterly inflation in the first quarter of 1995 is matched by accompanying movements in the real exchange rate when the Mexican peso started to float. From then onwards, the inflation rate and the RER series are subject to a downward trend. The other spike in Mexican inflation around the late 1980s occurred under the managed exchange rate system, in which case the RER did not respond as in the mid-1990s. In all, these two figures suggest a closer relationship between price changes and the RER in Mexico than in Chile.

[Figures 3 and 4 here]

The remaining sections formalize this stylized fact under several unit root tests and estimates of the autoregressive coefficient of the real exchange rate.

[3] These included: 1. an increase in the band's width from 10% in 1990 to 25% in 1997; 2. discounting a positive productivity factor in addition to foreign inflation when adjusting the band's center; 3. changing the foreign inflation definition; and 4. moving from a dollar reference to a basket of currencies (with the dollar, the yen, and the mark). Complementary policies included: regulations to capital inflows (reduced to 0 in September of 1998 and eliminated in April 2001) and sterilization of foreign exchange reserves. These were attempts to reduce the peso appreciation. In general, Chile has been mentioned in the literature as an example of targeting the real exchange rate. Calvo et al. (1995, p. 100), for example, refer to the July 1985 – January 1992 period in Chile as "no deviations from the rule", while Grier and Hernández Trillo (2004, p. 5) acknowledge this fact in a study on MXN and output: "short run real exchange rate targeting, as practiced by Chile, involves repeated adjustments to the nominal exchange rate."

III. Testing Strategy and Discussion

Provided individual series are I(1), empirical tests of long-run PPP are based on:

$$s_t = \alpha + \beta_1 p_t + \beta_2 p^*_t + \varepsilon_t \qquad (1),$$

where: s is the logarithm of the nominal exchange rate (domestic price of foreign currency), p is the logarithm of domestic prices, p* is the logarithm of foreign prices, and ε_t is the error term. Equation (1) has been subject to extensive cointegration tests, starting perhaps with Enders (1988). Froot and Rogoff (1995) refer to this as stage 3 of PPP tests. In this case, if the three individual series are I (1) and if there is a cointegrating vector representing a linear combination of them, long-run PPP is found. However, despite implying mean reversion, cointegration assumes a linear process for ε_t, meaning the adjustment process is both continuous and has a constant speed.[4]

Imposing the restrictions $\alpha = 0$, $\beta_1 = 1$ and $\beta_2 = -1$ on (1), the error term becomes a measure of the real exchange rate (q_t), as discussed in Xu (2003), for example. With these restrictions, it makes sense to measure deviations from parity as follows:

$$q_t \equiv s_t - p_t + p^*_t \qquad (2),$$

All series of real exchange rates (q) are first tested for a unit root using the ADF test. This procedure is referred to as "stage 2 of PPP tests" by Froot and Rogoff (1995). If one supposes

[4] Theoretical contributions, initiated by Dumas (1992), for example, have emphasized the importance of transaction costs, suggesting a nonlinear adjustment is more plausible. Michael et al. (1997) and Taylor et al. (2001) contain evidence for major currencies, while Ferreira and León-Ledesma (2005) discuss asymmetries in the real interest rate hypothesis (differences in real interest rates) for a set of emerging and developed countries using monthly data from 1995 to 2002.

some form of long-run PPP to hold, the real exchange rate should be stationary and the unit root null should be rejected. The ADF test procedure estimates:

$$\Delta q_t = \alpha_0 + \alpha_1 t + \beta_0 q_{t-1} + \sum_{j=1}^{k} \beta_j \Delta q_{t-j} + \nu_t \qquad (3),$$

where: α_0 is a constant; t is the time trend whenever the time trend is included in the estimation in levels[5]; q_t is the real exchange rate; Δq_t is the first-difference of q_t; α_1 and the β's are parameters to estimate; and ν_t is the stochastic disturbance with white-noise properties. The null hypothesis of a unit root is represented by $\beta_0 = 0$ and the ADF statistic is the value associated with the t-ratio on the β_0 coefficient. In practice the optimal lag-length (k) in this paper is determined by the sequential procedure suggested by Ng and Perron (1995). The choice of k in this fashion is expected to yield the desired white-noise properties on ν_t.

Elliott et al. (1996) perform asymptotic power calculations and show that the modified DF-GLS test can achieve substantial power gain over the testing procedure outlined in (3). The DF-GLS$^{\tau}$ test that allows for a time trend is based on:

$$\Delta q_t^{\tau} = \phi_0 q_{t-1}^{\tau} + \sum_{j=1}^{k} \phi_j \Delta q_{t-j}^{\tau} + \nu_t \qquad (4),$$

[5] The literature related to Balassa-Samuelson (BS) effects attributes a natural interpretation to the trend term in the ADF-type regression in (3). Allowing for trend is equivalent to accepting factors that have a systematic influence on the real exchange rates. As Sabaté et al. (2003) put it, this may be due to movements in the relative prices of traded over non-traded goods across borders due to BS effects and a demand side bias in favor of non-traded goods. But this may be also be due to the non-stationarity of real exchange rate for traded goods because of menu costs of pricing to market strategies.

where q_t^{τ} is the GLS-detrended series: the original q_t without either a constant and a constant and trend. One considers the same null hypothesis on the t-ratio tests on ϕ_0 but critical values change as reported in Elliott el al. (1996). The lag-length selection procedure below is based on the minimization of the Schwarz-Bayes information criterion as in the original paper.

We also employ the KPSS test as developed in Kwiatkowski (1992) with bandwidth set at 4 and Bartlett kernel. Finally, the more recently developed tests by Ng and Perron (2001) may be particularly useful when the autoregressive root is close to one or when there are negative moving average terms. For these tests, we perform estimation with both the modified Akaike information criterion (MAIC) and the Schwarz Bayesian information criterion (BIC) as advocated in Ng and Perron (2001). The tables below contain the results based on BIC, which tend to select lower lags than MAIC. As occurred in the G-7 inflation series reported in Ng and Perron (2001), evidence of stationarity herein is weaker with MAIC than with BIC.

Complementary to the unit root tests is the study on the persistence of the dynamics of real exchange rates. The reason is that the unit root null hypothesis of the test procedures above is tested against the alternative of stationary autoregressive (AR) model. In order to estimate the speed of convergence to PPP, the first-order autoregressive model on q_t is adopted under the assumption of independent identically distributed (i.i.d.) normal errors:

$$q_t = \alpha_0 + \alpha_1 q_{t-1} + \nu_t \qquad (5),$$

where the autoregressive parameter α_1 lies in the interval (-1, 1]. The half-life (HL) measures the time it takes for a deviation from PPP to dissipate by 50% and is calculated by HL = ABS[ln (0.5)/ln (α_1)]. Survey papers on long-horizon data, such as: Froot and Rogoff (1995) and Rogoff

(1996), report as the consensus in the literature that the half-life of a shock to the real exchange rate lies between 3 and 5 years. This slow speed of reversion to PPP is difficult to reconcile with the observed large short-run volatility of real exchange rates.

The problem with (5), however, is the presence of serial correlation. As discussed in Andrews and Chen (1994), the AR (p) model may be used, incorporating lagged first-differences to account for serial correlation. The AR (p) model, for t = 1, …, T, is the special case of (3):

$$q_t = \alpha_0 + \beta_0 q_{t-1} + \sum_{j=1}^{k} \beta_j \Delta q_{t-j} + \nu_t \qquad (6),$$

where we again use the general-to-specific lag selection procedure suggested by Ng and Perron (1995), with maximum lag set at k = 12 and 5% as the significance criterion for the last k term.

On the calculation of the half-lives itself, following Rossi (2005), we calculate the half-life taking into account the b (1) correction factor, which is shown to be equal to b (1) = 1 - $\Sigma\beta_i \Delta q_{t-i}$ (i = 1 to k) in the ADF-type regression. The b (1) correction factor enters the calculation of the half-life as: $h^* \equiv \max\{\ln(0.5b(1))/\ln(\alpha_1), 0\}$, which differs from the standard half-life obtained by $h_a \equiv \max\{\ln(0.5)/\ln(\alpha_1), 0\}$. Both half-lives ($h_a$ and h^*) will be reported in the next section and the 95% confidence intervals for h_a and h^* (respectively, h_{al}, h_{ah}, h^*_l, and h^*_h) will be calculated using a delta method approximation as suggested by Rossi (2005): $h_a \pm 1.96\sigma_{\alpha 1}$ $\{(\ln(0.5)/(\alpha_1))\ [\ln(\alpha_1)]^{-2}\}$, where $\sigma_{\alpha 1}$ is the estimate of the standard deviation of α_1. Since the half-life can not be negative, we impose a lower bound of zero.[6]

[6] Using quarterly data from non-EMU countries from 1973:3 to 2002:2 and from EMU countries from 1973:3 to 1998:2 and CPIs, Rossi (2005) reports point estimates of h* to be around 8 to 12 quarters for most currencies, while earlier research had focused on h_a estimation. Due to the absence of the correction factor, the latter generally underestimates the true half-life.

IV. Results

A. ARMA Models

Tables 1 and 2 contain ARMA (1,1) and ARMA (2,2) as a first pass on the time series processes for quarterly and monthly datasets, respectively. In table 1, MA terms are statistically significant for the Chilean peso, regardless of the domestic price index used.[7] In order to detect serial correlation, we employ the Ljung-Box Q (12) statistics for quarterly data and Q (36) for monthly data, as well as the Lagrange Multiplier Breusch-Godfrey statistic, which coincide in general. Any evidence of serial correlation implies a systematic movement in q_t that is not accounted for by the ARMA coefficients. For the Mexican peso, when the CPI is used as domestic price index, serial correlation is detected in both models, although the MA terms are weaker, implying rejection of the null of zero β-coefficients at 10% only. When the PPI is employed, there is no evidence of misspecification for the real value of the bilateral Mexican peso.

[Tables 1 and 2 here]

These results are confirmed in table 2 with monthly data for the same time span. Here the Chilean peso has MA terms statistically significant in almost all cases. While the Chilean peso equations seem to be devoid of serial correlation, the bilateral real Mexican peso equations are plagued by serial correlation problems in 3 out of the 4 cases. Employing the REER at the bottom of the table, however, one finds white-noise processes for all models, except for the

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[7] This is not entirely unheard of for Latin American currencies. Mollick (1999) provides evidence of significant MA terms (a –0.22 coefficient) in the very long run (1855-1990) for the Brazilian currency against the British pound and the U.S. dollar. MA terms are typically associated with seasonal factors in the data. Enders (1995, pp. 106-111), for example, estimates AR (1), AR (2), ARMA (1,1) and ARMA (1,2) for the U.S. wholesale price index and refers to an ARMA (1,1) with an additional MA coefficient at lag 4 to account for the possibility of seasonality.

ARMA (1,1) in Mexico, although the rejection is only at 10%. The worst ARMA specifications seem to be for the Mexican peso under either the PPI or the CPI as the domestic price index.[8]

Simpler AR (1) specifications seem to fit the Mexican data better than the Chilean data as will be discussed in detail below.

B. Unit Roots

Important for our purposes in running simple ARMA (1,1) tests is to detect large moving average terms in the real exchange rate, which could affect seriously the size of standard unit root tests according to Ng and Perron (2001). Table 3 contains various unit root tests for the Chilean and Mexican peso real exchange rates with respect to the U.S. dollar. The frequency of data is quarterly in this case. The table contains also the more efficient DF-GLS tests, as well as the complementary KPSS tests under the null of stationarity, and the Ng and Perron (2001) tests.

According to the tests based on PPI for Chile, the results are very supportive of I(1) series. ADF and DF-GLS equally do not reject the unit root null in levels while DF-GLS does reject it in first-differences at the 1% level. The Ng and Perron MZ tests confirm the latter at standard confidence levels. Similarly, the KPSS rejects the null of stationarity in levels but does not do so in first-differences. These results are completely insensitive to the inclusion of the deterministic trend in the regression, as can be seen comparing the first (with trend) with the second row, with the trend omitted. For the results under PPI for Mexico, the DF-GLS surprisingly does not reject the null in first-differences, contrasting with the ADF that suggests

[8] As further diagnostic tests, we conduct ARCH LM tests where the null, distributed as χ^2 (q) - where q is the number of squared residuals - is that the coefficients on lagged squared residuals are all zero. Under quarterly data, we do not reject the null in all cases, except for the MXN measured by the CPI: at 10% for the ARMA (1,1) and at 5% for the ARMA (2,2). Under monthly data, we do reject the null in all cases for the bilateral CLP at 1% levels and for multilateral MXN REER at the 10% level, but do not reject for any other REER. We conclude that evidence on ARCH terms is not strong in our estimates.

I(1) pattern. The Ng-Perron MZ tests confirm the rejection of stationarity for the tests without trend in levels. But the KPSS is consistent with I(1) inference, although there is also disagreement in view of the inclusion of the trend term: 0.84 against 0.11. Evidence of stationarity in first-differences for the Mexican peso is roughly the same as that in Chile for the PPI: 4 out of 5 cases.

For the CPI-based tests for Chile, the inference favors I(1) series, except for the ADF that is able to reject the null in levels, regardless of the inclusion of the trend term. For Mexican CPI, there is rejection of the unit root in levels (without trend) for the DF-GLS but the MZ tests are not able to reject the null in first-differences. The results from ADF and KPSS tests, however, are consistent with I(1) series. Overall, except for a few cases, the non-stationarity null is not rejected for the real exchange rate in levels. And, except for the tests with Mexican CPI, stationarity is achieved in first-differences. The newly developed tests by Ng and Perron (2001) confirm I(1) inference throughout except for the real exchange rate constructed with Mexican CPI. This class of unit root tests are particularly appropriate in cases when statistically significant MA terms are present and when there is near unit root behavior. As shown in table 1, there seems to be MA terms in the Chilean case: the ARMA (1,1) model for Chile, for example, detects fairly robust MA terms (0.268 for the PPI and 0.377 for the CPI) in table 1.

[Tables 3 and 4 here]

Employing monthly data, table 4 reports overwhelming evidence of I (1) processes for both Chilean and Mexican pesos. The only exceptions are the ADF tests for the Mexican peso, calculated with the PPI and the REER, which seem to suggest some stationarity in levels. However, all other three tests do not confirm this. The DF-GLS, KPSS and MZ tests uniformly do not reject the null in levels, which support the I(1) inference for both currencies under the three different exchange rate definitions. It is known that the DF-GLS and MZ tests, in particular, are superior to the ADF tests. Comparing to the quarterly data case, table 4 documents a more precise picture in favor of non-stationary processes in levels. This is consistent with a reduction in size distortions as the sample size increases from T = 96 to T = 288. One implication that follows from tables 3 and 4 for studies of cointegration analysis should be clear: one benefits from moving into a monthly dataset for these two real exchange rates as far as the evidence of I(1) processes becomes firmer.

C. Half-Lives

Table 5 contains simple autoregressive models for the quarterly dataset, which form the alternative hypothesis for the unit root testing procedure in the literature on real exchange rates as discussed in Froot and Rogoff (1995). For the AR (1) process for Chilean real exchange rate, the calculations imply very high 0.944 and 0.965 autoregressive coefficients, depending on whether PPI or CPI is used. As these estimates are potentially biased, the median unbiased (MU) estimator of Andrews (1993) is applied next. These methods are called *median unbiased* (MU) estimations of AR parameters in DF or ADF-type regressions and essentially deal with the problem of downward bias in the AR parameter.

It turns out that the MU estimator (α^{U}_{1}) clearly approaches 1: 0.990 for the PPI-based Chilean Peso real exchange rate and 1 for the CPI-based. The consequence for half-lives is an extremely slow convergence process for the Chilean peso, varying from 68.62 quarters (for the PPI) to ∞ (for the CPI). As the Ljung-Box and Breusch-Godfrey serial correlation tests detect serial correlation, however, we proceed with the more general AR (p) procedure for the Chilean peso real exchange rate. We conduct extensive search on additional autoregressive terms as recommended by Murray and Papell (2002) in order to account for serial correlation and employ the Ng and Perrron (1995) sequential test procedure to determine the optimal lag-length. We set the maximum number of lags in the quarterly case (in table 5) at k = 8 and in the monthly case (in table 6) at k = 24. In both datasets it is possible to obtain well-specified equations as there is no rejection of the null of no-serial correlation. Under the LM test, not reported in the table, there is only one rejection at the 10% level for the monthly Mexican peso under the PPI.

Below the AR (1) estimates in table 5, we report the AR (p) models and apply the algorithm in Andrews and Chen (1994) with one iteration when computing MU estimators for α_1. The resulting AR (p) models yield much lower coefficients for Chilean PPI (0.870) and CPI (0.936), compared to the AR (1) case. Note that there is no serial correlation according to Ljung-Box Q (.) tests (LM tests yield similar results) in these new AR (p) specifications. Applying the MU estimators due to Andrews and Chen (1994) yield a higher $\alpha^U{}_1$ for the Chilean peso compared to the biased α_1: 0.965 for PPI and 1.000 for CPI. We then get a fairly slow dissipation of shocks to the Chilean calculated real exchange rate at 19.47 quarters (slightly over 4.9 years) for the PPI. For the CPI, the corrected half-life tends to infinity, which is consistent with the random-walk hypothesis and no convergence whatsoever to PPP levels.

[Table 5 here]

We have seen that for Chilean PPI exchange rates the implied half-life is 12 quarters under the simpler AR (1) process. Application of conventional two-sided intervals (h_a) yields wide confidence intervals varying from 0 to 30.66 quarters. After taking into account serial correlation, the implied half-life becomes 4.98 quarters with h_a confidence intervals ranging from 1.52 to 8.44 quarters. When the correction factor b (1) is incorporated, the adjusted half-life increases to 6.77, with corresponding h* confidence intervals varying from 3.31 to 10.23

quarters. This is perhaps the only case that leads to discrepancy since the implied half-life based on the MU estimator is 19.47 quarters against 6.77 quarters in the corrected half-life calculation. In general, compared to the h_a calculation, increases in the magnitudes of the half-lives are observed for the other three real exchange rates in table 5. The increase is particularly high for the Chilean CPI case as we go from h_a to h*: from 10.48 quarters to 20.28.

The results for Mexican real exchange rates, however, present a contrasting pattern. The biased α_1 in table 5 are originally lower than the Chilean peso: for the AR (1) they are 0.905 for the PPI and 0.884 for the CPI. Computing the MU estimators for the AR (1) case yield 0.943 and 0.920 for the PPI and CPI, respectively. These are devoid of serial correlation according to the Ljung-Box Q (.) and LM tests and do imply 11.81 or 8.31 half-lives, respectively. These figures suggest half-lives between 2 and 3 years, in the lower band of the 3 to 5 years period discussed in Rogoff (1996) for industrial countries.[9] Applying the correction factor b (1) to the half-life calculation in the last column of table 1 yields 7.32 and 6.65 quarters, still below the "consensus".

Table 6 contains the estimates under monthly data, which reinforce the basic findings of the quarterly data. The Andrews (1993) and Andrews and Chen (1994) MU estimators include the same procedures as in the quarterly case. The only difference is that now the Ljung-Box test is evaluated at Q (36) to keep the same horizon with the quarterly case. Besides, the maximum lag for the AR (p) model selection procedure is also set at 24 in order to cover the same period as in the quarterly case (2 years).

[9] Although unnecessary given that AR (1) for MXN is not plagued by serial correlation, calculating more general AR (p) processes imply half-fives at 12.25 quarters and 5.70 quarters for the PPI and CPI Mexican peso bilateral real exchange rates, respectively. In years, these persistence measures imply almost 3 years for the PPI and 1 year and a half for the CPI.

In table 6 all Chilean peso rates, the two bilateral rates and the multilateral, imply infinite half-lives in the AR (1) and AR (p) cases alike. As in the quarterly data case, increases in the half-lives are observed for the real exchange rates in table 6 as the methodology by Rossi (2005) is adopted. For example, when the correction factor b (1) is incorporated, the adjusted half-life increases to 52.04 months from 36.13 for the PPI case, with corresponding confidence intervals running from 14.41 to 89.68 months. All other upper bands of the two confidence intervals for the Chilean peso are quite wide: 95.58 for the CPI and 132.81 for the multilateral REER.

[Table 6 here]

Since serial correlation is detected in the AR (1) models according to both Ljung-Box Q (.) and LM tests, the AR (p) models for the Mexican peso provide AR (11) for the PPI, AR (24) for the CPI and AR (10) for the REER. Similar to quarterly data, the results of half-lives are different across currencies in the following sense. Applying the MU estimators due to Andrews and Chen (1994) yield a higher α^U_1 compared to the biased α_1: 0.955 for PPI, 0.972 for CPI, and 0.974 for the REER. The implied half-lives are, respectively, for the AR (p) cases: 15.07, 24.40 and 26.35 months for the Mexican peso, which supports convergence to PPP levels.

In contrast to the Chilean peso, the half-lives and confidence intervals are much smaller for the Mexican peso. This finding is corroborated by the calculation of MU estimators and by confidence intervals for half-life using the correction factor in Rossi (2005). The half-life goes from 16.98 to 11.20 as we take into account serial correlation for Mexican PPI, and from 17.42 to 15.05 for Mexican CPI. When the correction factor b (1) is incorporated, the adjusted half-life increases to 24.81 from 11.20 for the Mexican PPI case, with corresponding confidence intervals running from 17.63 to 31.98 months. Similarly, the adjusted half-life increases to 20.01 from 15.05 for the Mexican PPI case, with corresponding confidence intervals going from 3.91 to

36.11 months. Therefore, the upper band of the Mexican CPI is not much beyond the 36 month-mark. For the multilateral case, the adjusted half-life increases to 35.89 from 15.77, with corresponding confidence intervals varying from 27.07 to 44.71 months. The *upper band* in this case is still between 3 and 4 years!

Taken together, evidence of mean reversion ($\alpha_1 < 1$) is clearly stronger for the Mexican peso RER. This holds across the three datasets investigated. The econometric techniques on the half-lives also provide consistent results with each other: whether the MU estimates by Andrews and Chen (1994) or the half-life calculations based on a correction factor by Rossi (2005).

How do our results relate to others in the literature? For the MXN real exchange rate, Taylor (2002) reports half-lives during his 1971 to 1996 floating rate period of just 1.1 year, much lower than the 3.6 years of the gold standard or the 6.2 years of the interwar period. The 1.1 year figure is considerably smaller than the mean and median half-lives around 2 to 3 years, a time frame considered "even more favorable to rapid PPP adjustment than most empirical studies." (Taylor, 2002, p. 145). There is also (pooled, with Mexico but without Chile) evidence of contrasting patterns between half-lives in industrial countries (from 2 to 5 years): "Most of the half-lives for developing countries are less than 3 years. Accordingly, the persistence in PPP deviations tends to be lower for developing countries than for industrial countries." Cheung and Lai (2000, p. 388). The only reference we know of half-lives for CLP finds smaller values than ours. Using a threshold autoregressive (TAR) model for 1810-2002 and splitting the sample into 1810-1973 and 1974-2000, "the half-life of the process falls from 2.1 to 1.6 years along both periods, which are even lower than the 2.8 years found for the full sample. In any case, the estimates are fairly lower than the traditional consensus, from 3 to 5 years, stated by Rogoff (1996)." (Calderón and Duncan, 2003, p. 121).

V. Concluding Remarks

Several papers, including Frenkel (1978), McNown and Wallace (1989), and Mahdavi and Zhou (1994), have shown that past German, Israeli and Latin American hyperinflation contribute to purchasing power parity (PPP) relatively to industrial countries with very stable inflation rates. Usually the approach in these studies has been to estimate any of the 3 stages of PPP tests in Froot and Rogoff (1995). None of these studies, however, has combined techniques in stage 2 with econometric advances in half-lives such as Andrews (1993) and Andrews and Chen (1994). Recent works in this vein by Cashin and McDermott (2003) are confined to industrial countries.

This paper attempts to fill this gap for two very representative real Latin American currencies: CLP and MXN. We find that the evidence on half-lives contrasts markedly across the two currencies: at 5 years or infinity for the Chilean peso and between 1 and 3 years for the Mexican peso. This set of results holds across the three RER datasets. The econometric techniques on the half-lives also provide consistent results with each other: whether the MU estimates by Andrews and Chen (1994) or the half-life calculations based on a correction factor by Rossi (2005). With the MXN implying a mean reversion pattern faster than the 3 to 5 years "consensus" by Rogoff (1996), our evidence on MXN is more consistent with research on groups of developing countries that report less than 3 years for the half-lives (Cheung and Lai, 2000). We present in this paper evidence that the abrupt depreciations in MXN and its relatively higher inflation record may have amplified monetary forces in the dynamics of the real exchange rate.

As extension, the incorporation of heterogeneity in disaggregated relative prices seems worthwhile. Recent work by Imbs et al. (2002) shows that this makes estimates of the half-lives to fall dramatically to little more than 1 year. We leave this research route to future work.

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Figure 1. Bilateral Chilean Peso Log Real Exchange Rates based on PPI and CPI.

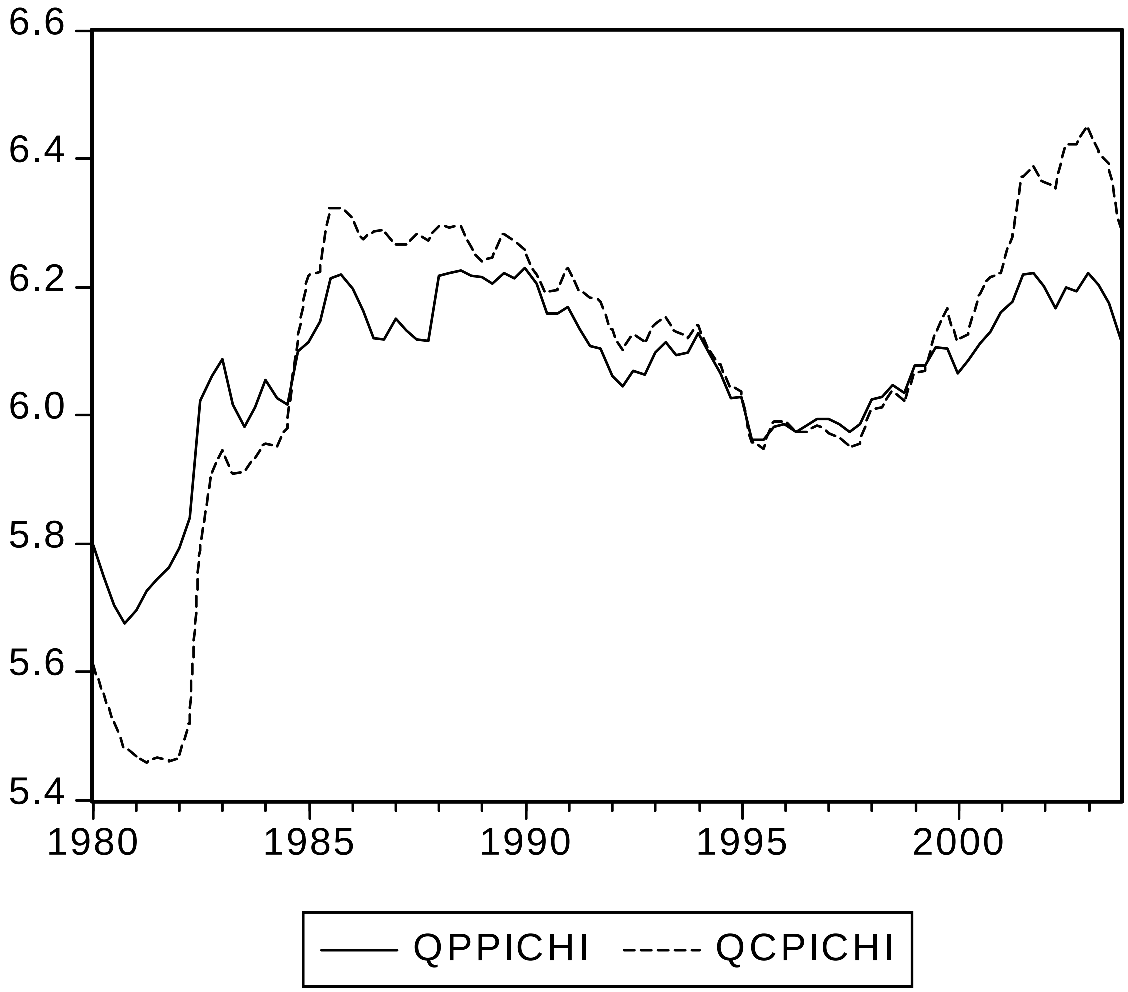

Figure 2. Bilateral Mexican Peso Log Real Exchange Rates based on PPI and CPI.

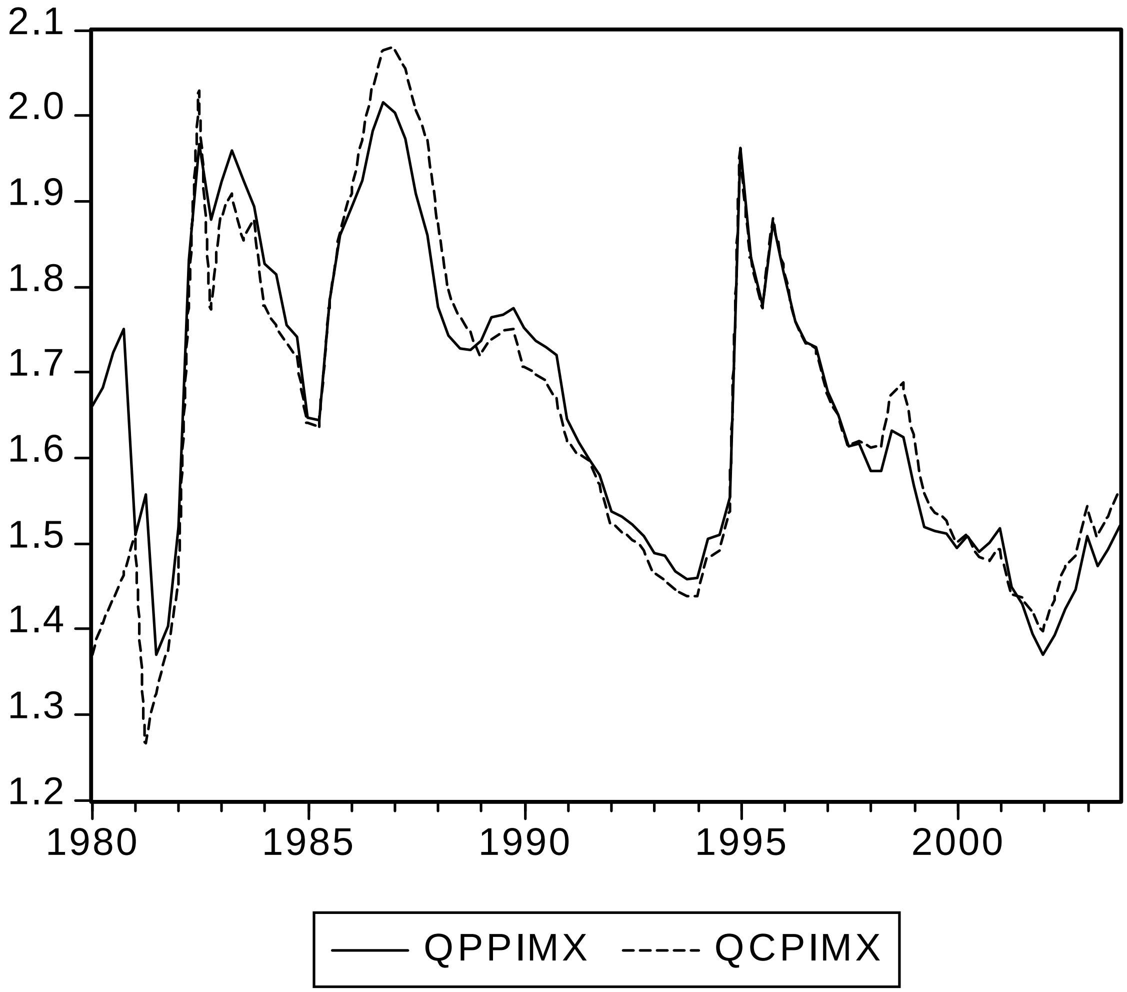

Figure 3. Chilean Peso Log Real Exchange Rates against CPI Quarterly Inflation Rates.

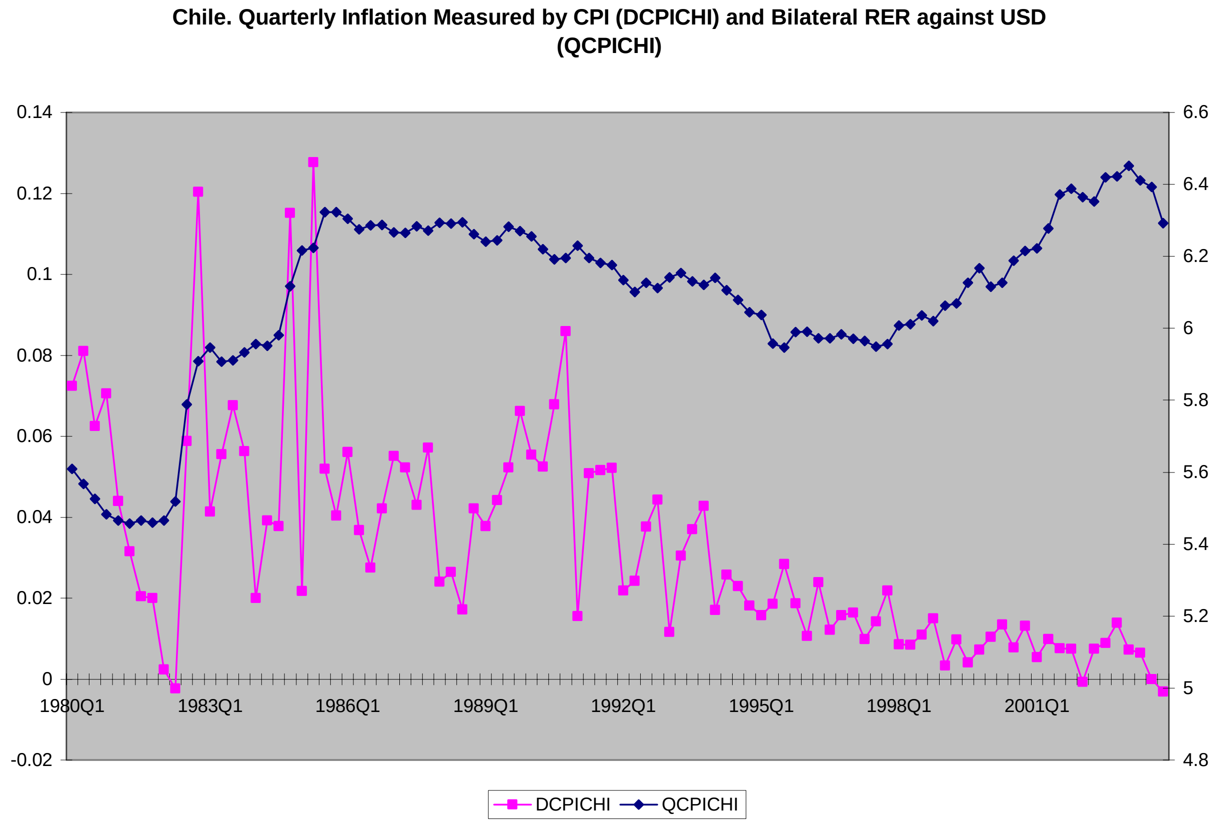

Figure 4. Mexican Peso Log Real Exchange Rates against CPI Quarterly Inflation Rates.

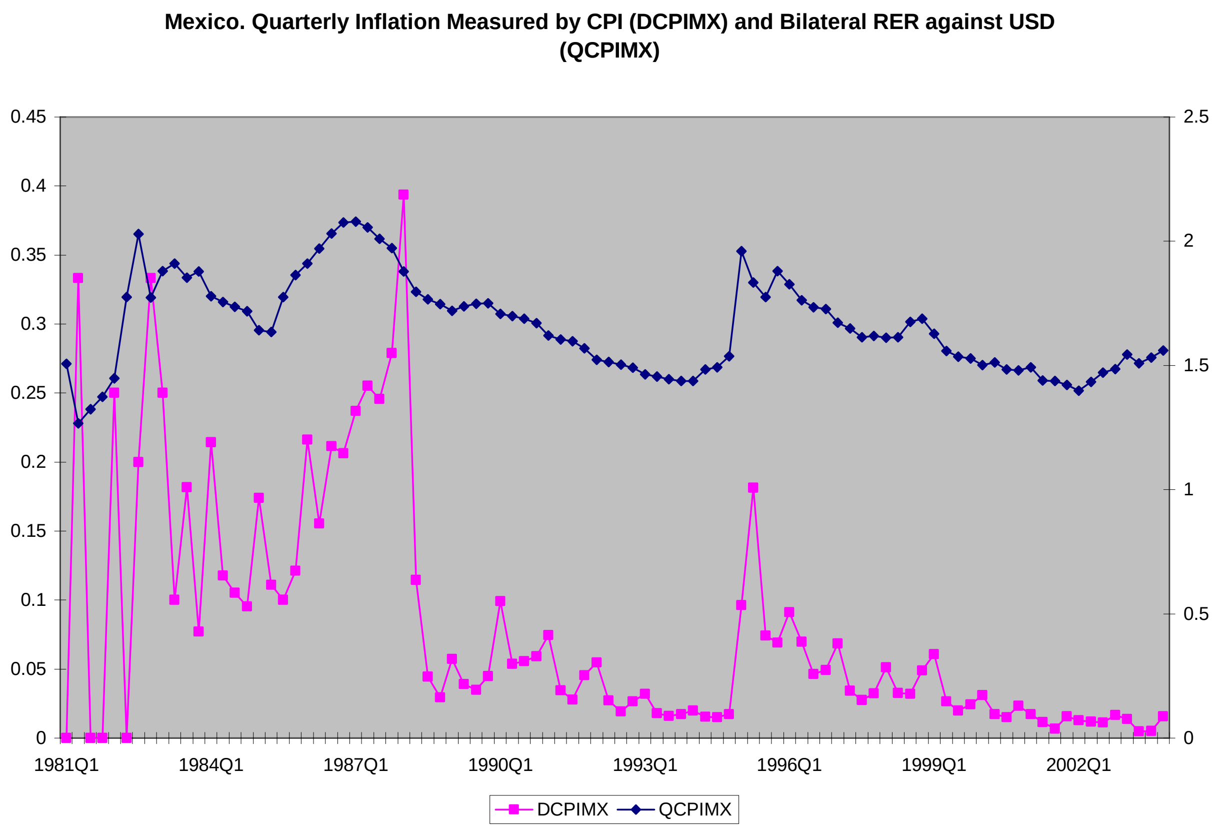

Table 1. Estimates of ARMA (1,1) and ARMA (2,2) Models under Quarterly Data:

ARMA (1,1): $q_t = \alpha_0 + \alpha_1 q_{t-1} + \beta_1 \varepsilon_{t-1} + \varpi_t$

ARMA (2,2): $q_t = \alpha_0 + \alpha_1 q_{t-1} + \alpha_2 q_{t-2} + \beta_1 \varepsilon_{t-1} + \beta_2 \varepsilon_{t-2} + \nu_t$

Real Exchange Rate Models	α_0	α_1	α_2	β_1	β_2	**Adj. R^2**	**Ljung -Box Q (12) test**	**LM test = N x R^2**
Chile - PPI								
ARMA (1,1)	6.112***	0.928***		0.268***		0.933	14.395	0.946
	(0.054)	(0.049)		(0.075)			[0.156]	[0.623]
ARMA (2,2)	6.111***	1.682***	-0.72***	-0.577***	-0.15*	0.936	9.884	1.049
	(0.024)	(0.102)	(0.095)	(0.097)	(0.089)		[0.273]	[0.592]
Chile - CPI								
ARMA (1,1)	6.229***	0.954***		0.377***		0.968	12.368	1.928
	(0.127)	(0.044)		(0.100)			[0.261]	[0.381]
ARMA (2,2)	6.116***	1.891***	-0.91***	-0.753***	-0.40***	0.976	14.78*	23.89***
	(0.035)	(0.062)	(0.061)	(0.125)	(0.135)		[0.063]	[0.00]
Mexico -PPI								
ARMA (1,1)	1.656***	0.877***		0.152		0.813	5.116	2.207
	(0.097)	(0.058)		(0.163)			[0.883]	[0.332]
ARMA (2,2)	1.655***	1.445**	-0.53	-0.431	0.06	0.812	3.152	2.387
	(0.062)	(0.648)	(0.546)	(0.730)	(0.183)		[0.317]	[0.303]
Mexico -CPI								
ARMA (1,1)	1.638***	0.820***		0.258*		0.781	16.60*	20.74***
	(0.066)	(0.081)		(0.141)			[0.084]	[0.00]
ARMA (2,2)	1.641***	0.385	0.40	0.718*	0.03	0.799	10.287	14.47***
	(0.074)	(0.326)	(0.247)	(0.414)	(0.226)		[0.245]	[0.0007]

Notes: Data are of quarterly frequency from 1980:01 to 2003:04. The symbols * [**] (***) attached to the figure indicate rejection of the null at the 10%, 5%, and 1% levels, respectively. As diagnostic tests, the Ljung-Box Q (12) test reports the Q statistics, under the null of no serial correlation, constructed from the autocovariances with probabilities adjusted for 2 ARMA terms. Also reported is the Breusch-Godfrey LM serial correlation test with ARMA errors of order up to 2. In the latter, the null hypothesis is of no serial correlation. The statistics are constructed as the number of observations (N) times the R^2. For the serial correlation LM test, the NR^2 statistic has an asymptotic χ^2 distribution under the null.

Table 2. Estimates of ARMA (1,1) and ARMA (2,2) Models under Monthly Data:

ARMA (1,1): $q_t = \alpha_0 + \alpha_1 q_{t-1} + \beta_1 \varepsilon_{t-1} + \varpi_t$

ARMA (2,2): $q_t = \alpha_0 + \alpha_1 q_{t-1} + \alpha_2 q_{t-2} + \beta_1 \varepsilon_{t-1} + \beta_2 \varepsilon_{t-2} + \nu_t$

Series and Models	α_0	α_1	α_2	β_1	β_2	Adj. R^2	Ljung-Box Q (36) test	LM test = N x R^2
Chile - PPI								
ARMA (1,1)	6.111***	0.978***		0.266***		0.979	33.025	0.405
	(0.061)	(0.014)		(0.089)			[0.515]	[0.817]
ARMA (2,2)	6.111***	0.013	0.944***	1.233***	0.260**	0.979	33.185	0.541
	(0.063)	(0.040)	(0.039)	(0.108)	(0.102)		[0.409]	[0.763]
Chile - CPI								
ARMA (1,1)	6.258***	0.987***		0.158		0.989	26.605	5.789*
	(0.161)	(0.012)		(0.217)			[0.813]	[0.055]
ARMA (2,2)	6.133***	1.968***	-0.97***	-0.883***	-0.109	0.990	24.496	3.798
	(0.011)	(0.011)	(0.011)	(0.217)	(0.217)		[0.826]	[0.150]
Mexico - PPI								
ARMA (1,1)	1.663***	0.947***		0.170**		0.924	47.850*	8.216**
	(0.061)	(0.022)		(0.085)			[0.058]	[0.016]
ARMA (2,2)	1.663***	1.170**	-0.198	-0.089	-0.160	0.925	38.012	3.988
	(0.083)	(0.533)	(0.516)	(0.554)	(0.137)		[0.214]	[0.136]
Mexico - CPI								
ARMA (1,1)	1.663***	0.952***		0.141		0.931	62.08***	13.02***
	(0.063)	(0.021)		(0.094)			[0.002]	[0.001]
ARMA (2,2)	1.669***	1.250*	-0.274	-0.194	-0.140	0.931	51.50**	14.79***
	(0.081)	(0.710)	(0.686)	(0.736)	(0.133)		[0.016]	[0.0006]
Chile-REER								
ARMA (1,1)	4.495***	0.983***		0.290***		0.980	35.922	0.296
	(0.100)	(0.013)		(0.052)			[0.378]	[0.862]
ARMA (2,2)	4.499***	0.320*	0.650***	0.959***	0.204***	0.979	36.877	1.187
	(0.096)	(0.172)	(0.171)	(0.172)	(0.077)		[0.253]	[0.552]
Mexico-REER								
ARMA (1,1)	4.496***	0.962***		0.264***		0.953	46.225*	4.586
	(0.101)	(0.017)		(0.074)			[0.079]	[0.101]
ARMA (2,2)	4.496***	0.755	0.209	0.443	-0.083	0.953	41.329	1.667
	(0.126)	(0.583)	(0.562)	(0.561)	(0.133)		[0.125]	[0.435]

Notes: Data are of monthly frequency from 1980:01 to 2003:12. The symbols * [**] (***) attached to the figure indicate rejection of the null at the 10%, 5%, and 1% levels, respectively. As diagnostic tests, the Ljung-Box Q (36) test reports the Q statistics, under the null of no serial correlation, constructed from the autocovariances with probabilities adjusted for 4 ARMA terms. Also reported is the Breusch-Godfrey LM serial correlation test with ARMA errors of order up to 2, constructed as the number of observations (N) times the R^2; NR^2 statistic has an asymptotic χ^2 distribution under the null.

Table 3. Unit Root Tests on (Bilateral) Real Exchange Rates under Quarterly Data.

Series	Trend?	ADF (k)	DF-GLS (k)	KPSS (4)	Ng-Perron MZ_{α} (k)	Ng-Perron MZ_t (k)
Chile - PPI						
q	Yes	-2.21 (12)	-1.29 (0)	0.26***	-3.20 (0)	-1.22 (0)
q	No	-2.21 (12)	-0.59 (0)	0.48**	-0.85 (0)	-0.52 (0)
Δq	No	-2.52 (11)	-2.66 (2)***	0.17	-13.52 (2)**	-2.59 (2)***
Mexico - PPI						
q	Yes	-3.04 (7)	-2.42 (0)	0.84***	-10.99 (0)	-2.34 (0)
q	No	-2.13 (0)	-2.12 (0)**	0.11	-8.58 (0)**	-2.04 (0)**
Δq	No	-8.80 (0)***	-0.88 (2)	0.05	-32.14 (2)***	-4.01 (2)***
Chile - CPI						
q	Yes	-3.55 (12)**	-1.62 (1)	0.27***	-5.19 (1)	-1.60 (1)
q	No	-3.27 (12)**	-0.79 (1)	0.64**	-1.14 (1)	-0.57 (1)
Δq	No	-6.06 (0)***	-2.98 (2)***	0.17	-16.95 (2)***	-2.85 (2)***
Mexico - CPI						
q	Yes	-1.69 (12)	-2.55 (0)	0.15**	-6.63 (0)	-1.81 (0)
q	No	-1.27 (12)	-2.28 (0)**	0.42*	-4.74 (0)	-1.53 (0)
Δq	No	-4.98 (11)***	-4.44 (2)***	0.04	-1.97 (2)	-0.88 (2)

Notes: Data are of quarterly frequency from 1980:01 to 2003:04 and are explained in the data section. For each country, **q** refers to the bilateral (against the U.S.) real exchange rate calculated using baskets of either producer or consumer price indexes taken from the IFS dataset. We include the deterministic trend only when testing the series in levels. ADF(k) refers to the Augmented Dickey-Fuller t-tests for unit roots, in which the null is that the series contains a unit root. The lag length (k) for ADF tests is chosen by the Campbell-Perron data dependent procedure, whose method is usually superior to k chosen by the information criterion, according to Ng and Perron (1995). The method starts with an upper bound, k_{max}=12, on k. If the last included lag is significant, choose k = k_{max}. If not, reduce k by one until the last lag becomes significant (we use the 5% value of the asymptotic normal distribution to assess significance of the last lag). If no lags are significant, then set k = 0. Next to the reported calculated t-value, in parenthesis is the selected lag length. DF-GLS (k) refers to the modified ADF test proposed by Elliott et al. (1996), with the Schwarz Bayesian Information Criterion (BIC) used for lag-length selection. The KPSS test follows Kwiatkowski et al. (1992), in which the null is that the series is stationary and k=4 is the used lag truncation parameter. We report two of the M-tests developed by Ng and Perron (2001) with BIC used for lag-length selection. The MZ_{α} and MZ_t can be viewed as modified versions of the Phillips and Perron (1988) Z_{α} and Z_t tests, which suffer from severe size distortions when the errors have a negative moving average (MA) root. The first step of the method is to construct the DF-GLS modified ADF test proposed by Elliott et al. (1996) and then to compute the MZ_{α} and MZ_t statistics as defined in Ng and Perron (2001). The symbols * [**] (***) attached to the figure indicate rejection of the null at the 10%, 5%, and 1% levels, respectively.

Table 4. Unit Root Tests on Real Exchange Rates under Monthly Data.

Series	Trend?	ADF (k)	DF-GLS (k)	KPSS (4)	Ng-Perron MZ_{α} (k)	Ng-Perron MZ_{t} (k)
Chile - PPI						
q	Yes	-2.00 (1)	-1.51 (0)	0.71***	-4.96 (1)	-1.53 (1)
q	No	-2.21 (1)	-0.67 (1)	0.74***	-1.18 (1)	-0.65 (1)
Δq	No	-12.91 (0)***	-12.30 (0)***	0.16	-172.24 (0)***	-9.27 (0)***
Mexico - PPI						
q	Yes	-5.66 (12)***	-1.81 (2)	0.22***	-6.72 (2)	-1.82 (2)
q	No	-3.24 (12)**	-1.65 (2)*	2.12***	-5.57 (2)	-1.67 (2)*
Δq	No	-4.15 (12)***	-5.50 (4)***	0.10	-28.24 (4)***	-3.76 (4)***
Chile - CPI						
q	Yes	-1.57 (1)	-1.11 (1)	0.75***	-2.61 (1)	-1.11 (1)
q	No	-1.88 (1)	-0.01 (1)	1.77***	0.03 (1)	0.02 (1)
Δq	No	-13.83 (0)***	-3.87 (4)***	0.22	-22.06 (4)***	-3.30 (4)***
Mexico - CPI						
q	Yes	-2.91 (5)	-1.71 (0)	0.41***	-5.84 (0)	-1.69 (0)
q	No	-2.52 (5)	-1.34 (0)	1.05***	-3.67 (0)	-1.33 (0)
Δq	No	-6.68 (4)***	-5.82 (4)***	0.13	-36.78 (4)***	-4.29 (4)***
Chile – REER						
q	Yes	-1.81 (1)	-1.78 (1)	1.07***	-6.34 (1)	-1.78 (1)
q	No	-1.69 (1)	-0.87 (1)	1.66***	-2.26 (1)	-0.89 (1)
Δq	No	-12.98 (0)***	-10.80 (0)***	0.10	-152.35 (0)***	-8.71 (0)***
Mexico – REER						
q	Yes	-3.42 (10)*	-2.07 (2)	0.68***	-8.48 (2)	-2.06 (2)
q	No	-3.06 (10)**	-1.74 (2)*	1.37***	-6.26 (2)*	-1.74 (2)
Δq	No	-6.06 (4)***	-2.59 (4)***	0.06	-9.07 (4)***	-2.03 (4)**

Notes: Data are of monthly frequency from 1980:01 to 2003:12 and are explained in the data section. For each country, **q** refers to either bilateral (against the U.S.) or multilateral (REER) real exchange rate calculated as explained in the data section 2. We include the deterministic trend only when testing the series in levels. ADF(k) refers to the Augmented Dickey-Fuller t-tests for unit roots, in which the null is that the series contains a unit root. The lag length (k) for ADF tests is chosen by the Campbell-Perron data dependent procedure, whose method is usually superior to k chosen by the information criterion, according to Ng and Perron (1995). The method starts with an upper bound, k_{max}=12, on k. If the last included lag is significant, choose k = k_{max}. If not, reduce k by one until the last lag becomes significant (we use the 5% value of the asymptotic normal distribution to assess significance of the last lag). If no lags are significant, then set k = 0. Next to the reported calculated t-value, in parenthesis is the selected lag length. DF-GLS (k) refers to the modified ADF test proposed by Elliott et al. (1996), with the Schwarz Bayesian Information Criterion (BIC) used for lag-length selection. The KPSS test follows Kwiatkowski et al. (1992), in which the null is that the series is stationary and k=4 is the used lag truncation parameter. We report two of the M-tests developed by Ng and Perron (2001) with BIC used for lag-length selection. The MZ_{α} and MZ_{t} can be viewed as modified versions of the Phillips and Perron (1988) Z_{α} and Z_{t} tests, which suffer from severe size distortions when the errors have a negative moving average (MA) root. The first step of the method is to construct the DF-GLS modified ADF test proposed by Elliott et al. (1996) and then to compute the MZ_{α} and MZ_{t} statistics as defined in Ng and Perron (2001). The symbols * [**] (***) attached to the figure indicate rejection of the null at the 10%, 5%, and 1% levels, respectively.

Table 5. Estimates of Half-Lives under Quarterly Data:

$$\text{AR (1): } q_t = \alpha_0 + \alpha_1 q_{t-1} + \varpi_t$$

$$\text{AR (p): } q_t = \alpha_0 + \alpha_1 q_{t-1} + \sum_{i=1}^{k} \beta_i \Delta q_{t-i} + \varpi_t$$

Real Exchange Rate Models	α_1	Implied Half-Life (95% C.I.)	Adj. R^2	Ljung-Box Q(12) test	α^U_1	Implied Unbiased Half-Life (quarters)	Half-Life with Correction (95% C.I.)
Chile - PPI							
AR (1)	0.944***	12.03	0.928	25.668***	0.990	68.62	
	(0.043)	(0, 30.66)		[0.007]			
AR (7)	0.870***	4.98	0.863	7.133	0.965	19.47	6.77
	(0.043)	(1.52, 8.44)		[0.849]			(3.31, 10.23)
Chile - CPI							
AR (1)	0.965***	19.46	0.963	21.616***	1.000	∞	
	(0.038)	(0.61, 61.60)		[0.028]			
AR (4)	0.936***	10.48	0.961	10.415	1.000	∞	20.28
	(0.031)	(0.19, 20.77)		[0.580]			(9.99, 30.57)
Mexico - PPI							
AR (1)	0.905***	6.94	0.811	9.060	0.943	11.81	
	(0.044)	(0.32, 13.57)		[0.616]			
AR (8)	0.900***	6.58	0.816	4.926	0.945	12.25	7.32
	(0.061)	(0, 14.87)		[0.960]			(0, 15.61)
Mexico - CPI							
AR (1)	0.884***	5.62	0.780	14.816	0.920	8.31	
	(0.052)	(0.37, 10.88)		[0.191]			
AR (3)	0.852***	4.33	0.785	9.001	0.885	5.70	6.65
	(0.062)	(0.47, 8.18)		[0.703]			(2.80, 10.51)

Notes: Data are of quarterly frequency from 1980:01 to 2003:04. The symbols * [**] (***) attached to the figure indicate rejection of the null at the 10%, 5%, and 1% levels, respectively. For the AR (p) model, a general to specific procedure is adopted with maximum lag set at 8 quarters. The implied half life is calculated according to the formula HL = $\ln(0.5)/\ln(\alpha_1)$ and the 95% confidence interval is a conventional two-sided interval. As diagnostic tests, the Ljung-Box Q (12) statistic at 12 lags is a test of the null hypothesis that there is no auto-correlation up to order 12. We also applied the Breusch-Godfrey LM serial correlation test NR^2 statistic that has an asymptotic χ^2 distribution under the null. As the results matched the Q(12), we omit these tests. The α^U_1 coefficient is the median unbiased (MU) estimator obtained from table II (for T+1 = 90) associated with model 2 (without the time trend as defined above) in Andrews (1993) for the AR (1) case and by the algorithm proposed by Andrews and Chen (1994) for the AR (p) case. One iteration is employed after the first round of estimation of the algorithm described in Andrews and Chen (1994, p. 191). In the last column, following Rossi (2005), we calculate the half-life taking into account the b (1) correction factor, which is shown to be equal to $b(1) = 1 - \Sigma\beta_i \Delta q_{t-i}$ (i = 1 to k) in the ADF-type regression. The b(1) correction factor enters the calculation of the half-life as: $h^* \equiv \max\{\ln(0.5b(1))/\ln(\alpha_1), 0\}$, which differs from the standard half-life obtained by $h_a \equiv \max\{\ln(0.5)/\ln(\alpha_1), 0\}$.

Table 6. Estimates of Half-Lives under Monthly Data:

$$AR\ (1):\ q_t = \alpha_0 + \alpha_1 q_{t-1} + \varpi_t$$

$$AR\ (p):\ q_t = \alpha_0 + \alpha_1 q_{t-1} + \sum_{i=1}^{k} \beta_i \Delta q_{t-i} + \varpi_t$$

Real Exchange Rate Models	α_1	Implied Half-Life (95% C.I.)	Adj. R^2	Ljung-Box Q(36)	α^U_1	Unbiased Half-Life (months)	Half-Life with Correction (95% C.I.)
Chile - PPI							
AR (1)	0.983***	40.43	0.978	60.694***	1.000	∞	
	(0.013)	(0, 101.54)		[0.000]			
AR (2)	0.981***	36.13	0.979	31.927	1.000	∞	52.04
	(0.010)	(0, 73.77)		[0.663]			(14.41, 89.68)
Chile - CPI							
AR (1)	0.989***	62.67	0.989	39.995	1.000	∞	
	(0.011)	(0, 186.17)		[0.258]			
AR (11)	0.980***	34.31	0.987	19.266	1.000	∞	61.62
	(0.010)	(0.34, 68.28)		[0.990]			(27.65, 95.58)
Mexico - PPI							
AR (1)	0.960***	16.98	0.923	55.186***	0.977	29.75	
	(0.017)	(2.54, 31.42)		[0.016]			
AR (12)	0.940***	11.20	0.930	17.238	0.955	15.07	24.81
	(0.019)	(4.03, 18.38)		[0.997]			(17.63, 31.98)
Mexico - CPI							
AR (1)	0.961***	17.42	0.931	72.059***	0.978	31.22	
	(0.017)	(2.24, 32.61)		[0.000]			
AR (24)	0.955***	15.05	0.934	23.482	0.972	24.40	20.01
	(0.024)	(0, 31.16)		[0.946]			(3.91, 36.11)
Chile - REER							
AR (1)	0.989***	62.67	0.978	68.730***	1.000	∞	
	(0.012)	(0, 197.40)		[0.000]			
AR (6)	0.986***	49.16	0.979	32.557	1.000	∞	56.56
	(0.011)	(0, 125.41)		[0.633]			(0, 132.81)
Mexico - REER							
AR (1)	0.975***	27.38	0.950	69.944***	0.994	115.52	
	(0.014)	(0, 57.81)		[0.000]			
AR (10)	0.957***	15.77	0.956	12.889	0.974	26.35	35.89
	(0.012)	(6.95, 24.59)		[1.000]			(27.07, 44.71)

Notes: Data are of monthly frequency from 1980:01 to 2003:12. The symbols * [**] (***) attached to the figure indicate rejection of the null at the 10%, 5%, and 1% levels, respectively. For the AR (p) model, a general to specific procedure is adopted with maximum lag set at 24 months. As diagnostic tests, the Ljung-Box Q (36) statistic at 36 lags is a test of the null hypothesis that there is no auto-correlation up to order 36. Also reported is the Breusch-Godfrey LM serial correlation test with ARMA errors of order up to 2, with the null hypothesis as no serial correlation. For the serial correlation LM test, the NR^2 statistic has an asymptotic χ^2 distribution under the null. The α^U_1 coefficient is the median unbiased (MU) estimator obtained from table II (for T+1 = 200) associated with model 2 (without the time trend as defined above) in Andrews (1993) for the AR (1) case and by the algorithm proposed by Andrews and Chen (1994) for the AR (p) case. One iteration is employed after the first round of estimation of the algorithm described in Andrews and Chen (1994, p. 191). In the last column, following Rossi (2005), we calculate the half-life taking into account the b (1) correction factor, which is shown to be equal to $b(1) = 1 - \Sigma\beta_i \Delta q_{t-i}$ (i = 1 to k) in the ADF-type regression and enters the calculation of the half-life as: $h^* \equiv \max\{\ln(0.5b(1))/\ln(\alpha_1), 0\}$, which differs from the standard half-life obtained by $h_a \equiv \max\{\ln(0.5)/\ln(\alpha_1), 0\}$.