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Deviations of Exchange Rates from Purchasing Power Parity: A Story Featuring Two Monetary Unions

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We examine the mean-reverting properties of real exchange rates, by comparing the unit root properties of a group of international real exchange rates with two groups of intranational real exchange rates. Strikingly, we find that while the international real rates taken as a group appear mean reverting, the intranational rates are not. This is consistent with the view that while nominal shocks may be mean reverting over the medium term, underlying real factors do generate long-term trends in real exchange rates. [JEL C12, C23, F31]

The proposition that exchange rates are volatile when allowed to float freely has become something of a stylized fact in the international finance literature (see, for example, Frenkel and Mussa, 1980; MacDonald and Taylor, 1992; and Frankel and Rose, 1995). Indeed, the volatility of exchange rates during the recent floating experience has led economists to advocate moving from an international monetary regime based on flexible exchange rates toward one based on greater exchange rate fixity (McKinnon, 1988; Mundell, 1992; and Williamson, 1987) and is also one of the central arguments made by proponents of greater monetary integration in Europe. The volatility of nominal exchange rates has also had implications for the behavior of real exchange rates. In particular, because prices in goods markets are generally regarded as being sticky (certainly in the short run), volatility in nominal exchange rates is transferred into comparable real exchange rates. This violation of purchasing power parity (PPP) may be viewed as a second stylized fact in international finance.

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The failure of PPP to hold continuously is well documented empirically (see the summaries in Froot and Rogoff, 1995, and MacDonald, 1995). However, there is now growing evidence to suggest that although PPP does not hold on a month-to-month or quarter-to-quarter basis, it does hold as a long-run phenomenon (see, for example, Edison, 1987; Frankel, 1988; and Diebold, Husted, and Rush, 1991). The main explanation for this follows on directly from our discussion in the previous paragraph and the perceived source of the deviations from PPP. If the predominant force upsetting the PPP relationship is nominal, this will have only a transitory effect on deviations from PPP (this is essentially the story in the seminal Dornbusch, 1976, model). If, however, the source of PPP disturbances are truly "real" in nature (as suggested by Stockman, 1987), we argue this will have a permanent, or more permanent, effect on the real exchange rate.[1]

In this paper, we propose a way of gaining a perspective on the importance of nominal shocks in generating deviations from PPP. We do this by comparing the behavior of relative prices across countries with those within countries.[2] While exchange rates across countries include both real and nominal disturbances, it appears reasonable to assume that relative prices movements within countries are dominated by real factors, with little or no nominal influence.[3] In this context, it is interesting to know whether relative prices within countries are dominated by long-run trends, and hence are nonstationary or not.

Our method involves constructing a panel data set for the real exchange rates of 20 countries and comparing the time-series properties of these data with comparable data sets within two monetary unions (namely, the United States and Canada). We confirm the findings of others that relative price variability within countries is considerably lower than across countries,[4] and that real exchange rates appear stationary, or mean reverting, across countries.[5] However, we also find that relative prices within countries are nonstationary. The implication is that underlying real factors can create long-run trends in relative prices even within a fairly homogeneous economic environment. The implication for real exchange rates is that, although they may appear stationary in the longer run compared with their short-term behavior, these results in all probability mask long-run trends caused by real behavior.

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[1]The issue of whether the mean reversion observed in long runs of data—a half-life of four years is the standard finding (see MacDonald, 1995)—is in fact consistent with purely nominal shocks is not uncontroversial. Rogoff (1996), for example, argues it is not. One way in which such reversion could be consistent with purely nominal shocks is if the initial real exchange rate deviation is not immediately offset because of the pricing-to-market policies pursued by multinational companies and the inability of agents to arbitrage away potentially profitable misalignments. That is the interpretation we offer.

[2]Previous work comparing the behavior of real exchange rates in inter- and intranational data sets includes Engel and Rogers (1996) and Engel, Hendrickson, and Rogers (1998).

[3]Although price connections might be different within and between countries, we believe that the fundamental distinguishing feature of an intra- and intercountry comparison is the absence of differential nominal disturbances within a monetary union.

[4]See, for example, Vaubel (1978), Eichengreen (1992), Bayoumi and Thomas (1995), Engel (1993), and Parsley (1996).

[5]See, for example, Frankel and Rose (1996) and MacDonald (1996a).

I. Panel Unit Root Methods

We use panel unit root methods to compare the time-series properties of real exchange rates within and across countries. Such tests have a clear statistical advantage over univariate tests, such as the Dickey-Fuller class of statistics, because they have greater power to reject the null of a unit root when it is in fact false. Panel unit root tests may be motivated by considering the following regression equation:[6]

$$\Delta q_{it} = \alpha_i + \delta q_{t-1} + \sum_i \gamma_i D_i + \sum_t \gamma_t D_t + \sum_i \beta_i t_i + \upsilon_{it}, \tag{1}$$

where q denotes a real exchange rate, i denotes a currency, D_i and D_t denote, respectively, country-specific and time-specific fixed effects dummy variables, t_i denotes a country-specific time trend, α_i, δ, γ_i, γ_t, and β_i are estimated coefficients, and υ_{it} is an error term.[7] Equation (1) is essentially the panel analogue to the standard Dickey-Fuller autoregression. Of particular interest is the magnitude of δ, which indicates the speed of mean reversion and its significance as judged by the estimated *t*-ratio. As Levin and Lin (1992 and 1993) have demonstrated, the critical values for the *t*-ratio are affected by the particular deterministic specification used.

In circumstances where all of the deterministic elements in equation (1) are excluded apart from the single constant term, α, Levin and Lin (1992) demonstrate that the *t*-statistic on δ converges to a standard normal distribution. Including individual specific effects—either $(\Sigma_i \gamma_i D_i)$ or $(\Sigma_i \beta_i t_i)$ or both—but excluding time-specific intercepts, Levin and Lin (1992) demonstrate that the *t*-ratio converges to a noncentral normal distribution, with substantial impact on the size of the unit root test (and they tabulate critical values). However, Levin and Lin (1993) argue that unless there are very strong grounds for exclusion, time-specific intercepts should always be included in these kinds of panel tests. The reason for this is that the inclusion of such dummies is equivalent to subtracting the cross-sectional average in each period. This subtraction may be dispensed with in cases where the units in the panel are independent of each other; however, in cases where this is not the case such a subtraction is vital to ensure independence across units.

In addition to facilitating the removal of time means, the panel methods of Levin and Lin (1993) have a number of other advantages, such as allowing the residual term to be heterogeneously distributed across individuals (in terms of both nonconstant variance and autocorrelation), rather than a white noise process. The testing method has the null hypothesis that each individual time series in the panel has a unit root, against the alternative that all individual units taken as a panel are stationary. The procedure consists of four steps, which we now briefly note (these steps do not correspond exactly to the steps in Levin and Lin).

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[6] A similar equation forms the basis of a cross-country panel study by Frankel and Rose (1996).

[7] Hence our tests are robust to the criticism made of earlier studies by Papell (1997). Additionally, our subtraction of cross-sectional means for each time period addresses the point made by O'Connell (1997) that the q_i series within a given panel are not independent. Husted and MacDonald (1997) have demonstrated that, having controlled for cross-sectional means, the use of a S.U.R.E.-type estimator makes little difference to the adjusted *t*-ratios reported in this paper.

The first step involves subtracting the cross-sectional mean from the observed exchange rate series. Thus we have q_{it}, where i runs from 1 to N and N denotes the total number of real exchange rates in the panel. We construct $\overline{q}_t = (1/N)\Sigma_{i=1}^{N} q_{it}$. In the following steps, the term q_{it} is interpreted as having been adjusted by $\overline{q}_t$.

Step 2 involves performing regression (2) and (3) for the de-meaned data:

$$\Delta q_{it} = \sum_{L=1}^{P_i} \hat{\pi}_{iL} \Delta q_{it-L} + \hat{a}_{mi} d_{mt} + \hat{e}_{it}, \tag{2}$$

$$q_{it-1} = \sum_{L=1}^{P_i} \hat{\phi}_{iL} \Delta q_{it-L} + \hat{a}_{mi} d_{mt} + \hat{v}_{it-1}; \tag{2'}$$

and then constructing the following regression equation:

$$\hat{e}_{it} = \delta_i \hat{v}_{it-1} + \varepsilon_{it}. \tag{3}$$

The t-ratio calculated on the basis of δ_i is the panel equivalent to an augmented Dickey-Fuller (ADF) statistic. To control for heterogeneity across individuals, both $\hat{e}_{it}$ and $\hat{v}_{it-1}$ are deflated by the regression standard error from equation (3); these adjusted errors are labeled $\tilde{e}_{it}$ and $\tilde{v}_{it-1}$. Under the null hypothesis, these normalized innovations should be independent of each other; this can be tested by running the following regression as step 3:

$$\tilde{e}_{it} = \delta \tilde{v}_{it-1} + \tilde{S}_{it}. \tag{4}$$

Under the null hypothesis that $\delta_i = 0$ for all $i = 1,\ldots N$, the asymptotic theory in Section 4 of Levin and Lin (1993) indicates that the regression t-statistic, t_δ, has a standard normal distribution in a specification with no deterministic terms, but diverges to negative infinity in models with deterministic terms. However, Levin and Lin (1993) demonstrate that the following adjusted t-ratio has an $N(0,1)$ distribution and the critical values of the standard normal distribution can be used to test the null hypothesis that $\delta_i = 0$ for all $i = 1,\ldots, N$:

$$t_\delta^* = \frac{t_\delta - N\tilde{T}\,\hat{S}_{NT}\sigma_\epsilon^{-2}\, RSE(\delta)\, \mu_{mT}^*}{\sigma_{m\tilde{T}}^*}. \tag{5}$$

The terms in equation (5), other than t_δ, are calculated in step 4. Specifically, $\hat{S} = (1/N)\Sigma_i^n \hat{q}_i$, where $\hat{q}_i = \hat{\sigma}_{qi}/\hat{\sigma}_{ei}$, $\hat{\sigma}_{ei}$ is the residual standard error from equation (4) and $\hat{\sigma}_{qi}$ is an estimate of the long-run standard deviation of q_i, $RSE(\delta)$ is an estimate of the reported standard error of the least-squares estimate of δ, $\hat{\sigma}_\varepsilon$ is the estimated standard error of regression (4), $\tilde{T} = (T - \overline{p} - 1)$ is the average number of observations per individual in the panel, and

$$\overline{p} = (1/N)\sum_{i=1}^{N} p_i \tag{6}$$

is the average lag order for the individual ADF statistics. $\sigma^*_{m\bar{T}}$ and $\mu^*_{m\bar{T}}$ represent the mean and standard deviation adjustments, respectively, and are tabulated in Table 1 of Levin and Lin (1993) for different deterministic specifications.

II. Data Sources and Data Description

In line with our earlier discussion we have constructed three annual data sets: an international data set and two intranational data sets. The international data set consists of two real bilateral exchange rates defined for 20 countries relative to the United States (the countries are listed in Table 1), constructed using relative wholesale and consumer prices (which are the most widely studied real exchange rates in the literature). The international data run from 1973 through 1993, and the wholesale and consumer prices and the exchange rates are taken from the IMF's *International Financial Statistics*.[8]

The two monetary unions we focus on are Canada and the United States. For the former country, real exchange rates are defined using consumption indices, while for the United States production indices are used. Although these series were chosen because of their availability, there is a debate in the literature regarding the most appropriate price series to use in defining a real exchange rate (see, for example, Frenkel, 1978). Since our chosen indices may be interpreted as representing two extreme forms of price series, they should help to determine if a particular intracountry result is driven by the choice of price index or is independent of the index used. More specifically, for Canada we have collected data on provincial nondurable consumption, and the real exchange rate is measured as (the log of) the relative price of a particular province with respect to Ontario. The Canadian sample period is 1972–94. The U.S. data consist of gross state product data for 48 states (we exclude Alaska and Hawaii) and the real exchange rates are constructed relative to New Jersey (again in logs).[9] The total U.S. sample period runs from 1963 through 1992. We have used this full sample but, to be consistent with the international data sample, we also constructed a subsample corresponding to the recent floating period.[10] We believe it is important to run our panel tests for a variety of samples since it is well known that the panel estimators we use are most efficient when the dimensions of the panel are approximately square; that is, when the cross-sectional dimensions are approximately equal to the time-series dimensions. For the international data set, this will be true for the recent floating period and it will also be approximately true for the U.S. data over the full sample period.

Before conducting formal tests, it is useful to examine some of the properties of the data graphically. Given our interest in the mean-reverting properties of the data, we compare the relationship between current and future movements in real exchange

[8]The wholesale price series is line 62, the consumer price line series is line 63, and the exchange rate is line ae. As our interest is in the low frequency characteristics of the data, annual data are sufficient.

[9]As the regressions use dummy variables for each year, the choice of numeraire has no impact on the results. New Jersey was selected as it has been used in the numeraire in some other studies using intrastate data from the United States.

[10]We also experimented with a subsample of 20 large states, to see if the results were sensitive to the number of regions being considered. As the results were very similar to those with the full sample, they are not reported.

rates. Accordingly, the data in the three panels were divided into successive five-year periods, and the change in the logarithm of the real exchange rate in the first year was compared with the average change over the next four years within each five-year period for each country/state/province.

The results from this exercise are shown in Figure 1. Two differences are immediately apparent between the international data set and those for the United States and Canada. The first is the much larger degree of real exchange rate movements in the international data, consistent with the notion that international relative prices are driven by larger underlying shocks than their intranational counterparts. The variability for the Canadian data set is also considerably smaller than that for the United States, presumably because of the much stronger forces toward equalization of the consumer prices used in the Canadian panel compared to the producer prices used in the U.S. panel. By contrast, in the international data set differences in behavior across alternative types of prices indices are minimal (not reported for the sake of brevity).

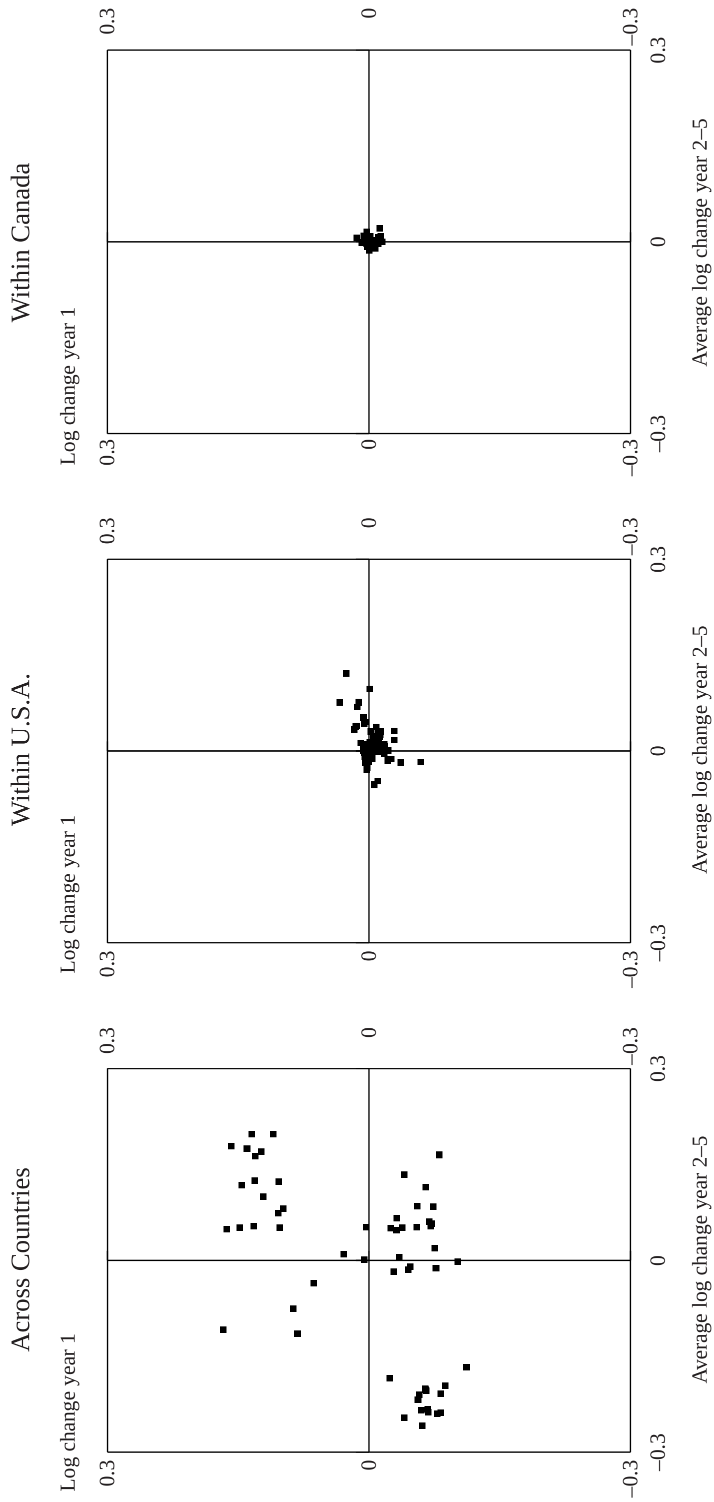
The second difference is in the predictability of future movements in relative prices. The international data show almost no correlation between current relative price movements and movements over the next four years, which is consistent with the weak mean reversion found in most studies of the international data. By contrast, current increases in relative prices between U.S. states are clearly positively correlated with further increases in relative prices in the future, implying that changes in relative prices have considerable momentum over time (the behavior across Canadian provinces is difficult to assess because of the much lower level of variability). In short, the intranational data show much less evidence of mean reversion. The next section examines this issue more formally using the panel unit root tests discussed earlier.

III. Univariate and Panel Unit Root Results

Before implementing the panel unit root tests we examine the univariate unit properties of each of the real exchange rates using standard ADF statistics. These results for our range of real exchange rates are reported in Tables 1 through 3. With very few exceptions the international data set, reported in Table 1, confirms the now standard result that on a univariate basis, and for the recent float, real exchange rates are nonstationary variables. Tables 2 and 3 confirm that this international result also holds for real exchange rates within our two monetary unions. What happens, though, when we take these groupings as panels? Two aspects of our panel results should be emphasized: first, the speed of adjustment, as represented by δ, and second, our estimated *t*-ratios. Table 4 reports panel unit root tests for our two international data sets (using CPIs and WPIs), the U.S. panel over the full sample period and the subperiod since 1973, and the Canadian panel.

The estimated adjustment speeds for our different regressions are reported in the rows labeled "δ" in Table 4 (the layout of the table is explained in the footnotes). It is noteworthy that the adjustment speeds in the international and intranational panels are negative and are therefore all indicative of mean reversion. However, adjustment is much more rapid in the international data sets relative to the national ones. For example, the average value across the latter regressions is –0.12, while the average

Figure 1. Behavior of Real Exchange Rates



Sources: See text for more details.

Table 1. Augmented Dickey-Fuller Unit Root Tests: Levels, International Results

	CPI		*WPI*	
Country	t_μ	t_τ	t_μ	t_τ
Australia	–1.98	–2.08	–1.21	–1.99
Austria	–2.28	–2.26	–2.05	–2.13
Belgium	–1.12	–0.79	–1.69	–2.51
Canada	–2.27	–1.46	–1.71	–1.71
Denmark	–1.54	–1.36	–1.58	–1.51
Finland	–4.17*	–4.42*	–2.91	–3.58
France	–3.58*	–5.68*	–1.64	–1.64
Germany	–1.84	–1.76	–1.95	–1.99
Greece	–1.84	–2.09	–1.50	–1.57
Ireland	–2.44	–3.87*	–0;89	–1.62
Italy	–0.46	–1.45	–0.91	–3.27
Japan	–1.77	–2.56	–1.44	–2.26
Netherlands	–1.87	–1.53	–1.22	–2.28
New Zealand	–2.05	–1.89	–2.14*	–3.05
Norway	–3.92*	–3.77*	–4.29*	–4.34*
South Korea	–1.80	–2.14	–2.28	–2.29
Spain	–1.46	–1.90	–0.83	–1.53
Sweden	–2.47	–2.44	–1.66	–0.88
Switzerland	–3.12*	–2.99	–2.89	–3.16
United Kingdom	–1.10	–1.61	–1.41	–1.45

Notes: The numbers in the columns labeled t_μ and t_τ are ADF *t*-ratios from an autoregression with, respectively, a constant and a constant plus a time trend included. An asterisk denotes significance at the 5 percent level.

for the international data sets is two and a half times greater at –0.29. These figures translate into half-lives of two and six years, respectively, for the international and intranational data sets. The average half-life from our international data sets is shorter than the estimates reported by Frankel and Rose (1996) (they report average half-lives of four years), but nevertheless reinforces the importance of using panel data when defining PPP deviations. The average half-life for the intranational data,

Table 2. Augmented Dickey-Fuller Unit Root Tests:
U.S. Results

State	t_μ	t_τ
New England		
Connecticut	–1.98	–1.76
Maine	–1.81	–1.68
Massachusetts	–1.53	–1.53
New Hampshire	–0.75	–0.99
Rhode Island	–2.08	–2.00
Vermont	–1.42	–0.91
Mid-East		
Delaware	–2.22	–1.14
Maryland	–1.19	–1.75
New York	–1.69	–1.60
Pennsylvania	–1.73	–2.61
Great Lakes		
Illinois	–1.53	–1.56
Indiana	–0.37	–0.86
Michigan	0.07	–0.91
Ohio	–2.62	–3.09
Wisconsin	–0.47	–1.13
Plains		
Iowa	–0.18	–1.18
Kansas	–0.73	–1.48
Minnesota	–1.65	–0.95
Missouri	–0.85	–1.53
Nebraska	–0.22	–1.10
North Dakota	–1.18	–1.32
South Dakota	–1.97	–1.61
Southeast		
Alabama	–1.96	–2.11
Arkansas	–1.34	–0.92
Florida	–0.95	–0.79
Georgia	–1.92	–0.93
Kentucky	–0.87	–1.59
Louisiana	–0.95	–1.03
Mississippi	–1.63	–1.71
North Carolina	–1.59	–1.07
South Carolina	–3.04*	–3.02
Tennessee	–0.39	–1.41
Virginia	–0.27	–1.56
West Virginia	–1.28	–1.90
Southwest		
Arizona	–1.31	–1.05
New Mexico	–0.95	–0.79
Oklahoma	–1.74	–1.25
Texas	–1.71	–1.39

Table 2. (concluded)

State	t_μ	t_τ
Rocky Mountains		
Colorado	–1.73	–1.50
Idaho	–2.02	–1.14
Montana	–1.76	–1.66
Utah	–1.62	–1.25
Wyoming	–1.73	–0.63
Far West		
California	–1.61	–1.26
Nevada	–1.97	–1.11
Oregon	–2.19	–0.72
Washington	–2.04	–1.83

Notes: See Table 1.

Table 3. Augmented Dickey-Fuller Unit Root Tests: Canadian Results

Province	t_μ	t_τ
Alberta	–1.41	–1.97
British Columbia	–1.68	–2.51
Manitoba	–0.75	–0.99
New Brunswick	–2.08	–2.00
Newfoundland	–1.42	–0.91
Nova Scotia	–2.22	–1.14
Prince Edward Island	–1.19	–1.75
Quebec	–1.69	–1.60
Saskatchewan	–1.73	–2.61

Notes: See Table 1.

although much shorter than the international value, is still longer than the average value culled from single-country estimates for the recent float (which would imply a very long half-life of about 20 to 30 years—see MacDonald, 1995). However, a crucial issue is whether the mean reversion exhibited in our panel data sets is statistically significant; that is, are the negative adjustment speeds significantly different from zero or not?

The estimated unadjusted *t*-ratio, that is, t_δ, is in all cases larger in absolute value than –4.0 and, in terms of the original Levin and Lin (1992) critical values, these *t*-ratios would be statistically significant. However, as we have noted, the unadjusted *t*-

Table 4. Panel Unit Root Tests

International Panel	*INT/CPI*	*INT/WPI*
t_δ	–8.23	–8.72
t^*_δ	–2.28	–2.58
2-tail	(0.02)	(0.00)
1-tail	(0.01)	(0.00)
δ	–0.276	–0.308
U.S. Panel	*US/47/full*	*US/47/sub*
t_δ	–10.11	–10.21
t^*_δ	–1.04	0.49
2-tail	(0.29)	(0.62)
1-tail	(0.14)	(0.31)
δ	–0.079	–0.146
Canadian Panel	*Province/Ontario*	
t_δ	–4.47	
t^*_δ	–0.39	
2-tail	(0.69)	
1-tail	(0.34)	
δ	–0.126	

Notes: The numbers in the rows labeled t_δ and t^*_δ are, respectively, the unadjusted and adjusted panel unit-root *t*-ratios, defined in the text. The latter statistic has a standard normal distribution; numbers in parentheses are marginal significance levels. The numbers in the rows labeled δ are the adjustment speeds defined in the text. The columns labeled *INT/CPI* and *INT/WPI* denote the international panels using, respectively, consumer and wholesale prices to define the real exchange rate. The columns labeled *US/47/full* and *US/47/sub* denote the U.S. panel samples over the full and subsample period (see text for further details). The column headed *Province/Ontario* denotes the Canadian real exchange rates defined for each province with respect to Ontario.

ratios are biased to minus infinity and it is not appropriate to draw inferences on the basis of these test statistics. Interestingly, the estimated adjusted *t*-ratios—the t^*_δ values—give a dramatically different picture. For all of the currency union samples the estimated value of t^*_δ is insignificantly different from zero, but for the international data set both real exchange rate data sets produce statistically significant adjusted *t*-ratios. Given that we use two very different price series for the monetary unions, we do not believe our results are a result of the particular series used. We offer an interpretation in the following concluding section.

IV. Conclusion

Recent empirical work on the behavior of exchange rates has gone through a number of distinct phases. The first phase involved testing the hypothesis that rates were a random walk, and hence unpredictable in the long run. More recent work indicates that while the random-walk model is a reasonably good approximation to short-run dynamics, real exchange rates show mean-reverting tendencies over the medium to long term.

The evidence in this paper can be seen as adding a further layer of complexity to this story. To abstract from the nominal factors, which are often thought to generate much of the short-term dynamics, we studied the behavior of relative prices across regions within a country. The results indicate that these relative prices have significant long-run trends. This implies that underlying real factors can create long-run trends in relative prices even in a fairly homogeneous environment. The implication we draw is that, while *nominal shocks* may be mean reverting over the medium term, generating the observed mean reversion in real exchange rates, this medium-term effect obscures the fact that underlying *real factors* generate long-term trends in real exchange rates. The next task for empirical researchers is to identify and quantify these effects.[11]

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