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# Purchasing Power Parity in High-Inflation Countries: A Cointegration Analysis of Integrated Variables with Trend Breaks

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This paper examines the long-run validity of purchasing power parity (PPP) for four high-inflation countries. The method of Zivot and Andrews (1992) is employed to detect the time-series behavior of the exchange rates and consumer price indices of these countries. We find that these variables are integrated with some trend breaks. We then utilize these data to test PPP using Johansen's (1988) multivariate cointegration technique. The cointegration tests are conducted with the correction of the finite sample bias and the adjustment for trend breaks. The results are consistent with the argument that, during the recent floating exchange-rate period, PPP holds well, at least in a weak form, in high-inflation countries where the general price level movement overshadows the factors causing deviations from PPP.

## 1. Introduction

Purchasing power parity (PPP) is an important building block for international economic modeling. The absolute version of PPP (APPP) posits a long-run relation between the bilateral exchange rate and price levels of two relevant countries, while relative PPP (RPPP) suggests comovements of changes in the exchange rate with the inflation differential of two countries. An important use of PPP is that it may serve as a guide for monetary authorities when they intervene in the foreign exchange market to move the exchange rate toward the level consistent with PPP.

Recent developments in time-series analysis provide some advanced techniques to examine the statistical behavior of economic series and to test the existence of a long-run relationship among integrated variables. Several recent studies have applied new econometric techniques to examine PPP. Among them, Taylor and McMahon (1988), Abuaf and Jorion (1990), Kim (1990), Ardeni and Lubian (1991), Choudhry, McNown, and Wallace (1991), Diebold, Husted, and Rush (1991), Glen (1992), and Cheung and Lai (1993a) utilize the unit root tests, variance ratio tests, and fractional differencing or fractional cointegration analysis and find some evidence favoring PPP. Their studies are for the 1920s floating exchange rate period, for the 1950s Canadian float using monthly or quarterly data, or for long historical periods of different ex-

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change-rate systems employing long low-frequency (annual) data. However, when using monthly or quarterly data and applying univariate unit root tests to real exchange rates or Engle–Granger (Engle and Granger 1987) bivariate cointegration tests to exchange rates and relative price levels, Baillie and Selover (1987), Corbae and Ouliaris (1988), Taylor (1988), Mark (1990), Layton and Stark (1990), and Flynn and Boucher (1993) fail to provide support for long-run PPP over the recent floating-rate period.

In more recent studies, Cheung and Lai (1993b), Kugler and Lenz (1993), and MacDonald (1993) utilize the multivariate cointegration methodology proposed by Johansen (1988) and Johansen and Juselius (1990) to test the long-run validity of PPP in a trivariate framework.[1] They find evidence more favorable to long-run PPP during the recent floating-rate period than the findings of other research applying the Engle–Granger regression methodology. They demonstrate that there is a long-run relationship between a number of bilateral exchange rates and their corresponding relative prices, although the hypothesis of proportionality of the exchange rate with respect to relative prices does not receive much support from the data.

Most of the cointegration studies of PPP use the data of industrial countries. There has been only limited effort examining the relevance of long-run PPP for less-developed countries (LDCs). Liu (1992) employs the cointegration analysis and long historical data (from the late 1940s to the end of the 1980s) that encompass periods of different exchange-rate arrangements to study PPP for some high-inflation countries of Latin America. His findings again do not provide general support for the proportionality hypothesis of PPP but are consistent with the hypothesis of the existence of a long-run relationship between some bilateral exchange rates and relevant price variables. Unfortunately, the results from other studies for LDCs using the data of the flexible exchange-rate period are not as favorable to PPP as those of Liu (1992). McNown and Wallace (1989) show some evidence of cointegration between the exchange rate and the wholesale price index (WPI) in two out of four high-inflation countries but find no evidence of cointegration in any of the four cases when the consumer price indices (CPIs) are employed. Bahmani-Oskooee (1993) uses the concept of an effective exchange rate to examine the experience of 25 LDCs, including both high-inflation and low-inflation countries. He finds little empirical support for PPP for most countries.

Mahdavi and Zhou (1994) notice that the time series properties of the variables of high-inflation LDCs might be different from those of industrial countries. After conducting the standard stationary tests for the data of 13 countries with moderately high to very high inflation rates over the flexible-rate period, they conclude that, for most of the countries with very high inflation, exchange rates and price ratios, with the U.S. as the base country, are integrated of order two, or I(2) for short. Similar findings are also reported in Bahmani-Oskooee (1993). Bahmani-Oskooee then considers the results that the residuals from the cointegrating regressions of I(2) variables are integrated to a degree less than two, even if they are still nonstationary, as the evidence supporting PPP. Unlike Bahmani-Oskooee (1993), Mahdavi and Zhou (1994) directly test two versions of PPP for two groups of countries using the Johansen maximum

[1] Johansen and Juselius (1992) test the PPP relation and the uncovered interest parity (UIP) relation for the United Kingdom in a five-dimensional system of equations (two prices, exchange rate, and two interest rates). Their results reject the hypothesis that the PPP relation is stationary by itself, but they are consistent with the hypothesis of a stationary PPP relation with a combination of the two interest rates.

likelihood methodology; that is, they apply the cointegration tests to the *levels* of the exchange rate and price ratio for the group of countries whose variables are found to be integrated of order one, or I(1) for short, to test for the absolute version of PPP. At the same time, they investigate the validity of relative PPP as a long-term equilibrium relationship between the change in the exchange rate and the inflation differential for another group of countries whose exchange-rate and price-ratio series seem to be I(2). They show support for RPPP for all countries in the second group. Yet APPP holds for only three out of eight countries in the first group when the wholesale price indices are used.

A question arises from the results and conclusions of Bahmani-Oskooee (1993) and Mahdavi and Zhou (1994). If the residuals from the cointegrating equations of I(2) variables are still nonstationary or if the RPPP relation is stationary while the APPP relation is not, does this imply that the APPP relation is basically nonstationary for high-inflation countries and therefore their exchange rates and price levels permanently wander from a stable long-run relationship? If this is the case, the claim by Mahdavi and Zhou that PPP holds in high-inflation countries is not well supported by their study, nor has the hypothesis that APPP holds as a long-run equilibrium relationship in LDCs over the flexible-rate period received much support from the existing studies, especially when the consumer price indices are utilized.

It has long been argued that deviations from APPP may result from a number of factors, including transport costs and trade restrictions, the existence of nontraded goods and services, relative price changes, differential speeds of adjustment in the currency exchange markets and the goods markets, as well as the problems of price-level measurement associated with aggregation and index constructions (see Edison 1987, p. 382; Melvin 1992, pp. 124–127). That PPP often holds better for the WPI pairs than the CPI pairs could be explained by the fact that the CPI does not include exported goods and thus is weighted more toward nontraded goods than is the WPI.

The effects of most of the factors listed above will be less important in the period during which a country experiences high inflation. It is generally believed that PPP should hold better in high-inflation countries where the disturbances to their economies are mostly monetary in origin and the relative price effects are shadowed by the general price level movements (see Melvin 1992, pp. 123–124). However, the existing studies mentioned above for the recent flexible-rate period have not offered strong evidence favoring APPP in less-developed, high-inflation countries. One possible reason for their failure to support APPP is that these studies might fail to detect the time-series properties of the variables of high-inflation countries, and consequently they are unable to appropriately model these variables when they conduct the tests.

For example, the issue of whether the exchange rates and price variables of some high-inflation countries (HICs) are I(2) is not beyond controversy. If these variables are I(2), their first differences would be nonstationary I(1) series. To have a visual analysis, in Figure 1, we plot the log differences of the U.S. consumer price index and the variables of four countries that experienced very high inflation in the last two decades. Among these countries, the exchange rates and price variables of Brazil, Israel, and Mexico are shown in Mahdavi and Zhou (1994) to be I(2), while the price variable of Zaire is found to be either I(1) or I(2). The data plotted in Figure 1 are taken from the *International Financial Statistics* (various issues) of the International Monetary Fund (IMF). The exchange rates (ERs) are the end-of-the-period market rates in terms of units of domestic currency per U.S. dollar and the price variables (P, domestic price, or USP, the U.S. price level) are the consumer price indices.[2] Due to the availability of the data, the sample periods for these countries are varied. When we take a close look at these plots of the first differences of the logged variables (dlnER, dlnP, and dlnUSP), two interesting characteristics of the data can be seen. One is the comovement pattern of dlnER and dlnP for each country and the other is the existence of some apparent trend breaks in the data. This is not beyond expectation because, during the sample periods, all these HICs experienced some dramatic changes in their economies. Naturally, one may ask the question: Are the first differences of these variables really nonstationary I(1) series or are they actually stationary but with some trend breaks?

Perron (1989), Christiano (1992), Banerjee, Lumsdaine, and Stock (1992), and Zivot and Andrews (1992) point out that the standard augmented Dickey–Fuller (ADF) tests (Dickey and Fuller 1981), which are used in Bahmani-Oskooee (1993) and Mahdavi and Zhou (1994), are not appropriate for the variables with apparent structural breaks. Therefore, it might be incorrect to claim that the first differences of the variables of some HICs are nonstationary (thus their levels are I(2)) on the basis of the results using the standard ADF tests when there are some notable trend breaks in these variables.

In this paper, I re-examine the time-series behavior of the exchange rates and price indices of the four high-inflation countries: Brazil, Israel, Mexico, and Zaire, for which Bahmani-Oskooee (1993) and Mahdavi and Zhou (1994) fail to show the existence of a stationary APPP relation. I use the method of Zivot and Andrews (1992), which allows us not only to examine the stationarity for the variables with a structural break but also to test for a possible break point rather than assuming it exists. I then conduct the cointegration tests for the long-run validity of APPP in these high-inflation countries, based on the findings of the Zivot–Andrews tests, and incorporate the correction of the finite sample bias and the adjustment for trend breaks in the tests. As far as I know, no other studies have been conducted taking the trend breaks into account in the cointegration analysis of PPP for high-inflation countries.

The rest of the paper is organized as follows. The next section briefly introduces the sequential ADF tests of Zivot and Andrews (1992) and applies them to the variables I study. In section 3, I offer some explanations for the existence of trend breaks in these variables, on the basis of the experiences of the relevant countries. I then use Johansen's multivariate cointegration methodology to investigate whether the APPP condition holds in the four high-inflation countries and report the results in section 4. The last section presents conclusions of this study.

## 2. Identifying the Order of Integration for the Variables with Trend Breaks

A time series $x_t$ is said to be nonstationary and integrated of order $d$, denoted by $x_t \sim I(d)$, if it achieves stationarity after being differenced $d$ times. In testing the nonstationarity of time series, the augmented Dickey–Fuller (ADF) tests (Dickey and Fuller 1981; Said and Dickey 1984) have been widely used.[3] However, Perron (1989, 1990) shows that the exis-

[2] I employ the CPI data rather than the WPIs for the following reasons: (i) In so doing, we may offer a valid comparison of our results to those of Mark (1990), Kugler and Lenz (1993), and others who use the CPIs in their studies of PPP for industrial countries. (ii) It would be interesting to see if my results are more favorable to APPP than the findings of other studies on high-inflation countries, which often fail to support APPP when the CPI data are utilized. (iii) The WPI data are not available for Zaire.

[3] Because the ADF tests are well known, the descriptions of the tests are omitted here.

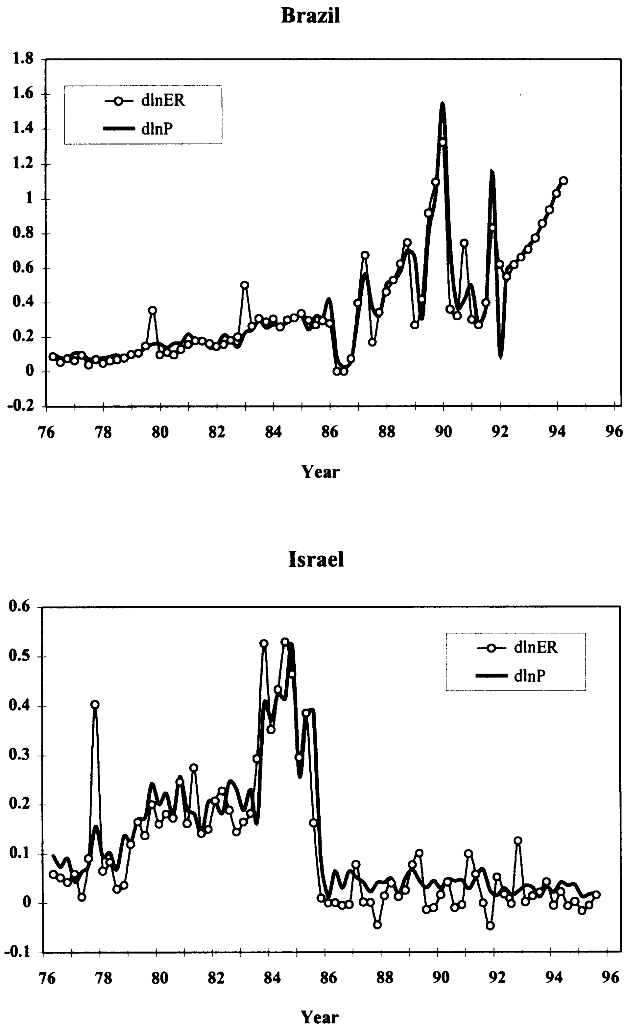
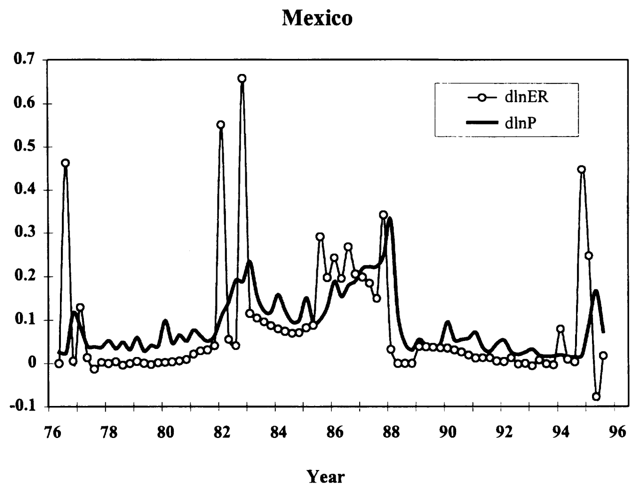
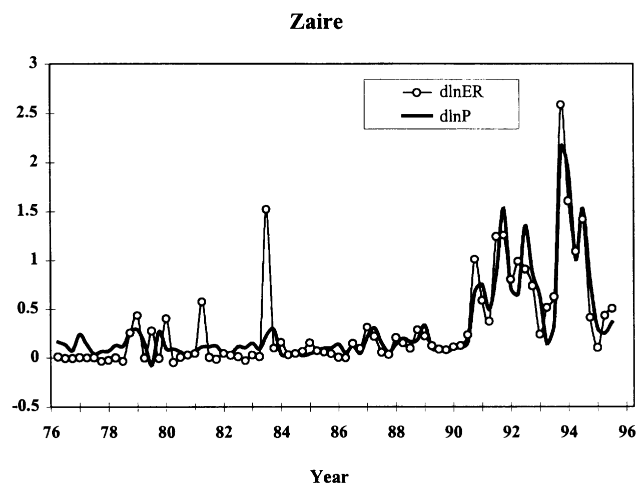



**Figure 1.** Changes of the Exchange Rates and Consumer Price Indices

tence of a structural change in the mean of a stationary time series biases the standard ADF tests toward nonrejection of nonstationarity. Since the variables plotted in Figure 1 appear to have such structural changes, it seems to be appropriate to employ the method of Zivot and Andrews (1992), called sequential ADF tests, to investigate the nonstationarity of the variables.

The sequential ADF tests developed by Zivot and Andrews (1992) may not only be used to identify the order of integration for the variables with a structural break, but also allow us to test a possible break point rather than assuming it exists. The null hypothesis of the tests is that the series $x_t$ is integrated, with the errors to be normal ARMA($p$, $q$) processes, without an exogenous structural break. The alternative hypothesis is that $x_t$ can be represented by a trend-stationary process with a trend break occurring at an unknown point in time. Basically, their tests are represented by the following augmented regression equations:



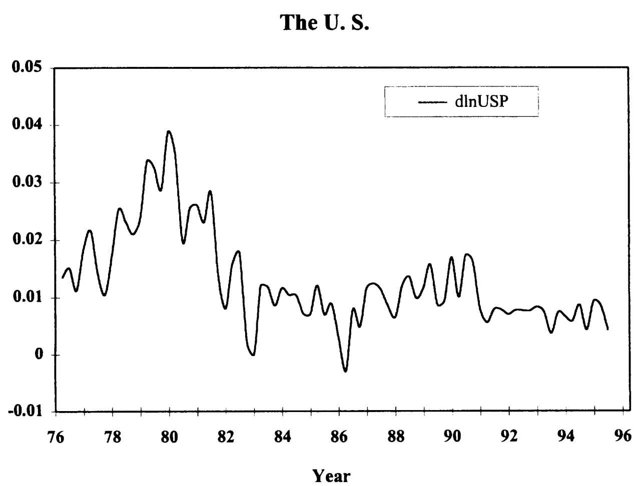




**Figure 1.** Continued

Model A: $$x_t = \mu + \beta t + \theta DU_t + \alpha\, x_{t-1} + \sum_{j=1}^{k} c_j \Delta x_{t-j} + e_t,$$

Model B: $$v_t = \alpha v_{t-1} + \sum_{j=1}^{k} c_j \Delta v_{t-j} + e_t,$$

Model C: $$x_t = \mu + \beta t + \theta DU_t + \gamma DT_t + \alpha x_{t-1} + \sum_{j=1}^{k} c_j \Delta x_{t-j} + e_t,$$

where the level dummy variable $DU_t = 1$ if $t > T_B$ (here $T_B$ denotes a possible break point) and zero otherwise. The slope dummy variable $DT_t = t - T_B$ if $t > T_B$ and zero otherwise. Models A and C are estimated by one-step regressions. The estimation of Model B has a two-step procedure, where the series $v_t$ is the residual series from a regression of $x_t$ on a constant, a time trend, and a slope dummy variable $DT_t$.[4] Note that if $DU_t$ is not included, Model A turns out to be the standard ADF test with a time trend, and if both $DU_t$ and the time trend are excluded, Model A becomes the standard ADF test without a time trend.

The sequential ADF procedure estimates a regression equation for every possible $T_B$ within the sample and calculates the $t$-statistics for the estimated coefficients. The null of nonstationarity is rejected in favor of the alternative of stationarity if $\alpha$ is significantly different from one. The chosen break point for each series is that $T_B$ for which the $t$-statistic for $H_0$: $\alpha = 1$ is minimized. The asymptotic critical values for these tests, which are greater (in absolute value) than those for the standard ADF tests, are tabulated in Zivot and Andrews (1992). Model specification (i.e., which of models A, B, or C is appropriate) is determined by first running each series on Model C, with the possibility of both a slope and a level break. Model C is chosen if both dummy variables are significant. If only the slope dummy variable is significant, Model B is estimated. If only the level dummy variable is significant, Model A is estimated.

For either standard ADF tests or sequential ADF tests, the choice of lag length $k$ may affect the test results. We follow the procedure suggested by Campbell and Perron (1991). Start with an upper bound, $k_{\max}$, 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 set $k_{\max} = 8$ for the quarterly data we use.

To illustrate the differences between standard ADF tests and sequential ADF tests, we first apply the standard ADF tests to the first (log) differences of the exchange rates and price variables, i.e., $x_t$ = dlnER or dlnP or dlnUSP, and report the results in Table 1. The results are similar to the findings of Mahdavi and Zhou (1994). The test statistics fail to reject the null hypothesis of nonstationarity. If we draw the conclusions based on these statistics, we may conclude that dlnER, dlnP, and dlnUSP are integrated variables, while lnER, lnP, and lnUSP are integrated of order greater than one. Due to the presence of trend breaks in the data, such conclusions could be false.

We then apply the tests of Zivot and Andrews to dlnER, dlnP, and dlnUSP. The results are presented in Table 2. The first four columns of Table 2 list the countries, the variables to be tested, the number of observations, and the chosen lag lengths. Note that Model A is selected for dlnUSP and for dlnER and dlnP of Brazil, Mexico, and Zaire; that is, in these cases, the slope dummy variable is found to be insignificant while the level dummy variable is significant. Model C is selected for the variables of Israel, i.e., both the level and the slope dummy variables

[4] Perron and Vogelsang (1991) give an explanation of why Model B is estimated differently from Models A and C.

**Table 1.** Standard ADF Tests

| Country | Variable | Lag $k$ | $t_\mu$ | Q(12) | $t_\tau$ | Q(12) |
|---|---|---|---|---|---|---|
| Brazil | dlnER | 3 | −0.75 | 5.85 | −3.14 | 5.21 |
| | dlnP | 3 | −0.83 | 7.93 | −3.15 | 8.54 |
| Israel | dlnER | 1 | −2.28 | 6.26 | −2.80 | 6.46 |
| | dlnP | 4 | −2.23 | 7.07 | −3.06 | 6.17 |
| Mexico | dlnER | 2 | −2.82 | 4.52 | −2.80 | 4.51 |
| | dlnP | 4 | −2.56 | 8.53 | −2.59 | 8.51 |
| Zaire | dlnER | 4 | −2.31 | 8.68 | −3.41 | 6.55 |
| | dlnP | 4 | −1.75 | 14.50 | −2.66 | 14.13 |
| U.S. | dlnUSP | 3 | −2.24 | 7.97 | −3.04 | 6.82 |
| | | | **5% Critical Values** | | | |
| Sample size $T$ = 50 | | | −2.93 | | −3.50 | |
| Sample size $T$ = 100 | | | −2.89 | | −3.45 | |

The sample period for Brazil is from the first quarter of 1976 to the second quarter of 1994. For all the other countries, the sample period runs from the first quarter of 1976 to the third quarter of 1995. $t_\mu$ is for the model with a constant term but no time trend. $t_\tau$ is for the model with a constant term and a time trend. The 5% critical values of $t_\mu$ and $t_\tau$ are taken from Fuller (1976). Q(12) is the Ljung-Box statistic for $H_0$: the residual series is white noise. The 5% critical values of Q(12) is 21.03.

are found to be significant. The Ljung–Box Q(12) statistics, reported in column 11, are unable to reject the null hypothesis that there is no serial correlation remaining in the residuals from the models.

The discovered dates of break points are illustrated in column 5 of Table 2. A trend break of dlnUSP that occurred at the third quarter of 1981 may reflect the switches of the Federal Reserve's monetary policy in the early 1980s. The discussion regarding the break points for the HIC variables will be given in the next section. In the 10th column of Table 2, the test statistics for the nonstationarity of variables are reported. The results show that, when we take the structural breaks into account in our testing, the null hypothesis of nonstationarity (i.e., $H_0$: $\alpha = 1$) is rejected at the 5% level (or less) of significance for all the first differences of the exchange rates and price indices using either the asymptotic critical values of Zivot and Andrews or the finite sample critical values listed in the last three columns of Table 2. These finite sample critical values are obtained through the Monte Carlo experiments under the assumption that the errors driving the series dlnER, dlnP, and dlnUSP are normal ARMA($p$, $q$) processes. Following Zivot and Andrews (1992), we determine $p$ and $q$ by fitting ARMA($p$, $q$) models to the first difference of each dlnER, dlnP, or dlnUSP (i.e., the second differences of lnER, lnP, or lnUSP) and using the model-selection criterion of Schwarz (1978) to choose the optimal ARMA($p$, $q$) with $p, q \leq 5$. The selected ARMA($p$, $q$) models are reported in Table 3. We then perform the Monte Carlo experiment described in Zivot and Andrews (1992, p. 262) to get the critical values for the finite-sample distributions based on 10,000 repetitions.[5] The results in Table 2 suggest that dlnER, dlnP, or dlnUSP are stationary with some trend breaks, and therefore the levels of the exchange rates and price variables of the four HICs and the U.S. are unlikely to be integrated of order two. This contrasts sharply with the results of Mahdavi and Zhou (1994) using the standard ADF tests.

[5] The results of the Monte Carlo experiments here for the sequential ADF tests and later for the cointegration tests are obtained through 10,000 replications using the GAUSS programming language.

**Table 2.** Sequential ADF Tests of Zivot and Andrews

| (1) Country | (2) Variable $x$ | (3) $T$ | (4) $k$ | (5) $\hat{T}_B$ | (6) $\hat{\mu}$ | (7) $\hat{\beta}$ | (8) $\hat{\theta}$ | (9) $\hat{\gamma}$ | (10) $\hat{\alpha}-1$ | (11) Q(12) | (12) Finite Sample Critical Values 1% | (13) 5% | (14) 10% |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | **Model A:** $x_t = \mu + \beta t + \theta DU_t + \alpha x_{t-1} + \Sigma_1^k c_j \Delta x_{t-j} + e_t$ | | | | | | | | |
| Brazil | dlnER | 74 | 0 | 87:4 | 0.006 | 0.005 | 0.155 | N.A. | −0.615 | 7.34 | −6.18 | −5.24 | −4.81 |
| | | | | | (0.110) | (2.273) | (2.152) | | (−5.411)* | | | | |
| | dlnP | 74 | 0 | 87:4 | 0.021 | 0.005 | 0.197 | N.A. | −0.706 | 9.03 | −6.37 | −5.54 | −5.15 |
| | | | | | (0.371) | (2.325) | (2.213) | | (−6.204)* | | | | |
| Mexico | dlnER | 79 | 0 | 87:4 | 0.012 | 0.004 | −0.219 | N.A. | −0.953 | 9.89 | −6.10 | −5.35 | −4.97 |
| | | | | | (0.347) | (2.998) | (−3.787) | | (−8.511)** | | | | |
| | dlnP | 79 | 4 | 88:1 | −0.003 | 0.002 | −0.112 | N.A. | −0.418 | 11.76 | −5.86 | −5.12 | −4.73 |
| | | | | | (−0.281) | (5.185) | (−6.196) | | (−5.791)* | | | | |
| Zaire | dlnER | 79 | 3 | 91:2 | 0.050 | 0.004 | 0.922 | N.A. | −1.360 | 10.80 | −5.85 | −5.04 | −4.72 |
| | | | | | (0.485) | (1.471) | (4.439) | | (−6.081)* | | | | |
| | dlnP | 79 | 3 | 91:2 | 0.069 | 0.003 | 0.875 | N.A. | −1.250 | 14.22 | −6.69 | −5.89 | −5.46 |
| | | | | | (0.912) | (1.429) | (4.891) | | (−6.298)* | | | | |
| U.S. | dlnUSP | 79 | 2 | 81:3 | 0.013 | 0.00001 | −0.009 | N.A. | −0.509 | 8.07 | −6.45 | −5.39 | −4.89 |
| | | | | | (4.964) | (0.267) | (−4.411) | | (−5.412)* | | | | |
| | | | | | **Model C:** $x_t = \mu + \beta t + \theta DU_t + \gamma DT_t + \alpha x_{t-1} + \Sigma_1^k c_j \Delta x_{t-j} + e_t$ | | | | | | | | |
| Israel | dlnER | 79 | 0 | 85:2 | 0.002 | 0.007 | −0.267 | −0.007 | −0.731 | 5.38 | −6.23 | −5.94 | −5.52 |
| | | | | | (0.061) | (5.000) | (−6.179) | (−4.831) | (−7.430)** | | | | |
| | dlnP | 79 | 3 | 85:3 | 0.010 | 0.008 | −0.282 | −0.008 | −0.833 | 12.67 | −6.67 | −5.80 | −5.39 |
| | | | | | (0.513) | (6.228) | (−8.009) | (−5.067) | (−7.841)** | | | | |

$T$ is the number of sample observations. $\hat{T}_B$ is the date of revealed break point. Numbers in parentheses are the $t$-statistics. The finite sample critical values are obtained through the Monte Carlo experiment under the assumption of normal ARMA($p$, $q$) errors. The selected ARMA($p$, $q$) models are reported in Table 3. The asymptotic critical values of Zivot and Andrews (1992) for Model A are −5.34, −4.80, and −4.58, corresponding to the 1%, 5%, and 10% significance levels, respectively. Those for Model C are −5.57, −5.08, and −4.82, respectively. N.A. = Not applicable. Also see notes to Table 1.

* Statistically significant at the 5% level for test of $\alpha = 1$ using the finite sample critical values.

** Statistically significant at the 1% level for test of $\alpha = 1$ using the finite sample critical values.

**Table 3.** Selected ARMA($p$, $q$) Models

| Country | Variable $x$ | Model | $\hat{\theta}_1$ | $\hat{\theta}_2$ | $\hat{\theta}_3$ | $\hat{\sigma}$ | Q(12) |
|---|---|---|---|---|---|---|---|
| | | | **Model:** $\Delta x_t = \mu + e_t + \Sigma_1^q \theta_i e_{t-i}$ | | | | |
| Brazil | dlnP | (0, 1) | −0.62 (−7.18) | N.A. | N.A. | 0.20 | 11.02 |
| Israel | dlnER | (0, 1) | −0.32 (−3.11) | N.A. | N.A. | 0.09 | 7.16 |
| | dlnP | (0, 1) | −0.30 (−2.81) | N.A. | N.A. | 0.06 | 14.60 |
| Mexico | dlnER | (0, 1) | −0.58 (−6.16) | N.A. | N.A. | 0.14 | 9.25 |
| U.S. | dlnUSP | (0, 3) | −0.23 (−2.21) | −0.42 (−4.23) | 0.31 (3.15) | 0.01 | 4.98 |

| Country | Variable $x$ | Model | $\hat{\phi}_1$ | $\hat{\phi}_2$ | $\hat{\phi}_3$ | $\hat{\phi}_4$ | $\hat{\sigma}$ | Q(12) |
|---|---|---|---|---|---|---|---|---|
| | | | **Model:** $\Delta x_t = \mu + \Sigma_1^p \phi_i \Delta x_{t-i} + e_t$ | | | | | |
| Brazil | dlnER | (2, 0) | −0.22 (−1.97) | −0.36 (−3.23) | N.A. | N.A. | 0.19 | 8.75 |
| Mexico | dlnP | (4, 0) | −0.16 (−1.25) | −0.18 (−1.32) | −0.20 (−1.56) | 0.28 (2.12) | 0.04 | 12.91 |
| Zaire | dlnER | (2, 0) | −0.53 (−3.76) | −0.28 (−2.09) | N.A. | N.A. | 0.41 | 11.01 |
| | dlnP | (4, 0) | −0.09 (−0.79) | −0.51 (−4.73) | 0.05 (0.45) | −0.31 (−2.58) | 0.30 | 15.43 |

All models were estimated using BOXJENK in RATS 3.1. Numbers in parentheses are the $t$-statistics. Also see notes to Table 1.

## 3. Trend Breaks and the Experiences of the Relevant Countries

During the sample period, all four high-inflation countries in our study have experienced some dramatic changes in their economies, especially in their inflation rates. The average annual rates of consumer price inflation of these countries have a range of 50 to 1600% over the sample period, and the record of their highest annual inflation ranges from 130 to 22,000% within the sample period. The governments of these countries made various efforts to try to stabilize their economies.

Brazil has had an inflation problem for decades, but the inflation worsened sharply following the debt crisis between 1981 and 1985. The Brazilian government announced its Cruzado Plan at the end of February 1986, when the inflation rate was threatening to rise above 20% a month. The Plan centered around a general freeze of prices and the exchange rate. However, by late 1986, the Plan became unsustainable. Price controls were abandoned in December 1986, and inflation reached a monthly rate of 26% in May and June of 1987. The acceleration of inflation led the government to adopt a new stabilization plan, known as the Bresser Plan, as an attempt to correct some of the errors of the Cruzado Plan. This program, too, was a failure: Inflation rose rapidly in the last quarter of 1987, and by 1988, the annual inflation rate approached nearly hyperinflation, at about 700%, accompanied by some sharp depreciations of the Brazilian cruzeiro. The date of the break point that we found for Brazil, the fourth quarter of 1987, corresponds to this period. The situation turned even worse in the late 1980s and the 1990s.

Like Brazil, Israel had long experienced acute, chronic inflation. A financial crisis in 1983 stimulated inflationary expectations and sharply increased the domestic debt of the public sector, resulting in a rapid increase in inflation and currency depreciation. On July 1, 1985, the Israeli government enacted an ambitious heterodox stabilization program, including a program of public expenditure cuts, mainly in defense, and a price freeze covering mostly the prices of consumer goods. The monetary authority also tightened the money supply and exchange-rate control. The U.S. provided an extra $1.5 billion in aid in 1985 and 1986 on top of its regular annual handout of around $3 billion. This took the pressure off Israel's balance of payments and made it easier for Israel to control its exchange rate. The exchange rate played a central role in Israel's stabilization strategy. Given the economy's relatively small size and open character, the authorities used the exchange rate quite deliberately to anchor the price level. The Israeli reform appears to have been a remarkable success. The nation's international reserve position improved and inflation has come down since 1985. This is consistent with the break points that we discovered for Israel at the second quarter of 1985 for the exchange rate and the third quarter of 1985 for the price index.

From the mid-1950s to the 1970s, Mexico was a paragon of financial stability and economic growth. The stability ended with the oil price increase of the 1970s because of fiscal extravagance resulting from the dramatic increase in revenue from oil exports. With added room in the budget, policies became highly expansionary, the currency became overvalued, and government borrowing increased. After a huge increase in government debt followed by the collapse of the financial system and the exchange rate, Mexico became insolvent in 1982, the first of the debtor countries to do so. Since 1987, when consumer price inflation reached nearly 160%, the top priority for the Mexican government had been reducing inflation. In December 1987, the Economic Solidarity Pact was announced. The Pact combined strict fiscal and monetary policies, wage and price controls, and an exchange-rate policy that allowed the peso to significantly appreciate in real terms against the dollar. The Pact and its successor, the Pact for Stability and Economic Growth, had succeeded in reducing inflation. Consumer price inflation was brought down to 10% in the second quarter of 1988, following a real appreciation of the peso. Correspondingly, the dates of break points for the Mexican exchange rate and price index are found to be the fourth quarter of 1987 and the first quarter of 1988, respectively.

The IMF and the World Bank supported Zaire with at least five stabilization programs between 1976 and 1987, but the custom for Zaire quickly became to make the first drawing of the loans from the IMF or the World Bank and then to drift away from the economic reform performance criteria. In January 1989, the government once more took steps to establish economic stability. A structural adjustment program, including the commitment to contain budget deficits, narrowed the gap between the official and black-market exchange rate to 10% by April 1989. Despite the country's reform efforts, the pace of economic activity had not accelerated sufficiently in 1989 to boost living standards, which had fallen each year for more than a decade. Inefficient and corruptly managed parastatals (semiautonomous, quasi-governmental, state-owned enterprises) had contributed to Zaire's troubled economic history and were a severe strain on the budget. By the end of 1989, it was apparent that the latest reforms were unsuccessful in promoting sustained economic expansion. Instead, Zairians experienced a massive drop in per capita income as inflation rose and the GDP growth rate fell. Economic indicators for 1990 were even more dismal. IMF credits had expired, and large public-expenditure deficits were expected in order to fund pay increases for government workers. The figures for 1991, in which a break point was disclosed at the second quarter, showed a GDP decline of 2.6%, a 2000%

rate of consumer price inflation, and further devaluation of the zaire against Western currencies. By 1992 and continuing into 1993, poverty and unemployment were widespread and hyperinflation was a permanent fixture.

From the above discussions, we see that the dates of the break points in the data exposed by the sequential ADF tests of Zivot and Andrews are generally consistent with the true experiences of these countries. This illustrates the strong power of the tests. Hence, further effort to detect the trend breaks in the variables seems unnecessary.

## 4. Testing for the Volatility of Long-Run PPP in High-Inflation Countries

PPP between the exchange rate and relative prices, with the U.S. as the base country, could be expressed as the following empirical relationship:

$$\beta_{er}\text{lnER}_t + \beta_p\text{lnP}_t + \beta_{usp}\text{lnUSP}_t = \xi_t,$$

where lnER, lnP, and lnUSP are the logs of the exchange rate, domestic price level, and the U.S. price level, respectively (as defined earlier), and $\xi$ is an error term reflecting deviations from PPP. When the exchange rate and price levels are integrated variables, for PPP to hold in the long run, $\xi_t$ should be stationary. Strictly speaking, the PPP condition requires $-\beta_{er} = \beta_p = -\beta_{usp}$ or $[\beta_{er}, \beta_p, \beta_{usp}] = [-1, 1, -1]$ when $\beta_{er}$ is normalized to be $-1$, i.e., the long-run proportionality between exchange rates and price levels. However, as argued by Edison (1987) and Taylor (1988), transportation costs and measurement error, as well as the existence of nontradable elements in measured price indices, may imply that $\beta_{er}$, $\beta_p$, and $\beta_{usp}$ are not equal to unity. Therefore, we refer to the situation where $\xi_t$ is stationary and the condition $[\beta_{er}, \beta_p, \beta_{usp}] = [-1, 1, -1]$ is satisfied as strong-form PPP. Weak-form PPP requires a stationary $\xi_t$, but the proportionality condition does not necessarily hold.

Cointegration analysis offers a natural way to test the existence of a long-run relationship among integrated variables. A set of I($d$) variables, $\mathbf{X}_t$, is said to be cointegrated if a linear combination of them, $Z_t = \boldsymbol{\beta}'\mathbf{X}_t$, is integrated of any order less than $d$. The vector $\boldsymbol{\beta}$ is referred to as the cointegrating vector. If $Z_t$ is found to be integrated of order zero (i.e., stationary), we may say that the variables in $\mathbf{X}_t$ do not drift too far apart and there exists a long-run equilibrium relationship among these variables.

Johansen's multivariate cointegration tests (Johansen 1988, 1992; Johansen and Juselius 1990) are utilized in this paper to test the long-run PPP relation between four high-inflation countries and the U.S. Since in section 2 the first differences of $\text{lnER}_t$, $\text{lnP}_t$, and $\text{lnUSP}_t$ are shown to be stationary with some trend breaks, we may directly apply the cointegration tests to the levels of the variables rather than to their first differences. The tests are conducted through a vector error-correction mechanism with the null hypothesis of no cointegration. Defining a vector $\mathbf{X}_t$ containing three variables, $\mathbf{X}_t' = [\text{lnER}_t, \text{lnP}_t, \text{lnUSP}_t]$, a vector error-correction model can be written as

$$\Delta\mathbf{X}_t = \sum_{j=1}^{k-1} \boldsymbol{\Gamma}_i\Delta\mathbf{X}_{t-j} + \boldsymbol{\Pi}\mathbf{X}_{t-1} + \boldsymbol{\mu}_0^* + \boldsymbol{\mu}_1^*t + \boldsymbol{\eta}\mathbf{D}_t + \boldsymbol{\epsilon}_t, \tag{4.1}$$

where $\boldsymbol{\epsilon}_t$ is a vector of independent Gaussian variables with mean zero and variance matrix $\boldsymbol{\Sigma}$. $\boldsymbol{\mu}_0^*$ is a constant term that implies that the process $\mathbf{X}_t$ has a linear trend. The model including $\boldsymbol{\mu}_1^*t$ allows for the possibility of quadratic trends in $\mathbf{X}_t$ (see Osterwald-Lenum, 1992, p. 471).

The vector $\mathbf{D}_t$ usually contains deterministic variables such as some dummies or other variables outside the cointegration space. They are included in the model to ensure that the disturbances $\boldsymbol{\epsilon}_t$ "are as close to being Gaussian as possible" (Pesaran and Pesaran, 1991, p. 85). The hypotheses of interest involve $\mathbf{\Pi}$; if the rank of $\mathbf{\Pi}$ is $r$, where $r \leq 2$, then $\mathbf{\Pi}$ can be decomposed into two $3 \times r$ matrices $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ such that $\mathbf{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$. The matrix $\boldsymbol{\beta}$ consists of $r$ linear, cointegrating vectors, while $\boldsymbol{\alpha}$ can be interpreted as a matrix of vector error-correction parameters. The Johansen method offers the test for the number of cointegrating vectors and gives consistent maximum likelihood estimates of the entire cointegrating matrix. For each country pair, the number of cointegrating vectors is determined using the maximum eigenvalue test statistic or $\lambda_{max}$.[6] If the results are consistent with the hypothesis of at least one cointegrating vector, the hypotheses regarding the relationship of the cointegrating coefficients $\boldsymbol{\beta}'$ could be tested using the maximum likelihood methodology.

Asymptotic critical values for the maximum eigenvalue test are tabulated in Osterwald-Lenum (1992). However, when the sample length is not long, the tests based on asymptotic critical values may be biased toward finding cointegration too often. Cheung and Lai (1993c) examine the finite sample properties of the Johansen tests and provide finite sample critical values for the case assuming there is no linear trend nor quadratic trend in the data generating processes (DGPs) but allowing the constant term $\boldsymbol{\mu}_0^*$ in the models of estimation.

Since in Table 2 we found a significant time trend in the first differences of the logged exchange rates and price indices of Brazil, Israel, and Mexico, we assume that the DGPs of lnER and lnP of these three countries have a quadratic trend, i.e.,

$$\mathbf{X}_t = \mathbf{X}_{t-1} + \boldsymbol{\mu}_0 + \boldsymbol{\mu}_1 t + \boldsymbol{\epsilon}_t, \tag{4.2}$$

which is equivalent to $\Delta\mathbf{X}_t = \boldsymbol{\mu}_0 + \boldsymbol{\mu}_1 t + \boldsymbol{\epsilon}_t$, with $\boldsymbol{\epsilon}_t$ being defined earlier for Equation 4.1, and include both a constant term and a time trend in the models for testing PPP. For Zaire, the coefficients of the time trend are found to be insignificant in its dlnER and dlnP; therefore, the DGPs of lnER and lnP of Zaire are assumed to have a linear trend, i.e.,

$$\mathbf{X}_t = \mathbf{X}_{t-1} + \boldsymbol{\mu}_0 + \boldsymbol{\epsilon}_t, \tag{4.3}$$

which is equivalent to $\Delta\mathbf{X}_t = \boldsymbol{\mu}_0 + \boldsymbol{\epsilon}_t$, and only a constant term is included in the models for testing the PPP relation between Zaire and the U.S. Based on the above assumptions, we use the methods suggested by Cheung and Lai (1993c) to perform the Monte Carlo experiment and produce the finite sample critical values, reported in Table 4, for the Johansen tests.

## *Testing for Cointegration*

The results of the cointegration tests without accounting for the effects of trend breaks (i.e, including no $\mathbf{D}_t$ in Eqn. 4.1) are presented in Table 4. The reported lag lengths $k$ are chosen on the basis of the Bayesian information criterion (BIC) (Schwarz 1978) as well as the criterion that the residuals from the model are not serially correlated; that is, among the models where the Ljung-Box Q(12) statistics are found to be smaller than their 5% critical values, $k$ is chosen corresponding to the one with the lowest BIC statistic. We set $k_{max} = 8$.

For the models expressed by Equation 4.1 and the DGPs assumed to be Equation 4.2, the

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[6] Johansen (1988) also proposes another likelihood ratio test known as the trace test for determining the number of cointegrating relationships. Of these two tests, however, the maximum eigenvalue test is expected to provide more clear cut results than the trace test (Johansen and Juselius 1990).

**Table 4.** Cointegration Tests

| Country Pair: | Brazil/U.S. | Israel/U.S. | Mexico/U.S. | Zaire/U.S. |
|---|---|---|---|---|
| $k$: | 4 | 6 | 5 | 6 |
| $\lambda_{max}$ ($H_0$: $r = 0$) | 16.47 | 35.50* | 16.10 | 19.95 |
| $Q(12)_{er}$ | 7.51 | 9.67 | 6.07 | 4.28 |
| $Q(12)_p$ | 10.35 | 8.29 | 17.93 | 6.04 |
| $Q(12)_{usp}$ | 6.40 | 6.86 | 8.69 | 6.33 |
| **Finite Sample Critical Values of $\lambda_{max}$** | | | | |
| 5% (1) | 28.10 | 31.13 | 29.64 | 26.79 |
| (2) | 29.09 | 31.37 | 30.00 | 27.55 |
| 10% (1) | 25.55 | 28.00 | 26.56 | 23.87 |
| (2) | 26.22 | 28.32 | 27.64 | 24.57 |
| **Estimated Cointegrating Vector for Israel/U.S.** | | | | |
| | $\beta_{er}$ | $\beta_p$ | | $\beta_{usp}$ |
| $\beta$ | −1.00 | 0.99 | | −0.14 |
| $H_0$: $[\beta_{er}, \beta_p, \beta_{usp}] = [-1, 1, -1]$ | | | $\chi^2(2) = 7.64$* | |
| $H_0$: $-\beta_{er} = \beta_p$ | | | $\chi^2(1) = 0.84$ | |

$r$ is the hypothesized number of cointegrating vectors. The reasons of having two sets of finite-sample critical values of $\lambda_{max}$ are given in the text (see section 4). The 5% critical values of Q(12) is 21.03 for $H_0$: the residual series is white noise. The subscript of Q(12) denotes the left-hand-side variable in the equation where the residual series is from. $\beta$'s are the parameters of the cointegrating vectors normalized on lnER. The 5% critical values of $\chi^2(1)$ and $\chi^2(2)$ are 3.84 and 5.99, respectively.
* Significance at the 5% level.

finite-sample distribution of the $\lambda_{max}$ test statistic is invariant with respect to the value of the drift $\boldsymbol{\mu}_0$, but is variant with the value of $\boldsymbol{\mu}_1$ in Equation 4.2. If there is no time trend in the model (i.e., $\boldsymbol{\mu}_1^* = 0$ in Eqn. 4.1) and the assumed DGPs are expressed by Equation 4.3, the distribution of the $\lambda_{max}$ test statistic would be variant with the value of $\boldsymbol{\mu}_0$ in Equation 4.3. Therefore, there are two sets of finite sample critical values listed in Table 4 for each country pair. For Brazil/U.S., Israel/U.S., and Mexico/U.S., the first set of critical values is obtained assuming $\boldsymbol{\mu}_1 = \mathbf{1}$, and the second set of critical values is produced by setting $\boldsymbol{\mu}_1$ in the DGPs equal to the coefficients of the time trend for dlnER, dlnP, and dlnUSP, displayed in Table 2. In the case of Zaire/U.S., the first and second set of critical values are corresponding to $\boldsymbol{\mu}_0 = \mathbf{1}$ and $\boldsymbol{\mu}_0$ equal to the coefficients of the constant term for dlnER, dlnP, and dlnUSP in Table 2, respectively. Although the assumed $\boldsymbol{\mu}_0$ and $\boldsymbol{\mu}_1$ might not be good proxies for the true $\boldsymbol{\mu}_0$ and $\boldsymbol{\mu}_1$, the two sets of the critical values indicate that, when $\boldsymbol{\mu}_0$ or $\boldsymbol{\mu}_1$ vary from one to the values close to zero, the critical values are not much affected.

Comparing the $\lambda_{max}$ test statistics for the null of no cointegration ($H_0$: $r = 0$) with the finite sample critical values, the null hypothesis could be rejected only for Israel/U.S. There is no supportive evidence for PPP for the other three country pairs. For the Israel/U.S. pair, where one cointegrating vector is found, we conduct some tests for the hypotheses on the cointegrating vector applying the procedures of Johansen and Juselius (1990, 1992). The tests are basically the likelihood ratio tests constructed from the estimated eigenvalues corresponding to the restricted and unrestricted models.

For the hypothesis of proportionality, discussed earlier in section 4, the likelihood ratio test statistics reject the null of $[\beta_{er}, \beta_p, \beta_{usp}] = [-1, 1, -1]$ but fail to reject the null of $-\beta_{er} = \beta_p$ (i.e., $\beta_{er} + \beta_p = 0$). The results suggest the existence of a proportional relationship in the long run between the exchange rate and the price level of Israel.

### *Testing for Cointegration with Adjustment for Trend Breaks*

Without taking the effects of trend breaks into consideration, the cointegration tests described in the above subsection provide little support for PPP in high-inflation countries. Cheung et al. (1995) argue that the existence of trend breaks in the stochastic processes may bias the cointegration test toward finding no cointegration too often if one fails to model the breaks when performing the test. When one includes some dummy variables in the model to capture the effects of trend breaks, the distributions of the test statistics would be different from those for the variables with no trend breaks.

Following Cheung et al. (1995), we set the null hypothesis of the tests to be no cointegration, while the alternative is that the variables are cointegrated with some trend breaks in the data. Hence, we allow $\mathbf{D}_t$ in Equation 4.1 to include the dummy variables, the slope and/or the level dummies, representing the structural breaks revealed in section 2.[7] They are: (i) $DU_t(81{:}3)$ and $DU_t(87{:}4)$ for the Brazil/U.S. pair; (ii) $DU_t(81{:}3)$, $DU_t(85{:}2)$, $DT_t(85{:}2)$, and $DU_t(85{:}3)$ for Israel/U.S.; (iii) $DU_t(81{:}3)$, $DU_t(87{:}4)$, and $DU_t(88{:}1)$ for Mexico/U.S.; and (iv) $DU_t(81{:}3)$ and $DU_t(91{:}2)$ for Zaire/U.S.[8] Correspondingly, the finite sample critical values for the Johansen cointegration tests in the presence of trend breaks are simulated through the Monte Carlo experiment.[9] Again, there are two sets of finite-sample critical values displayed in Table 5 for the reasons stated in the above subsection.

The results of cointegration tests with adjustment for trend breaks are summarized in Table 5. They provide more positive evidence for cointegration than those presented in Table 4. In all four cases, the hypothesis of no cointegration can be rejected at the 5% level of significance, indicating that weak-form PPP holds in these high-inflation countries.

When testing the hypothesis of proportionality, i.e., $H_0$: $[\beta_{er}, \beta_p, \beta_{usp}] = [-1, 1, -1]$, the likelihood ratio test statistics fail to reject this hypothesis for Israel/U.S. and Mexico/U.S. but reject it for Brazil/U.S. and Zaire/U.S. Furthermore, the test results fail to reject the null of $-\beta_{er} = \beta_p$ (i.e., $\beta_{er} + \beta_p = 0$) for three out of four cases. The above findings generally support weak-form PPP in high-inflation countries for all the countries in the study. There is also some evidence in favor of strong-form PPP for two out of four cases.[10]

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[7] It is worthwhile to point out that we do not add additional dummies like (0 0 . . . 0 1 0 . . . 0) into the model to capture some one-period jumps in the data of the exchange rates. As can be seen in Figure 1, these jumps often reflect the catch-ups of the exchange rates under the pressure of persistent inflation. They should be considered to be consistent with PPP. If we use some one-period dummies to capture these jumps in the exchange rates, it may distort the true PPP relation. Readers may argue that the trend breaks in the data may also reflect the PPP relation. Note that the trend breaks we discovered appear in both exchange rates and price variables. If we appropriately model these breaks, it would not cause distortion of the PPP relation. On the other hand, if we neglect these trend breaks as McNown and Wallace (1989), Bahmani-Oskooee (1993), and Mahdavi and Zhou (1994) did, the standard ADF tests would falsely evidence that the variables of high-inflation countries are I(2), and the cointegration tests would be unable to support the absolute version of PPP when it may actually hold.

[8] Note that we are unable to add both $DT_t(85{:}2)$ and $DT_t(85{:}3)$ dummies into the model for the Israel/U.S. pair. If we do so, we would receive an error message for the problem of serious multicollinearity. For this reason, only one of the two slope dummies is included.

[9] The author is grateful to Yin-Wong Cheung for his valuable suggestions and guidance in conducting these Monte Carlo experiments.

[10] The results obtained using break dummies have to be interpreted with caution. As mentioned in Cheung et al. (1995, p. 185), while the existence of trend breaks may bias empirical tests, "using dummy variables to capture them creates another problem." These breaks may not be fully exogenous, and the dummy variables for these breaks may capture nonstationarity in the data. In addition, the presence of trend breaks may affect the distributions of the test statistics for the hypothesis of proportionality. This second problem would not affect the conclusions regarding weak-form PPP, but may have effects on the tests for strong-form PPP. We will address this problem in a succeeding study.

**Table 5.** Cointegration Tests with Adjustments for Trend Breaks

| Country pair: *k*: | Brazil/U.S. 1 | Israel/U.S. 7 | Mexico/U.S. 4 | Zaire/U.S. 6 |
|---|---|---|---|---|
| $\lambda_{max}$ ($H_0$: $r = 0$) | 37.46* | 44.73* | 34.60* | 33.87* |
| $Q(12)_{er}$ | 15.02 | 16.03 | 10.11 | 15.12 |
| $Q(12)_{p}$ | 13.59 | 14.54 | 17.85 | 15.39 |
| $Q(12)_{usp}$ | 0.66 | 16.48 | 13.57 | 8.14 |
| **Finite Sample Critical Values of $\lambda_{max}$** | | | | |
| 5% (1) | 29.53 | 43.41 | 33.84 | 32.55 |
| (2) | 31.71 | 44.75 | 34.57 | 33.81 |
| 10% (1) | 26.80 | 39.63 | 30.62 | 29.31 |
| (2) | 28.94 | 40.66 | 31.32 | 30.43 |
| **Estimated Cointegrating Vectors** | | | | |
| $\beta_{er}$ | −1.00 | −1.00 | −1.00 | −1.00 |
| $\beta_{p}$ | 0.92 | 1.03 | 1.00 | 1.04 |
| $\beta_{usp}$ | −0.27 | −0.37 | −2.11 | 1.69 |
| **$\chi^2(2)$** | | | | |
| $H_0$: $[\beta_{er}, \beta_{p}, \beta_{usp}] = [-1, 1, -1]$ | 16.10* | 1.13 | 4.07 | 26.88* |
| **$\chi^2(1)$** | | | | |
| $-\beta_{er} = \beta_{p}$ | 14.90* | 0.25 | 0.001 | 1.14 |

See notes to Table 4.

## 5. Summary and Conclusions

With the development of cointegration techniques, the long-run validity of purchasing power parity has received more favorable evidence for industrial countries. However, although the PPP theory suggests that PPP is likely to hold well in high-inflation countries, the absolute version of PPP has not obtained much support from the existing studies of the less-developed high-inflation countries for the recent flexible exchange-rate period. The results are particularly unfavorable to PPP for these countries when the consumer price indices are utilized. One possibility is that the existing studies have not modeled the variables of HICs appropriately, as some of them claim that the first differences of the exchange rates and price variables of HICs are nonstationary and thus their levels are integrated of order two.

In this study, I use the method of Zivot and Andrews (1992) to re-examine the time series behavior of the exchange rates and consumer price indices of four high-inflation countries. I uncover that the first differences of these variables are stationary with some trend breaks. With the aid of the Zivot–Andrews tests, I figure out the dates of break points and find that they are consistent with the substantial changes that occurred in these economies. I then employ these data to conduct the cointegration analysis for the long-run validity of APPP in high-inflation countries and include the revealed structural breaks in our analysis.

The results from the Johansen cointegration tests with adjustment for trend breaks show the existence of a cointegrating relationship among the bilateral exchange rate, the U.S. price level, and the high-inflation country price variable for all four country pairs in this study. In addition, there is some evidence in favor of the hypothesis of proportionality (i.e., strong-form PPP) for a couple of cases. This study provides support for the argument that PPP holds well, at least in a weak form, in high-inflation countries where the general price level movement overshadows the factors causing deviations from PPP.

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