Lat Am Econ Rev (2017) 26:5
https://doi.org/10.1007/s40503-017-0042-9

Purchasing power parity in Mexico since 1933

Frederick H. Wallace[1]

Received: 2 March 2016 / Revised: 12 April 2017 / Accepted: 20 May 2017 /
Published online: 2 June 2017

Abstract A new approach to cointegration developed by Enders et al. (Cointegration tests using instrumental variables with an example of the U.K. demand for money. Unpublished working paper. http://wenders.people.ua.edu/time-series-methods.html, 2008) is applied to long-span, high-frequency data to test for purchasing power parity in the Mexico–US real exchange rate. Overall the empirical results suggest that purchasing power parity (PPP) holds for the study period. The evidence for PPP is stronger when structural breaks are allowed in the real exchange rate.

Keywords Purchasing power parity · Real exchange rate · Cointegration

JEL Classification C22 · F30

1 Introduction

Absolute purchasing power parity (PPP) between two countries holds if the cost of a market basket of tradable goods and services is the same in both countries in terms of a common currency. If there are no barriers to trade or transactions costs, any differences in the cost of the market basket are eliminated through arbitrage. Interestingly, Cassel (1918) advanced the PPP hypothesis as the norm under the

✉ Frederick H. Wallace
wallace.f@gust.edu.kw

[1] Department of Economics and Finance, Gulf University for Science and Technology, West Mishref, Kuwait

aforementioned assumptions then sought to explain the apparent deviation from PPP in Sweden and its trading partners as arising from trade barriers.

Empirical testing of PPP presents some difficulties, but surmountable ones. First, there are no data on the costs of a market basket of tradable goods and services. Instead researchers use price indexes, but indexes contain data on non-traded goods and services that are irrelevant in testing for PPP. Second, price indexes and nominal exchange rates are constructed by the averaging of observations drawn at different times and/or different locations. Such averaging can introduce bias into unit root tests for PPP (Taylor 2001). Third, evidence of mean reversion of the real exchange rate is regarded as evidence of purchasing power parity, but there is no guarantee that the mean to which a real exchange rate measure constructed using price indexes reverts is the PPP level. Of course there is no theoretical basis for mean reversion to a real rate different from the one associated with PPP either.

Purchasing power parity in the real exchange rate between Mexico and the United States is examined in this study. Focus on this bilateral relation is justified because most of Mexico's international trade has historically been with the US. This paper extends previous work on PPP for the Mexico–US relationship in two important ways. First, it makes use of the long-span, high-frequency data set developed by Ventosa-Santàularia et al. (2015), henceforth VGW, and updated to 2014. Wallace et al. (2011) used the 1930–1960 portion of this data set to test for PPP and Gómez-Zaldivar et al. (2013) studied PPP in Mexico with the data for 1969–2010, but this paper is the first after VGW to utilize the complete data set. Second, the study employs a recent approach to cointegration testing developed by Enders et al. (2008) to examine the PPP hypothesis. The Enders, Im, and Lee (henceforth EIL) methodology uses stationary instrumental variables for the integrated series of the empirical model to overcome the difficulties of inference posed by non-standard distributions of nonstationary variables. EIL show that the t statistics of the estimated model are asymptotically normal when nonstationary regressors are instrumented with stationary ones.

Taylor (2001) shows that bias against the PPP hypothesis can arise in tests of stationarity of the real exchange rate due to temporal aggregation of high-frequency data in the construction of the lower frequency price indexes. He demonstrates that the bias reduces the power of unit root tests, but finds that bias diminishes with longer data spans. In light of Taylor's work most studies of purchasing power parity have focused on industrial countries likely due to the availability of data over extended spans. The long-span data set used in this study, thus, mitigates the bias arising from temporal aggregation allowing tests of PPP for Mexico, a nation whose economic development has lagged that of the industrial countries.

The long-span high-frequency data set put together by VGW and updated for this study is a composite from several different sources. Monthly, historical nominal exchange rate and price index data for Mexico, 1930–1960, are obtained from Cárdenas (1994) who cites the Bank of Mexico as a source, although the central bank does not provide these historical figures through its statistical website. Post-1960 data are from annual reports of the Bank of Mexico, commonly referred to as Banxico, as well as the Banxico website.

2 Literature review

Due to the large number of PPP studies, this review focuses only on PPP research for Mexico. In studying real exchange rate variability and GDP in Mexico Ávalos Huerta and Hernández Trillo (1995) fail to find evidence of PPP in annual or quarterly data for the Mexico–US real exchange rate for 1961–1994. McLeod and Welch (1992) apply unit root and cointegration tests to quarterly data for the Mexico–US real exchange rate, 1960–1991. They find evidence of stationarity of the real rate and cointegration of the nominal exchange rate and price levels, results normally regarded as evidence of PPP between two countries. Despite their finding McLeod and Welch express skepticism regarding the presence of PPP. Santaella and Vela (2005) provide an excellent description and analysis of nominal exchange rate policy in Mexico during the tumultuous period of 1987–1994. Mejía Reyes and González Núñez (1996) test for absolute purchasing power parity in three measures of the Mexico–US real exchange rate. Two of the real exchange rate measures are constructed using annual data from 1940 to 1994 and the third for 1921–1994. The three measures differ in the price indexes used to calculate the real exchange rate. Using augmented Dickey–Fuller and Phillips–Perron unit root tests Mejía Reyes and González Núñez find weak evidence of stationarity of the real rate, thus support for the PPP hypothesis, in the three series. They also apply the Engle–Granger tests to the three series and conclude that the Mexican and US price levels are cointegrated for both of the time periods studied although the evidence is weaker for 1921–1994. Evidence from Johansen's cointegration test provides weak evidence of PPP only for the 1921–1994 period.

Hegwood and Papell (1998) argue that PPP requires reversion of the real exchange rate to a constant mean. Unit root tests showing stationarity of the real exchange rate in the presence of breaks are not evidence of PPP if shifts in the series imply reversion to a changing mean. They define such instances of reversion to an occasionally changing mean as quasi-purchasing power parity or QPPP. Noriega and Medina (2003) test for PPP in annual data for the peso–dollar real exchange rate for 1925–1994 allowing for an unknown number of structural breaks in the data. Noriega and Medina find that a structural break in the mean real exchange rate occurred in 1981; but the rate is stationary both before and after this date, evidence of QPPP.

Taylor (2002) tests for purchasing power parity in a data set composed of more than 100 annual observations for 20 countries, including Mexico. His tests indicate mean reversion, evidence of PPP, in Mexico and most of the other countries studied. Several other studies have applied different stationarity tests to the Taylor data and find support for PPP in most countries including Mexico (Wallace and Shelley 2006; Bahmani-Oskooee et al. 2007; Wallace 2008, 2013).

Gómez-Zaldivar et al. (2013) use a unit root test developed by Kapetanios (2005) allowing multiple, exogenously determined breaks to test for PPP in monthly data for the Mexico–US real exchange rate, 1969–2010. They test an empirical model allowing for a break in the mean, without trend, and another with breaks in the mean and trend. The number of breaks varies across the two empirical specifications. The version without trend corresponds to QPPP as defined by Hegwood and Papell. All alternatives reject the unit root null in favor of a stationary series with breaks. Most of the breaks occurred between 1982 and 1987, a period associated with shifts toward opening Mexico to additional foreign investment and expanding international trade.

As figure 1 in Ventosa et al. shows, the Mexico–US real exchange has been volatile over the period 1930–2012, but the volatility is most pronounced in the decade beginning 1980 during which exchange rate controls were adopted and banks nationalized (Wallace 1999 provides a summary of events). VGW apply the Kim, Leybourne, and Newbold unit root test which is robust to volatility changes and find evidence of purchasing power parity for the Mexico–US real exchange rate. The test also reveals a structural break in the error variance occurring in July 1976, shortly before Mexico switched from a fixed nominal exchange rate regime to a managed float.

3 Methodology

If purchasing power parity holds between Mexico and the United States then Eq. (1) follows:

$$c_t^{\mathrm{M}} - e_t = c_t, \tag{1}$$

where c_t^{M} is the log of the cost of a market basket of tradable goods and services denominated in Mexican pesos, c_t is the log of the US dollar cost of the same market basket in the United States, and e_t is the log of the nominal exchange rate in pesos per US dollar.

As noted earlier, tests of PPP use price indexes rather than actual costs of a market basket. Price indexes do not correspond to the cost terms in Eq. (1). Consequently, an empirical test of PPP must modify the relationship yielding Eq. (2), where the logged period t price indexes for Mexico $\left(p_t^{\mathrm{M}}\right)$ and the United States (p_t) replace the costs of the market basket in Eq. (1) and an error term, u_t, is included. Since price indexes are not monetary measures the intercept, β_0, and coefficient, β_1, may be different from their implied values of zero and one, respectively, in Eq. (1).

$$p_t^{\mathrm{M}} - e_t = \beta_0 + \beta_1 p_t + u_t. \tag{2}$$

Testing for purchasing power parity is not as simple as Eq. (2) might suggest, because price levels and nominal exchange rates are often integrated of order one, thus the regression results may be spurious. However, if purchasing power parity holds then there is a stable, long-term relationship between the price levels in the two countries when expressed in terms of a common currency. That is the two common currency price levels must be cointegrated if the individual series are I(1) with Eq. (2) expressing the long-run relationship.

Unit root test results, discussed below, indicate that the Mexican and US price indexes and the peso/dollar nominal exchange rate are integrated of order one. Thus if purchasing power parity holds, the three series must be cointegrated. Common single equation approaches to testing for cointegration include the error correction model (ECM) and the autoregressive distributed lag (ADL) variant. However, it is well known that such testing is problematic because of the presence of nuisance parameters (Pesavento 2004) and the dependence of the distribution of the relevant test statistic on the variables included in the model. The distribution will vary, for example, with the location of a structural break, an issue of particular importance in this study. A researcher can bootstrap errors to uncover the empirical distribution, but bootstrapping may also create some difficulties. Harris and Judge (1998) show that bootstrapped errors may result in poor test size in a model with both stationary and nonstationary covariates and at least one cointegrating vector, particularly in small samples.

Enders, Im, and Lee offer an alternative to bootstrapping in evaluating the test statistics from ECM and ADL models. They propose using stationary instrumental variables for any I(1) series in the models. They show that the test statistics from the ECM and ADL estimations are asymptotically normal when the nonstationary variables are appropriately instrumented.[1] Test statistics are free of the nuisance parameter problem and can be evaluated using the standard normal distribution, thus simplifying inference in cointegration models.

The EIL methodology is followed in this paper. Letting the Mexican price level denominated in dollars be denoted f_t^{M}, that is $p_t^{\mathrm{M}} - e_t = f_t^{\mathrm{M}}$, the basic ECM is given by Eq. (3):

$$\Delta f_t^{\mathrm{M}} = \delta_0 + \delta_1 \left(f_{t-1}^{\mathrm{M}} - \widehat{\beta}_0 - \widehat{\beta}_1 p_{t-1} \right) + \delta_2 \Delta p_t + v_t, \tag{3}$$

where Δ is the difference operator, the $^\wedge$ above a term indicates an estimated value, the expression in parentheses is the estimated residual, $\widehat{u}_{t-1}$, from Eq. (2) lagged one period, that is the error correction term, and v_t is a white noise error. This specification assumes that the US price level is weakly exogenous thus all adjustments to the error correction term occur in the dollar-denominated Mexican price level. The US economy is much larger than that of Mexico so the weak exogeneity assumption seems reasonable; nonetheless it is tested before estimation of Eq. (3).

Lags of Δf_t^{M} can be added to Eq. (3) to address autocorrelation and dummy variables can be introduced as needed. The innovation of EIL is to show that instrumenting $\widehat{u}_{t-1}$ and any other nonstationary covariates in the model yields test statistics that are asymptotically normally distributed. As an instrument for any nonstationary I(1) covariate, x_{t-1}, they propose using the m period difference, $x_{t-1} - x_{t-m-1}$. For example, in the basic ECM given by Eq. (3) the instrument for $\widehat{u}_{t-1}$ would be $\widehat{u}_{t-1} - \widehat{u}_{t-m-1}$. No theoretical guidance is provided for selecting m, but EIL suggest using the value of m that minimizes the sum of the squared residuals of the test equation. Empirical results from specifications for a range of values of m are reported below. The ADL variation of Eq. (3) is given by Eq. (4).

[1] Selection of the appropriate IV is addressed below. A concise description of the EIL methodology can be found in Wallace (2013). The reader interested in a more complete description and proofs should consult Enders et al. (2008).

$$\Delta f_t^{\mathrm{M}} = \alpha_0 + \delta_1 f_{t-1}^{\mathrm{M}} + \alpha_1 p_{t-1} + \delta_2 \Delta p_t + \upsilon_t, \tag{4}$$

where $\alpha_0 = \delta_0 - \delta_1 \widehat{\beta}_0$, $\alpha_1 = -\delta_1 \widehat{\beta}_1$, and υ_t is a white noise error term. Again dummies and lags of the dependent variable can be introduced. In the case of the basic ADL model in Eq. (4) the instruments are $f_{t-1}^{\mathrm{M}} - f_{t-m-1}^{\mathrm{M}}$ and $p_{t-1} - p_{t-m-1}$ for the nonstationary series f_{t-1}^{M} and p_{t-1}, respectively. Of interest in testing for purchasing power parity is the coefficient δ_1 in Eqs. (3) and (4). If the price levels in terms of US dollars in the two countries are cointegrated, then $\widehat{\delta}_1$ will be significantly negative so that the null and alternative hypotheses are $H_0 : \delta_1 \geq 0, H_1 : \delta_1 < 0$. The coefficient on the error correction term will be negative and significant if the short run change in the dollar-denominated Mexican price level responds to deviations from the long-run relation. For example, if $\widehat{u}_{t-1}$ is positive it signifies that the Mexican price level measured in dollars is above its long-run equilibrium value relative to the US price level. If PPP holds the Mexican price level should decline in the short run so the coefficient on $\widehat{u}_{t-1}$ will be significantly negative. Correctly assessing the significance of a coefficient from the ECM or ADL specifications with integrated regressors is the difficulty addressed by the EIL procedure.

4 Data

The updated VGW data set used in this study includes three different price index measures for Mexico. One price level series is the wholesale price index for Mexico City available for 1930m01–2000m12. Wholesale prices for Mexico City are the only historical data available for the country prior to 1969. A second (third) price level series is derived by appending the producer price index excluding (including) petroleum for 2001m01–2014m09 to the Mexico City wholesale price level. The nominal exchange rate is expressed as the peso price of a US dollar. The US producer price index series is taken from the statistical database of the Federal Reserve Bank of St. Louis. All other data are drawn from Cárdenas (1994), annual Banxico reports, and the website for the Bank of Mexico. More details can be found in Ventosa et al. The data are plotted in Figs. 1, 2, 3, and 4. Due to the similarity of the producer price series with and without the price of petroleum, only the series inclusive of the oil price is shown. All series display trends; thus, a deterministic trend is included in the unit root and cointegration tests discussed below.

The augmented Dickey–Fuller test, the GLS detrended version of the Dickey–Fuller test, and the KPSS test are employed to check for stationarity of all series.[2] The ADF and GLS version of the Dickey–Fuller test statistics indicate failure to reject the unit root null in the three Mexican price level series, the US price level, and the nominal exchange rate. These two tests uniformly reject the unit root null in the first difference of each series at the 5% level or better indicating that the five

[2] Tests on the logged series in levels include a constant and trend. Tests on the first differenced logged series include only a constant since plots of the differenced series do not show trends. The Schwarz criterion is used to select lag length in the ADF and DF-GLS tests.

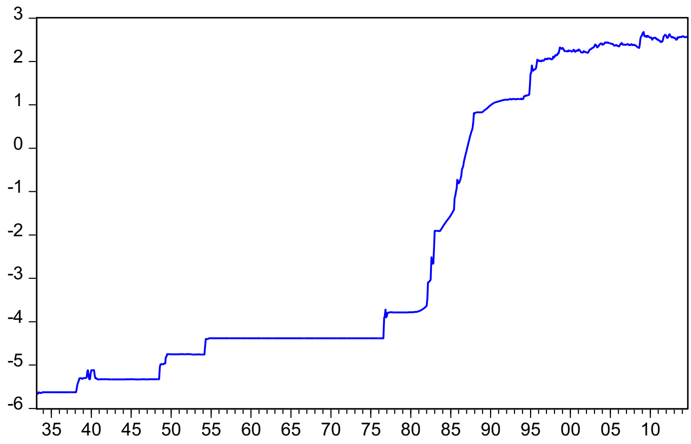

Fig. 1 Log nominal exchange rate, Pesos per US dollar 1933m03–2014m09

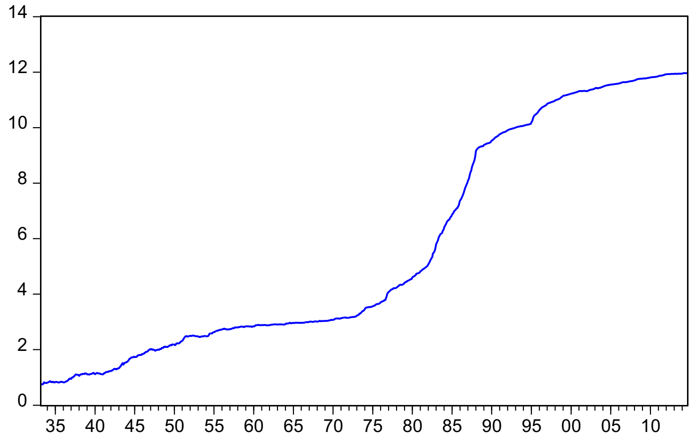

Fig. 2 Log producer price index-Mexico, 1933m03–2014m09. Includes petroleum

variables are integrated of order one. The Kwiatkowski, Phillips, Schmidt and Shin (KPSS) indicate rejections of the trend stationary null in both levels and first differences at the 5% level for the three Mexican price-level measures. Thus, the KPSS tests suggest that these three series are integrated of order two or higher. According to the KPSS test, the nominal exchange rate and the US price level are I(1) confirming the earlier finding.[3]

Structural breaks in the data set are likely especially since Mexico has experienced regime shifts since 1933. From 1933 to 1980 the nominal exchange rate was generally fixed interrupted by brief interludes of floating rates. Subsequently, the nominal exchange rate was allowed to fluctuate beginning with a managed float.

[3] For reasons addressed later in the paper the all unit root tests are applied to data series starting in 1933m03.

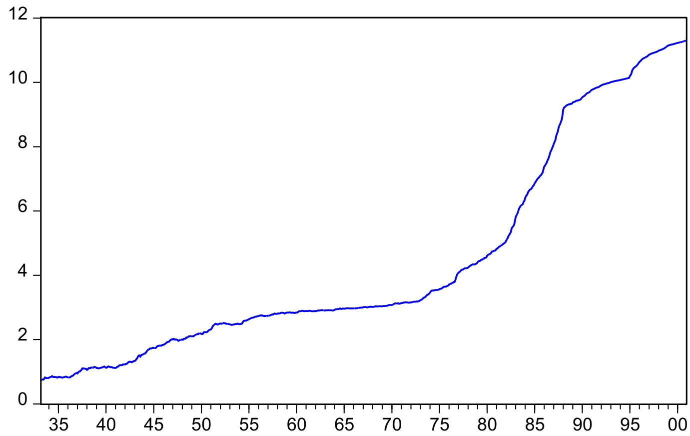

Fig. 3 Log wholesale price index, Mexico City 1933m03–2000m12

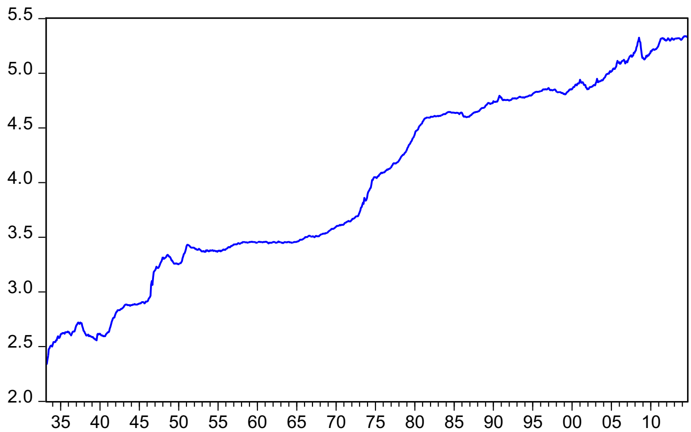

Fig. 4 Log US producer price index, 1933m03–2014m09

Since 1995 the Mexican government has allowed the peso/dollar rate to fluctuate with occasional interventions. In the mid-1980s trade policy shifted from a historical focus on import substitution to a more recent emphasis on exports.

The Lee and Strazicich (2003, 2013) henceforth LS, unit root test is used to test the null of a unit root in each series while allowing for structural breaks. Up to 40 lags are permitted with a general-to-simple procedure used to select lag length. The test is performed allowing for one or two breaks in both trend and intercept. The null hypothesis of a unit root with break cannot be rejected for any of the logged series under either number of breaks. Since no trend can be observed in the first differenced series, only intercept breaks are considered in testing for a second unit root. The null of a second unit root cannot be rejected for any of the three differenced Mexican price series allowing for one break; but when two breaks are permitted the LS tests reject the unit root null at the 5% level for all variables.

Despite some evidence against the I(1) hypothesis from the KPSS test and the one break version of the LS test for the logged Mexican price series, most of the evidence from unit root tests suggest that all series are integrated of first order and the remainder of the work is conducted under that assumption.

5 Results

The assumption of weak exogeneity of the US producer price index is tested in a vector ECM before proceeding with the EIL estimations. The Johansen test for cointegration indicates the presence of one cointegrating vector between the logs of the dollar-denominated Mexican price level and the US producer price index, regardless of the choice of Mexican price series. The Schwarz Bayesian criterion selects seven lags for a vector autoregression of the two variables so six are included in one version of the VECM. Another version of the VECM uses 12 lags of the endogenous variables to capture possible seasonal effects. Tests for autocorrelation indicate that the problem remains in the six lag VECM, but that autocorrelation is eliminated for the 12-lag version. All results reported below are based on the VECM with 12 lags.

The VECM, shown in Eqs. (5) and (6) below, has two endogenous variables, the change in the log of the dollar-denominated Mexican price level, Δf_t^{M}, and the change in the log of the US producer price index, Δp_t. The error correction portion is the term in parenthesis in both equations.

$$\Delta f_t^{\mathrm{M}} = c_1 + a_1\left(f_{t-1}^{M} - b_1 - b_2 p_{t-1}\right) + \sum_{i=1}^{12} d_{1i}\Delta f_{t-i}^{\mathrm{M}} + \sum_{i=1}^{12} h_{1i}\Delta p_{t-i} + e_{1t}, \tag{5}$$

$$\Delta p_t = c_2 + a_2\left(f_{t-1}^{\mathrm{M}} - b_1 - b_2 p_{t-1}\right) + \sum_{i=1}^{12} d_{2i}\Delta f_{t-i}^{\mathrm{M}} + \sum_{i=1}^{12} h_{2i}\Delta p_{t-i} + e_{2t}. \tag{6}$$

Weak exogeneity of the log of the US producer price index holds if the coefficient on the error correction term in Eq. (6), a_2, is not significantly different from zero. Three VECM are estimated, each using the dollar-denominated Mexican price level constructed from a different Mexican price series. Likelihood ratio tests of the null hypothesis $a_2 = 0$ have marginal significance levels ranging from 0.15 to 0.19 indicating failure to reject the null thus support for the weak exogeneity assumption.

Initial estimation of both the ECM and the ADL versions (Eqs. 3 and 4) using data for 1930m01–2014m09 shows that identifying empirical specifications of the ECM and ADL equations without autocorrelation requires the inclusion of very long lags of the dependent variable, regardless of the price index measure used. Interestingly, the need for long lags is obviated if the first 39 months of data are excluded from the sample; a specification with 12 lags of the dependent variable shows no evidence of autocorrelation when the period is restricted to

1933.03–2014.09 regardless of the price-level measure used in Eq. (2).[4] It is beyond the scope of the paper to investigate the reason(s) that removal of the first 39 months of data alters the specification so dramatically, but it may be due to unusual exchange rate or price level behavior caused by the removal of gold from circulation and the prohibition of the coinage of silver in Mexico during this period (see Wallace 1999 and sources cited therein). All estimation results reported below are for specifications of Eqs. (3) and (4) that include 12 lags of the dependent variable.

Table 1 reports the estimated value of δ_1 from the ECM in column 2 of the table and its estimated values in the IV specifications for $m = 14\ldots24$ in subsequent columns. The first column shows the error correction term used in the ECM. Zmex1 is $(f^{\mathrm{M}}_{t-1} - \widehat{\beta}_0 - \widehat{\beta}_1 p_{t-1})$ in Eq. (3) when f^{M}_{t-1} is calculated using the price index for Mexico that includes petroleum, and zmex2 is the error correction in Eq. (3) where the price index for Mexico excludes petroleum. Thus zmex1 and zmex2 are the estimated residuals obtained from the full sample, 1933m03–2014m09 estimations of Eq. (2). Zmex3 is the error correction term in Eq. (3) when using the wholesale price index for Mexico City for the period 1933m03 to 2000m12 to estimate Eq. (2).

The estimated value of δ_1 is negative and significant at the 1% level in the basic ECM, regardless of the particular Mexican price level series, indicating cointegration between the dollar-denominated Mexican and US price levels thus evidence of PPP. The remaining columns show estimates of δ_1 when error correction terms are instrumented for values of m ranging from 14 to 24. The choice of m has some effect on the size of $\widehat{\delta}_1$, but all are significant at the 5% level (critical value -1.645) and all but five of the estimated coefficients are significant at the 1% level. Thus, the results indicate that the US and Mexican price levels, denominated in dollars, are cointegrated a result supportive of purchasing power parity. The choice of Mexican price index used in the estimations makes very little difference in the results, for a given value of m the different indexes produce very similar estimates of δ_1. The estimated coefficient in bold and italicized font in the table is from the equation having the smallest sum of squared residuals (SSR) for the 11 IV specifications. The smallest SSR appears at the final value, $m = 24$, in the three different estimations.

In the ADL version of the empirical model δ_1 is the coefficient on the lagged value of the dollar-denominated Mexican price level, f^{M}_{t-1}. There are three different measures of the dollar-denominated price level in Mexico in Table 1: fmex1 includes the price of petroleum, fmex2 excludes petroleum price, and fmex3 is the wholesale price index for Mexico City. The two former measures span 1933m03 to 2014m09 while fmex3 ends in 2000m12. The estimated coefficients on the dollar-denominated Mexican price-level measures are displayed in Table 2. The basic ADL results, without instruments, suggest cointegration of the Mexican and US

[4] Once it was determined that the first few years of the study were responsible for difficulties in finding a satisfactory specification and those years eliminated, a general-to-simple procedure was followed. Starting with a specification for the ECM and ADL with 37 lags of the dependent variable, Wald tests for joint significance of the longest lags were carried out eliminating those lags that were not jointly significant at the 15% level or higher. Six different LM tests, gradually reducing the number of lags in the LM test, were undertaken for each ECM and ADL specification (without instruments). For both ADL and ECM 12 lags of $\Delta f^{\mathrm{M}}_{t}$ were sufficient to eliminate autocorrelation. Twelve lags were then imposed on the corresponding estimations with instruments.

Table 1 Estimation results for the error correction term coefficient: with and without instruments

ECM	No IV[a]	$m = 14$	$m = 15$	$m = 16$	$m = 17$	$m = 18$	$m = 19$	$m = 20$	$m = 21$	$m = 22$	$m = 23$	$m = 24$
ZMEX1	−0.0333	−0.1210	−0.1077	−0.0879	−0.0891	−0.0799	−0.0740	−0.0814	−0.0821	−0.0801	−0.0811	***−0.0716***
	−4.179	−2.078	−2.389	−2.307	−2.644	−2.657	−2.667	−3.080	−3.232	−3.288	−3.463	−3.199
ZMEX2	−0.0328	−0.1193	−0.1063	−0.0866	−0.0877	−0.0785	−0.0726	−0.0801	−0.0807	−0.0789	−0.0799	***−0.0705***
	−4.145	−2.054	−2.363	−2.278	−2.608	−2.614	−2.622	−3.036	−3.187	−3.246	−3.423	−3.159
ZMEX3	−0.0434	−0.1367	−0.1213	−0.0988	−0.0969	−0.0846	−0.0790	−0.0857	−0.0850	−0.0833	−0.0847	***−0.0753***
	−4.327	−2.219	−2.610	−2.535	−2.829	−2.770	−2.810	−3.218	−3.333	−3.409	−3.612	−3.352

[a] The 5% critical value for the ECMs without IVs is −3.218 and the 1% critical value is −3.804. The critical values are obtained using the response surfaces in Ericsson and MacKinnon (2002), as implemented in the program ECMtest.xls, version 1.0 (Ericsson and MacKinnon 2000). Since the test statistics from the IV estimations are asymptotically normal, the significant values are −1.645 at the 5% level and −2.327 at the 1% level

Table 2 Estimation results for the coefficient on the dollar-denominated Mexican price level: ADL model with and without instruments

ADL	No IV[a]	IV14	IV15	IV16	IV17	IV18	IV19	IV20	IV21	IV22	IV23	IV24
FMEX1	−0.0338	***−0.0693***	−0.0525	−0.0346	−0.0400	−0.0309	−0.0281	−0.0262	−0.0178	−0.0334	−0.0094	0.0050
	−4.212	−0.977	−0.729	−0.488	−0.567	−0.410	−0.360	−0.269	−0.147	−0.284	−0.0489	0.0185
FMEX2	−0.0333	***−0.0705***	−0.0531	−0.0351	−0.0401	−0.0309	−0.0281	−0.0268	−0.0193	−0.0355	−0.0132	−0.0012
	−4.179	−1.015	−0.755	−0.507	−0.585	−0.426	−0.375	−0.292	−0.170	−0.325	−0.076	−0.005
FMEX3	−0.0441	**−0.0436**	−0.0326	−0.0175	−0.0227	−0.0140	−0.0114	−0.0061	−0.0037	−0.0184	0.0050	0.0118
	−4.372	−0.478	−0.343	−0.188	−0.244	−0.143	−0.111	−0.049	−0.027	−0.137	0.025	0.049

[a] Critical values for the ADL model without IVs are based on the *t* statistic bounds test (unrestricted intercept, no trend) of Pesaran et al. (2001). A *t* statistic smaller than the I(1) bound of −3.53 (−4.10) indicates significance at the 5% (1%) level. The coefficient italicized in bold font is from the equation that minimizes the sum of the squared residuals

dollar-denominated price levels, evidence of PPP. The ADL model with instruments includes the same variables as the ECM so it is surprising that none of the $\widehat{\delta}_1$ in the IV specifications are significant; none of the test statistics are close to their normal distribution critical values. Apart from the significant coefficients on the three Mexican price-level measures in the ADL model without instruments, there is no evidence for purchasing power parity in the ADL results.

Long-span data are necessary for purchasing power parity tests because PPP is a long-run condition and because empirical evidence indicates that deviations from PPP have long half-lives.[5] As previously noted structural breaks arising from monetary policy, exchange rate, or trade policy regime changes are likely given the long time period spanned by the data. GVW argue that the commercial opening of Mexico, especially in the 1980s, likely changed the Mexico–US real exchange rate. They find evidence of PPP when allowing for mean shifts that they attribute to regime changes.

The contradictory results for the instrumented ECM and ADL estimations raise the question of whether structural breaks might explain the differences. This issue is investigated by including dummy variables to allow for mean shifts attributable to the structural breaks identified in GVW and VGW. The VGW approach identifies one break in the innovation variance and this occurs in July 1976. GVW find multiple breaks in the monthly mean real exchange rate during the 1969–2010 period. Dummies for the five breaks identified in Model D-5 of GVW and the single break in July 1976 are now included in the ECM and ADL models of Eqs. (3) and (4). Wald tests on the estimated dummy variable coefficients of the ECM and ADL models indicate that the dummies for 1979m12 and 1982m12 are jointly insignificant, so they are dropped from the ECM and ADL specifications as well as the IV estimations regardless of the Mexican price index used. Although the 1976m07 dummy is not significant in all estimations, it is significantly different from zero at the 5% level in some instances thus retained in all models. The final versions of the ECM and ADL models add dummies allowing intercept breaks in 1976m07, 1985m06, 1987m12, and 1998m09 to the earlier versions of Eqs. (3) and (4). Results are reported in Tables 3 (ECM) and 4 (ADL).

The inclusion of the dummy variables causes small changes in the estimated coefficients displayed in Table 1. The coefficients are now slightly smaller, but, as in the specification without dummy variables, all are significant regardless of the Mexican price index used and choice of m for the instrumental variable. Thus the same qualitative conclusion is obtained as from the instrumented ECM estimations without dummies; there is cointegration of the two common currency price levels, thus evidence of PPP.

Including four dummy variables changes the ADL results with instruments and leads to a different conclusion regarding purchasing power parity. As in the specifications without dummy variables the estimated value of δ_1 is negative in the

[5] Estimating half-lives is beyond the scope of this study, but estimates vary widely. VGW, using virtually the same data set as in this study find half-lives ranging from 1.37 to 2.41 years for Mexico depending on the time period studied. Rogoff (1996) surveys the literature on PPP and finds half-life estimates of 2.8–4.75 years.

Table 3 Estimation results for the error correction term coefficient: models with time period dummies

ECM	No IV[a]	IV14	IV15	IV16	IV17	IV18	IV19	IV20	IV21	IV22	IV23	IV24
ZMEX1	−0.0630	−0.1611	−0.1361	−0.1095	−0.1095	−0.0964	−0.0880	−0.0958	−0.0962	−0.0937	−0.0952	***−0.0854***
	−6.123	−2.217	−2.559	−2.471	−2.834	−2.842	−2.847	−3.286	−3.449	−3.510	−3.708	−3.441
ZMEX2	−0.0633	−0.1604	−0.1354	−0.1088	−0.1086	−0.0954	−0.0870	−0.0948	−0.0953	−0.0929	−0.0947	***−0.0849***
	−6.145	−2.197	−2.537	−2.447	−2.803	−2.803	−2.806	−3.247	−3.408	−3.472	−3.673	−3.406
ZMEX3	−0.0678	−0.2046	−0.1671	−0.1328	−0.1278	−0.1090	−0.1003	−0.1075	−0.1064	−0.1039	−0.1058	***−0.0954***
	−5.847	−2.279	−2.736	−2.692	−3.011	−2.959	−3.005	−3.434	−3.562	−3.649	−3.874	−3.623

[a] Critical values are affected by the presence of time period dummies hence are unknown. Normal distribution significant values are −1.645 at the 5% level and −2.327 at the 1% level for the IV estimations

Table 4 Estimation results for the coefficient on the dollar-denominated Mexican price level: ADL models with time period dummies

ADL	No IV[a]	IV14	IV15	IV16	IV17	IV18	IV19	IV20	IV21	IV22	IV23	IV24
FMEX1	−0.0672	−0.1212	−0.1111	−0.0924	−0.0956	−0.0848	−0.0785	−0.0872	−0.0889	−0.0893	−0.0915	***−0.0830***
	−6.298	−2.105	−2.310	−2.245	−2.647	−2.665	−2.724	−3.145	−3.304	−3.486	−3.615	−3.376
FMEX2	−0.0676	−0.1215	−0.1110	−0.0921	−0.0949	−0.0840	−0.0777	−0.0865	−0.0882	−0.0887	−0.0910	***−0.0825***
	−6.321	−2.108	−2.304	−2.231	−2.624	−2.633	−2.689	−3.112	−3.270	−3.452	−3.583	−3.344
FMEX3	−0.0679	−0.1443	−0.1397	−0.1170	−0.1153	−0.0992	−0.0922	−0.1016	−0.1029	−0.1029	−0.1067	***−0.0977***
	−5.838	−1.773	−2.105	−2.155	−2.521	−2.564	−2.681	−3.052	−3.217	−3.406	−3.562	−3.384

[a] Critical values are affected by the presence of time period dummies hence are unknown. Significant values are −1.645 at the 5% level and −2.327 at the 1% level for the IV estimations

standard ADL model without instruments for all three Mexican price-level measures. When using instruments, $\widehat{\delta}_1$ is significantly negative at the 5% level for all values of m and of comparable size to the estimated δ_1 in the corresponding ECM.[6] The ADL specifications with dummies suggest the change in the dollar-denominated price level in Mexico is negatively related to its long-run level the previous month, evidence that it adjusts to its purchasing power parity value. The significance of most of the time period dummies suggests that this long-run value has shifted over time. As in the ECM specifications, the sum of the squared residuals is minimized at the highest value of m used to create the instruments regardless of the Mexican price level variable used in the estimation.

The significant dummy variables show that the relationship of the dollar-denominated price levels of Mexico and the US underwent mean shifts during the study period. Ex post it is easy to find events that might cause mean shifts in the relationship of the price levels of the two countries, but proving causality is impossible. Nonetheless, there are some likely candidates. Ventosa-Santàularia et al. associate a break in the innovation variance in July 1976 with macroeconomic difficulties in Mexico culminating in a shift from a fixed nominal exchange rate vis-à-vis the US dollar to a managed float in September 1976. Gómez et al. note that a policy shift toward a more open economy in Mexico began in the 1980s with membership in the General Agreement on Tariffs and Trade and the start of negotiations for the North American Free Trade Agreement. Such a policy shift, they argue, would lead to a change in the traded/non-traded goods mix and cause a real exchange rate shift.

6 Summary and conclusions

The evidence on purchasing power parity in Mexico is mixed. Some researchers find that PPP holds, others do not, and others argue that it holds but that structural shifts in the relation have occurred. A recent approach to cointegration developed by Enders, Im, and Lee is used to test for PPP between the United States and Mexico. The EIL procedure uses stationary variables as instruments for nonstationary covariates in the error correction and ADL models. When instruments are used the test statistics are asymptotically normally distributed thus eliminating the nuisance parameter problem and avoiding the need to bootstrap errors.

Initial estimation of the basic ECM and its ADL variant yields conflicting evidence regarding PPP. All ECM specifications, instrumented or not, support the PPP hypothesis; none of the ADL specifications with instruments do. Since other researchers have found evidence of mean shifts in the Mexico–US exchange rate, dummy variables for six exogenously imposed breaks are included in the empirical models. Coefficients on four of the six breaks are significant in at least one variant

[6] As a robustness check on the results, both the error correction and autoregressive distributed lag models are also re-estimated replacing the four intercept break dummies with intercept and trend-shift dummies at dates identified by the Lee–Strazicich unit root tests. Results are very similar to those shown in Tables 3 and 4, hence are not reported.

of the ECM and ADL models. With or without IVs the results from the ECM with dummies provide evidence of PPP, as do the ECM without dummies. Interestingly, the ADL models with instruments and dummies also support PPP; all coefficients on the dollar-denominated Mexican price level are significantly negative, unlike the specifications without dummies. Overall the evidence indicates that PPP holds for the 1933m03–2014m09 period between the US and Mexico, but that the purchasing power parity relation between the two countries has changed over this time period.

References

Ávalos Huerta A, Hernández Trillo F (1995) Comportamiento del tipo de cambio real y desempeño económico en México. Econ Mex IV:239–263

Bahmani-Oskooee M, Kutan A, Zhou S (2007) A century of purchasing power parity: further evidence. Econ Bull 6:1–9

Cárdenas E (1994) La hacienda pública y la política económica 1929–1958. El Colegio de México and Fondo de Cultura Económica, México DF

Cassel G (1918) Abnormal deviations in international exchanges. Econ J 28:413–415

Enders W, Im KS, Lee J (2008) Cointegration tests using instrumental variables with an example of the U.K. demand for money. Unpublished working paper. http://wenders.people.ua.edu/time-series-methods.html. Accessed 20 Jan 2010

Ericsson NG, MacKinnon JG (2000) ECMtest.xls, Version 1.0

Ericsson NG, MacKinnon JG (2002) Distributions of error correction tests for cointegration. Econ J 5:285–318

Gómez-Zaldivar M, Ventosa-Santàularia D, Wallace FH (2013) The PPP hypothesis and structural breaks: the case of Mexico. Empir Econ 45:1351–1359. doi:10.1007/s00181-012-0653-6

Harris RID, Judge G (1998) Small sample testing for cointegration using the bootstrap approach. Econ Lett 58:31–37

Hegwood ND, Papell DH (1998) Quasi purchasing power parity. Int J Financ Econ 3:279–289. doi:10.1002/(SICI)1099-1158(199810)3:4<279:AID-IJFE83>3.0.CO;2-K

Kapetanios G (2005) Unit-root testing against the alternative hypothesis of up to m structural breaks. J Time Ser Anal 26:123–133. doi:10.1111/j.1467-9892.2005.00393.x

Lee J, Strazicich MC (2003) Lagrange multiplier unit root test with two structural breaks. Rev Econ Stat 85:1082–1089

Lee J, Strazicich MC (2013) Minimum LM unit root test with one structural break. Econ Bull 33:2483–2492

McLeod D, Welch JH (1992) El libre comercio y el peso. Econ Mex I:193–235

Mejía Reyes P, González Núñez JC (1996) La paridad de poder de compra en el largo plazo: El caso de México. Econ Mex V:37–62

Noriega A, Medina L (2003) Quasi-purchasing power parity: structural change in the Mexican peso/U.S. dollar real exchange rate. Estudios Económicos 18:227–236

Pesaran MH, Shin Y, Smith RJ (2001) Bounds testing approaches to the analysis of level relationships. J Appl Econ 16:289–326. doi:10.1002/jae.616

Pesavento E (2004) Analytical evaluation for the power of tests for the absence of cointegration. J Econ 122:349–384. doi:10.1016/j.jeconom.2003.10.025

Rogoff K (1996) The purchasing power parity puzzle. J Econ Lit XXXIV:647–668

Santaella JA, Vela AE (2005) The 1987 Mexican disinflation program: an exchange-rate-based stabilization? Econ Mex XIV:5–40

Taylor A (2001) Potential pitfalls for the purchasing power parity puzzle? Sampling and specification biases in mean-reversion tests of the law of one price. Econometrica 69:473–498

Taylor A (2002) A century of purchasing power parity. Rev Econ Stat 84:139–150

Ventosa-Santàularia D, Gómez-Zaldivar M, Wallace FH (2015) The real exchange rate, regime changes, and volatility shifts. Appl Econ 47:2445–2454. doi:10.1080/00036846.2015.1005821

Wallace FH (1999) Long run neutrality of money in the Mexican economy. Appl Econ Lett 6:637–639

Wallace FH (2008) Nonlinear unit root tests of PPP using long-horizon data. Econ Bull 6:1–8

Wallace FH (2013) Cointegration tests of purchasing power parity. Rev World Econ 149:779–802. doi:10.1007/s10290-013-0165-2

Wallace FH, Shelley GL (2006) An alternative test of purchasing power parity. Econ Lett 92:177–183. doi:10.1016/j.econlet.2006.01.033

Wallace FH, Shelley GL, Cabrera Castellanos LF (2011) La paridad de poder de compra en México (1930–1960). Trimest Econ 78:675–693