Some Further Evidence on the Law of One Price: The Law of One Price Still Holds

John Baffes

International trade models often postulate the existence of a representative price, i.e., the price which prevails at all markets. This is known as the "Law of One Price." In this paper, the law of one price is tested for seven commodities among four countries by explicitly considering transaction costs. The empirical evidence suggests that, in most cases, the law of one price cannot be rejected as a maintained hypothesis. Furthermore, for the remaining cases transaction costs seem to cause the failure.

Key words: cointegration, "Law of One Price," stationarity, transaction costs.

International trade models of agricultural commodities often postulate the existence of a representative price, i.e., a price which prevails at all markets. This is known as the "Law of One Price" (LOP). The LOP assumption is crucial, especially in empirical studies, because its violation may invalidate the studies' conclusions. For example, a model of optimal intervention in a specific commodity market, domestic or international, requires a representative price. If an inappropriate price is used, policy recommendations based on such a model will not have the expected effects. This paper focuses on the consistency of the LOP postulate with empirical evidence.

Traditionally, tests of the LOP have applied the following procedure: A commodity price in one country is regressed on the same commodity's price in another country. Then, the slope coefficient is restricted to equal one and (possibly) the intercept term is restricted to equal zero. If such restrictions are not rejected, the conclusion is that the LOP holds. However, such procedures have not yielded favorable results regarding the LOP (e.g., Isard and Richardson among others). Those tests have been criticized on several grounds. For example, Crouhy-Veyrac, Crouhy, and Melitz cite the omission of transfer costs; if these costs were taken into consideration, favorable results regarding the LOP would be derived. More recently, Goodwin, Grennes, and Wohlgenant argued that the price prevailing in one country should be compared with the expected price of the other country; further, they argued that transaction costs should be explicitly incorporated in the model. They used an expectations-augmented model by including transaction costs and found support for the LOP. Protopapadakis and Stoll (1986) tested whether the LOP holds by making an explicit distinction between the short and the long run. They found supportive evidence for the long-run LOP, which is consistent with their earlier findings (Protopapadakis and Stoll 1983). Ardeni challenged the stationarity properties of the prices, tested the LOP as a long-run relationship by using cointegration, and found that in most cases the LOP does not hold.

The objective of this study is to test whether the LOP holds in the long run. Although the procedure used here is the one suggested by Ardeni (i.e., stationarity tests and subsequently cointegration), the particular test differs in two respects: (*a*) the restricted version of the cointegration test is used instead, and (*b*) transportation costs receive special attention as a cause for failure of the LOP. The next section discusses the concepts of the order of integration and cointegration and how they relate to the LOP. Then, an alternative testing procedure is proposed. Empirical results are presented in the second section. The third section extends the model by incorporating transportation costs and discusses some methodological issues. Finally,

John Baffes is a graduate research assistant in the Department of Agricultural and Resource Economics, University of Maryland.

The author wishes to thank Pier Giorgio Ardeni, who very kindly provided most of the data used in the analysis. Helpful comments from an anonymous *Journal* referee are greatly appreciated. The author is solely responsible for any remaining errors.

the last section summarizes the major findings and presents directions for future research.

The Law of One Price

Consider two countries, engaging in international trade of one commodity. In its strict sense, the LOP states that the price of this commodity (expressed in common currency unit) should be the same in the two countries if the necessary adjustments with respect to transaction costs have been made. This is a fundamental principle of commodity arbitrage. In this study we test the LOP as a long-run relationship by exploring certain time-series properties of the variables of interest.

Testing the Law of One Price

A starting point for testing the LOP is the determination of the order of integration of the prices being considered. To define the order of integration, consider a series with a finite spectrum which is nonzero at all frequencies. This series is said to be weakly stationary or integrated of order zero, and is denoted as $I(0)$. If the series must be differenced once to become $I(0)$, it is then said to be integrated of order one, $I(1)$. In general, a series that must be differenced d times to become $I(0)$ is called $I(d)$. A major difference between an $I(0)$ and an $I(d)$ ($d > 0$) series is that the $I(0)$ series has a finite mean and variance, while for an $I(d)$ series those magnitudes do not exist. In what follows, the terms stationarity (or stationarity in levels) and $I(0)$ will be used interchangeably, although strictly speaking they do not coincide.[1]

Consider now the price of a specific commodity in two countries at time t expressed in common currency unit, denoted as x_{1t} and x_{2t}^*. Then, distinguish the following three cases:

(a) $x_{1t} \sim I(0)$ and $x_{2t}^* \sim I(0)$. Because both prices are $I(0)$, their means and variances exist; this in turn implies that the LOP holds as a long-run relationship as both prices fluctuate around their means.[2] Consideration of the weak version of the LOP implies possible differences in means. Such differences might reflect stationary components such as middleman profits, taxes, and tariffs. This case can be extended in several ways by considering stricter versions of the LOP. For example, one can form the ratio of the variances of prices and then test the hypothesis that this ratio is one. (See Campbell and Shiller for applications of such test.) Alternatively, if prices follow autoregressive (AR) processes, say $x_{1t} \sim AR(n)$ and $x_{2t}^* \sim AR(k)$, the hypothesis of whether $n = k$ can be tested. Finally, it is valid to regress x_{1t} on x_{2t}^* and test the restriction that the slope coefficient equals one and (possibly) the intercept term equals zero, since all properties of traditional testing techniques are applicable. This last test, however, is subject to the assumption that one price is exogenous because, if prices are simultaneously determined, endogeneity problems are present. Finally, it should be mentioned that prices are usually found to be at least $I(1)$.

(b) $x_{1t} \sim I(d)$, $x_{2t}^* \sim I(b)$, $d \neq b$. In this case, prices have different orders of integration and the LOP does not hold because at least one of either x_{1t} or x_{2t}^* will exhibit explosiveness. This can be understood if x_{1t} is $I(0)$ and x_{2t}^* is $I(1)$; i.e., x_{2t}^* contains explosive components which cannot be explained by x_{1t} alone.

(c) $x_{1t} \sim I(d)$, $x_{2t}^* \sim I(d)$, $d > 0$. Here, both prices have the same order of integration which is greater than zero. Hence, additional information is needed to examine the validity of the LOP. Such information is obtained from the theory of cointegration (Engle and Granger). In simple terms, cointegration states that, even though some explosive pattern characterizes both prices, there might exist a (unique) parameter which brings them together in the long run so that their linear combination is of a lower order of integration than the original series.[3] In such a case, x_{1t} and x_{2t}^* form a cointegrated system. Cointegration has been used extensively in similar concepts. For example Enders (1988), Corbae and Ouliaris, and Taylor and McMahon used cointegration to test the purchasing power parity hypothesis, while Hakkio and Rush, Enders (1989), and Coleman used cointegration to test the efficient market hypothesis in currency markets.

In general, to determine whether x_{1t} and x_{2t}^* are cointegrated, the following regression is formed:

[1] For the exact definition and properties of integrated series, see Granger (1981).

[2] In a similar context, Hamilton and Flavin tested the limitations of domestic government borrowing. In particular they found that U.S. surplus and U.S. internal debt are stationary and concluded that the U.S. government runs a balanced budget in expected value terms.

[3] A proof regarding the uniqueness of the cointegration parameter in the bivariate case can be found in Hendry (p. 202).

$$(1) \qquad x_{1t} = \mu + \beta x_{2t}^{*} + \varepsilon_t,$$

where μ and β denote parameters to be estimated. If ε_t is integrated of order $b(b < d)$, x_{1t} and x_{2t}^{*} are said to be cointegrated. For the specific case that $b = 0$, the LOP holds since both prices move together in the long run. (1) needs to be estimated only if β is unknown. This, however, is not the case when the LOP holds. For example, if the variables of interest were the price of a particular stock and its associated dividend or consumption and income, β would represent the rate of return or the marginal propensity to consume, which are unknown parameters. However, if parameter β is known from theory, as with the LOP, there is no need to estimate it. In particular, the LOP postulates that the cointegration parameter is one. In fact, unity is the only value of the cointegration parameter for which the LOP (i.e., the maintained hypothesis) holds. In this case, the cointegration test is transformed into a stationarity test of the difference between the two prices. (This is stated formally in a proposition and a proof reported in the appendix.)[4] Therefore, testing the LOP in the case that x_{1t} and x_{2t}^{*} are both $I(d)$, $d > 0$ is equivalent to taking the following difference:

$$(2) \qquad z_t = x_{1t} - x_{2t}^{*},$$

and testing whether z_t, is $I(0)$. Here the problem becomes a unit root test of a univariate case (Engle and Yoo). Testing for cointegration by taking the difference of two variables is common practice.[5] Admittedly, it is the case that if x_{1t} and x_{2t}^{*} are cointegrated, regression (1) "should give an excellent estimate of the true cointegration coefficient" (Granger 1986, p. 219) so that one would prefer (1) over (2). However, this is true in large samples only.[6] This entails that information from economic theory is utilized, leading to the preference of (2) over (1).

The intuitive interpretation of case (c), which is the most relevant since all prices are $I(1)$ as will be shown later, is presented in figures 1a through 2b. Figure 1a depicts the price of wheat in Canada and Australia (1966:2–1983:4 time period) and shows that both prices are characterized by an explosive pattern over the time period examined. However, when the price differential is considered (fig. 1b), the explosive patterns observed in figure 1a are no longer present because the price differential has a tendency to return to an earlier value (i.e., the mean) very often. Consider now the case of wool between U.K. and Australia (1966:1–1986:1 time period). Again, an explosive pattern is observed when the levels of prices are taken (fig. 2a). Contrary to the previous case, however, this explosive pattern is preserved when the price differential is considered (fig. 2b). As the empirical evidence will show later (table 3), those two sets of prices represent two polar cases with respect to cointegration properties.

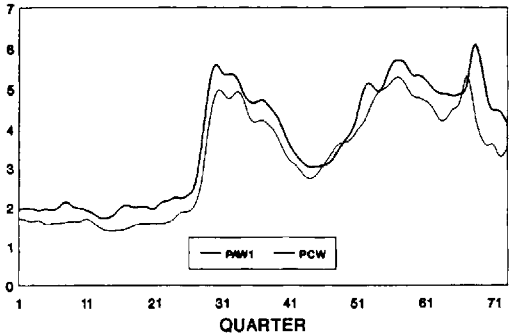

Figure 1a. Wheat price in Canada and Australia

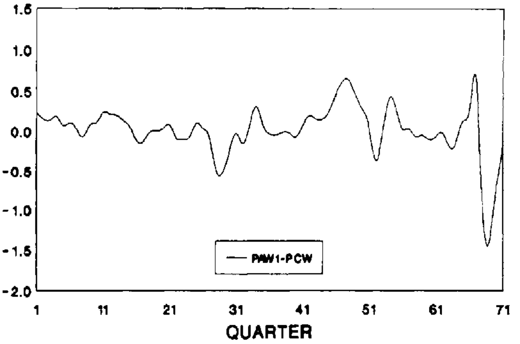

Figure 1b. Wheat price differential

---

[4] In a recent study, Ahking tested the validity of the purchasing power parity (PPP) principle between U.S. and U.K. by using cointegration. In commenting on the difference between the unrestricted and restricted version of the cointegration regression Ahking (p. 913) states: "It should be noted that cointegration of exchange rate and prices is necessary but not sufficient condition for long-run PPP. Long-run PPP further requires a unit cointegrating vector." So, the proposition reported in the appendix can be viewed as giving sufficient conditions regarding the validity of the LOP as a long-run relationship.

[5] For example, consider the term structure of the interest rate. The cointegration parameter of the short- and long-term interest rate equals one, as implied by the expectations theory (Campbell and Shiller). Corbae and Ouliaris have used the theoretical restriction to test the purchasing power parity. Other studies of a similar nature include Hall and Ambler.

[6] The OLS estimate of β from (1) is biased even asymptotically (Stock). Banerjee et al. have shown through Monte Carlo experiments that the bias can be of substantial magnitude sometimes.

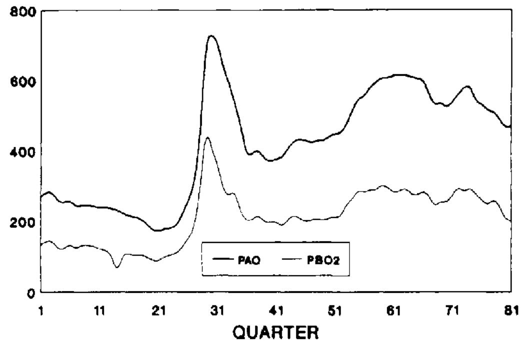

Figure 2a. Wool price in Australia and the U.K.

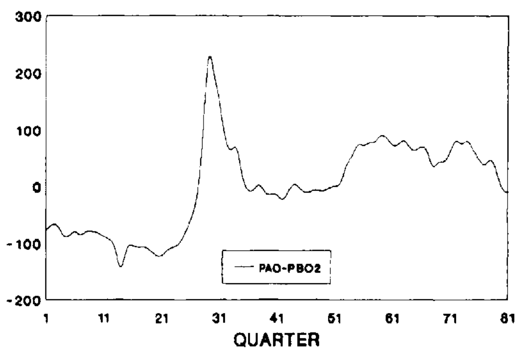

Figure 2b. Wool price differential

Determining the Order of Integration

To determine the sensitivity of the results with respect to stationarity properties, the following three procedures were employed: (*a*) Dickey-Fuller (DF), (*b*) augmented Dickey-Fuller (ADF), and (*c*) Durbin-Watson (DW). The DF test is based on the regression; $\Delta z_t = \mu + \beta z_{t-1} + \varepsilon_t$, where z_t denotes the variable being tested and Δ denotes the difference operator (i.e., $\Delta z_t = z_t - z_{t-1}$); μ and β are parameters to be estimated. The hypothesis tested is

$$\mathrm{H}_0\text{: } z_t \text{ is not } I(0) \text{ against } \mathrm{H}_1\text{: } z_t \text{ is } I(0).$$

H_0 is rejected if the estimate of β is negative and significantly different from zero. The ADF test accounts for the possibility tht ε_t is not white noise and is based on the regression: $\Delta z_t = \mu + \beta z_{t-1} + \Sigma_{i=1}^{\tau} \gamma_i \Delta z_{t-i} + \varepsilon_t$, where τ is selected so that ε_t is white noise; μ, β, and γ_i's are parameters to be estimated as before. Again, H_0 is rejected if β is negative and significantly different from zero. Finally, the DW test is based on the Durbin-Watson statistic of: $z_t = \mu + \varepsilon_t$. Low Durbin-Watson statistic indicates that z_t is not $I(0)$.

If H_0 is not rejected, we take first differences and repeat the same tests. In this case, the hypothesis becomes

$$\mathrm{H}_0\text{: } z_t \text{ is not } I(1) \text{ against } \mathrm{H}_1\text{: } z_t \text{ is } I(1).$$

Theoretically, this procedure is continued until H_0 is rejected in which case H_1 determines the order of integration. In practice, however, the procedure is terminated in first (and sometimes in second) differences because most economic variables are $I(1)$ and rarely $I(2)$.

Recently, several other procedures, which test for cointegration and impose less restrictions on the stochastic behavior of the error term of the cointegration regression, have been proposed (e.g., Perron; Phillips and Ouliaris). However, because the main purpose of this study is to test the LOP given the unitary restriction, these procedures were not applied.

Empirical Results

This section discusses empirical results which are in accordance with the testing procedure outlined above. The LOP was tested for the following commodities: wheat, tea, beef, sugar, wool, zinc, and tin, and countries: United States, Canada, Australia, and United Kingdom. The number of observations ranges from 71 to 118, representing quarterly average prices reported in the International Monetary Fund (*International Financial Statistics*). Because the data sets are identical, further details regarding the data and definitions of the variables can be found in Ardeni.

The first step regarding empirical analysis is the determination of the order of integration of prices. First, the levels of the prices were considered. Table 1 reports results for all three tests. The hypothesis that prices are not $I(0)$ was not rejected. All tests supported this conclusion, with the single exception of the *PBZ* and *PUZ* cases, for which the null hypothesis was rejected at the 10% level of significance (ADF test only). Second, first differences of the prices were considered. Results regarding stationarity tests for first differences are reported in table 2 and uniformly indicate that all prices are $I(1)$ (1% level of significance). To summarize, results from tables 1 and 2 show that the LOP for the specific commodities and time periods considered should be

Table 1. Stationarity Tests for Prices—Levels

Variable	Number of Observations	DF	ADF	DW
P_{AW1}	84	−1.54	−1.81	0.082
P_{AW1}	71	−0.41	−0.98	0.061
P_{AW2}	117	−1.45	−2.18	0.062
P_{AW2}	71	−1.43	−2.36	0.088
P_{UW}	71	−1.05	−1.73	0.057
P_{UW}	84	−1.52	−1.67	0.081
P_{UW}	117	−1.33	−1.46	0.060
P_{CW}	71	−1.38	−1.88	0.095
P_{BT}	84	−1.06	−1.26	0.173
P_{UT}	84	−0.70	−0.02	0.112
P_{AB}	89	−1.56	−1.49	0.064
P_{UB}	89	−1.29	−1.25	0.045
P_{AS}	80	−1.97	−1.97	0.160
P_{BS}	118	−1.39	−1.85	0.083
P_{BS}	80	−1.58	−2.00	0.116
P_{US}	118	−2.07	−2.47	0.149
P_{US}	80	−2.16	−2.01	0.205
P_{AO}	81	−1.38	−1.68	0.087
P_{BO1}	81	−1.05	−2.02	0.048
P_{BO2}	81	−1.42	−2.15	0.079
P_{BZ}	80	−1.88	−2.82*	0.153
P_{BZ}	89	−1.19	−1.57	0.030
P_{CZ}	80	−0.96	−1.22	0.026
P_{UZ}	89	−1.99	−2.72*	0.148
P_{UZ}	80	−0.99	−1.29	0.033
P_{BN}	83	−0.90	−0.93	0.030
P_{UN}	83	−0.82	−0.88	0.028

Notes: The first subscript identifies country: *A* is Australia; *U*, U.S.; *C*, Canada; and *B*, U.K. The second subscript identifies commodity: *W* is wheat; *T*, tea; *B*, beef; *S*, sugar; *O*, wool; *Z*, zinc; and *N*, tin. The numbers 1 and 2 following wheat for Australia and wool for U.K. indicate two different measures of prices.

A star (*) indicates rejection of the null hypothesis at the 10% level of significance (the critical value is −2.58).

The lag length regarding the ADF test was determined by the Akaike information criterion allowing up to a length of four lags.

In order to be consistent with the cointegration tests and at the same time use as many observations as possible, stationarity tests for some prices were conducted using more than one set of observations. For example, for P_{AW2} and P_{UW} there were 117 available observations, while for P_{CW} there were 71 observations only. For the stationarity tests of the pairs P_{AW2}–P_{CW} and P_{UW}–P_{CW}, we truncated P_{AW2} to 71. Hence, we tested for stationarity of P_{AW2} by using both 117 and 71 observations.

tested through cointegration as outlined in case (*c*) above.

Results regarding cointegration or alternatively stationarity of the price differential as defined in (2) are reported in table 3. In general, the results are favorable for the LOP. There is an agreement among all three tests as to the validity of the LOP. Considering the 5% level of significance, results can be summarized as follows: Out of a total of sixteen cases examined, the DF test does not reject the LOP in nine cases, the ADF test in ten cases, and the DW test in thirteen cases. In eight cases, all three tests indicate that the LOP cannot be rejected (5% level of significance). There is almost unanimous rejection of the LOP for wool [i.e., P_{AO}–P_{BO1} and P_{AO}–P_{BO2} are not $I(0)$] and in two of the wheat cases (P_{AW2}–P_{UW} and P_{AW2}–P_{CW}). Thus, rejection of the LOP is a price-specific problem rather than a general failure.[7]

Model Extensions

The primary focus of this section is to examine whether failure of the LOP for the P_{AW2}–P_{UW} and P_{AW2}–P_{CW} cases results from transportation costs. (The issue of transportation costs is discussed in Protopapadakis and Stoll 1983; Good-

[7] At this point one might ask why such differences between findings of this study and Ardeni's findings are present, given that the data set is identical and the testing procedures are similar. Ardeni considers the regression: $\ln(x_{1t}) = \alpha + \beta \ln(x^*_{2t}) + \varepsilon_t$ and tests for stationarity of ε_t, which upon rearranging terms becomes: $\varepsilon_t = \ln(x_{1t}/x^{*\beta}_{2t}) - \alpha$. The present study considers the difference: $z_t = (x_{1t} - x^*_{2t})$ and tests for stationarity of z_t; x_{1t} and x^*_{2t} denote levels of the prices. It is reasonable to expect that, even if z_t is $I(0)$, ε_t might not be $I(0)$ because of the nonlinear transformations performed upon x_{1t} and x^*_{2t}. Another reason for the differences in the findings is that the tabulated "*t*-values" utilized here are lower (in absolute terms) than the ones used in the cointegration regressions. This is the case because this study uses known cointegration parameters.

Table 2. Stationarity Tests for Prices—First Differences

Variable	Number of Observations	DF	ADF	DW
PAW1	84	−8.29	−5.28	1.854
PAW1	71	−4.48	−4.07	1.135
PAW2	117	−6.94	−4.42	1.195
PAW2	71	−5.06	−4.23	1.555
PUW	71	−4.67	−5.07	1.028
PUW	84	−8.23	−4.80	1.833
PUW	117	−9.79	−8.29	1.835
PCW	71	−7.34	−6.15	1.844
PBT	84	−8.36	−9.38	1.867
PUT	84	−8.06	−8.53	1.792
PAB	89	−9.18	−7.25	1.990
PUB	89	−8.58	−7.08	1.863
PAS	80	−9.39	−5.81	2.145
PBS	118	−8.47	−6.92	1.545
PBS	80	−6.92	−5.63	1.546
PUS	118	−7.76	−6.38	1.382
PUS	80	−6.23	−6.26	1.354
PAO	81	−7.03	−4.31	1.579
PBO1	81	−5.57	−4.36	1.171
PBO2	81	−5.32	−4.68	1.091
PBZ	80	−6.24	−5.45	1.364
PBZ	89	−5.54	−5.51	1.088
PCZ	80	−5.90	−4.71	1.332
PUZ	89	−6.72	−5.36	1.383
PUZ	80	−6.20	−4.71	1.376
PBN	83	−7.05	−6.72	1.538
PUN	83	−6.81	−6.47	1.472

Notes: for definitions of the variables see table 1.
The null hypothesis that the prices are not $I(1)$ is rejected for all prices at the 1% level of significance (significance levels and the corresponding critical values are reported in table 3).

Table 3. Cointegration Tests between Prices

Variable	Number of Observations	DF	ADF	DW
PAW1–PUW	84	−4.62***	−4.85***	0.844***
PAW1–PCW	71	−3.88***	−3.35**	1.033***
PAW2–PUW	117	−1.28	−2.00	0.094
PAW2–PCW	71	−1.82	−2.22	0.240
PUW–PCW	71	−3.45**	−3.82***	0.635**
PBT–PUT	84	−3.80***	−3.35**	0.594***
PAB–PUB	89	−3.18**	−2.76*	0.393***
PAS–PBS	80	−3.60***	−3.12**	0.614***
PAS–PUS	80	−3.21**	−2.91**	0.485***
PBS–PUS	118	−4.40***	−4.85***	0.574***
PAO–PBO1	81	−2.42	−1.89	0.282**
PAO–PBO2	81	−2.08	−2.23	0.190
PBZ–PCZ	80	−2.70*	−3.70***	0.361**
PBZ–PUZ	89	−2.63*	−3.91***	0.307**
PCZ–PUZ	80	−2.74*	−2.84*	0.367**
PBN–PUN	83	−6.24***	−3.66***	1.306***

Notes: Critical values for the DF and ADF tests (Fuller, $\hat{\tau}_\mu$ statistic, table 8.5.2) are:
100 observations: −3.51 (1%), −2.89 (5%), −2.58 (10%)
50 observations: −3.58 (1%), −2.93 (5%), −2.60 (10%)
Critical values for the DW tests (Sargan and Bhargava, table 1) are:
101 observations: 0.376 (1%), 0.259 (5%)
51 observations: 0.705 (1%), 0.493 (5%)
One asterisk indicates rejection at the 10% level; two asterisks indicate rejection at the 5% level; three asterisks indicate rejection at the 1% level.
Caution should be exercised when applying the DW test because the critical values are very sensitive to sample size.
For definitions of the variables see table 1.

win, Grennes, and Wohlgenant, among others.) For that issue to be investigated, freight rates were tested with respect to being cointegrated with the price differential. Because Australia, Canada, and The United States are net exporters of wheat, freight rates regarding the destinations Canada–Australia and U.S.–Australia were not available.[8] Instead, freight rates for the destinations St. Lawrence, Canada–U.K., North Pacific–U.K., U.S. Gulf–U.K., and West Australia–U.K. denoted as T_{CB}, T_{PB}, T_{GB}, and T_{AB}, respectively, were used. They represent quarterly averages reported in the International Wheat Council (*World Wheat Statistics*) and cover the 1957:1–1985:4 time period.

Because stationarity properties of freight rates are of interest in their own right, the order of integration of freight rates was determined for all four destinations and two sample sizes. Results are reported in table 4. The freight rate series between Gulf ports and U.K. (T_{GB}) is $I(0)$ at the 5% level of significance, as all three tests indicate. For the freight rates concerning Canada–U.K. (T_{CB}) and North Pacific–U.K. (T_{PB}), on the other hand, there is a disagreement among the three tests. The general conclusion, however, is that except of T_{CB}, the hypothesis that freight rates are not $I(0)$ is not uniformly rejected as occurred with the prices (compare with tables 1 and 2).

A second important result is the sensitivity of stationarity properties with respect to the time period examined. When we tested for stationarity using 68 observations, (essentially the second half of the sample size) instead of 112 as was the case before, all freight rates were found to be $I(1)$. That result points to the conclusion that the order of integration of freight rates depends on the time period considered.

To see whether transportation costs cause failure of the LOP, we tested for cointegration between freight rates and the price differentials P_{AW2}–P_{UW} and P_{AW2}–P_{CW}. Because the freight rates considered here are unlikely to coincide with the difference in the prices which is due to transportation costs, we tested for cointegration without restricting the cointegration parameter to one.[9] The results (table 5) showed weak evidence of cointegration. Specifically: while the DF test rejects cointegration (all t-values are less than -3.03, the 10% level of significance tabulated value), the ADF test indicates that cointegration exists in three cases at the 10% and one case at the 5% level of significance. With respect to the

[8] As one reviewer suggested, a way to overcome the problem of not having data on transportation costs between Canada and Australia is to take a ratio such as that of Australia-Japan and Canada-Japan freight charges. Tests (not reported here) between this ratio and the associated price differential showed no cointegration.

[9] We also tested for stationarity of the difference between freight rates and price differential. However, no evidence of cointegration was found.

Table 4. Stationarity Tests for Freight Rates

Variable	Number of Observations	DF	ADF	DW
Levels				
T_{AB}	112	−1.40	−1.95	0.084
T_{PB}	112	−3.45**	−2.51	0.313**
T_{CB}	112	−2.99**	−2.72*	0.252
T_{GB}	112	−3.34**	−3.31**	0.315**
T_{AB}	68	−1.49	−2.33	0.097
T_{PB}	68	−2.11	−1.82	0.292
T_{CB}	68	−2.07	−2.08	0.250
T_{GB}	68	−2.19	−2.66	0.271
Differences				
T_{AB}	112	−7.85	−6.05	1.129
T_{PB}	112	−11.85	−8.11	1.955
T_{CB}	112	−10.64	−6.47	1.893
T_{AB}	68	−5.39	−3.98	1.249
T_{PB}	68	−9.22	−4.83	2.282
T_{CB}	68	−8.11	−4.42	2.028
T_{GB}	68	−8.48	−4.16	2.103

Notes: The first subscript identifies source and the second destination: *A* is West Australia; *P*, North Pacific; *C*, St. Lawrence, Canada; *G*, U.S. Gulf Ports; *B*, U.K.).

Levels of significance are given in table 3. For the differences the null hypothesis that freight rates are not $I(1)$ is rejected at the 1% level of significance for all cases and all tests. Because T_{GB} (112 observations) is $I(0)$ we did not test for stationarity of the first difference.

Table 5. Cointegration Tests between Freight Rates and Price Differentials

Variables	Number of Observations	DF	ADF	DW
TAB, (*PAW2*–*PUW*)	112	−1.37	−2.11	0.110
TPB, (*PAW2*–*PUW*)	112	−1.44	−3.08*	0.105
TCB, (*PAW2*–*PUW*)	112	−1.66	−1.83	0.134
TAB, (*PAW2*–*PCW*)	68	−2.84	−3.25**	0.631
TPB, (*PAW2*–*PCW*)	68	−2.83	−2.40	0.630
TCB, (*PAW2*–*PCW*)	68	−2.85	−3.14*	0.637
TGB, (*PAW2*–*PCW*)	68	−2.84	−3.13*	0.642

Notes: Critical values for the cointegration tests (Engle and Granger, p. 269, table 2) are:
DF (100 obs.): −4.07 (1%), −3.37 (5%), −3.03 (10%)
ADF (100 obs.): −3.77 (1%), −3.17 (5%), −2.84 (10%)
DW (100 obs.): 0.511 (1%), 0.386 (5%), 0.322 (10%)
The critical values for DW, 50 observations (Engle and Yoo, p. 158, table 4) are 1.00 (1%), 0.78 (5%), and 0.69 (10%).

DW test, strong cointegration is found in four cases if we use the t-values corresponding to 100 observations, while no cointegration is found if the t-values corresponding to 50 observations are used. The sample consists of 68 observations.

To further investigate the link between transportation costs and failure of the LOP, the order of integration of freight rates as well as the price differential between *PAW2* and *PUW* were examined by considering only the first 62 observations (i.e., 1957:1–1972:2 period; note that results reported in tables 1 and 2 are based on the 1968:3–1986:1 period). Table 6 reports such results, which are in sharp contrast not only with respect to cointegration but also with respect to the order of integration of the individual price series. In particular, the null hypothesis that the price series are not $I(0)$ is not rejected at the 1% level as it was in the case before (table 1). Furthermore, when the price differential *PAW2*–*PUW* is considered, strong evidence of cointegration is found [given that the series are not $I(0)$]. Results regarding the order of integration of freight rates for the truncated sample strongly indicate that freight rates are $I(0)$.

To summarize, results from table 6 show that when freight rates are $I(0)$, the price differential is $I(0)$, which means that nonstationarity of freight rates may have been the main cause of nonstationarity of price differentials. Although this conclusion is not an outcome of a formal test, it does give additional evidence regarding the link between transportation costs and price differentials.

Explanations abound for the failure of the LOP even after taking into consideration transportation costs. For example, a nonstationary tax/subsidy might cause persistent divergence from long-run equilibrium. That may be the case if the tax/subsidy is exogenously determined, i.e., the explosiveness of the tax/subsidy is independent of the explosiveness of the price, hence the rejection of cointegration. Other reasons include heterogeneity of the data. For example, the prices examined may represent products of different quality, as might occur for wool.

Some important implications have emerged, especially with respect to international trade modeling. Consider the case of wheat. The fact that the differences *PAW1*–*PUW*, *PAW1*–*PCW*, and *PUW*–*PCW* are stationary implies that one would prefer *PAW1*, *PCW* or *PUW* over *PAW2*. This result is of particular importance because it directs researchers as to what price should be used and

Table 6. Stationarity Tests for Wheat and Freight Rates—Truncated Sample

Variable	Number of Observations	DF	ADF	DW
PAW2	62	−2.08	−3.19**	0.292
PUW	62	−3.21**	−2.83*	0.462
PAW2–*PUW*	62	−3.55**	−3.88***	0.599**
TAB	62	−5.14***	−3.08**	0.381
TPB	62	−7.01***	−4.66***	0.686**
TCB	62	−5.42***	−3.82***	0.520**
TGB	62	−5.56***	−3.22**	0.480

Notes: Levels of significance are given in table 3. The 62 observations of this test correspond to the 1957:1–1972:2 time period.

because it makes Isard's proposition (i.e., tests of the LOP would be insensitive to the particular selection of price) a testable hypothesis. Another implication is related to what commodities really can be modeled by assuming that a world price measure exists. For example, the results show that none of the wool prices can be used as a world price measure unless prices are decomposed and components which cause rejection of cointegration can be identified. On the other hand, for commodities such as sugar, the LOP assumption can be made rather safely. Finally, the simplicity of this test allows researchers to test the LOP rather than assume its existence before the price to be used is selected.

Concluding Comments

In this study the LOP was examined for seven commodities regarding four countries. The empirical results gave supportive evidence for the LOP for the specific commodities and time periods examined. In particular, the empirical results showed that (*a*) the failure of the LOP as a long-run relationship is a price-specific and time-period-specific problem rather than a general failure, and (*b*) a possible reason for the LOP failure is transportation costs. The results of this study are in agreement with recent studies (e.g., Officer; Protopapadakis and Stoll 1986; Goodwin, Grennes, and Wohlgenant; Goodwin 1990).

To further investigate the validity or denial of the LOP, a wider range of commodities must be examined. Furthermore, when the LOP does not hold, price decomposition could lead into the identification of the causes of the cointegration rejection and hence rejection of the LOP. Such issues await further research.

[Received November 1989; final revision received December 1990.]

References

Ahking, F. W. "Further Results on Long-Run Purchasing Power Parity in the 1920s." *Eur. Econ. Rev.* 34(1990):913–19.

Ambler, S. "Does Money Matter in Canada? Evidence from a Vector Error Correction Model." *Rev. Econ. and Statist.* 71(1989):651–58.

Ardeni, P. G. "Does the Law of One Price Really Hold for Commodity Prices?" *Amer. J. Agr. Econ.* 71(1989): 661–69.

Banerjee, A., J. J. Dolado, D. F. Hendry, and G. W. Smith. "Exploring Equilibrium Relationships in Econometrics through Static Models: Some Monte Carlo Evidence." *Oxford Bull. Econ. and Statist.* 48(1986):253–77.

Campbell, J. Y., and R. J. Shiller. "Cointegration and Tests of Present Value Models." *J. Polit. Econ.* 95 (1987):1061–88.

Coleman, M. "Cointegration-Based Tests of Daily Foreign Exchange Market Efficiency." *Econ. Letters* 32 (1990):53–59.

Corbae, D., and S. Ouliaris. "Cointegration and Tests of Purchasing Power Parity." *Rev. Econ. and Statist.* 70(1988):508–11.

Crouhy-Veyrac, L., M. Crouhy, and J. Melitz. "More About the Law of One Price." *Eur. Econ. Rev.* 18(1982):325–44.

Enders, W. "ARIMA and Cointegration Tests of PPP Under Fixed and Flexible Exchange Rate Regime." *Rev. Econ. and Statist.* 70(1988):504–7.

———. "Unit Roots and the Real Exchange Rate Before World War I: The Case of Britain and the USA." *J. Int. Money and Finan.* 8(1989):59–73.

Engle, R. F., and C. W. Granger. "Co-Integration and Error Correction: Representation, Estimation, and Testing." *Econometrica* 55(1987):251–76.

Engle, R. F., and B. S. Yoo. "Forecasting and Testing in Co-Integrated Systems." *J. Econometrics* 35(1987):143–59.

Fuller, W. A. *Introduction to Statistical Time Series*. New York: John Wiley & Sons, 1976.

Goodwin, B. K. "Empirically Testing the Law of One Price in an International Commodity Market: A Rational Expectations Application to the Natural Rubber Market." *Agr. Econ.* 4(1990):165–77.

Goodwin, B. K., T. Grennes, and M. K. Wohlgenant. "A Revised Test of the Law of One Price Using Rational Price Expectations." *Amer. J. Agr. Econ.* 72(1990):682–93.

——. "Testing the Law of One Price When Trade Takes Time." *J. Int. Money and Finan.* 5(1990):21–40.

Granger, C. W. "Developments in the Study of Cointegrated Economic Variables." *Oxford Bull. Econ. and Statist.* 48(1986):213–28.

——. "Some Properties of Time Series Data and Their Use in Econometric Model Specification." *J. Econometrics* 16(1981):121–30.

Hakkio, G.S., and M. Rush. "Market Efficiency and Cointegration: An Application to the Sterling and Deutschemark Exchange Markets." *J. Inter. Money and Finan.* 8(1989):75–88.

Hall, S. G. "An Application of the Granger & Engle Two-Step Estimation Procedure to United Kingdom Aggregate Wage Data." *Oxford Bull. Econ. and Statist.* 48(1986):229–40.

Hamilton, J. D., and M. A. Flavin. "On the Limitations of Government Borrowing. A Framework for Empirical Testing." *Amer. Econ. Rev.* 76(1986):808–19.

Hendry, D. F. "Econometric Modelling with Cointegrated Variables: An Overview." *Oxford Bull. Econ. and Statist.* 48(1986):201–12.

International Monetary Fund. *International Financial Statistics*. Selected issues.

International Wheat Council. *World Wheat Statistics*. Selected issues.

Isard, P. "How Far Can We Push the 'Law of One Price'?" *Amer. Econ. Rev.* 67(1977):942–48.

Officer, L. "The Law of One Price Cannot Be Rejected: Two Tests Based on Tradable/Nontradable Goods Dichotomy." *J. Macroecon.* 8(1986):159–82.

Perron, P. "The Great Crash, the Oil Price Shock and the Unit Root Hypothesis." *Econometrica* 57(1989):1361–1402.

Phillips, P. C. B., and S. Ouliaris. "Asymptotic Properties of Residual Based Tests for Cointegration." *Econometrica* 58(1990):165–93.

——. "Testing for Cointegration Using Principal Components Methods." *J. Econ. Dynamics and Control* 12(1988):205–30.

Protopapadakis, A. A., and H. R. Stoll. "Spot and Futures Prices and the Law of One Price." *J. Finan.* 38(1983):1431–55.

——. "The Law of One Price in International Commodity Markets: A Reformulation and Some Formal Tests." *J. Int. Money and Finance* 9(1986):335–60.

Richardson, J. D. "Some Empirical Evidence on Commodity Arbitrage and the Law of One Price." *J. Int. Econ.* 8(1978):341–51.

Sargan, J. D., and A. Bhargava. "Testing Residuals from Least Squares Regression for Being Generated by the Gaussian Random Walk." *Econometrica* 51(1983):153–74.

Stock, J. H. "Asymptotic Properties of Least Squares Estimators of Cointegrating Vectors." *Econometrica* 55(1987):1035–56.

Taylor, M. P., and P. C. McMahon. "Long-Run Purchasing Power Parity in the 1920s." *Eur. Econ. Rev.* 32(1988):179–97.

Appendix

PROPOSITION: *Let x_{1t} and x_{2t} denote the (nonstationary) prices of one commodity in countries 1 and 2. Further, let x_{2t}^* be the price in country 2 expressed in country 1's currency. Then, assuming that the LOP holds, testing for stationarity of the difference $x_{1t} - x_{2t}^*$ is a sufficient test for cointegration.*

Proof: Because of the LOP, x_{1t} and x_{2t} form a cointegrated system. Let $x_{1t} = ex_{2t} + \varepsilon_t$ be the cointegration regression where e represents the cointegration parameter. Under the LOP, e is the exchange rate (if it is constant) or the long-run counterpart of the exchange rate (if it is variable). The best (in the sense of least variance and unbiasedness) estimator of e, say $\hat{e}$, is the observed exchange rate. Expressing x_{2t} in terms of country 1's currency as $x_{2t}^* = \hat{e}x_{2t}$, and substituting it in the cointegration regression, yields: $x_{1t} = x_{2t}^* + \varepsilon_t$, which has a cointegration parameter of one. Thus, a stationarity test for ε_t is a sufficient test for cointegration. Q.E.D.