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The law of one price for transitional Ukraine

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Abstract

Five commodity prices in Kiev, Ukraine, are examined for their movements relative to US counterparts during Ukraine's early transition period of 1991–1992. Panel unit root tests are applied to the resulting commodity real exchange rates. Although the law of one price did not hold during the period, the commodity real exchange rates appear to have possessed deterministic trends that were in the direction of closing the initial considerable price gap. © 2001 Elsevier Science B.V. All rights reserved.

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1. Introduction

The law of one price postulates that the prices of similar goods expressed in the same currency will tend to equalize across countries, in the absence of barriers to trade and significant transportation costs. It is the foundation for purchasing power parity (PPP), which is central to many open economy models. Unlike PPP, however, the law of one price (LOP) has received relatively little empirical testing. As Rogoff (1996) summarizes, the available results are not very favorable to the LOP (e.g. Isard, 1977; Knetter, 1989; Engel and Rogers, 1996). In this paper the LOP is examined with a rather unusual data set, five food prices over an 11-month period in Kiev, Ukraine, during the early 1991–1992 period of Ukraine's transition to independence. We compare these prices with the prices of similar goods in the US.¹

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¹One other paper of which we are aware has examined pricing in transitional Ukraine. Conway (1999), using data from a later period than we do (1993–1996), examines price convergence among markets *within* Kiev for three commodities. He reports that price convergence among the markets occurred to varying degrees, but also that 'sizeable and sustained divergences from the law of one price' remained (Conway, 1999, p. 253).

2. Data

The Ukrainian food price data used in this paper are derived from on-site work in 1991 and 1992 by the Project on Economic Reform in Ukraine (PERU, 1992) of Harvard University. We focus on the semi-monthly ruble prices from the Bessarabskii market in Kiev of beef, potatoes, salo (similar to US-style bacon), tomatoes, and sour cream. Prices are per kilogram (per litre for sour cream) for the period June 15, 1991, through April 30, 1992 (observations are centered on the 15th and last day of each month). For comparable foreign food prices, we employ US data from the US Department of Labor, *CPI Detailed Report*, for beef and veal, white potatoes, sliced bacon, tomatoes, and natural yogurt. All are dollar prices after converting quantity units to kilograms or litres. These data are available monthly and are assumed to be centered on the 15th. End-of-month values are then obtained by interpolation. The ruble/dollar exchange rate data are black market buying rates from Goldberg (1992), the IMF (1997), and several *Memoranda* of the PERU. Although Ukraine declared independence in 1991, the ruble continued to be used until late 1992.²

3. Econometric analysis

We analyze the log ratios of the Ukrainian–US prices. These move from an average of -2.501 on June 15, 1991, to an average of -1.691 on April 30, 1992. Since the log ratios are well below 0.000 at all times, it is clear that the Ukrainian prices were relatively quite low and that the LOP did not hold. However, the gap apparently closed. We now evaluate the statistical significance of these movements.

Consider the following equation for the log ratios, p_t :

$$p_t = \alpha + (1 - \beta)p_{t-1} + \gamma t + e_t,$$

where e_t is a random error. The unit-root case is given by $\beta=0$. Unit root processes are not consistent with stable movements around or toward a long-run, LOP equilibrium. Meanwhile, stationary cases result if $0 < \beta < 2$. Stationarity around a zero mean is consistent with the LOP, but seems unlikely in our data. However, an upward linear trend could approximate the adjustment path, starting from below, to a long-run mean consistent with LOP. The partial adjustment model results when β is interpreted as the partial adjustment coefficient, γ equals 0, and $\alpha = \beta p_{LR}$, where p_{LR} is the long-run equilibrium relative price (zero under absolute LOP). Define p_0 as the initial price below zero, and assume $\gamma=0$ and $p_{LR}=0$. Then $p_t = (1 - \beta)^t p_0 + u_t$, where $u_t = \sum (1 - \beta)^i e_{t-i}$, $i = 0, 1, \dots, t-1$, which is equivalent to $u_t = (1 - \beta)u_{t-1} + e_t$ with $u_0 = 0$. Thus, prices follow an upward nonlinear trend with an autoregressive error process. With low β (slow adjustment), the curvature will be quite

²The period for the data is rather short by traditional testing standards. However, we hope that two unusual features of the data will make up for this. Firstly, our data set concerns a type of economy for which there is a relative scarcity of analysis: the transitional one. Secondly, our data set comes from a highly inflationary environment, unlike the environment for most other LOP-test data sets. Perhaps more support could be found for LOP under high inflation, just as for PPP in the same situation (e.g. Frenkel, 1978; McNown and Wallace, 1989). Nevertheless, it is clear we can only examine short-run, not long-run, LOP.

gentle, and the nonlinear trend over a given range can be well approximated by an upward linear trend (a first-degree Taylor expansion around t^* gives $(1 - \beta)^{t^*} p_0 + (1 - \beta)^{t^*} p_0 \ln(1 - \beta)(t - t^*)$ with a remainder involving powers of $\ln(1 - \beta)$ that are quite small with low β).

To determine which of the above processes might be driving the price ratios, we apply a series of unit root tests. The price ratios can be thought of as commodity-specific real exchange rates, and so their analysis is the analog of a PPP analysis using aggregate real exchange rates.

Although unit root tests are typically applied to long periods of data, there is nothing inherent in them that prevents their application to short periods. Indeed, Granger (1993) argues that there is no clear definition of what precisely is meant by ‘the long-run’. He holds that for the applied econometrician neither the macro nor microeconomic distinctions between the short and long run are useful. Nevertheless, small samples may suffer from size distortion and low power. To account for size distortion, we generate small-sample P -values rather than use asymptotic ones. This is done by applying each test procedure to 2000 Gaussian random walks for the sample sizes relevant to the paper. To increase power, we rely on pooled or panel unit root tests. Although seasonal effects could also be present, for the Ukrainian variables these are surely small relative to Ukraine’s high, fluctuating inflation rate at the time.

The panel tests we use are based on the Dickey–Fuller unit root test. One is the t_{δ}^* test of Levin and Lin (1993) (LL), and the other is the modified t -bar test of Im et al. (1997) (IPS). Both the LL and IPS tests generate statistics that asymptotically follow the standard normal distribution. Tests of the LL variety have recently been used in PPP analysis, with results somewhat favorable to PPP (e.g. Frankel and Rose, 1996; MacDonald, 1996; Papell, 1997).

With constants included, both panel procedures test unit root without drift against mean stationary processes. With linear trends also included, both procedures test unit root with drift against linear trend stationary processes. Both tests can also include time means, which would allow for a single nonlinear time trend of order $T - 1$ across all series. Finally, both tests include lagged differences to control for serial correlation. We choose the lag orders with the Ng and Perron (1995) recursive procedure.

The LL test imposes the same rate of mean reversion under the alternative for all individuals in the panel. Consequently, its alternative is that *all* the processes are mean or trend stationary. On the other hand, the IPS test is based on the mean of Dickey–Fuller statistics for each individual, and so its alternative includes the possibility that only some of the processes are stationary. Im et al. (1997) find that their test has better size and power properties than the LL test.

However, neither test is robust to contemporaneous error correlation. Specification of time means is a partial solution in either test, but rules out allowing differential serial correlation patterns among the series in the IPS test.³ To examine these issues under the null of unit root processes, we examined a first-difference VAR with three lags for our five series. In it, the lag coefficients differ quite a bit across the equations, but the contemporaneous error correlations are below values that are problematic in O’Connell’s (1998) results. Thus, it seems preferable not to specify time means for the IPS test.

The results of the panel unit root tests are given in Table 1. With constants only, neither test rejects the null. With constants and linear trends, the IPS test rejects at the 0.07 level, and the LL test at the 0.15 level. Although this LL outcome is a rather weak rejection, when time means instead of linear

³The GLS estimator proposed by O’Connell (1998) to deal with the contemporaneous correlation problem cannot do so either.

Table 1
Panel unit root test results for commodity real exchange rates

Test	Constants	Trends and constants	Time means and constants
LL	−0.008 (0.329)	0.145 (0.151)	−1.859 (0.032)
IPS	0.224 (0.560)	−2.275 (0.072)	

Small-sample *P*-values are in parentheses.

trends are specified, the LL test rejects at the 0.03 level. These results suggest that we can reject the unit root null in favor of either linear trend stationarity or a form of nonlinear trend stationarity estimated by the time means. Although the IPS result with trends does not rule out that some of the relative prices are still nonstationary, it seems unlikely that some would be stationary and others nonstationary.

Table 2 shows estimates from autoregressive specifications of the possible deterministic linear trend coefficients (with lag order determined recursively). The trend coefficients are positive and significant at the 0.05 level for the middle three commodities and at the 0.10 level for the first. The average semi-monthly change in the trend is 0.031 for all five, and 0.043 for the three significant at the 0.05 level.

For an estimate of the nonlinear trends if the time-means specification is the correct one, Table 3 shows estimates of the time means of the commodity real exchange rates. The first column of time means is from a pooled equation without lagged levels. Although this specification suffers from serial

Table 2
Estimated trends for the commodity real exchange rates

	Beef	Potatoes	Saló/bacon	Tomatoes	Sour cream/yogurt
Trend coefficient	0.006	0.032	0.061	0.037	0.017
(<i>t</i> -statistic)	(1.87)	(2.76)	(3.50)	(2.98)	(1.57)

Table 3
Estimated time means

Date	Mean ^a	Mean ^b	(<i>t</i> -stat)	Date	Mean ^a	Mean ^b	(<i>t</i> -stat)
6/15/91	0.000			11/30/91	0.051	0.338	(0.44)
6/30	0.038			12/15	0.285	0.402	(1.12)
7/15	0.136			12/30	0.639	0.530	(2.50)
7/30	0.231			1/15/92	0.993	0.603	(3.13)
8/15	0.296	0.296		1/30	1.087	0.448	(1.43)
8/30	0.263	0.238	(−0.62)	2/15	1.132	0.543	(2.24)
9/15	0.248	0.287	(−0.10)	2/28	1.098	0.538	(2.04)
9/30	0.199	0.253	(−0.46)	3/15	0.930	0.399	(0.82)
10/15	0.093	0.179	(−1.23)	3/30	0.675	0.318	(0.18)
10/30	−0.015	0.173	(−1.29)	4/15	0.730	0.592	(2.43)
11/15	−0.077	0.190	(−1.10)	4/30	0.808	0.510	(1.71)

^a Estimated with no lags.

^b Estimated with four lagged levels.

Table 4

Out-of-sample forecasts and observed values of logs of Ukrainian/US price ratios

	Beef	Potatoes	Salo/bacon	Tomatoes	Sour cream/yogurt
4/30/92, observed	−2.076	−2.182	−1.979	−1.115	−1.102
6/15/92, observed	−1.834		−1.640	−1.136	−1.036
6/15/92, forecast	−1.774		−1.538	−0.452	−0.885
(Standard error)	(0.230)		(0.261)	(0.398)	(0.342)

correlation, the estimates are still unbiased. The second column of time means is from an equation with lagged levels to control for serial correlation. However, it cannot give values for the first few time periods. Its estimates are linked with the column-one value for period five, and the t -statistics are with respect to subsequent changes from this benchmark. These results suggest a statistically significant average semi-monthly increase in real commodity exchange rates of 0.023 (column 2) to 0.037 (column 1), not too dissimilar from the average linear trends previously calculated for the individual series.

Both the linear and time-means trend specifications indicate substantial movement toward absolute LOP. However, the linear trend specification has the more natural interpretation as an approximation to an adjustment path toward the LOP.⁴

The available data from PERU contain for four of the five commodities a final observation 1.5 months beyond the end of the estimation period. We use these to evaluate out-of-sample forecasts for the linear trend models. Table 4 shows that in three of the available four cases the relative prices continued to move upward as forecasted by the models. Furthermore, the actual values for these three commodities are within one standard deviation of the forecasts.

4. Conclusions

Although the data and unit root tests indicate that the law of one price did not hold in Ukraine's early transition period, they do suggest that the five Ukrainian commodity real exchange rates can be modeled as deterministic trend stationary processes. Furthermore, the trends are in the direction of closing the very considerable gap that existed between the Ukrainian and US food prices in 1991–92. Our conclusions thus parallel those of Conway (1999), who found some convergence for prices *within* Kiev. We find that prices in this newly emerging market economy, although not in accord with the law of one price, did begin to conform somewhat more closely to international conditions.

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⁴The observed nonlinear pattern in the time means is unlikely to reflect the nonlinear function we noted earlier for the partial adjustment model.

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