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The law of one price: a test based on prices for selected inputs in New Zealand agriculture

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Abstract

The law of one price (LOOP) is an essential foundation of both the pure theory of trade and monetary theory. Strictly speaking, the law relates to prices of individual commodities. However, empirical tests of LOOP have often relied on aggregated data. In this paper, a model is derived and estimated using price data for 15 selected inputs in New Zealand agriculture. The results offer no support for the LOOP in the short run, and the results for the long run are mixed. It may be inappropriate, therefore, to assume that the LOOP holds generally in modelling exercises, particularly when models are used for policy purposes.

1. Introduction

Both orthodox theory and empirical evidence suggest that domestic prices of tradable goods depend on world prices and the exchange rate. However, there has been considerable debate about whether or not they depend on any other factors. The law of one price (LOOP) asserts that once known exogenous factors such as transport costs, marketing margins, tariffs and the monetary equivalents of non-tariff barriers are taken into account, commodity arbitrage ensure that the price of a perfectly substitutable commodity in one country is equal to the price of the same commodity in all other countries, after adjusting for the exchange rate. The LOOP is an essential element of orthodox pure theory of trade and is also an important foundation of monetary theory. Failure of the LOOP has some serious implications: for example, the 'small country' assumption is untenable, price competition matters, and the exchange rate is available as an instrument of macroeconomic policy.

The weight of empirical evidence points to rejection of the LOOP. Several econometric studies have concluded that the LOOP does not hold for many internationally traded goods (Isard, 1977; Kravis and Lipsey, 1978; Richardson, 1978; Protopapadakis and Stoll, 1983; Ardeni, 1989). (For a detailed survey, see Officer, 1990). These studies have considered a wide range of tradable goods including motor vehicles, consumer durables, and food. Moreover, the results have been similar for a variety of econometric mod-

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elling approaches. However, owing to the difficulty of obtaining suitable data, most recent work has relied on the use of highly aggregated price data: for example, export unit values, wholesale price indices, and consumer price indices. Use of such data has been justified by arguing that the LOOP would hold approximately for price indices if it holds for individual goods. This argument is strictly correct but irrelevant. It is true that if the LOOP holds at the level of individual goods then it will also hold for price indices, but it is not true that the LOOP observed at an aggregate level implies that the LOOP also holds at the level of individual goods. Observation of the LOOP at an aggregate level could be purely coincidental, for example because different countries have used different weighting schemes in aggregating prices of individual goods (Kravis and Lipsey, 1971).

In addition to the problems posed by using aggregate data, some of the econometric evidence is suspect because of inappropriate testing procedures (Ardeni, 1989). An example is Protopapadakis and Stoll (1983) whose regressions displayed highly autocorrelated residuals and low values for the coefficient of determination. This combination of test results is symptomatic of spurious regressions. Furthermore, recent literature on inference from regression emphasizes the need to test the usual assumption that the distribution of the unobservable error term in the regression equation is normally and independently distributed. This is because standard statistical test results may not even be asymptotically valid in the presence of non-normal residuals (see Arnold, 1980; Bera and Jarque, 1982).

The purpose of this paper is twofold. First, we build and estimate a model that uses disaggregated data that have been carefully adjusted to avoid inconsistencies in measurement of prices. Secondly, the testing procedures used recognize the importance of the normality assumption noted above. We believe that our methodology and results are quite robust.

The remainder of this paper is organized as follows. In the next section the model is explained. Then the testing procedures are outlined and test results are discussed. Next, the regression results are analysed. The paper concludes with a summary and suggestions for future research.

Table 1

Results of regression equation $\ln P\text{NZ}_t = \alpha + \beta \ln P\text{US}_t + u_t$

Variable	β	t on β	R^2	DW	Normality χ^2
Urea	0.8497 (0.0626)	13.544	0.85	0.87	4.029
Sulphate of ammonia	0.8297 (0.0510)	16.247	0.90	1.12	3.275
Superphosphate	0.9771 (0.0522)	18.711	0.92	0.65	0.624
Diesel	0.6061 (0.0650)	13.811	0.86	0.38	0.658
Grease	0.4181 (0.0618)	11.828	0.82	0.55	1.478
Oil	0.7099 (0.0630)	11.268	0.81	0.22	0.314
Barbed wire	0.0409 (0.0284)	36.525	0.98	1.32	1.125
24D	0.7693 (0.0548)	14.032	0.87	0.39	2.711
Cement	0.8664 (0.00274)	31.610	0.97	0.72	0.112
Paint	1.1824 (0.0267)	44.220	0.98	1.47	1.635
Fenceposts	0.4909 (0.0471)	10.422	0.78	0.27	3.548
Galvanised pipes	0.8198 (0.0225)	36.360	0.98	1.59	0.782
Mouldboard ploughs	0.7468 (0.0573)	13.033	0.84	1.52	197.80
Tractors 30-39 BHP	0.9917 (0.0542)	18.290	0.92	0.38	5.142
Tractors 50-59 BHP	1.1197 (0.0443)	25.241	0.95	0.46	2.207

Standard errors are given in parentheses.
Sample period 1956-1987.
t critical value = 2.042 ($P < 0.05$).
DW critical value = 1.345 (upper 1%).
Jarque-Bera χ^2 critical value = 5.991 ($P < 0.05$).

2. The model

For the purposes of the research reported here, 15 internationally traded inputs in agriculture were chosen as the focus, with price data being collected for New Zealand and the United States (see Table 1 for the list of inputs). Apart from the obvious fact the we are researchers resident in New Zealand, our interest in agricultural inputs arises from the fact that New Zealand is a small open economy (with exports comprising over 25% of expenditure on GDP and agricultural products comprising about 40% of total exports). A model using prices of inputs in New Zealand agriculture should be favourable to the LOOP.

A generalized version of the LOOP for a given homogeneous product with respect to New Zealand and the United States can be expressed as:

$$P\text{NZ}_t = (P_t^* e\text{NZ/US}_t)^{\beta} \tag{1}$$

where $P\text{NZ}$ is the price in New Zealand expressed in New Zealand currency, P^* is the price in the United States expressed in United States currency both at time period t, $e\text{NZ/US}$ is the exchange rate between the two currencies at time t, and β is a parameter representing the relationship between the two price levels. Perfect commodity arbitrage, perfect competition, zero protection and no transportation or marketing costs for trading between the two countries are assumed. To derive a more satisfactory expression of the LOOP, adjustments should be made to include the cost of import protection, transportation costs and marketing margins involved in the transmission of the product from producer to consumers. Therefore, the United States component is better expressed as:

$$P\text{US}_t = P_t^*(1 + a + d + m) e\text{NZ/US}_t \tag{2}$$

where $P\text{US}_t$ is the adjusted U.S. price of the commodity in terms of a common currency (New Zealand dollar), a is the ad valorem tariff in decimal form, d is the percentage transportation cost in decimal form, and m is the fixed percentage marketing margin in decimal form at time period t. This is similar to the model developed by Hazledine (1980), except for the addition of the marketing margin. The United States dollar price on the right-hand side of Eq. (2) represents the United States price of the commodity measured at the New Zealand border. In general form, we have:

$$P\text{NZ}_t = f(P\text{US}_t)^{\beta} \tag{3}$$

For estimation purposes, we can use:

$$\ln P\text{US}_t = \alpha + \beta \ln P\text{US}_t + u_t \tag{4}$$

where α and β are coefficients. Eq. (4) is an estimable form of Eq. (3) with the addition of a stochastic disturbance term (u_t) and transformation to a logarithmic form. The coefficient on $\ln P\text{US}_t$ can be interpreted as the static long-run price transmission elasticity, that is, it measures the degree to which a change in the United States price (measured at the New Zealand border) flows through to a change in the New Zealand price.

The data selected for this study were annual observations on prices paid by farmers in New Zealand and the United States for 15 internationally traded farm inputs between 1956 and 1987. The data for United States were collected from *Agricultural Prices* (annual), published by the National Agricultural Statistics Service, United States Department of Agriculture. The prime source of data for New Zealand was the *Annual Technical Paper Series (1)* on farm costs and prices published by the Ministry of Agriculture and Fisheries. Supplementary data were obtained from the *Farm Budget Manual*, published by Lincoln University, New Zealand.

Nominal prices needed to be adjusted in several ways to obtain consistent data. These included adjustment for common composition (for example, superphosphate and 24D are sold in different concentrations in the two countries), for common units of measurement, for state or national taxes, and for effective rates of protection. However, the prices of some commodities, such as tractors, were difficult to adjust because they were characterised by many variations (for example, with respect to transmission system, output power, weight, wheel size, engine capacity, fuel type, and accessories). Hence, for analytical convenience, two major categories of tractor, 30–39 BHP [1] and 50–59 BHP, were chosen and other characteristics were ignored. Detailed descriptions of the adjustments are available in Delpachitra (1990).

[1] BHP, British horsepower = 550 lbf ft s^{-1} ≈ 746 W.

Three assumptions were made in the use of data. First, it was assumed that New Zealand is a small country relative to the United States for the 15 commodities considered in the analysis. Thus, New Zealand was assumed to be a price taker. This is consistent with the LOOP. Secondly, it was assumed that protection imposed by New Zealand was equal to the tariff equivalent of the protection for each commodity and that no other barriers to trade existed between the two countries. Thirdly, transportation costs and marketing margins were assumed to be endogenous to the price level but with constant growth rates over time.

It is important to note that commodities in the sample data set included both branded and non-branded products. When there is branding, products are no longer homogeneous, and there is potential for price discrimination and product differentiation. For analytical purposes, paint, 24D, galvanized pipes, mouldboard ploughs and tractors were regarded as branded products, and fertiliser (urea, sulphate of ammonia, superphosphate), fuel and lubrication materials (diesel, oil, grease) and others (barbed wire, cement, and fenceposts) were considered as non-branded products. This classification is somewhat arbitrary because there is some branding even within the class of products that we have defined as non-branded (for example, fuel and lubricants).

Care was taken in this research to ensure that models were statistically robust. Model selection criteria specified by Harvey (1981), Gilbert (1986) and Hendry (1987) were generally used. The aim of model selection was to ensure that the estimated coefficients were econometrically acceptable. Two important points with respect to the model selection process need to be noted. First, the normality of the residual distribution was examined because any non-normality of the distribution biases the standard test results on the coefficients. In cases where outliers were observed, an attempt was made to relate them to the economic changes that occurred during the sample period. Secondly, New Zealand was assumed to be a small country in trading of the commodities under review. This removed the need for tests of exogeneity or validity conditioning.

3. Results

Empirical results for Eq. (4). Sample data were regressed by applying ordinary least squares to Eq. (4) for each of the 15 commodities. The results are set out in Table 1. Although most regression results appeared satisfactory, for all commodities the null hypothesis of no autocorrelation was rejected at $P^* < 0.05$ and was rejected for all but three commodities at $P^* < 0.01$ (paints, galvanised pipes and mouldboard ploughs). Hence, in general, the model appears to have been autocorrelated. Furthermore, the null hypothesis in favour of normality in the distribution of residuals was rejected ($P^* < 0.05$) for mouldboard ploughs. In most cases the estimates of price transmission elasticities were positive and close to 1. However, high R^2 and high t-values might have reflected bias caused by the presence of autocorrelation (and non-normality in the distribution of regression residuals in the case of mouldboard ploughs). Hence, the overall regression results indicated that the initial model needed to be respecified.

An alternative specification. Some empirical literature on pricing behaviour suggests that, in some markets, it takes a certain period of time to adjust to an equilibrium price level after a shock in the market (Stigler and Kindahl, 1970). Reasons include long-term buyer–seller contacts that favour existing buyers rather than new buyers, availability of substitutes, and sellers ability to build up buffer stocks that offset sudden changes in supplies and prices. Based on these findings, Stigler and Kindahl suggested that price changes across markets are correlated. These findings are further empirically supported by Carlton (1986) who found that some commodity prices take a long time to adjust, even between markets within the same country. These findings may also be applicable to the farm inputs considered in this analysis. For instance: demand and supply patterns may vary over time owing to seasonality; some commodities, such as fertilisers and agricultural chemicals are often stockpiled in order to avoid delays in transportation; and farm inputs are sometimes sold under medium or long-term contracts. Furthermore, owing to the physical nature of the commodities, arbitrage is not likely to be instantaneous.

To take some of these characteristics into account, it is necessary to hypothesize an adjustment process. Nerlove's partial adjustment model (Nerlove, 1958) is appropriate in this context. Suppose that there is an equilibrium price level ($P\text{NZ}_t^*$) in the market in New Zealand for a particular commodity. Under a given set of constraints, such as limited storage facilities and seasonality in supply and demand, the equilibrium price level may be hypothesized to be a function of some economic variable such as the United States price at the New Zealand border. The relationship might be expressed as:

$$P_t^{\text{NZ}*} = a + bP\text{US}_t \tag{5}$$

where a and b are parameters. Eq. (5) can be interpreted as a behavioural approximation of $P\text{NZ}_t^*$. Since $P\text{NZ}_t^*$ is not directly observable, the actual change in the price level ($P\text{NZ}_t - P\text{NZ}_{t-1}$) can be assumed to be some proportion of the

Table 2

Results of regression equation $\ln P\text{NZ}_t = \alpha + \gamma \ln P\text{NZ}_{t-1} + \beta \ln P\text{US}_t + e_t$

Variable	β	t	γ	t	R^2	χ^2	Durbin-H
Urea	0.4229	5.02	0.5614	5.97	0.94	2.584	0.0154
	0.0842)		(0.0939)				
Sulphate of ammonia	0.4639	4.75	0.4825	4.18	0.94	2.060	0.6999
	(0.0974)		(0.1152)				
Super-phosphate	0.4158	4.18	0.6131	6.03	0.96	0.494	1.4380
	(0.0993)		(0.1016)				
Diesel	0.2670	4.18	0.7695	11.18	0.97	1.621	0.9360
	(0.0637)		(0.0687)				
Grease	0.2362	3.26	0.6634	5.97	0.92	36.840	0.8067
	(0.0723)		(0.1109)				
Oil	0.1997	3.66	0.8108	11.09	0.96	1.450	1.8797
	(0.0544)		(0.0730)				
Barbed wire	0.4018	3.65	0.6329	5.89	0.99	3.440	0.1223
	(0.1100)		(0.1074)				
24D	0.2476	2.95	0.7121	6.88	0.95	18.630	2.5000
			(0.1034)				
Cement	0.3370	4.97	0.6650	8.13	0.99	0.068	1.3800
	(0.0678)		(0.0817)				
Paint	0.6508	4.49	0.4696	3.71	0.99	11.542	−0.5460
	(0.1447)		(0.1263)				
Fenceposts	0.1014	2.45	0.8771	11.15	0.96	5.350	0.8389
	(0.0412)		(0.0789)				
Galvanised pipes	0.4653	5.27	0.4580	4.12	0.98	0.678	−2.6817
	(0.0881)		(0.1109)				
Mouldboard ploughs	0.0615	1.61	0.9720	19.63	0.99	13.570	−0.9918
	(0.0381)		(0.0494)				
Tractors 30–39 BHP	0.1552	1.48	0.8629	8.32	0.97	6.395	1.4774
	(0.1048)		(0.1037)				
Tractors 50–59 BHP	0.0615	0.41	0.9537	7.33	0.98	1.834	0.7628
	(0.1459)		(0.1299)				

Sample period: 1956–1987.

Standard errors are given in parenthesis. Note that standard errors on β and γ coefficients were computed using Eqs. (9) and (8), respectively. Hence the long-run t-statistics are asymptotic.

Durbin-H critical value = 1.96, $P < 0.05$.

Jarque–Bera χ^2 critical value = 5.991, $P < 0.05$.

t critical value = 2.042, $P < 0.05$.

All the estimations were performed using SHAZAM software.

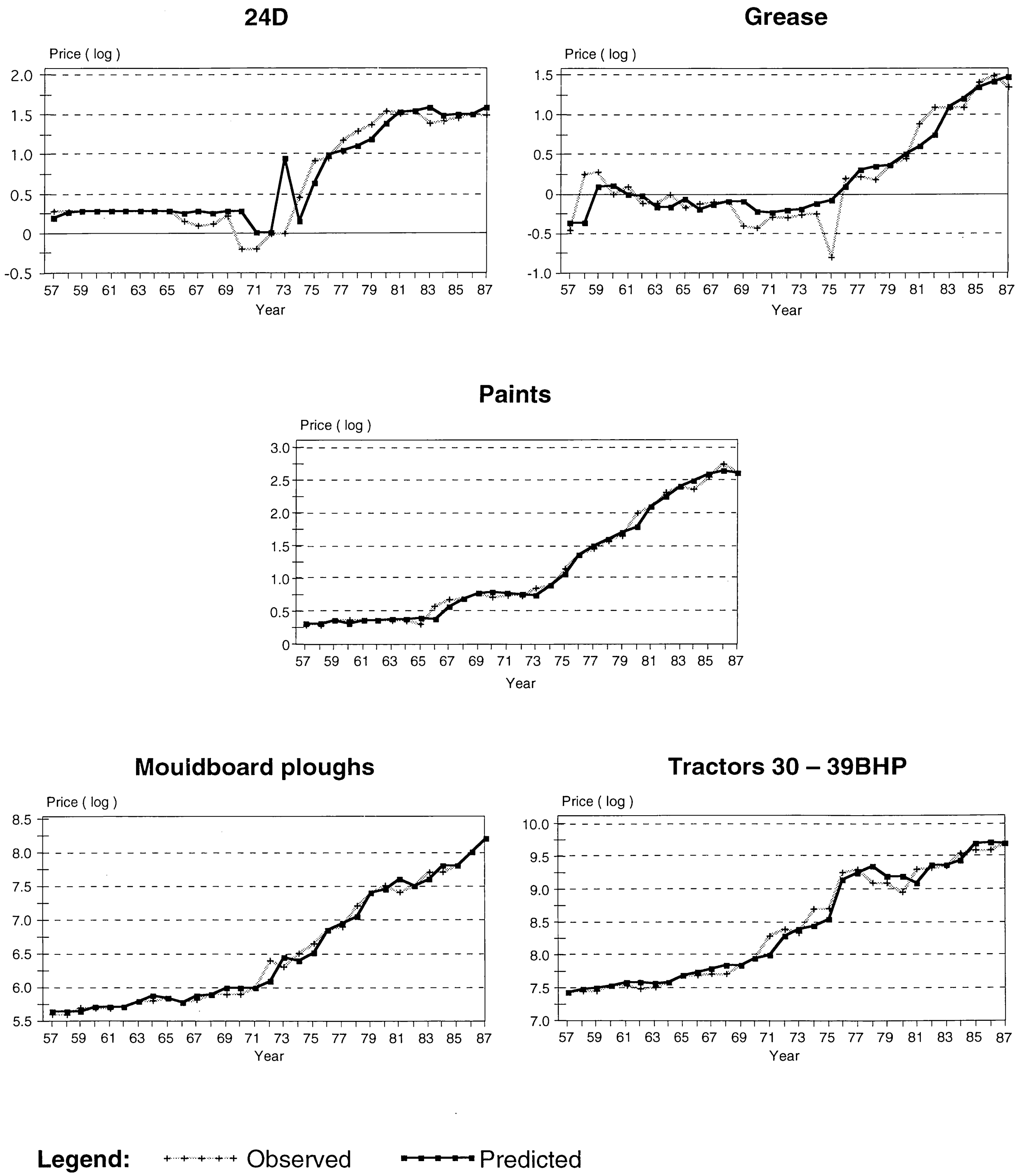

Fig. 1. Non-normality of Eq. (9).

desired change $(P\text{NZ}_t^* - P\text{NZ}_{t-1})$ in each time period. The proportion of desired change actually achieved is $(1-\lambda)$, where λ is a value which lies between 0 and 1. Accordingly, the partial adjustment mechanism can be expressed in two ways:

$$P\text{NZ}_t - P\text{NZ}_{t-1} = (1-\lambda)(P\text{NZ}^* - \lambda P\text{NZ}_{t-1}) + u_t \quad (6)$$

$$P\text{NZ}_t = (1-\lambda)P\text{NZ}_t^* + P\text{NZ}_{t-1} + u_t \quad (7)$$

Eq. (7) implies that the actual price level, $P\text{NZ}_t$, is a weighted average of the desired current level, $P\text{NZ}^*$, previous actual price, $P\text{NZ}_{t-1}$ and the random error effect, u_t. The value of λ provides some useful information on the adjustment mechanism because a situation where λ is equal to 0 implies instantaneous adjustment in prices, and a value closer to 1 implies a very slow adjustment process. The behavioural specification in Eq. (7) can be interpreted in its reduced form as:

$$P\text{NZ}_t = a(1-\lambda) + \lambda P\text{NZ}_{t-1} + b(1-\lambda)P\text{US}_t + u_t \quad (8)$$

Eq. (8) is non-linear in parameters but can be expressed in linear form as:

$$P\text{NZ}_t = \alpha + \delta P\text{NZ}_{t-1} + \beta P\text{US} + u_t \quad (9)$$

where $\alpha = a(1-\lambda)$, $\delta = \lambda$ and $\beta = b(1-\lambda)$, in which β is the short-run price adjustment coefficient, and $\gamma \cong \beta/(1-\delta)$ is the long-run price transmission elasticity. It can be seen that the new specification provides two measurements of price transmission elasticities which are of importance to the analysis. The partial adjustment model can also be used to examine the effects of lagged price levels on price adjustments.

Empirical results for Eq. (9). Eq. (9) was estimated for each commodity using Maximum Likelihood Estimation (MLE) which, in linear regression under normality conditions, coincides with the OLS estimations (see Gujarati, 1988). The derived coefficients were then used to compute long-run price transmission elasticities. Results are summarised in Table 2.

Note that, in these regressions, the Durbin-H statistic was used to test for autocorrelation because of the presence of lagged price as a regressor. According to the Durbin-H statistics, the null hypothesis in favour of no-autocorrelation of residuals was not rejected ($P^* < 0.05$), except for 24D and galvanised pipes. A mixture of results was observed in respect of the normality test. Among the 15 commodities, the null hypothesis of normality in the distribution of residuals was rejected ($P^* < 0.05$) for five commodities (grease, 24D, paints, mouldboard plough, and 30–39 BHP tractors). These commodities are examined below. The positive signs on coefficients (both in the short and long run) were consistent with the theoretical predictions.

To examine the causes of non-normality in residuals of the estimated equations for the above mentioned commodities, predicted and observed changes in price levels (over the previous year) in New Zealand were plotted. The results are presented in Fig. 1. For 24D, significant outliers were observed between 1969 and 1973. This period coincided with the commodity boom that took place in the early 1970s. The reason for the erratic behaviour of price movements during this period may have been the impact of heavy import protection, which increased the ability of New Zealand producers to control the domestic market. For grease, outliers were observed in 1958 and 1976. For paint, outliers were observed during 1968 and 1980. Outliers observed in both 50–59 BHP tractors and mouldboard ploughs (1972 and 1981) are similar in most instances, but for tractors, two additional outliers were observed in 1975 and 1976. Use of dummy variables to treat for non-normality in the distribution of the residuals was rejected because of the substantial variations in the timing of outliers that were observed among the commodities.

The t-statistics in Table 2 indicate that the β coefficient in all regressions was significantly different to 1 ($P^* < 0.05$) for all commodities except 50–59 BHP tractors. Although they are not reported in the table, t-statistics were also computed to test whether or not the β coefficients were significantly different to 1. In no case was the null hypothesis (H_0: $\beta = 1$) rejected ($P^* < 0.05$). Therefore, we concluded that for the commodities considered the LOOP does not hold in the short run. With respect to the γ coefficients, t-statistics in the table indicate that in all cases γ

was significantly different to 0. To test whether or not the γ coefficients were significantly different to 1, t-statistics were also used. They indicated that the null hypothesis H_0: $g = 1$ could not be rejected ($P^* < 0.05$) for nine commodities (urea, sulphate of ammonia, superphosphate, diesel, barbed wire, cement, paints, fenceposts and 50–59 BHP tractors). We concluded therefore that LOOP holds for some commodities in the long run.

4. Summary and Conclusions

The results of this study confirm previous empirical findings that the LOOP does not hold in the short run. Several reasons can be suggested for this result. As noted in the discussion of model specification above, prices of farm inputs are influenced by a variety of factors such as seasonality in demand and supply patterns, the existence of stockpiles that act as a buffer when market conditions change, and the existence of medium and long-term contracts. All of these factors could be expected to dampen price fluctuations in the short run. In contrast to the results for the short-run elasticities, the long-run results were mixed. It is important to not that the LOOP appears to hold for most non-branded commodities included in this research, but only for two of the branded commodities. This finding is quite consistent with economic theory which suggests that branding of products offers the potential for differentiation and price discrimination. Under these circumstances, opportunities for arbitrage are much reduced. There is also an empirical issue in this context. The LOOP assumes homogeneous products, but branding implies heterogeneity. Therefore the LOOP ought not to be assumed to hold generally, even in the long run. This is particularly important to bear in mind when models are used as a basis for policy advice.

Because the LOOP is such an important concept in the pure theory of trade and monetary theory, it will no doubt be the subject of continued empirical study. The research reported in this paper has highlighted the importance of identifying branded and non-branded commodities in such research. Furthermore, given that the LOOP appears to hold in the long run for non-branded commodities, it may be that expectations are important in price adjustment. Goodwin (1990) has provided a useful contribution in this area, but there is scope for further work.

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