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SOME EMPIRICAL EVIDENCE ON COMMODITY ARBITRAGE AND THE LAW OF ONE PRICE

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This paper draws three conclusions from a regression study of disaggregated commodity arbitrage between the U.S. and Canada: (1) Inability to detect commodity arbitrage characterizes a majority of commodity classes, which can potentially be described as nontradeables. (2) Commodity arbitrage is never perfect. (3) When commodity arbitrage is detected, Canadian prices invariably respond as much or more to the exchange rate as they do to U.S. prices.

1. Introduction

Commodity arbitrage has a prominent place in international finance. Many conclusions rest on the existence of at least some buyers and sellers who are willing to contract for the same good in several alternative national markets, and who choose markets according to relative costliness or profitability. Commodity arbitrage contributes to explaining the international transmission of inflation, the insulation properties of variable exchange rates, the exchange-rate consequences of disparate monetary growth rates, 'purchasing power parity' tendencies, and the consequences (if any) of policy-dictated changes in central rates and intervention. Extreme, perfect, and instantaneous commodity arbitrage underlies the view (christened 'global monetarism'[1]) that the boundaries of almost all markets are world-wide, with startling conclusions for the efficacy of monetary and exchange-rate policy when confronted with the implied 'law of one price'. At the other extreme, the transitory absence of commodity arbitrage can contribute to explaining the short-run volatility of exchange rates [Dornbusch (1976a, b, c)], and its permanent absence is the distinguishing characteristic of the class of commodities described as 'nontradables'.

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[1]The christening is usually ascribed to Whitman (1975). The view itself is usually ascribed to Arthur B. Laffer and Robert A. Mundell in numerous papers.

This paper summarizes some of the early results from a study of disaggregated commodity arbitrage between the U.S. and Canada. Its findings are consistent with the following conclusions. (1) Commodity arbitrage takes place, but not significantly for every commodity group. If the size of the Canadian nontradeables sector is measured by the frequency with which acceleration/deceleration of Canadian commodity prices appears unrelated to acceleration/deceleration of U.S. commodity prices and exchange rates, then nontradeables can scarcely be dismissed or ignored. (2) When commodity arbitrage does take place, it is never perfect. Even for comparatively homogeneous commodity groups (4- and 7-digit Standard Industrial Classifications), U.S. commodity-price inflation and exchange-rate changes feed only fractionally into Canadian commodity-price inflation. (3) Whether or not commodity arbitrage takes place, one can only rarely reject the hypothesis that however Canadian commodity prices respond to U.S. commodity prices, they respond symmetrically and comparably to the exchange rate.

Conclusion (3) differentiates this paper from recent work in the same spirit by Isard (1977), Bordo and Choudhri (1977), and Kravis, Lipsey, and Kalter (1977). It also contradicts the implied conclusions of some of the previous literature on disaggregated commodity arbitrage. Both Curtis (1971) and Dunn (1970, 1972) imply that while Canadian commodity prices do respond to U.S. commodity prices in the short run, they do not respond by as much – if at all – to the exchange rate. Rosenberg (1976) finds similar invariance to the exchange rate in his examination of steel-price differentials between the U.S., Europe, and Japan.

Estimation strategy differentiates this paper further from Isard (1977), Bordo and Choudhri (1977), and Kravis, Lipsey, and Kalter (1977). And the focus on disaggregation and attention to econometric properties set the paper apart from other research that has direct or indirect implications for commodity arbitrage [e.g. Kravis and Lipsey (1977a, b), Artus (1976), Isard (1976), Kwack (1976), Fieleke (1976), Taylor, Turnovsky, and Wilson (1972), and the large empirical literature on 'purchasing power parity' exchange rates, summarized by Officer (1976)].

2. Econometric considerations

Given any international commodity arbitrage, there will exist a relation between the Canadian and U.S. (or world) prices of similar commodities. For any particular commodity, the following relationship can be expected to hold as a multiplicative approximation:[2]

$$P_{\mathrm{C}} = \beta_0 E^{\beta_1} P_{\mathrm{US}}^{\beta_2} T^{\beta_3} R^{\beta_4}, \tag{1}$$

[2]The multiplicative character of the approximation implies desirable properties for the error term in the regression equivalent to (1). See footnote 8.

where

P_C = the Canadian dollar price of the product,
E = the Canadian dollar price of a U.S. dollar,
P_{US} = the U.S. dollar price of the product,
T = a measure of transfer costs (transport, insurance, tariffs, etc.),[3]
R = residual reasons why prices might differ between Canada and the U.S., and

$$\beta_0, \beta_1, \beta_2, \beta_3, \beta_4 = \text{parameters.}$$

Perfect commodity arbitrage coupled with perfect substitutability of the Canadian and U.S. products would suggest that

$$\beta_0, \beta_1, \beta_2, \beta_3 = 1,$$

$$\beta_4 = 0,^4$$

and equation (1) would become a statement of the 'law of one price'. Perfect nontradeability of the Canadian and American products – complete absence of commodity arbitrage – would suggest, by contrast, that

$$\beta_1, \beta_2, \beta_3 = 0.^5$$

Having parameterized extreme worlds in this way, it seems natural to contemplate time-series estimation of equation (1) for a sample of particular commodities, omitting R, and treating it as an unobserved random disturbance term (on the validity of this, see more below). For commodities that were very substitutable, and in which commodity arbitrage was very prominent, the estimated elements of the triplet $\{\beta_1, \beta_2, \beta_3\}$ should each lie close to one. For

[3]For a Canadian exportable, $T = 1+$ the costs of transferring a product from Canada to the U.S., expressed as a proportion of the product's U.S. price without these costs. For a Canadian importable, $T = 1+$ the costs of transferring a product from the U.S. to Canada, expressed again as a proportion of the U.S. price.

[4]Perfect commodity arbitrage would swamp all residual reasons why prices might differ, such as price discrimination on the part of producers, which can exist only when commodity arbitrage is imperfect. For a recent empirical attempt to discern price discrimination, see Kravis, Lipsey and Kalter (1977).

[5]As a referee has stressed, this claim is extreme if individual product pricing depends in part on the general price level, as implied by the imitative pricing in 'Scandinavian models,' usually ascribed to Edgren, Faxen, and Odhner (1969), or Aukrust (1970). Inflation rates even for perfect nontradeables could then have elements of parallelism ($\beta_2 > 0$). A pervasive rise in U.S. prices (of both tradeables and nontradeables) would be reflected in Canadian tradeables prices, ceteris paribus. And the rise in Canadian tradeables prices would induce sympathetic rises in Canadian nontradeables prices. For similar reasons, a change in the exchange rate under this assumption could have not only direct consequences for tradeables prices, but also indirect impacts on nontradeables prices ($\beta_1 > 0$). Bordo and Choudhri (1977) do find marked support for the dependence of individual product prices on the general price level.

products which were highly differentiated (by national origin), or which were not traded at all (because T fell outside the domain for which equation (1) was a useful approximation), or for which commodity arbitrage was imperfect, the estimated elements of the triplet $\{\beta_1, \beta_2, \beta_3\}$ should each lie close to zero.[6] If β_1, β_2 and β_3 were comparable in size, then the estimated triplet would provide a rough index of the comparative 'tradeability' or 'nontradeability' of various products. But it is an open question whether β_1, β_2 and β_3 are comparable in size. Much of Curtis' (1971), Dunn's (1970, 1972) and Rosenberg's (1976) work is addressed to explaining why they are not, and much of the empirical work that differentiates 'tariff elasticities' and 'exchange-rate elasticities' from 'price elasticities' [e.g. Wilson and Takacs (1977)] implies the same. On the other hand, traditional accounts of how variable exchange rates insulate a country from world inflation implicitly assume that β_1 is comparable in size to β_2, as does the 'purchasing power parity' (PPP) literature and the 'law of one price'.

Direct time-series estimation of equation (1) would probably be uninteresting, however. It describes a very strong implication of commodity arbitrage for price levels, akin to the absolute version of PPP theory. More probable, and therefore more interesting, is the weaker implication of commodity arbitrage for rates of change in prices, akin to the relative version of PPP theory. Second, straightforward estimation of equation (1) is inadvisable. Because of the well-accepted presence of trend in times-series data on prices and other economic variables, the disturbances in a statistical version of (1) can be expected to show high positive serial correlation. Failing to account for this would make statistical inference misleading. Third, time-series estimation of (1) is virtually impossible, given the usual absence of data on transport costs, to say nothing of the omission of R.

All three of these problems can be alleviated by estimating an equation that is *implied* by equation (1):

$$\Delta_1\Delta_k \ln P_{\mathrm{C}} = \beta_0' + \beta_1\Delta_1\Delta_k \ln E + \beta_2\Delta_1\Delta_k \ln P_{\mathrm{US}} + \varepsilon, \tag{2}$$

where Δ denotes the time difference operator:

$$\Delta_1 X = X^t - X^{t-1}, \qquad \Delta_k X = X^t - X^{t-k},$$

and ε incorporates terms in $\Delta_1\Delta_k \ln T$ and $\Delta_1\Delta_k \ln R$. Variables that are transformed by the operator product $\Delta_1\Delta_k$ appear somewhat formidable at first glance. But they are no more than a monotonic transformation of the period-

[6]Curiously, in this case, so should the coefficients of an inverted equation (1) that related P_{US} to P_{C} and E – even though equation (1)'s *implied* coefficient of P_{C}'s influence on P_{US} would be greater than 1 when β_2 was less than 1. On the reasonableness of keeping P_{C} dependent, as in equation (1), see the discussion of simultaneity below.

to-period rate of acceleration of prices (change in inflation rates) defined over k-period intervals. $\Delta_1\Delta_k \ln P_C \equiv \Delta_1 \ln(P_C^t/P_C^{t-k}) \equiv \Delta_1 \ln[1+$(the proportional rate of change in P_C over the past k periods)].[7] As such, they are interesting in their own right as a better measure of 'innovations' in prices and exchange rates than are inflation rates themselves, which tend to reflect the regular and pervasive inflationary trend of modern institutions.

The first problem – that commodity arbitrage is unlikely to be so strong that it binds price *levels* internationally (as opposed to inflation rates) – can be alleviated by choosing ks in equation (2) that correspond to the period over which measured inflation rates are believed to reflect real inflationary conditions reliably.

But the properties of ε are even more crucial in assigning estimation preference to equation (2) over equation (1), because they alleviate the problems of serial correlation and unobservable explanatory variables. First, ε can be taken to have a zero mean, since any true mean will be captured in the new intercept β_0'. Second, it seems difficult to argue convincingly that ε will *not* have a stable variance over time. ε includes, e.g., $\Delta_1 \ln(T^t/T^{t-k})$, a term related monotonically to acceleration/deceleration of transfer costs. It seems hard to defend any particular change over time in the variance of such acceleration/deceleration, even during the 1970s.[8] Third, and for similar reasons, it seems difficult to argue convincingly that there is any predictable or sizeable correlation between the omitted variables grouped in ε – $\Delta_1\Delta_k \ln T$ and $\Delta_1\Delta_k \ln R$ – and the included explanatory variables $\Delta_1\Delta_k \ln E$ and $\Delta_1\Delta_k \ln P_{US}$. In the absence of such correlation, ordinary-least-squares estimation of β_1 and β_2 in (2) will produce the same coefficients as if the omitted and unobserved variables had been included. Fourth, the suspicion of serial correlation in successive values of ε from equation (2) is not nearly as strong as it is for the comparable disturbances in equation (1)[9] (although more complex forms of serial dependence may have been inadvertently introduced).

One final econometric issue that confronts estimation of both equation (1) and equation (2) is simultaneity. If Canada were as large a country as the U.S., and if the price indexes in equation (2) were aggregate price indexes, then the following argument would seem very persuasive: 'E and P_{US} are determined simultaneously with P_C, and their transformed values cannot be assumed therefore to be independent of the values that ε takes on. When an unaccounted-for shock raises the rate of acceleration of Canadian prices (i.e. when $\varepsilon^t > 0$),

[7]Alternatively as Michael R. Darby has perceived, since $\Delta_1\Delta_k \ln P_C = \Delta_k \ln(P_C^t/P_C^{t-1})$, the variables are closely related to the change over k periods in the monthly inflation rate.

[8]The multiplicative specification of equation (1) buttresses the belief in a stable variance of ε over time in equation (2). Had equation (1) been specified linearly, a stable variance there would have implied progressively smaller transport costs and other elements of ε *relative* to commodity prices. Closer to reality seemed to be the assumption of a stable relationship of ε relative to commodity prices over the period.

[9]This suspicion was supported dramatically in a comparison of Durbin–Watson statistics.

one can expect that shock to feed back positively onto the exchange rate and U.S. prices through commodity arbitrage (i.e. one can expect $\Delta_1\Delta_k \ln E^t$ and and $\Delta_1\Delta_k \ln P^t_{US}$ to be raised from what they would have been otherwise). Causality runs in both directions. The resulting nonindependence of ε from the explanatory variables in E and P_{US} will cause simultaneity bias in the estimated coefficients and frustrate statistical inference.'

But these potential simultaneity problems are judged to be minimal in the actual versions of equation (2) estimated below. First, Canada is not as large a country as the U.S. It is no larger than one-tenth the economic size. Thus unaccounted-for shocks to the acceleration rate of Canadian prices should generally have only very small impacts on the acceleration rate of U.S. prices (or on $\Delta_1\Delta_k \ln P_{US}$). (Pulp and paper seems to be the only industry treated below that might qualify as an exception.) Second, equation (2) is estimated for highly disaggregated commodity groups, none of which looms extraordinarily large in overall Canadian commodity trade. Thus exogenous shocks to the acceleration rate of each particular commodity's price should have only very small impacts on the exchange rate (or on $\Delta_1\Delta_k \ln E$).

3. Empirical results

The empirical findings on disaggregated commodity arbitrage between Canada and the United States are summarized in the middle of table 1. The data base is described on the right. It consists of monthly observations on Canadian and (roughly comparable) U.S. price indexes over the period 1965 through 1974. The parameter k in equation (2) was set equal to 24, so that the results refer to relations discerned between month-to-month changes (Δ_1) in biannual ($\Delta_k = \Delta_{24}$) Canadian and U.S. commodity-price inflation and exchange-rate changes.[10]

Notable first in table 1 is the three-way classification of commodities: the presence of commodity arbitrage could be rejected with 95 percent confidence for the last 13 out of the 22 commodity groups, and in an even more conclusive way for the last 7 out of 22. In most cases, the classifications were not surprising: carbonated beverages, beer, cigarettes, and cement would have been chosen

[10]The particular commodity groups were chosen in such a way as to make them most comparable to Curtis' (1971) commodity groups. When Canadian and U.S. data seemed noncomparable for a particular commodity group, no data was collected, even though the group appeared in Curtis' list. P_C represents Canadian Industry Selling Prices from CANSIM, and were taken from Statistics Canada Publication No. 62–002, *Prices and Price Indexes*. P_{US} represents U.S. Industry Sector Price indexes for industries in which they were available, and were secured from the U.S. Department of Labor, Bureau of Labor Statistics, through the kind cooperation of Margaret S. Stotz, Chief, Division of Industrial Prices and Price Indexes. Otherwise, P_{US} represents U.S. wholesale price indexes, for the industry classifications noted. E represents the Canadian dollar price of U.S. dollars, and was taken from the Bank of Canada *Review*. I am indebted to Elizabeth Ruddick, formerly of the Economic Council of Canada, for conscientious compilation of this data and sound judgment on what belonged

a priori as U.S.–Canadian nontradeables. A case can be made, furthermore, that neither pig iron nor petroleum refining products (largely gasoline) is a tradeable – although their source products, iron ore and crude oil, are. Only several of the other nontradeables classifications are puzzling, and they do tend toward exhibiting the presence of at least *some* commodity arbitrage.

Second, it is notable that the 'law of one price' fails uniformly. The hypothesis of *perfect* commodity arbitrage is rejected with 95 percent confidence for *every* commodity group in table 1. (That is, the hypothesis $\beta_1 = \beta_2 = 1$ is rejected in all cases.) At this level of disaggregation, and for a horizon as long as two years, it seemed likely before the fact that perfect commodity arbitrage could have been accepted in at least some cases. As it turned out, not only are the estimated β_1 and β_2 – the elasticities of Canadian price with respect to the exchange rate and U.S. price – significantly different from 1, but their point estimates are not even close, as a rule. The implications of this finding are important for both analytical and empirical work in international economics. In particular, it may be unwise to invoke disaggregated commodity-price parity conditions, as frequently is done. It may be more descriptive to treat every country's products as differentiated from those of every other country, even when classified under similar headings. If domestic and foreign products are therefore imperfect substitutes, as is assumed in much empirical work in international economics, then analytical work that rests on small-country assumptions becomes suspect, and certain conclusions from the traditional elasticities-absorption approach to international finance are revived.

Third, it is remarkable to be generally *unable* to reject the hypothesis that exchange rates affect Canadian prices in precisely the same way and to the same degree that U.S. prices do. The hypothesis that $\beta_1 = \beta_2$ could be rejected in only three out of 22 cases. All three cases were tradeables (liquor, tractors, and pulp and paper) and, in all three, Canadian prices responded *more* to exchange rates than to U.S. prices! These findings directly contradict those of Curtis (1971) and Dunn (1970, 1972), who suggest explicitly and implicitly that Canadian commodity prices respond less (if at all) to exchange rates than to U.S. prices. The contradiction may be due in part to the measurement of inflation over a longer (two-year) horizon in this paper – neither Curtis nor Dunn denied that their conclusions would be weakened in the long run. Or the contradiction may be illusory, since neither Curtis nor Dunn subjected their conclusion to a formal statistical test.[11]

[11] Nor did either Curtis or Dunn test the null hypothesis that U.S. commodity-price inflation was *fully* transmitted to Canadian commodity prices, even under rigidly fixed exchange rates (the hypothesis that $\beta_2 = 1$). Dunn, and later Isard (1977) and Rosenberg (1976), pre-judge this hypothesis by accepting it. That is the econometric implication of their attempts to explain price *relatives* like P_C/P_{US} (or functions of them) by exchange rates and other variables. Curtis, by contrast, explains changes in Canadian prices by changes in U.S. prices and changes in the exchange rate in a manner akin to this paper. But he ignores the apparent implication of his own results that U.S. prices explain Canadian prices about as poorly as the exchange rate does.

Table 1

Equation (2) regression results: Estimates of parameters, hypothesis-testing, and data sources

Commodity group	Est. β_1[a]	Est. β_2[a]	Calc. t for $\beta_1=\beta_2$[b]	Calc. F[c]	D–W[d]	Canadian SIC code	U.S. SIC code	U.S. WPI code[e]	Observation period
Commodity arbitrage apparently present[f]									
Slaughtering and meat-packing products	0.43	0.52	−0.18	26.04	2.11	1010	2011+2031	02-21	65.1–73.12
	(0.49)	(0.07)							
Animal feeds	0.31	0.16	0.30	7.34	1.34	1230	2042	02-9	67.1–73.12
	(0.50)	(0.04)							
Distilled spirits	0.17	−0.62	3.27	6.09	2.25	1430	2085	02-61-02	67.1–73.12
	(0.14)	(0.18)							
Biscuits, crackers, and cookies	0.38	0.39	−0.03	6.34	2.05	1280	2052	–	67.1–74.6
	(0.27)	(0.12)							
Bakery products	0.2[illegible]	0.18	0.41	5.03	2.23	1290	2051	02-11-01	65.1–73.12
	(0.25)	(0.07)							
Leather tanning products	0.06	0.68	−1.25	12.25	2.20	1720	3111	–	67.1–73.12
	(0.46)	(0.14)							
Pulp and paper	0.74	0.33	2.46	27.97	2.16	2710	2611	09-11	65.1–73.12
	(0.16)	(0.05)							
Agricultural implements and tractors	0.44	0.02	2.79	4.70	1.97	3110	3552	11-11	65.1–73.12
	(0.15)	(0.04)							
Fertilizers	0.19	0.32	−0.36	5.01	1.86	3720	2871	–	67.1–74.6
	(0.37)	(0.10)							
Commodity arbitrage possibly present[f]									
Poultry products	1.16	0.09	0.97	1.76	2.92	1030	2015	–	65.1–74.6
	(1.07)	(0.08)							
Fruits and vegetables	0.18	0.09	0.44	1.03	2.71	1120	2033+2034 +2035+2037	02-4	65.1–73.12
	(0.17)	(0.09)							
Wines and brandy	−0.28	−0.21	−0.27	1.94	1.97	1470	2084	–	65.1–74.6
	(0.25)	(0.12)							
Wooden commercial furniture	−0.06	0.37	−1.55	1.91	2.31	2640 001	2521	12-21	67.1–73.12
	(0.19)	(0.19)							
Gypsum products	0.01	0.11	−0.43	1.94	1.56	3450	3275	–	65.1–74.6
	(0.22)	(0.06)							
Ready mixed concrete	0.04	−0.31	1.60	1.95	1.94	3480	3273	–	65.1–74.6
	(0.19)	(0.17)							

Table 1 (*continued*)

Commodity group	Est. β_1[a]	Est. β_2[a]	Calc. t for $\beta_1=\beta_2$[b]	Calc. F[c]	D–W[d]	Canadian SIC code	U.S. SIC code	U.S. WPI code[e]	Observation period
No significant evidence of commodity arbitrage[f]									
Chewing gum	0.13 (0.19)	−0.04 (0.05)	0.92	0.64	1.86	1310 001	2073	–	65.1–74.6
Nonalcoholic beverages	−0.10 (0.21)	−0.08 (0.19)	−0.06	0.19	2.25	1410	2086	02-62	67.1–73.12
Brewing products	0.03 (0.36)	0.23 (0.39)	−0.37	0.18	1.36	1450	2082	–	65.1–74.6
Cigarettes	0.04 (0.32)	0.01 (0.10)	0.12	0.01	2.15	1530 010	2111	–	65.1–74.6
Malleable pig iron	−0.23 (0.50)	0.19 (0.21)	−0.79	0.55	2.21	2910 016	3312	–	65.1–74.6
Cement	−0.16 (0.18)	0.00 (0.06)	−0.89	0.42	1.99	3410	3241	13-22-01-31	67.1–73.12
Petroleum refining products	0.21 (0.24)	−0.13 (0.12)	1.17	0.79	1.84	3651	2911	–	67.1–72.12

[a]Standard errors in parentheses. Estimated intercepts (β'_0) insignificantly different from zero with 95 percent confidence in 21 out of 22 cases.

[b]Calculated t on the hypothesis $\beta_1 = \beta_2$. High values of t suggest rejection of the hypothesis. Critical values of t for most of the commodity groups are 0.68 (50 percent significance), 1.66 (10 percent significance), 1.99 (5 percent significance), and 2.64 (1 percent significance), all for a two-tail test.

[c]Calculated F on the hypothesis $\beta_1 = \beta_2 = 0$. High values of F suggest rejection of the hypothesis. Critical values of F for most of the commodity groups are 3.10 (5 percent significance) and 4.85 (1 percent significance).

[d]Durbin–Watson statistic.

[e]When applicable (see footnote 10 of the text).

[f]Classification based on F test of the hypothesis in note c above, $\beta_1 = \beta_2 = 0$.

It is worth observing finally that this third conclusion is not revealed by casual inspection of the first two columns of table 1 – an inspection that leads one to reflect on how much less precise (as measured by standard error) the coefficient on the exchange rate is than the coefficient on U.S. prices. In this matter, precision and '*t*-values' are red herrings. They are only a part of the answer to the key question: whether or not $\beta_1 = \beta_2$. By not formally testing that hypothesis (nor allowing others to do so, as would have been possible had the covariance of the estimates of β_1 and β_2 been reported), earlier work appears to have been misleading. Whatever commodity arbitrage takes place, there appears to be little evidence that it responds differently to U.S. prices than to exchange rates – at least when measured over a two year period.[12]

[12]Supporting suggestive evidence from table 1 is that the estimated coefficient on the exchange rate is greater than that on U.S. prices in exactly half of the cases, and less in the other half. This is also true within the tradeables group. Furthermore, the estimated β_1 is inexplicably negative 5 times, one time *less* than the estimated β_2. Of these negative β_1 estimates, none are significantly different from zero with 95 percent confidence, whereas one of the negative estimated β_2's is, and two more come close.

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