How Wide Is the Border?

By Charles Engel and John H. Rogers *

We use CPI data for U.S. and Canadian cities for 14 categories of consumer prices to examine the nature of the deviations from the law of one price. The distance between cities explains a significant amount of the variation in the prices of similar goods in different cities. But the variation of the price is much higher for two cities located in different countries than for two equidistant cities in the same country. We explore some of the reasons for this finding. Sticky nominal prices appear to be one explanation but probably do not explain most of the border effect. (*JEL* F40, F41)

The failure of the law of one price in international trade has been widely documented (see Peter Isard [1977] for an early example). It should be no surprise that similar goods sold in different locations have different prices. Indeed, Gerard Debreu (1959 pp. 29–30) in *Theory of Value* defines goods to be different if they are not sold in the same place: ''Finally wheat available in Minneapolis and wheat available in Chicago play also entirely different economic roles for a flour mill which is to use them. Again, a good at a certain location and the same good at a different location are *different* economic objects, and the specification of the location at which it will be available is essential.'' Only when costs are borne to transport wheat from Chicago to Minneapolis will the miller in Minneapolis consider the Chicago wheat equivalent to the Minneapolis wheat. But can the international failure of the law of one price be attributed entirely to this segmentation of markets by physical distance, or are there other factors, such as nominal price-stickiness, that help to explain the failure?

Recent evidence suggests that not only are failures of the law of one price significant, but they play a dominant role in the behavior of real exchange rates. Engel (1993, 1995) and Rogers and Michael Jenkins (1995) examine the time-series behavior of prices of goods across and within countries. They find that the movement of prices of similar goods across borders accounts for much of the motion in real exchange rates. The variation in these prices appears to be far more significant in explaining real exchange rates than are movements in relative prices of different goods within a country's borders (such as nontraded to traded goods prices.)

We examine the importance of distance between locations where goods are sold and the presence of national borders separating locations in determining the degree of the failure of the law of one price. We employ consumer price data disaggregated into 14 categories of goods. We make use of data available for nine Canadian cities and 14 cities in the United States. The basic hypothesis is that the volatility of the price of similar goods between cities should be positively related to the distance between those cities; but holding distance constant, volatility should be higher between two cities separated by the national border.

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Our basic empirical results show that both distance and the border are significant in explaining price dispersion across locations. We provide a measure of how important the border is relative to distance—the "width" of the border. While distance is an economically significant determinant of price dispersion, the effect of the border relative to distance is extremely large. We explore some of the possible reasons why the border is so important, such as nominal price stickiness, integration of labor markets and trade barriers. Nominal price stickiness appears to account for a large portion of the border effect, but most of the effect is left unexplained.

I. Price Dispersion among Locations

The failure of prices of similar goods to equalize between sites is a sign that the markets are not completely integrated. There are several notions of market integration in the literature. It is helpful to enunciate a simple general framework that highlights the roles of distance and the border in determining price variation between locations.

Consider all final goods sold to consumers to be nontraded. (Kalyan K. Sanyal and Ronald W. Jones [1982] analyze a competitive model with this assumption.) Even the prices of goods that are normally classified as tradable, such as nonperishable commodities, must reflect costs of marketing and distribution, which are nontraded services. On the other hand, all goods contain a tradable intermediate component. If the final product is sold by a profit-maximizing monopolist in each location j, the price of good i is determined by

$$(1) \qquad p_j^i = \beta_j^i \alpha_j^i (w_j^i)^{\gamma_i} (q_j^i)^{1-\gamma_i}.$$

With a Cobb-Douglas technology, γ_i is the share of the nontraded service in final output. The price of the nontraded service, w_j^i, and the price of the traded intermediate input, q_j^i, are determined in competitive markets. The total productivity of the final-goods sector is measured by α_j^i. The markup over costs, β_j^i, is inversely related to the elasticity of demand.[1]

Geographical separation of markets provides one reason that the price of similar goods might vary across locations. Recent work in international trade, spearheaded by Paul Krugman (1991) and including empirical work by Jeffrey Frankel et al. (1995) and John McCallum (1995),[2] suggests that much of the pattern of international trade can be explained by geographical considerations. Countries are more likely to trade with neighbors because transportation costs are lower. Transportation costs may also be an explanation for the failure of the law of one price (as in Bernard Dumas [1992]). In equation (1), q_j^i may vary across locations if there are costs of transporting the tradable good. With the "iceberg" transportation costs of Krugman and others, price q_j^i is not necessarily equalized with the price in location k, q_k^i. The relative price could fluctuate in a range, $1/d_i \le q_j^i/q_k^i \le d_i$. The transportation cost, d_i, should depend positively on the distance between locations, so that the range of variation in q_j^i/q_k^i depends on that distance.[3]

It is also possible that places that are farther apart would have less similar cost structures, so that w_j^i/w_k^i and α_j^i/α_k^i might also vary more between more distant locations. From equation (1), these locations would have greater price dispersion.

However, we also entertain the possibility that price variation of similar goods over time might be higher if the cities lie across national borders, holding distance constant. The recent literature on pricing to market (e.g., Rudiger Dornbusch, 1987; Krugman, 1987; Avinash Dixit, 1989; Robert C. Feenstra, 1989; Kenneth A. Froot and Paul D. Klemperer, 1989; Michael N. Knetter, 1989, 1993; Kenneth Kasa, 1992) has examined markets that are segmented by borders.

There are a few reasons why the border might matter. Much of the pricing-to-market literature has emphasized that the markup, β_j^i, may be different across locations, and may

[1] If ε is the elasticity, $\beta = \varepsilon/(\varepsilon - 1)$.

[2] The McCallum paper is complementary to this one, in that it uses data from the states in the United States and provinces in Canada to measure the effects on the volume of trade of distance and crossing the national border.

[3] In the iceberg model, a fraction $1 - (1/d_i)$ of the good melts.

TABLE 1—CATEGORIES OF GOODS IN DISAGGREGATED CONSUMER PRICE INDEXES AND CITIES USED

Good	United States	Canada
1	Food at home	Food purchased from stores
2	Food away from home	Food purchased—restaurants
3	Alcoholic beverages	Alcoholic beverages
4	Shelter	Shelter − 0.2135 (water, fuel, and electricity)
5	Fuel and other utilities	Water, fuel, and electricity
6	Household furnishings and operations	Housing excluding shelter
7	Men's and boy's apparel	0.8058 (Men's wear) + 0.1942 (boy's wear)
8	Women's and girl's apparel	0.8355 (Women's wear) + 0.1645 (girl's wear)
9	Footwear	Footwear
10	Private transportation	Private transportation
11	Public transportation	Public transportation
12	Medical care	Health care
13	Personal care	Personal care
14	Entertainment	0.8567 (Recreation) + 0.1433 (reading material)

Note: The cities included are: Baltimore, Boston, Chicago, Dallas, Detroit, Houston, Los Angeles, Miami, New York, Philadelphia, Pittsburgh, San Francisco, St. Louis, and Washington DC; Calgary, Edmonton, Montreal, Ottawa, Quebec, Regina, Toronto, Vancouver, and Winnipeg.

vary with exchange rate changes. Alternatively, the markets for the nontraded marketing service might be more highly integrated on a national basis, so that w_j^i is more similar between two sites within a country than in two places separated by a border. These marketing services are likely to be highly labor-intensive. To the extent that the two national labor markets are more separated than are local labor markets within a country, there would be more variation in cross-border prices than in within-country prices. There might also be direct costs to crossing borders because of tariffs and other trade restrictions. In addition, there may be more homogeneity to the relative productivity shocks, α_j^i/α_k^i, for city pairs within the same country than for cross-border city pairs, so that, from equation (1), cross-border pairs have more price volatility.

An important reason why the border matters is unrelated to equation (1): the price of a consumer good might be sticky in terms of the currency of the country in which the good is sold. Goods sold in the United States might have sticky prices in U.S. dollar terms, and goods sold in Canada might have sticky prices in Canadian dollar terms. The nominal exchange rate is, in fact, highly variable. In this case, the cross-border prices would fluctuate along with the exchange rate, but the within-country prices would be fairly stable. Price stickiness may be dependent upon market segmentation. It would be easier for a producer in one location to resist attempts to undercut his fixed nominal price if markets were separated.

The sticky-price explanation is a natural one that has been addressed in previous literature. Our test is in part inspired by Michael Mussa (1986), who noted that the variance of the real exchange rate based on all goods in the consumer price index is larger for Toronto versus Chicago, Vancouver versus Chicago, Toronto versus Los Angeles, and Vancouver versus Los Angeles than it is for Toronto versus Vancouver and Chicago versus Los Angeles when there are floating exchange rates between the United States and Canada. He attributes this pattern to sticky prices or, in his terms, nominal exchange-regime nonneutrality. Within the recent literature on pricing to market, Richard C. Marston (1990) and Alberto Giovannini (1988) specifically consider the role of nominal price stickiness.

II. Distance and the Border

A. *The Regressions*

We use consumer price data from 23 North American cities for 14 disaggregated consumer price indexes. The data cover the period from June 1978 to December 1994. Table 1 lists the goods and cities in our study. The Data Appendix provides more detail on the construction of the price data.

For our purposes, it is natural to choose the United States and Canada as the countries to study. First, the countries share a border. Were it not for the country borders, one would expect more trade to occur between Toronto and New York than between New York and Los Angeles. Indeed, there are no other examples of adjacent market economies that are as large in area (so that there can be significant distances between major cities within a country). Also, trade has been relatively free between the two countries. If the border matters, it is unlikely that it matters because of trade restrictions. The facts that both countries are mostly English-speaking and have similar cultural and political traditions suggest that there is likely to be more cross-border labor migration than between most countries.

We hypothesize that the volatility of the prices of similar goods sold in different locations is related to the distance between the locations and other explanatory variables, including a dummy variable for whether the cities are in different countries.

Let $P^i_{j,k}$ be the log of the price of good i in location j relative to the price of good i in location k. (All prices are converted into U.S. dollars using a monthly average exchange rate before taking relative prices.) We take the difference in the log of the relative price between time t and $t - 2$ as our measure of $P^i_{j,k}$. We take the two-month difference because, for some of our U.S. cities, the price data are only reported every other month. We calculate its volatility as the standard deviation.[4]

We also consider a filtered measure of $P^i_{j,k}$. We regress the log of the relative price on 12 seasonal dummies and six monthly lags.[5] We then take the two-month-ahead in-sample forecast error from this regression as our measure of $P^i_{j,k}$. The in-sample forecast errors cover the period from February 1979 to December 1994. Qualitatively, the results were very similar, irrespective of our measure of prices. We report regressions for the two-month difference of the logs because these numbers are easily reproducible.

For each good i, there are 228 city pairs for which we have observations.[6] For each city pair, we calculate our measure of volatility using the time series on relative prices. Then, we conduct our analysis based on the cross section of 228 volatility measures.

Table 2 reports selected summary statistics. For each of the 14 goods, we report the average standard deviation for pairs of cities that are (a) both in the United States, (b) both in Canada, and (c) one in each country. Table 2 also reports the average distance between those cities. The table reveals that the volatility of prices between U.S. city pairs is generally slightly higher than that between Canadian city pairs, but cross-border city pairs have much higher volatility. However, cross-border city pairs are farther apart, on average, as well.

These generalizations from Table 2 do not apply to goods 7, 8, and 9: men's and boys' apparel, women's and girls' apparel, and footwear. For these goods, the variance of prices across U.S. cities is substantial. In fact, on average, it is greater for the U.S. city pairs than for the cross-border city pairs, and far greater than for the Canadian city pairs. These are the only three goods that exhibit this pattern.

The apparel goods are different from the other goods in several respects: (i) This category of goods probably has some of the most product differentiation. (ii) The prices of these goods are very seasonal.[7] (iii) Compared to other goods, a large fraction of clothing is imported from outside of United States and Canada.

[4] We also performed all of our tests using the spread between the 10th and 90th percentiles as the measure of volatility. Our results were essentially identical to the results reported here.

[5] In the case of the data that are bimonthly, we regress the log of the relative price on three bimonthly lags and six seasonal dummies.

[6] We do not attempt to match the U.S. cities whose data is reported in odd-numbered months with the even-month cities.

[7] However, we note that the apparel commodities show the same pattern of volatility when we use the filtered data, which presumably take out seasonals.

TABLE 2—AVERAGE PRICE VOLATILITY

	City pairs		
Good	U.S.–U.S.	Canada–Canada	U.S.–Canada
1	0.0139	0.0198	0.0247
2	0.0130	0.0100	0.0214
3	0.0185	0.0149	0.0271
4	0.0217	0.0085	0.0250
5	0.0486	0.0279	0.0498
6	0.0203	0.0097	0.0236
7	0.0483	0.0167	0.0461
8	0.0880	0.0178	0.0813
9	0.0618	0.0192	0.0505
10	0.0111	0.0186	0.0260
11	0.0443	0.0240	0.0628
12	0.0133	0.0190	0.0259
13	0.0258	0.0143	0.0271
14	0.0203	0.0083	0.0232
1–14	0.0321	0.0163	0.0367
Distance (miles):	1,024 (66 pairs)	1,124 (36 pairs)	1,346 (126 pairs)

Notes: Entries give the mean value of price volatility across all intercity combinations within the United States, within Canada, and across the U.S.–Canadian border, respectively. The measure of volatility is the standard deviation of the relative price series. Prices are measured as two-month differences. The average distance between cities is given in the final row. The sample period is September 1978–December 1994.

Our regressions attempt to explain $V(P^i_{j,k})$, the volatility of $P^i_{j,k}$. We estimate

$$(2) \qquad V(P^i_{j,k}) = \beta^i_1 r_{j,k} + \beta^i_2 B_{j,k} + \sum_{m=1}^{n} \gamma^i_m D_m + u_{j,k}$$

where $r_{j,k}$ is the log of the distance between locations. As in the gravity model of trade, we posit a concave relationship between relative-price volatility and distance. $B_{j,k}$ is a dummy variable for whether locations j and k are in different countries. For reasons we have explained, we expect the coefficient on this variable to be positive. The regression error is denoted as $u_{j,k}$. Note this is a cross-section regression.

We also include a dummy variable in equation (2) for each city in our sample, D_m. That is, for city pair (j, k) the dummy variables for city j and city k take on values of 1. There are a few reasons why we allow the level of the standard deviation to vary from city to city. First, there may be idiosyncratic measurement error or seasonalities in some cities that make their prices more volatile on average. Second, for the cities that report prices only bimonthly, there may be additional volatility that is introduced by measurement error from the less frequent observation of prices. Third, as Table 2 indicates, there seems to be somewhat higher average volatility for U.S. cities than for Canadian cities. This may be because the United States is a more heterogeneous country. Either labor markets or goods markets may be less integrated, so there can be greater discrepancies in prices between locations. Alternatively, there may be differences in methodologies for recording prices that lead to greater discrepancies in prices between locations in one country compared to the other.[8]

Table 3 reports our regressions for each of the 14 goods. We find strong evidence that distance is helpful in explaining price dispersion

[8] We could impose the restriction that the coefficient on the dummy for all U.S. cities be equal, and that it be equal for all Canadian cities. In all of the regressions we report here, that restriction is strongly rejected.

TABLE 3—REGRESSIONS RELATING PRICE VOLATILITY TO DISTANCE AND THE BORDER

Good	Log distance	Border	Adjusted R^2
1	4.95	7.50	0.94
	(2.32)	(0.18)	
2	1.84	9.71	0.97
	(0.89)	(0.11)	
3	3.50	9.98	0.93
	(2.80)	(0.22)	
4	8.37	9.42	0.93
	(1.78)	(0.21)	
5	35.7	10.5	0.81
	(6.88)	(0.74)	
6	1.11	8.26	0.97
	(0.97)	(0.12)	
7	10.5	12.9	0.96
	(2.79)	(0.34)	
8	28.1	26.4	0.93
	(7.34)	(0.89)	
9	7.74	9.20	0.97
	(3.23)	(0.36)	
10	9.80	10.8	0.95
	(2.19)	(0.20)	
11	32.9	27.3	0.87
	(7.70)	(0.95)	
12	−1.25	9.66	0.97
	(2.23)	(0.23)	
13	0.02	6.70	0.94
	(1.67)	(0.18)	
14	5.08	8.58	0.97
	(1.13)	(0.14)	
1–14	10.6	11.9	0.77
	(3.25)	(0.42)	

Notes: All regressions contain as explanatory variables dummies for each of the 23 individual cities, in addition to the variables listed in the cell. Heteroscedasticity-consistent standard errors (Halbert White, 1980) are reported in parentheses. Coefficients and standard errors on log distance are multiplied by 10^4, while those for "border" are multiplied by 10^3. The dependent variable is the standard deviation of the two-month difference in the relative price. Standard deviations are computed over the sample period from September 1978 to December 1994. There are 228 observations in each regression.

across cities. The coefficient on the log of distance is positive for 13 of the 14 goods, and it is significant at the 5-percent level in ten of the regressions.[9] In the one case in which the sign is wrong, the coefficient is not significantly different from zero. In most of the cases, the t statistics are very large.

The coefficients on the dummy variable for the border are of the hypothesized sign and highly significant for all 14 of the goods. The interpretation of the coefficient on the border dummy in this regression is the difference between the average standard deviation of prices for city pairs that lie across the border less the average for pairs that lie within one of the two countries, taking into account the effect of distance.

We note that the model works well even for the apparel commodities. The excess volatility for U.S. apparel derives from a few cities, but with city dummy variables included, distance and the border still have significant explanatory power.

We test for the restrictions that the coefficients on distance are the same in all regressions and the coefficients on the border dummy are the same in all regressions. The test statistics (not reported) are large, and the restrictions are very strongly rejected. Nonetheless, we report the results for the regressions pooling the data across all goods. Because we allow a separate intercept term for each good and for all but one city, the coefficients reported for distance and the border dummy in the pooled regression are simply the average of the coefficients across the 14 goods. Thus, the pooled regression provides a useful summary of the relationship between price dispersion and the explanatory variables. The last row of Table 3 reports the pooled results for all goods. We find that the coefficients on distance and the border dummy are highly significant and of the hypothesized sign.

The results using the filtered measure for prices are recorded as specification 1 in Table 4 and are very similar to those for the two-month differences. Distance has a positive effect on price dispersion in all regressions and is significant for eight of the 14 goods. The coefficient on the border dummy is positive and significant in all cases. If we restrict our tests to just those cities for which we have monthly data, our results are virtually unchanged qualitatively. (These results are not reported.)

[9] We calculated bootstrapped distributions for the t statistics in the first line of Table 3. The inference from the bootstrapped distributions is approximately the same as from the t distribution. Details are in an econometric appendix available from the authors upon request.

TABLE 4—ALTERNATIVE SPECIFICATIONS OF PRICE VOLATILITY REGRESSIONS

	Specification 1			Specification 2			
Good	Log distance	Border	Adjusted R^2	Distance	Distance squared	Border	Adjusted R^2
1	4.32	6.61	0.93	2.08	−6.53	7.53	0.93
	(1.92)	(0.17)		(0.92)	(2.93)	(0.18)	
2	2.26	9.81	0.97	1.00	−3.57	9.72	0.97
	(0.84)	(0.12)		(0.42)	(1.69)	(0.11)	
3	2.34	9.94	0.93	2.21	−7.88	10.0	0.93
	(2.79)	(0.21)		(1.04)	(3.34)	(0.22)	
4	7.00	9.96	0.95	3.56	−11.1	9.45	0.94
	(1.58)	(0.18)		(0.76)	(2.78)	(0.20)	
5	28.7	7.48	0.78	11.7	−33.8	10.7	0.78
	(5.29)	(0.59)		(3.62)	(13.0)	(0.73)	
6	1.21	8.90	0.97	0.48	−1.54	8.27	0.97
	(0.95)	(0.12)		(0.43)	(1.60)	(0.12)	
7	2.40	10.8	0.96	4.20	−13.1	13.0	0.96
	(2.40)	(0.32)		(1.07)	(3.86)	(0.34)	
8	12.2	17.0	0.97	8.76	−24.5	26.6	0.93
	(3.24)	(0.47)		(2.95)	(10.6)	(0.88)	
9	4.98	9.72	0.97	4.04	−12.9	9.19	0.97
	(3.04)	(0.33)		(1.30)	(4.62)	(0.35)	
10	9.04	11.0	0.96	4.17	−13.4	10.9	0.95
	(1.97)	(0.18)		(0.82)	(2.71)	(0.20)	
11	22.2	24.2	0.98	7.97	−13.6	27.2	0.88
	(4.91)	(0.65)		(2.91)	(13.0)	(0.93)	
12	0.25	8.51	0.98	−0.42	0.75	9.68	0.97
	(2.05)	(0.19)		(0.98)	(3.25)	(0.23)	
13	1.31	7.02	0.93	0.78	−3.23	6.69	0.94
	(1.77)	(0.19)		(0.80)	(2.76)	(0.18)	
14	3.13	9.75	0.98	2.43	−7.69	8.58	0.97
	(0.94)	(0.11)		(0.39)	(1.34)	(0.13)	
1–14	7.24	10.8	0.77	3.79	−10.9	12.0	0.77
	(2.73)	(0.35)		(1.36)	(4.68)	(0.42)	

Notes: All regressions contain as explanatory variables a dummy for each of the 23 individual cities, in addition to the variables listed in the cell. Heteroscedasticity-consistent standard errors (White, 1980) are reported in parentheses. Coefficients and standard errors on log distance, border, distance, and distance squared are multiplied by 10^4, 10^3, 10^6, and 10^{10}, respectively. In specification 1, the dependent variable is the standard deviation of the two-month-ahead forecast error from the filtered relative price. In specification 2, the dependent variable is the standard deviation of the two-month difference in the relative price. Standard deviations are computed over the sample period from September 1978 to December 1994. There are 228 observations in each regression.

Regression results when the distance function is quadratic, rather than logarithmic, are reported as specification 2 in Table 4. This specification is interesting because it allows a test for our assumption of a concave distance relationship. In fact, we find that distance has a positive effect on price variability in 13 of the 14 regressions and is significant at the 5-percent level in 11 of those regressions. Furthermore, in all 13 regressions where distance has a positive effect, the square of distance has a negative effect. It is significantly negative for the 11 goods that have a significantly positive distance effect. This is what we would expect if the distance relationship were concave. Once again, in this specification, the border dummy is positive and significant in all cases.

Although we report White's (1980) heteroscedasticity-consistent standard errors, we also specifically allow for the possibility that the variance of the error term might be greater for more distant cities. The first specification in Table 5 reports results when the left- and right-hand-side variables are all deflated by the log of distance, so that the standard deviation of the regression error is modeled as being

TABLE 5—ASSESSING THE ROLE OF DISTANCE

	Specification 1			Specification 2		Specification 3	
		Border					
Good	Constant	Log distance	Adjusted R^2	Log distance	Adjusted R^2	Log distance	Adjusted R^2
1	2.84	0.96	0.79	0.26	0.83	1.99	0.61
	(0.22)	(0.05)		(0.12)		(0.53)	
2	1.79	1.30	0.88	0.30	0.93	0.36	0.89
	(0.14)	(0.03)		(0.11)		(0.11)	
3	2.95	1.31	0.77	0.39	0.91	1.67	0.20
	(0.30)	(0.06)		(0.21)		(0.72)	
4	3.06	1.27	0.86	1.50	0.67	0.96	0.49
	(0.28)	(0.05)		(0.36)		(0.27)	
5	5.04	1.39	0.77	6.06	0.76	4.04	0.80
	(0.33)	(0.13)		(1.00)		(0.84)	
6	5.66	1.03	0.81	−0.05	0.87	50.2	0.58
	(0.40)	(0.05)		(0.18)		(13.0)	
7	10.2	1.59	0.87	0.38	0.85	1.91	0.55
	(0.69)	(0.10)		(0.40)		(0.47)	
8	19.1	3.41	0.89	2.29	0.85	1.57	0.90
	(1.38)	(0.19)		(0.74)		(0.37)	
9	10.2	0.95	0.87	−0.31	0.88	1.85	0.70
	(0.96)	(0.12)		(0.72)		(0.49)	
10	1.50	1.52	0.89	1.45	0.79	2.25	0.81
	(0.17)	(0.05)		(0.24)		(0.40)	
11	7.11	3.74	0.82	1.08	0.74	5.49	0.89
	(0.63)	(0.15)		(0.51)		(0.78)	
12	2.32	1.22	0.89	0.14	0.79	0.31	0.99
	(0.25)	(0.07)		(0.16)		(0.31)	
13	4.42	0.76	0.73	0.36	0.85	0.18	0.89
	(0.50)	(0.06)		(0.26)		(0.13)	
14	4.14	1.13	0.85	0.33	0.88	1.70	0.92
	(0.31)	(0.04)		(0.17)		(0.09)	
1–14	19.8	12.3	0.77	1.01	0.91	1.77	0.61
	(1.19)	(0.41)		(0.32)		(0.26)	

Notes: Heteroscedasticity-consistent standard errors (White, 1980) are reported in parenthesis. Specification 1 is the same as the specification in Table 3, with all variables deflated by the log of distance. In specification 2, the standard deviation of the two-month difference in the relative price for within-U.S. pairs is regressed on the log of distance and 14 individual U.S. city dummies. In specification 3, the standard deviation of the two-month difference in the relative price for within-Canada pairs is regressed on the log of distance and nine individual Canadian city dummies. All coefficients and standard errors have been multiplied by 1,000. Standard deviations are computed over the sample period from September 1978 to December 1994. There are 228 observations in each regression.

proportional to the log of distance between cities. That is, we estimate

$$V(P^i_{j,k})/r_{j,k} = \beta^i_1 + \beta^i_2(B_{j,k}/r_{j,k}) + \sum_{m=1}^{n} \gamma^i_m(D_m/r_{j,k}) + v_{j,k}.$$

The constant terms and the coefficients on the deflated border dummy are positive, as predicted, and highly significant in the regressions for each of the 14 goods.

We try several extensions to test the robustness of our results. In order to conserve space, we do not report these results. One variation is to alter the period covered by the data. We eliminate the early 1980's from our sample, using only data starting in September 1985. Over this later period the U.S. dollar experienced large swings in its value. There was virtually no change in the results in these regressions from the ones using the entire sample.

We also split the sample at January 1990, when the Canadian–U.S. Free Trade Agreement went into effect. If trade barriers are an important reason why the border variable is economically significant in explaining price dispersion, one would expect that the magnitude of this variable would decline after 1989. In fact, we found a slight tendency in the opposite direction: the estimated border coefficients were usually larger in the post-1989 period.

In general, there was very little difference in our full-sample estimates and our post-September 1985 and post-January 1990 results. Distance and the border dummy had positive coefficients for the same goods in all three samples. Not surprisingly, the *t* statistics were smaller in the shorter samples.

One other convex specification of the distance variable we tried is one in which we hypothesize that, after a certain critical distance (arbitrarily chosen to be 1,700 miles), additional distance does not contribute at all to volatility. In this model, there is a linear relation between volatility and distance for distances up to 1,700 miles, and then after 1,700 miles the derivative of volatility with respect to distance is zero. This model performs almost identically to the log-distance function in terms of the number of correct signs on coefficient estimates, the degree of significance, the adjusted R^2, and the magnitude of the coefficients on the border dummy.[10]

B. *How Important are Distance and the Border?*

We have seen that physical distance plays a significant role in explaining the failure in the law of one price between two locations. But physical distance alone does not explain the variability in prices of similar goods if the two locations are in different countries—the border matters.

We would like to get an idea of the economic significance of the border relative to distance in determining price dispersion. One way to do this is by examining the average coefficients on log distance and the border dummy from the regression in Table 3, which equal the reported coefficients for the pooled regression. There, the coefficient on the border is 11.9×10^{-3}, and on the log of distance it is 10.6×10^{-4}. Thus, crossing the border adds 11.9×10^{-3} to the average standard deviation of prices between pairs of cities. In order to generate that much volatility by distance, the cities would have to be 75,000 miles apart.[11] This calculation indicates that crossing the border adds substantially to volatility. Actually, this statistic may overstate the economic importance of the border, given that the natural log function is concave, and given the imprecision of the estimate of the coefficient on log distance. The 95-percent confidence interval for the distance coefficient is $(5.3 \times 10^{-4}, 15.9 \times 10^{-4})$. If we were to use the upper end of the confidence interval as the measure of the impact of distance, then crossing the border is equivalent to 1,780 miles of distance between cities. The effect of distance may also be understated if the log-distance function is not the appropriate one.

This statistic may not be meaningful if distance does not contribute much to the dispersion of prices—but that is not the case. Consider the price dispersion for cross-border pairs of cities. From Table 2, the average standard deviation is 0.0367. The border, which adds 0.0119 to the standard deviation of cross-border pairs, accounts for 32.4 percent of this. The average log distance between cross-border pairs is 7.03, so on average distance adds 0.00745 to the standard deviation, which is 20.3 percent of the total.

Table 5 also reports the results of regressing the price dispersion on the log of distance (and city dummy variables) when we use only U.S. cities (second specification of Table 5) and only Canadian cities (third specification). We note that for U.S. cities distance has the hypothesized positive coefficient for 12 of the 14 cities and is significant at the 5-percent level for eight of the goods (and significant at the 6-percent level for two more). When all 14 goods are used jointly, the effect of distance is positive

[10] We also included a dummy variable for pairs of cities in the same province or state. Inclusion of this dummy did not appreciably alter our results.

[11] Calculated as $\exp[(11.9 \times 10^{-3})/(10.6 \times 10^{-4})]$ miles.

and highly significant. For the Canadian cities, distance has a positive effect for all 14 goods, and it is significant for 13 of those goods. Thus, if we do not consider the effect of the border at all, we find that distance has strong explanatory power for price dispersion.

C. *Why Does the Border Matter So Much?*

Crossing the border adds significantly to price dispersion. In the Introduction, we proposed several reasons why the border would matter. Here we attempt to distinguish between a few of them.

We note that we have already tried a direct test for trade barriers and found that the size of the border coefficient was not diminished when the free-trade agreement between the United States and Canada went into effect. This, of course, does not rule out the possibility that informal trade barriers account for the price dispersion.

We suggested that labor markets might be more homogeneous within countries, so that w_j^i/w_k^i is less variable for city pairs (j, k) within a country than for cross-border pairs. We can investigate this hypothesis by seeing whether the explanatory power of the border dummy is affected by introducing relative wage volatility into the regression.

For each city, we construct a real wage as the average hourly wage for manufacturing employees (which is available for each city in the United States and by province in Canada) divided by the aggregate CPI for that city. We then calculate for each city pair the standard deviation of the two-month difference in the log of the relative real wages.

We add this wage-dispersion variable to our first regression, equation (2). These results are reported as the first specification in Table 6. As we expect, the wage dispersion coefficient is generally positive and significant. The coefficient is positive for 13 of the goods, and significant for ten.

However, the size of the border coefficient is not much affected by inclusion of the wage-dispersion variable. Apparently the border's importance does not arise because of the homogeneity of the labor markets within countries; but the distance coefficients are generally smaller and less significant. As we discuss in the Introduction, one of the reasons distance matters for intercity price dispersion is that more distant cities have less-integrated labor markets. The results from this regression bear out that hypothesis.

We investigate whether the sticky-price explanation for the importance of the border has power. In all of the regressions we have reported, if P_f is the U.S. dollar price of good f in a U.S. city, and P_f^* is the price in the Canadian city, the relative price is (the log of) P_f/SP_f^*, where S is the exchange rate. If P_f and P_f^* are sticky, then P_f/SP_f^* will fluctuate as S fluctuates. The border will be significant because it picks up the effect of the fluctuating exchange rate.

However, if we calculate the relative prices of good f between cities as relative real prices, then the nominal exchange rate will not appear in the calculation. That is, call P_f/P the real price of good f in the U.S. city, where P is an aggregate price index for that city, and P_f^*/P^* is the real price of good f in the Canadian city. Then the relative intercity price is $(P_f/P)/(P_f^*/P^*)$. If nominal price stickiness were the reason the border matters when we use P_f/SP_f^* as the measure of relative prices, then it should not be significant when we use $(P_f/P)/(P_f^*/P^*)$.

When the log of $(P_f/P)/(P_f^*/P^*)$ is taken to be the relative price, the filtered measure of prices is a better measure than the two-month difference. The log of $(P_f/P)/(P_f^*/P^*)$ appears to be stationary for all of our goods, so the two-month difference would be an over-differenced series.

The second specification reported in Table 6 is for the regressions when prices for the individual goods in each city are taken relative to the CPI for all goods in that city. The standard deviation of the filtered prices is regressed on the log of distance, the border dummy, and individual city dummies, so the explanatory variables are the same as in equation (2). We find that the coefficients on distance are all positive, and generally significant. The coefficients on the border dummy are all positive and highly significant. Thus, even without the nominal exchange rate in the calculation of cross-border prices, the border matters.

TABLE 6—ASSESSING WHY THE BORDER MATTERS

	Specification 1			Specification 2		Specification 3	
Good	Log distance	Border	SD of real wage	Log distance	Border	Log distance	Border
1	1.56 (1.15)	6.74 (0.23)	0.28 (0.08)	6.60 (1.87)	2.04 (0.14)	5.71 (2.08)	3.22 (0.16)
2	0.62 (1.02)	9.44 (0.13)	0.10 (0.03)	2.95 (0.93)	1.98 (0.11)	1.74 (0.60)	4.32 (0.09)
3	−0.84 (1.56)	9.01 (0.27)	0.36 (0.09)	5.44 (2.69)	3.46 (0.17)	4.47 (2.70)	4.28 (0.19)
4	5.72 (1.78)	8.83 (0.23)	0.22 (0.05)	4.99 (1.09)	2.66 (0.12)	6.69 (1.64)	6.15 (0.16)
5	31.5 (7.16)	9.53 (0.94)	0.35 (0.19)	28.6 (5.15)	3.47 (0.59)	28.6 (5.59)	4.19 (0.63)
6	−0.76 (1.01)	7.84 (0.15)	0.16 (0.03)	2.66 (0.87)	1.66 (0.10)	1.75 (0.85)	3.79 (0.11)
7	9.33 (2.97)	12.6 (0.46)	0.10 (0.11)	2.06 (2.42)	6.78 (0.32)	2.84 (2.38)	7.61 (0.31)
8	34.5 (7.97)	27.9 (1.12)	−0.54 (0.23)	11.0 (3.46)	13.0 (0.49)	13.2 (3.06)	13.7 (0.46)
9	6.15 (3.14)	8.84 (0.45)	0.13 (0.11)	5.44 (2.86)	5.34 (0.31)	4.64 (3.02)	5.93 (0.32)
10	7.03 (1.75)	10.2 (0.23)	0.23 (0.08)	8.34 (1.53)	3.15 (0.15)	9.82 (1.92)	4.66 (0.16)
11	24.6 (6.82)	25.4 (1.11)	0.68 (0.22)	23.2 (4.98)	21.3 (0.66)	21.8 (5.18)	21.8 (0.67)
12	−4.60 (1.73)	8.91 (0.26)	0.28 (0.06)	3.17 (1.70)	1.99 (0.16)	1.36 (1.78)	3.89 (0.18)
13	−0.16 (1.80)	6.66 (0.23)	0.01 (0.06)	1.53 (1.39)	1.62 (0.16)	2.22 (1.48)	3.17 (0.17)
14	3.21 (1.08)	8.16 (0.14)	0.15 (0.03)	4.98 (0.99)	2.11 (0.12)	3.70 (0.93)	3.99 (0.12)
1–14	8.43 (3.22)	11.4 (0.52)	0.18 (0.11)	7.93 (2.68)	5.04 (0.35)	7.76 (2.76)	6.48 (0.36)

Notes: All regressions contain a dummy for each of the 23 individual cities, in addition to the variables listed in the cell. Heteroscedasticity-consistent standard errors (White, 1980) are reported in parentheses. Coefficients and standard erorrs on log distance (border) are multiplied by 10,000 (1,000). Specification 1 is the same as the specification in Table 3 but adds the standard deviation of the two-month difference in the intercity real wage. Specifications 2 and 3 use a measure of the real price of each good: in specification 2, the individual goods prices are deflated by the city's overall CPI, while in specification 3 the deflator is the national PPI. The measure of volatility in each case is the standard deviation of the two-month-ahead forecast error from the filtered relative price, over the sample period from September 1978 to December 1994. There are 228 observations in each regression. The adjusted R^2 estimates, not reported in order to save space, were never less than 0.77

How much does the border matter in this regression as compared to the regressions in which relative prices are calculated as the log of P_f/SP_f^*? From the first specification reported in Table 4, the coefficient on the border dummy using the filtered measure of the log of P_f/SP_f^* when all 14 goods are aggregated in a single regression is 10.8×10^{-3}.

(Recall that the coefficients reported for the pooled regressions are the averages of the coefficients for the regressions for each of the 14 goods.) The average standard deviation for all cross-border city pairs using the filtered measure is 32.4×10^{-3}, so the border accounts for 33.3 percent of that standard deviation. When the log of $(P_f/P)/(P_f^*/P^*)$ is used as the measure of the relative price, the coefficient on the border dummy for the regression using all goods is 5.04×10^{-3} (last row of the second specification in Table 6). That compares to an average standard deviation of 26.6×10^{-3} for cross-border city pairs. Thus, the border accounts for only 18.9 percent of that standard deviation. Hence, when we drop the nominal exchange rate from our calculation of intercity prices, the percentage of the cross-border standard deviation accounted for by the border drops from 33.3 percent to 18.9 percent. We might conclude that the sticky-price story accounts for this difference; but we note that the border still accounts for a fairly large portion of the cross-border dispersion even after taking into account the role of sticky prices.

We also consider calculating the individual goods prices in each city relative to the national-level producer price index. The third row of Table 5 reports regressions using these prices, again taking the filtered measure of the log of $(P_f/P)/(P_f^*/P^*)$. We note that the results are qualitatively similar to the previous regression. Here, in the regression that uses all 14 goods, the coefficient on the border dummy is 6.48×10^{-3} (last row of the third specification in Table 6). The average standard deviation for cross-border city pairs with this measure of relative prices is 28.1×10^{-3}, so the border accounts for 23.1 percent of the total. This is still less than the 33.3 percent of the total when we use the log of P_f/SP_f^* as the measure of relative prices, but only about 30-percent less. Therefore, we can tentatively conclude that our sticky-nominal-prices story can explain about 30 percent of the border size.

We have not been able to explain fully why the border matters so much for intercity price dispersion. We have cast some doubt on the notion that formal trade barriers can explain it, while leaving open the possibility that informal barriers are significant. The hypothesis that wage costs are more homogeneous within countries does not seem to explain the border's importance. Sticky nominal prices do seem to account for a significant portion of the magnitude of the border effect, but apparently less than half. Other possibilities that we have not explored include differences in demand elasticities in the United States and Canada (which has received attention in the pricing-to-market literature) and homogeneity of productivity shocks within countries in the nontraded sectors (so that α_j^i/α_k^i from equation (1) has less dispersion within countries than between countries).

III. Conclusions

The major message of our empirical results is not just that the border matters for relative price variability; it is that both distance and the border matter. The literature on pricing to market has emphasized that, when markets are segmented, price discrimination can occur. The finding that distance is important in explaining price differences between locations lends support to this literature and the associated work on geography and trade. But our findings seem to suggest that there is more than standard price-discrimination behavior involved in cross-border price movements.

To the extent that our results indicate sticky nominal prices, they also shed some light on the price-setting process. We have found that the distance between markets influences prices, suggesting that price-setters take into account prices of nearby competitors. It is probably not too far-fetched to infer that firms would respond more to changes in prices of near substitutes, whether the nearness is in geographical space or product space. A reasonable model of price stickiness must take into account how isolated the market is for the product of the price setter. There appears to be potential for a marriage of the new-Keynesian literature on menu costs and the new trade literature emphasizing the role of geography.

Nominal price stickiness cannot account for all of the price dispersion between markets, however. The results of this paper confirm McCallum's (1995) finding that, despite the relative openness of the U.S.–Canadian border, the markets are still segmented.

Data Appendix

Our data for the United States was obtained from the Bureau of Labor Statistics. The 14 goods from the United States are listed on the left-hand side of Table 1. All of the price and wage data (for both countries) are seasonally unadjusted.

We use comparable price and wage data for Canada that were obtained from Statistics Canada. There is not always an exact match between the price indexes available in Canada and those available in the United States. However, we were able to construct indexes for the 14 categories of goods in Canada, in some cases by using even more disaggregated Canadian indexes. For example, the U.S. data contain a series on men's and boy's apparel. There is no comparable series in Canada. However, we can obtain from Canada individual series on men's wear and on boy's wear. We then construct a men's and boy's apparel series for Canada by taking a weighted average of the men's wear series and the boy's wear series.[12] This type of construction was needed to arrive at five of the 14 Canadian price series. Table 1 indicates how these series were derived.

These categories of goods are mutually exclusive. Together they comprise 94.6 percent of purchases (using the weights in the U.S. consumer price index).

Monthly price data were used for nine Canadian cities: Calgary, Edmonton, Montreal, Ottawa, Quebec, Regina, Toronto, Vancouver, and Winnipeg. Monthly price data for the United States are available for four "core" cities: New York, Philadelphia, Chicago, and Los Angeles. In addition, for five cities, data are released in even-numbered months: Dallas, Detroit, Houston, Pittsburgh, and San Francisco. For five other cities, there are data available in odd-numbered months: Baltimore, Boston, Miami, St. Louis, and Washington.[13]

Consumer price data are closer to being monthly average data than point-in-time data. Typically to get the price of a single product, several outlets are sampled during the month. The outlets are not all sampled on the same day. The change in the price of the product from the previous month is calculated as the average change across the various outlets. For the cities that report data every second month, the prices are for the second month of the interval (rather than an average across both months).

In order to nullify a potential bias, we use a monthly average (U.S. dollar)/(Canadian dollar) exchange rate from the Citibase tape. Averaging tends to reduce the volatility of the series. Thus, if we were to use an exchange rate at a specific point in time, but use price data which is essentially averaged, we would introduce volatility into our measure of cross-border prices. That is compensated for by taking the monthly average exchange rate.

For each good, we calculated the intercity relative prices. Thus, when we are using only the Canadian cities and the core U.S. cities, for each good there are 78 intercity prices (13 cities × 12/2). Adding the five even-month U.S. cities adds another 75 prices, and adding the five odd-month U.S. cities adds another 75 prices.

We also use data on the distance between cities. We use two separate measures of distance, both obtained from the Automap (version 2) software. One measure is the great-circle distance, and the other is the quickest-driving-time distance. Our results were not affected by the choice of distance measure, so all results reported use the great-circle distance.

[12] The weights come from the current weights used in the U.S. consumer price index, which we obtained from the Bureau of Labor Statistics.

[13] Data for Cleveland are available every other month. However, the data switched from being odd-month to even-month in the middle of our sample. Also, at the beginning of the sample, Detroit data were monthly, but switched to even-month, while the reverse is true for San Francisco. We make use only of the even-month data for these two cities.

REFERENCES

Debreu, Gerard. *Theory of value.* New Haven, CT: Yale University Press, 1959.

Dixit, Avinash. "Hysteresis, Import Penetration, and Exchange Rate Pass-Through." *Quarterly Journal of Economics,* May 1989, *104*(2), pp. 205–28.

Dornbusch, Rudiger. "Exchange Rates and Prices." *American Economic Review,* March 1987, *77*(1), pp. 93–106.

Dumas, Bernard. "Dynamic Equilibrium and the Real Exchange Rate in a Spatially Separated World." *Review of Financial Studies,* 1992, *5*(2), pp. 153–180.

Engel, Charles. "Real Exchange Rates and Relative Prices: An Empirical Investigation." *Journal of Monetary Economics,* August 1993, *32*(1), pp. 35–50.

______. "Accounting for U.S. Real Exchange Rate Changes." National Bureau of Economic Research (Cambridge, MA) Working Paper No. 5394, December 1995.

Feenstra, Robert C. "Symmetric Pass-Through of Tariffs and Exchange Rates under Imperfect Competition: An Empirical Test." *Journal of International Economics,* August 1989, *27*(1/2), pp. 25–45.

Frankel, Jeffrey; Stein, Ernesto and Wei, Shang-Jin. "Trading Blocs and the Americas: The Natural, the Unnatural and the Super-natural." *Journal of Development Economics,* June 1995, *47*(1), pp. 61–96.

Froot, Kenneth A. and Klemperer, Paul D. "Exchange Rate Pass-Through when Market Share Matters." *American Economic Review,* September 1989, *79*(4), pp. 637–54.

Giovannini, Alberto. "Exchange Rates and Traded Goods Prices." *Journal of International Economics,* February 1988, *24*(1/2), pp. 45–68.

Isard, Peter. "How Far Can We Push the Law of One Price?" *American Economic Review,* December 1977, *67*(5), pp. 942–48.

Kasa, Kenneth. "Adjustment Costs and Pricing-to-Market: Theory and Evidence." *Journal of International Economics,* February 1992, *32*(1/2), pp. 1–30.

Knetter, Michael N. "Price Discrimination by U.S. and German Exporters." *American Economic Review,* March 1989, *79*(1), pp. 198–210.

______. "International Comparisons of Pricing-to-Market Behavior." *American Economic Review,* June 1993, *83*(3), pp. 473–86.

Krugman, Paul. "Pricing to Market When the Exchange Rate Changes," in Sven W. Arndt and J. David Richardson, eds., *Real-financial linkages among open economies.* Cambridge, MA: MIT Press, 1987, pp. 49–70.

______. "Increasing Returns and Economic Geography." *Journal of Political Economy,* June 1991, *99*(3), pp. 483–99.

Marston, Richard C. "Pricing to Market in Japanese Manufacturing." *Journal of International Economics,* November 1990, *29*(3/4), pp. 217–36.

McCallum, John. "National Borders Matter: Regional Trade Patterns in North America." *American Economic Review,* June 1995, *85*(3), pp. 615–23.

Mussa, Michael. "Nominal Exchange Rate Regimes and the Behavior of Real Exchange Rates: Evidence and Implications." *Carnegie-Rochester Conference Series on Public Policy,* Autumn 1986, *25*, pp. 117–214.

Rogers, John H. and Jenkins, Michael. "Haircuts or Hysteresis? Sources of Movements in Real Exchange Rates." *Journal of International Economics,* May 1995, *38*(3/4), pp. 339–60.

Sanyal, Kalyan K. and Jones, Ronald W. "The Theory of Trade in Middle Products." *American Economic Review,* March 1982, *72*(1), pp. 16–31.

White, Halbert. "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity." *Econometrica,* May 1980, *48*(4), pp. 817–38.