Relative Price Volatility:

What Role Does the Border Play?

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Abstract:

We reexamine the effect of the U.S./Canadian border on integration of markets. The paper updates work from our earlier paper, Engel and Rogers (1996). We consider alternative measures of deviations from the law of one price. We pay special attention to the effect of the U.S.-Canada free trade agreement on market integration. Our conclusions are unchanged: markets in the U.S. and Canada are more segmented than can be explained by the physical distance between the locations. Formal trade barriers do not appear to explain much of that segmentation.

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Popular accounts of the international economy stress the increased "globalization" of economic activity. The press and other media highlight the increased openness of goods markets and capital markets across borders, and the burgeoning volume of trade flows and international capital flows. While there is little doubt that markets are opening worldwide, curiously academic work has begun to focus more on the amount of market segmentation, particularly in goods markets. One strand of the literature (exemplified by Krugman (1991)) stresses the role of geography in determining international trade flows. This new line of research finds its ancestry in the "gravity" model of trade.[1] A related area of research stresses that segmentation of goods markets internationally allows monopolistic firms to price discriminate across national markets. Krugman (1987) presents some of the early theoretical analysis, which has been bolstered by a wealth of empirical evidence.[2]

Empirical work stemming from study of the gravity model has used the volume of trade flows between two locations as a measure of integration between those locations.[3] When there are few impediments to trade, either imposed by policy, culture or geography, then goods should flow freely between countries, cities or provinces. Our contention, however, is that trade flows are a problematic measure of the degree of market integration. Traditional trade theory that assumes no impediments to trade still predicts the volume of trade flows between two regions depends on such things as factor endowments or the degree to which economies of scale are exploited. The absence of costs to trading does not imply that there will be an unlimited, or even necessarily a large volume of trade. So, to use the volume of trade flows as a measure of how openness has changed over time, one must carefully monitor other determinants of trade flows to ascertain whether the change in trade volumes can be attributed to changes in integration or other factors.

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[1] See, for example, Pöyhönen (1963), Pullianen (1963) or Linneman (1966). Helliwell (1998) traces the lineage of this area of research.
[2] See Goldberg and Knetter (1997) for an excellent survey of the literature.

Our approach to measuring market integration is to examine prices of goods and services. A fundamental proposition of economic theory is that in the absence of transactions costs, identical goods must sell for the same price. A comprehensive measure of how well two markets are integrated is how closely prices move together in those markets. Prices will fail to equalize when there are barriers to the free movement of goods. These barriers might be natural barriers, such as the geographic distance between regions emphasized by the gravity literature. But there may be other barriers that are man-made. Tariffs and other formal trade barriers lead to inequality of consumer prices between locations, as do informal trade barriers. These latter include, for example, marketing agreements, or tradition, which tend to leave foreign goods on an unequal footing with domestically produced goods in the consumer market.

Here we extend our earlier research, particularly Engel and Rogers (1996), that investigates the relative importance of these types of barriers to market integration. We examine the behavior of prices of fourteen categories of consumer goods and services among 14 cities in the United States, and 10 provinces in Canada. We ask why the price of a particular category of goods in one location fluctuates relative to the price of similar goods in a different location. We relate a measure of this volatility to various explanatory variables, including the distance between the pair of locations, a dummy variable for whether the locations are in different countries, and a variable meant to capture different labor market conditions in the two locations.

There is related work that investigates market integration by looking at the adjustment of goods prices. A series of papers by Engel and Rogers separately (Engel (1993, 1995), Rogers and Jenkins (1995)) documents that markets are very poorly integrated if final goods prices are the benchmark. Failures of the law of one price account for the vast majority of real exchange rate movements in

[3] Recent empirical studies of the gravity model include Bergstrand (1985, 1989), Frankel and Wei (1994), Frankel, Stein and Wei (1994), McCallum (1995) and Helliwell (1996).

the short run and longer for industrialized countries. Another co-authored paper, Engel and Rogers (1998), looks at the behavior of a number of categories of goods over a large sample of countries. Relative to Engel and Rogers (1996), the later paper uses data from more locations – twenty-three countries and eight North American cities – but fewer goods – only eight categories. We find that relative price volatility (the volatility of prices of similar goods across locations) is a function primarily of exchange rate volatility and distance. Other similar work includes that of Wei and Parsley (1995), who find that the speed of convergence to PPP also depends on the distance between locations.[4]

Two recent contributions examine the convergence of prices within the U.S. Parsley and Wei (1996) find that the speed of convergence of prices of goods sold in cities within the U.S. is lower the more distant the city pairs. Nonetheless, convergence tends to be much faster than is found in international price data, suggesting that national borders somehow affect the speed of convergence. O'Connell and Wei (1997) model prices as a non-linear process, in which the speed of adjustment is greater when relative prices lie outside of some band. This model, which fares well empirically, is meant to capture the effect of transportation costs. Goods arbitrage will not occur when prices across locations are close, even if they are not equal. But, when they diverge greatly, market forces act more rapidly to cause prices to converge.

Our work on international and intra-national pricing can be considered the complement of the recent work on trade flows within and between countries. McCallum (1995), Helliwell (1996), Wei (1996) and Wolf (1997) all find that the volume of trade between countries is significantly less than the volume of trade within countries, taking into account other determinants of trade volumes such as distance and size of the trading unit.

[4] Although Engel, Hendrickson and Rogers (1997) find no such bias in a panel study involving eight cities from four countries on two continents.

Here, we are particularly interested in the effect of the U.S.-Canadian free-trade agreement on the volatility of these relative prices. Using estimates from a gravity model, Helliwell (1998) finds that the bias in trade between Canadian provinces (relative to trade between provinces and U.S. states) fell significantly after the free-trade agreement. If indeed trade restrictions were a chief obstacle to market integration, then we should see prices moving more closely together after the free-trade agreement than before. So, we investigate how prices behave before January 1990, when the free-trade agreement went into effect, compared with how they behave afterwards.

Our basic empirical results show that while distance is a significant deterrent to market integration, national borders impose a much more important barrier. But, trade barriers do not seem to explain the border effect. As in Engel and Rogers (1996), we conclude that there are likely two significant reasons why the border matters so much. First, to the degree that any two markets are segmented, there is opportunity for pricing to market. If, in addition, there is nominal price stickiness and prices are set in the consumers' currencies, then Canadian and U.S. prices can diverge greatly in the short run when the nominal exchange rate fluctuates. However, this nominal price-stickiness effect does not account for all of the border effect. We posit that there are national markets for consumer goods, established by tradition, by national distribution networks, and national marketing campaigns. These national markets would lead to deviations in U.S. and Canadian prices even in the absence of nominal price stickiness.

One issue that needs to be addressed up front is our use of consumer prices for some goods and services that are traditionally classified as "non-traded". In the model of Engel and Rogers (1996), all consumer goods have a non-traded component. This non-traded component comes from location-specific costs of marketing and distribution services and other local services. It is probably not useful to classify goods simply as "traded" and "non-traded". There is a continuum depending on the degree of input of local services. So, there are at least two determinants of how closely linked consumer markets are between locations: the degree of the barriers to shipment of the traded goods component, and the size of the non-traded component in any consumer good. But, also, as we discuss in Engel and Rogers (1996), there is a third determinant. Even if the non-traded service accounts for a large fraction of the cost of the good, the costs of these services is not necessarily independent across locations. For example, in a two-factor two-sector model of international trade, if one good is traded and one factor is traded, then factor prices for both factors will be equalized across locations. In turn, the price of non-traded goods is equalized across countries.[5]

A fourth factor that determines whether prices equalize is the degree of monopoly mark-up. In standard models of price discrimination, in regions where the elasticity of demand is lower, the mark-up will be higher. This is the avenue highlighted by much of the work on pricing to market.

Following Engel and Rogers (1996), we can write the price of some good in location *i* as:

$$p_i = \mu_i w_i^{\gamma} q_i^{1-\gamma},$$

where μ_i is the mark-up, w_i is the cost of the non-traded service, q_i is the price of the traded component, and γ is the share of the non-traded component in total costs.

So, we can write:

$$\ln(p_i) - \ln(p_j) = \ln(\mu_i / \mu_j) + \gamma \ln(w_i / w_j) + (1-\gamma)\ln(q_i / q_j).$$

Assuming each of the components on the right-hand-side is independent of the others, we have the variance of the relative price given by:

$$Var(\ln(p_i) - \ln(p_j)) = Var(\ln(\mu_i / \mu_j)) + \gamma^2 Var((\ln(w_i / w_j)) + (1-\gamma)^2 Var(\ln(q_i / q_j)).$$

So, the variance of the relative price increases as the variance of the relative mark-ups increases, as the variance of the relative costs of the non-traded service increases, and as the variance of the

[5] For example, the model of Engel and Kletzer (1989) has this structure.

relative price of the traded good increases. The latter variance is not zero because of transportation costs and other barriers to shipment of traded goods. Then, assuming,

$$Var((\ln(w_i/w_j)) > Var(\ln(q_i/q_j)),$$

the variance of the relative price of the final product also increases as the share of the non-traded component increases.

In our empirical work below, we find that the presence of a national border separating two locations is a significant determinant of the variance of the relative price of the final product. There is not an obvious link between the fact that our data contain prices on many "non-traded" goods, and the presence of the border effect. Most hypotheses of why goods are not traded would imply that goods are not traded across distant locations, whether or not there is a national border between the locations. So, for example, if transportation costs are high, then the good might not be traded between cities or provinces that are far apart. But, it is not clear why the national border should matter. The border effect cannot be easily explained by simply noting that some of the goods in our sample are non-traded. What is required is an explanation of why goods are more non-traded across national borders than among locations within a country. That is, why are international markets more segmented than intra-national markets, taking into account distance?

1. Econometric Model and Data

1.a Model specification

Let q_{ij} be the log of the relative price of some good between locations i and j. We model this price as following a stationary 6th-order autoregressive process, with twelve monthly seasonals.[6] We take the standard deviation of the 12-month ahead forecast as the dependent variable to be

[6] For prices that are measured bi-monthly, we assume an AR(4) with six seasonal dummies.

explained by factors such as the distance between location *i* and *j* and a dummy variable for whether the locations are in the same country or in a different country.

In practice, we find qualitatively essentially the same results whether we use the standard deviation of the 12-month forecast error from the stationary model, or if we use the standard deviation of the first-difference of q_{ij}.[7] This may seem surprising at first. We certainly expect q_{ij} to be a stationary variable. There is probably no plausible economic theory that would suggest that q_{ij} could have a unit root. But in practice, even if q_{ij} is stationary, it is so persistent that there is very little evidence of convergence for any of the price series. So, for most of the regressions we estimate, we use the standard deviation of the first-difference of q_{ij} as the dependent variable, because it enhances the reproducibility of our findings.

Let $V(q_{ij})$ be one of the measures of volatility of q_{ij} described above. Our basic regressions are of the form:

(1) $V(q_{ij}) = \beta_1 d_{ij} + \beta_2 B_{ij} + \sum_{k=1}^{m} \lambda_k D_k$.

Here, d_{ij} is the log of the distance between locations *i* and *j*; B_{ij} is a dummy variable which takes on the value of one if locations *i* and *j* are in different countries; and D_k are dummy variables for each location which take on a value of one for D_i and D_j and a value of zero for all other locations. We expect to find $\beta_1 > 0$ and $\beta_2 > 0$.

For each good there are 251 city pairs. We consider the standard deviation of, for example, New York-Los Angeles prices to be a separate observation from the New York-Chicago and Chicago-Philadelphia observations. In general, if there are *G+1* cities, then there are only *G*

[7] We actually use two-month differences, because some of our data is bi-monthly.

independent relative prices. Yet we calculate *G(G+1)/2* standard deviations.[8] It is helpful to think of our estimation as a two step procedure, in which we estimate a model for the levels of relative prices in the first step, and then estimate a model for the covariance matrix of those relative prices in the second step.

Consider the model in which we take the standard deviation of the first difference of q_{ij}. Define $\varepsilon_{ij,t} = q_{ij,t} - q_{ij,t-1}$. Pick one location as the reference, and call it location *0*. We can consider the $\varepsilon_{i0,t}$ as the first-differences of the G independent prices. The other $\varepsilon_{ij,t}$ are related to $\varepsilon_{i0,t}$ by $\varepsilon_{ij,t} = \varepsilon_{i0,t} - \varepsilon_{j0,t}$.

Equation (1), however, can be thought of as a model that estimates the elements of the covariance matrix of the $\varepsilon_{i0,t}$. Let $Var(\varepsilon_{ij})$ be the variance of $\varepsilon_{ij,t}$. Then

$$Var(\varepsilon_{ij}) = Var(\varepsilon_{i0}) + Var(\varepsilon_{j0}) - 2 \cdot Cov(\varepsilon_{i0}, \varepsilon_{j0}),$$

where Cov refers to the covariance. It follows that

$$Cov(\varepsilon_{i0}, \varepsilon_{j0}) = \frac{1}{2}(Var(\varepsilon_{i0}) + Var(\varepsilon_{j0}) - Var(\varepsilon_{ij})).$$

So, all of the covariances, $Cov(\varepsilon_{i0}, \varepsilon_{j0})$, can be obtained from estimates of the variances of the relative prices. Thus, estimates of the *G(G+1)/2* elements of the covariance matrix of $\varepsilon_{i0,t}$ can be obtained from estimates of the *G(G+1)/2* variances of $\varepsilon_{ij,t}$. Equation (1), then, can be considered a model that provides estimates of the elements of the covariance matrix of $\varepsilon_{i0,t}$.

In words, if Chicago is the base location, even after we calculate the variance of the Chicago/New York and the Chicago/Los Angeles relative prices, we are still interested in the variance of the New York/Los Angeles relative price because this variance is needed to calculate

[8] Actually, we calculate fewer standard deviations because we do not attempt to measure the standard deviation of the

the covariance of the Chicago/New York and the Chicago/Los Angeles relative prices. So the New York/Los Angeles relative price is not redundant.

In one set of regressions, rather than using a measure of the volatility of q_{ij} as the dependent variable, we use the correlation of the logs of p_i and p_j (where, recall, $q_{ij} \equiv \ln(p_i) - \ln(p_j)$). Let each of these nominal prices share a common nominal shock and independent idiosyncratic shocks:

$$\Delta \ln(p_{it}) = u_{it} + v_t ,$$

$$\Delta \ln(p_{jt}) = u_{jt} + v_t .$$

Noting that $\varepsilon_{ij,t} = u_{it} - u_{jt}$, then $Var(\varepsilon_{ij}) = Var(u_i) + Var(u_j)$. But, the correlation of $\Delta \ln(p_{it})$ and $\Delta \ln(p_{jt})$ can be written as $\left(1 + \dfrac{Var(u_i)}{Var(v)}\right)^{-\frac{1}{2}} \left(1 + \dfrac{Var(u_j)}{Var(v)}\right)^{-\frac{1}{2}}$. So, the correlation measure scales the variance of the independent idiosyncratic shocks by the variance of the common nominal shock. As the idiosyncratic shocks are larger relative to the common shock, the correlation declines.

For this specification, we estimate the model:

(2) $corr(p_i, p_j) = \alpha_1 d_{ij} + \alpha_2 B_{ij} + \sum_{k=1}^{m} \theta_k D_k$,

where the right-hand side variables are defined as in equation (1). We now expect to find $\alpha_1 < 0$ and $\alpha_2 < 0$.

1.b. Data

Table 1 displays the fourteen categories of goods for which we have data. The data are disaggregated into these categories because that is the most disaggregated data that is publicly available for U.S. cities. The Canadian sub-indexes that are reported do not match exactly with

prices from U.S. cities measured in even-numbered months relative to those measured in odd-numbered months.

U.S. indexes, so we construct sub-indexes for Canadian categories from even more disaggregated data. The data are monthly for each of the ten Canadian provinces, and for four U.S. cities – New York, Los Angeles, Chicago and Philadelphia. We have bi-monthly data for ten other U.S. cities. For five of these cities the data are measured in odd-numbered months, and the others in even-numbered months.

The data are monthly from September 1978 to December 1997. This data is similar to the data we used in Engel and Rogers (1996). The categories of consumer prices are identical. The data in that paper started in September 1978 and ended in December 1994. The data is for the same fourteen U.S. cities. Our earlier paper had data on consumer prices for nine Canadian cities, but Statistics Canada ceased publishing disaggregated price data by city at the end of 1994. So, in this paper we use price data by Canadian province.

Distance between U.S. cities is calculated using the great circle distance. For Canadian provinces, we measure great circle distances from the largest city of each province.[9]

Table 2 presents some summary statistics for each of the fourteen categories of goods, as well as for the overall CPI (labeled "All" in Table 2.A.) The first set of statistics is for the standard deviation of the first difference of the relative prices.[10] The statistics are reported for all within-U.S. city pairs, for all within-Canada province pairs, and for all U.S. city – Canadian province pairs. In all but three cases, the standard deviation for the cross-border location pairs is greater than for either of the within-country pairs. We will investigate whether this arises because cross-border locations are, on average, farther apart, or whether the border itself contributes to this standard deviation. The three exceptions, in which the largest standard deviation is for U.S. city pairs, are the three

[9] Newfoundland – St. John's; Prince Edward Island – Charlottetown; Nova Scotia – Halifax; New Brunswick – St. John; Quebec – Montreal; Ontario – Toronto; Manitoba – Winnepeg; Saskatoon – Regina; Alberta – Calgary; British Columbia – Vancouver.
[10] The standard deviations are multiplied by 100.

clothing categories: men's and boys' apparel; women's and girls' apparel; and footwear. It may be notable that only a small fraction of the goods sold in these categories is produced either in the U.S. or Canada. Finally, we note from Table 2.A that in all but three cases, the smallest standard deviation is for within-Canada province pairs. This tendency for volatility to be lower in Canada may reflect greater integration of local markets within Canada, or it may reflect some sort of difference in how prices are constructed in Canada versus the U.S. We will take account of this in our regressions in section 2.

The second set of statistics reported in Table 2.A is the correlation coefficients of prices for U.S. city pairs, Canadian province pairs and cross-border location pairs. Here we note a similar pattern as was found with the standard deviation measures. The cross-border pairs appear to be the least integrated, as evidenced by their lower correlations. In this case, the correlations are lowest for cross-border pairs for all categories except footwear, household furnishings and operations, and personal care.

The third set of statistics reported in Table 2 .A is the standard deviation of the 12-month forecast error from the model described at the beginning of section 1.a. Here we note that for all goods except women's and girls' clothing, the standard deviation is largest for the location pairs that lie on opposite sides of the U.S./Canada border. For all goods except private transportation and medical care, the intra-Canadian relative prices are the least volatile.

The first columns of Table 2.B report standard deviations of the first differences of relative prices for three sub-periods: 1978-1989, 1990-1993 and 1994-1997. The earlier sub-period is prior to the U.S./Canada free-trade agreement. That agreement went into effect in January 1990. Allowing a few years for adjustment, the 1994-1997 period represents a time span in which there were fewer formal trade restrictions than in the 1980s. If it is formal trade restrictions that lead to less market integration across the national border, then we would expect to see that national borders matter less in the 1994-1997 period than in the 1978-1989 period. In fact, we note very little difference. As reported in the last row, taking all 14 goods together, the standard deviation falls from 3.63 percent to 3.54 percent from 1978-89 to 1994-97. The standard deviation is higher in the earlier period for seven of the goods, but higher in the latter period for the other seven.

2. Basic Regression Results

Table 3 reports three versions of estimates of equation (1) and one version of equation (2). The first set of estimates attempts to replicate the estimates in Engel and Rogers (1996). It uses the same time period. We note that in this paper, Canadian price data is by province, while in the 1996 paper it was by city. In this first set of estimates, the dependent variable is the standard deviation of the two-month differences in the relative price between pairs of locations. The sample period in this regression is identical to the 1996 paper: September 1978 to December 1994.

The second specification is a version of equation (2). The dependent variable is the correlation of the first-difference of p_i and p_j for locations *i* and *j*. The sample period is September 1978 to December 1994, so the second specification differs from the first only in the definition of the dependent variable.

The third specification is identical to the first, except the model is estimated for an updated data set: September 1978 to December 1997.

The fourth specification is estimated on the same sample as the third, but the dependent variable is the standard deviation of the 12-month ahead forecast error from the AR(6) model with seasonal dummy variables.

All four specifications qualitatively tell almost identical stories. First, we focus on the common elements of the four specifications. Then we briefly mention some minor differences. Note that Table 1 only reports the coefficient estimates for the distance and border dummy variables, and not the individual location dummy variables.

1. The final row of Table 3 reports regression results for a regression that pools data for all fourteen goods. For all specifications, distance between locations is statistically significant and of the correct sign. Apparently distance explains part of the segmentation of markets.

2. In the pooled regressions, the border dummy variable is highly significant, and of the correct sign for all specifications. Even taking into account distance, U.S./Canadian location pairs are less integrated than city pairs within the U.S., or province pairs within Canada.

3. The first fifteen rows report the results for regressions for the overall CPI and each of the fourteen individual goods. In every single regression, the border coefficient is significant and of the expected sign.

4. Distance is usually of the expected sign and significant. But, for some goods this tends not to be true: Distance is always statistically insignificant (and sometimes of the wrong sign) for alcoholic beverages; household furnishings and operations; men's and boys' clothing; footwear; medical care; and personal care. Distance is always significant and of the correct sign for food at home; shelter; fuel and other utilities; private transportation; and entertainment.

5. There is no apparent relationship between the "tradability" of the good and the size of either the distance or border coefficients.

These regression results tend to confirm those in Engel and Rogers (1996). One slight difference is that in the regressions for individual goods, the distance variable is less frequently significant and of the correct sign. This is probably because in the current study, the data for

Canada is provincial, as opposed to city, data. Measuring distance between provinces in Canada, or between provinces in Canada and cities in the U.S. is more ambiguous than calculations of distances between cities in the earlier study.

We note that there is virtually no difference between specifications 1 and 3 – the specification corresponding to the Engel and Rogers (1996) sample period (9/78 to 12/94) and the updated sample (9/78 to 12/97).

We also note in the three specifications that use the standard deviation as a measure of volatility (specifications 1, 3 and 4), the ratio of the coefficient on log of distance and the border is about the same. This suggests that the influence of the border (relative to distance) does not depend on the specification or the sample period.

Now we turn to the minor differences across the four specifications. Distance is significant using the 12-month forecast errors, but for no other specification, for food away from home. Distance is significant in specifications 2 and 4 (but not 1 and 3) for women's and girls' apparel.

As we noted in our earlier paper, the border coefficient is not only highly statistically significant, but it is also very large in economic terms. To see this, note that the average distance between a pair of locations in our sample is around 1200 miles. So, the average effect of distance on the standard deviation is $0.0684 \cdot \ln(1200) = 0.485$. The effect of crossing the border is 1.19, which is about four times the effect of the average distance. The estimated coefficient on the border (in the pooled regression) in this paper for the two-month changes in relative prices is almost identical to the estimate in Engel and Rogers (1996). However, the estimated coefficient on the log of distance is about two-thirds the estimate in our earlier paper. We hypothesize that this arises because our earlier paper had data on Canadian city prices, for which distance can be precisely measured, while this paper uses price data by Canadian province for which distance measures are necessarily cruder.

3. What Explains the Large Border Effect?

In this section we investigate possible explanations for the large border effect. Our first candidate hypothesis is that the border matters because of formal trade restrictions. Helliwell (1998) cites evidence that for trade flows, the border effect declines noticeably in the 1990s, which he attributes to the implementation of the U.S./Canadian free trade agreement in the early part of the decade.

Table 4 presents regressions over sub-periods 1978-89, 1990-93 and 1994-97. Of particular interest is comparison of the first column, which is for the period prior to the free-trade agreement, to the last column, which is for a period in which the free trade agreement is in effect and sufficient time has passed for the economies to adjust to the new laws. The first thing that we note from that pooled regression results is that the coefficient on the border dummy drops by about 20 percent in the two sub-periods: 1.23 in the early period, and 1.00 in the later period.

If we look by good, the border coefficient does drop for 11 of the 14 goods from the earlier to the later period. The decline is generally small, but statistically significant. The largest decline occurs for the clothing items – men's and boys' clothing; women's and girls' clothing; and footwear – and for private transportation, public transportation and fuel and other utilities.

Helliwell (1998) notes that the effect of distance on trade is shrinking over time. So, if we want to measure the effect of the border on market integration, we should compare it to the effect of distance on integration. By this measure, the border is no more important in the 1994-97 sub-period than in the pre-FTA period. Looking at the pooled regression, while the border coefficient falls by about 20 percent, so does the coefficient on distance, which falls from 6.35 to 4.99 (and is not statistically significant). In the 1990s, it appears that the effects of both distance and the border in segmenting markets have fallen about equally.

So, following Helliwell's logic, the decline of the border effect in the 1990s is probably not caused by the free-trade agreement according to our regressions. While the border effect has declined, the size of the border effect has declined in proportion to the decline of the distance effect. Markets are becoming more integrated, perhaps because of improvements in the efficiency of transportation of goods, or communications, or marketing. But, the removal of trade barriers cannot explain the simultaneous diminution in the importance of both the border and distance effects.

Since the removal of trade barriers does not do much to diminish the border effect, we turn to other possible explanations in Table 5. One possibility is that labor markets may be more homogeneous within countries. As we noted in section 1, the variability of the relative price between two locations will depend on the variability of the relative price of non-traded services, which we approximate with relative wages between locations. We investigate this hypothesis by including the standard deviation of the relative wage into our regression (1) to see if it reduces the explanatory power of the border dummy.

The results of those regressions (which add the standard deviation of the relative wage as an explanatory variable to specification 3 from Table 3) are reported in the first column of Table 5. According to our theory, the relative price volatility ought to increase when the relative wage volatility is greater. But, in the pooled regression, we actually find the opposite result – the relative wage variable has a negative sign (though it is statistically insignificant.) It does not change the border coefficient at all, and has a very minor effect on the distance coefficient. In the regressions for the individual goods, the relative wage variable is generally insignificant and has the wrong sign for seven of the fourteen goods.

When we construct the price of a good for a U.S. city relative to a Canadian province, the prices are converted into a common currency using the U.S. dollar/ Canadian dollar exchange rate. One possible explanation for the large border coefficient is nominal price stickiness. Specifically, suppose that U.S. prices tend to be sticky in U.S. dollar terms, and Canadian prices are sticky in Canadian dollar terms. The exchange rate is much more volatile than goods prices. So, when we examine cross-border prices, we should expect a lot of volatility – the goods prices, in their own currencies, do not move much but the exchange rate does. Relative prices for location pairs within a country – either the U.S. or Canada – of course do not involve an exchange rate, and so are likely to be fairly stable because the prices are sticky in their own currencies. So, the cross-border prices may exhibit high volatility because they contain the volatile exchange rate, while within-country prices do not.

We would like to find out whether the exchange rate volatility accounts for all of the border effect. One way to get at this is to construct cross-border relative prices that do not involve the exchange rate. We would not want to take the Canadian price of a good in Ontario, for example, relative to the U.S. dollar price of a good in New York, because that relative price would then be in units of Canadian dollars per U.S. dollars, while all of the within-country prices would be unit-free.

Instead, we first express the prices of all goods in a particular location relative to the overall price index of that location. For example, we take food at home (good 1) in Ontario relative to the overall CPI in Ontario. We can then compare this relative price to a similar price in a different location, such as the price of food at home in New York relative to the overall CPI in New York. By constructing relative relative prices such as this, we can compare prices across all locations without using nominal exchange rates. If the sticky-price cum volatile-nominal-exchange-rate explanation accounted for all of the border effect, then the border coefficient should not be significant in a regression such as equation (1) with relative prices constructed in this way.

We present the coefficient estimates from such regressions in the second panel of Table 5. The coefficient estimates from this regression are not directly comparable to any previous regressions since the dependent variable is constructed differently. But, the notable point about these regressions is that the border coefficients are still highly significant for the pooled data and for every good. Apparently we cannot explain the market segmentation between the U.S. and Canada solely on the basis of sticky nominal prices.

4. Conclusions

Using the metric of price dispersion, markets are segmented between the U.S. and Canada to a much greater extent than can be explained by the distance between the locations. It appears that formal trade restrictions do not explain the segmentation. Differences in labor costs also do not help explain the relative price dispersion across locations. While there may be a role for sticky nominal prices in conjunction with a volatile nominal exchange rate, that avenue does not explain the entire border effect.

What is left? It appears that consumer markets are, to a great degree, national markets. Perhaps this is because distribution networks are organized nationally. Perhaps it is because marketing efforts are conducted on a national basis. Perhaps there are non-tariff legal barriers to movement of goods between U.S. and Canadian markets. Perhaps tastes, shaped by custom and advertising, are differentiated across national borders.

It is a significant understatement to observe that prices are an important economic variable. At this stage, we do not have a clear understanding of what determines differences in prices across locations. This continues to be an important and rich area for research.

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Table 1: Categories of Goods in Dis-aggregated Consumer Price Indices and Locations Used

Good	United States	Canada
1	Food at Home	Food purchased from stores
2	Food away from home	Food purchased from restaurants
3	Alcoholic beverages	Alcoholic beverages
4	Shelter	Shelter
5	Fuel and other utilities	Water, fuel, and electricity
6	Household furnishings and operations	Housing operations and furnishings
7	Men's and boys' apparel	0.8058*(Men's clothing) + 0.1942*(children's clothing)
8	Women's and girls' apparel	0.8355*(Women's clothing) + 0.1645*(children's clothing)
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)

The U.S. cities in the sample are: Baltimore, Boston, Chicago, Dallas, Detroit, Houston, Los Angeles, Miami, New York, Philadelphia, Pittsburgh, San Francisco, St. Louis, and Washington D.C. The Canadian provinces are: New Foundland, Prince Edward Island, Nova Scotia, New Brunswick, Quebec, Ontario, Manitoba, Saskatchewan, Alberta and British Columbia.

Table 2: Summary Statistics

A. Three different measures of L.O.P. deviations, full sample

	$SD(\Delta p_{jk})$			$Corr(\Delta p_j, \Delta p_k)$			$SD(P_{jk})$		
Good:	U-U	C-C	U-C	U-U	C-C	U-C	U-U	C-C	U-C
All	0.83	0.52	1.80	0.54	0.65	0.35	1.35	1.08	1.85
1	1.42	1.68	2.48	0.46	0.47	0.30	1.99	2.48	4.71
2	1.23	0.95	2.03	0.19	0.58	0.15	2.13	1.78	4.53
3	1.80	1.50	2.51	0.29	0.49	0.15	2.85	2.56	4.87
4	2.02	0.84	2.40	0.27	0.51	0.13	3.41	2.03	4.96
5	5.03	2.86	5.18	0.30	0.36	0.07	4.96	4.89	7.04
6	2.08	0.98	2.32	0.11	0.33	0.12	2.86	1.47	4.39
7	5.41	2.34	5.05	0.28	0.39	0.13	5.59	2.47	6.23
8	9.15	2.58	8.60	0.44	0.42	0.25	9.23	2.59	9.31
9	6.31	2.66	5.27	0.16	0.40	0.18	7.81	3.18	7.58
10	1.12	1.56	2.50	0.77	0.52	0.30	1.77	2.32	4.97
11	4.62	1.82	6.55	0.37	0.93	0.07	7.45	2.55	9.21
12	1.29	1.73	2.37	0.21	0.48	0.12	2.23	3.42	5.13
13	2.60	1.16	2.67	0.03	0.43	0.07	4.10	1.87	5.35
14	2.00	0.84	2.35	0.08	0.68	0.04	2.97	1.13	4.45
Pooled (1-14)	3.29	1.68	3.73	0.32	0.51	0.16	4.24	2.48	5.91

B. Relative price variability over three sub-periods

Good:	SD(Δp_{jk}), 1978-89 U-U	C-C	U-C	SD(Δp_{jk}), 1990-93 U-U	C-C	U-C	SD(Δp_{jk}), 1994-97 U-U	C-C	U-C
All	0.92	0.52	1.82	0.70	0.55	1.89	0.61	0.45	1.65
1	1.29	1.65	2.42	1.63	1.75	2.68	1.55	1.61	2.39
2	1.46	0.97	2.00	0.71	0.88	2.31	0.70	0.77	1.70
3	1.76	1.73	2.61	1.92	1.13	2.40	1.70	1.01	2.21
4	2.44	0.88	2.68	1.26	0.85	2.13	0.99	0.65	1.72
5	4.80	2.81	5.06	5.11	2.90	5.38	5.28	2.69	5.05
6	1.90	1.00	2.23	2.28	0.95	2.47	2.32	0.91	2.37
7	4.83	1.86	4.48	5.92	2.61	5.88	5.85	2.79	5.06
8	8.31	1.75	7.86	9.89	2.65	9.51	9.74	3.56	8.58
9	5.83	2.05	4.76	7.22	2.93	5.90	6.77	3.66	5.82
10	1.08	1.62	2.61	1.19	1.63	2.63	1.14	1.20	2.03
11	4.20	1.97	6.75	4.99	1.37	5.74	5.22	1.59	6.46
12	1.40	2.14	2.60	1.17	0.80	2.03	0.98	0.68	1.81
13	2.42	1.08	2.51	2.87	1.31	2.95	2.76	1.19	2.71
14	1.98	0.77	2.31	2.05	0.80	2.29	1.86	0.94	2.46
Pooled (1-14)	3.12	1.59	3.63	3.44	1.61	3.88	3.26	1.62	3.54

Notes: Entries in each column give the mean values across all inter-city combinations within the U.S. (labeled U-U), within Canada (C-C), and across the U.S.-Canadian border (U-C), respectively.

(A) In column (1), the measure of volatility is the standard deviation of the relative price series. Column (2) gives the average correlation between the price in city j and city k. In both cases, prices are measured as two-month differences. In column (3), the measure of volatility is the standard deviation of the 12-month ahead forecast error of the relative price, based on a sixth-order uni-variate auto-regression. The sample period is September 1978-December 1997.

(B) These three columns repeat the calculation of column 1 in part A of the table, over the sub-periods 9/78-12/89, 1/90-12/93, and 1/93-12/97, respectively.

(*) The average distance between cities is 1070 miles for the intra-U.S. pairs, 1343 miles for the intra-Canada pairs, and 1428 miles for the cross-border pairs. The standard deviation of the two-month log-difference in the nominal exchange rate is 1.57, 1.56, 1.62, and 1.51 for the full sample, and the 1978-89, 1990-93 and 1994-97 sub-periods, respectively.

Table 3: Updating the Main Regression of Engel-Rogers (1996)

	Specification 1		Specification 2		Specification 3		Specification 4	
Good	Log Distance	Border	Log Distance	Border	Log Distance	Border	Log Distance	Border
All	4.47	1.13	-4.00	-0.27	4.02	1.11	11.4	2.84
	(0.73)	(.008)	(0.68)	(.008)	(0.68)	(.007)	(3.03)	(0.04)
1	6.68	0.88	-4.86	-0.16	6.72	0.86	9.22	2.48
	(1.55)	(.018)	(0.85)	(.008)	(1.62)	(.016)	(2.96)	(0.04)
2	2.10	0.98	-1.03	-0.27	1.95	0.94	7.60	2.70
	(1.54)	(.014)	(0.86)	(.010)	(1.32)	(.012)	(2.97)	(0.04)
3	2.52	0.89	0.05	-0.26	2.15	0.87	0.81	2.30
	(1.89)	(.022)	(0.91)	(0.10)	(1.86)	(.020)	(3.98)	(0.06)
4	10.6	0.96	-5.94	-0.27	9.75	0.93	31.1	2.34
	(1.89)	(.020)	(1.06)	(.013)	(1.75)	(.018)	(5.71)	(0.07)
5	37.4	1.20	-10.1	-0.27	34.1	1.11	55.1	1.99
	(7.57)	(.082)	(2.01)	(.020)	(7.90)	(.079)	(6.95)	(0.09)
6	-2.61	0.83	-0.38	-0.13	-1.20	0.79	-1.59	2.40
	(1.48)	(.014)	(0.83)	(.011)	(1.29)	(.014)	(3.26)	(0.04)
7	3.97	1.28	-1.99	-0.27	4.19	1.10	6.76	2.36
	(4.33)	(.048)	(1.46)	(.017)	(3.87)	(.046)	(4.35)	(0.07)
8	3.34	2.89	-2.08	-0.24	6.44	2.58	18.7	3.39
	(11.0)	(.152)	(1.13)	(.016)	(9.74)	(.137)	(10.1)	(0.12)
9	3.19	0.76	-1.60	-0.09	-0.98	0.69	-5.18	2.07
	(5.14)	(.056)	(1.11)	(.011)	(4.50)	(.049)	(6.38)	(0.08)
10	12.8	1.18	-4.60	-0.39	12.3	1.11	15.8	2.89
	(1.67)	(.017)	(0.93)	(.009)	(1.51)	(.015)	(2.88)	(0.04)
11	15.4	3.19	-0.44	-0.54	17.7	3.15	-0.99	3.81
	(5.29)	(.090)	(0.92)	(.011)	(5.22)	(.086)	(8.81)	(0.11)
12	-2.31	0.90	1.55	-0.26	-2.00	0.87	-0.88	2.52
	(1.81)	(.023)	(1.23)	(.015)	(1.64)	(.021)	(6.34)	(0.07)
13	-0.17	0.77	-0.95	-0.18	0.65	0.74	-1.94	2.49
	(1.86)	(.018)	(1.09)	(.011)	(1.70)	(.017)	(5.60)	(0.056)
14	3.53	0.91	-2.14	-0.28	4.14	0.92	6.77	2.57
	(1.29)	(.017)	(0.89)	(.012)	(1.43)	(.017)	(3.79)	(0.042)
1-14	6.88	1.26	-2.60	-0.26	6.84	1.19	10.1	2.59
	(3.41)	(.041)	(0.48)	(.006)	(3.22)	(.039)	(3.99)	(.049)

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 parenthesis. There are 251 observations in each regression. Coefficients and standard errors on log distance are multiplied by 100. In specification 1, the dependent variable is the standard deviation of the two-month difference in the relative price. Standard deviations are computed over the sample period 9/78 to 12/94, as in Engel and Rogers (1996). In Specification 2, the dependent variable is the correlation between ΔP_{ij} and ΔP_{ik}, the two-month change in the price of good i in cities j and k, respectively. Correlations are computed over the sample period 9/78 to 12/94. Specification 3 is analogous to specification 1, but computes standard deviations over the extended sample period 9/78 to 12/97. In specification 4, the dependent variable is the standard deviation of the time-series of 12-month ahead forecast errors of the relative price, based on a sixth-order uni-variate auto-regression estimated beginning in 3/79 and ending in 12/96.

Table 4: Analysis of the Sub-Periods

	Pre-FTA (1978-89)		1990-93		1994-97	
Good	Log Distance	Border	Log Distance	Border	Log Distance	Border
All	5.10	1.07	2.74	1.24	-0.04	1.13
	(1.00)	(0.01)	(1.11)	(0.01)	(0.99)	(0.01)
1	6.04	0.87	9.31	0.94	8.48	0.70
	(1.52)	(0.02)	(3.13)	(0.04)	(4.06)	(0.03)
2	2.82	0.77	-0.37	1.59	-0.12	0.99
	(1.78)	(0.02)	(1.99)	(0.02)	(1.38)	(0.02)
3	4.05	0.88	0.73	0.86	-3.09	0.79
	(1.95)	(0.03)	(3.61)	(0.03)	(3.08)	(0.03)
4	11.9	0.96	6.21	1.01	1.03	0.92
	(2.38)	(0.02)	(1.77)	(0.02)	(1.50)	(0.02)
5	43.7	1.20	35.9	1.13	17.1	0.76
	(7.78)	(0.08)	(12.1)	(0.12)	(13.2)	(0.09)
6	-2.42	0.77	-0.75	0.88	1.57	0.73
	(4.71)	(0.02)	(2.80)	(0.03)	(3.06)	(0.03)
7	-2.43	1.10	12.9	1.50	5.35	0.69
	(4.71)	(0.05)	(7.79)	(0.09)	(7.41)	(0.09)
8	-2.89	2.72	10.7	3.09	14.8	1.78
	(12.3)	(0.17)	(16.4)	(0.20)	(12.1)	(0.15)
9	2.00	0.80	5.47	0.67	-10.9	0.38
	(5.05)	(0.06)	(8.91)	(0.09)	(6.51)	(0.09)
10	12.2	1.20	16.9	1.12	9.32	0.82
	(1.59)	(0.02)	(3.02)	(0.03)	(2.31)	(0.02)
11	14.6	3.58	14.1	2.41	20.6	2.57
	(6.36)	(0.11)	(6.03)	(0.09)	(9.91)	(0.11)
12	-0.85	0.83	-6.44	1.10	-1.15	0.97
	(1.93)	(0.02)	(1.94)	(0.03)	(1.32)	(0.01)
13	-1.56	0.72	2.10	0.83	6.84	0.62
	(3.19)	(0.02)	(2.92)	(0.04)	(3.35)	(0.04)
14	1.72	0.88	6.34	0.89	7.15	1.10
	(1.58)	(0.02)	(2.51)	(0.03)	(3.47)	(0.04)
1-14	6.35	1.23	8.08	1.29	4.99	1.00
	(3.53)	(0.04)	(4.12)	(0.05)	(3.30)	(0.04)

Notes: These regressions replicate that of column 1 in Table 3, over the sub-periods 9/78-12/89, 1/90-12/93, and 1/94-12/97, respectively.

Table 5: Assessing the Importance of the Border

	Specification 1			Specification 2	
Good	Log Distance	Border	SD of Real Wage	Log Distance	Border
All	4.07 (0.73)	1.11 (.007)	-0.24 (0.45)	7.86 (2.97)	0.71 (.036)
1	5.14 (1.36)	0.84 (.015)	7.34 (1.18)	8.96 (1.48)	0.29 (.015)
2	1.95 (1.40)	0.94 (.013)	-0.02 (1.40)	3.80 (1.26)	0.20 (.010)
3	1.64 (1.98)	0.86 (.020)	2.36 (1.23)	3.54 (1.99)	0.40 (.017)
4	9.97 (1.85)	0.93 (.018)	-1.05 (1.28)	6.37 (1.07)	0.22 (.013)
5	34.4 (8.50)	1.12 (.080)	-2.05 (5.86)	32.2 (7.40)	0.74 (.077)
6	-1.98 (1.31)	0.78 (.014)	3.64 (1.73)	1.56 (1.09)	0.17 (.013)
7	6.54 (4.03)	1.14 (.048)	-9.65 (2.61)	5.66 (3.69)	0.91 (.043)
8	10.7 (10.3)	2.64 (.142)	-17.6 (6.33)	11.2 (8.85)	2.53 (.122)
9	-1.31 (4.77)	0.69 (.050)	1.51 (3.02)	0.83 (4.36)	0.44 (.051)
10	11.2 (1.50)	1.09 (.016)	4.73 (0.93)	9.31 (1.22)	0.36 (.013)
11	15.3 (5.61)	3.11 (.094)	11.0 (6.28)	19.6 (4.98)	2.86 (.085)
12	-2.10 (1.69)	0.87 (.022)	0.46 (1.23)	0.13 (1.17)	0.27 (.015)
13	1.11 (1.88)	0.74 (.017)	-2.13 (1.18)	1.55 (1.60)	0.20 (.016)
14	4.16 (1.51)	0.92 (.017)	-0.06 (1.11)	5.29 (1.38)	0.35 (.016)
1-14	6.91 (3.26)	1.19 (.041)	-0.29 (3.50)	7.86 (3.03)	0.71 (.038)

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 parenthesis. There are 251 observations in each regression. Coefficients and standard errors on log distance and the wage variable are multiplied by 100. Specification 1 adds to specification 3 of Table 3 the standard deviation of the two-month change in real wages. In specification 2, the dependent variable is the standard deviation of the two-month change in the relative real price, estimated over the full sample.

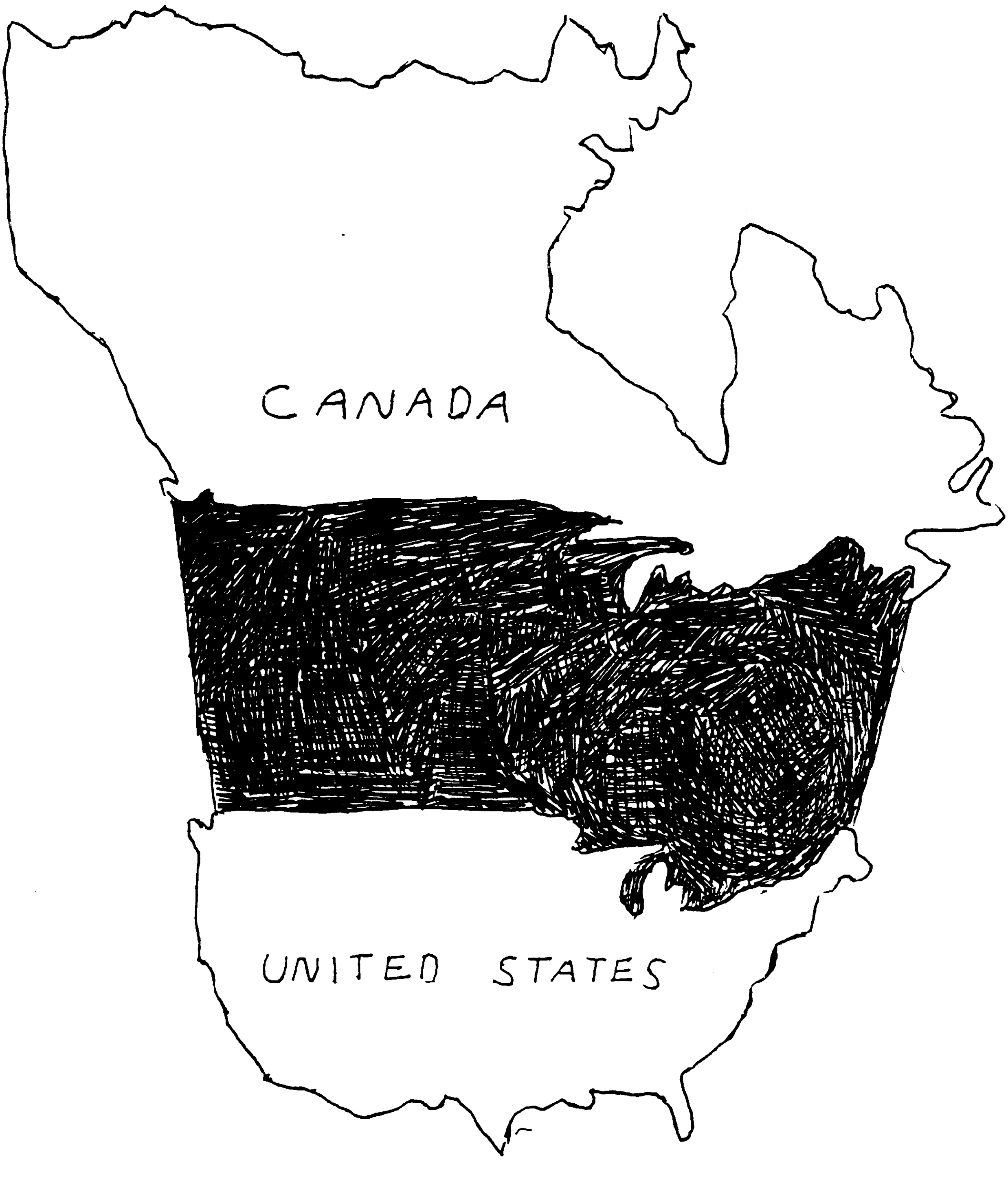

FIGURE 1