BORDER EFFECT OR COUNTRY EFFECT? SEATTLE MAY NOT BE SO FAR FROM VANCOUVER AFTER ALL

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ABSTRACT

This paper reexamines the evidence on the border effect, the finding that the border drives a wedge between domestic and foreign prices. We argue that if there is cross-country heterogeneity in the distribution of within-country price differentials, there is no clear benchmark from which to gauge the effect of a border. In the absence of a structural model or a (natural) experiment it is impossible to separate the "border" effect from the effect of trading with a country with a different distribution of prices. We show that the border effect identified by Engel and Rogers (1996) is entirely driven by the difference in the distribution of prices within the US and Canada.

Keywords: border effect, law of one price, purchasing power parity, price adjustment.
JEL classification: F3, F40, F41.

1 Introduction

In an important paper, Engel and Rogers (1996) estimate the impact of an international border on price dispersion across US and Canadian cities. After controlling for distance and other factors, they conclude that the economic impact of the border between the U.S. and Canada is equivalent to shipping a good 75,000 miles (Table 3, p. 1117, Engel and Rogers (1996)).[1] Numerous subsequent studies estimate similarly impressive border coefficients. Parsley and Wei (2001) find a larger border effect in U.S.-Japanese price data. Using data on quantities, McCallum (1995) finds that intranational trade flows are, ceteris paribus, 22 times larger than international trade flows. In a similar vein, *provincial* borders in Canada (Helliwell and Verdier, 2001) and *state* borders in the United States (Wolf, 2000) account for a significant fraction of the decreased trade flows across provinces and states relative to trade flows within states and provinces. Furthermore, Ceglowski (2003) finds that *provincial* borders in Canada account for a significant fraction of the discrepancy of prices across provinces. The presence of large border effects has important welfare implications and Obstfeld and Rogoff (2000) include the border effect in their list of the major puzzles in international economics.

Evidence of a border effect in and of itself is not surprising. However, the magnitude of these estimated border effects is surprisingly – many would say unbelievably – large. If the results are to be taken literally, it is odd that the U.S. and Canada remain important bilateral trading partners in the presence of such a large barrier to trade. The border coefficient between the U.S. and another of its key trading partners, Japan, is estimated by Parsley and Wei (2001) to be equivalent to 43,000 trillion miles (note that the distance to the Moon is a mere 238,900 miles). Likewise, the findings of large impediments to trade between states and provinces *within* a country are also difficult to believe.[2] Finding a significant border effect where it should not be found, and finding coefficients that are orders of magnitude larger than one can plausibly defend, raises doubts about the validity of the empirical methodology used to isolate the border effect.

In this paper we reexamine the identification strategy employed in Engel and Rogers (1996) (henceforth ER) and other studies that estimate border frictions. In this line of research, border effects are measured by the difference between the within-country dispersion and the cross-country dispersion of some economic variable, such as the price of a particular commodity,

[1] If citation count is an indication of impact, Engel and Rogers (1996) is one of the most frequently cited studies published by the *American Economic Review*. According to the Social Sciences Citation Index the paper is in the very top percentiles of papers published in the top tier of economics journals.

[2] For example, there are explicit legal norms (e.g., U.S. constitution) prohibiting limitation of intranational trade.

the speed of convergence to the law of one price, or trade volumes, to list just a few examples. A perceived advantage of this metric for border frictions was that it provided a model-free, easy-to-interpret summary statistic of the magnitude of international frictions. We argue that in many empirically relevant cases this measure of border frictions is contaminated with factors unrelated to the border. For example, if relative price variability across locations within the same country differs systematically country by country, the border effect measured by a regression comparing within-country and cross-country price dispersion will be confounded by the divergence between two countries' internal price distributions. We call this the *country heterogeneity effect.* Moreover, we show that, when there is country heterogeneity, border effect type regressions cannot separate the border frictions from the effect of trading with a country with a different distribution of prices.

The following example illustrates the pitfalls of estimating border equations. Consider a dataset with price data from four cities: Seattle, Chicago, Vancouver and Calgary. Suppose that, after conditioning on distance and adjusting for the exchange rate, the price of some good *k* in Vancouver and Calgary is identical, but the price of the good is quite different in Seattle and is yet more different in Chicago. If one were to test whether the pair-wise price differentials of intra-country prices (in this example Vancouver-Calgary and Chicago-Seattle) were statistically significantly different from the cross-border prices (Vancouver-Chicago, Vancouver-Seattle, Calgary-Seattle and Calgary-Chicago) one would find that there are wider deviations from parity in the cross-border pairs than in the sample of within-country pairs. However, that difference would be driven by the disparate behavior of prices in the US (Seattle and Chicago). It would have nothing to do with the friction of crossing a border. One could, of course, include city dummies to pick up the dispersion of prices across U.S. cities. However, as we will show below, when the dispersion across cities is larger in one country relative to the other, it becomes impossible to separate cross-country differences from border frictions. This example is not just a hypothetical pathology. Our statistical analysis will show that this example is a reasonable characterization of US-Canada data. The variance across cities within the U.S. is much greater than the variance across cities within Canada. Given this country heterogeneity, a border regression based on comparing prices in the U.S. and Canada will produce a U.S./Canada border effect driven by the high variance within the U.S.

To illustrate the problem of isolating the true impact of an international border, we develop a simple model of trade with transportation costs and location-specific cost shocks. We simulate the model to make a few points: 1) country heterogeneity can have a critical effect on measures of border frictions; 2) the methodology typically used in the literature may yield a significant border coefficient even when no border friction is present, 3) the border coefficient estimated from one country's perspective is not sufficient for establishing that there is a border effect; 4) within-country variation can be a poor benchmark for measuring the border effect because intra- and international price dispersion are determined simultaneously with very important general equilibrium effects. The (counter) examples we present are too simple to be taken as definitive models of price-setting behavior across locations, but these examples demonstrate that estimating a border coefficient is a joint test of the existence of a border friction and the structural model that generates prices in different locations. The overriding message is that the border coefficient that emerges from tests comparing within-country prices to cross-border prices tells us little about actual border effects in the absence of a fully articulated structural model or a (natural) experiment.

Although we focus on identification of the border effect in price data, our critique applies to other analyses of border frictions. Our arguments suggest that a comparison of within- and cross-country statistics is insufficient for identifying border frictions. This is true irrespective of whether one compares aggregated measures such as price indices or highly disaggregated measures such as unique product code prices or trade flows. Likewise, comparisons of derivative statistics, such as the speed of convergence to the law of one price, do not yield a causal effect of crossing the border. Furthermore, temporal variation in an estimated border coefficient based on differences between within- and cross-country statistics does not necessarily mean that there is variation in the size of border frictions. Identification of the border effect is much more subtle and requires credible theory-based restrictions or (natural) experiments. This paper is a warning to users of reduced-form border effect regressions that the conventional interpretation of estimated coefficients can be highly misleading.

The structure of the paper is as follows. In section 2, we discuss the econometric specification that is conventionally used to estimate the border effect using U.S./Canada price data analyzed in previous studies and briefly replicate the results reported in ER. In section 3, we discuss potential identification problems (specifically, the country heterogeneity effect) in using the standard econometric specification to estimate the border effect. In Section 4, we show the sensitivity of the border effect to the presence of cross-country heterogeneity. In Section 5, we present some numerical examples to illustrate the effect of cross-country differences in trade costs on the estimation of a border coefficient. We conclude in Section 6.

2 Within- and Cross-Country Price Variation

Define P_{it}^{k} as the price of good *k* in location *i* at time *t* and S_t as the exchange rate that converts prices from location *j*'s currency to location *i*'s currency. Then $q_{ijt}^{k} = \ln(S_t P_{it}^{k} / P_{jt}^{k})$ is the log of the real exchange rate. The variable of interest in ER and in much of the subsequent literature on border effects is the standard deviation of the real exchange rate variation over time for a given good *k* between locations *i* and *j*. We denote this standard deviation with $\sigma(q_{ijt}^{k})$. Note that after collapsing the time series information into $\sigma(q_{ijt}^{k})$, we have only the cross-sectional variation in the volatility of the real exchange rates across goods and locations.

Because measured prices in this literature are typically price indices rather than actual transaction prices, ER and the subsequent literature use the standard deviation of *changes* in the real exchange rate. Taking differences in the real exchange rate is appropriate for price indices and implies that one is testing relative rather than absolute purchasing power parity. Differencing the data also helps reduce the persistence of the real exchange rate.

To give the reader a sense of the magnitude of the volatility of real exchange rates, we use the data set studied by ER and report a few summary statistics in Table 1. ER's data cover 14 categories of goods, nine Canadian cities and 14 U.S. cities. We denote pairs of cities by *UU* (*US-US*), *UC* (*US-Canada*) and *CC* (*Canada-Canada*). The time span for each city pair varies from 1978-1992 at minimum to 1976-1995 at maximum.[3] Because for some cities the price data are released bimonthly, we follow ER and use differences over two-month intervals so that the main variable in our analysis is $\sigma(\Delta_2 q_{ijt}^{k})$ where $\Delta_2 x_t = x_t - x_{t-2}$. Examination of the table immediately indicates that, on average, *CC* pairs are considerably less volatile than *UU* pairs and treating these two groups as homogenous may be inappropriate. For example, $\sigma(\Delta_2 q_{ijt}^{k})$ equals 0.031 for *UU* pairs and 0.016 for *CC* pairs.[4] It is especially remarkable that the volatility of *UU*

[3] Detailed discussion of the data can be found in Engel and Rogers (1996). The data are available at Charles Engel's website: http://www.ssc.wisc.edu/~cengel/Data/Border/BorderData.htm.

[4] Engel and Rogers (footnote 8, p. 1116, 1996) note that the dispersion of price differentials is different for the U.S. and Canada. They, however, do not elaborate on possible implications of such a difference.

pairs is similar to the volatility of *UC* pairs, given that the latter reflects the variability of exchange rates in addition to the dispersion of prices across cities.[5]

Because distance is typically a robust determinant of trade frictions and hence price dispersion, we also illustrate the distribution of $\sigma(\Delta_2 q_{ijt}^k)$ for each type of city pair, unconditional and conditional on distance (Figure 1). We depict the distribution using a box plot, which shows the median (the bar in the middle of the rectangle), 25th and 75th percentiles (the base and top of the rectangle) as well as the range of the data (the length of the line). Note that the standard deviation conditional on distance need not be positive. After controlling for distance, we still find that *CC* pairs systematically have smaller volatility than *UU* and *UC* pairs which have similar volatility levels.

We also observe similar patterns when we use alternative measures of price dispersion. Consider first the volatility of price differentials measured by the relative exchange rate $RER_{ijt}^k = \left((P_{it}^k / P_{it}) \big/ (P_{jt}^k / P_{jt})\right)$ where P_{it}^k is the price of good k in city i at time t, P_{it} is the Consumer Price Index for city i at time t. The advantage of this measure is that it does not depend on the volatile nominal exchange rate. In this case, *UU* and *UC* pairs have practically identical volatilities while *CC* pairs are approximately 50% less volatile. Likewise, if we use residuals from regressions of the real exchange rate q_{ijt}^k on its lags, we still find that volatility for *UU* and *UC* pairs is roughly the double of volatility for *CC* pairs.

To compare differences in the volatility of real exchange rates across types of city pairs after controlling for observable and time invariant characteristics of city pairs, ER employ the following specification:

$$\sigma(\Delta_2 q_{ijt}^k) = \beta_0 + \beta_1 \ln d_{ij} + \beta_2 Border_{ij} + \phi_k + \alpha_i + \alpha_j + \varepsilon_{ij}^k , \qquad (1)$$

where d_{ij} is the distance between locations i and j, $Border_{ij}$ is a dummy variable equal to one if locations are separated by a border and zero otherwise, α_i is the cost of trade specific to location

[5] The large volatility of *UU* pairs relative to *CC* pairs suggests that there are large idiosyncratic departures from the law of one price across US cities, and that these departures are dramatically larger in the US than in Canada. This asymmetry between the US and Canada remains an important puzzle for future research. Possible explanations could be differences in pricing strategies across countries and differences in compositions of consumed goods (e.g., Canada could have a greater overlap in consumption baskets across locations than the US; see section 5 for an illustration). The differences in statistical practices appear to be too minor (see e.g. OECD 2002) to explain so different levels of within-country price volatility in the US and Canada.

i, ϕ_k is the cost of trade specific to good k, ε_{ij} is a random time-invariant component in the cost of trade between locations i and j. Here distance d_{ij} proxies for shipping costs, costs of acquiring information, etc.

ER and the literature that followed in this vein interpret the coefficient on the border dummy as a measure of frictions associated with crossing the border, i.e., the border effect. To provide a sense of the "width" of the border, ER use the distance equivalent of the border effect:[6]

$$Border\ Effect = \exp(\beta_2 / \beta_1)\ . \tag{2}$$

In the next section, we reexamine identification of the border effect in specification (1) and find that the border dummy does not necessarily measure the border effect and that it is possible to find a large statistically significant coefficient on the border dummy when there is no friction associated with crossing the border. To highlight the difference between actual and measured border effects, in subsequent discussion we distinguish between the genuine border friction, which we call the "border effect," and the estimated coefficient on the border dummy, which we will refer to as the "border coefficient."

3 Identification of the border effect

For concreteness, suppose that the two countries are Canada and the U.S. and cities are the relevant geographical units. The border coefficient in specification (1) measures how much a *ceteris paribus* transition from an intra-national city pair (*UU* or *CC*) to an international city pair (*UC*) raises the average volatility of the price differential. Intuitively, if the border coefficient correctly identifies the border effect, it should pick up the variation that is specific to *UC* pairs, and should not include factors that stem from differences between *UU* and *CC* pairs. To put this in the language of the program evaluation/treatment effects literature, the border coefficient should pick up the effect associated with the treatment of crossing the border. If, in the absence of the border treatment, city-pairs differ for other (non-border related) reasons, one should condition on this heterogeneity to isolate the effect of the treatment. The treatment effect

[6] Parsley and Wei (2001) suggest an alternative measure of the border effect: $\overline{d} \times \exp(\beta_2/\beta_1 - 1)$ where $\overline{d}$ is the average distance between cities. Since this measure is a monotonic transformation of $\exp(\beta_2/\beta_1)$, our qualitative conclusions do not change if we use this alternative measure. Likewise, since variation in β_1 across specifications is relatively small, our qualitative conclusions do not change if we focus only on β_2.

literature is replete with examples of how ignoring such heterogeneity can bias the estimate of the treatment effect (see Smith (2004) for a discussion and examples).

We return to the ER specification to show how the dispersion for each city pair is decomposed into a city and country contribution. To simplify exposition, we omit for now the error term and controls for distance and commodities. For all pairs, the volatility of the real exchange rate is described by:

$$\sigma_{ij} = \beta UC_{ij} + \gamma_U UU_{ij} + \gamma_C CC_{ij} + \sum_{s=1}^{N} \alpha_s D_s , \tag{3}$$

where *UC*, *UU*, *CC* are dummy variables for *UC*, *UU*, and *CC* city pairs, D_s is a city dummy equal to one if $s = i$ or $s = j$ and zero otherwise and N is the number of cities. Equation (3) shows that the volatility of prices can be orthogonalized into four components: β, γ_U, γ_C and the city dummies. The coefficient γ_U reflects institutional and structural features of the data specific to cities in the United States, γ_C reflects institutional and structural features of the data specific to cities in Canada, the city dummies reflect variation that is city-specific and not captured in the country effects, and β reflects the variation specific to crossing the border.

Without loss of generality, suppose that 1,…,k cities are in Canada. The average volatility for types of city pairs is then given by:

US-Canada: $$\bar{\sigma}_{UC} = \beta + \bar{\alpha}_U + \bar{\alpha}_C , \tag{4}$$

US-US: $$\bar{\sigma}_{UU} = \gamma_U + 2\bar{\alpha}_U , \tag{5}$$

Canada-Canada: $$\bar{\sigma}_{CC} = \gamma_C + 2\bar{\alpha}_C , \tag{6}$$

where $\bar{\alpha}_C = \frac{1}{k}\sum_{s=1}^{k} \alpha_s$ is the average city effect for Canadian cities and $\bar{\alpha}_U = \frac{1}{N-k}\sum_{s=k+1}^{N} \alpha_s$ is the average city effect for US cities. Since by definition city effects reflect city specific deviation from national characteristics, the average city effect can be set to zero in each country, i.e., $\bar{\alpha}_C = \bar{\alpha}_U = 0$.

Unfortunately, while equation (3) provides an exact decomposition of variance in theory, we cannot estimate equation (3) in practice because the *UU* and *CC* dummies are collinear with the set of city dummies and the *UC* dummy so that $(\beta, \gamma_U, \gamma_C)$ and $\{\alpha_s\}_{s=1}^{N}$ are not identified separately. Specifically,

$$CC_{ij} = -\tfrac{1}{2}UC_{ij} + \tfrac{1}{2}\sum\nolimits_{s=1}^{k} D_s ,$$

$$UU_{ij} = -\tfrac{1}{2}UC_{ij} + \tfrac{1}{2}\sum\nolimits_{s=k+1}^{N} D_s . \tag{8}$$

Equations (7) and (8) make clear that the country-specific component of price variation shows up in two places, in the cross-border effect and in the city effects. ER proceed by substituting equations (7) and (8) into equation (3), effectively eliminating the country dummies CC_{ij} and UU_{ij} so that:

$$\begin{aligned}
\sigma_{ij} \quad &= \beta UC_{ij} + \gamma_U UU_{ij} + \gamma_C CC_{ij} + \sum\nolimits_{s=1}^{N} \alpha_s D_s \\
&= \beta UC_{ij} + \gamma_U(-\tfrac{1}{2}UC_{ij} + \tfrac{1}{2}\sum\nolimits_{s=k+1}^{N} D_s) + \gamma_C(-\tfrac{1}{2}UC_{ij} + \tfrac{1}{2}\sum\nolimits_{s=1}^{k} D_s) + \sum\nolimits_{s=1}^{N} \alpha_s D_s \\
&= [\beta - \tfrac{1}{2}(\gamma_U + \gamma_C)]UC_{ij} + \sum\nolimits_{s=1}^{k} \left(\tfrac{1}{2}\gamma_C + \alpha_s\right) D_s + \sum\nolimits_{s=k+1}^{N} \left(\tfrac{1}{2}\gamma_U + \alpha_s\right) D_s .
\end{aligned} \tag{9}$$

The country heterogeneity effect is now absorbed partially in the city-fixed effects but it also appears in the border coefficient. Under this specification the coefficient on the border coefficient (the term in the first bracket) measures the increase in cross-border volatility relative to the *average* volatility of the intra-country pairs (this benchmark is the omitted category).

Why isn't the average of intra-country volatility a good benchmark for measuring the price impact of crossing the border? Take the case where $\gamma_C < \gamma_U$, i.e. prices within Canada have very little dispersion relative to prices within the U.S. As shown in the first row of Table 1, the standard deviation of the real exchange rate of Canadian city pairs is in fact about half the volatility of that for US city pairs, and the dispersion of cross-border pairs is similar to that of US city pairs.[7] Using the ER specification, the estimated border coefficient – relative to the two-country average of intra-national volatilities – is large and positive, even when from the U.S. perspective crossing the border has almost no impact on the variance of the real exchange rate. To understand what is going on, suppose one were to ignore the information on intra-Canadian prices, and simply compare US-US pairs with cross-border pairs. In this case, the border coefficient would be small. However, one cannot tell whether the small border coefficient means that crossing the border has a minimal impact on volatility, or whether it reflects the fact that

[7] This is true not only for the pooled sample of prices, but also for 11 of the 14 individual goods categories considered by Engel and Rogers.

Americans are trading with Canadian partners that have less dispersed prices. From the perspective of Canadians, crossing the border dramatically increases the variance of the real exchange rate. If one were to repeat the exercise and ignore the intra-US data, the estimated border coefficient based on *CC* and *UC* pairs would be extremely large. However, one does not know whether the large coefficient occurs because crossing the border actually leads to greater price dispersion, or whether this is because Canadians are trading with US cities that have more disparate prices. The point is that one needs *both* countries' price variation to identify the border. Furthermore and most importantly, one needs guidance from economic theory to assign the appropriate weight to each country's price variability.

In Section 5, we show that the choice of weights is far from obvious even in a standard trade model. In fact, numerical simulations in which there is no border friction but there is cross-country heterogeneity suggest that the best benchmark could put all weight on the country with the smallest difference between within- and cross-country price volatility. Even this highly conservative estimate can overstate the magnitude of the border effect. In addition, within-country variation itself may be affected by the imposition of a border further compromising within-country variation as a good benchmark for the border. That is, within-country price dispersion may serve as a poor benchmark because price dispersion between and within countries is determined simultaneously in general equilibrium and therefore one can grossly overstate or understate the size of the border friction because of endogenous amplification or attenuation effects. Hence, in the absence of a structural model that identifies particular control variables, such as factors affecting markups, transportation, and production costs, or a controlled/natural experiment, in which a country moves from autarky to trade, we do not know if the arithmetic average, or any other combination of within-country variances, is the appropriate benchmark for evaluating the effect of the border.[8]

We propose an alternative decomposition into country and city effects. To be clear, the alternative decomposition does not solve the weighting problem – rather it reveals the magnitude of the cross-country heterogeneity problem by explicitly estimating the country dummies.

[8] Although natural experiments are unlikely to provide truly random borders, they can supply alternative sources of identification. For example, Morshed (2007) utilizes historical price data to investigate how splitting Pakistan into Pakistan and Bangladesh affected price dispersion between locations that became separated by the border. He finds that price dispersion between Pakistan and Bangladesh was different from the price dispersion within Bangladesh and Pakistan before and after the split. Thus, even with no explicit border price dispersion can be different across regions from the price dispersion within regions. Furthermore, he reports that cross-region price dispersion did not change much when regions became separate countries. Results of this study strongly indicate that within country price dispersion could be a poor benchmark for cross-border price dispersion (see also our Monte Carlo simulations in Section 5).

Denote deviations of the city effect from the national mean with $\hat{\alpha}_s = \alpha_s - \bar{\alpha}_U$ if s is a US city and $\tilde{\alpha}_s = \alpha_s - \bar{\alpha}_C$ if s is a Canadian city. Using equations (7) and (8), one can rearrange terms in equation (3) as follows:

$$
\begin{aligned}
\sigma_{ij} &= \beta UC_{ij} + \gamma_U UU_{ij} + \gamma_C CC_{ij} + \sum\nolimits_{s=1}^{N} \alpha_s D_s \\
&= \beta UC_{ij} + \gamma_U UU_{ij} + \gamma_C CC_{ij} + \sum\nolimits_{s=1}^{k} (\alpha_s - \bar{\alpha}_C + \bar{\alpha}_C) D_s + \sum\nolimits_{s=k+1}^{N} (\alpha_s - \bar{\alpha}_U + \bar{\alpha}_U) D_s \\
&= \beta UC_{ij} + \gamma_U UU_{ij} + \gamma_C CC_{ij} + \sum\nolimits_{s=1}^{k} \tilde{\alpha}_s D_s + \bar{\alpha}_C \sum\nolimits_{s=1}^{k} D_s + \sum\nolimits_{s=k+1}^{N} \hat{\alpha}_s D_s + \bar{\alpha}_U \sum\nolimits_{s=k+1}^{N} D_s \\
&= \beta UC_{ij} + \gamma_U UU_{ij} + \gamma_C CC_{ij} + \sum\nolimits_{s=1}^{k} \tilde{\alpha}_s D_s + \bar{\alpha}_C (2CC_{ij} + UC_{ij}) + \sum\nolimits_{s=k+1}^{N} \hat{\alpha}_s D_s + \bar{\alpha}_U (2UU_{ij} + UC_{ij}) \\
&= (\beta + \bar{\alpha}_C + \bar{\alpha}_U) UC_{ij} + (\gamma_U + 2\bar{\alpha}_U) UU_{ij} + (\gamma_C + 2\bar{\alpha}_C) CC_{ij} + \sum\nolimits_{s=1}^{k} \tilde{\alpha}_s D_s + \sum\nolimits_{s=k+1}^{N} \hat{\alpha}_s D_s \\
&= b_{UC} UC_{ij} + b_{CC} CC_{ij} + b_{UU} UU_{ij} + \sum\nolimits_{s=1}^{k} \tilde{\alpha}_s D_s + \sum\nolimits_{s=k+1}^{N} \hat{\alpha}_s D_s \\
&= const + (b_{UC} - b_{UU}) UC_{ij} + (b_{CC} - b_{UU}) CC_{ij} + \sum\nolimits_{s=1}^{k} \tilde{\alpha}_s D_s + \sum\nolimits_{s=k+1}^{N} \hat{\alpha}_s D_s \,, \qquad (10)
\end{aligned}
$$

By construction $\sum\nolimits_{s=1}^{k} \tilde{\alpha}_s = 0$ and $\sum\nolimits_{s=k+1}^{N} \hat{\alpha}_s = 0$, so dummies CC_{ij} and UU_{ij} are not collinear with other right-hand side variables.[9] In the last line, we use *UU* pairs as the (omitted) benchmark.

Because rearrangement in equations (10) and (9) is purely *algebraic*, equations (10) and (9) have the same explanatory power (e.g., the same R^2) and one can back out coefficients in equation (10) from equation (9) and vice versa. The main advantage of equation (10) is that b_{UC}, b_{UU}, and b_{CC}—the coefficients on *UC*, *UU* and *CC*—now measure the average volatility for US-Canada, US-US, and Canada-Canada city pairs so that the estimated coefficients are directly related to the objects of interest in equations (4)-(6). Given the convention that $\bar{\alpha}_C = \bar{\alpha}_U = 0$, the differences $b_{UC} - b_{CC} = \beta - \gamma_C$ and $b_{UC} - b_{UU} = \beta - \gamma_U$ provide meaningful estimates of increases in the volatility of the real exchange rate when one goes from an intra-national city pair to an international city pair. To estimate equation (10), we augment specification (1) with the *CC* dummy as follows:

$$\sigma(q_{ij,t}^{k}) = \beta_0 + \phi_k + \alpha_i + \alpha_j + \beta_1 \ln d_{ij} + \beta_2 Border_{ij} + \beta_3 CC_{ij} + \varepsilon_{ij}^{k} \,, \qquad (11)$$

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[9] Because city effects sum up to zero, the addition of the *CC* dummy does not require that we drop an arbitrary city dummy. Also note that if average city effects are not zero then identification of the border effect is even more complicated.

where *Border* is the dummy variable for US-Canada pairs, city-specific effects are constrained as in equation (10), β_2 measures the increase in the dispersion in the transition from a US-US pair to US-Canada pair and $\beta_2 - \beta_3$ measures the increase for the transition from a Canada-Canada pair to a US-Canada pair holding everything else constant. In general, the benchmark (untreated) group is heterogeneous (i.e., $\beta_3 \neq 0$) and, therefore, the researcher should compute the border coefficient relative to each of the US-US and Canada-Canada benchmarks. To reiterate, specification (11) exposes but does not resolve the identification problem we discuss above.

We next show that there is a significant difference between within-country volatility measures for the Canada and the U.S., and failing to account for that difference dramatically affects one's view of the magnitude of the border.

4 Re-examination of the Border Effect Estimates

To replicate ER's results, we estimate specification (1) and report results in the first column of Table 2. To preserve space and keep the analysis focused, in what follows we report only results for regressions pooling across goods. We find that that the coefficients on the border dummy and distance are all positive. Specifically, in the first panel the estimates are β_1=1.076 and β_2=12.026 so that the border effect is equivalent $\exp(\beta_2 / \beta_1) = 71{,}438$ km, consistent with ER's estimates (see Table 3 in ER).[10]

ER argue that the large border coefficient measured by $\exp(\beta_2 / \beta_1)$ cannot be completely explained by the volatility of the nominal exchange rate. They find that the border coefficient remains large when price differentials are measured by the relative exchange rate RER_{ijt}^k. We reproduce their large estimate of the border coefficient for $\sigma(\Delta_2 RER_{ijt}^k)$ in column 2, Table 2. In this case, the implied distance equivalent is 845 km.

By pooling *UU* and *CC* pairs, the researcher estimates a border coefficient that is mixed with country heterogeneity. To demonstrate how important this fact is, we estimate specification (11) and report the results in the first column of Table 3. We find that the key coefficient β_2 in specification (1) is sensitive to the inclusion of the *CC* dummy. In particular, the coefficient β_2 for the pooled regression remains statistically significant but drops from 12.026 to 4.148 thus

[10] Our point estimates of β_2, β_1 are close to ER's. However, because the border effect is Border Effect = $\exp(\beta_2/\beta_1)$, even a minor variation in estimates of β_2, β_1 can result in large changes in the estimates of the border effect. In any case, the order of magnitude is the same.

making the distance equivalent of the border effect implied by the border coefficient, here estimated from the US perspective, fall from 71,438 km to 47 km. By the same token, one can estimate a border coefficient of 108 million km for Canada.

The finding that the border coefficients are so dramatically different for Canada and the US is a direct consequence of the large differences in σ_{UU} and σ_{CC} – the within-country deviations from the law of one price. The border coefficient for the US is small because US-Canada price differentials have about the same dispersion as US-US price differentials. But from this data alone it is impossible to tell whether this is because the border itself is unimportant, or because cross border trade reflects transactions with Canadian cities that happen to exhibit very little cross-city price dispersion. Conversely, the very large border coefficient for Canada could be because crossing the border is costly, or because Canadians are trading with US cities that exhibit more price dispersion than Canadian cities do. Because we do not know how to combine within-country price distributions to obtain a clean benchmark for isolating the border effect, the strongest, and not very informative, statement we can make in this framework is that the point estimate of distance equivalent of the border coefficient is somewhere between 47 kilometers and 108 million kilometers.[11] More importantly, because within this framework we cannot separate country heterogeneity and the border effect, we cannot rule out that the genuine border effect is zero.

As we show in Section 3, under the ER specification differences in the volatility of real exchange rate for *UU* and *CC* pairs are absorbed by the city dummies. Inspection of coefficients on the city dummies in Table 4 shows the relationship between country and city effects. Canadian cities (the shaded rows) are systematically negative while the coefficients on city dummies for U.S. cities are systematically positive. Once we control for the differences across countries (specification (11); Table 4, column 2), the volatility of *UC* and *CC* pairs is measured by the coefficient on the *Border* and *CC* dummy and the coefficients on city dummies do not have any pattern. We should note that it is not possible to distinguish on econometric grounds between the ER specification and the one we propose because they are just rearrangements of one another. However, we believe that given the nature of the country-specific variation in the data, our specification is more informative since it reveals this heterogeneity. At a minimum, our

[11] There are large differences across goods in terms of the magnitude of the deviations from the law of one price and the ratios of σ_{UC} to σ_{UU} and σ_{CC}. The addition of a country dummy changes the estimated border coefficients dramatically for 11 of the 14 goods. In nine of those cases, the "border" from the Canadian perspective is larger than from the US perspective. In two cases the US "border" appears larger. In the three cases where the country dummy had no effect, the border coefficient is estimated to be zero under either the ER specification or ours.

results suggest that the estimated border coefficients are highly sensitive to the assumptions one makes about the source of variation and the relevant benchmark for measuring the border effect.

We reach the same conclusion for the relative exchange rate RER^k_{ijt}: ER's findings for the relative exchange rate are also affected by the country heterogeneity effect. Indeed, when the *CC* dummy is included (column 2, Table 3), distance equivalent of the border coefficient shrinks from 845 km to less than zero from the US perspective and increases to 2.3 million km from the Canadian perspective. The true border effect is under-identified.

5 A Model of Country Heterogeneity and the Border Effect

We have argued that the border regression typically used to measure the importance of cross-border deviations from the law of one price hinges critically on the extent of deviations from the law of one price *within* countries. In this section we consider a simple model of trade in the spirit of Eaton and Kortum (2002) and Anderson and van Wincoop (2004) and investigate whether within-country variation is an appropriate benchmark. We show that in presence of country heterogeneity the standard border regression methodology yields positive coefficients on the border dummies even when no border friction exists. We also show that the correct identification of the border varies across models – thus there is no easy, model-free way of isolating the border effect from price differentials alone.

Consider two countries *U* and *C* that trade in a variety of goods. Each country has *N* cities. Consider good *k*. Let x_{it} be the log cost of producing the good in location *i* at time *t*. Suppose that the cost evolves according to $x_{it} = \rho x_{i,t-1} + \upsilon_{it}$ where $\upsilon_{it} \sim iid\ N(0,\theta^2)$ are location-specific shocks uncorrelated across goods and locations. The price of buying from another location includes the cost of producing the good at that location and the cost of trade between locations. Consumers will search across all locations, including their own city location, and will buy each good from its lowest cost (inclusive of the trade cost) source: so that the price paid in location *i* is given by:

$$p_{it} = \min\{x_{1t} + \tau_{1i}, \ldots, x_{i-1,t} + \tau_{i-1,i}, x_{it}, x_{i+1,t} + \tau_{i+1,i}, \ldots, x_{Nt} + \tau_{Ni}\} \tag{12}$$

where τ_{ij} is the cost of trade between locations *i* and *j*. In the numerical simulations, we look at the price distribution across locations for a randomly chosen good *k* drawn from the continuum of goods.

To make this theoretical model comparable to our empirical specification, we assume that the cost of trade is $\tau_{CC} + \alpha_i + \alpha_j$ for *CC* pairs and $\tau_{UU} + \alpha_i + \alpha_j$ for *UU* pairs where α_i is idiosyncratic cost of trading with location *i*. By setting $\tau_{CC} < \tau_{UU}$, the model will generate greater price dispersion in *UU* pairs than in *CC* pairs. The cost of trading with an international partner is a function of the intranational trading costs and a border effect, $\tau_{UC} = \eta\tau_{CC} + (1-\eta)\tau_{UU} + B + \alpha_i + \alpha_j$ for *UC* pairs. Parameter η measures the relative importance of *intra*national trade costs when applied to *inter*national trade. The border effect *B* captures all impediments to international trade. Setting $B = 0$ corresponds to no border effect. In the experiments below we vary the specification of the cross-border trading costs by changing η and *B*. We require a non-degenerate distribution of α_i to ensure that the distribution of prices across cities is non-degenerate. Without this assumption there is no price dispersion because all cities within a country will buy from the same location. We draw α_i from U(0,*D*), where *D* is a positive constant.

Because there is no closed-form joint density of prices for any two locations,[12] we simulate the model and compute temporal dispersion of price differentials for types of city pairs. We set $N = 25$, $T = 1000$ (number of time series observations), $\tau_{CC} = 0.1$, $\theta = 0.15$, $\rho = 0.9$, and $D = 0.2$. This calibration ensures that cities buy from locations in other countries. We vary τ_{UU} from 0.1 to 0.5 to show how country heterogeneity affects the estimate of the border effect. We consider two scenarios: $\eta = 0$ and $\eta = 1$.[13]

First, we want to analyze the no-border-effect benchmark, i.e., the case with $B = 0$. Once prices are computed according to (12), we calculate real exchange rates between locations, i.e., $q_{ijt} = p_{it} - p_{jt}$. Then we compute the $\sigma(q_{ijt})$ and compare the standard deviations across types of city pairs (i.e., *CC*, *UU*, *UC*). We plot the distributions of the resulting price dispersion in Figure 2 (Panel A) illustrates the case when $\eta = 0$ (international trade costs are equal to *U*'s internal trade cost) and Panel B illustrates $\eta = 1$ (international trade costs are equal to *C*'s internal trade cost). We report mean price dispersion by types of city pairs and by cost of trade τ_{UU} in columns (2)-(7), Panels A and B, Table 5. The table and figures show that the dispersion of prices for *UU* pairs (i.e., within country *U*) and for *UC* pairs (i.e., for international pairs of cities in countries *U* and *C*) increase as τ_{UU} increases. Moreover, price dispersion for *UC* pairs increases in line with price dispersion for *UU* pairs.

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[12] In short, to estimate the standard deviation of price dispersion, we need to know joint distribution of extreme value statistics from two correlated distributions (options available to cities are clearly correlated since they can buy from the same set of locations). Finding such joint distribution is a very hard task.

[13] Results are qualitatively similar when $\eta \in (0,1)$.

We learn several lessons from this simulation. First, one can find that the border coefficient is positive and significant based on differences in price dispersion even when there is no border effect. Specifically, $\bar{\sigma}_{UC} \approx \max\{\bar{\sigma}_{UU}, \bar{\sigma}_{CC}\}$ and international price dispersion is driven by within-country price dispersion in country *U*. ER's method ($BE = \bar{\sigma}_{UC} - \frac{1}{2}(\bar{\sigma}_{UU} + \bar{\sigma}_{CC})$) finds a large border coefficient when there is no border effect and the bias in ER's estimate increases with country heterogeneity, i.e., with τ_{UU} and, consequently, with $\bar{\sigma}_{UU}$. ER's estimate appears to be correct only if there is no country heterogeneity, i.e., $\tau_{UU} = \tau_{CC}$. Second, there could be a very large border coefficient from one country's perspective even when there is essentially zero border coefficient from the perspective of the other country. Columns (6) and (7) show the two border coefficients for *U* and *C*: $b_{UC} - b_{CC} = \bar{\sigma}_{UC} - \bar{\sigma}_{CC}$ and $b_{UC} - b_{UU} = \bar{\sigma}_{UC} - \bar{\sigma}_{UU}$. Clearly, from *C*'s perspective, the border effect measured by the appropriate border coefficient is enormous, while from *U*'s perspective there is no border effect. This case is similar to the pattern we observe in the US-Canada case when *CC* dispersion is about 50% less than *UU* and *UC* dispersion. Hence, a large coefficient on the border dummy for one of the countries is not evidence of a border effect. Third, even our conservative estimate based on taking the minimum (across countries) border coefficients may overstate the border effect and, hence, even the smallest of the border estimates could be an upper bound.

Now we impose a border between *U* and *C* by increasing *B* from 0 to 0.1. This allows us to look at the impact of a border on cross-border and within-country price dispersion. We report the resulting price dispersions in columns (8)-(10) and the corresponding border effect estimates in columns (11)-(13). Finally, in column (14) we report $\Delta\bar{\sigma}_{UC}$,which is the change in the $\bar{\sigma}_{UC}$ between the no-border case and small-border case. In the ER approach, $\Delta\bar{\sigma}_{UC}$ is the ideal measure of the border effect. However, $\Delta\bar{\sigma}_{UC}$ is not uniformly positive and, in fact, is negative for moderate τ_{UU}. Hence, knowing $\Delta\bar{\sigma}_{UC}$ is not enough to infer the size of the border effect. Furthermore, estimates of the border effect based on the no-border benchmark constructed from within country variation is much larger than $\Delta\bar{\sigma}_{UC}$, although our approach of taking minimum value again yields the smallest bias. Note that the estimate of the border effect based on the border coefficient is consistent across approaches only if there is no country heterogeneity, i.e., $\tau_{UU} = \tau_{CC}$. We conclude that within-country variation serves as a poor benchmark for computing the border effect and one has to know τ_{UU}, τ_{CC} and other fundamental parameters as well as the model of price determination to uncover the size of the border.

To emphasize the last point, we present an example when a costly border diverts international trade toward intranational trade and within-country dispersion can endogenously decrease or increase with the introduction of a border. We continue working with our model but now restrict the number of production locations to be less than the number of consumption locations, which is equal to the number of cities. Such a situation can arise if there are increasing returns to scale in production, e.g., due to fixed costs of having an active production unit. Suppose that the central planner minimizes total cost which includes the variable and fixed cost of production and transportation cost. Without loss of generality, assume that only n production units operate at optimum. Hence, price in location i is determined as

$$p_{it} = \min\{x_{i_1 t} + \tau_{i_1}, ..., x_{i_n, t} + \tau_{i_n}\}, \tag{13}$$

where $i_1,\ldots,i_n$ index production locations. Also suppose that the number of cities in Canada is smaller than the number of cities in the US.

Since $n < N$, cities form clusters around production locations. If cities always buy from the same production center, the time series dispersion of the price differential is zero. Because marginal cost fluctuates over time, some cities switch suppliers and this induces price dispersion. Clearly, cities that buy from different suppliers have time-varying price differentials and, thus, the time-series standard deviation of the price differential is greater than zero. In brief, within cluster price dispersion is zero, but there is positive dispersion between clusters. One interpretation of this model is that it generates regional markets around production centers – correcting for distance does not fully correct for this regionalization as some cities that are relatively close to each other may buy from a different production center. By creating regional markets, this model allows us to look at the impact of asymmetry in size. Big countries will have more regional markets and hence more dispersion in prices, while small countries will have fewer markets and therefore less dispersion in prices.

We simulate the model with 15 Canadian and 21 US cities, so that N=36. In the benchmark specification we set $\tau_{CC} = \tau_{UC} = \tau_{UU} = 0$, $\theta = 0.4, \rho = 0, D = 0$ and assume that the trade cost is equal to 0.1 per unit of distance. Since location of production units depends on the geography of the cities, we assume that cities are located as in Figure 3. We present the results in Table 6.

Because Canada (small country) has a smaller number of production centers than the US (large country) and hence there is a greater overlap in the nomenclature of products consumed across locations in Canada than in the US, price dispersion in Canada is smaller than price dispersion in the US. Furthermore, because average distances between cities within a country are smaller than the average distances to cities in the other country, cities within a country are more likely to buy from a production unit within its country and, hence, international price dispersion is greater than intra-national price dispersion. Again, this model will generate a significant border coefficient which measures differences in intra- and intercountry price dispersion even when there is no friction in crossing the border.

Now, because the number of production locations is fixed, drawing a border or raising the cost of trade within the US forces Canada to have fewer clusters (production units) and, therefore, to have smaller dispersion of price differentials. In fact, if a country has only one production unit, the (time-series) dispersion of price differentials is zero. Basically, the border induces more within-country coordination and less cross-country coordination of prices and international flows of goods are diverted to within-country flows because some Canadian (US) cities are now forced to switch from US (Canadian) suppliers to Canadian (US) suppliers. This endogenous response of price dispersion is similar to the endogenous response of trade flows in Anderson and van Wincoop (2003) to changes in border frictions. Again, $\Delta\bar{\sigma}_{UC}$ varies with τ_{UU} and border effects based on the benchmark of within-country price dispersion are biased upwards. Even the conservative estimate overstates the size of the border effect.

The main messages of this simulation, however, are that within-country and cross-border price dispersion is determined simultaneously and that ignoring equilibrium effects can result in an enormous overstatement of the border effect. Put differently, the introduction of a small border friction can induce large changes in within- and cross-border price dispersion. Static models of price determination will interpret these differences as large border effects.

In summary, without a structural model that specifies market structure, transportation and production costs it is impossible to identify the border effect. To reiterate, our (counter) examples lead to several important conclusions. First, country heterogeneity can play a crucial role in measuring the border effect. Second, the standard methodology based on comparison of between and within country price dispersion can yield significantly upwardly biased estimates of the border effect. Third, finding a significant border coefficient from some country's perspective does not entail a genuine border friction. Fourth, international price dispersion (relative to either the unobserved no-border case or within country variation) does not indicate the size of the border effect without information about other fundamental parameters of price determination. Finally, although a structural model is required to have a clean interpretation of the border estimate, the simulations show how sensitive conclusions about border effects can be even to minor changes in model specification. In light of these results, natural experiments such as Morshed (2007) appear to provide cleaner identification and more robust estimates of the border effect.[14]

6 Concluding Remarks

The border effects estimated from price data are often implausibly large. Moreover, significant border effects are found where they should not be found. We take these facts as indication that the commonly applied methodology makes assumptions that do not hold in the data. We show that when there is country heterogeneity in prices, there is no clear benchmark from which to gauge the effect of the border. In effect, the standard regression used to estimate border effects is under-identified; it is impossible to separate the border from the effect of trading with a country with a different distribution of prices. We return to the data studied by Engel and Rogers and show that in the U.S.-Canada sample, adjusting for cross-country heterogeneity produces two border coefficients: the first one implies border distance equivalent of 47 km, from the US perspective, and the second one implies border distance equivalent of over 108 million km, from the Canadian perspective. However, *neither* coefficient is a true estimate of the border frictions as both coefficients reflect the combined effect of the border *and* the effect of cross-county heterogeneity on trade prices. When within-country heterogeneity in prices is significant, as it is in the case of the United States and Canada, the border coefficient that emerges from tests comparing within-country prices to cross-border prices tells us little about actual border effects. Our results strongly suggest that reduced-form coefficients in border-effect regressions should not be generally interpreted as measures of border frictions irrespective of what variable is used in the analysis. Simple comparison of within and cross country statistics cannot credibly identify the impact of the border.

[14] Time series variation of price differentials could be another source for identifying the border effect. However, time series information is not likely to provide model-free estimates of the border frictions. For example, the sources and/or characteristics of shocks to prices as well as the mechanisms that bring prices back to equilibrium can vary across countries and hence country heterogeneity in another guise will confound border effect estimates. Thus, even with time series information, one has to impose model-based restrictions to identify the border frictions.

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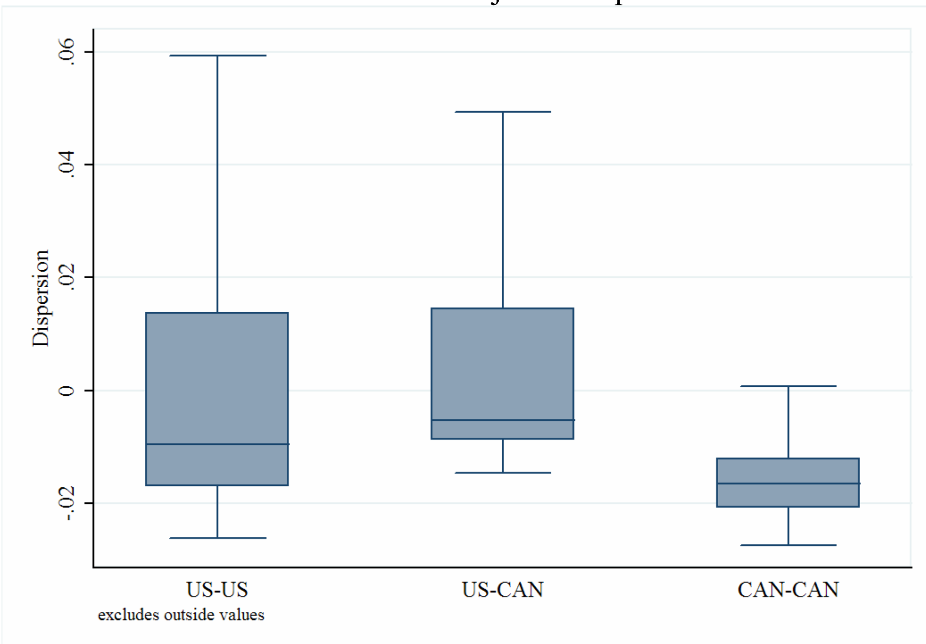
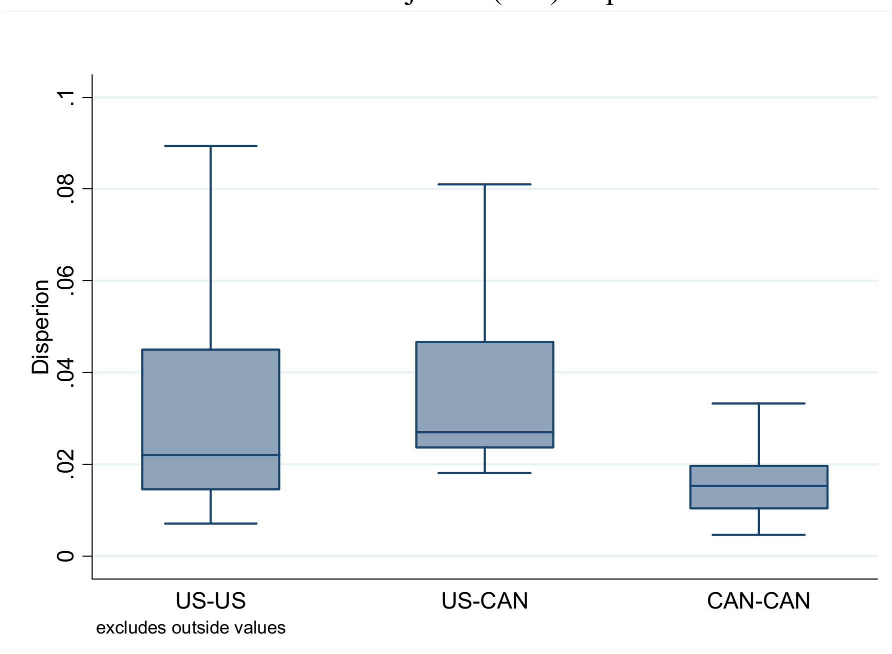
Figure 1. Distribution of dispersion of changes in real exchange rate.
Panel A: distance-adjusted dispersion



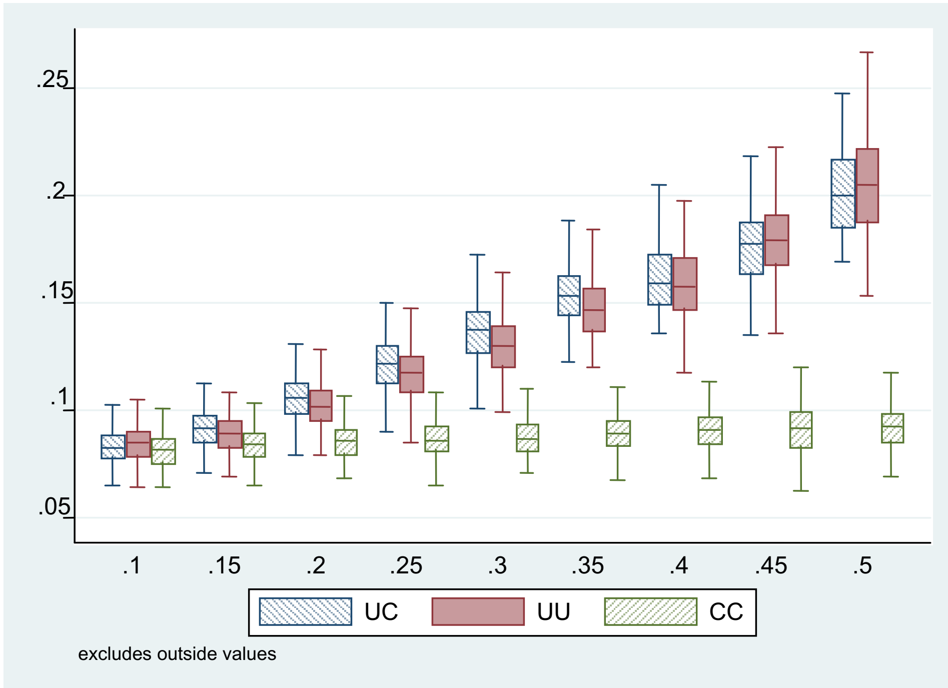
Panel B: non adjusted (raw) dispersion



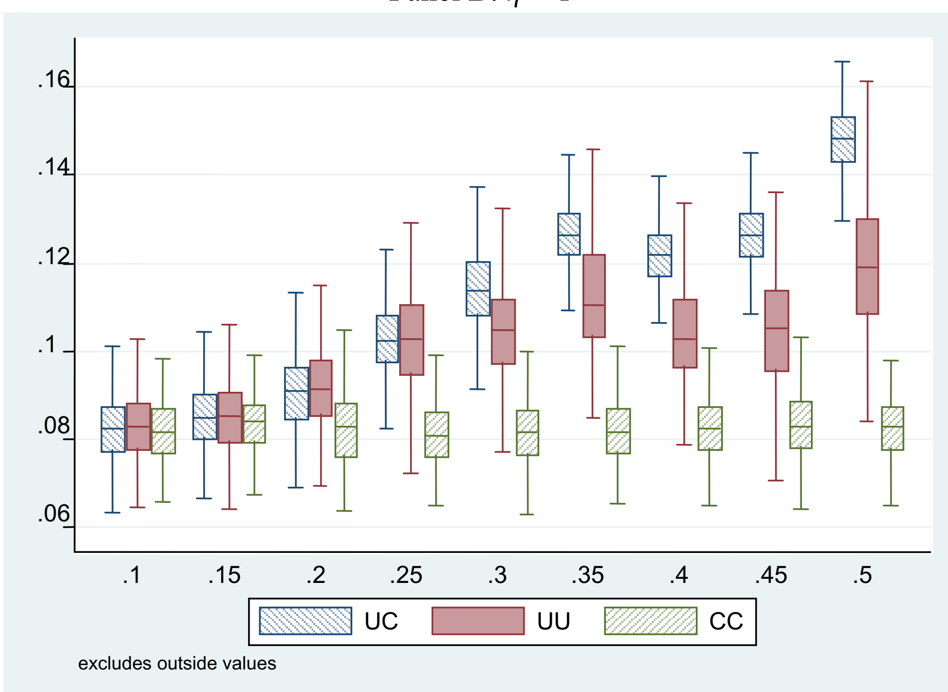
Note: Distance-adjusted dispersion is computed as the residual in the regression of dispersion on log distance between cities. Because the distance-adjusted dispersion is measured as a residual, it may take negative values. Whiskers extend from the box to the upper and lower adjacent values and are capped with an adjacent line. Adjacent values are calculated utilizing the interquartile range (IQR). The upper adjacent value is the largest data value that is less than or equal to the third quartile plus 1.5 X IQR and the lower adjacent value is the smallest data value that is greater than or equal to the first quartile minus 1.5 times IQR. Values exceeding the upper and lower adjacent values (outside values) are not presented.

Figure 2. Simulated dispersion of real exchange rate by types of city pairs.

Panel A: $\eta = 0$

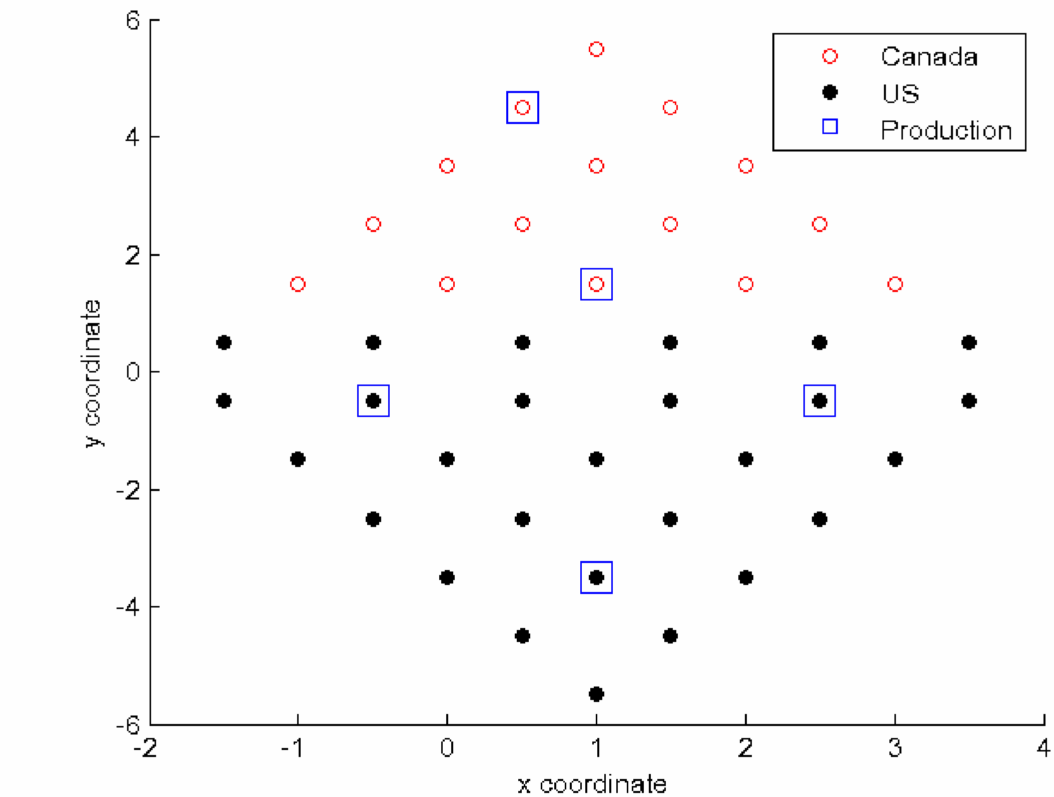


Panel B: $\eta = 1$



Note: Standard deviation is on vertical axis. Cost of trade τ_{UU} is on horizontal axis. See text for further details.

Figure 3. Geography of city locations.



Note: this figure presents assumed geographic locations for the simulation. Optimal locations for production units are derived for the benchmark case $\tau_{CC} = \tau_{UC} = \tau_{UU} = 0$. The cost of transposition is 0.1 per unit of distance. Optimal production locations minimize production and transportation costs before introducing the border.

Table 1. Descriptive statistics by types of city pairs: US-Canada

	US-US pairs (1)	US-Canada pairs (2)	Canada-Canada pairs (3)	All pairs (4)
Standard deviation of changes in real exchange rate, $\sigma(\Delta_2 q_{ijt}^k)$	0.031	0.037	0.016	0.032
Standard deviation of changes in real exchange rate adjusted for distance	0.000	0.004	-0.015	
Standard deviation of innovations in real exchange rate	0.028	0.031	0.014	0.028
Standard deviation of changes in *relative* exchange rate, $\sigma(\Delta_2 RER_{ijt}^k)$	0.032	0.032	0.016	0.029
Log(*distance*)	7.17	7.53	7.21	7.35
N	1,140	1,764	504	3,408

Note: the table reports mean values of the presented variables. Average is taken over city pairs. $\sigma(x)$ is time series standard deviation of variable x. In the measure of price dispersion adjusted for distance, the price dispersion for US-Canada and Canada-Canada pairs is relative to the price dispersion for US-US pairs, where the latter is normalized to zero. Innovations in the real exchange rate are taken to be the residuals from the regression of real exchange rate on its lags. See text for further details and for the definitions of the variables..

Table 2. Estimates of the border effect: US-Canada.

	Dependent variable is standard deviation of changes in real exchange rate, $\sigma(\Delta_2 q_{ijt}^k)$	Dependent variable is standard deviation of changes in relative exchange rate, $\sigma(\Delta_2 RER_{ijt}^k)$
	(1)	(2)
Log(*distance*)	1.076***	1.044***
	(0.269)	(0.266)
Border	12.026***	7.036***
	(0.363)	(0.357)
R^2	0.78	0.78
Border effect (km)	**71,438**	**845**

Note: The estimated specification is equation (1). Coefficients are multiplied by 1000. *CC* is the dummy variable for Canada-Canada city pairs, *Border* is the dummy variable for US-Canada pairs (*UC* dummy). Omitted category is US-US city pairs. Robust standard errors are in parentheses. *, **, and *** are significant at 10%, 5%, and 1% respectively. City and good category dummies are included but not reported.

Table 3. Alternative estimates of the border effect: US-Canada.

	Dependent variable is standard deviation of changes in real exchange rate, $\sigma(\Delta_2 q_{ijt}^k)$	Dependent variable is standard deviation of changes in relative exchange rate, $\sigma(\Delta_2 RER_{ijt}^k)$
	(1)	(2)
Log(*distance*)	1.076***	1.044***
	(0.269)	(0.266)
Border	4.148***	-1.161***
	(0.393)	(0.388)
CC	-15.756***	-16.393***
	(0.529)	(0.522)
R^2	0.78	0.78
Border from the US perspective	**47**	**-3**
Border from the Canadian perspective	**108.2 mln**	**2.3 mln**

Note: The estimated specification is equation (11). Coefficients are multiplied by 1000. *CC* is the dummy variable for Canada-Canada city pairs, *Border* is the dummy variable for US-Canada pairs (*UC* dummy). Omitted category is US-US city pairs. Robust standard errors are in parentheses. *, **, and *** are significant at 10%, 5%, and 1% respectively. City and good category dummies are included but not reported.

Table 4. Border effect and country/city dummies: US-Canada.

	ER specification Equation (1)	Alternative specification Equation (11)
	(1)	(2)
Log(*distance*)	1.076***	1.076***
	(0.269)	(0.269)
Border	12.026***	4.148***
	(0.363)	(0.393)
CC		-15.756***
		(0.529)
Baltimore	5.329***	2.246***
	(0.606)	(0.596)
Boston	5.264***	2.181***
	(0.598)	(0.587)
Calgary	-4.135***	0.660
	(0.563)	(0.539)
Chicago	3.975***	0.892
	(0.559)	(0.554)
Dallas	4.620***	1.538**
	(0.629)	(0.615)
Detroit	1.309**	-1.773***
	(0.564)	(0.558)
Edmonton	-4.854***	-0.059
	(0.564)	(0.540)
Houston	4.329***	1.246**
	(0.631)	(0.618)
Los Angeles	-0.637	-3.719***
	(0.581)	(0.576)
Montreal	-5.259***	-0.463
	(0.564)	(0.540)
Miami	3.838***	0.755
	(0.608)	(0.598)
New York	-1.468***	-4.550***
	(0.564)	(0.559)
Ottawa	-5.428***	-0.633
	(0.566)	(0.543)
Philadelphia	3.556***	0.473
	(0.567)	(0.562)
Pittsburgh	2.135***	-0.948
	(0.632)	(0.618)
Quebec	-5.297***	-0.501
	(0.559)	(0.535)
Regina	-2.596***	2.199***
	(0.559)	(0.535)
San Francisco	1.870***	-1.212**
	(0.584)	(0.579)
St. Louis	5.767***	2.684***
	(0.601)	(0.591)
Toronto	-5.170***	-0.374
	(0.569)	(0.546)
Vancouver	-5.528***	-0.733
	(0.572)	(0.549)
Winnipeg	-4.891***	-0.096
	(0.558)	(0.534)
Washington, D.C.	3.271***	0.188
	(0.604)	(0.594)
Constant	18.337***	24.502***
	(1.954)	(1.959)

Note: Dependent variable is the standard deviation of changes in the real exchange rate, $\sigma(\Delta_2 q_{ijt}^k)$. *CC* is the dummy variable for Canada-Canada city pairs, *Border* is the dummy variable for US-Canada pairs (*UC* dummy). Omitted category is US-US city pairs. Coefficients are multiplied by 1000. Robust standard errors are in parentheses. *, **, and *** are significant at 10%, 5%, and 1% respectively. Columns 1 and 2 correspond to column 1 in Table 2 and column 1 in Table 3 respectively. See the note for Table 2 and the text for further details.

Table 5. Simulated dispersion of real exchange rate by types of city pairs.

	Border B = 0						*Border B* = 0.1						True border
				Implied estimate of the border effect						Implied estimate of the border effect			
τ_{UU}	$\bar{\sigma}_{CC}$	$\bar{\sigma}_{UU}$	$\bar{\sigma}_{UC}$	ER	GT US	GT Canada	$\bar{\sigma}_{CC}$	$\bar{\sigma}_{UU}$	$\bar{\sigma}_{UC}$	ER	GT-US	GT-Canada	$\Delta\bar{\sigma}_{UC}$ (10)-(4)
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)
							Panel A: η = 0						
0.1	0.082	0.088	0.088	0.003	0.000	0.006	0.089	0.087	0.119	0.031	0.032	0.030	0.031
0.2	0.092	0.106	0.130	0.031	0.024	0.038	0.094	0.111	0.160	0.057	0.049	0.066	0.030
0.3	0.098	0.147	0.184	0.062	0.037	0.086	0.092	0.149	0.188	0.067	0.039	0.095	0.004
0.4	0.100	0.188	0.229	0.085	0.041	0.130	0.100	0.181	0.214	0.074	0.034	0.114	-0.015
0.5	0.100	0.239	0.268	0.099	0.029	0.169	0.098	0.234	0.245	0.079	0.010	0.147	-0.024
							Panel B: η = 1						
0.1	0.083	0.087	0.085	0.000	-0.002	0.002	0.090	0.087	0.119	0.031	0.032	0.029	0.034
0.2	0.086	0.097	0.094	0.002	-0.003	0.008	0.089	0.106	0.110	0.012	0.004	0.021	0.016
0.3	0.084	0.118	0.118	0.017	0.000	0.033	0.088	0.128	0.110	0.002	-0.018	0.022	-0.007
0.4	0.087	0.120	0.139	0.035	0.019	0.051	0.094	0.139	0.120	0.004	-0.019	0.026	-0.019
0.5	0.086	0.139	0.159	0.047	0.021	0.073	0.089	0.163	0.141	0.015	-0.022	0.052	-0.018

Note: $\bar{\sigma}_{UU}, \bar{\sigma}_{CC}, \bar{\sigma}_{UC}$ is the mean dispersion of real exchange rate for *UU*, *CC*, and *UC* city pairs. Columns (5)-(6) and (11)-(14) report the implied border effect according to Engel and Rogers (1996) and our (GT; from each country's perspective) methods to measure the border effect. The last column (14) is equal to column (10) minus column (4) and it measures the increase in the international price dispersion with an increase of *B* from 0 to 0.1. Prices are determined according to (12).

Table 6. Simulated dispersion of real exchange rate by types of city pairs: production clusters

	Border B = 0							*Border B* = 0.1							
					Implied estimate of the border effect							Implied estimate of the border effect			True border
τ_{UU}	$\bar{\sigma}_{CC}$	$\bar{\sigma}_{UU}$	$\bar{\sigma}_{UC}$	Production locations in Canada	ER	GT US	GT Canada	$\bar{\sigma}_{CC}$	$\bar{\sigma}_{UU}$	$\bar{\sigma}_{UC}$	Production locations in Canada	ER	GT-US	GT-Canada	$\Delta\bar{\sigma}_{UC}$ (11)-(4)
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
Panel A: η = 0															
0	0.082	0.096	0.125	2	0.036	0.029	0.044	0.077	0.102	0.138	2	0.049	0.036	0.061	0.013
0.1	0.077	0.102	0.133	2	0.043	0.031	0.056	0.074	0.103	0.141	2	0.053	0.039	0.068	0.009
0.2	0.074	0.102	0.136	2	0.048	0.034	0.062	0.073	0.103	0.142	2	0.055	0.040	0.069	0.006
0.3	0.073	0.102	0.137	2	0.049	0.035	0.064	0.000	0.103	0.148	1	0.096	0.044	0.147	0.011
0.4	0.000	0.103	0.144	1	0.093	0.041	0.144	0.000	0.103	0.148	1	0.096	0.044	0.148	0.003
0.5	0.000	0.103	0.144	1	0.093	0.042	0.144	0.000	0.103	0.148	1	0.096	0.044	0.148	0.003
Panel B: η = 1															
0	0.082	0.096	0.125	2	0.036	0.029	0.044	0.077	0.102	0.138	2	0.049	0.036	0.061	0.013
0.1	0.087	0.098	0.119	3	0.026	0.021	0.032	0.077	0.102	0.133	2	0.043	0.031	0.056	0.014
0.2	0.093	0.093	0.106	3	0.013	0.013	0.013	0.087	0.099	0.120	3	0.027	0.021	0.033	0.014
0.3	0.093	0.086	0.100	3	0.011	0.015	0.007	0.093	0.093	0.106	3	0.013	0.013	0.013	0.006
0.4	0.093	0.073	0.095	3	0.012	0.022	0.002	0.093	0.086	0.100	3	0.011	0.015	0.007	0.005
0.5	0.093	0.068	0.094	3	0.013	0.026	0.001	0.093	0.073	0.095	3	0.012	0.022	0.002	0.001

Note: $\bar{\sigma}_{UU}, \bar{\sigma}_{CC}, \bar{\sigma}_{UC}$ is the mean dispersion of real exchange rate for *UU*, *CC*, and *UC* city pairs. Columns (5)-(6) and (9)-(11) report the implied border effect according to Engel and Rogers (1996) and our (GT; from each country's perspective) methods to measure the border effect. The last column (16) is equal to column (11) minus column (4) and it measures the increase in the international price dispersion with an increase of *B* from 0 to 0.1. The number of production clusters is 5. The number of city/consumption locations is 21 in US and 15 in Canada. City locations are presented in Figure 3. Production locations minimize production and transportation costs. See text for further details. Prices are determined according to (13).