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Explaining the border effect: the role of exchange rate variability, shipping costs, and geography

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Abstract

This paper exploits a three-dimensional panel data set of prices on 27 traded goods, over 88 quarters, across 96 cities in the US and Japan. We show that a simple average of good-level real exchange rates tracks the nominal exchange rate well, suggesting strong evidence of sticky prices. Focusing on dispersion in prices between city-pairs, we find that crossing the US–Japan ‘Border’ is equivalent to adding as much as 43 000 trillion miles to the cross-country volatility of relative prices. We turn next to economic explanations for this so-called border effect and to its dynamics. Distance, unit-shipping costs, and exchange rate variability, collectively explain a substantial portion of the observed international market segmentation. Relative wage variability, on the other hand, has little independent impact on segmentation. © 2001 Elsevier Science B.V. All rights reserved.

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1. Introduction

International markets have been more segmented than intra-national markets for at least as long as there have been political borders. Despite technological

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advances and negotiated reductions in barriers, market segmentation continues to exist. This is not news. The extent of present day segmentation when quantitatively documented however, is striking. In a seminal paper that looks at price volatility, Engel and Rogers (1996) showed that the dispersion of prices of similar goods increases with the distance between city pairs, a pattern that holds even within a country. However, when the price comparisons cross national political boundaries (the US and Canada in their example), the dispersion of prices goes far beyond distance (and hence transportation costs): crossing the US–Canada border is equivalent to crossing a distance of 75 000 miles. This is surprising given that formal trade barriers between these two countries are low — and declining over time — and physical barriers to trade between the northern US states and the southern Canadian provinces are presumably less important than those existing among east and west coast US cities. Moreover, differences in culture and legal systems between these two countries also appear small.

Whatever the reason for the sizable border effect, its existence is at least consistent with the literature on the speed of convergence to the law of one price (LOP) or purchasing power parity (PPP). For example, studies of convergence of real exchange rates using cross-country evidence (e.g., Frankel and Rose, 1996, among many others) have settled down on a near-consensus of 3–5 years for the half-life of PPP deviations. In contrast, Parsley and Wei (1996) estimated that the half-life of deviations from the LOP is only about 1 year for purely intra-US prices. Moreover, there is an analogue in studies using international trade quantity data. McCallum (1995), Wei (1996), and Helliwell (1998) each find substantial home bias in trade.

Crucini et al. (2000) provide an interesting recent twist based on a large cross-section of goods prices in European capital cities in 1985. They find that while CPI- or WPI-based log real exchange rates may be far away from zero, the simple average of good-level log real exchange rates was actually fairly close to zero. In other words, the equally weighted average of goods prices in local currencies between two European cities, say, Paris and Bonn, is a good predictor of the nominal exchange rate in that year. This suggests that markets (in Europe at least) may, in fact, be more integrated, and borders may matter less than studies examining the variability of price differences would suggest. Of course, exchange rates were managed among the European countries in their sample, and both the physical distance between countries and policy-induced trade barriers were low.¹ The border effect could be more significant between country pairs that do not have such favorable conditions.

In this paper, we exploit a three-dimensional panel data set of prices for 27

¹Despite formal exchange rate agreements, intra-European currency movements were surprisingly large over longer periods. For example, the 1980–84 percentage changes (versus the Deutschmark) were: Italy (29.5), France (32.2), Spain (36.2), Belgium (24.5), and Portugal (98.6), compared to a yen/dollar change of only 23.7% for that period. Source: IFS 1999 Yearbook.

commodity-level goods (e.g., one box of facial tissue, 175 count), in 88 quarters (1976:1–1997:4), in 96 cities in Japan and the United States. Each of the 27 goods is selected so that we can match the definition of the good reasonably well between the two countries.²

We have several objectives. First, we examine the behavior of the average good-level real exchange rate for the US and Japan — the counterpart to the measure examined in the Crucini et al. (2000) paper. Our data set allows us to ask two questions that the earlier paper cannot address. Does the average exchange rate between countries stray farther away from zero than that between cities within a country? And second, is there any tendency for the average exchange rate to move closer towards zero over time?

Second, we examine the infamous border effect, which is related to the dispersion of the real exchange rate. The border effect is defined as the extra dispersion in prices between cities in different countries beyond what can be explained by physical distance — the counterpart to the measure studied by Engel and Rogers (1996).³ Our innovation is on understanding its dynamics. We ask two related questions. First, is there any evidence that the Japan–US ‘Border’ narrows over time? And second, is there evidence linking the evolution of the border effect with plausible economic candidates (e.g., the unit cost of international transportation)?

In contrast to Crucini et al., we present evidence that the mean of good-level international log real exchange rates is substantially more volatile, and farther away from zero on average, than the comparable mean of intra-national log real exchange rates. We also show that the simple average of good-level real exchange rates tracks the nominal exchange rate closely. This seems to be very strong evidence of sticky prices in local currencies. We turn next to economic explanations for this so-called ‘Border’ effect. Focusing on variability in good-level real exchange rates, we confirm previous findings that international borders matter a great deal. However, there is evidence that the border effect between Japan and the US declines over time in our sample. Furthermore, distance, shipping costs, and exchange rate variability collectively explain a substantial portion of the border effect.

2. Data

The source for the Japanese data is the *Annual Report on the Retail Price Survey*, published by the Statistics Bureau of the Management and Coordination

²A subset of the US data has been examined in Parsley and Wei (1996), and O’Connell and Wei (2000).

³Of course the US and Japan are not actually contiguous. We nonetheless continue to refer to the effect of international market segmentation on price dispersion as the ‘Border’ effect.

Agency of the Government of Japan. This print publication contains the prices of a large number of goods and services (~700) for a sample of Japanese cities (~70) on a monthly basis for the year. For this study we selected the first month of each quarter to obtain a time match with our US data set; to assure geographic coverage and comparability with the US sample, we also selected 48 Japanese cities. There is still a slight time mismatch, however. The US data are generally sampled 7–10 days prior to the Japanese data. For every quarter in our sample (1976.1–1997.4), all 48 Japanese cities were part of the sample.

The source for the US data is the *Cost of Living Index* published by the American Chamber of Commerce Researchers Association. This data set is described in more detail in Parsley and Wei (1996). Briefly, for this study we selected 48 US cities and the 27 traded goods most closely resembling those available in the Japanese *Annual Report*. Each quarterly issue of *Cost of Living Index* reports prices from a cross-section of US cities (currently exceeding 300). We selected US cities that appeared in roughly 90% of the quarterly surveys. The working paper version of this paper contains a more complete description of the cities and goods we study (Parsley and Wei, 2000). Prior to conducting our analysis we scaled the prices to further insure the units for each good were comparable between the two countries.

3. Statistical results

3.1. The mean of good-level real exchange rates

Crucini et al. (2000) note that even though value-weighted average deviations from LOP over goods can be large, for the sample of European cities (in 1985) the equally weighted average was remarkably close to zero for that year. We will see if this result is something specific to their sample of countries, which were under a fixed exchange rate arrangement, or to the particular year for which they have data.

In this paper, we focus only on those goods most clearly in the traded goods category, in part, to abstract from the Balassa–Samuelson effect. Of course, the retail price of any good could have tradable and non-tradable components. We will come back to this issue later. We attempt to limit variations in individual goods themselves through our matching process.

We choose one benchmark city from Japan (Tokushima) and one from the US (Louisville). These are ‘centrally located’ cities in their respective countries. This produces a sample of 189 city pairs in total.⁴ We repeated all of the analysis of this

⁴To arrive at 189 city-pairs, note that there are 47 intra-Japan city-pairs, 47 intra-US city-pairs, and 95 cross-country city-pairs: (48) Tokushima benchmark — each US city, plus (48) each Japanese city — Louisville benchmark, minus (1), since Tokushima — Louisville would be included twice.

paper using a different set of benchmark cities (Osaka and Houston) and found the results were not sensitive to this choice. Note this procedure still produces (without missing values) roughly 5100 good-level real exchange rates each period, or nearly 450 000 time-series observations. Ultimately, we study the evolution of these distributions of real exchange rates on a year-by-year basis.

Let $P(i,k,t)$ be the US dollar price of good k in city i at time t . For a given city pair (i,j) and a given good k at a time t , we could define a good-level log real exchange rate

$$r(ij,k,t) = \ln P(i,k,t) - \ln P(j,k,t).$$

We find it informative to study and compare the distributions of three types of good-level log exchange rates: (1) $r(ij,k,t)$ for intra-national US city-pairs, (2) $r(ij,k,t)$ for intra-national Japanese city-pairs, and (3) $r(ij,k,t)$ for international city-pairs.

Fig. 1 plots the empirical kernel density estimate of the log average real exchange rate for each of these three comparisons (within Japan, within the US,

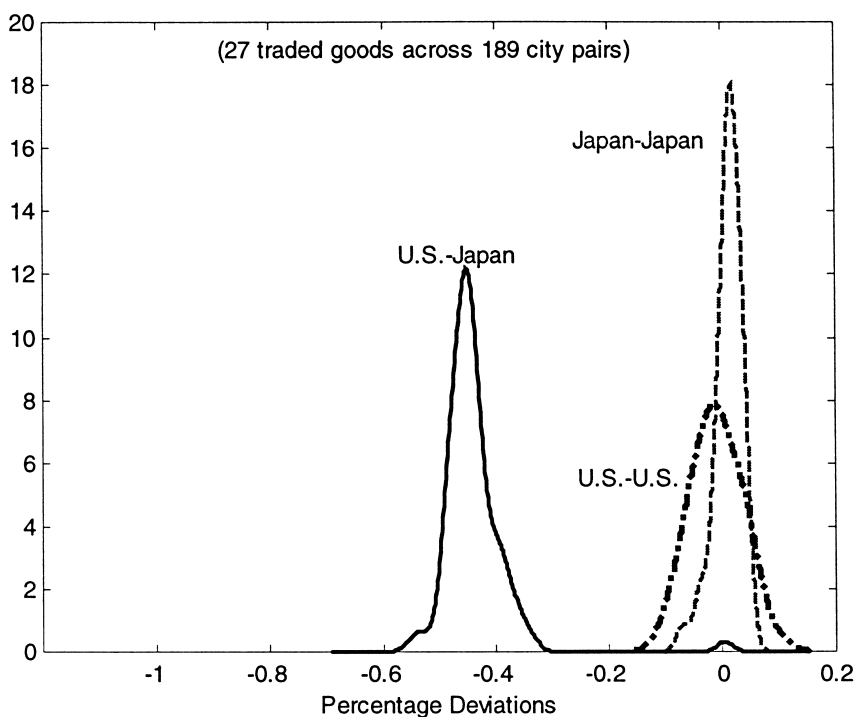


Fig. 1. Empirical density functions of good-level log real exchange rate in 1985.

and between the US and Japan) for 1985, the same year as used by Crucini et al. (2000). Several features of the figure stand out. First, the within country densities are more closely centered on zero (a function of the benchmark city). Note that Japanese prices are less dispersed than those in the US. This is possibly due to the relative sizes of the two countries; the greater average distance between cities in the US may allow prices to vary more. Judging by this figure, deviations from the LOP within a country do not appear extraordinary. And second, the US–Japan density function is centered to the left of zero. This means that in 1985 most Japanese prices were higher than US prices. It also suggests the Crucini et al. (2000) finding may be specific to Europe.⁵

In Fig. 2, we repeat the exercise for 1990. The comparison with Fig. 1 is striking. The between-country distribution has diverged from the two within-country distributions. Japanese prices expressed in US dollars have risen even more relative to US prices. The violation of the law of one price became even more severe.

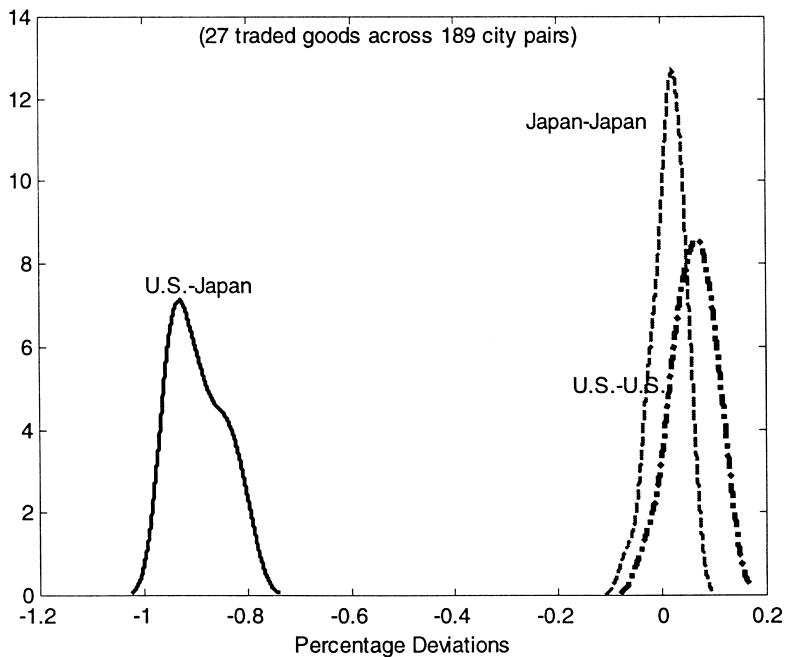


Fig. 2. Empirical density functions of good-level log real exchange rate in 1990.

⁵Other potential reasons for the difference between our results and those in Crucini et al. (2000) include: greater differences in goods internationally, than intra-nationally, in our sample; measurement error introduced by our rescaling procedure (e.g., the price of a 7-oz bottle of shampoo is probably not seven-tenths the price of a 10-oz bottle); and the fact that their cross-section is much more extensive.

This suggests that there may not be a trend decline in the average violation of the law of one price for traded goods. Of course, we naturally should be cautious in making a time series inference based on observations at two points in time. So we now turn to some time series evidence. Let us define the average within-US log real exchange rate at time t , $\bar{r}(us,t)$, as the average of $r(ij,k,t)$ over all goods and all city-pairs within the US. We can define $\bar{r}(japan,t)$ and $\bar{r}(us-japan,t)$ in an analogous way.

Fig. 3 plots the three average log real exchange rates over time (1976–1997), respectively. It is clear that the intra-national average log real exchange rates (or percentage deviation of prices of the same good between two cities), i.e., within both the US and Japan, are fairly close to zero. In fact they vary within plus/minus 5–7% in each of the 22 years in our sample. In comparison, the average percentage deviation between US and Japan makes much larger gyrations, from a minimum of 40% in 1982 to a maximum of 130% in 1995.

We cannot fail to notice that the time series path of the average log real exchange rate between the US and Japan resembles the log of the *nominal* yen/dollar exchange rate. We formally tested this hypothesis by regressing the first difference in the log average real exchange rate on a constant and the first

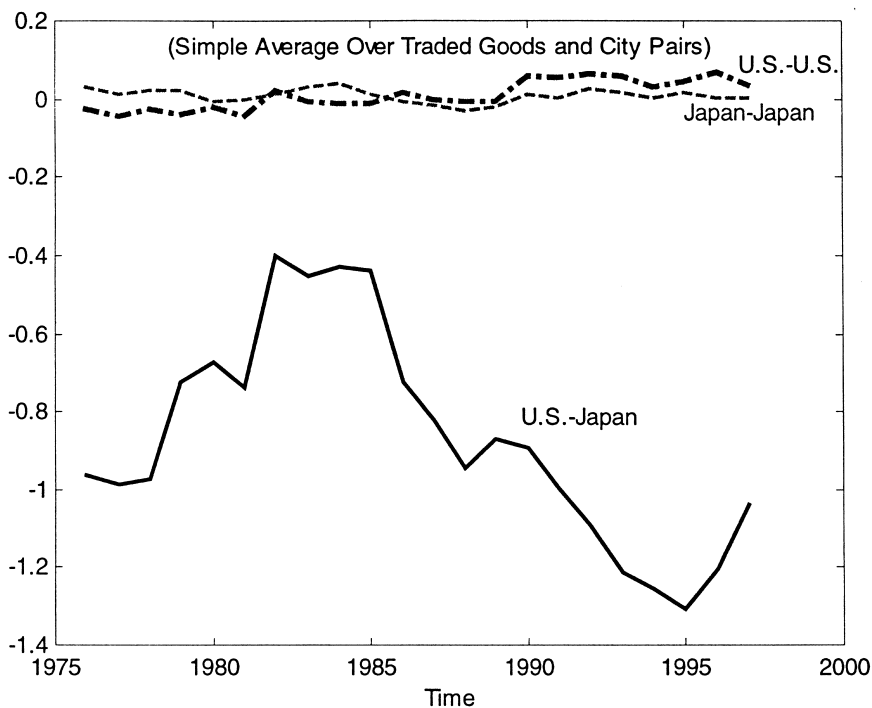


Fig. 3. Average of good-level log real exchange rates.

difference in the log nominal exchange rate.⁶ In accord with our expectations, the nominal exchange rate explains much of the variation — the adjusted R^2 of the equation is 0.49, and the coefficient on the nominal exchange rate is estimated at 0.62 with a standard error of 0.14. This seems to us very strong evidence that sticky prices in local currencies (as opposed to relative price of non-tradables), is a big part of CPI-based real exchange rate movements. This, from a different angle, confirms the finding in Rogers and Jenkins (1995).

Finally, we note that deviations from the law of one price in 1985 (or any year during 1980–86) were smaller than either earlier or later years. Hence the Crucini et al. (2000) finding may also be attributable to the particular year they study.⁷

3.2. Dispersion in intra-national versus international price differences

We would like to know whether international market integration has increased over time (or equivalently, whether the border effect has diminished). Clearly, the evidence in the previous sub-sections is that the average violation of the law of one price does not have a downward trend. However the range in which the violations take place, or the zone of no-arbitrage, could nonetheless narrow over time. In this section, we turn to an explicit investigation of the border effect.

The logic of no-arbitrage imposes two inequality constraints on the prices of an identical good, k , in two different locations, i and j . Let $C(ij,t)$ be the cost of engaging in arbitrage activity for transporting and selling one unit of good k from location i to j (or the reverse). Then, the price in one location plus the cost of arbitrage has to be at least as great as the price of the same good in another location.

$$\ln P(i,k,t) + \ln C(ij,t) \geq \ln P(j,k,t)$$

and

$$\ln P(j,k,t) + \ln C(ij,t) \geq \ln P(i,k,t)$$

Collectively, they imply that

$$-\ln C(ij,t) \leq \ln P(i,k,t) - \ln P(j,k,t) \leq \ln C(ij,t)$$

As long as a given price differential between the two locations satisfies these inequalities, it will not trigger arbitrage. To put it differently, within the zone of

⁶The rate was taken from the International Financial Statistics March 1999 CD (line ae).

⁷We also examined the absolute deviations from the LOP. Presentation of these results here is omitted to conserve space. Briefly, for our data we see the same pattern as in Fig. 3. Within each country, the mean absolute deviations are between 10 and 15% (again, somewhat larger in the US than in Japan). However, the cross-country mean absolute deviations are several times as large, between 75 and 140%.

no-arbitrage, the price differential can potentially take on an infinite number of possible values. The no-arbitrage story can be made more formal (see, e.g., O’Connell and Wei, 2000). The exact details need not concern us here. What is important for this paper is that both the simple no-arbitrage story above, and its formalization developed in O’Connell and Wei (2000) suggest that a given cost of arbitrage defines only the range in which price differences can occur, but not necessarily any particular realization of the difference. Therefore, in our empirical specification we use as our dependent variable, some measure of the possible range of price differences for a given city-pair. For robustness, we consider two measures of this range: (1) the standard deviation over many realizations of the log price difference, and (2) the inter-quartile range between the 75th and 25th quartiles in the empirical distribution of all price differences between a given city-pair.

3.3. The galactic border between Japan and the US

Let the change in the real exchange rate for good k in city i , relative to benchmark city j , be

$$Q(ij,k,t) \equiv \Delta \ln P(i,k,t) - \Delta \ln P(j,k,t)$$

Note $P(i,k,t)$ and $P(j,k,t)$ are expressed in a common currency.

The costs of arbitrage can have many components. For example, Samuelson’s (1954) ‘iceberg’ model introduces geography in a straightforward fashion. According to this model transportation costs should depend positively on the distance between locations, so that the variation of relative prices also increases with the distance. Secondly, sticky goods prices imply that nominal exchange rate variability would translate into variability of cross-country goods prices. A third important difference between intra-national and international city-pairs is the potential existence of non-traded inputs (e.g., labor) and its effect on relative prices. Engel and Rogers (1996) hypothesize that relative wages are less variable within countries than they are for cross-border city-pairs. Empirically, however, they find that inclusion of relative wage variability has little impact on the border effect.

Our plan is to examine these and other influences on relative price variability over time. As a starting point, however, we begin by reproducing the Engel–Rogers analysis of the border effect, using our US–Japan data set. Specifically, we regress the standard deviation of the change in the real exchange rate, $V(Q(\cdot))$, on the distance between locations and a border dummy

$$V(Q(ij,k)) = \beta_1 \ln(dist_{ij}) + \beta_2 Border_{ij} + \text{a constant, city, and good dummies} + \epsilon_{ij} \quad (1)$$

where $dist_{ij}$ is the greater-circle distance between cities i and j , and $Border_{ij}$ is a

dummy variable that equals 1 if cities i and j are in different countries. The great circle distance is computed by using the latitude and longitude of each city in our sample. The source for the Japanese latitude and longitude data is the United Nations, and the source for the US is the US Naval Observatory.⁸ Note that this regression will have (without missing values) 5103 observations (27 goods \times 189 city-pairs).

The point estimates in column 1 of Table 1 confirm that price dispersion increases with distance and that the border effect is important for explaining cross-country price dispersion (Fig. 4). We report heteroscedasticity-consistent standard errors in parentheses below the estimates. We note that the strength of the distance effect between Japan and the US is somewhat weaker than that for Canada and the US reported by Engel and Rogers.

Engel and Rogers calculate that the US–Canadian border is equivalent to adding as much as 75 000 = $\exp(\beta_2/\beta_1)$ miles does to the cross-country volatility. Performing a similar calculation, our point estimates imply a much larger effect for

Table 1
Explaining the average border effect^a

	Engel–Rogers benchmark regression	Eq. (1)	Eq. (2)	Eq. (3)	Eq. (4)	Eq. (5)
Log distance	0.0022 (0.0017)	0.0049 (0.0018)	0.0049 (0.0018)	0.0049 (0.0018)	0.0049 (0.0018)	0.0049 (0.0017)
Border	0.0649 (0.0055)	0.0717 (0.0057)	0.0601 (0.0059)	0.0154 (0.0066)	0.0291 (0.0064)	0.0219 (0.0066)
Nominal exchange			0.2625	0.2458	−0.4183	0.2773
Rate variability			(0.0434)	(0.0484)	(0.2184)	(0.0472)
Unit shipping costs				0.7509 (0.0571)	0.6476 (0.0676)	0.6747 (0.0559)
Wage variability					0.6461 (0.2120)	
Relative wage inflation variability						1.6713 (0.2245)
City dummies	Yes	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	0.78	0.53	0.53	0.56	0.56	0.56
Number of observations	5065	3820	3820	3820	3820	3820

^a Dependent variable: standard deviation of log price differential; Tokushima–Louisville benchmark cities.

⁸The latitude and longitude information is available at <http://www.un.org/Depts/unsd/demog/392.htm>, for the Japanese data, and, <http://www.touchplate.com/location.html>, for the US data. Formulas for computing greater circle distance are provided in The American Practical Navigator (1977).

the Japan–US case; the number is roughly 6.5 trillion miles, or about 70 000 times the distance from the Earth to the Sun.⁹

In these calculations, however, when one changes the units of distance measurement from (say) miles to kilometers, the interpretation of the distance equivalent changes, from 75 000 miles to 75 000 km. This is because the point estimates of β_1 and β_2 , in a log linear regression such as Eq. (1) are unaffected by a change in measurement. Clearly, this is not reasonable. An alternative way to compute the distance equivalent of the border effect is to ask how much extra distance we need to add to the average distance between the two countries to generate as much price dispersion as we actually observe internationally. Specifically, the distance equivalent of the border effect is the value of Z that solves the following equation:

$$\beta_1 \ln(\overline{\text{distance}} + Z) = \beta_2 + \beta_1 \ln(\overline{\text{distance}}) \quad (2)$$

where $\overline{\text{distance}}$ is the average distance between the US–Japan city-pairs (6627 miles) and β_1 , β_2 are coefficient estimates, say from regressions reported in Table 3.

Notice that this new way of computing the distance equivalent uses the average distance (in whatever units) explicitly in the calculation. Solving this equation for Z yields $Z = \overline{\text{distance}} \times \exp(\beta_2/\beta_1 - 1)$. Substituting the values of β_1 , β_2 and average distance reported by Engel–Rogers, raises their border estimate (between Canada and the US) significantly, from 75 000 miles to 101 million miles. Similarly, using our parameter estimates in Table 1 the ‘Border’ between Japan and the US becomes nearly 43 000 trillion miles! The ‘Border’ is indeed remarkably wide.

Of course, Japan and the US are farther apart than Canada and the US. In fact, the average distance between our international city-pairs is over six times that between the US and Canadian cities studied by Engel and Rogers; however, the greater separation between cities in our sample is only a small part of the story. Other candidate explanations include the fact that the yen/dollar exchange rate has been a lot more volatile than the Canadian dollar/US dollar rate, and the relative wage differential is also likely to be more variable between Japan and the US. We turn to these issues next.

⁹For this calculation to be valid it is necessary that the distance coefficient applies equally to internal and cross-border distances. For our data we cannot reject this null hypothesis at usual significance levels. Note, we do not intend to imply that the cost functions for space travel are similar to those for travel between the US and Japan. We resort to intra-galactic comparisons merely to put these huge distances into some perspective. An alternative (and more terrestrial) way to relate to the 6.5 trillion miles is that it is ~ 130 million times around the globe. Finally, note that the implied distance equivalent is sensitive to small changes in because distance enters the regression in logs.

3.4. Economic influences on the border

A major objective of this study is to examine the evolution of the border effect and determine whether it is influenced by identifiable economic factors. Towards that end, we examine price dispersion year-by-year. More formally, we adopt a measure of the range of possible differentials that is specific to a given city-pair and year. We make it year-specific by pooling over information from the 27 goods and four quarters in a given year.

Recall the change in the real exchange rate (for good k) relative to benchmark city j is:

$$Q(ij,k,t) \equiv \Delta \ln P(i,k,t) - \Delta \ln P(j,k,t)$$

where ij represents a city-pair, and t is time.

Prior to calculating variability we remove the good-specific fixed effects by regressing the vector of Q values on individual good dummies (for Q values over all goods and all quarters in that year, for that city-pair). Let $q(ij,k,t)$ be the residuals from that regression. We compute the standard deviation of q as our measure of variability. As noted above, for robustness, we later adopt an alternative measure of dispersion across cities — the inter-quartile range, defined as the 75th–25th percentiles of the distribution of q . Fig. 4 plots the average (across relevant city-pairs) dependent variable through time.

We begin our investigation by estimating:

$$V(q(ij,k,t)) = \beta_1 \ln(\text{dist}_{ij}) + \beta_2 \text{Border}_{ij} + \text{a constant and city dummies} + \epsilon_{ij,t} \quad (3)$$

Note this regression involves 189 city-pairs, each with 22 time periods, and individual good effects have been removed as described above.¹⁰

The second column in Table 1 reports results from this regression. We again confirm that price dispersion increases with distance and that a border effect exists between the US and Japan. Both estimates are of the hypothesized sign and statistically significant. Using our revised calculation procedure from Eq. (2), the ‘Border’ adds as much as 15 billion miles does to the within-country volatility; again, a very large number.

¹⁰Unlike the Engel–Rogers type regression (reported in column 1 of Table 1), this specification has a time series dimension. In column 1, the variability of relative prices is measured across time for a given city-pair and good; hence the maximum number of observations in the pooled regression will be equal to the number of goods \times number of city-pairs, or $27 \times 189 = 5103$. To study the evolution of the border effect, we compute the variability across (de-meaned) goods for each quarter in a given year (hence for a given city-pair, each year’s variability is computed using 4×27 observations). In estimation, we pool across city-pairs; hence the maximum number of observations in the pooled regression will be equal to the number of time periods \times number of city-pairs, or $22 \times 189 = 4158$.

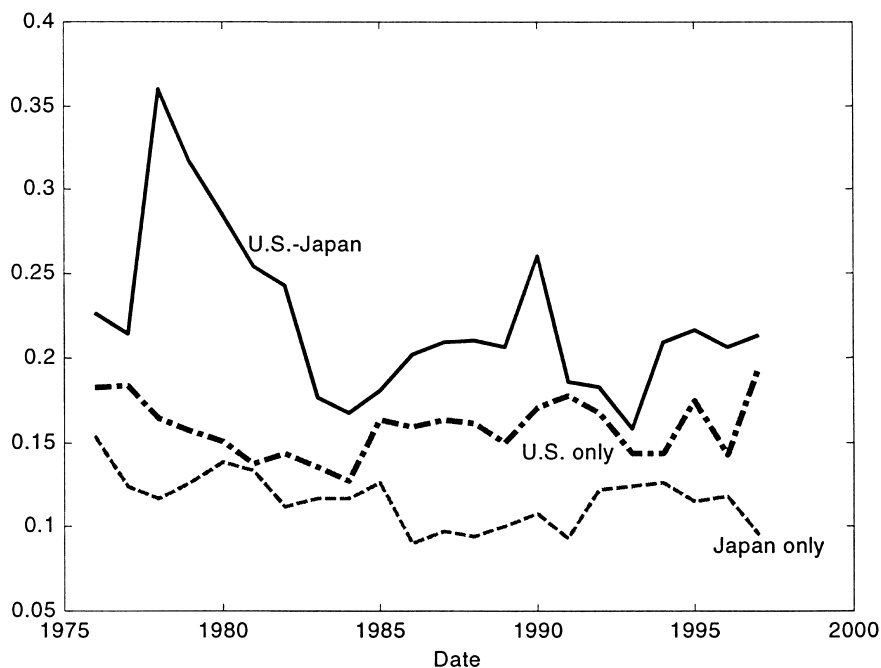


Fig. 4. Price dispersion averaged over relevant city-pairs.

The next three specifications in Table 1 examine potential economic explanations for this sizeable effect. We make an attempt to measure *explicitly and directly* three such factors: the unit costs of transportation and insurance, the variability of nominal exchange rate, and the variability of the relative wage differential.¹¹

We begin by considering exchange rate volatility. Exchange rate volatility is defined as the standard deviation of changes in the (log) nominal exchange rate. The results are reported in the third column of Table 1. As expected, exchange rate volatility has a positive and significant effect on cross-country price dispersion. More importantly however, note that the coefficient estimate of the border dummy declines. Apparently, exchange rate volatility has an important impact, but does not completely account for the border effect.

Next we turn to shipping and insurance costs. We hypothesize that the log of the shipping and insurance cost is the sum of two components: one depends on the log of distance, which has already been included in the regression, and the other is the cost per unit of distance. We concentrate on the second component here. For the international part of the unit cost, we use information on the difference between

¹¹Engel and Rogers (1996) examine the variability of the wage differential explicitly, but they infer the effect of exchange rate volatility only indirectly.

c.i.f. and f.o.b. values of bilateral US trade with Japan as a percentage of the total f.o.b. value.¹² Specifically we collect data on (1) unit shipping and insurance costs on US exports to Japan, and (2) unit shipping and insurance costs on Japanese exports to the US. Our measure of shipping and insurance costs is the average of (1) and (2). For the domestic (i.e., Japan only or US only) part of the unit cost, we have no direct observations. In this case, we assign a value equal to one-half the minimum of the international shipping cost. This is arbitrary. However, in Parsley and Wei (2000), we present an example based on quotes from United Parcel Service and the US Postal Service that the ratio of domestic to international shipping costs over comparable distance is between 0.3 and 0.7 (see Table A3). Additionally, we note that assigning a value of zero would exaggerate the transportation cost between international city-pairs (and hence might artificially explain too much of the border effect).

In the fourth column (labeled column 3) of Table 1 we add our measure of unit-shipment costs to the specification. As expected, the coefficient estimate is positive, and the estimate is highly statistically significant. Moreover, adding shipping costs has resulted in a further drop in the border estimate, and a further increase in the equation's \bar{R}^2 . Combined, these two variables account for a substantial portion of the border effect.

In the last two columns we consider two measures of the variability of relative wages. Here we are trying to get at the non-traded component of goods prices. For international city-pairs this variable is defined as the standard deviation during the year of the difference in the US and Japanese change in log (common currency) wage rates. Since this international component of wage variability is highly correlated with nominal exchange rate variability (correlation coefficient=0.98), we also consider (in the final column) an alternative measure — the difference in national currency denominated wage inflation rates. This variable isolates that part of relative wage variability not accounted for by nominal exchange rate movements.

For the intra-national component of both measures of relative wage variability we collect annual data from the BLS on average earnings in manufacturing for all the US states in our sample.¹³ Then, for each year we take the cross-sectional standard deviation of the first difference in logs of these state-level data. We assume Japanese intra-national wage variability is equal to that for the US. In the regression reported in the penultimate column, the coefficient on the variability of the wage difference is positive and statistically significant. However, now the coefficient on nominal exchange rate variability has turned negative, presumably due to the collinearity. Hence in the final column we add our second measure of

¹²We obtained the data from various issues of the Direction of Trade publication of the IMF.

¹³The source for the international wage index data is the IFS, lines 65 for Japan, and 65ey for the US. The source for the US state level data is the US Bureau of Labor Statistics, <http://146.142.4.24/cgi-bin/surveymost?sa>.

wage variability. Note that, while the coefficients are positive and statistically significant on both measures of wage variability, the border coefficient actually increases in both regressions (compared to Eq. (3)).

3.5. *Declining border?*

Table 1 documents a positive effect of distance on relative price variability, and a positive border effect between Japan and the US. The table also demonstrates that the border effect is positively related to economic factors, in particular to exchange rate variability and unit-shipping costs. These results refer to the ‘average’ border effect, i.e., over the sample period as a whole. In this section, we ask whether the border effect has changed over time, and have economic factors contributed to this evolution. To get at these questions we augment our basic specification (Eq. (3)) with a linear trend term and two trend-interaction terms: one for border, and one for distance.¹⁴ In this specification the coefficient on the border dummy now captures the border effect at the beginning of our sample, 1976. These results are reported in column 1 of Table 2.

The negative estimate for the trend/border interaction term suggests that the border effect is declining over time at about 0.4% per year. The coefficient estimate on log distance is slightly smaller than in Table 1 and is no longer statistically significant. Over time there has been a statistically significant trend decline in relative price variability for intra-national as well as international city-pairs.

We proceed as before by sequentially considering economic factors that vary through time. In column 2 we add nominal exchange rate variability. As in Table 1, the point estimate on the border dummy declines. All other coefficient estimates remain virtually unchanged. In column 3 we add unit-shipping costs. As with nominal exchange rate variability, adding unit-shipping costs to the specification results in a decline in the border dummy’s coefficient estimate. Finally, we add wage variability in the last two columns in the table. The results are qualitatively the same as in Table 1. In particular, there is little evidence of an independent impact of wage variability on the border effect.

To summarize, Table 2 corroborates several findings from Table 1. In particular, the table demonstrates that a positive border effect between Japan and the US existed in 1976. The table also confirms that the border effect is positively related to economic factors, especially to exchange rate variability and to unit shipping costs. One new finding in Table 2 is that international market segmentation is declining over time. More importantly, the table documents that the economic factors we consider go only part of the way toward understanding the evolution of the border effect through time. Ultimately both the trend decline in relative price

¹⁴We also considered a squared log distance term, but it was never significant in any of our specifications.

Table 2
Explaining the border effect through time^a

	Eq. (1)	Eq. (2)	Eq. (3)	Eq. (4)	Eq. (5)
Log distance	0.0017 (0.0023)	0.0017 (0.0023)	0.0018 (0.0023)	0.0018 (0.0023)	0.0018 (0.0023)
Border	0.1164 (0.0081)	0.0986 (0.0080)	0.0644 (0.0086)	0.0896 (0.0099)	0.0647 (0.0086)
Trend	-0.0023 (0.0007)	-0.0022 (0.0007)	-0.0022 (0.0007)	-0.0025 (0.0007)	-0.0022 (0.0007)
Trend×log distance	0.0003 (0.0001)	0.0003 (0.0001)	0.0003 (0.0001)	0.0003 (0.0001)	0.0003 (0.0001)
Trend×border	-0.0040 (0.0005)	-0.0043 (0.0005)	-0.0040 (0.0005)	-0.0038 (0.0005)	-0.0041 (0.0005)
Nominal exchange rate variability		0.4722 (0.0442)	0.4454 (0.0445)	-0.8786 (0.2098)	0.4452 (0.0445)
Unit shipping costs			0.5303 (0.0565)	0.3192 (0.0685)	0.5326 (0.0572)
Wage variability				1.2922 (0.2046)	
Relative wage inflation variability					-0.0872 (0.1898)
City dummies	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	0.60	0.61	0.62	0.63	0.62
Number of observations	3820	3820	3820	3820	3820

^a Dependent variable: standard deviation of log price differential; Tokushima–Louisville benchmark cities.

variability common to intra- and international city-pairs, and the relatively faster trend decline in relative price variability specific to international city-pairs, remain unexplained. We turn next to some robustness checks.

3.6. Extensions and robustness checks

In the regressions so far, we stack data for different city-pairs. Potentially correlated errors across city-pairs for the same year could lead to underestimated standard errors. In an effort to address this issue we implement a systems-estimation using the seemingly unrelated regressions method.

We select the first 10 international city-pairs containing no missing values using the Tokushima and Louisville benchmark. Thus the resulting system has 20 equations.¹⁵ We allow the intercept to be different in each equation, and hence, to be different for each city-pair. We impose the restriction that the coefficients on all other regressors are the same. With this specification, all time invariant and

¹⁵Using all city-pairs would lead to a singular variance–covariance matrix.

Table 3
Seemingly unrelated regressions estimation^a

	Eq. (1)	Eq. (2)	Eq. (3)	Eq. (4)
Trend	-0.0047 (0.0001)	-0.0045 (0.0002)	-0.0042 (0.0003)	-0.0046 (0.0003)
Nominal exchange Rate variability	0.4682 (0.0206)	0.4227 (0.0595)	0.4172 (0.0761)	0.4200 (0.0862)
Unit shipping costs		0.5965 (0.0672)	0.5359 (0.0878)	0.7564 (0.1036)
Wage variability			1.2842 (0.6987)	
Relative wage inflation variability				4.0439 (0.7957)
Equation specific intercepts	Yes	Yes	Yes	Yes
Average adjusted R^2	0.211	0.225	0.243	0.239
Number of equations	20	20	20	20
Number of observations	440 (=20×22)	440	440	440

^a Tokushima–Louisville benchmark.

city-pair-specific effects (e.g., distance and border effects) will be absorbed in the 20 intercepts.

In Table 3 we proceed sequentially as before, beginning with exchange rate variability. First note that the estimate of the trend effect (-0.0047) on international market segmentation is somewhat smaller than that implied by Table 2 ($-0.0063 = -0.0023 - 0.0040$). Next, nominal exchange rate variability is added; the reported coefficient estimate is virtually identical to that in the OLS regressions, and remains highly statistically significant. In column 2, we add unit-shipping costs. The coefficient on this variable is essentially the same as before. In the final columns we add wage variability. Overall, the results mirror those in Table 2 except that here — for international city pairs — relative wage inflation variability (measured exclusive of nominal exchange rate variability) is positively and statistically significantly related to relative price variability. Hence we conclude that, at least for international city pairs, all three economic variables are important in segmenting international markets.

We try two more extensions to test the robustness of our results. First, we repeat the analysis of Table 1 using a different measure of the variability of relative prices — the inter-quartile range of the distribution. A second robustness test is to repeat the analysis using two different benchmark cities. We selected Osaka and Houston, partly because Houston, like Louisville had only two quarters of missing values. Space considerations prohibit a full discussion of these results here; they are available in Parsley and Wei (2000). Here, we simply note that the basic findings reported in Tables 1–3 are unaffected by either the alternative definition of the dependent variable, or by the alternative choice of benchmark cities.

4. Concluding remarks

This paper exploits a three-dimensional panel data-set of prices on 27 traded goods, over 88 quarters, across 96 cities in the US and Japan. We present evidence that the distribution of intra-national real exchange rates is substantially less volatile and on average closer to zero, than the comparable distribution for international relative prices. We also show that an equally weighted average of good-level real exchange rates tracks the nominal exchange rate well, suggesting strong evidence of sticky prices.

We turn next to economic explanations for this so-called border effect and to its dynamics. Focusing on dispersion in prices between city-pairs, we confirm previous findings that crossing national borders adds significantly to price dispersion. Using our point estimates crossing the US–Japan ‘Border’ is equivalent to adding as much as 43 000 trillion miles to the cross-country volatility of relative prices. We examine several potential economic influences on the border effect. In our calculations, the estimated border effect declines substantially after controlling for the effects of distance, unit-shipping costs, and exchange rate variability. We find evidence of a declining trend in international market segmentation that remains even after controlling for unit-shipping costs and exchange rate variability. Finally, we also conclude that relative wage variability has little independent impact on the segmentation of international markets.

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