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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 U.S. 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 U.S.-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.

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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 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 U.S. and Canada in their example), the dispersion of prices goes far beyond distance (and hence transportation costs): crossing the U.S.-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 U.S. states and the southern Canadian provinces are presumably less important than those existing among east and west coast U.S. 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). 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 three to five years for the half-life of PPP deviations. This is in strong contrast to half-life estimates based on purely intra-U.S. prices. Parsley and Wei (1996) estimated that the half-life of deviations from the LOP is only about one year. They show that the convergence rate does slow down as the (physical) distance between price observations increases. However, despite the fact that the distance between international price observations tends to be greater than that for prices observed intra-nationally, they find that distance alone cannot explain the difference in convergence rates.

There is an analogue in studies using international trade quantity data to this price-data-based PPP, or LOP, literature. Using the value of exports and imports, McCallum (1995) showed that trade between Canadian provinces is 2200% larger than between Canadian provinces and U.S. states of similar distance (and sizes). Helliwell (1998) and Wei (1996) showed that the home bias in the goods market is equally non-negligible when they examine trade between and within OECD countries.

Crucini, Telmer, and Zachariadis (1999) 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 among the European countries in their sample were managed, 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 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 U.S. 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.

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[1] Despite formal exchange rate agreements, intra-European currency movements were surprisingly large over longer periods. For example, the 1980-84 percentage changes (versus the deutchmark) 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 percent for that period. Source: IFS 1999 Yearbook.

[2] A subset of the U.S. data has been examined in Parsley and Wei (1996), and O'Connell and Wei (1997).

[3] Of course the U.S. 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.

We ask two related questions. First, is there any evidence that the Japan-U.S. "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 a 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 U.S. 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 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 U.S. data set; to assure geographic coverage and comparability with the U.S. sample, we also selected 48 Japanese cities. There is still a slight time mismatch however. The U.S. data are generally sampled seven to ten days prior to the Japanese data. For every quarter in our sample (1976.1 – 1997.4), all forty-eight Japanese cities were part of the sample.

The source for the U.S. 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 forty-eight U.S. cities and the twenty-seven 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 U.S. cities (currently exceeding 300). We selected U.S. cities that appeared in roughly 90 percent of the quarterly surveys. This data set is described in more detail in Parsley and Wei (1996). Briefly, for this study we selected forty-eight U.S. cities and the twenty-seven 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 U.S. cities (currently exceeding 300). We selected U.S. cities that appeared in roughly 90 percent of the quarterly surveys. Appendix tables A1 and A2 list the goods and cities in the U.S. and in Japan that we include. 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 United States (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 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 U.S. 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

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

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

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 U.S. city-pairs, (2) $r(ij,k,t)$ for intra-national Japanese city-pairs, and (3) $r(ij,k,t)$ for international city-pairs.

Figure 1 plots the empirical kernel density estimate of the log average real exchange rate for each of our three comparisons (within Japan, within the U.S., and between the U.S. 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 United States. This is possibly due to the relative sizes of the two countries; the greater average distance between cities in the U.S. may allow prices to vary more. Judging by this figure, deviations from the LOP within a country do not appear extraordinary. And second, the U.S.-Japan density function is centered to the left of zero. This means that in 1985 most Japanese prices were higher than U.S. prices. It also suggests the Crucini, et al. (2000) finding may be specific to Europe[5].

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

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-U.S. 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 U.S. We can define $\bar{r}(japan,t)$ and $\bar{r}(us-japan,t)$ in an analogous way.

The left panel of appendix table A3 presents, and Figure 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 U.S. and Japan, are fairly close to zero. In fact they vary within plus/minus 5-7

[5] Other 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 seven ounce bottle of shampoo is probably not seven-tenths the price of a ten ounce bottle); and the fact that their cross section is much more extensive.

percent in each of the twenty-two years in our sample. In comparison, the average percentage deviation between U.S. 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 U.S. 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 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 .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 – although we note that the comparison between our U.S.-Japan sample and their European sample may not be entirely appropriate since the two data sets are not strictly comparable.

It may be also useful to gauge absolute deviations from the LOP. For a given city-pair *(i,j)*, a given good *k*, and a time period *t*, the absolute deviation is defined as:

$$X(ij,k,t) = \left|P(i,k,t) - P(j,k,t)\right|$$

Let $X(us,t)$ = the mean absolute deviation for the U.S. at time t be

$\mathrm{X}(\mathrm{us},\mathrm{t}) = \frac{1}{\mathrm{KN}}\sum_{\mathrm{ij,k}}\mathrm{X}(\mathrm{ij},\mathrm{k},\mathrm{t})$, where the sum is over all *(i,j)* U.S. pairs, and over all goods *k*.

We can define $X(japan,t)$, and $X(us - japan,t)$ analogously. In the right-hand-side of appendix table A3, we present evidence on the mean absolute percentage deviation from the LOP. Once again, we see the same pattern. Within each country, the mean absolute deviations are between 10-15% (somewhat larger in the United States than in Japan). However, the cross-country mean absolute deviations are several times as large, between 75%-140%.

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

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 violation can 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 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 no-arbitrage story can be made more formal. O'Connell and Wei (2000) present a continuous-time model in which an arbitrager solves an explicit optimization problem. The exact dynamics of the percentage price difference series depends on the structure of the arbitrage cost. They present three cases. In case 1, arbitrage involves a constant variable cost, but zero fixed cost (the so-called "iceberg" assumption on transport cost). The price difference follows a bounded Brownian motion process. The two bounds are determined by the variable arbitrage cost. Each time the price difference hits one of the boundaries an infinitesimal amount of arbitrage takes place to bring the difference back to just within the band. In case 2, arbitrage involves a constant fixed cost, but zero variable cost. Each time the price difference hits one of the boundaries, a discreet amount of arbitrage activity takes place to bring the price difference to the center of the no-arbitrage band (namely the point of zero price difference). And in case 3, arbitrage involves both fixed and variable cost. A constrained Brownian motion process with four boundaries can characterize the price difference series: two outer and two inner boundaries. Whenever the price difference hits one of the outer boundaries, a discreet amount of arbitrage activity takes place to bring it to the closest inner boundary.

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 use 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 U.S.

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. Appendix table A4 presents summary statistics on the dispersion of $\mathrm{Q(ij,k,t)}$. We are especially interested in intra-national versus international comparisons. In the table, we report averages for Japanese-only, and U.S.-only city pairs, and we similarly average over all cross-country city pairs. Looking across the columns we see that as suggested by Figures 1 and 2, the percentage deviations within Japan or within the U.S. are smaller than for the international city pairs.

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 U.S.-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,

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

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 United States is the U.S. Naval Observatory.[7] Note that this regression will have (without missing values) 5103 observations (27 goods x 189 city-pairs).

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

Engel and Rogers calculate that the U.S.-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 the Japan-U.S. case; the number is roughly 6.5 trillion miles, or about 70,000 times the distance from the Earth to the Sun.[8]

In these calculations however, when one changes the units of distance measurement from

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[7] The 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 U.S. data.

[8] For 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 U.S. 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 β_1, β_2 because distance enters the regression in logs.

(say) miles to kilometers, the interpretation of the distance equivalent changes, from 75,000 miles to 75,000 kilometers. This is because the point estimates of β_1 and β_2, in a log linear regression such as equation (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:

$$\beta_1 \, ln\left(\overline{distance} + Z\right) = \beta_2 + \beta_1 \, ln\left(\overline{distance}\right), \tag{2}$$

where $\overline{distance}$ is the average distance between the U.S.-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 $Z = \overline{distance} * \left(exp(\beta_2/\beta_1) - 1\right)$. Substituting the values of β_1, β_2 and average distance reported by Engel-Rogers, raises their border estimate (between Canada and the U.S.) significantly, from 75,000 miles to 101 million miles. Similarly, using our parameter estimates in Table 1 the "Border" between Japan and the U.S. becomes nearly 43,000 trillion miles! The "Border" is indeed remarkably wide.

Of course, Japan and the U.S. are farther apart than Canada and the U.S. In fact, the average distance between our international city-pairs is over six times that between the U.S. 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/U.S. dollar rate, and the relative wage differential is also likely to be more variable between Japan and the U.S. We turn to these issues next.

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 twenty-seven 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*'s on individual good dummies (for *Q*'s 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 (see Table 4 below) adopt an alternative measure of dispersion across cities – the inter-quartile range, defined as the 75th – 25th percentiles of the distribution of q.

We begin our investigation by estimating:

$$V\big(q(ij,k,t)\big) = \beta_1 \, ln(dist_{ij}) + \beta_2 Border_{ij} + \textit{a constant and city dummies} + \varepsilon_{ij,t}\,. \tag{3}$$

Note this regression involves 189 city-pairs, each with twenty-two time periods, and individual good effects have been removed as described above.[9]

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 United States and Japan. Both estimates are of the hypothesized sign and statistically significant. Using our revised calculation procedure from equation (2), the "Border" adds as much as 15 billion miles does to the within-country volatility; again, a very large number.

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.[10]

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

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[9] Unlike 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 x number of city-pairs, or 27x189=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 4x27 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 x number of city-pairs, or 22x189=4158.

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 c.i.f. and f.o.b. values of bilateral U.S. trade with Japan as a percentage of the total f.o.b. value.[11] Specifically we collect data on (1) unit shipping and insurance costs on U.S. exports to Japan, and (2) unit shipping and insurance costs on Japanese exports to the U.S. Our measure of shipping and insurance costs is the average of (1) and (2). For the domestic (i.e., Japan only or U.S. 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 appendix table A5, we present an example based on quotes from United Parcel Service and the U.S. Postal Service that the ratio of domestic to international shipping costs over comparable distance is between .3 and .7. 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-shipping 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 $\overline{R}^2$. 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 U.S. 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 = .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

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

[11] We obtained the data from various issues of the Direction of Trade publication of the IMF.

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 U.S. states in our sample.[12] 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 U.S. 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 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 Equation 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 U.S. 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 (equation 3) with a linear trend term and two trend-interaction terms: one for border, and one for distance.[13] 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 percent 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

[12] The source for the international wage index data is the IFS, lines 65 for Japan, and 65ey for the U.S. The source for the U.S. state level data is the U.S. Bureau of Labor Statistics, http://146.142.4.24/cgi-bin/surveymost?sa.

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

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 U.S. 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 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 ten international city-pairs containing no missing values using the Tokushima and Louisville benchmark. Thus the resulting system has twenty equations.[14] 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 city-pair specific effects (e.g., distance and border effects) will be absorbed in the twenty 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

[14] Using all city-pairs would lead to a singular variance-covariance matrix.

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. Once again, the coefficient on this variable is essentially the same as before. In the final columns we add wage variability. 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 thee 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. These results are presented in Table 4. 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. These results are reported in Table 5.

The basic findings reported earlier are unaffected by either the alternative definition of the dependent variable, or by the alternative choice of benchmark cities. In particular, international price dispersion is significantly greater than intra-national price dispersion, even after controlling for distance; i.e., a border effect exists. Also, sequentially adding shipping costs and exchange rate volatility produces a smaller and smaller border effect. In our inter-quartile range regressions (Table 4) the coefficient on the trend-border interaction term is insignificant, but continues to have a negative coefficient. In these two alternative specifications the coefficient on wage variability again fails to reduce the border effect. Finally, we again conclude that despite a substantial decline in the estimated border effect due to the inclusion of these economic factors, a statistically significant effect remains.

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 U.S. 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 a simple 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 U.S.-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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Table 1. Explaining the Average Border Effect
Dependent Variable: Standard Deviation of Log Price Differential
Tokushima-Louisville Benchmark Cities

	Engel-Rogers Benchmark Regression	***Equation 1***	***Equation 2***	***Equation 3***	***Equation 4***	***Equation 5***
Log Distance	0.0022	0.0049	0.0049	0.0049	0.0049	0.0049
	(0.0017)	(0.0018)	(0.0018)	(0.0018)	(0.0018)	(0.0017)
Border	0.0649	0.0717	0.0601	0.0154	0.0291	0.0219
	(0.0055)	(0.0057)	(0.0059)	(0.0066)	(0.0064)	(0.0066)
Nominal Exchange Rate Variability			0.2625	0.2458	-0.4183	0.2773
			(0.0434)	(0.0484)	(0.2184)	(0.0472)
Unit Shipping Costs				0.7509	0.6476	0.6747
				(0.0571)	(0.0676)	(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

Table 2. Explaining the Border Effect Through Time
Dependent Variable: Standard Deviation of Log Price Differential
Tokushima-Louisville Benchmark Cities

	Equation 1	*Equation 2*	*Equation 3*	*Equation 4*	*Equation 5*
Log Distance	0.0017	0.0017	0.0018	0.0018	0.0018
	(0.0023)	(0.0023)	(0.0023)	(0.0023)	(0.0023)
Border	0.1164	0.0986	0.0644	0.0896	0.0647
	(0.0081)	(0.0080)	(0.0086)	(0.0099)	(0.0086)
Trend	-0.0023	-0.0022	-0.0022	-0.0025	-0.0022
	(0.0007)	(0.0007)	(0.0007)	(0.0007)	(0.0007)
Trend*Log Distance	0.0003	0.0003	0.0003	0.0003	0.0003
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Trend*Border	-0.0040	-0.0043	-0.0040	-0.0038	-0.0041
	(0.0005)	(0.0005)	(0.0005)	(0.0005)	(0.0005)
Nominal Exchange Rate Variability		0.4722	0.4454	-0.8786	0.4452
		(0.0442)	(0.0445)	(0.2098)	(0.0445)
Unit Shipping Costs			0.5303	0.3192	0.5326
			(0.0565)	(0.0685)	(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

Table 3. Seemingly Unrelated Regressions Estimation
Tokushima-Louisville Benchmark

	Equation 1	*Equation 2*	*Equation 3*	*Equation 4*
Trend	-0.0047	-0.0045	-0.0042	-0.0046
	(0.0001)	(0.0002)	(0.0003)	(0.0003)
Nominal Exchange Rate Variability	0.4682	0.4227	0.4172	0.4200
	(0.0206)	(0.0595)	(0.0761)	(0.0862)
Unit Shipping Costs		0.5965	0.5359	0.7564
		(0.0672)	(0.0878)	(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	.211	.225	.243	.239
Number of equations	20	20	20	20
Number of observations	440 (=20x22)	440	440	440

Table 4. Explaining the Border Effect Through Time
Dependent Variable: Inter-quartile Range of Log Price Differential
Tokushima-Louisville Benchmark Cities

	Equation 1	*Equation 2*	*Equation 3*	*Equation 4*	*Equation 5*
Log Distance	0.0085	0.0086	0.0086	0.0086	0.0086
	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)
Border	0.0734	0.0620	0.0533	0.0519	0.0519
	(0.0070)	(0.0075)	(0.0081)	(0.0081)	(0.0081)
Trend	0.0041	0.0041	0.0041	0.0040	0.0040
	(0.0006)	(0.0006)	(0.0006)	(0.0006)	(0.0006)
Trend*Log Distance	-0.0005	-0.0005	-0.0005	-0.0005	-0.0005
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Trend*Border	-0.0001	-0.0002	-0.0002	-0.0002	-0.0002
	(0.0005)	(0.0005)	(0.0005)	(0.0005)	(0.0005)
Nominal Exchange Rate Variability		0.2762	0.2660	0.2729	0.2729
		(0.0473)	(0.0463)	(0.0465)	(0.0465)
Unit Shipping Costs			0.1372	0.1406	0.1406
			(0.0483)	(0.0491)	(0.0491)
Wage Variability				-0.1552	
				(0.1660)	
Relative Wage Inflation Variability					-0.1552
					(0.1660)
City Dummies	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	0.72	0.72	0.72	0.74	0.74
Number of observations	3840	3840	3840	3840	3840

Table 5. Explaining the Border Effect Through Time
Dependent Variable: Standard Deviation of Log Price Differential
Osaka-Houston Benchmark Cities

	Equation 1	*Equation 2*	*Equation 3*	*Equation 4*	*Equation 5*
Log Distance	0.0059	0.0059	0.0059	0.0059	0.0059
	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)
Border	0.0999	0.0788	0.0518	0.0697	0.0535
	(0.0069)	(0.0068)	(0.0074)	(0.0093)	(0.0074)
Trend	-0.0005	-0.0005	-0.0005	-0.0006	-0.0005
	(0.0006)	(0.0006)	(0.0006)	(0.0006)	(0.0006)
Trend*Log Distance	0.0000	0.0000	0.0000	0.0000	0.0000
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Trend* Border	-0.0023	-0.0026	-0.0024	-0.0023	-0.0025
	(0.0005)	(0.0005)	(0.0006)	(0.0005)	(0.0005)
Nominal Exchange Rate Variability		0.5611	0.5398	-0.4019	0.5864
		(0.0448)	(0.0436)	(0.2090)	(0.0379)
Unit Shipping Costs			0.4207	0.2706	0.3879
			(0.0506)	(0.0640)	(0.0514
Wage Variability				0.9192	
				(0.2007)	
Relative Wage Inflation Variability					0.0858
					(0.2486)
City Dummies	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	0.65	0.66	0.67	0.67	0.67
Number of observations	3833	3833	3833	3833	3833

Table A1. Correspondence of Japanese and United States Goods

Good	*Japanese Prices*	*U.S. Prices*
1	Canned tuna fish, in oil, #4 can, 80 g	Canned tuna (6.5 oz)
2	Beef loin (100g)	Steak (lb)
3	Beef shoulder (100g)	Ground Beef (lb)
4	Chicken, broiler, leg (100g)	Whole Chicken
5	Bacon, side, (100g)	Bacon (lb); Sausage
6	Fresh milk in carton (1000ml)	Milk (1/2 gal)
7	Processed cheese in carton, 'Snow brand Hokkaido cheese' (225 g)	Parmesan Cheese
8	Hen eggs, 1 kg	Eggs (1 dozen, large)
9	Lettuce, head	Lettuce, head
10	White potatoes, 1 kg	Potatoes, white or red
11	Tomatoes, 1 kg	Canned tomatoes, Del Monte or Green Giant
12	Bananas, 1 kg	Bananas, (lb)
13	Margarine, 1 carton	Margarine (lb)
14	Sugar, white, packaged 1 kg	Sugar, white, packaged (5 lb)
15	Instant coffee	Ground coffee, (2 lb), Maxwell House, Folgers
16	100% fruit drinks, Valencia orange juice, in cartons (1000 ml)	Canned orange juice (6 oz)
17	Cola Drinks, canned, (350 ml)	Soft drink (2 ltr)
18	Whisky, imported	Liquor (Seagrams 7 Crown; J&B scotch)
19	Wine, 1 bottle	Wine (1.5 liter)
20	Beer, in restaurant	Beer in store (6 pack)
21	Tissue (facial), 1 pouch	Facial tissue, 175 count box
22	Laundry detergent, for cotton, hemp, rayon and synthetic fiber, high density, in box (1.25 kg)	Washing powder (49 oz), Tide, Bold, or Cheer
23	Men's slacks, denim jeans, 100% cotton, 29~31"	Jeans, Levis
24	Men's long sleeve business shirts	Man's shirt, Arrow or Van Heusen
25	Men's briefs, 100% cotton, ordinary quality	Men's briefs, package of 3
26	Shampoo, Kao Essential, 220 ml	Shampoo, VO-5, 15 oz
27	Toothpaste, 170g, Denter Lion	Toothpaste, Crest or Colgate, 6 oz.

Table A2. List of Japanese and United States Cities

	Japanese Cities	*U.S. Cities*
1	Sapporo	Birmingham AL
2	Aomori	Mobile AL
3	Morioka	Blythe CA
4	Sendai	Indio CA
5	Akita	Palm Springs CA
6	Yamagata	Denver CO
7	Fukushima	Lakeland FL
8	Utsunomiya	Boise ID
9	Maebashi	Champaign-Urbana IL
10	Urawa	Peoria IL
11	Chiba	Ft. Wayne IN
12	Ku-area of Tokyo	Indianapolis IN
13	Yokohama	Cedar Rapids IA
14	Niigata	Lexington KY
15	Toyama	Louisville KY
16	Kanazawa	Baton Rouge LA
17	Fukui	Lafayette LA
18	Kofu	New Orleans LA
19	Nagano	Benton Harbor MI
20	Gifu	Traverse City MI
21	Shizouka	Columbus MS
22	Nagoya	St. Jopseph MO
23	Tsu	St. Louis MO
24	Otsu	Falls City NE
25	Kyoto	Hastings NE
26	Osaka	Omaha NE
27	Kobe	Reno, Sparks NV
28	Himeji	Newark NJ
29	Itami	New York NY
30	Nara	Hickory NC
31	Wakayama	Columbus OH
32	Tottori	Altoona PA
33	Matsue	Rapid City SD
34	Okayama	Vermillion SD
35	Hiroshima	Chattanooga TN
36	Yamaguchi	Knoxville TN
37	Tokushima	Abilene TX
38	Takamatsu	EL Paso TX
39	Matsuyama	Ft. Worth TX
40	Kochi	Houston TX
41	Fukuoka	Lubbock TX
42	Saga	Salt Lake city UT
43	Nagasaki	Charleston WV
44	Kumamoto	Appleton WI
45	Oita	Eau Claire WI
46	Miyazaki	Madison WI
47	Kagoshima	Oshkosh WI
48	Naha	Casper WY

Table A3. International vs. Intra-national Shipping Costs
Over Comparable Distance

		Distance (1,000 miles)	Air (UPS)	Sea/Ground (U.S. Postal)
(1)	Boston – Lisbon	3.2	$226.50	$34.30
(2)	DC – Caracas	2.0	$401.75	$34.20
(3)	Average of (1) and (2)	2.6	$314.13	$34.20
(4)	Boston – San Diego	26	$95.25	$24.21
(5)	Domestic/International Cost Ratios (row 4/row 3)	1.0	0.30	0.71

Notes:

1. UPS Shipping Cost, 20 lb. Package, 24x20x20 inches, Domestic 2 day air, International 4 day air, comparable class service.
   Source: http://www.ups.com/using/services/rave/rate.html

2. U.S. Postal Service Postal Cost, 20 lb. Package 24x20x20 inches, Domestic shipment, 4-5 days; International shipment, 4-6 weeks.

Table A4. Measures of Intra-national and International Price Deviations
(Simple Average over Traded Goods and City Pairs)

	Average of Good-Level Log Real Exchange Rates			Average of Good-Level Absolute Percentage Deviations		
Year	***Japan***	***U.S.***	***U.S -Japan***	***Japan***	***U.S.***	***U.S. Japan***
1976	0.0312	-0.0267	-0.9617	0.1230	0.1494	0.9968
1977	0.0125	-0.0463	-0.9885	0.1111	0.1406	1.0315
1978	0.0228	-0.0240	-0.9713	0.1195	0.1276	1.2087
1979	0.0227	-0.0388	-0.7251	0.1105	0.1105	0.9866
1980	-0.0054	-0.0185	-0.6738	0.1072	0.1167	0.9661
1981	-0.0021	-0.0418	-0.7369	0.1036	0.1368	1.0010
1982	0.0135	0.0215	-0.3997	0.0949	0.1229	0.7805
1983	0.0304	-0.0066	-0.4518	0.1055	0.1227	0.7889
1984	0.0425	-0.0102	-0.4279	0.1047	0.1326	0.7662
1985	0.0110	-0.0096	-0.4405	0.1300	0.1453	0.7967
1986	-0.0053	0.0157	-0.7256	0.1262	0.1272	0.9892
1987	-0.0181	-0.0036	-0.8221	0.1266	0.1465	1.0568
1988	-0.0304	-0.0080	-0.9454	0.1234	0.1416	1.1606
1989	-0.0220	-0.0044	-0.8720	0.1260	0.1379	1.0930
1990	0.0125	0.0595	-0.8949	0.1131	0.1498	1.1171
1991	0.0008	0.0523	-0.9954	0.1065	0.1727	1.1939
1992	0.0259	0.0618	-1.0900	0.1130	0.1591	1.2682
1993	0.0187	0.0587	-1.2123	0.1065	0.1561	1.3657
1994	0.0034	0.0308	-1.2523	0.1055	0.1436	1.3920
1995	0.0160	0.0440	-1.3056	0.1115	0.1506	1.3953
1996	0.0045	0.0665	-1.2036	0.1153	0.1523	1.2183
1997	0.0045	0.0361	-1.0406	0.1135	0.1549	1.1801
Average	0.0086	0.0095	-0.8699	0.1135	0.1408	1.0797

Table A5. Variability in Relative Prices
Tokushima-Louisville benchmark city

	Std. Dev. of the diff. in log prices			Interquartile range of the diff. in log prices		
Year	***Japan only***	***U.S. only***	***U.S-Japan***	***Japan only***	***U.S. only***	***U.S.-Japan***
1976	0.1541	0.1828	0.2264	0.0548	0.1945	0.2061
1977	0.1238	0.1836	0.2139	0.0436	0.1721	0.1977
1978	0.1167	0.1638	0.3596	0.0358	0.1587	0.2540
1979	0.1261	0.1567	0.3166	0.0282	0.1552	0.1702
1980	0.1378	0.1502	0.2850	0.0704	0.1459	0.2102
1981	0.1332	0.1375	0.2545	0.0507	0.1445	0.1861
1982	0.1120	0.1431	0.2431	0.0315	0.1511	0.1736
1983	0.1171	0.1355	0.1766	0.0409	0.1467	0.1764
1984	0.1169	0.1270	0.1667	0.0311	0.1445	0.1525
1985	0.1262	0.1630	0.1805	0.0549	0.1489	0.1709
1986	0.0897	0.1593	0.2021	0.0426	0.1596	0.1443
1987	0.0974	0.1634	0.2096	0.0415	0.1550	0.1605
1988	0.0938	0.1604	0.2105	0.0377	0.1500	0.2081
1989	0.1004	0.1499	0.2065	0.0600	0.1552	0.1753
1990	0.1073	0.1696	0.2608	0.0788	0.1773	0.2295
1991	0.0929	0.1773	0.1857	0.0630	0.1728	0.1690
1992	0.1222	0.1672	0.1826	0.0733	0.1665	0.1793
1993	0.1236	0.1431	0.1579	0.0687	0.1459	0.1534
1994	0.1257	0.1430	0.2089	0.0709	0.1418	0.1910
1995	0.1142	0.1747	0.2165	0.0766	0.1580	0.2438
1996	0.1173	0.1428	0.2059	0.0841	0.1539	0.1411
1997	0.0955	0.1940	0.2132	0.0724	0.2041	0.2035
Average	0.1156	0.1585	0.2219	0.0551	0.1592	0.1862

Figure 1

Empirical Density Functions of Good-Level Log Real Exchange Rate in 1985

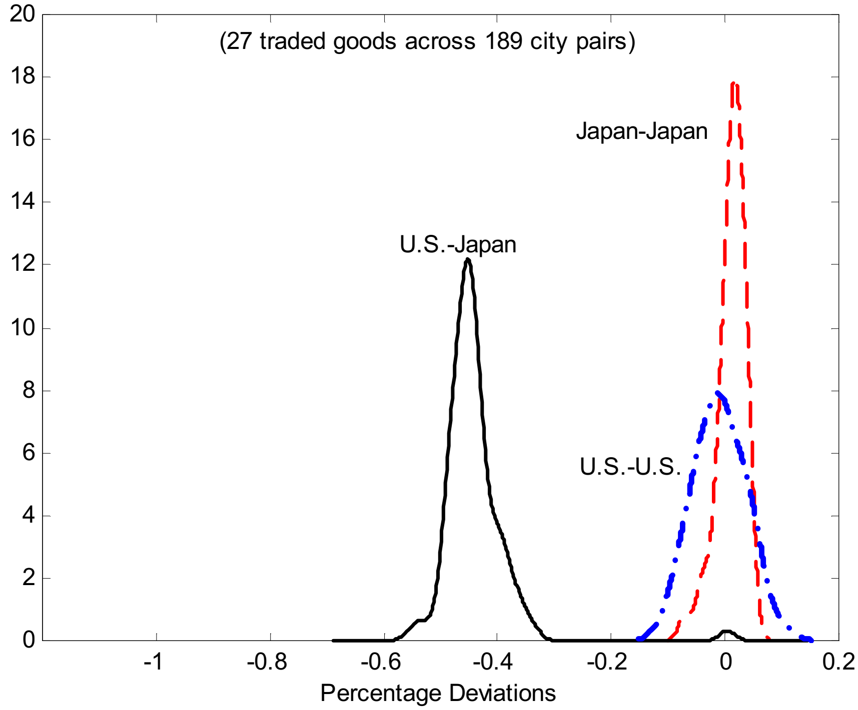

Figure 2

Empirical Density Functions of Good-Level Log Real Exchange Rate in 1990

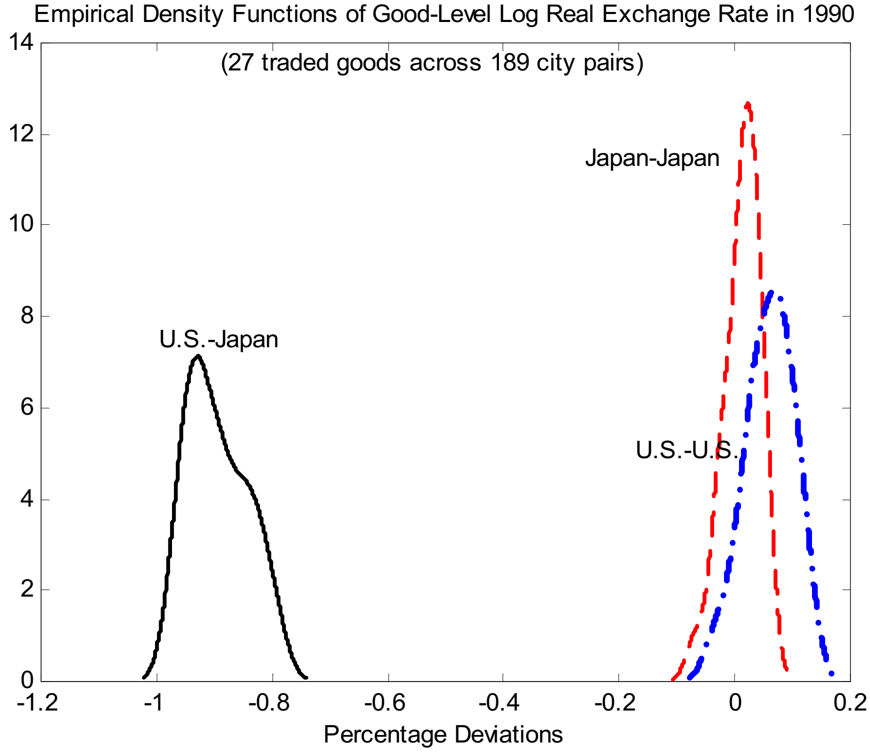

Figure 3: Average of Good-Level Log Real Exchange Rates

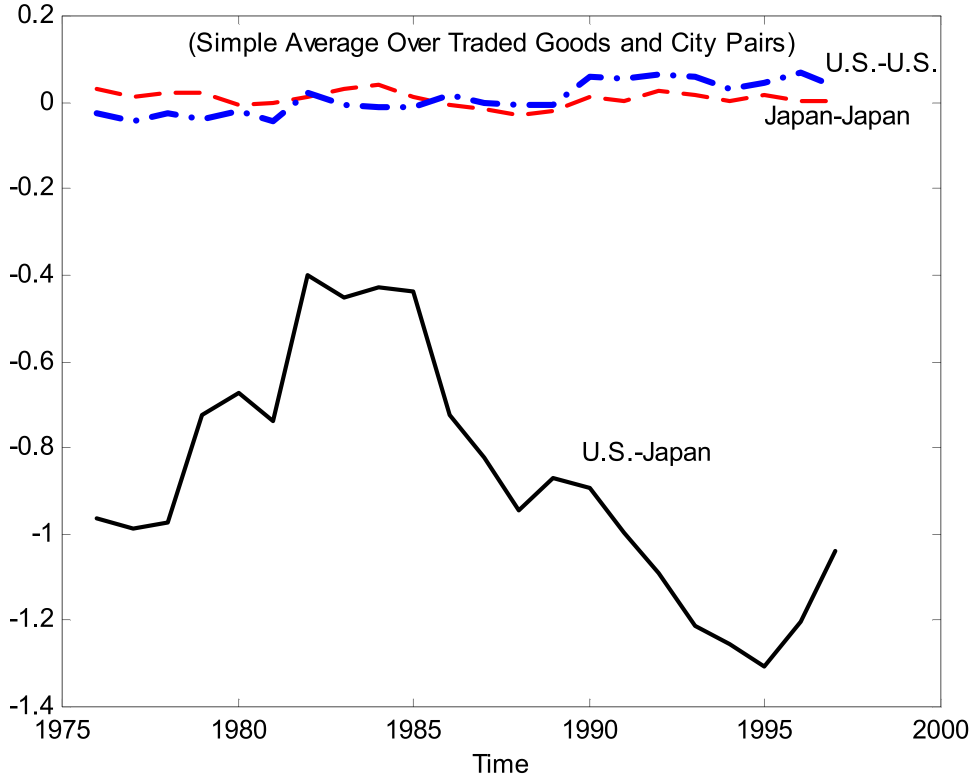

Figure 4: Price Dispersion Averaged over Relevant City Pairs