Border, Border, Wide and Far,
How We Wonder What You Are

David C. Parsley
Owen Graduate School of Management, Vanderbilt University

and

Shang-Jin Wei
Harvard University, NBER

June 1999

Abstract

This paper exploits a three dimensional panel data set of prices on 25 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 an equally-weighted average of goods-level real exchange rates tracks the nominal exchange rate well, suggesting strong evidence of sticky prices.

We turn next to economic explanations for the dynamics of this so-called “Border” effect. Focusing on dispersion in prices between city pairs, we confirm previous findings that crossing national borders adds significantly the price dispersion. Using our point estimates crossing the U.S.-Japan “Border” is equivalent to adding 13 billion miles to the cross-country volatility of relative prices. We make a direct and explicit inference on the influence of shipping costs, distance, exchange rate and relative wage variability on the “Border” effect. In our calculations, the “Border” effect diappears after controlling for these additional variables.

1. Introduction

International markets have been more segmented than intra-national markets for at least as long as the Swiss have eaten cheese. And it continues to be so today. So that is not news. The extent of segmentation, when quantitatively documented, appears striking nonetheless. 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 striking 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 province 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 the other 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 the real exchange rate based on a value-weighted price index (e.g. CPI) may be far away from zero, the equally-weighted average of individual goods-level real exchange rates was actually fairly close to one. 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, among the European countries in their sample, the exchange rates were (nearly) fixed, and the physical distance and policy-induced trade barriers were low. The "Border" effect could be more significant between country pairs that do not have the same environment.

In this paper, we exploit a three-dimensional panel data set of prices for 25 commodity-level goods (e.g., one box of facial tissue, 175 count), in 88 quarters (1975:1-1996:4), in 96 cities in Japan and the United States. Each of the 25 goods is selected so that we can match the definition of the good reasonably well between the two countries[1].

We aim for several objectives. First, we examine the behavior of the average goods-level real exchange rate for the U.S. and Japan – the counterpart to the measure examined in the Crucini, et al. (1999) 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

---

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

distance – the counterpart to the measure studied by Engel and Rogers (1996).[2] Our innovation is on understanding its dynamics. We ask two related questions. First, is there any evidence that the “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 goods-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 equally-weighted average of goods-level real exchange rates tracks closely the nominal exchange rate. 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 goods-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, shipping costs, distance, exchange rate and relative wage variability collectively explain a substantial portion of the “Border” effect.

2. Data

Appendix Table 1 provides a brief description of the goods and their correspondence between Japan and the United States. In some cases the correspondence is remarkable given there was no single price reporting authority. In other cases we see that there is a basic, if not exact coherence in the cross-country 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. There is still a slight time mismatch however. The U.S. data are generally sampled seven to ten days

---

[2] 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.

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-five 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 Table 2 lists the U.S. and Japanese cities we include.

3. Statistical results

3.1 The mean of goods-level real exchange rates

Crucini et al. (1999) note that even though value weighted average deviations from LOP over goods can be big 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 and one from the United States. The benchmark cities are Tokushima for Japan, and Louisville for the United States. These are ‘centrally located’ and ‘representative’ cities of their respective countries. This produces a sample of 189 city pairs in total.

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 4700 goods-level real exchange rates, each period, or over 400,000 time-series observations. We obtain a further scale reduction by conducting our analysis at the annual frequency. Annual observations have the additional benefit of allowing U.S. to examine more non-price economic determinants than otherwise possible since these often are observed less frequently.

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 goods-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 goods-level log exchange rates: $r(ij,k,t)$ over all city-pairs within the U.S., $r(ij,k,t)$ over all city-pairs within Japan, and $r(ij,k,t)$ over all city-pairs where city *i* is in the U.S. and city *j* in Japan.

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 1980. 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 somewhat less disperse than those in the U.S. This is possibly due to the relative sizes of the two countries. 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 most Japanese prices in 1980 were higher than in the U.S.

In Figure 2, we repeat the exercise for 1990. The comparison with Figure 1 is striking. The between country distribution no longer overlaps with either of the two within country distributions. Japanese dollar denominated prices 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 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, $r(us,t)$, as the average of $r(ij,k,t)$ over all goods and all city pairs within the U.S. We can define $r(japan,t)$ and $r(us-japan,t)$ in an analogous way.

The left panel of Table 1 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 U.S. and Japan, are fairly close to zero. In fact they vary within plus/minus ten percent in each of the twenty-two years in our sample. In comparison, the average log real exchange rate between U.S. and Japan makes much larger gyrations, from over 15% in 1982 to –75% in 1995.

We cannot fail to notice that the time series path of the average log real exchange rate between 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.[3] In accord with our expecations, the nominal exchange rate explains much of the variation– the adjusted R^2 of the equation is .47, and the coefficient on the nominal exchange rate is estimated at 0.78 with a standard error of 0.18. In fact, it is not possible to reject the hypothesis that the coefficient estimate is 1.0 even at the 15% level. 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).

This also suggests a potential reconciliation with the Crucini et al. (1999) finding. Namely, within their sample of European countries in 1985 exchange rates were tied by the European Monetary system – hence exchange rate movements were bounded. This is consistent with our intra-national samples (Japan-Japan, and U.S-U.S.) evidence where nominal exchange rates are fixed. Incidentally, there is also the possibility, partly borne out by evidence presented in Figure 3, that 1985 was a special year. The deviations from

---

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

the law of one price in 1985 (or any year during 1980-86) were smaller than either earlier or later years.

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

$X(us,t) = \frac{1}{KN}\sum_{ij,k} X(ij,k,t)$, 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 Table 1, and in Figure 4, 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 50%-75%.

3.2 Declining border?

We would like to understand if 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 dynamics 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. Because of this, we think that the best specification to explore these restrictions is to look at how the standard deviation of price differentials (or other measure of range of possible price differentials) may be linked to factors related to the cost of arbitrage. This is the approach that Engel and Rogers (1996) adopted.

More formally, we adopt a measure of 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-five goods and four quarters in a given year, in the in order to study the evolution of the "Border" effect. Let the change in the real exchange rate relative to benchmark city j be

$Q(ij,k,t) \equiv \Delta \ln P(i,k,t) - \Delta \ln P(j,k,t)$, where ij represents a city pair.

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.

Later, for robustness checks, we also adopt an alternative measure of dispersion across cities – the interquartile range (75th-25th fractiles of the distribution). In our regression analysis we stack the twenty-two annual observations of the variability of q, and pool all city pairs into each regression.

3.3 Results and Inferences

Table 2 presents summary statistics on our dependent variable. We summarize the data by averaging over all intra-national city pairs for the 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. Transport costs, costs of non-mobile factors (labor), or exchange rate variability are all likely candidates. 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 varies with the distance. Introducing sticky goods prices explicitly demonstrates the impact of exchange rate variability on real exchange rates measured between countries. A third important difference between intra-national city pairs 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 within country relative wages are less variable than that for cross-border city pairs. Their econometric evidence generally supports this hypothesis. Empirically however they find that inclusion of relative wage variability has little impact on the "Border" effect. We wish to check whether this result holds between the United States and Japan.

We begin our investigation of this "Border" effect with a benchmark regression posited by Engel and Rogers (1996). We estimate

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

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.[4] This regression involves 189 city pairs (using Tokushima and Louisville as the benchmark cities) each with twenty-two time periods.

The first column of Table 3 reports results from this regression. We confirm that price dispersion increases with distance and that the "Border" effect is economically and statistically significant. Both estimates are of the hypothesized sign. Compared to Engel and Rogers, both effects are somewhat larger for our data. In fact, Engel and Rogers calculate that the "Border" adds as much as 75,000 miles does to the cross-country volatility. In our case, the number is roughly 13 billion miles.[5] Of course, Japan and the U.S. are further apart (the actual distance between Osaka and Houston is 6,891 miles) than between Canada and the U.S. 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, this is not the whole story. Whereas distance accounts for roughly thirty-two percent of the variability of relative prices between the U.S. and Canada, in our case distance contributes less than two percent of the variability of relative prices between the U.S. and Japan. Moreover, 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. In the present study, the "Border" effect contributes roughly forty percent to the standard deviation. The question remains whether this effect declines over time and is influenced by identifiable economic factors.

In column 2 we add three regressors: a time trend, and interaction terms between the time trend and distance and the border dummy variable. Of these three new variables, only the trend/border interaction variable is significant. Its negative sign suggests that the "Border" effect is declining over time, albeit slowly. The coefficient on the border dummy now captures the "Border" effect at the beginning of our sample, 1976. We re-run the regression, dropping the insignificant interaction terms. These results, which are presented in the third column in the table, leave our conclusions unaffected.

---

[4] The latitude and longitude information is available on the world wide web 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.

The next two specifications examine economic factors differing by country; hence they represent potential economic explanations for a “Border” 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. The Engel and Rogers paper (1996) captures the variability of the wage differential explicitly, but infers the effect of exchange rate volatility only indirectly. We present summary statistics on these variables in Table 4, and graph them in Figure 6.

One effect we should expect to be relevant for cross-country price volatility is 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, our measure is the difference between c.i.f. and f.o.b. values of U.S. imports as a percentage of the total f.o.b. value.[6] For the domestic 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 admittedly arbitrary, but plausible. We note that assigning a value of zero would overly penalize international city pairs (and hence might explain too much of the “Border” effect).

Another of the potential explanations alluded to above is exchange rate variability. Figure 3 above, and the regression of the average real goods-level exchange rate on the nominal exchange rate suggests sticky prices. Thus we include the standard deviation of daily changes in the (log of) the nominal exchange rate as an additional regressor.[7]

We add both unit shipping costs and exchange rate volatility to the regression. This perturbation to our basic specification has a major effect. Exchange rate volatility has a positive effect on cross-country price variability as expected, and the coefficient estimate is highly statistically significant. According to the table an increase in daily exchange rate volatility of one-percent, is associated with an 18 percent increase in annual cross-country

---

[5] This is calculated as $\exp(\beta_2/\beta_1)$. For comparison, the distance from the earth to the Sun is roughly 94 million miles. This large distance could be a consequence of entering distance in log form.

[6] We obtained the U.S. import data from the March 1999 IFS CD (lines 71 and 71v).

[7] The source for the daily exchange rate data is the St. Louis Federal Reserve web site.

price volatility. Adding this variable also has a positive impact on the estimate of shipping costs. According to this specification, reducing shipping costs reduces cross-country price variability by nearly three times as much as the reduction in shipping costs. The stunning result in the table however is that now the "Border" effect has reversed sign. In other words, these two variables have more than explained the "Border" effect!

The next variable we want to control for is variability of the differences in wages. Here we are trying to get at the non-traded component of goods prices. This variable is defined, analogous to the dependent variable, as the standard deviation of the first difference of the log of the Japanese wage minus the first difference of the log of the U.S. wage.[8] The coefficient on the variability of the wage difference are positive and statistically significant. Moreover, adding this variable to our equation further reduces the "Border" effect, and raises the impact of shipping costs slightly.

In sum, the three variables collectively have not only reduced the "Border" effect, but also have reversed the sign of the effect. In other words, we have more than explained the "Border" effect. It is important to know if any of this is robust against more scrutiny.

3.4 Extensions

In the regressions so far a number of the independent variables are stacked for each city pair included in the regression. It is possible that this procedure would understate the standard errors in our regressions. Moreover, more efficient parameter estimates are attainable by explicitly considering the cross-equation error correlations. In an effort to focus more precisely on the determinants of the "Border" and to see if they are truly significant once we allow for possible correlation of the error terms across city pairs and across time, we implement systems estimation using the seemingly unrelated regressions method.

In our sample we have 95 international city pairs and only twenty-two time series observations. Hence it is not possible to obtain parameter estimates since the variance-covariance matrix will be singular. Additionally, missing values will affect the estimation

---

[8] The source for this data is also the March 1999 IFS CD (lines 65 for Japan, and 65ey for the U.S.).

of the cross-equation correlations. We address these issues by selecting the first ten international city pairs containing no missing values using each of the Tokushima and Louisville benchmarks. Thus the resulting system has twenty equations. 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.

We proceed sequentially as before by beginning with a benchmark specification, given in column 1 of Table 5. We first estimate the effects of shipping costs. The point estimate is slightly lower than in the OLS regressions but is still positive and statistically significant. Exchange rate volatility enters the equation with virtually the same point estimate as in the OLS equations. In the final specification we add wage variability. Again, the coefficient estimate on wage variability is virtually unchanged while its statistical significance increases tremendously. However now, the economic impact of shipping costs on real exchange rate variability is reduced.

We try several extensions to test the robustness of our results. First, we repeat the analysis using the interquartile range (75th-25th percentiles of the distribution) of our real exchange rate variability measure. These results are presented in Table 6

A second robustness test was 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, and partly because we wanted to select larger, coastal cities for the analysis. These results are reported in Table 7.

The basic findings reported earlier are qualitatively valid in these alternative definitions of the dependent variable, alternative estimation method, and alternative choice of the benchmark cities. In particular, the border effect adds significantly to price dispersion, and shipping costs, exchange rate volatility and relative wage volatility are all related to the border effect.

4. Concluding remarks

This paper exploits a three dimensional panel data set of prices on 25 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 an equally-weighted average of goods-level real exchange rates tracks the nominal exchange rate well, suggesting strong evidence of sticky prices.

We turn next to economic explanations for the dynamics of this so-called “Border” effect. Focusing on dispersion in prices between city pairs, we confirm previous findings that crossing national borders adds significantly the price dispersion. Using our point estimates crossing the U.S.-Japan “Border” is equivalent to adding 13 billion miles to the cross-country volatility. We make a direct and explicit inference on the influence of shipping costs, distance, exchange rate and relative wage variability on the “Border” effect. In our calculations, the “Border” effect diappears after controlling for these additional variables.

REFERENCES

Balassa, Bela, 1964, "The Purchasing-Power parity Doctrine: A Reappraisal," **Journal of Political Economy**, pp. 584-96.

Crucini, Mario, Chris Telmer, and Marios Zachariadis, 1999, “Dispersion in International Prices”, University of Pennsylvania working paper.

Engel, Charles, 1993, "Real Exchange Rates and Relative Prices: An Empirical Investigation," **Journal of Monetary Economics**, 32:35-50.

Engel, Charles, and John Rogers, 1996, "How Wide is the Border?" **American Economic Review,** 86:5, pp. 1112-1125.

Frankel, Jeffrey, and Andrew Rose, 1996, “A Panel Project on Purchasing Power Parity: Mean Reversion within and between Countries”, **Journal of International Economics,** 40:1-2, pp. 209-224.

Froot, Kenneth, and Kenneth Rogoff, 1995, "Perspectives on PPP and Long-Run Real Exchange Rates," Handbook of International Economics**,** Volume 3. New York, Elsevier, pp. 1647-1688.

Helliwell, John, 1998, How Much Do National Borders Matter? DC: Brookings Institution Press.

O’Connell, Paul, and Shang-Jin Wei, 1997, “The Bigger They Are, The Harder They Fall: How Price Differences Across U.S. Cities Are Arbitraged?” NBER Working Paper 6089, July.

Rogers, John, and Michael Jenkins, 1994, “Haircuts of Hysteresis? Sources of Movements in Real Exchange Rates, **Journal of International Economics**, 38(3-4), May 1995, pages 339-60

Samuelson, Paul, 1954, "Theoretical Notes on Trade Problem," **Review of Economics and Statistics,** 46: 145-164.

Statistics Bureau, Management and Coordination Agency, Government of Japan, Annual Report on the Retail Price Survey, various issues.

The American Practical Navigator, 1977, Volume 1: Washington: U.S. Defense Mapping Agency.

Wei, Shang-jin, 1996, “Intra-national versus International Trade: How Stubborn Are Nations in Global Integration?” NBER Working Paper 5531, April.

Table 1. Measures of Intra-national and International Price Deviations
(Simple Average over Traded Goods and City Pairs)

	Average of Goods-Level Log Real Exchange Rates			Average of Goods-Level Absolute Percentage Deviations		
Year	***Japan***	***U.S.***	***U.S –Japan***	***Japan***	***U.S.***	***U.S. -Japan***
1976	0.031	-0.031	-0.355	0.1230	0.1474	0.6990
1977	0.013	-0.049	-0.382	0.1111	0.1414	0.6915
1978	0.023	-0.025	-0.366	0.1195	0.1284	0.5921
1979	0.023	-0.041	-0.109	0.1105	0.1100	0.5222
1980	0.005	-0.018	-0.089	0.1072	0.1164	0.6542
1981	0.002	-0.041	-0.147	0.1036	0.1362	0.6616
1982	0.014	0.023	0.172	0.0949	0.1234	0.5902
1983	0.030	-0.008	0.119	0.1055	0.1233	0.5903
1984	0.043	-0.010	0.143	0.1047	0.1326	0.5871
1985	0.011	-0.008	0.129	0.1300	0.1447	0.6264
1986	0.005	0.015	-0.153	0.1262	0.1274	0.5567
1987	0.018	-0.004	-0.252	0.1266	0.1474	0.5601
1988	0.030	-0.007	-0.376	0.1234	0.1417	0.5867
1989	0.022	-0.003	-0.293	0.1260	0.1378	0.5750
1990	0.013	0.059	-0.318	0.1131	0.1500	0.6067
1991	0.001	0.054	-0.414	0.1065	0.1740	0.6130
1992	0.026	0.061	-0.555	0.1130	0.1584	0.6583
1993	0.019	0.058	-0.636	0.1065	0.1553	0.7182
1994	0.003	0.030	-0.676	0.1055	0.1422	0.7302
1995	0.016	0.044	-0.727	0.1115	0.1512	0.7461
1996	0.005	0.066	-0.627	0.1153	0.1524	0.6855
1997	0.005	0.036	-0.462	0.1135	0.1554	0.5753
Average	0.009	0.009	-0.290	0.1135	0.1408	0.6285

Table 2. 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	0.0924	0.2012	0.2034	0.0423	0.2249	0.2086
1977	0.1076	0.1628	0.2112	0.0388	0.1710	0.1860
1978	0.1008	0.1553	0.3397	0.0354	0.1520	0.2342
1979	0.1072	0.1379	0.2608	0.0296	0.1337	0.1532
1980	0.0976	0.1355	0.2524	0.0501	0.1343	0.2051
1981	0.0887	0.1208	0.2360	0.0337	0.1208	0.1693
1982	0.0894	0.1161	0.2147	0.0220	0.1060	0.1539
1983	0.0836	0.1265	0.1842	0.0256	0.1209	0.1669
1984	0.0826	0.1283	0.1512	0.0246	0.1257	0.1213
1985	0.0805	0.1381	0.1538	0.0248	0.1458	0.1655
1986	0.0734	0.1427	0.1924	0.0180	0.1538	0.1371
1987	0.0884	0.1262	0.1987	0.0238	0.1374	0.1436
1988	0.0816	0.1398	0.1970	0.0292	0.1289	0.1881
1989	0.0835	0.1445	0.1857	0.0372	0.1427	0.1741
1990	0.0726	0.1350	0.2411	0.0410	0.1472	0.2052
1991	0.0694	0.1416	0.1815	0.0315	0.1420	0.1496
1992	0.0877	0.1566	0.1643	0.0318	0.1564	0.1750
1993	0.0867	0.1543	0.1597	0.0301	0.1564	0.1635
1994	0.0970	0.1507	0.2086	0.0411	0.1611	0.1841
1995	0.1015	0.1452	0.2293	0.0545	0.1460	0.2190
1996	0.0950	0.1207	0.1902	0.0553	0.1312	0.1378
1997	0.0888	0.1979	0.2205	0.0555	0.1962	0.1792
Average	0.0889	0.1444	0.2080	0.0349	0.1433	0.1720

Table 3. Explaining the Border Effect
Dependent Variable: Standard Deviation of Log Price Differential
Tokushima-Louisville Benchmark cities

	Equation 1	***Equation 2***	***Equation 3***	***Equation 4***	***Equation 5***
Log Distance	0.0035	0.0029	0.0036	0.0036	0.0038
	(0.0017)	(0.0009)	(0.0017)	(0.0015)	(0.0014)
Border	0.0815	0.1084	0.1092	-0.1046	-0.1733
	(0.0055)	(0.0055)	(0.0055)	(0.0137)	(0.0129)
Trend		-0.0006			
		(0.0009)			
Trend*Log Distance		0.0001			
		(0.0002)			
Trend* Border		-0.0024	-0.0025	-0.0006	0.0031
		(0.0005)	(0.0001)	(0.0003)	(0.0034)
Shipping Costs				2.7163	3.2061
				(0.2707)	(0.2489)
Nominal Exchange Rate Variability				18.0790	17.8740
				(0.5645)	(0.5185)
Wage Variability		1.7441			
		(0.0663)			
Adjusted R^2	0.612	0.643	0.643	0.720	0.764
N. of observations	3820	3820	3820	3820	3820

Table 4. Shipping Costs, Exchange Rate Variability, and Wage Variability

Year	***Shipping Costs***	***Exchange Rate Variability***	***Wage Variability***
1976	0.0633	0.0021	0.0536
1977	0.0593	0.0036	0.0545
1978	0.0567	0.0082	0.0513
1979	0.0568	0.0066	0.0419
1980	0.0493	0.0074	0.0635
1981	0.0474	0.0070	0.0480
1982	0.0448	0.0076	0.0251
1983	0.0458	0.0051	0.0196
1984	0.0474	0.0043	0.0207
1985	0.0474	0.0058	0.0184
1986	0.0531	0.0071	0.0076
1987	0.0447	0.0065	0.0089
1988	0.0409	0.0064	0.0183
1989	0.0417	0.0069	0.0156
1990	0.0438	0.0066	0.0251
1991	0.0408	0.0061	0.0224
1992	0.0400	0.0053	0.0169
1993	0.0395	0.0070	0.0182
1994	0.0383	0.0056	0.0142
1995	0.0368	0.0088	0.0160
1996	0.0336	0.0048	0.0222
1997	0.0327	0.0072	0.0195
Average	0.0456	0.0062	0.0273

- Shipping Costs are defined as the percentage difference between the value of U.S. imports on a c.i.f. basis and an f.o.b. basis
- Exchange Rate Variability is defined as the standard deviation over the year of ln(xt)-ln(xt-1)
- Wage Variability is defined as the standard deviation of dln(wagej)-dln(wageus)

Table 5. Robustness Checks: SUR Estimation
Tokushima-Louisville Benchmark

	Equation 1	***Equation 2***	***Equation 3***
Shipping Costs	2.2932	3.2870	0.8198
	(0.2555)	(0.1146)	(0.0596)
Nominal Exchange Rate Variability		17.7292	16.9050
		(0.6079)	(0.2419)
Wage Variability			1.8105
			(0.0282)
Average adjusted R^2	.052	.272	.493
N. of equations	20		
N. of observations	22		

Note: equation specific intercepts not reported.

Table 6. Explaining the Border Effect
Dependent Variable: Interquartile Range of Log Price Differential
Tokushima-Louisville Benchmark cities

	Equation 1	***Equation 2***	***Equation 3***	***Equation 4***	***Equation 5***
Log Distance	-0.0034	0.0011	-0.0003	-0.0003	-0.0002
	(0.0016)	(0.0023)	(0.0016)	(0.0015)	(0.0015)
Border	0.0825	0.0912	0.0924	-0.0221	-0.0766
	(0.0050)	(0.0076)	(0.0052)	(0.0144)	(0.0140)
Trend		0.0010			
		(0.0009)			
Trend*Log Distance		-0.0012			
		(0.0002)			
Trend* Border		-0.0006	-0.0007	0.0011	0.0041
		(0.0005)	(0.0001)	(0.0003)	(0.0037)
Shipping Costs				1.9465	2.1927
				(0.2842)	(0.2714)
Nominal Exchange Rate Variability				5.7914	5.6283
				(0.5926)	(0.5655)
Wage Variability					1.3859
					(0.0723)
Adjusted R^2	0.711	0.713	0.713	0.721	0.746
N. of observations	3820	3820	3820	3820	3820

Table 7. Explaining the Border Effect
Dependent Variable: Standard Deviation of Changes in Real Goods-Level Exchange rates
Osaka-Houston Benchmark Cities

	Equation 1	***Equation 2***	***Equation 3***	***Equation 4***	***Equation 5***
Log Distance	0.0039	0.0029	0.0041	0.0040	0.0042
	(0.0017)	(0.0025)	(0.0017)	(0.0015)	(0.0014)
Border	0.0815	0.1117	0.1098	-0.0905	-0.1578
	(0.0055)	(0.0081)	(0.0055)	(0.0138)	(0.0130)
Trend		-0.0008			
		(0.0009)			
Trend*Log Distance		0.0001			
		(0.0002)			
Trend* Border		-0.0027	-0.0026	-0.0011	0.0026
		(0.0005)	(0.0001)	(0.0003)	(0.0003)
Shipping Costs				2.3663	2.6723
				(0.2722)	(0.2515)
Nominal Exchange Rate Variability				18.4121	18.2113
				(0.5681)	(0.5244)
Wage Variability					1.7060
					(0.0670)
Adjusted R^2	0.613	0.644	0.644	0.722	0.764
N. of observations	3818	3818	3818	3818	3818

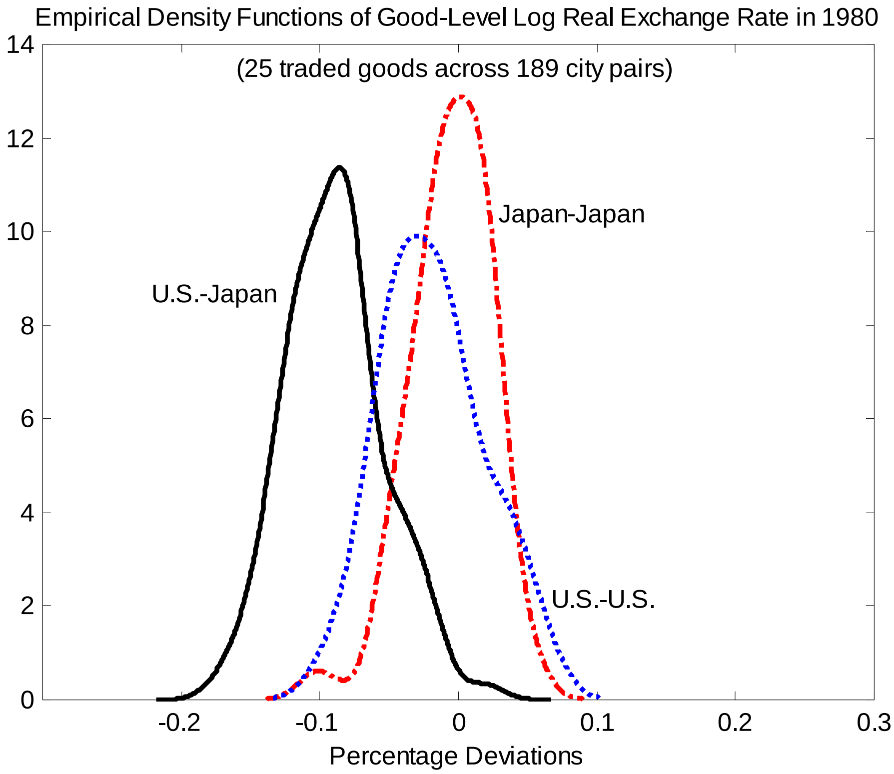
Table A1. Correspondence of Japanese and United States Goods

Good	***Japanese Prices***	***U.S. Prices***
1	Canned tuna fish, in oil, #4 can, 80kg	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

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

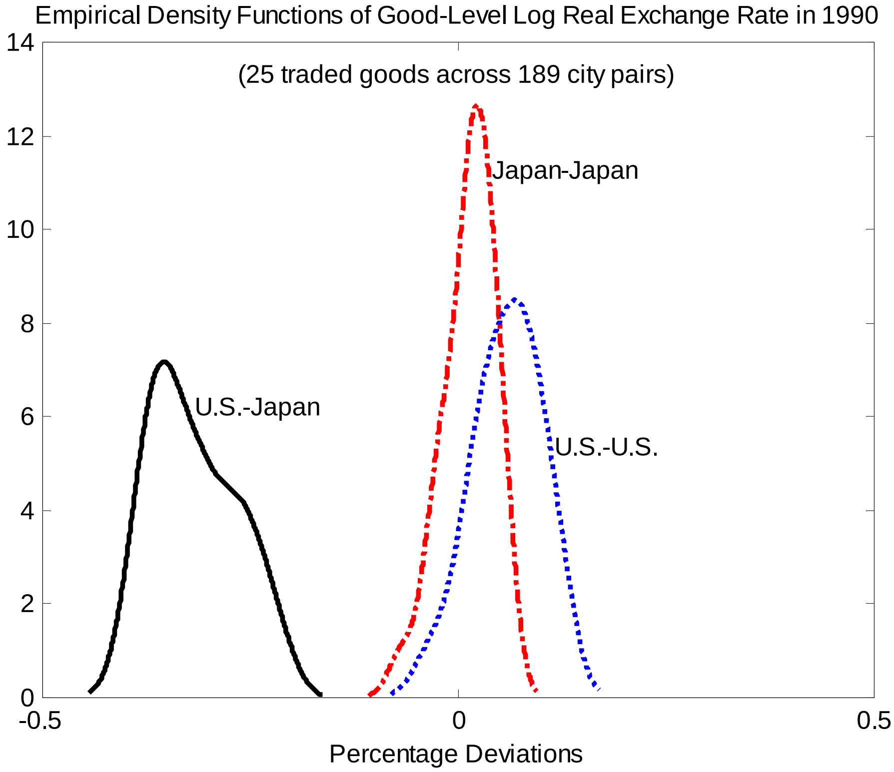
Figure 1



Empirical Density Functions of Good-Level Log Real Exchange Rate in 1980

(25 traded goods across 189 city pairs)

Figure 2



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

(25 traded goods across 189 city pairs)

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

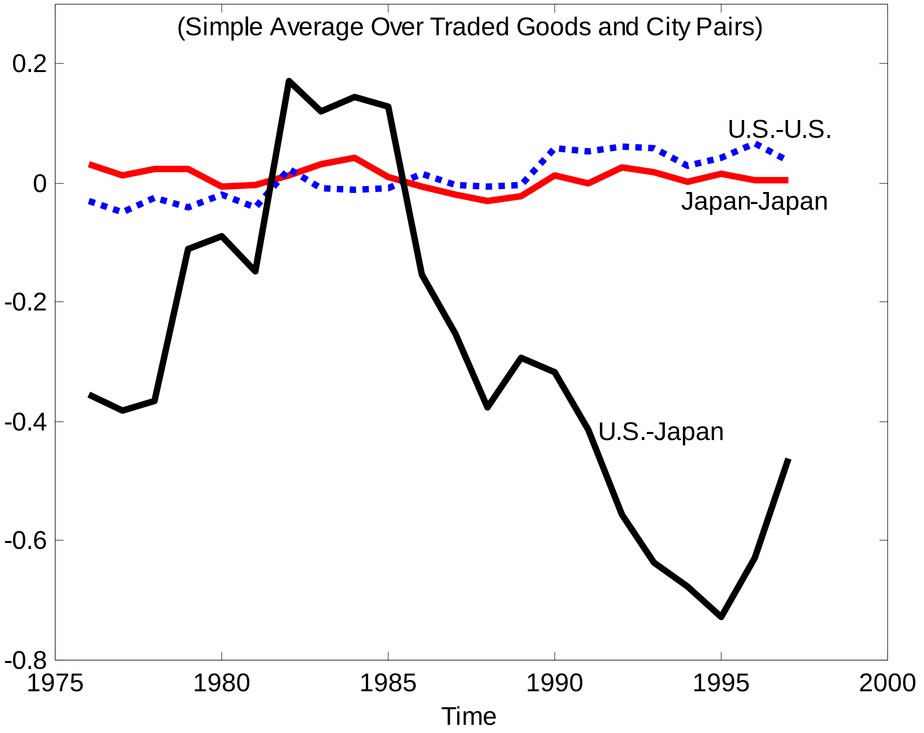


Figure 4: Average of Good-Level Absolute Percentage Deviations

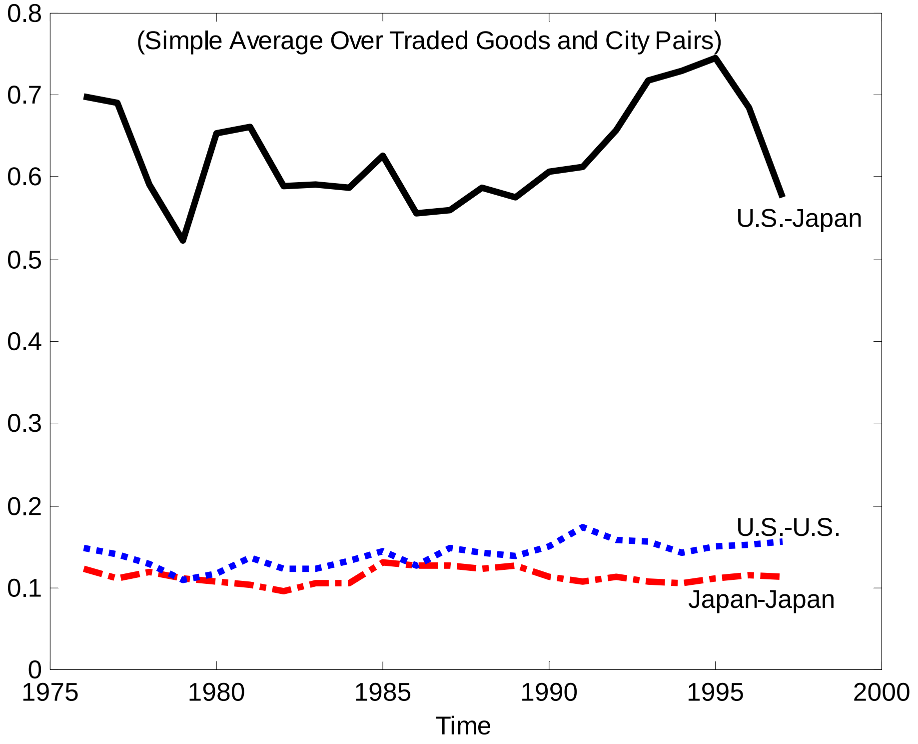


Figure 5 Price Dispersion Averaged over Relevant City Pairs

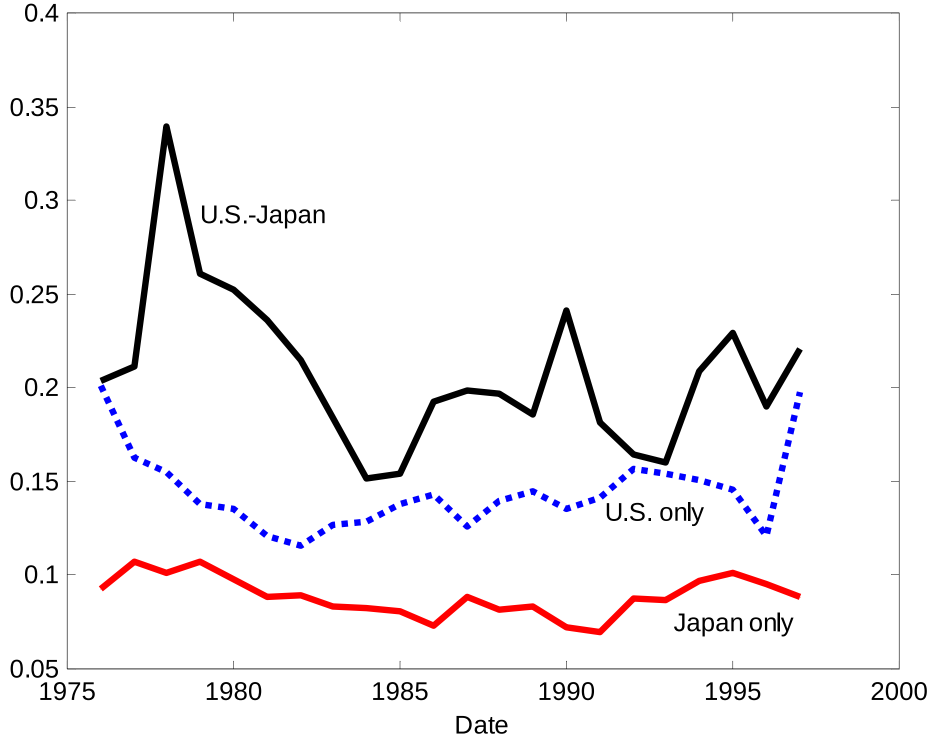


Figure 6a Unit Shipping Costs

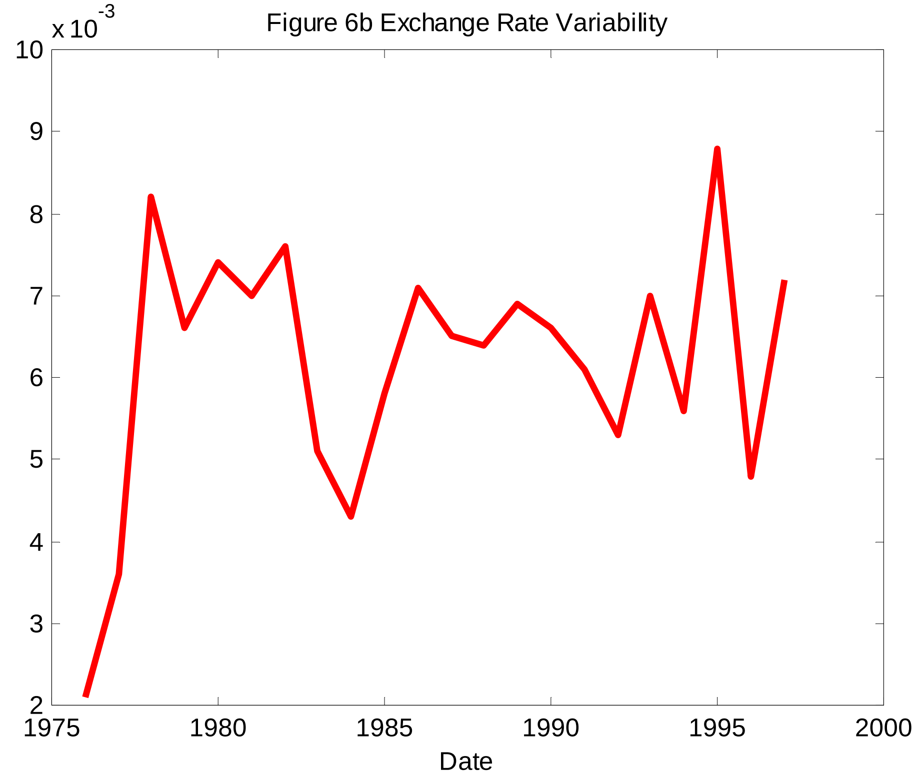


Figure 6b Exchange Rate Variability

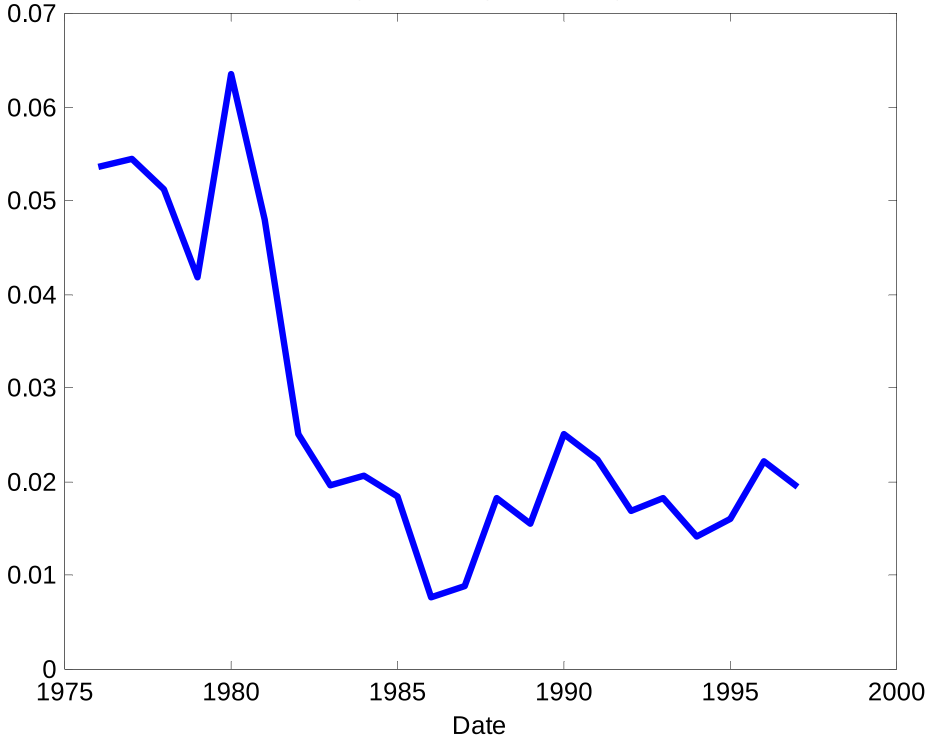


Figure 6c Wage Variability