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Abstract: Using consumer price indexes from cities in the U.S., Canada and Mexico, we estimate the "border effect" on U.S.-Mexican relative prices and find that it is nearly an order of magnitude larger than for U.S.-Canadian prices. However, during a very stable sub-period in Mexico (May 1988 to November 1994), the "width" of the U.S.-Mexican border *falls dramatically* and becomes approximately equal to the U.S.-Canadian border. We then show that when consideration is limited to cities lying geographically very close to the U.S.-Mexican border -- San Diego, Los Angeles, Houston, Dallas, Tijuana, Mexicali, Juarez, and Matamoros -- the border width falls compared to that estimated with the full sample of U.S. and Mexican cities, but *falls only very slightly*. We also present evidence that the border effect in U.S.-Mexican prices is not primarily due to the border effect in U.S.-Mexican wages. Finally, using the prices of 276 highly dis-aggregated goods and services, we estimate the variability of relative prices of different items within Mexican cities. This measure of relative price variability declines during the stable peso sub-period, but by less than the decline in nominal and real (i.e., CPI-based) exchange rate variability. Our results are strong evidence of a "nominal border effect" in relative prices within NAFTA, but also indicate that real side influences are important.

JEL classification: F3, F4
Keywords: relative prices, exchange rates, purchasing power parity

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I. Introduction

Evidence of an increased trend towards globalization abounds. Worldwide, exports as a percent of GDP have grown dramatically since 1950. Regional trade agreements have significantly reduced legal impediments to cross-border flows of goods and services, and in Europe especially, factors of production. Boosted by the North American Free Trade Agreement (NAFTA), U.S. trade rose from 15% of GDP in the early 1990s to nearly 20% in 2000. Recent financial crises in Mexico and Asia underscore the speed with which shocks can be transmitted across national asset markets.

Recent research has shown that, in spite of the trend towards globalization, consumer prices are not nearly as equalized across countries as one would expect in a world of highly integrated goods markets. In perfectly-integrated markets, prices of similar goods ought to be equalized, when those prices are denominated in a common currency. If the price in one location rose substantially above that in another, market forces would tend to move prices back towards equality. However, empirical studies uniformly find large deviations from such a benchmark.

The extent to which prices of similar goods fail to equalize *across countries* has recently been quantified against a baseline of failures across regions *within countries*. Following this metric, Engel and Rogers (1996) examine the variability of the relative price of similar goods, using consumer price data from 23 cities in the U.S. and Canada. They show that relative price variability is positively and significantly related to the distance between cities. But, accounting for the effect of distance, relative price variability is strikingly larger for cities that lie across the border than for cities that lie within either country. As a pedagogical device, Engel and Rogers (1996) dub this the “width of the border”.

There are several possible explanations for this large border effect in relative prices. These include tariffs and non-tariff barriers to trade, the presence of non-tradeable goods and services embodied in final consumer goods prices, and relatively less homogenous labor markets or distribution networks across countries than within countries. Engel and Rogers (2000) label the contribution of such factors (toward explaining the large border effect) the "real border effect". It is analogous to the border effects in the trade volume literature [e.g., McCallum (1995) and Helliwell (1996)]. An alternative explanation relies on nominal exchange rate variability with sticky final goods prices. The hypothesis is that prices in all cities are sticky when denominated in the local currency. When the nominal exchange rate fluctuates, so do relative goods prices for cross-border city pairs; relative price variability for within-country city pairs is unaffected, however. Engel and Rogers (2000) label this a "nominal border effect".[1]

In this paper we use consumer price data from Mexican cities, along with the U.S. and Canadian city price data as in Engel and Rogers (1996), to estimate border effects on relative prices within the NAFTA countries. Of particular interest is the "width" of the U.S.-Mexican border. The Mexican data provide a laboratory experiment of sorts, and with the data we attempt to draw sharp conclusions. We begin by showing that the border effect in U.S.-Mexican prices is nearly an order of magnitude larger than that for the (already-found-to-be-large) effect in U.S.-Canadian prices, over the full sample period 1980-1997.

There are several reasons to believe that the border effects involving Mexican prices ought to be larger than those in the U.S.-Canada data. Before NAFTA, trade between the U.S.

[1]They also cast doubt on whether the welfare costs of deviations from the law of one price due to this source are as large as the costs from the "real border effects".

and Canada was less restricted than trade between the U.S. and Mexico. The U.S.-Canada free-trade agreement preceded NAFTA by four years, and U.S.-Canadian automobile trade had been unrestricted for decades before that. Marketing and distribution networks in the U.S. are more similar to those in Canada than Mexico, perhaps because English is the primary language of both the U.S. and Canada. In addition, labor mobility is likely to be greater between the U.S. and Canada than between the U.S. and Mexico, illegal immigration aside.[2]

In order to shed light on the possible explanations for the relatively large border effect on U.S.-Mexican prices, we restrict our sample in two ways. First, we limit consideration to the sub-period May 1988- November 1994, known as *El Pacto*. During this period the peso/USD exchange rate was quite stable, with a standard deviation about equal to that of the CD/USD exchange rate. This sub-period also coincides with the advent of the important U.S.-Canada Free Trade Agreement. We find that during this sub-period the large border effect in U.S.-Mexican relative prices *falls dramatically* – to a level approximately equal to that of U.S.-Canadian prices (while the U.S.-Canada border actually widens somewhat).[3]

Second, we consider a limited set of cross-border cities, each lying geographically very close to one another and so subject to more similar regional supply or real demand shocks. This sample includes San Diego, Los Angeles, Houston and Dallas, and four true Mexican border towns: Tijuana, Mexicali, Juarez, and Matamoros. The Mexican cities all lie well within the "frontier zone" through which goods have been allowed to enter U.S. markets in a relatively

[2]Although U.S. and Mexican labor markets may be more integrated than is commonly believed [Hanson and Spilimbergo (1999) and Robertson (2000)].

[3]Consistent with this, Mendoza (2000) finds a large drop in the importance of nominal exchange rate volatility during this sub-period in variance decompositions of the peso-U.S. dollar real exchange rate.

unrestricted fashion for many years. When the sample includes only this sub-set of cities, the U.S.-Mexico border width falls compared to that estimated with the full sample of 28 U.S. and Mexican cities, but *falls only very slightly*. Thus the goods markets of the border cities do not appear to be much more integrated than the full sample of cities.

This result contrasts sharply with Robertson's (2000) findings that *labor* market integration with the U.S. is considerably higher for Mexican border towns than for towns in the interior of Mexico. It suggests that the border effect in U.S.-Mexican prices does not arise primarily from a relative lack of labor market integration. We confirm this directly: using Robertson's data on manufacturing wages in Mexican cities, along with manufacturing wage data from our U.S. locations, we show that the large border effect in prices remains even when we account for the presence of a large border effect in U.S.-Mexican wages.

Although we find the results described so far to be very informative, they are obtained using data that is limited in two important ways. First, the data are only for the aggregate consumer basket.[4] To see if our results using aggregate CPIs are being driven by movements in relative prices of different goods within cities, we examine price data on 276 very narrow categories such as "eggs" and "funeral services". Second, our data is in the form of price indexes rather than actual goods prices. Our measure of the deviation from PPP is the standard deviation of changes in the log of the relative price (index) across locations *j* and *k*. A finding that this measure of price variability is low indicates that percentage changes in the price of the market basket in location *k* relative to location *j* are small. Numerically this could occur because (1) the

[4]Engel and Rogers (1996), for example, use city-level data on fourteen broad sub-categories of goods, such as "food at home", "footwear", and "transportation."

"absolute law of one price holds", so that the difference in the price of all goods in locations *k* and *j* is close to zero; (2) the market basket price in one location *k* is roughly proportional to the price in location *j*, so that the relative price is nearly constant; or, (3) because price changes in cities *k* and *j* are themselves nearly constant. Were we to use price levels rather than price indexes, we would be able to distinguish between these three possibilities.

With these data limitations in mind, we interpret our results as strong evidence of the role of a nominal border effect in relative prices within NAFTA. An alternative interpretation is that, due to real-side events, the variability of the equilibrium peso-dollar real exchange rate fell during the sub-period 1988-94,[5] but was subsequently reversed after 1994. We put forth evidence to suggest that this explanation has some merit, but probably explains less of the border effect than the one that relies on sticky local-currency goods prices.

II. Data

We use consumer price data from 38 North American cities, beginning in January 1980.[6] As described in Table 1A, our "Full Sample" uses the monthly CPI over the period January 1980 to December 1997. We construct 678 bilateral relative prices from these 38 locations. In addition, as described in Table 1B, we examine a "Border Towns Sample", consisting of CPI data from eight U.S. and Mexican locations over the period January 1984 - December 1997. This allows construction of 28 bilateral relative prices. These data are semi-annual and begin in 1984,

[5] And that this drop in *real* exchange rate variability was simultaneously transmitted to lower *nominal* exchange rate variability, perhaps through monetary policy.

[6] Data sources are listed below Table 1.

due to the availability of data for San Diego. Figure 1 presents a time-line depicting the various sample periods used in the paper.[7]

Summary Statistics

Let *P(j,k)* be the log of the CPI in location *j* relative to that in location *k*. All prices are converted into U.S. dollars using a monthly average exchange rate before taking relative prices. We consider two-month changes in relative prices for the full sample (as did Engel and Rogers (1996)), because the price data for several U.S. cities is only reported every other month, and six-month changes for data used in the border towns sample because San Diego's CPI is only reported twice a year. We also examine the robustness of our results using 48-month differences.

We construct a measure of price volatility for each pair of locations, and base our analysis on the cross-section of the volatility measures. We calculate volatility as the standard deviation. Summary statistics are listed in Table 2 for all city pairs and for eight subsets of location pairs, namely those that are (1) both within the same country (labeled intra-national) (2) both in the United States (labeled US-US), (3) both within Canada, (4) both within Mexico, (5) one in one country and one in a foreign country (labeled inter-national), (6) one in the U.S., one in Canada (US-CA), (7) one in the U.S. and one in Mexico, and (8) one in Canada, one in Mexico.

Table 2A reports summary statistics for the full sample of cities, for two different time periods: 1980:1-97:12 and 1988:5-94:11. The first column reports the average standard

[7] The "full" sample period ends in 1997 because the U.S. Bureau of Labor Statistics revised the CPI in 1998 to take into account both demographic changes and new expenditure patterns. As described in Table 1A, the BLS stopped reporting price data for St. Louis and Pittsburgh after December 1997 and combined Washington, DC and Baltimore into one price area. In addition, the BLS started publishing CPI data for San Francisco and Philadelphia in even months only instead of on a monthly basis. Also at this time, Miami switched from having data reported in odd months to even months, while reporting for Dallas switched in the reverse manner. Notice that San Diego is the only city included in the "border towns" sample that is not also in the "full" sample of cities.

deviation of $\Delta P(j,k)$ across all cities *j* and *k*, for the longer sample period. The average volatility of cross-border pairs is 6.17, about 6 times larger than the average volatility for within-country pairs, 1.03. Within countries, price volatility is larger on average for Mexican city pairs, with a standard deviation of 1.44, than for U.S. (0.81) or Canadian city pairs (0.53). This may indicate that it is more costly to transport goods between Mexican cities than between cities within the U.S. or Canada. Looking across countries, average price volatility for the cross-section of U.S.-Mexican and Canadian-Mexican pairs is 4 ½ times larger than for the U.S.-Canadian pairs. Finally, as the third column shows, the nominal peso-dollar exchange rate is about 6 times more volatile than the US dollar-Canadian dollar rate for the full sample period. Note from the "Distance" column that the inter-national city pairs are on average slightly farther apart than are the intra-national pairs. We account for both nominal exchange rate volatility and distance in the regressions below.

During the truncated sample period from May 1988 - November 1994, known as *El Pacto*, the volatility of the Mexican exchange rate decreased substantially, in conjunction with an overall improvement in macroeconomic and financial conditions in Mexico. A comparison of columns 3 and 4 of Table 2A reveals the large drop in nominal exchange rate variability over the two sub-periods. Indeed during *El Pacto*, the peso-US dollar exchange rate is actually *less* volatile than the US dollar-Canadian dollar rate, with standard deviations of 1.37 versus 1.57, respectively! This stabilization of peso nominal exchange rates is mirrored by a drop in relative price variability for all pairs involving Mexican cities, as seen by comparing columns 1 and 2.

Table 2B presents a comparison of summary statistics for the full sample of cities and the border towns sub-sample, for the period 1984-1997. We might expect the variability of relative prices to be much smaller for cities that are on the U.S.-Mexico border, as these border cities are likely to be more integrated than the average pair of U.S. and Mexican cities. Robertson (2000) looks at the transmission of (aggregate) U.S. wage shocks to Mexico, and finds strong evidence that border cities are more highly integrated with the U.S. than are cities in the interior of Mexico. For goods prices, however, it does not appear that the border region is much more integrated, according to columns 1 and 3 of table 2B. The average standard deviation of relative price changes among the border towns is equal to 9.75, which is not much smaller than in the full sample of U.S. and Mexican cities (11.2). This is true despite the fact that the average distance between cities is only 700 miles in the sample of border cities and 1572 miles in the full sample of cities.

In appendix tables A-1 and A-2, we complete the comparison of summary statistics over different sub-periods and cross-sections, and look at 48-month horizons. These tables confirm that relative price volatility is dramatically lower during the period 1988:5-1994:11 than over the entire sample period, in both "full" and "border towns" cross-sections: relative price and nominal exchange rate changes are each about 3 times less volatile during the sub-period. The tables also show that the results are very robust to calculating volatilities using 48-month changes instead of two-month and six-month changes. Because PPP should hold better at longer horizons than short ones, we might have expected the volatility of cross-border relative price changes to be closer to the intra-national city pairs at the longer horizon. However, the variability of relative price changes at the 48-month horizon is six times larger for international city pairs (21.2) than for intra-national pairs (3.82), just as it was for the short-horizon changes.

III. Regressions

Following Engel and Rogers (1996), our regressions attempt to explain $V(\Delta P(j,k))$, the volatility (standard deviation) of $\Delta P(j,k)$.[8] We hypothesize that the volatility of the prices of similar goods sold in different locations is related to the distance between the locations and other explanatory variables, including the exchange rate and a dummy variable, Border, for whether the cities are in different countries.

We estimate regressions of the form:

$$V(\Delta P(j,k)) = \sum \alpha(m) D(m) + \beta r(j,k) + X\theta + u(j,k)$$

where $D(m)$ is a dummy variable for each city in our sample, $r(j,k)$ is the log distance between cities j and k, and X is a vector of explanatory variables that differ across specifications. The error is denoted $u(j,k)$. Note that all regressions are cross-sectional. We first calculate $V(\Delta P)$ over the full-sample of cities and the entire time period, and then over different sub-periods and different sub-sets of cities, as depicted in Figure 1.

As in the gravity model of trade, we posit a concave relationship between relative-price volatility and distance. We expect β to be positive. Transport costs should be larger the greater the distance between locations. In addition, more proximate locations are likely to be subjected

[8]In earlier work, Engel and Rogers (1996, 1999) also consider several alternative proxies for deviations from PPP, including root mean square errors instead of standard deviations and having a stationary autoregressive representation for P(j,k) instead of differencing the series. In the former case, results were essentially identical, since for most cities the difference in the drift terms is very close to zero [Engel-Rogers (1999)]. Results are also very robust to using the residuals from a stationary AR(6) representation of relative prices as the proxy for deviations from PPP [Engel-Rogers (1996)].

to more similar real supply and demand shocks.

We include a separate dummy variable for each city in our sample, D(m), so that for city pair (*j*,*k*) the dummy variables for city *j* and city *k* take on the values of 1. The inclusion of separate dummies for each individual location allows the standard deviation of price changes to vary from city to city. An empirical motivation for this comes from table 2, which indicates somewhat greater relative price volatility for Mexican cities than U.S. or Canadian cities. This might reflect differences across countries in methodologies for gathering price data. For instance, the U.S. cities that report prices only bi-monthly may have additional volatility that is introduced by measurement error from the less frequent observation of prices.

The variables included in X differ across specifications. We are particularly interested in whether there is a border effect, thus we have included the dummy variable "Border", which takes on a value of unity if cities *j* and *k* are in different countries. We also measure the importance of each border separately, by including in X individual border dummies, US-Canada, US-Mexico, and Canada-Mexico (which sum to Border). We expect the coefficients on the border dummies to be positive, and are interested in comparing the width of the U.S.-Mexican border to the U.S.-Canadian border.

One reason the border dummies might be important is because nominal prices are sticky in local currency terms, and hence are not adjusted optimally in response to fluctuations in the nominal exchange rate [see Betts and Devereux (2000) and Devereux and Engel (1998)]. One way to examine this channel is to include in *X* the volatility (standard deviation) of nominal exchange rate changes between locations *j* and *k*, $V(\Delta s(j,k))$. Including this variable in addition to Border allows us to go beyond what we were able to do in Engel and Rogers (1996). In the

U.S.-Canadian data set used in that paper, there is no distinction between the Border dummy and nominal exchange rate variability, since all cross-border pairs have the same nominal exchange rate. With the addition of data from Mexico, we are able to examine the importance of the Border dummy while accounting for the effect of nominal exchange rate variability.

There are several other potential explanations for the large border effect in relative prices (that is, the "country" effect on relative price variability, holding constant the effect of distance). First, there may be important barriers to trade. Although there are no longer many formal barriers to trade within the NAFTA countries, there were earlier in the sample. There are also informal trade barriers, even after NAFTA. Second, marketing and distribution networks may be more homogenous within countries than across borders, perhaps in part because of language. Third, because tastes are different, and because markets can be segmented due to our previous considerations, prices can differ across locations, even for identical goods. If tastes are more homogenous within countries than across the border, this will contribute to large positive estimates on the border dummies. Fourth, labor markets are undoubtedly more integrated within countries than across the border, suggesting that there should be a large border effect on important input prices. We shed light on the importance of these factors in accounting for the large border effect on relative prices.

Full-Sample Regression Results

The main results of the paper are in Tables 3A and 3B. Tables 4-8 assess robustness. Table 3A presents regression results for the full sample of cities, over the period 1980-1997, with the variables in 2-month changes. The first column presents the results of regressing the standard deviation of the log relative price on log(Distance), Border, and 28 individual location dummies (whose values are not reported). This regression is therefore identical to the main regression in Engel and Rogers (1996). The coefficient on distance is 1.13, which is significant at the 5 percent level. The coefficient on the Border dummy is 4.81 (standard error of 0.18), slightly more than four times the effect of distance. Notice from the bottom panel of Table 3A that distance is significant in explaining relative price variation even within countries.[9]

In the second specification we replace Border with its constituent parts: a dummy variable indicating whether the pair lies across the U.S.-Canada, U.S.-Mexico or Canada-Mexico border. We know that the coefficients on these three dummies will sum to 4.81, the coefficient on Border from specification #1. We expect that the border effect for pairs including one Mexican city is larger than the border effect in U.S.-Canadian city prices, which Engel and Rogers (1996) found to be quite large. Column 2 tells us how much larger, and gives us a glimpse of "How wide is the Rio Grande?". The coefficient on the U.S.-Mexico border dummy is 6.70, more than 25 times larger than the effect of distance. The coefficient on the Canada-Mexico dummy is of the same magnitude. The U.S.-Canada dummy is estimated to be 1.02, which is still highly significant given the standard error of 0.06. This is similar to the findings in Engel and Rogers (1996), who report an average border coefficient of 1.19 across the fourteen goods studied.

We would like to understand what causes the relatively large border effect for pairs involving Mexican cities, having noted several potential sources above. As a first attempt, consider specifications 3 and 4 of Table 3A, which introduce a measure of nominal exchange rate volatility, $V(\Delta s(j,k))$, to the regression. For specification 3, the explanatory variables are

[9]Consistent with Engel and Rogers (1996, 1999) and Cecchetti, Mark, and Sonora (2000).

log(Distance), Border, V(Δs(*j*,*k*)), and the 28 individual city dummies. In the fourth specification we replace Border with the US-MX and CA-MX dummy variables. Once the measure of nominal exchange rate volatility is included in the model, the effect of distance is substantially weaker, and all of the border dummies lose their significance entirely. For example, in specification 4, the coefficient on the standard deviation of the exchange rate is 0.66 with a standard error of 0.04, while the estimate on US-MX falls from 6.70 to 0.26 (with standard error 0.37). Thus it seems that a very large part of the border effect is from variable nominal exchange rates under sticky prices.

Analysis of the Sub-Samples

Table 3B presents the regression results for the truncated sample period May 1988 to November 1994. The regression specifications mirror those of Table 3A. During the stable peso period we see that the coefficients on the Border dummy (specification 1) or the US-MX and CA-MX dummies (specification 2) are notably smaller and less significant than in the entire sample period. The coefficient on the US-MX dummy variable falls from 6.70 to 1.20, nearly the same as the coefficient on the US-CA dummy (1.21).[10] The border effects in the shorter sample period are only about one-third the size of the distance effect, as opposed to being several times *larger* than the distance effect in the full period.

Geography

Having shown that the size of the border effect in U.S.-Mexican relative prices drops dramatically during *El Pacto*, we now estimate the border effect using only the sub-set of

[10]Notice that the coefficient on US-CA actually increases (from 1.02 to 1.21), despite the fact that the sub-period coincides with the beginning of the US-Canada Free Trade Agreement.

locations comprising our “border towns sample” (San Diego, Los Angeles, Houston, Dallas, Tijuana, Mexicali, Juarez, and Matamoros). These cities are geographically much closer to one another than are the cities in the full sample, and are presumably subject to relatively similar regional supply or real demand shocks. In addition, the Mexican border towns are all in the frontier zone through which goods have passed relatively freely for years before the NAFTA, and whose labor markets Robertson (2000) found to be well-integrated with U.S. labor markets.

Table 4 presents results for the stable peso period, 1988:5-1994:11.[11] Columns 1 and 2 contain results for the full sample of cities, while column 3 contains the results for the limited sample of border cities. The coefficient on Border in specification 1, or on US-MX in specification 2, is slightly smaller in the border towns sample, dropping to 1.93 from over 2 ½ in the full sample. This suggests that, to the extent that the geographically-proximate cities are indeed subject to similar regional supply and demand shocks, the effects of such shocks on relative price variability is not nearly as large as the effect of nominal exchange rate fluctuations.

Trade Barriers

We would also like to gauge the importance of trade barriers in explaining the large border effect in U.S.-Mexican relative prices. Given the advent of NAFTA in 1993, it is natural to think of estimating the border effect on samples of data before and after the agreement. Using 1993 as a cut-off date for determining the sub-periods might understate the effect of trade barriers, however, both because NAFTA was implemented gradually and because of the instability in the Mexican economy in the aftermath of the December 1994 peso crisis.

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[11] In appendix table A-3 we show that the following results are robust to considering the entire sample period, 1984-1997, instead of just the period of *El Pacto*.

Hence, we examine a "post-crisis NAFTA" period, January 1996 - December 1999, which begins a few years after the free trade agreement went into effect and after the period of unusually large exchange rate variability that accompanied the peso crisis.[12] Table 5 contains the results. As a baseline, the first column displays the results from the "full" sample period, 1984-1997, with all of the available cities. The next two columns contain results from the stable peso period 1988-1994 and post-crisis NAFTA period 1996-99, respectively. In the full sample, the coefficient on the US-MX dummy is 9.76, consistent with our earlier results. According to the third column, this drops by about 50% in the post-crisis NAFTA period, to 4.88. Although this is a large drop in the estimated border coefficient, the estimate is still significantly larger than the coefficient of 2.72 during the stable peso period, shown in the second column. These results suggest that the reduction in trade barriers that occurred in the first few years of the NAFTA did have an effect on relative price volatility between the United States and Mexico, but not as large an effect as a stable nominal exchange rate had during *El Pacto*.

Long-Horizon Changes

We also ran regressions from the full sample of cities and time period, but using 48-month changes in relative prices and nominal exchange rates. If purchasing power parity holds more nearly at long horizons, we might detect a drop in the estimated border effects in 48-month changes from the 2-month and 6-month changes we've considered so far. The results are listed in Table 6. Border is still large and significant, as are distance and nominal exchange rate variability. However, unlike the results observed earlier, with the 48-month changes, Border

[12]We use 6-month changes in prices, and a more limited set of cities than the full sample, due to data availability, as noted above. During the post-crisis NAFTA sub-period, the standard deviation of the 6-month change in the nominal exchange rate was 4.19, compared to 2.19 during *El Pacto*.

continues to be highly significant, even after $V(\Delta s(j,k))$ is added to the regression. This suggests that the factors leading to the "real border effect" discussed by Engel and Rogers (2000) become relatively more important at longer horizons.

Labor Market Segmentation

Finally, we attempt to estimate how much of the large border effect on U.S.-Mexican relative prices is due to a border effect in relative wages. We use Robertson's (2000) data on manufacturing wages in six Mexican cities, and data from the website of the U.S. Bureau of Labor Statistics on manufacturing wages in the U.S. states. These states correspond to the cities for which we have price data (the BLS does not publish wage data by city).

Table 7a displays by location the average hourly wage, in U.S. dollars, over the period 1987-1998 (the available sample for the Mexican wage data). Clearly, there is a large border effect, with wages in the U.S. about 10 times larger than in Mexico. To see if the large U.S.-Mexican border effect in relative CPIs is due to the border effect in relative wages, we add to our regressions W(j,k), the relative hourly manufacturing wage, in U.S. dollars, across locations j and k.[13] Table 7b presents the results of specifications with distance, the US-MX border dummy, individual location dummies (not shown), and with and without W(j,k). The dependent variable is the 6-month change in relative prices. We report results for two sample periods, 1987:I-98:II and 1988:II-1994:II.

According to table 7b, the US-MX border dummy is positive and significant in all specifications. The coefficient on W(j,k) is also positive and significant, with t-statistics of 2.7

[13]Because we have data on wages themselves, rather than a wage index, we put the relative wage rate into the regression, rather than the variability of the change in the relative wage. The results are unchanged if we use $V[\Delta W(j,k)]$ in the regression instead.

and 4.0. However, the addition of W(j,k) to the regression does nothing to affect our earlier conclusions: the border effect on relative prices is large, even when we account for wages (columns 1 vs. 2 and 3 vs. 4), and the border effect is strongly influenced by the degree of nominal exchange rate variability (columns 1 or 2 vs. 3 or 4).

Disaggregated Relative Prices Within Tijuana and Mexico City

We would like to have more direct evidence on how much of the large border effect on U.S.-Mexican relative prices is a real border effect. Such evidence is difficult to come by if we restrict ourselves to using only data on the aggregate CPI. To understand why, consider the following insights that have been exploited by Engel (1993, 1999), Rogers and Jenkins (1995), Betts and Kehoe (1999), and Mendoza (2000). These authors note that in a simplified two-good setting, with β denoting the share in the overall price index of the “non-tradeable” good, relative prices between locations j and k, q(j,k), can be written as:

$$q(j,k) = p(j) + s(j,k) - p(k) = [p_T(j) + s(j,k) - p_T(k)] + \beta[(p_N(k) - p_T(k)) - (p_N(j) - p_T(j))]$$

$$q(j,k) = x(j,k) + y(y(k),y(j))$$

Thus the real exchange rate is the sum of two components, the common currency price of tradeables across locations, labeled x(*j*,*k*) and the relative price of non-tradeables to tradeables within locations, y(*k*) and y(*j*).

As discussed in the papers above, models of real exchange rate determination can conveniently be classified into two groups. In models of real exchange rate determination without nominal rigidities, movements in q(*j*,*k*) are accounted for by movements in y. In models with sticky prices, variation in q(*j*,*k*) is due to movements in x(*j*,*k*), as s(*j*,*k*) varies while p_T and p_N fluctuate very little.[14]

Clearly we cannot rule out a large real border effect -- reflected in large movements in the y component -- by examining only aggregate CPI data. The strong positive correlation between the size of the border effect and the degree of nominal exchange rate variability that we have documented so far, could result from the endogenous response of nominal exchange rate fluctuations to real side factors. In the extreme, traded goods prices in Mexico could have adjusted by an amount sufficient to offset the documented movements in the nominal exchange rate, so that *all* of the variability in q was due to the real factors that influence the y component. If such factors were decisive in accounting for the border effect, *we ought to see large movements in the relative prices of different goods within Mexico, and a decline in these relative price movements during the stable peso period.*[15]

To investigate this possibility, we examine Mexican price index data on 276 very narrow categories from three locations: Mexico City, Tijuana, and Mexico's city-wide average (referred to simply as "Mexico"). The items comprise the entire CPI at the most disaggregated level. The items are listed in the Appendix table, arranged into 8 sub-components of the CPI ("Food, beverages, and tobacco"; "Clothing"; "Transport" etc.).

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[14]Using various proxies for traded goods prices, Engel and Rogers-Jenkins find that for the G-7 countries, the x component accounts for the vast majority of the variability of the real exchange rate. These authors argue that this is strong evidence of the importance of sticky-prices in explaining deviations from the law of one price. On the other hand, Mendoza and Betts-Kehoe suggest that this result might not be completely robust to Mexico, at least not in all sub-periods (Mendoza) and for all types of goods prices (Betts and Kehoe).

[15]In principle, we might also expect to see large relative price movements of different goods within U.S. and Canadian cities, but the changes in the size of the border effect documented above are clearly due to changes in the behavior of Mexican prices.

We start by calculating for each item the standard deviation of the two-month change in its price relative to each of the other 275 items in the sample.[16] The mean of the 275 relative price volatilities produces a within-city measure of average relative price variability, by item.

In Figure 2, we plot this measure of average relative price volatility for each of the 276 items. The top panel shows results for the average of all Mexican cities, the middle and lower panels depict results for Mexico City and Tijuana, respectively. For ease of exposition, we display to the right of the individual item results, the mean of the 276 measures of average relative price variability.[17] This is the bar labeled "Average of the Relative Prices". We also display on the far right of each panel the standard deviation of the two-month change in the nominal peso/dollar exchange rate. This is the bar labeled "s". Finally, for the reasons discussed above, we compare results from the full period (lightly-shaded in the figures) to those of the stable peso period (shaded in dark).[18]

Each of the panels in figure 2 paints the same story: a noticeable decline in within-city relative price variability during the stable peso period compared to the full sample. With the exception of a few of the relative prices in Category I (Food, beverages, and tobacco), which experienced an increase in variability in the sub-period, the decline in relative price variability is across the board. Also, notice that the decline is fairly large, for many of the items on the order of 30 to 40 percent. Given the rapid stabilization of Mexico's aggregate CPI inflation rate, from well above 100% in 1987 to 20% in 1989 and under 10% in 1993, the drop in relative price

[16] We also performed the analysis on six-month price changes. The results are robust.

[17] Thus, for each location, this is the mean of all 75,900 (=276x275) relative price volatilities in the sample.

[18] The complete set of data begins in 1982, so the full sample period is 1982:1-1997:12 in this case.

variability within Mexico is consistent with the large literature on relative prices and inflation.[19]

Figure 2 thus suggests that the factors identified with a "real border effect" are important. However, the figure also indicates that the decline in within-location relative price variability from the full sample period to the stable-peso sub-period is not nearly as large as the ten-fold decline in nominal exchange rate variability.[20] This suggests that the large drop in the border effect in U.S.-Mexican relative prices during 1988-94 is not primarily accounted for by a drop in the variability of the equilibrium real exchange rate.

IV. Conclusion

We use consumer price indexes from cities in the U.S., Canada and Mexico, to quantify the extent to which prices fail to equalize *across countries* against a baseline of the size of failures across regions *within countries*. That is, we compute the Engel-Rogers "width of the border" measures for Mexico. We show that the border effect in U.S.-Mexican prices is nearly an order of magnitude larger than in U.S.-Canadian prices, over the sample period 1980-1997.

We then examine subsets of the CPI data, and incorporate data on wages and highly disaggregated goods prices for Mexican cities, as a way of presenting evidence on alternative explanations of the large border effect for pairs involving Mexican cities. These explanations include sticky prices and variable nominal exchange rates; formal or informal barriers to trade; and labor markets, marketing networks and distribution networks that are more homogenous

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[19]See Vining and Elwertowski (1976), Parks (1978), Cecchetti (1985), Domberger (1987), Lach and Tsiddon (1992), Bomberger and Makinen (1993), Parsley (1996), and DeBelle and Lamont (1997).

[20]The largest drop in variability for *any* of the 75,900 relative prices was four-fold in the case of Mexico City (from 0.30 in the full sample to 0.075 in the sub-period), three-fold for Tijuana, and 2 ½ fold for Mexico.

within countries than across borders.

We interpret our results as strong evidence of a “nominal border effect” in relative prices within NAFTA, but take seriously that proposition that the results are in part consistent with a “real border effect”. In particular, we present evidence suggesting that there was a decline in the variability of the equilibrium peso-dollar real exchange rate during the sub-period 1988-94, compared to the full sample. Given the crisis of December 1994 through 1995, this should not be too controversial. Nonetheless, although the “real borders effect” explanation has some merit, it probably explains less of the border effect than the one that relies on sticky local-currency goods prices.

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Table 1: Locations and CPI Data Availability

A. "Full Sample":

Country:	Location	Availability	
United States		*1980-1997*	*1998-present*
	Chicago	Monthly	Monthly
	Los Angeles	Monthly	Monthly
	New York	Monthly	Monthly
	Philadelphia	Monthly	Even months
	Baltimore	Odd months	NA
	Boston	Odd months	Odd months
	Miami	Odd months	Even months
	St. Louis	Odd months	NA
	Washington, D.C	Odd months	NA
	Dallas	Even months	Odd months
	Detroit	Even months	Even months
	Houston	Even months	Even months
	Pittsburgh	Even months	NA
	San Francisco	Even months	Even months
	Baltimore/Washington	NA	Odd months
Canada (monthly, 1980:1-present)	Alberta	Nova Scotia	
	British Columbia	Ontario	
	Manitoba	Prince Edward Island	
	New Brunswick	Quebec	
	New Foundland	Saskatchewan	
Mexico (monthly, 1980:1-present)	Acapulco	Mexicali	
	Chihuahua	Mexico City	
	Guadalajara	Monterrey	
	Hermosillo	Tampico	
	Juarez	Tijuana	
	Matamoros	Veracruz	
	Merida	Villahermosa	

B. Border Towns Sample: semi-annual data, 1984:I-1997:II

Country:	Location of Price Index Used	
United States	Dallas	Los Angeles
	Houston	San Diego
Mexico	Juarez	Mexicali
	Matamoros	Tijuana

Notes: Price data is from CANSIM (http://www.statcan.ca/start.html), the Bank of Mexico (http://www.banxico.org.mx/siteBanxicoINGLES/index.html), the Bureau of Labor Statistics (http://stats.bls.gov/), and nominal exchange rate data are taken from the International Monetary Fund's *International Financial Statistics*.

Table 2A: Summary Statistics for "Full Sample" of cities, 2-Month Changes

Pairs:	Std. Dev. ΔP(j,k) 80:1-97:12	Std. Dev. ΔP(j,k) 88:5-94:11	Std. Dev. ΔS(j,k) 80:1-97:12	Std. Dev. ΔS(j,k) 88:5-94:11	Distance	#obs
All	4.64	1.61	5.18	1.13	1518	678
Intra-national	1.03	0.67	0.00	0.00	1001	202
US-US	0.81	0.70	0.00	0.00	1071	66
CA-CA	0.53	0.53	0.00	0.00	1343	45
MX-MX	1.44	0.72	0.00	0.00	782	91
Inter-national	6.17	2.01	7.37	1.61	1737	476
US-CA	1.78	1.84	1.54	1.57	1428	140
US-MX	8.02	1.94	9.75	1.37	1572	196
CA-MX	7.97	2.28	9.88	2.00	2277	140

Table 2B: Summary Statistics for 6-Month Changes, 1984-1997

Pairs:	Full Sample of cities		Border Towns Sample			
	Std. Dev. ΔP(j,k)	Std. Dev. ΔS(j,k)	Std. Dev. ΔP(j,k)	Std. Dev. ΔS(j,k)	Dist.	#obs
All	6.56	9.31	6.04	9.89	750	28
Intra-national	1.27	0.00	1.10	0.00	818	12
US-US	0.97	0.00	0.85	0.00	904	6
CA-CA	0.80	0.00	---	---	---	---
MX-MX	1.72	0.00	1.35	0.00	731	6
Inter-national	8.81	13.3	9.75	17.3	700	16
US-CA	3.20	2.85	---	---	---	---
US-MX	11.2	17.3	9.75	17.3	700	16
CA-MX	11.0	18.0	---	---	---	---

Notes: Columns display the mean values of the standard deviation of changes in the relative price between location j and k, P(j,k) and the change in the nominal exchange rate S(j,k), distance (in miles), and the number of observations. Prices are in U.S. dollars. Listed by row is the sample of cities used in the calculations. The first row uses all locations; US-US indicates that only the within-US city pairs are used; CA-CA and MX-MX are the analogues for Canada and Mexico; intra-national indicates that only pairs of cities within countries are used in the calculations; and international indicates that only cross-border pairs are used.

Table 3A: Regression Results for the "Full Sample"

Using all pairs of cities

Specification	1	2	3	4
Log Distance	1.13 (0.12)	0.26 (0.03)	0.25 (0.03)	0.26 (0.03)
Border	4.81 (0.18)	---	-0.04 (0.07)	---
US-CA	---	1.02 (0.06)	---	---
US-MX	---	6.70 (0.05)	---	0.26 (0.37)
CA-MX	---	6.73 (0.07)	---	0.20 (0.37)
V(Δs(j,k))	---	---	0.69 (0.01)	0.66 (0.04)
Adj. R^2	.72	.98	.98	.98

Using only within-country city pairs

Specification	USUS	CACA	MXMX
Log Distance	4.66 (1.30)	3.54 (0.48)	19.7 (3.95)
Adj. R^2	.68	.75	.59

Notes: The sample period is 1980:1-97:12, for the full set of cities. The dependent variable is the standard deviation of the 2-month change in the log relative price. The independent variables in the top panel are: the log of distance between cities in the particular pair (in miles); Border, which equals unity if the cities in the pair lie across an international border; US-CA (US-MX or CA-MX) if the cities lie across the U.S.-Canadian (U.S.-Mexican or Canadian-Mexican) border; and V(Δs(j,k)), the standard deviation of the 2-month change in the nominal exchange rate. All specifications include 38 individual city dummies. In the bottom panel, regressions are run only for the within-country pairs indicated in the top row. Coefficients and standard errors on log distance have been multiplied by 100.

Table 3B: *El Pacto* (Stable Peso) Period, 1988:5-94:11
Full Sample of Cities, 2-Month Changes

Specification	1	2	3	4
Log Distance	8.56	3.16	4.86	3.16
	(1.01)	(0.58)	(0.62)	(0.58)
Border	1.29	---	0.17	---
	(0.01)		(0.03)	
US-CA	---	1.21	---	---
		(0.01)		
US-MX	---	1.20	---	0.14
		(0.01)		(0.01)
CA-MX	---	1.63	---	0.08
		(0.01)		(0.01)
V(Δs(j,k))	---	---	0.72	0.77
			(0.02)	(0.01)
Adj. R^2	.94	.98	.98	.98

Notes: as in table 3A, with the sample period now 1988:5-94:11.

Table 4: *El Pacto* Period:
Comparison of the two different samples of cities using 6-Month Changes

	Full Sample		Border Towns
Specification	1	2	3
Log Distance	19.2	4.92	5.44
	(2.31)	(1.32)	(4.46)
Border	2.67	---	1.93
	(0.03)		(0.08)
US-CA	---	2.36	---
		(0.02)	
US-MX	---	2.55	---
		(0.02)	
CA-MX	---	3.41	---
		(0.03)	
V(Δs(j,k))	---	---	---
Adj. R^2	0.93	0.98	0.95

Notes: as in table 3A, with the sample period 1988:5-94:11 and using 6-month changes in the log relative prices and nominal exchange rates.

Table 5: Assessing the Effects of NAFTA:
Extended Full Sample using Semi-Annual Data

Specification	1984:I-1997:II	1988:II-1994:II	1996:I-1999:II
Log Distance	15.6	5.51	0.57
	(2.25)	(1.25)	(0.98)
US-CA	2.25	2.22	2.49
	(4.23)	(0.03)	(0.02)
US-MX	9.76	2.72	4.88
	(0.04)	(0.02)	(0.02)
CA-MX	9.61	3.60	2.26
	(0.05)	(0.03)	(0.02)
Adj. R^2	0.995	0.98	0.99

Notes: The dependent variable is the standard deviation of the 6-month change in the log relative price. The independent variables are as described above. The sample period is indicated in the top row. All specifications include individual city dummies. Coefficients and standard errors on log distance have been multiplied by 100.

Table 6: Regression Results for 48-Month Changes
("Full Sample")

Specification	1	2	3
Log Distance	1.93	0.62	0.40
	(0.24)	(0.11)	(0.13)
Border	17.0	---	9.03
	(0.35)		(0.28)
US-CA	---	10.0	---
		(0.21)	
US-MX	---	21.7	---
		(0.19)	
CA-MX	---	18.0	---
		(0.23)	
$V(\Delta s(j,k))$	---	---	0.13
			(.003)
Adj. R^2	0.86	0.97	0.96

Notes: as in table 3A, using 48-month changes in the log relative prices and nominal exchange rates. In addition, the coefficients and standard errors on log distance have not been multiplied by 100.

Table 7a: Manufacturing Wage Rates

Location	Wage ($)	Location	Wage ($)
California	12.1	Juarez	1.33
Florida	9.66	Guadalajara	1.36
Illinois	12.1	Matamoros	1.76
Maryland	12.5	Mexico City	1.60
Massachussets	12.0	Monterrey	1.66
Michigan	15.2	Tijuana	1.62
New York	11.8		
Pennsylvania	12.0		
Texas	11.0		

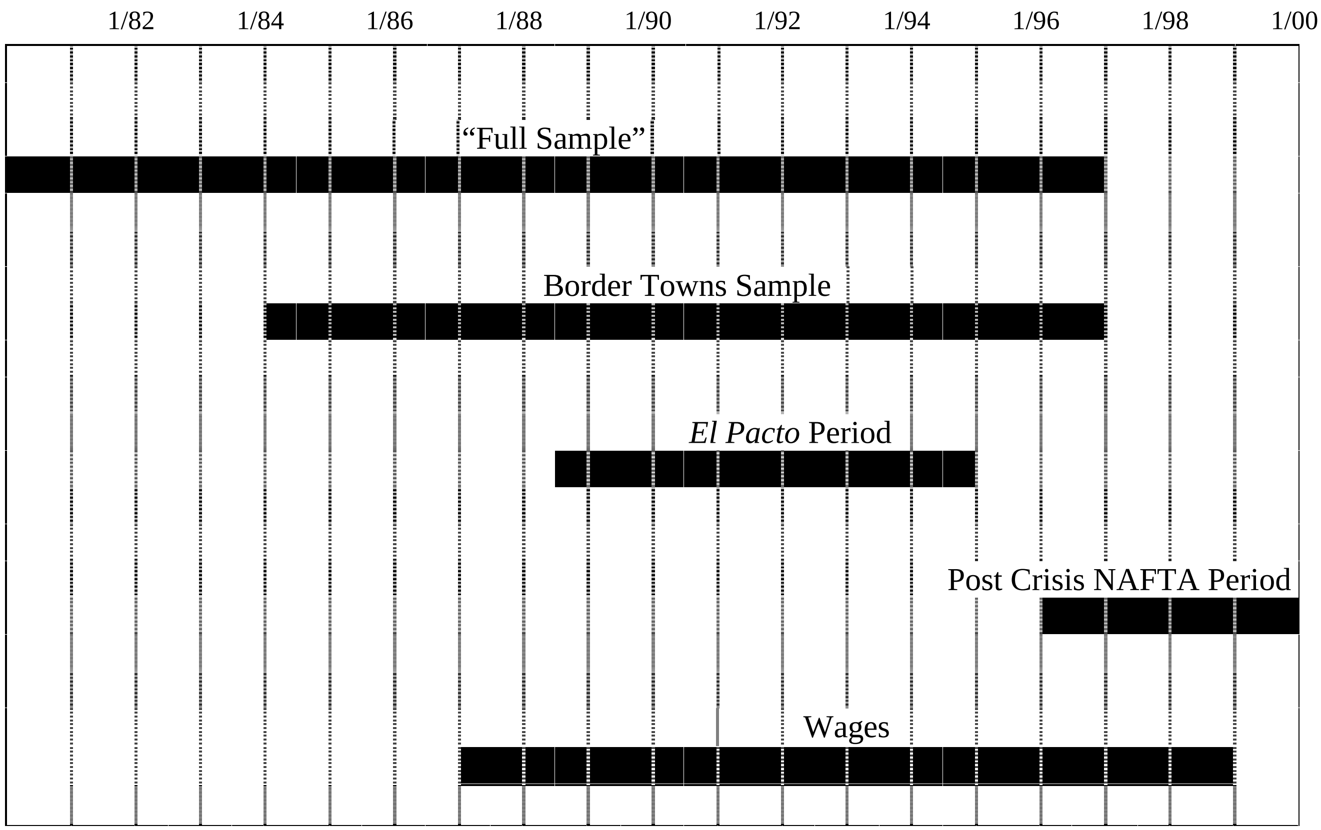
Notes: Average hourly wage rate for manufacturing workers, 1987-1998, in U.S. dollars

Table 7b: Assessing the Role of Wages

Specification	1987:I-1998:II		1988:II-1994:II	
	1	2	3	4
Log Distance	7.47	6.03	4.21	2.39
	(3.24)	(3.23)	(2.48)	(2.36)
Border	8.61	8.07	2.77	2.09
	(0.05)	(0.25)	(0.04)	(0.18)
W(j,k)	---	0.08	---	0.12
		(0.03)		(0.03)
Adj. R^2	0.998	0.998	0.99	0.99

Notes: Regressions are run for two different time periods: 1987:I-1998:II (the period for which we have wage data) and 1988:II-1994:II (the *El Pacto* period). The dependent variable is the standard deviation of the 6-month change in the log relative price. W(j,k) is the relative wage between each location in U.S. dollars. 15 individual location dummies listed in Table 7a are also included in the regresssion. Coefficients and standard errors on log distance have been multiplied by 100.

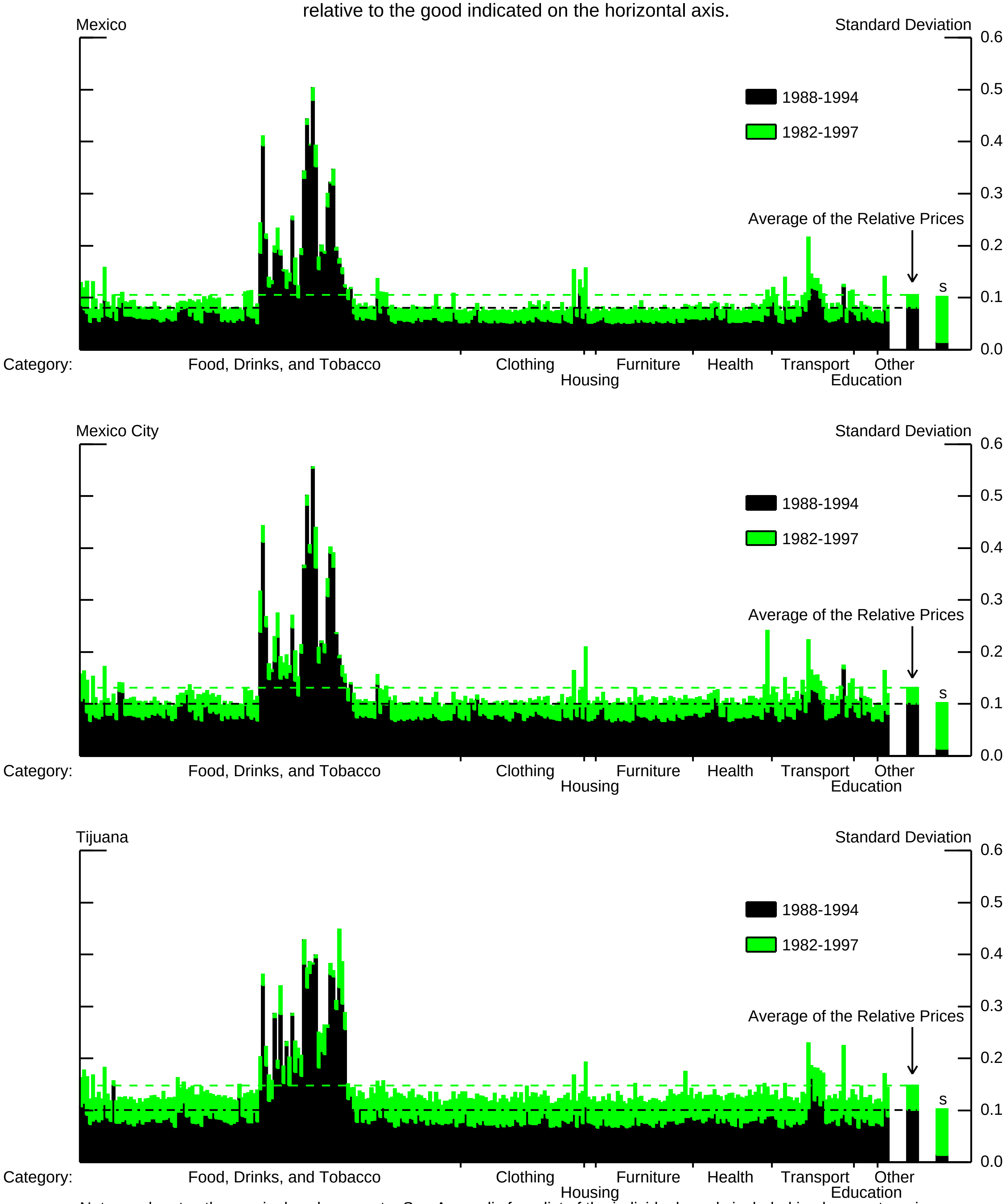
Figure 1: Description of Sample Periods



Notes: Full Sample is the longest available sample period for which there are data from the maximum number of cities. Border Towns Sample is the period for which there are available data from eight locations on the U.S.-Mexico border. *El Pacto* is the stable peso period from 1988:5-1994:11. The Post Crisis NAFTA sample includes all cities with continuously available data. Wages is the sample period, 1987:I-1998:II, determined by the availability of wage data for Mexican cities (Robertson 2000).

Figure 2: Relative Price Volatility within Mexico, Mexico City, and Tijuana

Average variability of the two-month change in the price of each good
relative to the good indicated on the horizontal axis.



Notes: s denotes the nominal exchange rate. See Appendix for a list of the individual goods included in above categories.

Table A-1: Summary Statistics for 6-Month Changes, 1988:5-1994:11

Pairs:	Full Sample		Border Towns Sample	
	Std. Dev. ΔP(j,k)	Std. Dev. ΔS(j,k)	Std. Dev. ΔP(j,k)	Std. Dev. ΔS(j,k)
All	2.98	3.10	2.00	2.37
Intra-national	1.04	0.00	0.90	0.00
US-US	0.98	0.00	0.85	0.00
CA-CA	0.87	0.00	---	---
MX-MX	1.16	0.00	0.95	0.00
Inter-national	3.80	4.41	2.81	4.15
US-CA	3.31	3.10	---	---
US-MX	3.67	4.15	2.81	4.15
CA-MX	4.48	6.10	---	---

Table A-2: Summary Statistics for 48-Month Changes

Pairs:	Full Sample (1980-1997)		Border Towns Sample (1984-1997)	
	Std. Dev. ΔP(j,k)	Std. Dev. ΔS(j,k)	Std. Dev. ΔP(j,k)	Std. Dev. ΔS(j,k)
All	16.0	47.0	14.7	50.9
Intra-national	3.82	0.00	3.70	0.00
US-US	2.31	0.00	3.24	0.00
CA-CA	2.20	0.00	---	---
MX-MX	5.71	0.00	4.16	0.00
Inter-national	21.2	66.9	22.9	89.0
US-CA	12.6	11.4	---	---
US-MX	26.2	89.0	22.9	89.0
CA-MX	22.6	91.6	---	---

Notes: see notes to tables 2A and 2B.

Table A-3: Regression Results for 1984-1997, 6-Month Changes

	Full Sample			Border Towns	
Specification	1	2	3	4	5
Log Distance	128.7 (15.3)	15.6 (2.25)	12.8 (2.43)	3.43 (3.91)	3.43 (3.91)
Border	7.22 (0.23)	---	0.88 (0.05)	8.66 (0.07)	---
US-CA	---	2.25 (4.23)	---	---	---
US-MX	---	9.76 (0.04)	---	---	---
CA-MX	---	9.61 (0.05)	---	---	---
V(Δs(j,k))	---	---	0.50 (.003)	---	0.50 (.004)
Adj. R^2	0.75	.995	0.99	.998	.998

Notes: as in table 3A, with the sample period 1984-97 and using 6-month changes in the log relative prices and nominal exchange rates.

Appendix: List of Disaggregated Goods in Mexico, Mexico City, and Tijuana

I. Alimentos, bebidas y tabaco

Tortilla de maíz
Masa de maíz
Harina de maíz
Fécula de maíz
Pan blanco
Pan de caja
Pan dulce
Pastelillos y pasteles
Harinas de trigo
Pasta para sopa
Galletas populares
Arroz
Cereales en hojuela
Pollo entero
Pollo en piezas
Pulpa de cerdo
Chuleta
Lomo
Pierna
Bistec de res
Carne molida de res
Cortes especiales de res
Retazo
Hígado de res
Otras vísceras de res
Jamón
Tocino
Chorizo
Salchichas
Pastel de carne
Carnes ahumadas o enchiladas
Carnes secas
Otros embutidos
Huachinango
Robalo y mero
Mojarra
Otros pescados
Camarón
Otros mariscos
Atún en lata
Sardina en lata
Otros pescados y mariscos en conserva
Leche pasteurizada envasada
Leche sin envasar
Leche en polvo
Leche evaporada
Leche condensada
Leche maternizada
Crema de leche
Mantequilla
Queso amarillo
Queso chihuahua o manchego
Queso fresco
Otros quesos
Yoghurt
Helados
Huevo
Aceite vegetal
Manteca vegetal
Margarina
Manteca de cerdo
Naranja
Limón
Toronja
Plátano tabasco
Otros plátanos
Melón
Papaya
Sandía
Piña
Uva
Manzana
Aguacate
Mango
Pera
Guayaba
Jitomate
Tomate verde
Chile serrano
Chile poblano
Cebolla
Ajo
Papa
Zanahoria
Chícharo
Calabacita
Chayote
Pepino
Col
Lechuga
Elote
Frijol
Chile seco
Otras legumbres secas
Chiles procesados
Puré de tomate
Verduras envasadas
Frutas y legumbres preparadas para bebés
Sopas enlatadas

Jugos o néctares envasados
Mermeladas
Azúcar
Café soluble
Café tostado
Refrescos envasados
Sal
Concentrado de pollo
Pimienta
Mostaza
Mayonesa
Chocolate en tableta
Chocolate en polvo
Dulces y caramelos
Concentrados para refrescos
Gelatina en polvo
Cajetas
Miel de abeja
Papas fritas y similares
Carnitas
Barbacoa o birria
Pollos rostizados
Cerveza
Vino de mesa
Brandy
Ron
Tequila
Otros licores
Cigarrillos

II. Ropa, calzado y accesorios

Camisas
Camisetas
Calzoncillos
Calcetines
Pantalón hombre base algodón
Pantalón hombre otros materiales
Trajes
Otras prendas para hombre
Blusas para mujer
Medias y pantimedias
Ropa interior para mujer
Pantalón mujer base algodón
Pantalón mujer otros materiales
Otras prendas para mujer
Vestido para mujer
Falda para mujer
Conjunto para mujer
Pantalón niño base algodón
Pantalón niño otros materiales
Blusa para niño
Ropa interior para niño
Vestido para niña
Ropa interior para niña
Traje para bebé
Camiseta para bebé
Suéter para niño
Suéter para niña
Chamarras
Abrigos
Sombreros
Uniforme para niño
Uniforme para niña
Zapatos para hombre
Zapatos para mujer
Zapatos para niños
Zapatos tenis
Servicio de tintorería y lavandería
Reparación de calzado
Bolsas, maletas y cinturones
Relojes
Joyas y bisutería

III. Vivienda

Renta de vivienda
Electricidad
Gas doméstico
Otros combustibles

IV. Muebles aparatos y accesorios domésticos

Estufas
Antecomedores
Calentadores para agua
Muebles para cocina
Recamaras
Colchones
Comedores
Salas
Refrigeradores
Lavadoras de ropa
Planchas eléctricas
Licuadoras
Maquinas de coser
Televisores y videocaseteras
Radios y grabadoras
Equipos mudulares
Cerillos
Velas y veladoras
Focos
Sabanas
Colchas
Cobijas
Toallas
Cortinas
Hilos y estambres
Detergentes y productos similares

Jabón para lavar
Blanqueadores y limpiadores
Desodorantes ambientales
Loza y cristalería
Baterías de cocina
Utensilios de plástico para el hogar
Escobas

V. Salud y cuidado personal

Analgésicos
Antigripales
Expectorantes y descongestivos
Antibióticos
Gastrointestinales
Anticonceptivos y hormonales
Nutricionales
Consulta médica
Operación quirúrgica y partos
Hospitalización
Cuidado dental
Análisis
Corte de cabello
Sala de belleza
Servicio de baño
Jabón de tocador
Pasta dental
Productos para el cabello
Desodorantes personales
Lociones y perfumes
Cremas para la piel
Artículos de maquillaje
Navajas y maquinas de afeitar
Papel Higiénico
Servilletas de papel
Toallas sanitarias
Pañales

VI. Transporte

Taxi
Metro o transporte eléctrico
Autobús foráneo
Ferrocarril
Cantinas
Cuotas licencias y otros documentos
Servicios funerarios
Transporte aéreo
Automóviles
Bicicletas
Gasolina
Aceites lubricantes
Neumáticos
Acumuladores
Otras refacciones
Mantenimiento de automóvil
Estacionamiento
Seguro de automóvil
Tenencia de automóvil

VII. Educación y esparcimiento

Jardín de niños y guardería
Primaria
Secundaria
Preparatoria
Universidad
Libros de texto
Otros libros
Cuadernos y carpetas
Plumas, lápices y otros
Hoteles
Cine
Espectáculos deportivos
Centro nocturno
Club deportivo
Periódicos y revistas
Revistas
Artículos deportivos
Juguetes
Discos y casetes
Instrumentos musicales y otros
Material y aparatos fotográficos

VIII. Otros servicios

Restaurantes, bares y similares
Loncherías
Cafeterías