Deviations from the Law of One Price: Sources and Welfare Costs

Charles Engel
University of Washington and NBER

John H. Rogers
Board of Governors of the Federal Reserve System

Abstract

We explore failures of the law of one price across European cities, attempting to move beyond a "first-generation" of papers that document very large border effects. We document two very distinct types of border effects embedded in relative prices. The first is a "real barriers effect", caused by various barriers to trade. The second is a sticky-consumer-price cum volatile exchange-rate effect. Both are shown to be important empirically, the second type especially so. For the first type of effect, the larger is the border effect the larger is the deadweight welfare loss. But for the second type of effect the issue is not so clear. While the existence of a border effect of the second type does imply economic distortions, it is not necessarily true, comparing two border effects across two pairs of countries, that a larger border effect of this type implies a larger deadweight welfare loss.

First Draft: June 3, 1999
This Draft: July 29, 1999

Charles Engel
Department of Economics
University of Washington
Seattle, WA 98195-3330
E-mail: cmengel@u.washington.edu

John H. Rogers
International Finance Division
Stop 22
Board of Governors
of the Federal Reserve System
Washington, DC 20551
E-mail: john.h.rogers@frb.gov

A spate of recent research has improved our understanding of the size, sources, and implications of the well-documented deviations from the law of one price internationally (Isard, 1977). Concerning size, one strand of the literature performs variance decompositions of real exchange rates. These studies find that in mature economies law of one price deviations are responsible for most of the variance of real exchange rates.[1] Another strategy has been to compare movements in goods prices across national borders to price movements between different regions within a country. Engel and Rogers (1996) demonstrate that prices of similar goods between U.S. and Canadian cities are systematically more variable than prices between equi-distant cities in the same country. By this "width of the border" metric, international failures of the law of one price are large. The Engel-Rogers finding is consistent with other research showing that goods, labor, and capital flow more readily between regions within a country than across borders (see McCallum, 1995 and papers in the volume edited by Hess and Van Wincoop, 1999).

Several insights have also been brought to bear on the possible sources of these failures of the law of one price. The potential sources include tariffs and non-tariff barriers to trade, transportation costs, non-traded inputs such as marketing and other distribution services that are a part of final goods prices, and variable nominal exchange rates under sticky prices. One explanation of this last source is that exporters "price to market". That is, producers selling abroad set prices in the consumers' currency rather than its own. Under local currency pricing (LCP), changes in nominal exchange rates do not affect goods prices in the local market, i.e., there is zero pass-through of exchange rate changes. Several papers present evidence of local currency pricing, including Giovannini (1988), Marston (1990), and Knetter (1993). Feenstra and Kendall (1997) find that a significant portion of observed deviations from purchasing power parity is attributable to incomplete exchange rate pass-through as a result of local currency pricing. Our results are consistent with Feenstra and Kendall's.

Interest in the extent to which local currency pricing can account for empirical failures of the law of one price has surged for two related reasons. First, with the launching of the Euro on January 1, 1999 the currencies of European Monetary Union members became irrevocably fixed on their way toward eventually disappearing from circulation. Across the Atlantic, advocates of a "dollar bloc" have recently proposed that some combination of Argentina, Mexico and Canada relinquish their national currencies and adopt the U.S. dollar. This would replace current arrangements that have produced different degrees of exchange rate flexibility, as did EMU. There have also been calls for a Japanese yen bloc in Asia. Deviations from the law of one price would be reduced under such unified currency blocs if local currency pricing is indeed a significant factor.

Second, several recent theoretical papers have argued that assumptions concerning price-setting behavior can affect debate about the merits of fixed versus floating exchange rate systems. Engel (1998) and Devereux and Engel (1998) point out that traditional arguments in favor of floating rates all assume producer-currency pricing (PCP), where prices are set in the currency of the producer. Under PCP, prices faced by consumers in an export market fluctuate with changes in the nominal exchange rate, so that there is complete pass-through of exchange rates to domestic prices; but, under local-currency pricing (LCP), prices are set in the currency of the consumer and there is zero pass-through of exchange rates to local prices. Tille (1998) and Bachetta and Van Wincoop (1999) also consider price-setting of this form, in dynamic general-equilibrium optimizing models, and highlight the role of deviations from the law of one price that result from different forms of price determination.

[1] Rogers and Jenkins (1995) and Engel (1999). A related literature is on purchasing power parity; see Rogoff (1996).

In this paper we undertake an empirical exploration of failures of the law of one price across European cities. Our principal focus is the role of local currency pricing and floating exchange rates in accounting for international failures of the law of one price. In Engel and Rogers (1996) we suggest that local currency pricing may be an important factor in explaining the large border effect in consumer price data from U.S. and Canadian cities. The present study is a much more direct examination of local currency pricing than the earlier paper. Using the framework of Devereux and Engel, we derive an expression for the variance of relative prices across locations in terms of the variance of nominal exchange rates and transportation costs. We use this model for two purposes. First, it motivates our regressions. In section III, we investigate the determinants of the volatility of real exchange rates among cities in Europe. We derive an expression that relates real exchange rate volatility to the volatility of nominal exchange rates, and to factors that influence transportation costs. That expression is very similar to the equations estimated in Engel and Rogers (1996), and is the basis for our regressions of section III.

We estimate this equation using consumer price data from European cities in 11 countries over the period 1981-97. The European data set has many advantages over that used in our earlier work. The Engel-Rogers (1996) data set consists of observations from U.S. and Canadian cities only. There is no distinction between the empirically-important border dummy (unity for city pairs lying across the border) and a measure of nominal exchange rate variability, since all cross-border pairs have the same nominal exchange rate. In the European data set we have city price data from several countries, and hence are able to include both a border dummy variable and a measure of nominal exchange rate variability in a regression explaining the variability of (common-currency) prices across cities. This allows us to assess the role of local currency pricing and variable nominal exchange rates directly rather than having to rely on the indirect methods used in Engel and Rogers (1996). Our results indicate that most of the failures of the law of one price are attributable to local-currency pricing, but transportation costs and other barriers are also important explanatory factors. We find that, even taking into account nominal exchange rate variability, the border continues to have a positive and significant effect on real exchange rate variability. However, these effects are small compared to the local currency pricing effect.

The other purpose of our model is to explore the welfare implications of our empirical findings. Devereux and Engel (1998) explore this issue extensively and methodically, but here we focus on a particular misconception about fixed exchange rates. When nominal prices are sticky in consumers' currencies, then nominal exchange rate fluctuations can lead to severe misalignment of goods prices between countries. There is an inefficiency since consumers in different countries pay different prices for the same good. One may be tempted to infer from our regressions that fixed exchange rates are more desirable than floating rates on the grounds that price deviations across borders are minimized, thus minimizing this inefficiency. This logic is not entirely correct. As in Devereux and Engel, under floating exchange rates and local-currency pricing, there is an economic inefficiency that results from consumers paying different prices in the two locations, even when transportation costs are zero. The law of one price holds if exchange rates are fixed and transport costs are zero, so that inefficiency is eliminated under fixed exchange rates. Even so, fixed exchange rates do not necessarily increase welfare because fixed rates increase aggregate risk by increasing swings in world income. Thus, even if the welfare losses from failures of the law of one price were eliminated by fixing exchange rates, this is not necessarily welfare-improving. The inefficiency (and hence the welfare loss) is due to price stickiness, and the choice of the exchange rate regime does not necessarily alter that. This analysis is undertaken in section IV.

The paper attempts to move beyond the above-mentioned research that made the important contribution of documenting a "border" effect. Our main point is that there are two very distinct types of border effects embedded in relative prices. One is a sort of "real barriers effect", caused by various barriers to trade. It is analogous to the border effects in the trade volume literature [e.g., McCallum (1995) and Helliwell (1996)]. This effect appears to be empirically significant in our European data set. The second is a sticky-consumer-price cum volatile exchange-rate effect. It is even more important empirically. For the first effect, it is clear that the larger the border effect the larger the deadweight welfare loss. The second effect is not so clear. While the existence of a border effect of the second type does imply economic distortions, it is not necessarily true, comparing two border effects across two pairs of countries, that a larger border effect of this type implies a larger deadweight welfare loss.

I. Data

We use consumer price indexes from 55 locations in Europe.[2] The data are monthly, covering the period March 1981 to July 1997. Table 1 lists the locations. Notice that these span eleven countries. In four cases, Germany, Italy, Spain, and Switzerland, this includes price data from different locations within the country. The German data were obtained from the statistical offices of the individual Lander (state). The Italian data come from regular publications of the official government statistical agency ISTAT. For Spain, the data are on-line in the monthly statistical bulletin of Spain's national statistical institute INE, while the Federal Office for Statistics provided the data for Switzerland. We also use the CPI for Paris, which we obtained from the database of the Bank for International Settlements. For the other six countries--Austria, Belgium, Denmark, Luxembourg, Netherlands, and Portugal--we were unable to obtain data from cities within the country. We rely on the national CPIs in those cases. These data were taken from the IMFs International Financial Statistics database, as were the nominal exchange rates used in the study.

Using price indexes from these 55 locations, we construct 1485 (=55 x 54/2) bilateral relative prices. In addition, our sample of eleven countries implies that the cross-border location pairs lie across one of 55 (=11 x 10/2) national borders (that are not necessarily adjacent).

There are a variety of arrangements determining the nominal exchange rates of our 55 country-pairs. Belgium and Luxembourg share a common currency. Germany was at the heart of the Exchange Rate Mechanism (ERM) of the European Monetary System, and so had a formal policy of fixed exchange rates with France, Austria, Belgium, Denmark, Luxembourg, and the Netherlands during the sample. Each of these countries was included in the first wave of entrants into the Euro, launched just after our sample ends. In contrast, Italy was a member of the ERM until 1992 (and later re-joined), while Spain and Portugal joined the ERM in 1994. There is also some variety in countries' participation in the free-trade area, the European Union (EU). Switzerland has remained out of any formal arrangements on either exchange rates or trade, while Spain and Portugal are relatively recent

[2] Price data at a less aggregated level is not available on a comparable basis across these locations. Engel (1999) shows that essentially all of the variability in relative CPIs across countries is due to variability in traded-goods prices across countries, rather than variability in the ratio of non-traded to traded goods prices within countries.

members of the EU, joining in the mid-eighties. All other countries were members at the beginning of our sample in 1981.

Related to these differences in formal participation in exchange rate arrangements are the different inflation performance of these countries. As seen from Table 1, average annual inflation rates range from 13 percent in Portugal, nearly 8 percent in Italy and Spain, to 3 percent or under in Germany, the Netherlands, and Switzerland. In spite of these cross-country differences in the period averages, the inflation rates of all countries converged noticeably over the sample, so that by 1996 inflation was less than 4 percent everywhere.

The eleven countries used in this study also differ along geographic, linguistic, and cultural lines. Germany, lying at the geographic center of Europe, shares a common border with seven of the other countries. On the other hand, Portugal and Denmark lie at the extremes, and share a common border with only one other country in the sample. We take explicit account of geographic considerations such as common border, as well as physical distance between locations. Finally, although languages are common to some extent across all countries, French and German are spoken in a relatively large number of countries in the sample, while Portuguese is rather specific to Portugal. These factors contribute, in varying degrees, to the openness of the countries. The final column of Table 1 displays one measure of openness, the imports-to-GDP ratio. These range from above ninety percent for Luxembourg to the low twenties for Italy, France, and Spain.

Relative Prices: Summary Statistics

Denote the log of the CPI in location j relative to that in location k as $P(j,k)$. All prices are denominated in U.S. dollars. We are interested in explaining the volatility of changes in $P(j,k)$. We consider one-month changes in relative prices, $\Delta P(j,k)$, but also examine the robustness of our results to using 12-month differences. We measure volatility as the variance.[3]

We construct a measure of volatility for each of our 1485 location-pairs. The analysis is then based on the cross-section of 1485 volatility measures. Table 2 presents summary statistics. The first row reports averages for all pairs of locations. In the rows that follow, we report averages for pairs of locations that are (*i*) both within Germany, (*ii*) both within Italy, (*iii*) both within Spain, (*iv*) both within Switzerland, (*v*) both within the same country (labeled intra-national), and (*vi*) one in one country and one in a foreign country (labeled inter-national).

The first column of Table 2 reports the average variance of $\Delta P(j,k)$, for three different horizons, 1-month, 12-months, and 48-months. For 1-month changes, the average volatility of cross-border pairs is 2.76, more than sixteen times larger than the average variance of within-country pairs, 0.17. The within-Germany pairs exhibit the lowest average volatility, followed in order by those of Switzerland, Italy, and Spain. But even the volatility of relative price changes across Spanish cities, equal to 0.23, is considerably lower than the average volatility of the cross-border pairs. At longer horizons, we observe a similar pattern: cross-border relative price changes are more volatile that intra-national price changes by at least an order of magnitude.

[3] In past work, we have looked at the robustness to using root mean squared errors as an alternative measure of volatility. Because the drifts in relative price changes are all near zero, the results are essentially identical to those using variances.

The next column reports the volatility of nominal exchange rate changes, $\Delta s(j,k)$, again for horizons of 1-months, 12-months, and 48 months. For cross-border pairs, nominal exchange rate variability is practically identical to relative price variability, a result that is well established in the literature. For example, at the 48-month horizon the average volatility of cross-border relative price changes is 159.8, while the average volatility of nominal exchange rate changes is 159.0. As seen in the next column, which reports the average distance between locations, we also see that cross-border pairs are typically more distant than the within-country pairs. In order to sort out the relative influence of these factors, we examine the regression evidence.

Regression Specifications

In the rest of the paper we present the results of our attempt to explain $V(\Delta P(j,k))$, the variance of $\Delta P(j,k)$. We estimate regressions of the form:

$$V(\Delta P(j,k)) = \sum \alpha(m)D(m) + \beta_1 r(j,k) + \beta_2 r^2(j,k) + X\theta + u(j,k)$$

where $D(m)$ is a dummy variable for each city in our sample, $r(j,k)$ is the distance between cities j and k, and X is a vector of explanatory variables that differ across different specifications. Our approach is not standard, in that we do not attempt to explain real exchange rate movements as arising from real demand and supply factors that alter relative prices of different goods (such as the price of tradable to non-traded goods.) Our earlier empirical studies (Engel (1993, 1999), Rogers and Jenkins (1995) and Engel and Rogers (1996)) strongly demonstrate the predominance of deviations from the law of one price in accounting for movements in real exchange rates, at least among industrialized countries. So, in our model, time-varying deviations from the law of one price drive real exchange rate movements. We abstract from any of the traditional explanations that rely on changes in relative supply and demand for goods and services. Before describing the results, we show in the next section how to derive this regression equation from a theoretical model of local currency pricing.

II. A Model of Deviations from the Law of One Price

We present a simple extension of the model of Devereux and Engel (1998). The representative consumer in the home country is assumed to maximize

$$U_t = E_t\left(\sum_{s=t}^{\infty} \beta^{s-t} u_s\right), \qquad 0 < \beta < 1 \tag{1}$$

where

$$u_s = \frac{1}{1-\rho} C_s^{1-\rho} + \chi \ln\left(M_s \big/ P_s \right) - \eta L_s, \qquad \rho > 0$$

C is a consumption index that is a geometric average of home and foreign consumption:

$$C = \frac{C_h^n C_f^{1-n}}{n^n (1-n)^{1-n}}.$$

We assume that there are n identical individuals in the home country, $0 < n < 1$. C_h and C_f are Dixit-Stiglitz constant-elasticity-of–substitution indexes over consumption of goods produced at home and in the foreign country, respectively.

$$C_h = \left[n^{-1/\lambda} \int_0^n C_h(i)^{\lambda-1/\lambda} di \right]^{\lambda/\lambda-1}; \qquad C_f = \left[(1-n)^{-1/\lambda} \int_n^1 C_f(i)^{\lambda-1/\lambda} di \right]^{\lambda/\lambda-1}$$

The elasticity of substitution between goods produced within a country is λ, which we assume to be greater than 1. There is a unit elasticity of substitution between the home goods and foreign goods indexes. M/P are domestic real balances, and L is the labor supply of the representative home agent.[4]

The price index that appears in the utility function deflating nominal balances is defined by

$$P = P_h^n P_f^{1-n}$$

where

$$P_h = \left[\frac{1}{n} \int_0^n P_h(i)^{1-\lambda} di \right]^{1/1-\lambda}, \qquad P_f = \left[\frac{1}{1-n} \int_n^1 P_f(i)^{1-\lambda} di \right]^{1/1-\lambda}.$$

There are $1-n$ identical individuals in the foreign country, with preferences similar to home country residents. The terms in the utility function involving consumption are identical in the home and foreign countries. The functional form for real balances and labor are the same as for home country residents, but for foreign residents they are functions of foreign real balances and foreign labor supply.

We assume complete asset markets. Specifically, residents of each country can purchase state-contingent nominal bonds. The government increases the money supply with direct transfers. There is no government spending or taxes. The Appendix gives details of the money supply processes.

Firms are monopolistic. Firm i in the home country sells $nC_{ht}(i)$ of its product to home-country residents in period t and $(1-n)C_{ht}^*(i)$ to the foreign consumers. We assume that each unit of output sold to home residents requires $1+k_t(i)$ units of labor, where $k_t(i)$ is an *i.i.d.* random variable, with mean $\bar{k}$ (which is the same for all *i*). $k_t(i)$ is distributed independently of $k_t(j)$, $i \neq j$. One

[4] This utility function is the special case of that examined in Devereux and Engel (1998). This special case is useful because it allows closed-form solutions of the model with no approximations. It leads to some implications in this model that are not entirely satisfying: nominal interest rates are constant in equilibrium, and wages are equalized across countries. But, the model is simplified enough that it allows us to explain some of the welfare issues much more clearly.

unit of output sold to foreigners requires $1+k_t^*(i)$ units of labor to produce, where $k_t^*(i)$ is also an *i.i.d.* random variable, with mean $\bar{k}^*$, and $k_t^*(i)$ is distributed independently of $k_t^*(j)$, $i \neq j$.

We can think of $k_t(i)$ and $k_t^*(i)$ as incorporating "iceberg" transportation costs. Under that interpretation, it takes one unit of labor to produce goods for sale to either home or foreign residents. But, only a fraction $\frac{1}{1+k_t(i)}$ of the good arrives when it is shipped to home country residents, and only a fraction $\frac{1}{1+k_t^*(i)}$ arrives when the good is shipped abroad. We assume that the mean cost of shipping abroad is greater than the mean cost of shipping domestically, so $\bar{k}^* > \bar{k}$. The prices that consumers pay are for the delivered good, so the producer bears the cost of shipping (which, of course, is incorporated in the price.)

We assume that prices must be set one period in advance. Firms set a price in domestic currency terms for home residents and in foreign currency for foreign consumers. Firms face a competitive labor market, and are wage-takers. Firm managers maximize the expected utility of firm owners. We assume that domestic residents own domestic firms. Obstfeld and Rogoff (1998) show that, given there is no real intertemporal dimension to the firms' decisions, the optimization problem reduces to choosing $P_{ht}(i)$ and $P_{ht}^*(i)$ to maximize expected profits next period discounted by a utility-based discount factor:

$$E_{t-1}\left[\frac{\beta C_t^{-\rho} P_{t-1}}{C_{t-1}^{-\rho} P_t}\left[nC_{ht}(i)\left(P_{ht}(i)-W_t(1+k_t(i))\right)+(1-n)C_{ht}^*(i)\left(S_t P_{ht}^*(i)-W_t(1+k_t^*(i))\right)\right]\right]. \qquad (2)$$

Foreign firms face a similar optimization problem, choosing $P_{ft}(i)$ and $P_{ft}^*(i)$ to maximize:

$$E_{t-1}\left[\frac{\beta C_t^{*-\rho} P_{t-1}^*}{C_{t-1}^{*-\rho} P_t^*}\left[nC_{ft}(i)\left(\frac{P_{ft}(i)}{S_t}-W_t^*(1+k_t^*(i))\right)+(1-n)C_{ft}^*(i)\left(P_{ht}^*(i)-W_t^*(1+k_t(i))\right)\right]\right]. \qquad (3)$$

In equilibrium, there is an optimal risk-sharing condition:

$$\frac{S_t P_t^*}{P_t} = \left(\frac{C_t}{C_t^*}\right)^{\rho}. \qquad (4)$$

The Appendix derives the solutions to the optimization problems of consumers and managers, and solves for equilibrium values of the endogenous variables. From equation (A.22) and (A.23):

$$p_t = m_{t-1} + n\kappa_{t-1} + (1-n)\kappa_{t-1}^* + \frac{1}{\rho}\sigma_m^2 + \ln\left(\frac{(1-\mu\beta)\lambda\eta}{\mu\chi(\lambda-1)}\right), \qquad (5)$$

$$p_t^* = m_{t-1}^* + n\kappa_{t-1}^* + (1-n)\kappa_{t-1} + \frac{1}{\rho}\sigma_{m^*}^2 + \ln\left(\frac{(1-\mu^*\beta)\lambda\eta}{\mu^*\chi(\lambda-1)}\right). \tag{6}$$

where

$$\kappa_{t-1} = \ln(1+E_{t-1}(k_t(i))), \qquad \kappa_{t-1}^* = \ln(1+E_{t-1}(k_t^*(i))).$$

Lower-case letters represent the natural logs of their upper-case counterparts. Note that κ_{t-1} and κ_{t-1}^* are assumed to be independent of the specific firm. We assume $\kappa_{t-1}^* > \kappa_{t-1}$, so the home country price index will be higher the greater the proportion of goods is imported.

As Obstfeld and Rogoff (1996, p. 582-583; 1998, p. 38-40) have shown, when the money supply follows a random walk and real balances enter the utility function logarithmically (as we have assumed), the nominal interest rate will be constant. In this model, an increase in the home money supply will lower real interest rates but increase expected inflation. The two effects exactly offset, so the nominal interest rate does not change. With nominal interest rates constant, and uncovered interest parity holding up to a constant, the expected change in the log of the exchange rate must be a constant. The log of the exchange rate follows a random walk. As discussed in the Appendix, we assume that money supplies also follow a random walk; thus, money is expected to be neutral after prices adjust in one period. So, a one-percent increase in the home money supply today is expected to lead to a one-percent depreciation of the currency next period. But, given the random walk behavior of exchange rates, this implies the current exchange rate rises one percent. There is no overshooting of exchange rates. Equation (A.5) in the appendix gives us:

$$s_t = m_t - m_t^* + \ln\left(\frac{1-\mu\beta}{1-\mu^*\beta}\right) \tag{7}$$

From equations (5)-(7) we can derive the following expression for the real exchange rate:

$$q_t \equiv s_t + p_t^* - p_t = s_t - s_{t-1} + (1-2n)(\kappa_{t-1} - \kappa_{t-1}^*) + \frac{1}{\rho}(\sigma_{m^*}^2 - \sigma_m^2) + \ln(\mu/\mu^*). \tag{8}$$

The variance of the change in the real exchange rate is given by:

$$Var(\Delta q_t) = 2Var(\Delta s_t) + (1-2n)^2 Var(\Delta(\kappa_{t-1} - \kappa_{t-1}^*)), \tag{9}$$

where Δ is the first-difference operator.

The effect of the variance of the nominal exchange rate on the real exchange rate is magnified by a factor of two. This occurs because prices adjust fully in one period. The variance of the real exchange rate in period *t* is affected one-for-one by the variance of the nominal exchange rate in period *t* because prices are preset (in period *t-1*) and do not respond to exchange-rate shocks. But, prices in period *t* adjust completely to period *t-1* monetary shocks, so period *t-1* exchange rate changes are reflected one-for-one in period *t* prices. In the real world, prices adjust more slowly. So, we would not expect the real exchange-rate variance to magnify the nominal exchange-rate variance so much.

Actual changes in transportation costs do not affect prices, only changes in the conditional expectations of transportation costs. That is because prices are preset, so producers cannot adjust prices in response to shocks to transport costs. In equation (9), if the countries are of equal size ($n = \frac{1}{2}$), these costs do not appear at all. The greater the share of the consumer basket that is imported, the higher the price of that basket because imported goods incorporate higher transportation costs. But when countries are of equal size, these costs affect price levels equally in both countries.

III. Regressions

As noted above, we estimate regressions of the form:

$$V(\Delta P(j,k)) = \sum \alpha(m)D(m) + \beta_1 r(j,k) + \beta_2 r^2(j,k) + X\theta + u(j,k) \qquad (10)$$

where $D(m)$ is a dummy variable for each city in our sample, $r(j,k)$ is the distance between cities j and k, and X is a vector of explanatory variables that differ across different specifications. Note that all regressions are cross-sectional, with 1485 observations. The inclusion of separate dummies for each individual location allows the variance of price changes to vary from city to city. One reason for this comes from Table 2, which indicates somewhat higher average volatility for Spanish and Italian cities than German or Swiss cities. This could be because Spain, for example, is a relatively heterogenous country. Spain's labor markets or goods markets may be less integrated than those of Switzerland, so there can be greater discrepancies in prices between cities in Spain. Aggregate supply shocks may vary more across cities in Spain and Italy than Switzerland. Alternatively, there may be differences across countries in methodologies for gathering price data that lead to greater discrepancies in prices between locations in one country than another.

We use a couple of measures to proxy for the variation in expected transport costs. First, we hypothesize that this variance will be larger the greater the distance between locations. As in the gravity model of trade, we postulate a concave relationship between distance and relative price volatility, and so expect that $\beta_1 > 0$ and $\beta_2 < 0$.[5] We interpret "transportation costs" liberally to include any factors that make it more costly to sell in one location compared to another. For example, there may be trade barriers between countries. A large portion of the cost that consumers pay is for the marketing and distribution of the good. Producers may find it more costly to market in locations that are distant from where the production takes place and corporate decisions are made. Likewise, distribution networks may be more difficult to organize and monitor in distant locales.

The variables included in X differ across specifications. We are particularly interested in whether there is a border effect. That is, even taking into account distance between two locations, and

[5] We also ran regressions with a logarithmic, rather than quadratic, functional form for distance. We found log(*distance*) to be positive and significant, as we have in our earlier work.

the nominal exchange rate variance, are there other factors that cause prices in cities in different countries to vary? For reasons outlined in the model above, and considering the results of Engel and Rogers (1996), we expect the variability of prices between cities that lie across a border to be higher than those between cities within a country, even after accounting for the effect of distance. Therefore, we include in X a dummy variable, Border, that takes on a value of unity if cities j and k are in different countries. This dummy variable may capture some of the formal and informal international barriers to trade. We typically find Border to be positive and significant.

One reason the dummy variable Border might be important is because of local currency pricing by producers, as shown above. To test this, we include in X the variance of nominal exchange rate changes between locations j and k, $V(\Delta s(j,k))$. We stress the usefulness of including this variable in addition to the Border dummy. This allows us to assess the role of exchange rate variability / sticky prices *directly*. In Engel and Rogers (1996) we had to rely on indirect attempts to assess this channel. This is because in our earlier data set, where we had observations from U.S. and Canadian cities only, there is no distinction between the Border dummy variable and nominal exchange rate variability, since all cross-border pairs have the same nominal exchange rate. Including both regressors would be redundant in that data set. In the current case we have city price data from several countries, and hence are able to include both Border and $V(\Delta s(j,k))$.

There are of course reasons other than sticky prices why the border matters. Although there are no formal barriers to trade among most countries in the sample for most of the period (Switzerland is the exception), there may be informal barriers. Furthermore, marketing and distribution networks may be less homogenous across borders than within countries, perhaps in part because a different language is spoken in the foreign country. We consider how much of the variance of relative prices continues to be explained by the Border dummy once we have included nominal exchange rate volatility in the regression, as well as controls for language and a common border. Finally, we measure the importance of each of the several borders in our sample, by including in X individual border dummies (e.g., Germany-Italy, Germany-Spain, etc.).[6] The relative sizes of these may provide clues as to why the border effect is large, and allows us to go beyond what we were able to do in Engel and Rogers (1996).

Regression Results

The first column of Table 3A presents the results of regressing the variance of the 1-month change in the log relative price on Distance, Distance squared, Border, and 55 individual location dummies (one for each of our 49 cities and for Austria, Belgium, Denmark, Luxembourg, Netherlands, and Portugal). All coefficients have the anticipated sign and are significant at the 5 percent level. The coefficient on the Border dummy is 2.85 with a standard error of 0.06. The interpretation of this coefficient is the difference between the average variance of relative prices for city pairs that lie across a border less the average for pairs that lie within a country, taking into account the effect of distance and city-specific characteristics. The positive and significant estimated Border effect confirms the results for the U.S. and Canada documented by Engel and Rogers (1996): crossing an international border adds considerable volatility to relative city prices, even after accounting for the effects of distance and city-specific characteristics.

---

[6] This necessitates dropping both Border and $V(\Delta(s(j,k)))$.

There are many possible explanations for the strong border effect, one of which emphasizes the role of local currency pricing in a world of variable nominal exchange rates. Specification 2 adds to the regression the variability of one-month nominal exchange rate changes, which of course is zero for all intra-national pairs. An implication of the theoretical model above, which assumes complete local currency pricing and full price adjustment after one period, is that this coefficient should be 2.00.[7] In the real world prices adjust more slowly than one period (month), of course, leading us to expect the coefficient to be less than 2.0, but with complete local currency pricing it will still be greater than 1.00. For our sample, the coefficient on nominal exchange rate variability is 0.92. The fact that this is significantly less than unity (the standard error is 0.005) suggests that there is not complete local currency pricing (irrespective of the length of nominal price stickiness), as local prices respond somewhat to exchange rate changes.

Although nominal exchange rate variability does not magnify real exchange rate variability by as much as our model of local currency pricing implies, the addition of $V(\Delta s(j,k))$ substantially weakens the effect of the Border dummy, whose point estimate falls from 2.85 in specification 1 to 0.21 in specification 2. This suggests that a very large part of the border effect is from variable nominal exchange rates under sticky prices. However, even with $V(\Delta s(j,k))$ in the regression, the Border dummy remains positive and significant with a t-statistic exceeding 10. In specification 3, we add two dummies variables, *Adjacent*, which is unity if there is a common border separating the city pair and *Language*, which is unity if the cities lie in different countries that speak a common language. We hypothesize that each of these dummies will be negatively related to relative price variability. As seen in Table 3A, the adjacency dummy is negative and significant, while Language is insignificant. There is very little effect on the rest of the estimated coefficient.

Specification 4 replaces the variance of the nominal exchange rate and the Border, Adjacency, and Language dummies with individual dummies for each border (Germany-Italy, Germany-Spain, ..., Netherlands-Portugal), of which there are 55 in our sample.[8] Distance and distance squared remain of the expected sign and are significant in this specification. We also report the 5 largest and smallest of the 55 estimated individual border dummies. The smallest are Belgium-Luxembourg, Germany-Austria, Germany-Netherlands, Austria-Netherlands, and Denmark-Luxembourg. The largest are Italy-Switzerland, Spain-Switzerland, Switzerland-Portugal, Italy-Portugal, and Italy-Luxembourg. Notice that the largest border coefficient, 3.98 for Italy-Switzerland, is nearly 40 times larger than the smallest (0.11 for Belgium-Luxembourg).

What explains the relative sizes of these individual border effects? Nominal exchange rate variability certainly comes to mind, but there are other possibilities. Even in our pre-Euro sample period, Belgium and Luxembourg shared a common currency and so have had *zero* nominal exchange rate variability. The currencies of Germany, Austria, and the Netherlands were tied together relatively

[7] Under the assumptions of complete local currency pricing and full price adjustment after one period, nominal exchange rates, s_t, and relative prices $(p - p^*)$ are uncorrelated but have equal unconditional variances. Thus, in a regression explaining the cross-sectional variation in the volatility of real exchange rates, $(s - p + p^*)$, the coefficient on the volatility of nominal exchange rates exceeds unity.

[8] With $n=11$ countries, there are $[n^*(n-1)/2]=55$ individual border dummies.

closely. This may not be everything, however. According to Table 1, Belgium and Luxembourg are the two most open countries in our sample, as measured by imports-to-GDP ratio. Also, in the latter three countries, German is a commonly-spoken language; these three are also likely to experience similar supply shocks due to their geographic proximity and the similarity of their economic bases. At the other extreme, all five of the largest border dummies involve Italy and/or Switzerland. Switzerland is the only country in our sample that is not a member of the free-trade EU area. In addition, Switzerland did not participate in the exchange rate mechanism of the EMS (nor is it a prospective EMU member) and so has experienced relatively large fluctuations in the value of its currency.

Figure 1 presents a scatter plot of the estimated border coefficients against the variance of the nominal exchange rate. The tight, positive relationship that is evident from the plot confirms our earlier evidence on the importance of nominal exchange rate variability.

Figure 2 plots the border coefficients against a frequently-used measure of openness to trade, the average imports-to-GDP ratio of the two countries. Table 1 lists this ratio for each country in the sample. Note that this measure only partially coincides with *a priori* beliefs about the degree of barriers to bilateral trade. For example, trade between France and Italy, two original members of the European Union, is by all accounts more open than trade between Portugal and Switzerland. But, the trade openness measure for the France-Italy pair is .21, which is less than the value of .365 for Switzerland-Portugal. Because things such as geographical factors are important in determining a country's import/GDP ratio above and beyond the influence of formal trade barriers, one should think of the horizontal axis in Figure 2 as reflecting how open the economies are on average.[9] Ocular inspection reveals some evidence of a negative relationship between the size of the border effect and the imports/GDP ratio, but one that is not nearly as tight as in Figure 1.

The three specifications in Table 3B repeat the regressions of specifications 1, 3 and 4, but using 48-month changes instead of one-month changes in relative prices and exchange rates. The effects of the Border dummy and nominal exchange rate variability are comparable to those reported in the earlier specifications. However, distance becomes slightly less significant. For example, in specification 2 distance is insignificantly different from zero in the case of the 48-month changes, but has a t-statistic of around five in the one-month case. This result is interesting because it suggests that the effects of transportation costs, and not those of local currency pricing, tend to diminish at longer horizons (to the extent that distance proxies for transportation costs).

Analysis of the Sub-Periods

The results so far were all computed over the full sample period, 1981:3-1997:7. Table 4 presents summary statistics for three different sub-periods: 1981:3-86:12, 1987:12-92:7, and 1992:11-97:7, and regression results for the 1987:12-92:7 sub-period. The sub-periods correspond roughly to the changing fortunes of the ERM. The so-called Hard-EMS period began in 1987 after some

[9] There are alternatives to using the imports-to-GDP ratio in our calculation of trade openness. Direct measures of trade barriers are periodically compiled by the United Nations Commission on Trade and Development (UNCTAD). These country-wide measures are constructed using, e.g., the average tariff rate or the percentage of goods which are subjected to some type of non-tariff barrier. Also, Harrigan (1996) presents model-based estimates of openness in manufacturing trade among OECD nations. The fact that Switzerland is not included in Harrigan's group of countries makes his measures unusable for us.

turbulence in foreign exchange markets in the early to mid-1980s. This period witnessed no re-alignments of official parities. However, the general calm in the foreign exchange market was broken abruptly in September 1992, when the U.K. and Italy withdrew from the ERM.

The first panel of Table 4 presents the average $V(\Delta P(j,k))$ and $V(\Delta s(j,k))$ for the three sub-periods and the full sample. The variability of relative price changes for within-country pairs falls noticeably across the sub-periods. The average for all intra-national pairs, for example, falls from 0.22 in the pre-1987 sub-period to 0.16 in the Hard-EMS period, to 0.07 post-1992. This decline is evidence of increased economic integration within countries, which is likely to be caused by advancements in transportation, communication, etc. Despite the apparent increased integration within countries, there is no tendency toward increased convergence of *inter*-national relative prices, as indicated by the average variances of 2.30, 1.06, and 3.65 for the three sub-periods. Instead, this pattern follows that of nominal exchange rate variability, as displayed in the final columns of the panel. The average variance of the nominal exchange rate change for our three sub-periods is 2.18, 0.71, and 3.62. This prompts the question: are the factors leading to increased intra-national integration not at work at the international level or is it simply that greater exchange rate variability masks what is otherwise a tendency toward increased international integration?

We present some evidence on this in the bottom panel of Table 4, which presents regression results for the Hard-EMS sub-period. We choose this period since it is associated with significantly lower variability in nominal exchange rates. We repeat specifications 1, 2, and 4 from Table 3A. Distance is notably smaller and less significant than in the full sample. The estimated Border coefficient is still positive and significant, but is only about one-third its size of the full sample regressions. This suggests that the forces leading to greater market integration of different regions within countries are at work to an even greater extent between countries. Interestingly, the coefficient on $V(\Delta s(j,k))$ is significantly greater than 1.00, at 1.24. Finally, the individual Border effects shown in specification 3 are considerably smaller than in the full sample.

IV. Welfare implications of exchange-rate arrangements

The derivation of the welfare expressions in the Appendix (equation (A.20)) does not rely on any assumption about the correlation of home money shocks with foreign money shocks. Expected utility depends on the variance of home and foreign money shocks, but not on their correlation. One implication is that if we compare utility under floating in the symmetric case (in which the variance of home and foreign money supplies are equal) to utility under fixed exchange rates (in which the home money supply is perfectly correlated with the foreign money supply) the welfare comparison depends only on the degree of monetary variance under the two systems. If money supply shocks have the same variance under fixed and floating exchange rates, then welfare is equal under the two exchange-rate systems.

This result definitely depends on some special assumptions we have made here. Our utility function is a special case of the more general formulation in Devereux and Engel (1998). That paper showed that floating rates dominate fixed rates in welfare terms when monetary variances are equal in the two regimes. Van Wincoop and Bacchetta (1999) have even more general welfare functions in a two-period model of local-currency pricing, and find the welfare comparisons can go in favor of either fixed exchange rates or floating rates.

But this simple model is perfect for illustrating a point we wish to emphasize. Under floating exchange rates and local-currency pricing, there is clearly an economic inefficiency because consumers are paying different prices in the two locations even when transportation costs are zero. The law of one price holds if exchange rates are fixed and transport costs are zero, so that inefficiency is eliminated under fixed rates. Even so, fixed exchange rates do not necessarily increase welfare.

Intuitively, under floating exchange rates expected utility is diminished because of the pricing inefficiency. Since residents in different locations face different prices under some states, there is idiosyncratic risk in equilibrium. The risk-sharing condition, equation (4), is identical to the condition derived by Backus and Smith (1993) in their model with non-traded goods. Aggregate consumption levels are not equalized in all states because consumers face different aggregate price levels. Under fixed exchange rates in this model (with no transport costs), the law of one price holds and idiosyncratic risk is eliminated in equilibrium. Equation (4) shows that when the law of one price (and, hence, purchasing power parity) holds, consumption is equal at home and abroad in all states. But, fixed exchange rates increase aggregate risk. Monetary policies at home and abroad are perfectly correlated under fixed exchange rates. Since, in sticky-price models, swings in the money supply affect output, swings in world income are enhanced under fixed exchange rates. The trade-off, then, is that fixed exchange rates eliminate idiosyncratic risk at the cost of increasing aggregate risk. In this model, those two effects exactly cancel (if the monetary variances are the same under fixed and floating), so welfare is the same under the two exchange-rate systems.

As we discussed above, in equilibrium, nominal interest rates are constant. From equation (A.12), we get

$$\rho c_t = m_t - p_t + \ln(1 - \mu\beta/\chi) . \tag{11}$$

We note from equation (5) that the price level depends not only on the level of the money supply and the expected cost shocks, but also on the variance of the money supply. As the recent literature has explained (see Bacchetta and van Wincoop (1998), Obstfeld and Rogoff (1998) and Devereux and Engel (1998)), the variance of the money supply affects the level of prices through the risk aversion of firm-owners. Firm managers must take into account the risk aversion of firm owners in setting prices. Monetary variance increases the variance of consumption in sticky-price models, and so enters the firm's pricing decisions. However, the price level does not depend on the correlation of the money shocks, and so is the same under fixed and floating exchange rates. The mean and variance of home consumption is related only to the mean and variance of home money supply, and similarly for foreign consumption and foreign money. The independence of real consumption from the exchange-rate regime recalls Baxter and Stockman's (1989) empirical study that showed that the volatility of many real economic variables (including consumption) do not depend on the volatility of real exchange rates.

Suppose money (and therefore consumption) in each country could take on only two values, high (H) and low (L). Under floating exchange rates, let us suppose there are four possible states, each equally likely. State 1 is H-H (that is high home consumption, high foreign consumption). State 2 is H-L; state 3 is L-H; and, state 4 is L-L. Given that the exchange rate is just proportional to the money supplies (equation (11)), in states 1 and 4 the law of one price holds even though prices are preset. But in states 2 and 3, there are deviations from the law of one price. In those states, a planner could reallocate goods to improve worldwide welfare in the Pareto sense.

Now, suppose under fixed exchange rates the home country alters its money supply to equal the foreign money supply in all states. So, state 1 is H-H; state 2 is L-L; state 3 is H-H; and, state 4 is L-L. There is no change in the allocations in states 1 and 4. But now in states 2 and 3 the money supplies and consumption levels are equal, and the law of one price holds. Has welfare been improved? Clearly not (at least from looking at the consumption part of the utility function.) The mean and variance of consumption in both countries is exactly the same as under floating exchange rates. There is no pricing inefficiency under fixed exchange rates, but there is greater variance in world consumption.

The choice of monetary regime affects the amount of aggregate risk in the economy. While fixed exchange rates eliminate the pricing inefficiency (in other words, eliminate the idiosyncratic risk in states 2 and 3), they increase the variance of aggregate average world consumption. The underlying inefficiency here is the stickiness of prices. That inefficiency cannot be eliminated by the choice of exchange-rate system.

While the Appendix derives the general expression for welfare in this model, it is convenient to examine the special case in which $n = \frac{1}{2}$ and $\sigma_m^2 = \sigma_{m^*}^2 \equiv \sigma^2$. Equation (A.21) evaluates expected utility as:

$$U = \frac{1+\rho(\lambda-1)}{\lambda(1-\rho)} \exp\left(\frac{1-\rho}{2\rho} \left(2\ln\left(\frac{\lambda-1}{\lambda\eta} \right) - \kappa - \kappa^* - \frac{1}{\rho}\sigma^2 \right) \right), \qquad (12)$$

where

$$\kappa = \ln(1+\bar{k}), \qquad \kappa^* = \ln(1+\bar{k}^*),$$

and $\bar{k}$ and $\bar{k}^*$ are the unconditional means of domestic and international transport costs, respectively.

Equation (12) shows that higher transport costs lower welfare. We expect transport costs between locations to be greater than transport costs within a location, so $\kappa^* > \kappa$. When transport costs are greater, the deviations from the law of one price will be greater and there will be lower utility.

However, deviations from the law of one price that occur because prices are preset in consumers' currencies and exchange rates respond to current monetary shocks do not necessarily translate directly into welfare losses. Equation (12) shows that welfare is lower the greater is the variance of money supplies. But this expression holds for any value of the correlation of home and foreign moneys. Hence, it is monetary variability, not exchange-rate variability (and the implied deviations from the law of one price) that lower welfare. It is easy to show that welfare under flexible prices is given by equation (12) with $\sigma^2 = 0$. In this model, the welfare loss from sticky prices is directly related to the variance of money shocks.

This has direct implications for the interpretation of our empirical results. We find that the large border effect can be primarily attributed to the volatility of nominal exchange rates. If the border effect declines because exchange rates become permanently fixed, that does not necessarily imply that welfare will increase. The only clear-cut welfare improvement that accompanies reductions in law of one price deviations comes from the reduction of real barriers to trade.

It is fair to say that one of the principle motivations for forming the EMU and moving to a single currency is that member countries hoped that some discipline might be imposed on their own monetary policies. If indeed the EMU succeeds in reducing overall monetary variance, then it will have a beneficial welfare effect.

V. Conclusions

We explore failures of the law of one price across European cities, using consumer price data from European cities over the period 1981-97. We attempt to move beyond a "first-generation" of research papers (by us and others) that empirically document very large border effects. There are two very distinct types of border effects embedded in relative prices. The first type, a "real barriers effect" caused by various barriers to trade, is analogous to the border effects in the literature on trade volumes. The second is a sticky-consumer-price cum volatile exchange-rate effect. Both effects are shown to be important empirically, the second type especially so. For the first type of effect, the larger is the border effect the larger is the deadweight welfare loss. But for the second type of effect the issue is not so clear. While the existence of a border effect of the second type does imply economic distortions, it is not necessarily true, comparing two border effects across two pairs of countries, that a larger border effect of this type implies a larger deadweight welfare loss.

We establish this general point through four particular contributions. First is our empirical attempt to identify the *sources* of the large border effects in international comparisons of consumer prices. Because of our unique data set on consumer prices from 55 locations across 11 European countries, we are able to fairly conclusively identify that most of the border effects are coming from local currency pricing with fluctuating nominal exchange rates. Our results thus provide support for the focus on local currency pricing taken in several recent theoretical papers examining the welfare implications of different exchange rate regimes.

Second, we find that even taking into account nominal exchange rate variability (as well as language and geographic considerations), the border still has a positive and significant effect on real exchange rate variability. However, these effects are small compared to the local currency pricing effect. We speculate that the remaining border effect reflects factors like cross-country differences in the national marketing and distribution systems. In addition, we confirm earlier finding that distance and city-specific factors are also significant.

The third contribution of the paper is to show how to derive our regression equation for the variance of relative prices in a dynamic optimizing model. This equation is similar to the one estimated in Engel and Rogers (1996). We use a version of the model in Devereux and Engel (1998) for this purpose.

Finally, we link our empirical findings to Devereux and Engel's (1998) welfare results on fixed versus floating exchange rates. In the model, failures of the law of one price imply welfare losses due to an inefficiency associated with consumers in different locations paying different prices for the same good. Our empirical work attributes most of the failure of the law of one price to local currency pricing and floating exchange rates. But even if failures of the law of one price were eliminated through fixed exchange rates, it does not necessarily eliminate the inefficiency. Fixed exchange rates would not eliminate the welfare loss, even though the welfare loss would no longer come from failures of the law of one price. Under fixed rates, the law of one price holds, but the welfare loss would come from volatility of consumption. It is worth emphasizing that the welfare loss (the inefficiency) is due to price stickiness, and the choice of the exchange rate regime does not necessarily alter that.

References

Bacchetta, Philippe and Eric van Wincoop, 1998, Does exchange rate stability increase trade and capital flows?, Federal Reserve Bank of New York, working paper no. 9818.

Bacchetta, Philippe and Eric van Wincoop, 1999, Does exchange rate stability increase trade and welfare?, Federal Reserve Bank of New York, mimeo.

Backus, David K., and Gregor W. Smith, 1993, Consumption and real exchange rates in dynamic economies with non-traded goods, Journal of International Economics 35, 297-316.

Baxter, Marianne, and Alan C. Stockman, 1989, Business cycles and the exchange-rate regime: Some international evidence, Journal of Monetary Economics 23, 377-400.

Devereux, Michael B. and Charles Engel, 1998, Fixed vs. floating exchange rates: How price setting affects the optimal choice of exchange-rate regime, National Bureau of Economic Research, working paper no. 6867.

Engel, Charles, 1993, Real exchange rates and relative prices: An empirical investigation, Journal of Monetary Economics 32, 35-50.

Engel, Charles, 1998, A retrial in the case against the EMU, manuscript, University of Washington.

Engel, Charles, 1999, Accounting for U.S. real exchange rate changes, Journal of Political Economy 107, 507-538.

Engel, Charles and John H. Rogers, 1996, How wide is the border?, American Economic Review 86, 1112-1125.

Feenstra, Robert and Jon Kendall, 1997, Pass-through of exchange rates and purchasing power parity, Journal of International Economics 43, 237-261.

Giovannini, Alberto, 1988, Exchange rates and traded goods prices, Journal of International Economics 24, 45-68.

Harrigan, James, 1996, Openness to trade in manufactures in the OECD, Journal of International Economics 40, 23-40.

Helliwell, John, 1996, Do national borders matter for Quebec's trade?, Canadian Journal of Economics 29, 507-522.

Hess, Gregory and Eric van Wincoop, eds., 1999, Intra-national macroeconomics, Cambridge University Press.

Isard, Peter, 1977, How far can we push the law of one price?, American Economic Review 67, 942-948.

Knetter, Michael, 1993, International comparisons of pricing to market behavior, American Economic Review 83, 473-486.

Marston, Richard, 1990, Pricing to market in Japanese manufacturing, Journal of International Economics 29, 217-236.

McCallum, John, 1995, National borders matter: regional trade patterns in North America, American Economic Review 85, 615-623.

Obstfeld, Maurice and Kenneth Rogoff, 1996, Foundations of International Macroeconomics (Cambridge: MIT Press).

Obstfeld, Maurice and Kenneth Rogoff, 1998, Risk and exchange rates, National Bureau of Economic Research, working paper no. 6694.

Rogoff, Kenneth, 1996, The purchasing power parity puzzle, Journal of Economic Literature 34, 647-668.

Rogers, John H. and Michael Jenkins, 1995, Haircuts or hysteresis? Sources of movements in real exchange rates, Journal of International Economics 38, 339-360.

Tille, Cedric, 1998, The international and domestic welfare effects of monetary policy under pricing-to-market, working paper, Federal Reserve Bank of New York.

Appendix

The optimal intratemporal consumption choices take on simple forms:

$$C_h(i) = \frac{1}{n}\left[\frac{P_h(i)}{P_h}\right]^{-\lambda} C_h, \qquad C_f(i) = \frac{1}{1-n}\left[\frac{P_f(i)}{P_f}\right]^{-\lambda} C_f,$$

$$P_h C_h = nPC, \qquad P_f C_f = (1-n)PC,$$

and,

$$\int_0^n P_h(i)C_h(i)di = P_h C_h, \qquad \int_n^1 P_f(i)C_f(i)di = P_f C_f.$$

Money demand is given by

$$\frac{M_t}{P_t} = \frac{\chi C_t^{\rho}}{1-d_t}, \tag{A.1}$$

where d_t, the inverse of the gross nominal interest rate, is given by:

$$d_t = E_t\left(\beta \frac{C_{t+1}^{-\rho} P_t}{C_t^{-\rho} P_{t+1}}\right). \tag{A.2}$$

The optimal tradeoff between consumption and leisure is given by:

$$W_t = \eta P_t C_t^{\rho}. \tag{A.3}$$

Obstfeld and Rogoff (1996, p. 582-583; 1998, p. 38-40) show that equations (A.1) and (A.2), combined with the random walk assumption on money supplies given in equation (2) imply nominal interest rates are constant. As they explain, monetary expansion at home causes home real interest rates to fall, but causes expected inflation to rise. When money balances enter the utility function logarithmically, the two effects exactly cancel out. We find:

$$C_t^{\rho} = \left(\frac{1-\mu\beta}{\chi}\right)\frac{M_t}{P_t}. \tag{A.4}$$

A set of equations analogous to (A.1)-(A.4) hold in the foreign country.

Equation (A.4) and its foreign counterpart, along with the risk-sharing condition (6) imply:

$$S_t = \frac{M_t(1-\mu\beta)}{M_t^*(1-\mu^*\beta)}. \tag{A.5}$$

Under floating exchange rates, we assume that in each country the log of the money supply follows a random walk. In the home country

$$m_{t+1} - m_t = -\ln(\mu) + \tfrac{1}{2}\sigma_m^2 + \upsilon_{t+1},$$

where υ_t is an independent and identically distributed normal random variable with mean of zero and variance of σ_m^2. This equation implies

$$E_t\left({M_t}/{M_{t+1}}\right) = \mu .$$

A similar set of equations describe the foreign money supply process under floating exchange rates. The foreign money supply shock, υ_t^* is assumed to be independent of υ_t when exchange rates are flexible. We will assume under fixed exchange rates that monetary policies in the two countries are perfectly harmonized, so as to keep the exchange rate fixed at unity. This means υ_t^* is assumed to have a correlation of one with υ_t.

Now we turn to the firms' optimization problem. From the intratemporal consumption choices, we have:

$$C_{ht}(i) = (P_{ht}(i))^{-\lambda} P_{ht}^{\lambda-1} P_t C_t,$$

and

$$C_{ht}^*(i) = (P_{ht}^*(i))^{-\lambda} P_{ht}^{*\lambda-1} P_t^* C_t^* .$$

Because the production function is linear, the firm's problem can be separated into the problem of choosing $P_{ht}(i)$ to maximize the discounted profits from sales domestically and choosing $P_{ht}^*(i)$ to maximize discounted profits from foreign sales. Substituting the above equations into the firm's objective function, and recalling that time *t* prices are in the time *t-1* information set, the firm's objective in choosing $P_{ht}(i)$ is to maximize:

$$E_{t-1}\left(C_t^{1-\rho}\left((P_{ht}(i))^{1-\lambda} - W_t(P_{ht}(i))^{-\lambda}(1+k_t(i))\right)\right).$$

To choose $P_{ht}^*(i)$, it maximizes:

$$E_{t-1}\left(C_t^{-\rho}C_t^*\left(S_t(P_{ht}^*(i))^{1-\lambda} - W_t(P_{ht}^*(i))^{-\lambda}(1+k_t^*(i))\right)\right).$$

The optimal prices are given by:

$$P_{ht}(i)=\frac{\lambda}{\lambda-1}\frac{E_{t-1}\left(C_t^{1-\rho}W_t(1+k_t(i))\right)}{E_{t-1}\left(C_t^{1-\rho}\right)},$$

and,

$$P_{ht}^{*}=\frac{\lambda}{\lambda-1}\frac{E_{t-1}\left(C_t^{-\rho}C_t^{*}W_t(1+k_t^{*}(i))\right)}{E_{t-1}\left(C_t^{-\rho}C_t^{*}S_t\right)}. \tag{A.6}$$

Equation (A.4) shows that shocks to C_t arise only from monetary shocks, which are assumed to be independent of $d_t(i)$. Using equation (A.3) for wages, we can rewrite the domestic price as:

$$P_{ht}(i)=P_{ht}=\frac{\lambda\eta}{\lambda-1}P_t\frac{(1+\bar{k})E_{t-1}(C_t)}{E_{t-1}\left(C_t^{1-\rho}\right)}. \tag{A.7}$$

Similarly, using the analog to equation (A.4) for foreign wages, and using the risk-sharing condition (6), we can rewrite the expression for foreign price of the home good as:

$$P_{ht}^{*}(i)=P_{ht}^{*}=\frac{\lambda\eta}{\lambda-1}P_t^{*}\frac{(1+\bar{k}^{*})E_{t-1}\left(C_t^{*}\right)}{E_{t-1}\left(C_t^{*1-\rho}\right)}. \tag{A.8}$$

Analogously, foreign firms set prices as:

$$P_{ft}(i)=P_{ft}=\frac{\lambda\eta}{\lambda-1}P_t\frac{(1+\bar{k}^{*})E_{t-1}(C_t)}{E_{t-1}\left(C_t^{1-\rho}\right)}, \tag{A.9}$$

and,

$$P_{ft}^{*}(i)=P_{ft}^{*}=\frac{\lambda\eta}{\lambda-1}P_t^{*}\frac{(1+\bar{k})E_{t-1}\left(C_t^{*}\right)}{E_{t-1}\left(C_t^{*1-\rho}\right)}. \tag{A.10}$$

Note that $P_{ft}=P_{ht}\frac{(1+\bar{k}^{*})}{(1+\bar{k})}$ and $P_{ht}^{*}=P_{ft}^{*}\frac{(1+\bar{k}^{*})}{(1+\bar{k})}$. We can therefore write

$$P_t=P_{ht}^{n}P_{ft}^{1-n}=P_{ht}\left(\frac{(1+\bar{k}^{*})}{(1+\bar{k})}\right)^{1-n}=P_t\frac{\lambda\eta}{\lambda-1}\frac{E_{t-1}(C_t)}{E_{t-1}\left(C_t^{1-\rho}\right)}(1+\bar{k})^{n}(1+\bar{k}^{*})^{1-n}.$$

This implies

$$E_{t-1}\left(C_t^{1-\rho}\right)=\frac{\lambda\eta}{\lambda-1}E_{t-1}\left(C_t\right)(1+\bar{k})^n(1+\bar{k}^*)^{1-n}. \quad \text{(A.11)}$$

Now note that from equation (A.4) we have

$$\rho c_t = m_t - p_t + \ln(1-\mu\beta/\chi), \quad \text{(A.12)}$$

where lower-case letters are logs of their upper-case counterparts. Since money is log-normally distributed, consumption is as well. Then we can write equation (A.11) as

$$(1-\rho)E_{t-1}(c_t)+\frac{(1-\rho)^2}{2}\sigma_c^2 = n\kappa+(1-n)\kappa^* + \ln\left(\frac{\lambda\eta}{\lambda-1}\right)+E_{t-1}(c_t)+\frac{1}{2}\sigma_c^2,$$

where

$$\kappa=\ln(1+\bar{k}), \qquad \kappa^*=\ln(1+\bar{k}^*).$$

From equation (A.12), $\sigma_c^2=\frac{1}{\rho^2}\sigma_m^2$, so we can solve out:

$$E_{t-1}(c_t)=\frac{-n}{\rho}\kappa-\frac{1-n}{\rho}\kappa^*-\frac{1}{\rho}\ln\left(\frac{\lambda\eta}{\lambda-1}\right)+\frac{\rho-2}{2\rho^2}\sigma_m^2. \quad \text{(A.13)}$$

In levels,

$$E(C)=(1+\bar{k})^{-n/\rho}(1+\bar{k}^*)^{n-1/\rho}\left(\frac{\lambda-1}{\lambda\eta}\right)^{1/\rho}\exp\left(\frac{\rho-1}{2\rho^2}\sigma_m^2\right). \quad \text{(A.14)}$$

Using equation (A.11) we can then write

$$\frac{1}{1-\rho}E(C^{1-\rho})=\left(\frac{1}{1-\rho}\right)V_h, \quad \text{(A.15)}$$

where

$$V_h\equiv\left(\frac{\lambda-1}{\eta\lambda}\right)^{1-\rho/\rho}(1+\bar{k})^{n(\rho-1)/\rho}(1+\bar{k}^*)^{(1-n)(\rho-1)/\rho}\left(\exp\left(\frac{\rho-1}{2\rho^2}\sigma_m^2\right)\right) \quad \text{(A.16)}$$

Analogously, we can derive an expression for expected foreign consumption to be used below:

$$E(C^{*}) = (1+\bar{k})^{\frac{n-1}{\rho}}(1+\bar{k}^{*})^{\frac{-n}{\rho}}\left(\frac{\lambda-1}{\lambda\eta}\right)^{\frac{1}{\rho}} \exp\left(\frac{\rho-1}{2\rho^{2}}\sigma^{2}_{m^{*}}\right). \tag{A.17}$$

and,

$$\frac{1}{1-\rho}E(C^{*1-\rho}) = \left(\frac{1}{1-\rho}\right)V_{f}$$

where

$$V_{f} \equiv \left(\frac{\lambda-1}{\eta\lambda}\right)^{\frac{1-\rho}{\rho}}(1+\bar{k})^{\frac{(1-n)(\rho-1)}{\rho}}(1+\bar{k}^{*})^{\frac{n(\rho-1)}{\rho}}\left(\exp\left(\frac{\rho-1}{2\rho^{2}}\sigma^{2}_{m^{*}}\right)\right). \tag{A.18}$$

Turning to the labor market equilibrium condition, note that $P_{ht}(i) = P_{ht}$ implies that $nC_{ht}(i) = C_{ht}$ and, similarly, $nC^{*}_{ht}(i) = C^{*}_{ht}$. Equilibrium in the goods and labor market in the home country requires:

$$nL_{t} = \int_{0}^{n} nC_{ht}(i)(1+k_{t}(i))di + \int_{0}^{n}(1-n)C^{*}_{ht}(i)(1+k^{*}_{t}(i))di\,. \tag{A.19}$$

and in the foreign country:

$$(1-n)L_{t} = \int_{0}^{1-n} nC_{ft}(i)(1+k^{*}_{t}(i))di + \int_{0}^{1-n}(1-n)C^{*}_{ft}(i)(1+k_{t}(i))di\,. \tag{A.19*}$$

Equation (A.19) can be written:

$$\begin{aligned} nL_{t} &= C_{ht}\int_{0}^{n}(1+k_{t}(i))di + \tfrac{1-n}{n}C^{*}_{ht}\int_{0}^{n}(1+k^{*}_{t}(i))di \\ &= n\frac{P_{t}C_{t}}{P_{ht}}\int_{0}^{n}(1+k_{t}(i))di + (1-n)\frac{P^{*}_{t}C^{*}_{t}}{P^{*}_{ht}}\int_{0}^{n}(1+k^{*}_{t}(i))di \end{aligned}$$

Taking expectations, we can write:

$$-\eta E(L) = -\eta\left[n\left(\frac{(1+\bar{k}^{*})}{(1+\bar{k})}\right)^{1-n}E(C_{t})(1+\bar{k}) + (1-n)\left(\frac{(1+\bar{k})}{(1+\bar{k}^{*})}\right)^{1-n}E(C^{*}_{t})(1+\bar{k}^{*})\right].$$

Using equations (A.14) and (A.17) and the definitions in (A.16) and (A.18), we derive:

$$-\eta E(L) = -\left(\frac{\lambda-1}{\lambda}\right)(nV_h + (1-n)V_f).$$

which can be used along with equation (A.15) to derive utility in the home country:

$$U = \frac{1}{1-\rho}V_h - \left(\frac{\lambda-1}{\lambda}\right)(nV_h + (1-n)V_f). \tag{A.20}$$

Note that the derivation of utility did not depend on the correlation of domestic and monetary shocks. Thus, utility under fixed exchange rates does not depend on the perfect correlation of domestic and monetary shocks. The only modification to expression (A.20) that is needed to derive utility under fixed exchange rate is that the simplification that $\sigma_m^2 = \sigma_{m^*}^2$ can be used.

We can further note that in the case where $n = \frac{1}{2}$ and $\sigma_m^2 = \sigma_{m^*}^2 \equiv \sigma^2$, then $V_h = V_f$ and the expression for utility simplifies to:

$$U = \frac{1+\rho(\lambda-1)}{\lambda(1-\rho)}\exp\left(\frac{1-\rho}{2\rho}\left(2\ln\left(\frac{\lambda-1}{\lambda\eta}\right) - \kappa - \kappa^* - \frac{1}{\rho}\sigma^2\right)\right). \tag{A.21}$$

We shall also make use of the solution for goods prices. Since P_t is in the time *t-1* information set, using equations (A.12) and (A.13) we have

$$\begin{aligned} p_t &= E_{t-1}\left(m_t - \rho c_t + \ln\left(\frac{1-\mu\beta}{\chi}\right)\right) \\ &= m_{t-1} + n\kappa_{t-1} + (1-n)\kappa_{t-1}^* + \frac{1}{\rho}\sigma_m^2 + \ln\left(\frac{(1-\mu\beta)\lambda\eta}{\mu\chi(\lambda-1)}\right) \end{aligned} \tag{A.22}$$

In writing equation (A.22) we have allowed for a slight generalization – that the shocks to transportation costs have a conditional mean that is not constant. So, in equation (A.22) we are using the following notation:

$$\kappa_{t-1} = \ln(1 + E_{t-1}(k_t(i))), \qquad \kappa_{t-1}^* = \ln(1 + E_{t-1}(k_t^*(i))),$$

where we continue to assume that the means are identical across firms. (We can still interpret $\kappa = \ln(1+\bar{k})$ and $\kappa^* = \ln(1+\bar{k}^*)$ as unconditional means. So, allowing time-varying conditional means in transportation costs does not alter the unconditional expectation of utility given in equations (A.20) or (A.21).) Analogously, for the foreign country, we have:

$$p_t^* = m_{t-1}^* + n\kappa_{t-1}^* + (1-n)\kappa_{t-1} + \frac{1}{\rho}\sigma_{m^*}^2 + \ln\left(\frac{(1-\mu^*\beta)\lambda\eta}{\mu^*\chi(\lambda-1)}\right). \tag{A.23}$$

Table 1 -- Locations of CPI Data

Country:	Location of Price Index used		Avg. Inflation	Imports/GDP
Germany	Baden-Wurttemberg (Stuttgart) Bavaria (Munich) Berlin	Hesse (Frankfurt) Nordrein-Westf. (Dusseldorf) Saarland (Saarbrucken)	2.85	0.27
Italy	Ancona Aosta Bari Bologna Cagliari Campobasso Florence Genoa L'Aquila Milan	Naples Palermo Perugia Potenza Reggio Calabria Rome Turin Trento Trieste Venice	7.90	0.20
Spain	Andalucia (Seville) Aragon (Zaragoza) Baleares (Palma) Canarias Cantabria (Santanda) Castilla la Mancha (Toledo) Castilla y Leon (Leon) Cataluna (Barcelona) Ceuta y Melilla (Ceuta)	Comunid. Valencia (Valencia) Extremadura (Badajoz) Galicia (La Coruna) La Rioja (Logrono) Madrid Murcia Navarra (Pamplona) Pais Vasco (Vizcaya) Principado Asturias (Gijon)	7.69	0.21
Switzerland	Basle Bern	Geneva Zurich	3.17	0.34
France	Paris		4.80	0.22
Austria	Austria (Vienna)		3.33	0.38
Belgium	Belgium (Brussels)		3.76	0.66
Denmark	Denmark (Copenhagen)		4.46	0.32
Luxembourg	Luxembourg		3.75	0.92
Netherlands	Netherlands (The Hague)		2.52	0.50
Portugal	Portugal (Lisbon)		13.3	0.39

Notes: The table lists the locations for our data on consumer price indexes. In those cases in which prices are taken from a state (as in Germany), region (as in Spain), or from the country as a whole, the city listed in parenthesis is that used to calculate distances. All data are monthly from 1981:3-1997:7.
Inflation rates and import-to-GDP ratios are annual averages over the period 1981-96.

Table 2; Summary Statistics

Pairs:	Variance $\Delta P(j,k)$			Variance $\Delta s(j,k)$			Distance	# obs.
	1-mo.	12-mo.	48-mo.	1-mo.	12-mo.	48-mo.		
All	2.12	39.7	121.3	1.97	40.1	120.0	653	1485
GE-GE	0.02	0.21	1.24	0.00	0.00	0.00	209	15
IT-IT	0.14	0.99	2.80	0.00	0.00	0.00	263	190
SP-SP	0.23	1.01	3.08	0.00	0.00	0.00	417	153
SW-SW	0.07	0.37	1.56	0.00	0.00	0.00	80	6
intra-national	0.17	0.96	2.83	0.00	0.00	0.00	322	364
inter-national	2.76	52.3	159.8	2.62	53.1	159.0	760	1121

Notes: Columns display the mean values of the variance of changes (both 1-month, 12-month, and 48-month changes) in the relative price between location j and k, $\Delta P(j,k)$, and the change in the nominal exchange rate $\Delta s(j,k)$, 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 city pairs in all 55 locations; GE-GE indicates that only the within-Germany city pairs are used; IT-IT, SP-SP, and SW-SW are the analogies for the Italian, Spanish, and Swiss cities; intra-national indicates that only pairs of cities within countries are used; and inter-national indicates that only cross-border pairs are used in the calculations. The data are monthly from 1981:3-1997:7.

Table 3A; Regressions Explaining Relative Price Variability

Specification:	1	2	3	4
Distance[1]	5.20*	1.05*	0.88*	0.81*
	(1.28)	(0.21)	(0.21)	(0.21)
Distance Squared[2]	-3.12*	-0.52*	-0.49*	-0.58*
	(0.59)	(0.11)	(0.11)	(0.11)
Border	2.85*	0.21*	0.23*	---
	(0.06)	(0.02)	(0.02)	
Var (Δs(j,k))	---	0.92*	0.92*	---
		(0.005)	(0.009)	
Adjacent	---	---	-0.03*	---
			(0.009)	
Language	---	---	0.006	---
			(0.02)	
Five smallest Border coeffs. (point estimates)	---	---	---	BE-LU (0.11*) GE-AU (0.18*) GE-NE (0.19*) AU-NE (0.38*) DE-LU (0.53*)
Five largest Border coeffs. (point estimates)	---	---	---	IT-SW (3.98*) SP-SW (3.90*) SW-PO (3.89*) IT-PO (2.91*) IT-LU (2.87*)
Adj. R^2	.89	.99	.99	.99

Notes: The sample period is 1981:3-1997:7. There are 1485 observations in each regression. Heteroscedasticity-consistent standard errors are in parenthesis. The dependent variable is the variance of the 1-month change in the log relative price. The independent variables are:

Distance = the distance between cities in the particular pair (in miles);
Var (Δs(j,k)) = the variance of the 1-month change in the nominal exchange rate;
Border = unity if the cities in the pair lie across an international border;
Adjacent = unity if there is a common border separating the city pair;
Language = unity if the cities lie in different countries that speak a common language.

Specifications 1-3 also include all 55 individual location dummies (one for each of our 49 cities and one for Austria, Belgium, Denmark, Luxembourg, Netherlands, and Portugal). Specification 4 includes all border dummies as well as all 55 individual location dummies. Mnemonics used for the border dummies are from the following country list: GE=Germany, SP=Spain, IT=Italy, SW=Switzerland, FR=France, AU=Austria, BE=Belgium, DE=Denmark, LU=Luxembourg, NE=Netherlands, and PO=Portugal. Thus, in specification 4, for example, the dummy variable IT-SW is unity for location pairs that lie across the Italian-Swiss border and zero for all other location pairs. A (*) indicates the coefficient is significant at the 5% level.

(1) coefficients and standard error have been multiplied by 10^4; (2) the coefficient and standard error have been multiplied by 10^7;

Table 3B; Regressions Explaining Relative Price Variability – 48-Month Changes

Specification:	1	2	3
Distance[1]	7.58*	0.46	-1.25*
	(2.15)	(1.01)	(0.33)
Distance Squared[2]	-3.26*	-2.09*	-0.11
	(1.00)	(0.40)	(0.19)
Border	120.8*	36.8*	---
	(8.18)	(4.34)	
Var (Δs(j,k))	---	0.93*	---
		(0.03)	
Adjacent	---	28.8*	---
		(6.19)	
Language	---	-66.2*	---
		(10.8)	
Five smallest Border coeffs. (point estimates)	---	---	BE-LU (5.02*) BE-FR (5.83*) GE-LU (7.73*) GE-AU (7.93*) FR-LU (8.09*)
Five largest border coeffs. (point estimates)	---	---	SP-SW (289.2*) IT-SW (266.1*) IT-NE (257.1*) SP-NE (256.9*) IT-PO (248.0*)
Adj. R^2	.50	.91	.99

Notes: The sample period is 1981:3-1997:7. There are 1485 observations in each regression. Heteroscedasticity-consistent standard errors are in parenthesis. The dependent variable is the variance of the 48-month change in the log relative price. The independent variables are defined in Table 3.
(*) indicates the coefficient is significant at the 5% level.
(1) coefficients and standard error have been multiplied by 10^2; (2) the coefficient and standard error have been multiplied by 10^5

Table 4; Analysis of the Sub-Periods

A. Summary Statistics

	Variance ΔP(j,k)				Variance Δs(j,k)			
Pairs:	81:3-86:12	87:12-92:7	92:11-97:7	Full	81:3-86:12	87:12-92:7	92:11-97:7	Full
All	1.59	0.77	2.27	1.78	1.24	0.40	2.06	1.25
GE-GE	0.02	0.02	0.03	0.02	0.00	0.00	0.00	0.00
IT-IT	0.18	0.11	0.09	0.13	0.00	0.00	0.00	0.00
SP-SP	0.32	0.28	0.05	0.22	0.00	0.00	0.00	0.00
SW-SW	0.11	0.06	0.03	0.07	0.00	0.00	0.00	0.00
Intra-national	0.22	0.16	0.07	0.16	0.00	0.00	0.00	0.00
Inter-national	2.30	1.06	3.65	2.69	2.18	0.71	3.62	2.52

B. Regressions Explaining Relative Price Variability

Specification:	1	2	3
Distance[1]	3.31*	0.30	0.76*
	(1.00)	(0.30)	(0.26)
Distance Squared[2]	1.08*	0.05	-0.34*
	(0.47)	(0.17)	(0.14)
Border	0.76*	0.07*	---
	(0.05)	(0.02)	
Var (Δs(j,k))	---	1.24*	---
		(0.03)	
Five smallest border coeffs. (point estimates)	---	---	BE-LU (0.09*) GE-BE (0.11*) GE-LU (0.11*) FR-BE (0.11*) GE-NE (0.14*)
Five largest border coeffs. (point estimates)	---	---	SP-SW (2.64*) SW-PO (2.37*) SP-NE (1.50*) SW-AU (1.47*) SW-NE (1.44*)
Adj. R^2	.79	.98	.99

Notes: See notes to Table 3. All results are for one-month changes. The sample period for the regressions of panel B is 1987:12-1992:7.

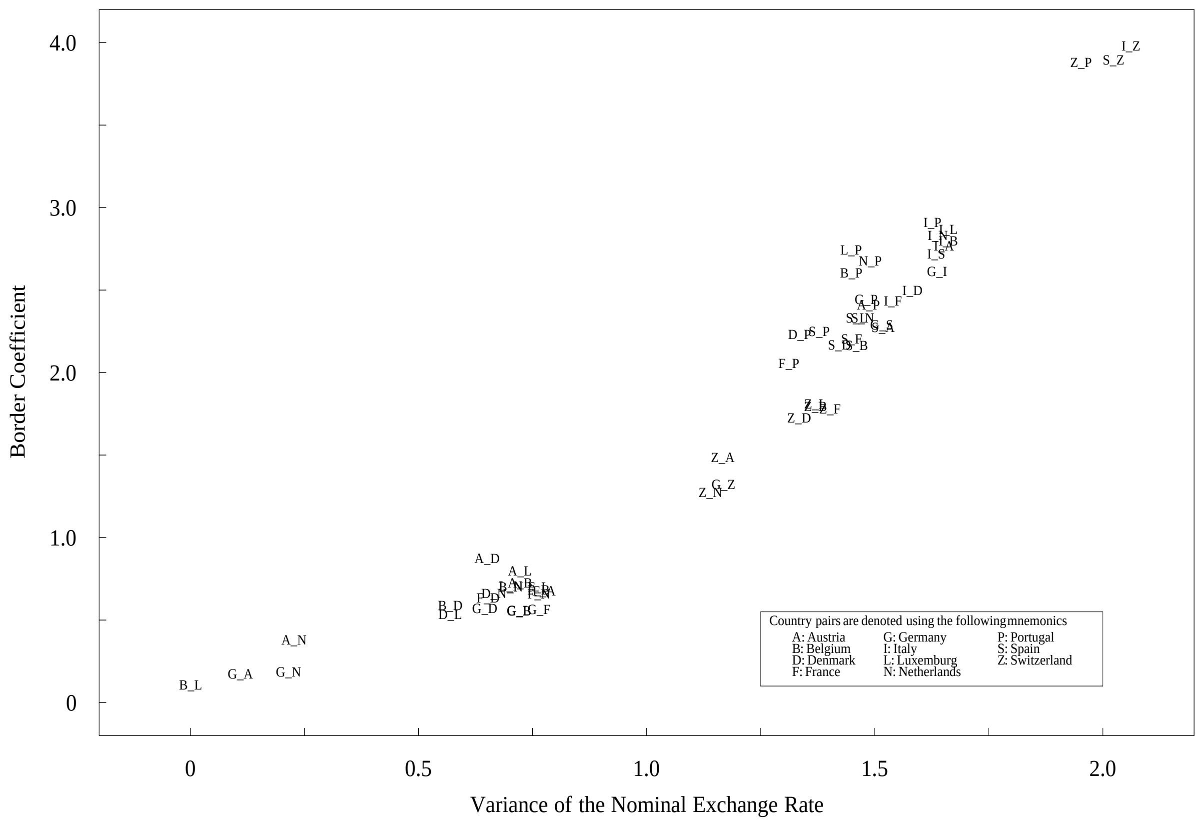

Figure 1: Estimated Border Coefficient vs. Exchange Rate Variability. On the vertical axis is the coefficient estimate on the individual border dummies in specification 3 of Table 3. The horizontal axis depicts the variance of the one-month change in the nominal exchange rate.

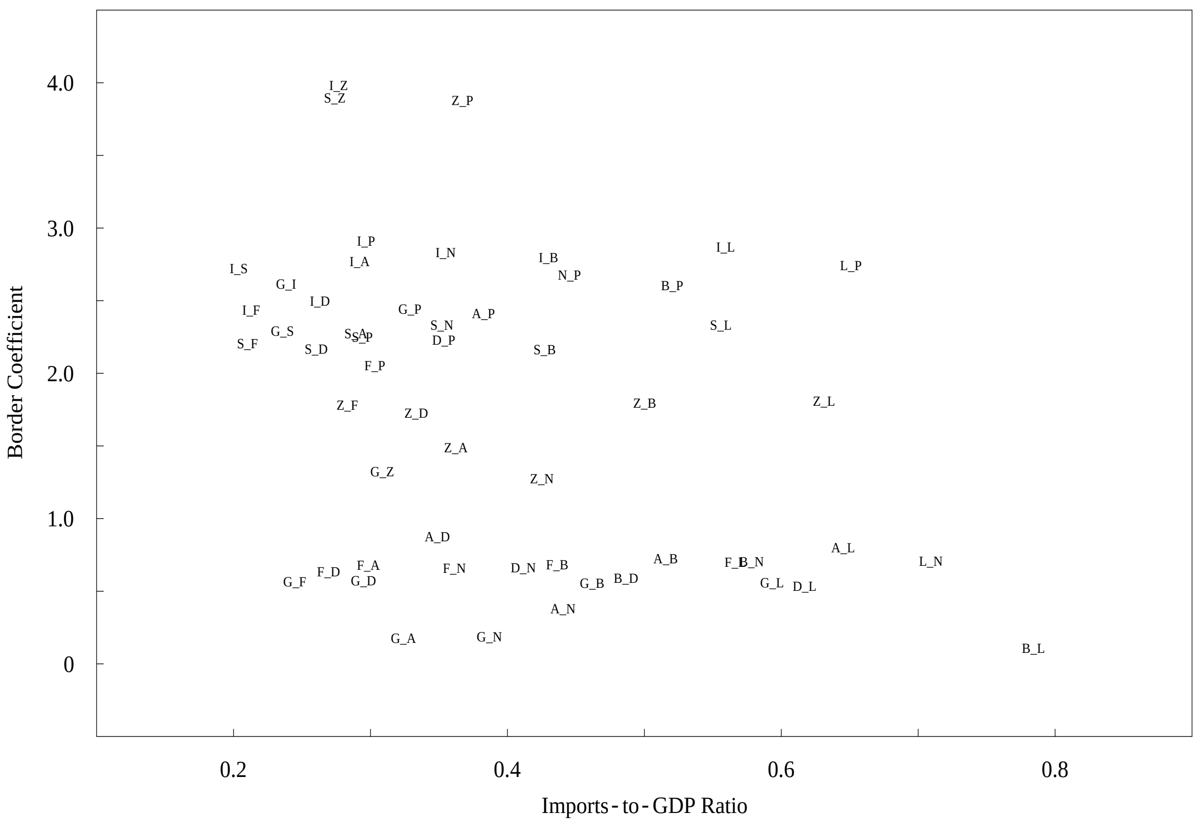

Figure 2: Estimated Border Coefficient vs. Import/GDP Ratio. Along the vertical axis is the coefficient estimate on the individual border dummies in specification 3 of Table 3. The horizontal axis depicts the average ratio of imports to GDP.