Europe Without Borders?
The Effect of the EMU on Relative Prices

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Abstract

Proponents of monetary union argue that establishing a single currency area will decrease market segmentation, lowering relative price volatility. A study by Engel and Rogers (1996) looked at how distance and the presence of national borders affected relative price volatility between cities in the U.S. and Canada, and found that crossing the border is equivalent to at least 1700 miles of distance between cities in the same country. This paper extends this analysis to cities in Western Europe and finds two key results. Both distance and national borders are significant determinants of relative price volatility, indicating that relative prices are significantly more volatile across borders than within European countries. This result is robust to a variety of potential explanations such as uneven sampling bias, idiosyncratic shocks, and incomplete exchange rate pass-through. Turning our attention to cross-border price volatility before and after the formation of the EMU, the effects vary by country size. Within the EMU, cross-border price volatility has not changed between the "small" countries, but has fallen significantly between the large EMU countries. Between the EMU and the UK, cross-border volatility has increased between the UK and the small EMU countries, but there has been no significant change between the UK and the large EMU countries.

Keywords: Law of one price, economic integration, exchange rates, European Monetary Union

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"Anything can be made anywhere on the face of the earth and sold everywhere else on the face of the earth."
– Lester Thurow (1996) in *The Future of Capitalism*

"National borders have effectively disappeared and, along with them, the economic logic that made them useful lines of demarcation in the first place."
– Kenichi Ohmae (1990) in *The Borderless World*

I. Introduction

Much has been made about the increased globalization of the international economy. Free trade zones such as NAFTA and currency unions such as the EMU have purportedly reduced the economic significance of national borders. Consequently, markets have become less segmented, with physical distance between locations becoming less of a factor. But is this really the case? If indeed it is, then we should observe the same good selling for the same price in all locations, the "law of one price."

Many empirical studies have refuted the law of one price, or shown that relative price convergence is glacially slow. Studies documenting the failure of the law of one price include Isard (1977), Giovannini (1988), and Knetter (1989, 1993), who show that prices for even highly tradable and homogenous goods fail to equalize across locations. Other studies have found little evidence that prices between locations even move in the same direction. An implication of the law of one price is that if prices for the same good across locations do move in the same direction, then real exchange rates constructed from these prices should exhibit mean reversion. Rogers and Jenkins (1995) and Jenkins (1997) construct real exchange rates from commodity prices and are unable to reject a unit root in the majority of these cases, implying no mean reversion.[1]

Why is it that such an economically plausible theory does not seem to hold? Two culprits that emerge are market segmentation brought on by physical distance and the

[1] Note that the failure to reject a unit root does not necessarily imply the presence of a unit root. Several studies have argued that the power of unit root tests close to unity is weak. More powerful tests introduced in Elliot, Rothenberg, and Stock (1996) and Elliot and Pesavento (2001) have lent greater support to mean reversion in real exchange rates, albeit at a fairly slow pace.

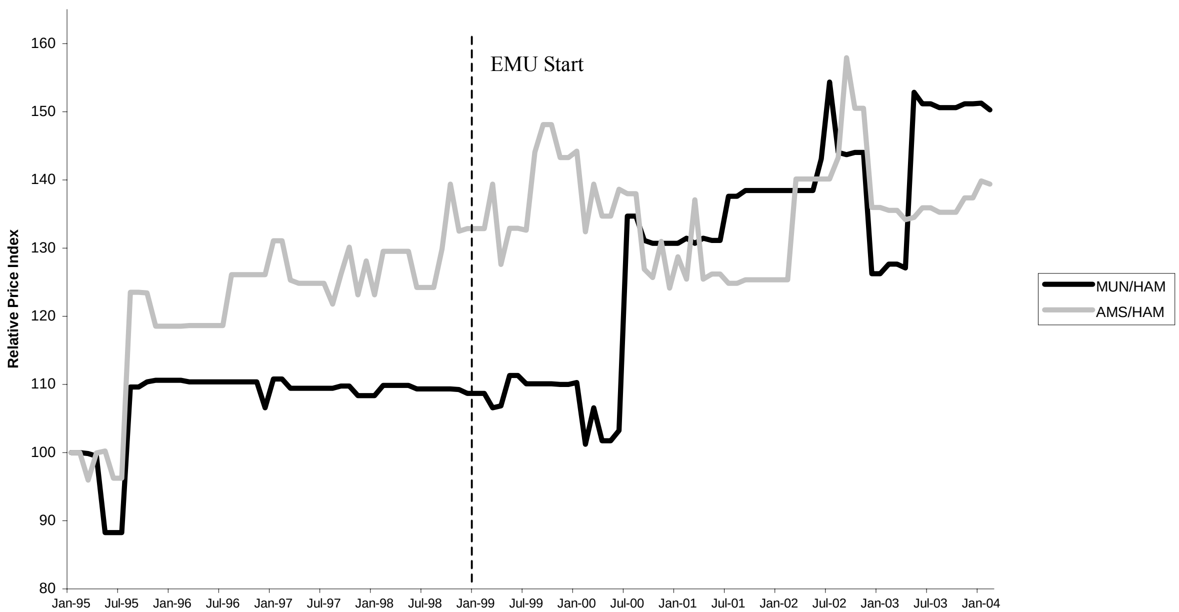
presence of national borders. Distance increases segmentation mainly through transportation costs, which limit the arbitrage opportunities that drive both theories.[2] O'Connell and Wei (1997) provide evidence of limited arbitrage by estimating a threshold model in which the speed of price convergence is greater when initial price differences lie outside some band.[3] Furthermore, markets that are further apart tend to be less similar than those that are closer together, creating different demands and therefore different prices for the same good.

Is there any reason to believe that the observed market segmentation is any higher across countries than within countries? Parsley and Wei (1996) find that convergence rates to the LOP are significantly higher between cities within the US than those found in cross-country studies. National borders may be significant for several reasons. The presence of trade barriers such as tariffs would tend to limit arbitrage in the same way as transportation costs. Another possibility is that residents of one country may have an affinity for domestically produced goods, the "home bias" in trade[4]. Differential tax schemes across countries may also have an effect (the same good may receive a different tax treatment in one country than it does in another.) Finally, there may be national standards that create natural market segmentation. Such examples include right-side steering wheels, 220-volt outlets, or a warranty void outside the country of purchase.

The basic issues examined in this paper are the role of national borders in segmenting markets and how this "border effect" has changed since the inception of the EMU. If borders were irrelevant, we would expect that the only impediment to fully integrated markets would be transportation costs. One way to examine this issue is to look at price levels in given markets. Figure 1 shows a relative price index for two pairs of cities, one pair separated by a border and the other pair within the same country[5]. If the border were not an issue, we would expect that the pair of cities located closer together would exhibit a greater degree of price stability than the pair located further

[2] Davis (1998) estimates transport costs as a percentage of imports in the US to 4.8%, larger than industry level tariffs. Limao and Venables (2000) estimate the elasticity of trade to transport costs –2.5%.
[3] The presence of a threshold model may indicate that LOP convergence rates may be even slower than previously thought. Taylor (2001) finds that linear AR models display an upward bias in estimating the speed of convergence when compared to the true threshold model in Monte Carlo simulations.
[4] An excellent example of this is McCallum (1995).
[5] The relative price is defined as the per diem rate for city 1 divided by the per diem rate for city 2. The index is constructed by using the January 1995 relative price as the base.

apart. In this example, Hamburg is 366km (229 miles) away from Amsterdam and 611km (382 miles) from Munich. Thus, the relative price index between Hamburg and Amsterdam should exhibit at least the same stability as the price index between Hamburg and Munich. As is evident from Figure 1, however, the Amsterdam to Hamburg relative price exhibits much greater volatility than the Munich to Hamburg price. Interestingly, this difference is volatility is reduced when restricting attention to dates after 1999 (and especially following July of 2000), perhaps indicating the impact of monetary union on reducing border effects.



Hamburg (GER) to Amsterdam (NED) = 366 km

Hamburg (GER) to Munich (GER) = 611 km

The seminal work by Engel and Rogers (1996) examined the role of distance and national borders in explaining these phenomena. They examined deviations from the law of one price for 14 categories of goods, looking at the relative prices between nine Canadian and 14 U.S. cities. They find that while both distance and the border are significant determinants of relative price volatility, crossing the border alone has the same effect as over 1,700 miles of distance between cities. Several explanations for this effect emerge, most notably incomplete pass through of exchange rates. Changes in nominal exchange rates are often the result of factors in financial markets and not in goods markets. These exchange rate changes are not completely passed through to nominal prices. Because prices are more rigid than nominal exchange rates, there is a tendency for cross border prices to be more volatile once they have been converted into a common currency. However, filtering out this effect only reduces (but does not eliminate) the border effect.

This paper is similar in spirit to the work by Engel and Rogers (1996). Whereas their original study focused on the U.S. and Canada, my work will examine prices in a large number of cities across Western Europe. The price variable is the daily per diem rate for two categories of goods published monthly by the U.S. State Department for 201 cities in 16 countries. As all prices are quoted in the same currency (U.S. Dollars), the problem of sticky prices/volatile exchange rates is mitigated[6]. For several measures of relative price volatility, both distance and the border are found to be significant determinants. A positive and significant border effect indicates that cross border price volatility is significantly higher than the volatility of prices within the two countries sharing a border. The positive border effects found in this study indicate that markets in Europe are still quite segmented[7].

A significant European border effect is consistent with the results found in Engel and Rogers (2001). Their study examined the distance and border effects (and potential explanations) across cities in Western Europe between 1981 and 1997 for a wide range of goods. My study is complementary to this work, in that I make use of a different data set and test for the effects of monetary union by including more recent data.

Has the adoption of the euro eliminated the border effect? In a world of fully integrated markets, borders remain economically relevant chiefly as lines demarcating currency usage. Given that the adoption of a common currency reduces the transaction costs of cross-border trade (thereby increasing arbitrage opportunities) and eliminates incomplete exchange rate pass through, we should observe a smaller border effect within the EMU. Testing this claim, we find that the effect of the EMU on cross-border price volatility varies by the size of the countries on either side of the border. Within the

[6] This claim will be tested later. While it is true that all prices are quoted in the same currency, this does not necessarily eliminate the problem of incomplete exchange rate pass through. The prices were simply converted to dollars when they were recorded, so there may very well be an exchange rate component of the relative price series.

[7] A recent study by Rogers (2001) supports this finding, although he does find some evidence of price convergence in that the "low price" EU members have seen increased inflation.

EMU, the larger EMU countries have seen a reduction in cross-border price volatility, whereas cross-border volatility for the smaller countries has not significantly changed. Restricting our attention to the border effect between the UK and the EMU countries, there has been an increase in cross-border volatility between the UK and the smaller EMU countries, but no significant change between the UK and the larger EMU countries.

The rest of the paper will proceed as follows. Section II describes the dataset and discusses the strengths and weaknesses of using this data for the analysis. Section III develops an empirical model to test the importance of national borders. The original Engel and Rogers specification is presented and refined to allow for even sampling, idiosyncratic shocks and incomplete exchange rate pass through. Estimation results are given for the original specification and the specification incorporating these refinements. Section IV looks at the impact of the EMU, breaking down the border and distance effects by country pairs. Section V is conclusion.

II. The Data

The data set is based on per diem rates published by the U.S. State Department for employees living abroad[8]. The rates are broken down into two categories, "lodging" and "meals and incidental expenses." The total per diem is the sum of these two rates. The following quote describes how the series is constructed:

> "The maximum per diem rates for foreign counties are based on costs reported in the Hotel and Restaurant Section (Form DSP-23W) submitted by foreign posts. This report includes prices for hotel rooms and meals at facilities representative of moderately prices and suitable hotels and restaurants most frequently used by typical federal travelers. The lodging portion of the allowance is based on average reported costs for a single room, including any mandatory service charges and taxes. Where breakfast or other meals are included in the reported lodging charge, these costs have been adjusted to exclude meals. The meal portion is based on the costs of an average breakfast, lunch, and dinner at facilities typically used by employees at that location, including taxes, service charges, and customary tips. The M&IE rate is based on these meal costs plus an additional amount, equal to 10% of the combined lodging and meal costs, to cover incidental travel expenses."
> (U.S. State Department, Office of Allowances. http://www.state.gov/m/a/als/prdm/)

The price series thus applies to highly non-tradable goods. We would expect to see less price dispersion among more tradable goods. However, these goods do have tradable intermediate components, and the theory being tested is that prices move in the same direction, not that they equalize across locations. The major drawback to using these prices is that any estimates of a border effect will be an upper bound, given the limited arbitrage opportunity. These prices will, however, give a lower bound on the effect of monetary union on reducing the border effect, since these goods are among the last for which we would expect to see price convergence. Thus, we can be more confident about any evidence of euro-zone price convergence when using this dataset than one with more tradable goods.

The sample has monthly observations covering the period 1/95 – 6/02 for 201 cities in 16 countries[9]. Data was collected for cities that had uninterrupted coverage over the entire sample period and any city with seasonal prices for lodging and meals (typically resort destinations) was dropped from the data set. The dependent variable is the relative price, so we are interested in city pairs. In general, if there are N cities, then there are only N-1 independent prices. However, it is necessary to calculate N(N-1)/2 relative price volatilities. This is because we not only need to calculate the relative price variances, but the covariance as well. For example, even after calculating the variance of the London/Paris and the London/Rome relative prices, we still need to calculate the variance of the Paris/Rome relative price, as this is necessary in calculating the covariance of the London/Paris and London/Rome relative prices. Thus, with 201 cities, there are 20,100 city pairs to be estimated. A summary of the data is presented below:

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[8] The choice of this data set over other sources of price data on cities in Europe (such as the Economist Intelligence Unit data on living expenses) was made because of the large number of cities *within* each country with this data. Significant within country price heterogeneity is essential to getting accurate estimates of the border effect. Having a large number of cities comes at a cost, however, limited scope of goods. Given that the emphasis is on price volatility, it seems better to have more cities and fewer goods than fewer cities and more goods.

[9] The sample includes city-pairs that do not share a border. While it is true that there is not a "border" separating Belgium and Spain (there are in fact *two* borders), all of the features that make the border significant are still present. The key point here is that the cities are not in the same country. A complete listing of all cities is given in the appendix.

Table 1: Summary Statistics for same-country city pairs

	# of cities	# of city pairs	Avg. Distance (km)	Avg. St. Dev. (x1000)
Austria	4	6	214.10	30.86
Belgium	15	105	79.98	21.71
Denmark	1	-	-	-
Finland	1	-	-	-
France	16	120	393.08	40.65
Germany	65	2080	285.20	40.59
Ireland	3	3	126.97	46.60
Italy	22	231	347.70	44.33
Luxembourg	1	-	-	-
Netherlands	10	45	54.39	41.82
Norway	1	-	-	-
Portugal	2	1	274.4	70.40
Spain	22	231	588.66	31.61
Sweden	1	-	-	-
Switzerland	3	3	148.97	37.12
United Kingdom	34	561	270.59	32.81
Cross-border pairs	-	16714	952.96	50.32
Total	201	20100	834.72	48.30

The first column lists the number of cities in each country, while the second column indicates the number of city pairs. Thus, there are zero city pairs within Finland, but we can form 200 city pairs with Helsinki. The third column lists the average distance between locations in each country. Distance is calculated by obtaining latitude and longitude coordinates for each location and directly computing the distance between city pairs[10]. The fourth column lists the average standard deviation of the relative prices (given by the daily per diem rate.) Comparing the values for the individual countries to that for just the cross border pairs, it appears that prices are more stable within countries,

[10] This calculation is made by first approximating the shape of the earth as spherical. Then, once the geographic coordinates are converted into radians, the following formula for calculating the surface distance between two points on a sphere:

$$6378 * \arccos\left[\sin(lat1) * \sin(lat2) + \cos(lat1) * \cos(lat2) * \cos(lon2 - lon1)\right]$$

perhaps indicating the importance of borders. However, given that the average distance for all pairs is higher than for the within country pairs, this may also have an influence.

III. The Border Effect

The summary statistics presented in Table 1 indicate that the relative price volatility for cross-border pairs is greater than that for pairs of cities within the same country. Is this difference significant, and how much is due to the fact that cross-border pairs are generally located further apart (implying higher transport costs between these markets) than pairs within these countries? To answer these questions, we must develop a model that can be tested empirically.

For ease of exposition, assume there are only two countries, home (h) and foreign (f). Let the relative price between two cities *i* and *j* that are separated by a border be:

$$p_{ij,t}^{c} = p_{i,t} - p_{j,t} \tag{1}$$

The relative price between two cities *k* and *l* that are within the same country is:

$$p_{kl,t}^{w} = p_{k,t} - p_{l,t} \tag{2}$$

Thus, relative price volatility will be the volatility of $p_{ij,t}^{c}$ or $p_{kl,t}^{w}$, depending on whether or not the pair of cities are separated by a border. Pooling all city pairs together, then the excess volatility of cross-border prices over within country prices is given by:

$$\Phi = AVG[Var(p_{ij,t}^{c})] - AVG[Var(p_{kl,t}^{w})] \tag{3}$$

A large value of Φ indicates that prices across a border are more volatile than prices within the countries on either side of this border. A test of the hypothesis that national borders are irrelevant to market segmentation would be a test of whether or not Φ is different from zero.

This approximation is fairly accurate, generally within 30km of the actual distance. The city center is chosen as the reference point for the latitude and longitude coordinates.

III-A A Comparison to Engel and Rogers (1996)

As a first pass, it is useful to compare the border effect found for the EU to that found for the US and Canada by Engel and Rogers. The econometric specification they use to measure relative volatility is given by[11]:

$$V(p_{ij,t}) = \alpha D_{ij} + \beta X_{ij} + \sum_{k=1}^{K} \gamma_k C_k + u_{ij} \tag{4}$$

$D_{i,j}$ is the log distance between location *i* and *j*, $X_{i,j}$ is a dummy variable equal to 1 if *i* and *j* are separated by a border, and C_k is city-specific indicator variable for each city in the sample. Inclusion of C_k picks up any idiosyncratic effects for city pairs not covered by distance or the border (such as differences in markups.) The coefficient on the border dummy will measure Φ, the excess volatility of cross-border prices, once distance and city fixed effects have been controlled for. In estimating relative price volatility, we can consider several methods. One candidate is the standard deviation of the first-difference relative price defined as:

$$p_{ij,t} = \ln\left(p_{i,t} / p_{j,t}\right) - \ln\left(p_{i,t-1} / p_{j,t-1}\right) \tag{5}$$

A drawback of using this measure is that it implicitly assumes that the relative price has a unit root. Engel and Rogers suggest modeling the relative price as a 6th order autoregressive process with 12 monthly seasonals. Then, the volatility measure is the standard deviation of the 12-month ahead forecast error. With this dataset, there are no qualitative differences in the results using either measure[12]. For ease of reproduction the first differenced relative price measure will be reported hereafter. That the end results were nearly identical may indicate that even though relative prices may not follow a random walk, they are persistent enough to act as though they were.

Table 2 presents some initial results from the estimated model (equation 4). Results are given for three categories of goods: lodging, meals, and the sum of these two price series. The estimation results are compared to those found by Engel and Rogers

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[11] This specification is essentially a gravity equation, which has proven to very successful in estimates of international trade (c.f. Bergstrand, 1985).

[12] The mean absolute deviation was also used as an alternative measure of relative price volatility. The sign and significance of the border and distance effects were the same with either measure, although the mean absolute deviation yielded a smaller distance effect.

(1996) for comparable goods. The relative price is computed as the first difference of the log relative price.

Table 2: Effect of distance and the border relative to Engel and Rogers

	Per Diem	E.R. Pooled Estimates	Lodging	E.R. "Shelter"	Meals	E.R. "Food away from home"
Log Distance	2.08	1.06	2.59	0.84	2.05	0.18
	(0.19)	(0.33)	(0.24)	(0.18)	(0.16)	(0.89)
Border	11.37	11.9	9.95	9.42	10.56	9.71
	(0.26)	(0.42)	(0.32)	(0.21)	(0.24)	(0.11)
Border Width[13]	74 miles	1,740 miles	18 miles	4,178 miles	60 miles	362 miles
Adj. R^2	0.89	0.77	0.87	0.93	0.86	0.97

** White heteroskedasticity consistent standard errors reported in parentheses. All coefficients are multiplied by 10^3. Unshaded columns give the estimation results for the lodging, meals, and per diem price series from my data, while shaded columns give the estimates for comparable goods from Table 3 of Engel and Rogers (1996). All coefficients are significant at the 1% level, and the border width is calculated from the upper bound on the distance effect.*

Looking first at the estimation results for the per diem rate, both distance and the border prove to be significant determinants of relative price volatility. Surprisingly, the border coefficient proves to be nearly identical to that for the pooled price series of Engel and Rogers. The robustness of these results indicates that not only does the border matter, this phenomena extends over a wider range of countries than just the US and Canada. In terms of the significance of the border relative to that of distance, the border has a larger effect than distance, but not in the same magnitude as that found by Engel and Rogers. They find that crossing the border is equivalent to approximately 1700 miles of distance between cities in the same country. I find that on average (over all countries in the sample), the border is 74 miles wide.

Both the distance and border effects remain significant when looking at the lodging and meals price series, supporting the results found by Engel and Rogers. We would expect that "meals" is a more tradable good than "lodging", implying a smaller distance effect for this good. This does, in fact, hold both for this dataset and the original Engel and Rogers study.[14] Surprisingly, the border effect is larger for the meals category than for the lodging category, although this difference is only significant at the 10% level. This may indicate that the border is picking up possible hysteresis in price setting caused by exchange rate changes. Finally, this study finds much narrower border "widths" than those found in Engel and Rogers (1996). This is due to larger distance effects in my sample than those found between the US and Canada, and not because of smaller border effects. This implies that that marginal effect of a mile of distance on relative price volatility between two cities in the same European country is greater than that of a mile of distance between two cities located within the US or Canada. For example, these estimates imply that relative price volatility between Berlin and Frankfurt should be higher than that found between New York and Washington D.C, given that the distance between these city pairs is approximately the same. This may indicate that markets are more segmented within European countries than within the U.S., a surprising result.

III-B Idiosyncratic Shocks

The analysis up to this point has focused on the excess volatility of cross-border prices. Implicit in this has been the fact that cross-border prices show less co-movement than within-country prices. Thus, a useful way to look at relative prices, as suggested by Engel and Rogers (1998), is in terms of their common and idiosyncratic shocks. Assume that we can model nominal prices as[15]:

$$p_{i,t} = \Psi(L) p_{i,t-1} = \upsilon_{i,t} + \omega_t \tag{6a}$$

$$p_{j,t} = \Psi(L) p_{i,t-1} = \upsilon_{j,t} + \omega_t \tag{6b}$$

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[13] A simple calculation of the "width of the border" would be exp[11.37/2.075]. Given that the natural log function is concave and the relative imprecision of the distance coefficient, it may be better to err on the side of caution by using the upper bound of the distance coefficient. The width would then be exp[11.37/(2.08+0.31)] = 118 km = 74 miles. The Engel and Rogers estimate of the border width uses this method. As a comparison, the lower bound on border width is reported here.

[14] The difference between the meals and lodging coefficient estimates from my data is significantly different from zero at the 5% level.

[15] While prices are modeled here as having a higher order autoregressive process, the possibility of a unit root is not ruled out. In fact, the volatility specification using the first difference of the relative price assumes a unit root. These equations are adapted from Engel and Rogers (1998).

Each price series is influenced by an idiosyncratic shock ($\upsilon_{i,t}$) and a common shock that affects all prices (ω_t). The correlation of the two price series is then:

$$Corr\left(\Psi(L)p_{i,t},\Psi(L)p_{j,t}\right)=\frac{Var(\omega)}{\sqrt{\left[Var(\upsilon_i)+Var(\omega)\right]\left[Var(\upsilon_j)+Var(\omega)\right]}} \tag{7}$$

This can be rearranged to:

$$Corr\left(\Psi(L)p_{i,t},\Psi(L)p_{j,t}\right)=\left(1+\frac{Var(\upsilon_i)}{Var(\omega)}\right)^{-\frac{1}{2}}\left(1+\frac{Var(\upsilon_j)}{Var(\omega)}\right)^{-\frac{1}{2}} \tag{8}$$

The size of the idiosyncratic shock relative to the common shock determines the correlation between the two price series. To see this explicitly, consider the partial derivative of the correlation measure with respect to the variance of the idiosyncratic shock:

$$\frac{\partial Corr(\cdot)}{\partial Var(\upsilon_i)}=-\frac{1}{2*Var(\omega)}\left(1+\frac{Var(\upsilon_i)}{Var(\omega)}\right)^{-\frac{3}{2}}\left(1+\frac{Var(\upsilon_j)}{Var(\omega)}\right)^{-\frac{1}{2}}<0 \tag{9}$$

As the idiosyncratic shocks increase relative to the common shocks, the price correlations fall.

How much can this hypothesis explain the differential in price covariance? To determine this, we must isolate the innovation part of price changes from the deterministic. This is accomplished by fitting a higher-order autoregressive process with seasonal dummies to each price series[16]. The covariance of price shocks was then taken to be the average pair-wise covariance between the residuals from each fitted price series. To determine the importance of the differential shock hypothesis in explaining excess volatility, the following relation was estimated:

$$V(p_{i,j})=\alpha D_{i,j}+\beta X_{i,j}+\delta SCV_{i,j}+\sum_{k=1}^{K}\gamma_k C_k+u_{i,j} \tag{10}$$

Where $SCV_{i,j}$ is the price shock covariance between locations *i* and *j*. If the price shocks between two locations are highly correlated, then the prices between those locations should also be highly correlated (and thus have less relative volatility). Given this, the

[16] Specifically, each price series was estimated with a 6th order AR process with 12 monthly seasonal dummies. Using a data-dependent lag selection method with Akaike information criteria, the optimal lag length was between 4 and 6 over all price series, so to err on the side of caution, 6 lags were used.

coefficient on SCV should be negative. By including this variable, any remaining distance or border effects will be filtered from any differential response to shocks between cities.

III-C Uneven Sampling Bias

Before estimating the model accounting for differential price shocks, we must first correct for another potentially serious bias due to uneven sampling of locations from each country. Cheung and Lai (2001) show that the bias in the excess volatility of cross-border pairs over within country pairs (what the border effect is picking up) is dependent on the number of locations sampled from each country. Recall the definition of excess cross border price volatility between two countries, home and foreign, presented in equation 3:

$$\Phi = AVG[Var(p^{c}_{ij,t})] - AVG[Var(p^{w}_{kl,t})]$$

When computing $AVG[Var(p^{c}_{ij,t})]$, every cross-border pair contains one observation from the home country, and one observation from the foreign country. The average variance of within country pairs, however, will be a weighted average of the relative price volatility within the home $Var(p^{h}_{kl,t})$ and foreign $Var(p^{f}_{kl,t})$ countries, with the weights being the fraction of locations sampled from each country.

To see if there is a bias from sampling different numbers of cities from each country, let us compare the expected border effect from even sampling (the same number of cities samples from home and foreign) to the border effect under uneven sampling. Let $E[Var(p^{c}_{ij,t})]$, $E[Var(p^{h}_{kl,t})]$, and $E[Var(p^{f}_{kl,t})]$ be the mean volatilities of the cross-border, home country, and foreign country relative prices respectively. The number of locations sampled from the home country is n_h and the number sampled from the foreign country is n_f. When $n_h = n_f$, the expected excess volatility of cross border prices is:

$$E(\Phi) = E[Var(p^{c}_{ij,t})] - \frac{E[Var(p^{h}_{kl,t})] + E[Var(p^{f}_{kl,t})]}{2} \qquad (11)$$

When $n_h \neq n_f$, the expected border effect is:

$$E(\Phi^{*}) = E[Var(p^{c}_{ij,t})] - \frac{n_h E[Var(p^{h}_{kl,t})] + n_f E[Var(p^{f}_{kl,t})]}{n_h + n_f} \qquad (12)$$

Using even sampling as the benchmark, the expected bias from uneven sampling can be found by subtracting (11) from (12)

$$E(\Phi^* - \Phi) = \frac{n_f - n_h}{n_f + n_h} * \frac{E[Var(p_{kl,t}^h)] - E[Var(p_{kl,t}^f)]}{2} \qquad (13)$$

Unless $E[Var(p_{kl,t}^h)] = E[Var(p_{kl,t}^f)]$, the bias is non-zero whenever $n_h \neq n_f$. The direction of the bias will be determined by how the relative price volatility for two locations in the home country compares to the relative price volatility for two locations in the foreign country. For example, if $E[Var(q_t^h)]$ is greater than $E[Var(q_t^f)]$, and there are more home countries sampled than foreign ($n_h > n_f$), the expected bias will be negative, indicating that the excess volatility of cross border relative prices is being underestimated.

With this in mind, the dataset was re-sampled to obtain an equal number of locations from each country. In order to maintain a wide range of countries, the ten largest cities (as measured by average population over the sample period) were selected from each country. If a country did not have ten cities in the dataset, it was dropped. The new sample had monthly price observations for the ten largest cities in Belgium, France, Germany, Italy, the Netherlands, Spain, and the U.K. From these 70 cities, we can create a total of 2,415 city-pairs. Of these, 315 are within country pairs and 2,100 are cross-border pairs. Summary statistics for this "even" sample are presented below:

Table 3: Summary Statistics for Even Sample

	# of cities	# of city pairs	Avg. Distance (km)	Avg. St. Dev. (x1000)
Belgium	10	45	78.58	20.28
France	10	45	346.22	45.90
Germany	10	45	337.80	49.74
Italy	10	45	438.45	52.58
The Netherlands	10	45	54.39	41.82
Spain	10	45	486.17	28.28
U.K.	10	45	277.13	49.21
Cross-Border Pairs	-	2100	918.33	51.76
Entire Sample	70	2415	836.17	50.37

The even sample preserves the cross-country heterogeneity that made the dataset appealing in the first place. Countries differ both in terms of geographic size and population density. Unfortunately, we are left with only one country that does not use the euro as its currency (the UK), but some useful inferences may still be made. In terms of relative price volatility, we see a similar pattern emerge for the even sample as for the uneven sample. There appears to be a positive (albeit weak) correlation between distance and relative price volatility, with countries that have cities located closer together experiencing lower volatility than those whose cities are located further apart. In general, within-country relative price volatility is lower than cross-border volatility. Although the average cross-border volatility is nearly identical in the even and uneven samples, the uneven sample has a much larger number of cities, which may indicate that uneven sampling may indeed be causing a bias.

Table 4 lists the estimation results from the entire sample and those using the even sample. Both the original econometric specification given in equation 4 and the specification incorporating idiosyncratic shocks (10) are presented.

Table 4 – Even vs. Uneven Sampling and Idiosyncratic Shocks

Specification	Distance ($x10^3$)	Border ($x10^3$)	SCV	Border Width	R^2
All Cities I	2.08* (0.19)	11.37* (0.26)	-	118 km	0.89
Even Sample I	0.47 (0.33)	9.98* (0.73)	-	20,992 km	0.90
All Cities II	0.87* (0.10)	5.58* (0.17)	-23.75* (0.33)	216 km	0.95
Even Sample II	0.68* (0.23)	3.85* (0.46)	-21.24* (0.78)	38 km	0.96

An asterisk indicates significance at the 1% level. Specification I is the original Engel and Rogers specification presented in equation 4. Specification II incorporates the covariance of idiosyncratic shocks to the model as an explanatory variable. All errors are heteroskedasticity consistent.

The evidence suggests that uneven sampling has created an upward bias in both the border and distance effects. Comparing the coefficient estimates from uneven and even sampling, the magnitudes of the border and distance effects universally fall with even sampling. Despite this, the border is still highly significant in both specifications. Given that even sampling uses a much smaller sample size, this is a striking result. The distance effect falls in both magnitude and significance when moving to an even sample. It is still positive, but has less explanatory power. Inclusion of idiosyncratic shocks as an explanatory variable would appear to be a wise decision. It has the expected sign and increases the R^2 value from 0.89 to 0.95. Once unique price shocks have been accounted for, the border is still significant, while the distance effect falls in both magnitude and significance. This would seem to indicate that the distance effect had been picking up differences in market conditions between locations. Once differential price shocks (which may reflect differences in market conditions) are accounted for, the distance effect is reduced. The continued significance of the border effect implies that this variable is picking up something other than just differential market conditions.

III-D Incomplete Exchange Rate Pass Through

Up to this point, the role of exchange rates in determining excess volatility of cross-border prices has been ignored. All prices in the dataset are quoted in a common currency, so at a first glance, this seems reasonable. Central to many of the arguments in the literature however, is the interaction between sticky prices and volatile exchange rates. For example, consider prices in Marseille, Paris, and London. It may be the case that the relative *real* price volatility between Marseille and Paris is no greater than between Paris and London, indicating no border effect. However, we are given nominal prices. To compare prices between France and the UK, all prices must be converted into a common currency. In order to measure the relative volatility between Paris and London, the London price must be converted to euros. Doing this adds another source of volatility: the exchange rate. As numerous studies have shown the failure of PPP, converting prices in London from pounds to euros does in fact add an artificial element of volatility. As mentioned above, all prices in this dataset are quoted in dollars. This does not mean that the actual prices in each city are listed in dollars, rather that the prices themselves were converted into dollars before being reported. Should this eliminate the exchange rate bias? Let us return to the two-country model used to construct the measure of excess cross-border price volatility. Before being converted into a common currency, the log-relative price between two cross-border cities is:

$$p^{c}_{ij,t} = p^{h}_{i,t} - p^{f}_{j,t} - e_t \tag{14}$$

where e_t is the log nominal exchange rate between home and foreign currency. The variance of this relative price is:

$$Var(p^{c}_{ij,t}) = Var(p^{h}_{i,t} - p^{f}_{j,t}) + Var(e_t) - 2Cov(e_t, p^{h}_{i,t} - p^{f}_{j,t}) \tag{15}$$

The variance of relative prices within the same country is:

$$Var(p^{w}_{kl,t}) = Var(p^{w}_{k,t} - p^{w}_{l,t}) \tag{16}$$

The excess volatility of cross-border relative prices is:

$$\Phi = AVG[Var(q^{c}_{ij,t})] - AVG[Var(q^{w}_{km,t})]$$

Excess volatility can further be broken down into two components, one isolating the effect of the exchange rate and the other isolating the effect of price changes.

$$\Phi = \xi + \sigma \tag{17}$$

The price change component is defined as:

$$\sigma = AVG[Var(p^{h}_{i,t} - p^{f}_{j,t})] - AVG[Var(p^{w}_{k,t} - p^{w}_{l,t})] \tag{18}$$

The exchange rate component is defined as:

$$\xi = AVG[Var(e_t)] - 2 * AVG[Cov(e_t, p^{h}_{i,t} - p^{f}_{j,t})] \tag{19}$$

The exchange rate component includes exchange rate variability as well as the covariance between the exchange rate and the cross-border relative price. The latter term allows for adjustment in prices in response to changes in the exchange rate. Given that prices are relatively sticky and there is incomplete exchange rate pass through, one would expect this term to be relatively small in comparison to exchange rate variability. If this is indeed the case, then $\xi > 0$, and a high degree of exchange rate variability will lead to higher excess volatility. To incorporate this effect, exchange rate volatility is added as a regressor.

The example above only had two countries, and thus only one exchange rate. This dataset, however, has many countries, requiring many exchange rates. For each country, the monthly local currency to US dollar exchange rate is collected. For each pair, the exchange rate is defined as the ratio of the first city's exchange rate to the exchange rate of the second city. The variance of the log exchange rate is then calculated over the entire sample period.

The sample period covers a significant structural break in exchange rates, the adoption of the euro by twelve countries in the sample on January 1st, 1999. Not factoring this event into account could lead to seriously biased estimates of the exchange rate effect. Consider the relative price between Rome and London. Pre-euro, the lire to pound exchange rate was a very large number. Given this, the variance of the exchange rate was also very high. Post-euro, however, the exchange rate was a much smaller number since each pound would buy a lot less euros than it did lire. Exchange rate variability was much smaller than before, even if the variation in real terms were no different. To correct for this, the sample was split into two periods, pre and post euro (from 1/1995 to 12/1998 and from 1/1999 to 6/2002) and an exchange rate index for each period was constructed, using 1/1995 as the base for the first period and 1/1999 as the base for the second. The following relation was then estimated:

$$V(p_{i,j}^{S}) = \alpha D_{i,j} + \beta X_{i,j} + \delta SCV_{i,j}^{S} + \lambda EV_{i,j}^{S} + \sum_{k=1}^{K} \gamma_k C_k + u_{i,j}^{S} \qquad (20)$$

The superscript "*S*" refers to the sub-sample the price series comes from. $EV_{i,j}$ is the log exchange rate volatility between cities *i* and *j*. Based on the analysis above, we would expect higher nominal exchange rate volatility to translate into higher cross-border price volatility. Table 5 gives the estimated coefficients from equation 20 for the periods before and after the introduction of the euro, using the even sample of cities.

Table 5: Correcting for Exchange Rate Volatility

	Pre-EMU (1/1995 – 12/1998)	**Post-EMU (1/1999 – 6/2002)**
Distance ($x10^3$)	0.90*	-0.63**
	(0.25)	(0.30)
Border ($x10^3$)	6.05*	7.03*
	(0.75)	(0.62)
SCV	-17.18*	-22.68*
	(5.76)	(0.94)
EV ($x10^4$)	0.23*	1.80*
	(0.06)	(0.93)
Adj. R^2	0.95	0.94

** Heteroskedasticity consistent standard errors in parentheses. Significance at the 1% and 5% levels given by * and ** respectively. The price series used is the per diem rate (sum of lodging and meals).*

Even after incorporating exchange rate volatility, the border effect remains positive and significant. This is a striking result given that incomplete exchange rate pass through has been proposed as one of the key reasons why cross border prices are more volatile than within country prices. The impact of exchange rate volatility on relative price volatility increases in the post-euro period. This result is driven by the fact that exchange rate volatility in the post-euro period comes from movements between the euro and the non-euro currencies (in this sample, just the pound). If an increase in exchange rate volatility has a greater effect on relative price volatility in the post-euro period than in the pre-euro period, then it would appear that monetary union has insulated the member countries from incomplete exchange rate pass through at the expense of the non-members experiencing cross-border prices that are more sensitive to movements in the exchange rate.

How much of the border effect can be explained by incomplete exchange rate pass through? To quantify this effect, we can compare the estimated border effects before and after incorporating for exchange rate volatility. Table 6 gives these estimated border effects for the pre and post euro periods:

Table 6: Exchange Rate Volatility and the Border Effect

	Pre-EMU	Post-EMU
Border w/o EV	6.83 (0.86)	7.51 (0.65)
Border with EV	6.05 (0.75)	7.03 (0.62)
Average Volatility of Cross-Border Pairs	43.47	57.14

** The row titled "Border w/o EV" gives the estimated coefficient on the border dummy from equation 10. "Border with EV" gives the coefficient on the border dummy from equation 20. The average cross-border volatility is the average standard deviation of all cross-border relative prices, calculated before and after the adoption of the euro. All numbers have been multiplied by 10^3.*

In the pre-EMU period, the average relative price volatility was 43.47 x 10^{-3}. Based on this, the presence of a border can explain 15.7% of cross-border price volatility when we do not account for exchange rate volatility.[17] Allowing for exchange rate volatility, the border can explain 13.9% of cross-border price volatility. Thus, exchange

[17] This calculation was made by dividing the border effect by the average relative price volatility.

rate volatility may account for nearly 12% of the estimated border effect prior to the formation of the EMU. In the post-EMU period, the border effect can explain 13.1% of cross border price volatility before accounting for exchange rate volatility, and 12.3% after accounting for this variable. In the post-EMU period, incomplete exchange rate pass through may account for only 6% of the estimated border effect. That the importance of incomplete exchange rate pass through declines in the post-EMU period is hardly surprising, given that most of the countries in this sample are using the same currency.

IV. The Effects of Monetary Union

One of the most useful features of this data set is the ability to look at relative price volatility before and after the establishment of monetary union. Before moving on to the effects of monetary union on relative price volatility, it will be useful to look at some of the reasons why establishing a single currency would reduce relative price volatility. The basic idea for this can be found in the work on optimum currency areas by Mundell (1961) and McKinnon (1963). In the context of this study, the adoption of the euro as an international transaction currency will insulate prices from exchange rate volatility. Such a result is presented in Devereux, Engel, and Tille (1999) who use a simple sticky-price intertemporal model to show how consumer prices respond to exchange rate changes if traded goods prices were quoted in euros as opposed to dollars. In this setup, menu costs limit price changes in response to exchange rate movements. Consider the following example. A U.S. firm selling goods in Europe facing twelve different currencies would need to have twelve different pricing rules[18]. Because of the high transaction costs associated with this, firms will stabilize prices in terms of U.S. dollars. This in turn, will cause consumer prices (which are set in the local currency) to be more sensitive to changes in the dollar exchange rate. If however, the firm faced only one currency, transaction costs would be lower, and more firms would adhere to a “euro pricing” rule further lowering price volatility.

[18] This could apply to a firm selling any type of good whose price (or the prices of its inputs) are quoted in dollars.

We can estimate how the border and distance effects have changed since the inception of the EMU by taking advantage of the fact that both the distance and presence of a national border between two cities does not change over time. Recall the econometric specification given in equation (10)[19]

$$V(p_{i,j}) = \alpha D_{i,j} + \beta X_{i,j} + \delta SCV_{i,j} + \sum_{k=1}^{K} \gamma_k C_k + u_{i,j}$$

Splitting the sample into a pre-euro period (A) and a post-euro period (B) gives:

$$V^A(p_{i,j}) = \alpha^A D_{i,j} + \beta^A X_{i,j} + \delta^A SCV_{i,j}^A + \sum_{k=1}^{K} \gamma_k^A C_k + u_{i,j}^A \tag{21a}$$

$$V^B(p_{i,j}) = \alpha^B D_{i,j} + \beta^B X_{i,j} + \delta^B SCV_{i,j}^B + \sum_{k=1}^{K} \gamma_K^B C_k + u_{i,j}^B \tag{21b}$$

The distance and border variables do not change over time. The price shock covariance term will, however. If we subtract the two volatility measures, we obtain the following expression:

$$\begin{aligned} V^B(p_{i,j}) - V^A(p_{i,j}) &= (\alpha^B - \alpha^A) D_{i,j} + (\beta^B - \beta^A) X_{i,j} + \delta^B SCV_{i,j}^B - \delta^A SCV_{i,j}^A \\ &+ \sum_{k=1}^{K} (\gamma_k^B - \gamma_k^A) C_k + (u_{i,j}^B - u_{i,j}^A) \end{aligned} \tag{22}$$

If monetary union has had no effect, the coefficients on distance and the border from estimating the above relation should be statistically insignificant. A positive distance or border coefficient would indicate an increase in this effect in the post-EMU period, whereas a negative coefficient would indicate a decrease. Thus, a negative coefficient on the border dummy would indicate a decrease in excess cross-border price volatility in the post-EMU period.

Table 7 gives the estimated coefficients from equation 22 for both the entire sample and for only pairs of cities located within the EMU. Looking at all city pairs, the border effect has increased, but this increase is not significantly different from zero. This would suggest that average cross-border price volatility in excess of within country volatility has not changed in the post-EMU period. The distance effect has, however,

[19] It is implausible to estimate equation 20 here, since there will be perfect collinearity between exchange rate volatility and the border when looking at EMU country pairs. Given the small impact of this variable on the border effect, omitting this variable may not be too problematic.

fallen significantly, indicating that arbitrage barriers have fallen. Restricting our attention to city pairs within the EMU, we see that there has been a decrease in both the border and distance effects, although the change in the border effect is not significantly different from zero.

Table 7 – Differential Distance and Border Effects

	All Countries	EMU Pairs
Change in the Border Effect (x10^3)	1.10	-0.41
	(0.93)	(0.93)
Change in the Distance Effect (x10^3)	-1.60*	-1.49*
	(0.34)	(0.33)

*Significance at the 1% level is given by *. "EMU pairs" refers to the relative price over city pairs located within the euro-zone. All errors are heteroskedasticity consistent.*

The estimation results above pooled many countries. It may be useful to look at the effect of monetary union on a country-by-country basis. Table 8 gives the differential volatility specification given in equation 22 for all country pairs. The estimation results are grouped by EMU membership and the economic size of the countries on either side of the border.[20] Table 8A gives the differential border and distance effects between the UK and the EMU countries, broken down by economic size. The border effects between the UK and the small EMU countries have significantly increased in the post-EMU period, indicating higher cross-border price volatility between these countries. There does not appear to be any significant change in the border effect between the UK and the large EMU countries, however. Turning our attention within the EMU, tables 8B, 8C, and 8D give the differential distance and border effects across the EMU, broken down by the size of the country on either side of the border. From table 8B, we can see that the border effects between the France, Germany, and Italy have all significantly decreased in the post-EMU period, indicating that the volatility of cross-border prices has been falling relative to the volatility of within country prices. Turning our attention to the border effects across the small EMU countries of Belgium, the Netherlands, and Spain, we see varied results. Cross border price volatility between Belgium and the Netherlands has

[20] A country's "size" classification was made based on its GDP over the sample period. Since there are six EMU countries in the sample, the three countries with the highest GDP (France, Germany, and Italy) are classified as large, while the three countries with the smallest GDP's (Belgium, the Netherlands, and Spain) are classified as small. Reclassifying Italy as a small country or Spain as a large country will not fundamentally change the results, however.

actually increased, that between Belgium and Spain has significantly decreased, and there has not been a significant change between the Netherlands and Spain. Whether these differential effects are due to increased convergence of the Spanish economy in the post-EMU period or some idiosyncrasy of one of these countries is a topic for future research. Finally, table 8D gives the differential distance and border effects between the large and small EMU countries. While the border effects appear to have decreased in general, none of these changes are significantly different from zero.

Table 8A: The Effect of the EMU on Borders with the UK

	Small Countries			Large Countries		
	UK – BEL	UK – NED	UK – ESP	UK – FRA	UK – GER	UK – ITA
Border	12.88*	5.66***	6.21***	-0.43	5.19	1.86
	(2.62)	(2.93)	(3.42)	(3.38)	(3.50)	(3.15)
Distance	0.92	1.29	0.87	0.70	3.25***	-0.47
	(0.85)	(1.52)	(1.52)	(1.79)	(1.88)	(1.65)

Table 8B: The Effect of the EMU on Intra-EMU Borders: Large vs. Large

	FRA – GER	FRA – ITA	GER – ITA
Border	**-5.14***	**-4.33***	**-5.45***
	(1.92)	(1.92)	(1.92)
Distance	-0.03	**-2.59*****	-1.55
	(1.48)	(1.37)	(1.73)

Table 8C: The Effect of the EMU on Intra-EMU Borders: Small vs. Small

	BEL – NLD	BEL – ESP	NLD – ESP
Border	2.20**	-4.78*	-2.04
	(1.10)	(1.75)	(3.31)
Distance	-0.54	-0.93	-1.06
	(0.77)	(0.80)	(1.37)

Table 8D: The Effect of the EMU on Intra-EMU Borders: Large vs. Small

	FRA – BEL	GER – BEL	FRA – NED	GER – NED	FRA – ESP	GER – ESP
Border	2.02	-0.08	-2.55	-0.43	-2.74	-0.71
	(1.91)	(1.46)	(2.88)	(1.70)	(1.98)	(2.20)
Distance	-1.28	0.97	-0.54	0.27	0.58	0.20
	(0.72)	(0.95)	(1.28)	(1.21)	(1.42)	(1.48)

Estimates of equation 10 on a country-by-country basis. The price series is the daily per diem rate (the sum of lodging and meals.) The reported coefficients give the differential distance and border effects. A positive coefficient indicates that the effect post-euro is greater than it was pre-euro. 1%, 5%, and 10% significance levels are given by *, **, and * respectively.*

V. Conclusion

The two main goals of this paper are to extend the Engel and Rogers border effect to the European Union and to examine what (if any) effect monetary union has had on the importance of the border. Both national borders and distance are significant determinants of relative price volatility, and the border does appear to be more important than distance. However, the magnitude of this difference is not nearly as large as that found in Engel and Rogers original (1996) study. They conservatively estimate that crossing the U.S. – Canada border is equivalent to adding over 1,700 miles of distance between cities in the same country. My estimates are not nearly as grandiose. I estimate that crossing the border is equivalent to anywhere from 20 to 75 miles of distance between cities. Thus, the border matters, but hardly any more than physical distance. Still, when considering that a border is nothing more than an imaginary line on a map, and that the border effect remains even after accounting for differential price shocks, trade barriers, distance, and exchange rate variability, this is a considerable result.

Proponents of monetary union argue that it will usher in a new age of price stability and economic prosperity across the currency area. Has this really been the case? Has price volatility decreased, and if so, has this been the result of monetary union? The results in this paper suggest that the effects of the EMU on cross-border price volatility depend on the size of the countries on either side of the border. Cross-border price volatility between the UK and the smaller EMU countries has significantly increased in the post-EMU era, but there has been no significant change between the UK and the larger EMU countries. This may indicate that the smaller EMU countries have turned inward, resulting in increased market segmentation between these nations and countries outside the EMU, while the larger EMU countries have managed to avoid this. Turning our attention to the effect of the EMU on cross-border price volatility within the union, we see that the country size effect re-emerges, but in reverse. In this case, cross-border price volatility between the large EMU countries has been significantly reduced while there has been little evidence of significant reductions between the small countries, or between large and small countries.

Several policy implications with regard to European countries outside of the EMU (most notably the UK) emerge from this study. Having stable cross-border prices is important for several reasons. Increased cross-border price stability has been shown to promote trade and investment between nations. Secondly, the potential problem of common monetary policy across a monetary union is mitigated if prices across the monetary union are stable. If one of the effects of the EMU has been to stabilize cross-border relative prices, then we can argue that two of the benefits of EMU membership are increased trade and cross-border investment and reduced uncertainty over asymmetric effects of a common monetary policy. As these are two benefits of EMU membership, they can also be considered as costs of not joining the union.

The results of this study suggest that there has been an increase in cross-border price volatility between the UK and the small EMU countries, but no change between the UK and the large EMU countries. Given that the majority of the UK's cross-border transactions are with the larger EMU countries, the small country border effect increase should not be a significant cost of non-membership. However, given that the large EMU countries have seen increased market integration amongst themselves, this may very well be a benefit the UK is missing out on from not joining the EMU. The decreased cross-border price volatility between France and Germany implies greater trade and investment between these nations. Had the UK been a member of the EMU, would cross-border volatility between the UK and Germany also fallen? This is a desirable result for those concerned with international trade and foreign investment, but less important for more domestically oriented goals. Thus, the results of this study will pertain more to open countries that rely on international trade and inflows of foreign capital, both of which are aided by international price stability. Given that most (if not all) of the countries in Western Europe fit this description, the benefits of joining the EMU may indeed outweigh the costs.

VI. References

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Appendix A – List of Cities, Uneven Sample

Country	City
Austria	Innsbruck
	Linz
	Salzburg
	Vienna
Belgium	Antwerp
	Bertrix
	Brugge
	Brussels
	Diegem
	Florennes
	Gent
	Gosselies
	Herstal
	Leuven
	Liege
	Mons
	Ostende
	Shape/Chievres
	Zaventem
Denmark	Copenhagen
Finland	Helsinki
France	Aix-en-Provence
	Bordeaux
	Clermont-Ferrand
	Lyon
	Marignane
	Marseille
	Metz
	Montpellier
	Mulhouse
	Nancy
	Nice
	Paris
	Strasbourg
	Suresnes
	Toulouse
Germany	Aachen
	Bad Honnef
	Berlin
	Boeblingen
	Bonames
	Bonn
	Bremen
	Chemnitz
	Cologne
	Delmenhorst
	Dresden
	Duesseldorf
	Echterdingen
	Erfurt
	Erlangen
	Eschborn
	Essen
	Esslingen
	Frankfurt am Main

Country	City
Germany	Fuerth
	Garmisch-Partenkirchen
	Germering
	Giebelstadt
	Hamburg
	Hannover
	Heidelberg
	Herongen
	Herrsching
	Herzogenaurach
	Hoescht
	Ingolstadt
	Kaiserslautern-Landkreis
	Kalkar
	Kitzingen
	Koenigswinter
	Konstanz
	Kornswestheim
	Leipzig
	Ludwigsburg
	Moenchen-Gladbach
	Munich
	Nellingen
	Neu Ulm
	Niederbachem
	Nuernberg
	Oberammergau
	Offenbach
	Rhoendorf
	Roedelheim
	Rostock Warnemuende
	Saarbrucken
	Schwabach
	Schwerin
	Sembach
	Sindelfingen
	Starnberg
	Stuttgart
	Sylt Island
	Tuebingen
	Twisteden
	Ulm
	Wahn
	Wuerzburg
	Zirndorf
Ireland	Adare
	Dublin
	Limerick
Italy	Bari
	Bologna
	Ferrara
	Florence
	Genoa
	La Spezia
	Milan
	Modena
	Naples
	Palermo

Country	City
Italy	Pisa
	Ravenna
	Reggio Emilia
	Rimini
	Rome
	Siena
	Taormina
	Trieste
	Turin
	Venice
	Verona
	Vicenza
Luxembourg	Luxembourg
The Netherlands	Amsterdam
	The Hague
	Leiden
	Lisse
	Noordwijk
	Papendrecht
	Rotterdam
	Schiphol
	Utrecht
	Ypenberg
Norway	Oslo
Portugal	Lisbon
	Oporto
Spain	Barcelona
	Bilbao
	Fuengirola
	Gerona
	Getafe
	La Coruña
	Lerida
	Logroño
	Madrid
	Malaga
	Marbella
	Oviedo
	San Sebastian
	Santander
	Santiago de Compostela
	Seville
	Tarragona
	Torrejon
	Torremolinos
	Valencia
	Vitoria
	Zaragoza
Sweden	Stockholm
Switzerland	Bern
	Geneva
	Zurich

Country	City
United Kingdom	Aberdeen
	Beaconsfield
	Belfast
	Birmingham
	Bournemouth
	Brighton
	Bristol
	Bury St. Edmunds
	Cambridge
	Canterbury
	Cardiff
	Caversham
	Crawley
	Dover
	Edinburgh
	Ft. Halstead
	Gatwick
	Glasgow
	Harrogate
	High Wycombe
	Horley
	Inverness
	Liverpool
	London
	Manchester
	Menwith Hill
	Nottingham
	Oxford
	Plymouth
	Poole
	Portsmouth
	Reading
	Southampton
	Winchester

Appendix B – List of Cities, Even Sample

Country	City	Country	City
Belgium	Antwerp	Netherlands	Amsterdam
	Brugge		The Hague
	Brussels		Leiden
	Gent		Lisse
	Herstal		Noordwijk
	Leuven		Papendrecht
	Liege		Rotterdam
	Mons		Schiphol
	Ostende		Utrecht
	Zaventem		Ypenburg
France	Aix-en-Provence	Spain	Barcelona
	Bordeaux		Bilbao
	Clermont-Ferrand		La Coruña
	Lyon		Madrid
	Marseille		Malaga
	Montpellier		Oviedo
	Nice		Seville
	Paris		Valencia
	Strasbourg		Vitoria
	Toulouse		Zaragoza
Germany	Berlin	United Kingdom	Birmingham
	Bremen		Bristol
	Cologne		Cardiff
	Duesseldorf		Edinburgh
	Essen		Glasgow
	Frankfurt am Main		Liverpool
	Hamburg		London
	Hannover		Manchester
	Munich		Nottingham
	Stuttgart		Plymouth
Italy	Bari		
	Bologna		
	Florence		
	Genoa		
	Milan		
	Naples		
	Palermo		
	Rome		
	Turin		
	Venice		

Appendix C – Distance and Border Effect Matrices

C1: Daily Per Diem Rate, January 1995 – June 2002

Distance Effects

	Belgium	**France**	**Germany**	**Italy**	**Netherlands**	**Spain**	**UK**
Belgium	-	0.32 (0.87)	0.002 (0.99)	0.92 (1.08)	0.77 (0.97)	0.88 (0.88)	1.32 (0.98)
France	0.32 (0.87)	-	-1.37** (0.60)	0.23 (0.89)	-0.13 (0.87)	0.58 (0.58)	0.32 (0.73)
Germany	0.002 (0.99)	-1.37** (0.60)	-	-0.76 (0.74)	-0.24 (0.69)	-0.54 (0.73)	-0.19 (0.84)
Italy	0.92 (1.08)	0.23 (0.89)	-0.76 (0.74)	-	0.22 (1.09)	0.65 (0.94)	0.33 (0.92)
Netherlands	0.77 (0.97)	-0.13 (0.87)	-0.24 (0.69)	0.22 (1.09)	-	0.32 (0.75)	1.40 (1.35)
Spain	0.88 (0.88)	0.58 (0.58)	-0.54 (0.73)	0.65 (0.94)	0.32 (0.75)	-	0.33 (0.72)
UK	1.32 (0.98)	0.32 (0.73)	-0.19 (0.84)	0.33 (0.92)	1.40 (1.35)	0.33 (0.72)	-

Border Effects

	Belgium	**France**	**Germany**	**Italy**	**Netherlands**	**Spain**	**UK**
Belgium	-	4.03* (1.48)	4.83* (1.06)	6.99* (2.06)	2.41** (1.27)	3.71** (1.76)	8.89* (2.10)
France	4.03* (1.48)	-	3.26* (0.80)	3.12* (0.79)	0.62 (1.88)	2.10** (0.80)	2.48*** (1.34)
Germany	4.83* (1.06)	3.26* (0.80)	-	5.73* (0.93)	2.82* (1.04)	5.17* (1.17)	8.20* (1.70)
Italy	6.99* (2.06)	3.12* (0.79)	5.73* (0.93)	-	3.55 (2.68)	5.28* (1.25)	3.91** (1.86)
Netherlands	2.41*** (1.27)	0.62 (1.88)	2.82* (1.04)	3.55 (2.68)	-	1.70 (1.77)	0.45 (3.79)
Spain	3.71** (1.76)	2.10** (0.80)	5.17* (1.17)	5.28* (1.25)	1.70 (1.77)	-	4.67* (1.75)
UK	8.89* (2.10)	2.48*** (1.34)	8.20* (1.70)	3.91** (1.86)	0.45 (3.79)	4.67* (1.75)	-

*Distance and border effects are coefficient estimates (x10^3) from equation 10 in the text (even sample incorporating price shocks). Relative price volatility is defined as the standard deviation of the log first-differenced per diem rate over the entire sample period. Significance at the 10, 5, and 1% levels given by ***, **, and * respectively.*

C2: Distance Effects Before and After the Introduction of the Euro

Distance Effects, Pre-Euro (January 1995 – December 1998)

	Belgium	**France**	**Germany**	**Italy**	**Netherlands**	**Spain**	**UK**
Belgium	-	1.00 (0.80)	-0.12 (0.99)	1.73*** (0.91)	1.37 (1.00)	1.75** (0.86)	0.18 (0.80)
France	1.00 (0.80)	-	-0.73 (0.95)	1.49 (1.03)	0.32 (1.04)	1.25 (0.79)	-0.11 (0.77)
Germany	-0.12 (0.99)	-0.73 (0.95)	-	0.45 (1.21)	-0.17 (0.95)	-0.16 (1.15)	-1.41 (0.86)
Italy	1.73*** (0.91)	1.49 (1.03)	0.45 (1.21)	-	1.25 (1.30)	1.53 (1.01)	0.49 (0.95)
Netherlands	1.37 (1.00)	0.32 (1.04)	-0.17 (0.95)	1.25 (1.30)	-	1.52 (1.25)	0.01 (0.91)
Spain	1.75** (0.86)	1.25 (0.79)	-0.16 (1.15)	1.53 (1.01)	1.52 (1.25)	-	-0.05 (0.72)
UK	0.18 (0.80)	-0.11 (0.77)	-1.41 (0.86)	0.49 (0.95)	0.01 (0.91)	-0.05 (0.72)	-

Distance Effects, Post-Euro (January 1999 – June 2002)

	Belgium	**France**	**Germany**	**Italy**	**Netherlands**	**Spain**	**UK**
Belgium	-	-0.27 (1.04)	0.74 (1.00)	0.48 (1.19)	1.00 (0.87)	0.91 (0.92)	1.64 (1.16)
France	-0.27 (1.04)	-	-0.78 (1.03)	-1.62 (1.18)	0.10 (0.96)	1.82 (1.16)	0.58 (1.39)
Germany	0.74 (1.00)	-0.78 (1.03)	-	-1.30 (1.12)	0.15 (0.92)	-0.20 (0.88)	1.80 (1.54)
Italy	0.48 (1.19)	-1.62 (1.18)	-1.30 (1.12)	-	-0.28 (1.11)	0.59 (1.38)	-0.07 (1.42)
Netherlands	1.00 (0.87)	0.10 (0.96)	0.15 (0.92)	-0.28 (1.11)	-	1.18 (0.90)	1.80 (1.50)
Spain	0.91 (0.92)	1.82 (1.16)	-0.20 (0.88)	0.59 (1.38)	1.18 (0.90)	-	0.92 (1.51)
UK	1.64 (1.16)	0.58 (1.39)	1.80 (1.54)	-0.07 (1.42)	1.80 (1.50)	0.92 (1.51)	-

*Distance effects are coefficient estimates($x10^3$) from equation 10 in the text (even sample incorporating price shocks). Relative price volatility is defined as the standard deviation of the log first-differenced per diem rate. Significance at the 10, 5, and 1% levels given by ***, **, and * respectively.*

C3: Border Effects Before and After the Introduction of the Euro

Border Effects, Pre-Euro (January 1995 – December 1998)

	Belgium	**France**	**Germany**	**Italy**	**Netherlands**	**Spain**	**UK**
Belgium	-	7.01* (1.53)	5.16* (1.35)	6.65* (1.80)	2.17 (1.35)	5.33* (1.80)	4.31* (1.21)
France	7.01* (1.53)	-	9.71* (1.14)	8.71* (1.02)	6.01* (2.22)	5.42* (1.09)	4.72* (1.15)
Germany	5.16* (1.35)	9.71* (1.14)	-	10.85* (1.370)	4.84* (1.48)	5.36* (1.74)	9.13* (1.41)
Italy	6.65* (1.80)	8.71* (1.02)	10.85* (1.370)	-	4.33 (2.95)	5.24* (1.23)	5.60* (1.53)
Netherlands	2.17 (1.35)	6.01* (2.22)	4.84* (1.48)	4.33 (2.95)	-	2.88 (3.04)	3.52** (1.57)
Spain	5.33* (1.80)	5.42* (1.09)	5.36* (1.74)	5.24* (1.23)	2.88 (3.04)	-	3.54** (1.37)
UK	4.31* (1.21)	4.72* (1.15)	9.13* (1.41)	5.60* (1.53)	3.52** (1.57)	3.54** (1.37)	-

Border Effects, Post Euro (January 1999 – June 2002)

	Belgium	**France**	**Germany**	**Italy**	**Netherlands**	**Spain**	**UK**
Belgium	-	9.26* (1.96)	4.88* (1.08)	7.54* (2.23)	4.00* (1.07)	0.82 (1.92)	20.23* (2.27)
France	9.26* (1.96)	-	4.82* (1.42)	5.19* (1.27)	3.03 (2.00)	3.15** (1.58)	4.85*** (2.47)
Germany	4.88* (1.08)	4.82* (1.42)	-	5.04* (1.26)	4.48* (1.10)	4.14* (1.43)	13.72* (3.23)
Italy	7.54* (2.23)	5.19* (1.27)	5.04* (1.26)	-	7.55* (2.49)	5.41* (1.80)	7.63* (2.87)
Netherlands	4.00* (1.07)	3.03 (2.00)	4.48* (1.10)	7.55* (2.49)	-	-1.21 (2.22)	8.75* (2.83)
Spain	0.82 (1.92)	3.15** (1.58)	4.14* (1.43)	5.41* (1.80)	-1.21 (2.22)	-	12.23* (2.87)
UK	20.23* (2.27)	4.85*** (2.47)	13.72* (3.23)	7.63* (2.87)	8.75* (2.83)	12.23* (2.87)	-

*Distance effects are coefficient estimates($x10^3$) from equation 10 in the text (even sample incorporating price shocks). Relative price volatility is defined as the standard deviation of the log first-differenced per diem rate. Significance at the 10, 5, and 1% levels given by ***, **, and * respectively.*