CENTER FOR FINANCIAL STUDIES

No. 2001/07 - revised

How Wide Are European Borders? On the Integration Effects of Monetary Unions

Guenter Beck / Axel A. Weber

Center for Financial Studies
an der Johann Wolfgang Goethe-Universität ▪ Taunusanlage 6 ▪ D-60329 Frankfurt am Main
Tel: (+49)069/242941-0 ▪ Fax: (+49)069/242941-77 ▪ E-Mail: ifk@ifk-cfs.de ▪ Internet: http://www.ifk-cfs.de

Center for Financial Studies

The *Center for Financial Studies* is a nonprofit research organization, supported by an association of more than 120 banks, insurance companies, industrial corporations and public institutions. Established in 1968 and closely affiliated with the University of Frankfurt, it provides a strong link between the financial community and academia.

The CFS Working Paper Series presents the result of scientific research on selected topics in the field of money, banking and finance. The authors were either participants in the Center´s Research Fellow Program or members of one of the Center´s Research Projects.

If you would like to know more about the *Center for Financial Studies*, please let us know of your interest.

Prof. Dr. Jan Pieter Krahnen | Prof. Volker Wieland, Ph.D.

Center for Financial Studies
an der Johann Wolfgang Goethe-Universität ▪ Taunusanlage 6 ▪ D-60329 Frankfurt am Main
Tel: (+49)069/242941-0 ▪ Fax: (+49)069/242941-77 ▪ E-Mail: ifk@ifk-cfs.de ▪ Internet: http://www.ifk-cfs.de

CFS Working Paper No. 2001/07 - revised

How Wide Are European Borders? On the Integration Effects of Monetary Unions†

Guenter W. Beck††
Goethe University Frankfurt and CFS
gbeck@wiwi.uni-frankfurt.de

Axel A. Weber
Cologne University and CEPR
weber@wiso.uni-koeln.de

This Version: April 2003
First Version: January 2001

Abstract:
We use consumer price data for 81 European cities (in Germany, Austria, Finland, Italy, Spain, Portugal and Switzerland) to study the impact of the introduction of the euro on goods market integration. Employing both aggregated and disaggregated consumer price index (CPI) data we confirm previous results which showed that the distance between European cities explains a significant amount of the variation in the prices of similar goods in different locations. We also find that the variation of relative prices is much higher for two cities located in different countries than for two equidistant cities in the same country. Under the EMU, the elimination of nominal exchange rate volatility has largely reduced these border effects, but distance and border still matter for intra-European relative price volatility.

JEL classification: F02, F40, F41

Keywords: Relative price volatility, real exchange rates, law of one price, purchasing power parity, goods market integration, border effects, European Monetary Union (EMU), regional diversity.

†We are grateful for helpful comments from Manfred Neumann, Juergen von Hagen, Ignazio Angeloni and seminar participants at the European Central Bank and the annual EEA and ESEM meetings in Lausanne and Venice in 2001 and 2002. This paper is part of a CFS research program on 'Local Prices and Aggregate Monetary Policy'. Financial support by the CFS is gratefully acknowledged. Of course, the authors are responsible for any remaining errors.
†† Corresponding author. Tel.: ++49-69-24294126. E-mail: gbeck@wiwi.uni-frankfurt.de

1 Introduction

Recent research has aimed at improving our understanding of the magnitude and determinants of deviations from purchasing power parity (PPP) and the law of one price (LOOP). One branch of the literature estimates the half-lives of real exchange rates. For most countries and time periods, real exchange rates are found to be highly persistent, with deviations from PPP amongst industrialized nations having half-lives of several years. A second approach focuses on the comparison of movements in goods prices across national borders to price movements between different regions within a country. A seminal paper by Engel and Rogers (1996) finds that both distance and the border are significant in explaining relative price dispersion in fourteen U.S. and nine Canadian locations. They show that (i) relative price variability increases with distance within each country and (ii) U.S.-Canadian relative price variability is significantly larger than within-country variability. The authors provide a useful measure of how important the border is relative to distance the 'width of the border'. Their estimates suggest that crossing the U.S.-Canadian border is equivalent to 75,000 miles of distance, i.e., in order to generate the same degree of relative price volatility by distance within a country, the cities would have to be 75,000 miles apart. By this 'width of the border' metric, international failures of the LOOP are large.

The role of borders and geography has increasingly received more attention in economics and a number of recent papers have discovered evidence of such border effects for various alternative categories of goods[1] and for additional locations. Engel et al. (1997), Parsley and Wei (2001a) and Parsley and Wei (2001b) use data from North America, Asia and Europe to study intra-national, intra-continental and intra-planetary deviations from the LOOP, whilst Engel and Rogers (2001) and Hufbauer et al. (2001) focus exclusively on European locations. In all of these studies only a few intra-national locations are used and the prime focus is on national data with cities being identified as the nations capitals.

In this paper, we examine the importance of both distance and national borders between locations in determining the degree of the failure PPP and the LOOP in Europe. We employ both aggregated consumer price index (CPI) data and disaggregated data for ten categories of consumer goods. We make use of regional data available within Europe for seven West German, six East German, twenty Austrian, five Finnish, twenty Italian, eighteen Spanish, seven Portuguese and four Swiss cities. These data are taken from the SPATDAT databank,[2] which is by far

[1]See Crucini et al. (2000) and O'Connell and Wei (1997) for a broad range of goods prices.

[2]SPATDAT is a CFS databank with spatial consumer price data for sub-national regions/districts/cities from North American countries (U.S.A., Canada, Mexico), South American countries (Argentina, Brazil, Bolivia and Columbia), European countries (Austria, Finland, Ger-

the largest cross-sectional data set used in this literature for Europe to date.

The specific focus of the current paper is on the integration effects arising from the formation of monetary unions. For this purpose we study the German and European monetary unification process in more detail. With the fall of the Berlin wall in September 1989 formerly divided East and West Germany de facto became one single country. German Economic and Monetary Unification (GEMU) occurred shortly afterwards in July 1990, and in October 1990 political unification followed. Whilst these events jump-started a process of economic integration, it is interesting to examine at what speed economic convergence and market integration took place. The present paper looks at relative price volatility across German cities in comparison to Austrian and Swiss locations in order to determine whether or not an East-West gap (or shadow-border effect) persisted even during GEMU.

The second process of monetary unification we consider is the launching of the euro on January 1, 1999, when the currencies of the member countries of the European Monetary Union (EMU) became irrevocably fixed on their way towards eventually disappearing from circulation in January 2002. As in the case of GEMU above, the effect of the EMU on convergence and market integration will be studied by looking at the persistence of relative price volatility across 81 European cities in Germany, Austria, Finland, Italy, Spain, Portugal and Switzerland.

Our estimation equations are similar to the ones used in Engel and Rogers (1996) and Engel and Rogers (2001): The dependent variable is the variance of changes in the log of real exchange rate across cities, and among the explanatory variables are distance and border dummy variables. Since our European data set has city price data from several countries we are able to include, in addition to distance, 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 separately the role of nominal exchange rate variability and the effects of a border. Our results indicate that most of the failures of PPP/the LOOP are attributable to currency volatility in conjunction with rigid nominal prices, but other barriers are also important explanatory factors. We find that, even taking into account nominal exchange rate variability, distance between cities and the border continue to have positive and significant effects on real exchange rate variability. In the words of Devereux and Engel (1998) this shows that observed border effects are largely 'nominal' and only a smaller part is 'real'. We also show that including

many, Italy, Portugal, Spain and Switzerland), Asian countries (India, Indonesia, Japan, Korea, Malaysia, Philippines, Taiwan and Thailand) and 'Pacific' countries (Australia and New Zealand). Both aggregated CPI data and data for a large number of disaggregated categories of consumer goods have been collected. Regional coverage within Europe is fairly broad (twenty Austrian, five Finnish, up to thirteen German, eighty Italian, eighteen Spanish, seven Portuguese and four Swiss locations are available).

nominal exchange rate volatility in a multi-country setting to control for the effect of currency volatility on observed border estimates can lead to biases in the estimates of other included variables. Our data provide us with a more elegant way to assess the 'real' size of the border. When we split the sample period into a pre-EMU and an EMU subperiod we find that border estimates across EMU member countries drop drastically (by about 80%) after January 1999. However, also in the EMU border estimates remain highly significant across all countries. As nominal exchange rate volatility has been extinguished across EMU member countries these results indicate that real factors play an important role for observed market segmentations.
The rest of the paper is organized as follows. In the next section, we will describe our data set of regional European CPI data. In section 3 we report some descriptive statistics and in section 4 we will shortly describe our estimation approach. Section 5 examines the presence and relative size of border effects across major European countries. In section 6, we split our data sample into a pre-EMU and an EMU sample to examine the impact of the introduction of the euro on integration. In section 7, we examine how integration between West and East Germany has evolved after the re-unification. Section 8 concludes.

2 Data

To study the impact of monetary unions on observed border effects we have - in the spirit of Engel and Rogers (1996) and other studies mentioned above - compiled a large set of European regional consumer price data.[3] This data set contains both aggregated and disaggregated CPI data and comprises a total of 86 locations. Table A of section B lists these locations. As one can see there, we are using regional data from seven European countries, namely Germany (East and West), Austria, Finland, Italy, Spain, Portugal and Switzerland.[4] The data are monthly, the covered period for the total index is January 1991 to December 2002, the disaggregated data span the period from January 1995 to December 2002. In principle, we could use disaggregated data for all countries for which we have available total index data. There is, however, one important reason why we restrict our analysis of disaggregated data to a subsample of countries. Only for the EMU countries Germany, Finland, Italy, Spain and Portugal do our disaggregated data follow an identical classification scheme (the COICOP classification scheme). Though we also have data on subcategories for Austrian and Swiss regions available, these data follow a

[3]In the following, only a short description of the used data is given. For a more detailed description of the employed data, see section A.

[4]For the Benelux countries (Belgium, Luxembourg, the Netherlands), France, Denmark and Ireland we were unable to obtain sub-national data. Rather than following Engel and Rogers (2001) and using the national CPIs in those cases, we chose to exclude these countries from our analysis.

different classification scheme and are therefore not used here. As table A shows, all data are retrieved from official sources, data integrity should not be a problem therefore. The nominal exchange rates used to construct relative prices of regions that are separated by a national border were taken from the IMFs International Financial Statistics database. As the price data are usually collected throughout the respective sample period we use monthly averages instead of end-of-period data. One novelty of the present paper is the use of Austrian, Finnish and Portuguese data from locations within these countries and the use of East German in addition to West German regional data. Another novelty is the use of disaggregated data on consumer price indices for European cities.

For studying the impact of the EMU on European goods markets we make use of data from 81 out of the 86 available locations, excluding the East German data. Out of these data we can construct a total of 3240 ($= 81 * 80/2$) bilateral relative price series. Our sample of seven countries implies that the cross-border city pairs lie across one out of 21 ($= 7 * 6/2$) national borders (that are not necessarily adjacent). There are two types of exchange rate arrangements determining the nominal exchange rates of our 21 country pairs. Germany was at the heart of the Exchange Rate Mechanism (ERM) of the European Monetary System (EMS), and adopted a policy of fixed but adjustable exchange rates with Austria, Finland, Italy, Spain and Portugal during the sample. Each of these six countries was included in the first wave of entrants into the EMU, launched in January 1999. Furthermore, all of these countries participated in the free-trade area of the European Union (EU). Switzerland is the only country in our sample that has remained out of any formal arrangements on either exchange rates (ERM, EMU) or trade (EU).[5] In addition to the total index data we are using data on COICOP subcategories from 57 locations from Germany, Finland, Italy, Spain and Portugal. Analogously to our procedure for the total index we construct a total of $57 * 56/2 = 1596$ relative price series. The inclusion of five countries allows us to study the impact of the euro on 10 ($= 5 * 4/2$) different borders. Following Engel and Rogers (1996), we also use the disaggregated data to construct a measure for the 'width' of European borders. The analysis of subperiods shows us whether this 'width' has reduced since January 1999.

In the second part of the paper, we combine CPI data from six East-German lo-

[5]The seven countries used in this study also differ along geographic, linguistic, and cultural lines. In our sample, Austria and Switzerland share a common border with three of the other countries (Germany, Italy and each other), whilst Germany has two adjacent neighbors (Austria and Switzerland). Portugal and Spain only share a common border with each other and Finland does not share a border with any of the other countries in the sample. Note that our study takes explicit account of geographic considerations such as common borders or physical distance between locations. Finally, it is worth mentioning that common language factors may also matter. For example, German is spoken in three countries in our sample (Germany, Austria and Switzerland), while Finnish, Italian, Spanish and Portuguese are languages specific to these countries. Like geography, these cultural factors may also contribute to economic integration between countries.

cations with corresponding data from seven West German, twenty Austrian and four Swiss locations. Thus, this sample consists of a total of 37 locations based on which we construct 666 bilateral relative prices. By considering East Germany as an 'independent' country (and thus introducing a West-East German shadow border variable) we can study the dynamics of East German integration both relative to West Germany and the other German speaking countries.

3 Summary Statistics on Relative Volatility and Distance

To measure the degree of integration across goods markets, we follow the literature and take the standard deviation of monthly changes in bilateral relative prices as our base measure, also denoted as volatility measure 1, for integration.[6] Let q_{ij} denote the log of the CPI in location i relative to that in location j. All prices are denominated in the same currency. Then our measure for relative price volatility is obtained as the root of the sample variance, $V(\Delta q_{ij})$ of two-month changes in relative prices, Δq_{ij}. To get some intuition of the size and the regional dispersion of relative price volatility across European regions, table 1 presents some summary statistics. In the column denoted 'all', we report the mean value (and its standard deviation across locations) of relative price dispersion for all included location pairs. For the total index (denoted as 'allit') 3240 relative price measures are available, for the COICOP subgroups ('food', etc.) 1596 observations are included. The columns denoted as 'intra-nat.' and 'internat.' report means and standard deviations (across locations) for two subsamples. In the column 'intra-nat.', numbers for those regional pairs are reported where both locations lie within the same country. In the column 'internat.', summary results are reported for region pairs where the two locations lie in different countries.

When looking at the total sample (column 'all'), we see that there are large differences in reported values across individual goods categories: Relative dispersion is lowest for furniture (with a value of 9.59) and is highest for clothing (with a value of 32.19). The other categories (incl. the total index) lie in between these two values. When comparing intra-national and international means, we can observe that - for all goods categories - average relative price dispersion is considerably lower for intra-national region pairs than for international region pairs. In line with the existing literature, this indicates that intra-national European markets are more integrated

[6] In addition to this measure we test the robustness of our results by employing two alternative measures: The first is the spread between the 10th and the 90th percentile of the distribution of relative price changes (volatility measure 2). The second is the standard deviation of the one-period ahead in-sample forecast error from an estimated AR(6) process (volatility measure 3).

than international European goods markets, i.e., European national borders matter for integration.

However, there is another explanation. In the last row of table 1, we report the average distance (in km) across locations. Distance is used as a proxy for transportation (or more generally transaction) costs of goods arbitrage. Assuming that transportation costs prevent arbitrage from equalizing prices across locations, the observed differences in average relative price dispersion between intra-national and international location pairs could also be caused by differences in transaction costs of arbitrage. As the last row of table 1 shows, average distances between locations within a country are drastically lower than between locations that are separated by a national border.

In tables B and C we provide further descriptive statistics. In these two tables, we report detailed results on the volatility of relative prices for all bilateral country pairs that are included in the respective samples. Table B contains descriptive statistics for 28 groups of location pairs from our sample of total index data. These 28 groups consist of seven groups of within-country city pairs (ge-ge, au-au, fi-fi, it-it, sp-sp, po-po and ch-ch[7]) as well as 21 groups of cross-border city pairs (ge-au, ge-fi, ge-it, ge-sp, ge-po, ge-ch, au-it, au-fi, au-sp, au-po, au-ch, fi-it, fi-sp, fi-po, fi-ch, it-sp, it-po, it-ch, sp-po, sp-ch and po-ch). As we saw above, the average volatility of cross-border pairs is typically considerably larger than the average variance of within-country pairs. The within-Germany city pairs exhibit the lowest average volatility (1.93). Relative price volatility is slightly higher in Finland, followed by Switzerland, Italy, Spain, Austria and Portugal. Note that the volatility of relative prices across Portuguese cities, equal to 5.66, is even slightly higher than the German-Austrian cross-border volatility (5.53), but except for this one case we typically find that within-country volatility is considerably lower than the average relative price volatility of the cross-border city pairs.

Columns three, four, seven and eight of table B display our measures for distance and nominal exchange rate volatility. There is an obvious correspondence between relative price volatility, distance and nominal exchange rate volatility. Looking at the relation between relative price volatility and distance we can see that the more volatile cross-border city pairs are typically more distant than the within-country pairs. However, there are many cases where higher distance is not related with higher relative price dispersion. Particularly for all bilateral Swiss country pairs and the German/Austrian/Finnish-Italian location pairs, relative price volatilities are much higher than average distances between these countries would suggest (given the evidence from other countries). Thus, we can already conclude that other factors than transaction costs of arbitrage will play a role in explaining relative price volatility

[7] The used country short names are explained in table A.

across locations.
One prime candidate for explaining the observed relative price volatility patterns is nominal exchange rate volatility in conjunction with short-run rigid national prices. This variable particularly seems to be the perfect candidate for explaining the high relative Swiss volatilities. Although Switzerland is a direct neighbor of Germany and Austria, its relative price volatility with these countries is by far larger than that of these countries with Finland, Spain or Portugal which are much further apart. However, as Switzerland is not a member of the ERM (nor the EU) nominal exchange rate volatility is not bound by any formal exchange rate arrangement and thus can be expected to be of considerable size.[8] A comparison of real and nominal exchange rate dispersion in table B confirms that the volatility of nominal exchange rate changes, Δs_{ij}, in many cases is closely linked to relative price variability for cross-border city pairs. This result is well established in the literature.[9] As changes in real exchange rates (Δq) are the sum of changes in the nominal exchange rate (Δs) and the change in relative national prices ($\Delta p - \Delta p^{\star}$), in the presence of short-run rigid national prices we would expect the change in the real exchange rates to be equal to the change in nominal exchange rates. In our sample, results from table B illustrate very close links between real and nominal exchange rate volatility for all cross-border combinations for Finland, Italy, Spain, Portugal and Switzerland. However, for Germany the two-month average volatility of cross-border relative price changes vis-a-vis Austria is 4.11, whilst the average volatility of nominal exchange rate changes is 0.41, which is 10 times smaller. Additionally, the link between two-month changes in real and nominal exchange rates is not perfect. In most cases, nominal exchange rate volatility is 10% or more smaller than real exchange rate volatility which is even true for bilateral Swiss location pairs. This suggests that nominal exchange rate volatility alone cannot explain relative price volatility. It additionally points to a problem that the existing literature on goods market integration has had but hasn't been able to deal with adequately: As table B shows, nominal exchange rate volatility (in conjunction with short-run rigid prices) plays a key role in explaining cross-country relative price dispersion. However, none study has thus far been able to show whether estimated border effects will vanish after having adequately controlled for nominal exchange rate volatility. When turning to our regression results for the EMU subperiod (1999.01 - 2002.12) we will be able to provide a satisfactory solution to this problem as nominal exchange rate volatility does no longer play a role across countries that adopted the euro.
The detailed descriptive statistics for the COICOP subcategories (table C) show

[8]The corresponding width-of-the-border measure may therefore be interpreted to be a welfare measure for Switzerland staying out of the EU and the EMU.
[9]See, e.g., Mussa (1986) for reference.

basically the same patterns of relative price volatility across European countries that we found for the total index. However, the link between relative volatility and nominal exchange rate volatility is much loser for COICOP subcategories than for the total index. This suggests that the link between relative price and nominal exchange rate volatility is less obvious than commonly expected. This finding additionally provides evidence for the hypothesis that nominal exchange rate volatility alone cannot explain the existence of border effects across national markets. In order to sort out the relative influence of these factors quantitatively, we now turn to our regression evidence.

4 Methodology

Engel and Rogers (1996) and others examine the hypothesis that the volatility of the prices of similar goods sold in different locations is related to the distance between the locations and other explanatory variables, including a dummy variable for whether the cities are in different countries. In the analysis below we present the results of estimating regression equations of the form:[10]

$$V(\Delta q_{ij}) = \sum \alpha(c)D(c) + \beta \ln(d_{ij}) + \delta B_{ij} + \gamma V(\Delta s_{ij}) + u_{ij}, \tag{1}$$

where $D(c)$ is a dummy variable for each city in our sample, d_{ij} is the distance between cities i and j, B_{ij} is a dummy variable for each national border that separates cities i and j, and $V(\Delta s_{ij})$ is a measure of nominal exchange rate volatility between cities i and j located in different countries. Note that all regressions are cross-sectional, with 3240 observations when total index data are employed and 1596 observations when COICOP subcategories are used. The inclusion of separate dummies for each individual location allows the variance of price changes to vary from city to city. That is, for city pair (j,k) the dummy variables for city j and city k take on values of 1. There are a few reasons why we allow the level of the standard deviation to vary from city to city. First, there may be idiosyncratic measurement error or seasonalities in some cities that make their prices more volatile on average. Second, as tables B and C indicate, there are notable differences in average within-country volatilities across countries. As, e.g., table B shows average volatility for Austrian and Portuguese cities than for German or Finnish cities. This may be because Portugal and Austria are more heterogeneous countries. Either labor markets

[10] In our sensitivity analysis, we also report results when a quadratic distance function is employed instead of the log of distance, i.e., when the regression equation takes the form:

$$V(\Delta q_{ij}) = \sum \alpha(c)D(c) + \beta_1 d_{ij} + \beta_2 d_{ij}^2 + \delta B_{ij} + \gamma V(\Delta s_{ij}) + u_{ij}.$$

or goods markets may be less integrated, so there can be greater discrepancies in prices between locations. Alternatively, there may be differences in methodologies for recording prices that lead to greater discrepancies in prices between locations in one country compared to the other.

Following Engel and Rogers (1996), we assume that relative price volatility will be larger the greater the distance d_{ij} between locations, due to transportation costs of arbitrage. The key argument here is that in the presence of transportation costs prices in one location are not necessarily equalized with prices in another location, and that the relative price could fluctuate in a range which is likely to be a function of the transportation cost and hence the distance between the locations. Equation (1) thus postulates that more distant locations would have greater price dispersion and that this relation is log-linear ($\beta > 0$). In our sensitivity analysis, we also employ a quadratic distance function that postulates a concave relationship between distance and relative price volatility. As mentioned several times, we interpret 'transportation costs' liberally to include any factors that make it more costly to sell goods in one location compared to another.[11]

We are furthermore particularly interested in whether there is a border effect. 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 and nominal exchange rate volatility. The recent literature on pricing-to-market has examined markets that are segmented by borders. There are a few reasons why the border might matter. Much of the pricing-to-market literature has emphasized that the mark-up may be different across locations and may vary with exchange rate changes. There might also be direct costs to crossing borders because of tariffs and other trade restrictions. In addition, there may be more homogeneity in relative productivity shocks for city pairs within the same country than for cross-border city pairs, so that, from equations (1) cross-border pairs have more price volatility. Another important reason why the border matters is unrelated to equation (1): The price of a consumer good might be sticky in terms of the currency of the country in which the good is sold. Goods sold in Germany might have sticky prices in German mark terms, and goods sold in Italy might have sticky prices in Italian lira terms, whilst the nominal exchange rate is highly variable. In this case, the cross-border prices would fluctuate along with the exchange rate, but the within-country prices would be fairly stable. To capture this effect, we include a border dummy variable, B_{ij}, that takes on a value of unity if cities i and j are in different countries. This border dummy is likely to capture both formal and informal international barriers to trade.

[11] For example, there may be trade barriers or marketing and distribution costs.

5 Evidence on the Relative Size of Border Effects Across European Countries

In the spirit of Engel and Rogers (1996), the first version of equation (1) that we are testing takes the form

$$V(\Delta q_{ij}) = \sum \alpha(c)D(c) + \beta \ln(d_{ij}) + \delta B + u_{ij},$$

i.e., we are using an aggregate border dummy for all international location pairs and do not involve any other variable (apart from distance) that might have an explanatory power for observed volatilities in relative prices. Engel and Rogers (1996) use this specification to derive a measure for the 'width' of the U.S.-Canadian border. They find that the U.S.-Canadian border has an average width of 75,000 miles. This result has found considerable attention in the literature and since then forms a benchmark measure for any study that examines the degree of integration across markets. However, there are some caveats to take into account. First, the log-linear distance specification employed makes results very sensitive with respect to even small changes in either the estimate of the distance and/or border coefficient. In this line, Engel and Rogers (1996) show that the border width declines to only 1,780 miles when the upper end of the confidence interval on the point estimate of the distance coefficient is used as the measure of the impact of distance. Secondly, we cannot be sure that a log-linear specification appropriately reflects the true relationship between relative price volatility and transportation costs. Thirdly, the use of distance as a proxy for transportation costs only works if a relationship between distance and actual transportation costs exists (which is very likely) and if this relationship remains relatively stable across the respectively considered sample period. Particularly the latter point is problematic as transportation costs across locations are very likely subject to time-varying influences (such as different environmental legislations in different countries) whereas distance between locations remains constant across time.

The upper panel of table 2 presents results from estimating the above equation using the total index data. The result on distance shows that transportation costs have both a significant and a positive impact on integration of markets. However, its influence is considerably smaller than that of the border (by a factor of 26). The coefficient on the border is also highly significant in illustrating that national European markets are, despite intensive political and economic efforts in the past, strongly segmented internationally.

Looking at the results for the COICOP subcategories, we can see that in most cases distance coefficients are positive and significant. Additionally, the border coefficients are again not only strongly significant but also much larger in size than the distance coefficients. This shows that national borders have relatively more importance for the volatility of relative prices across locations that are separated by a national border than transportation costs do.

While the distance coefficient is positive and significant in most cases, for two subcategories (communication and recreation) the coefficient on distance is either insignificant or even negative. There are two possible explanations. First, in both categories the relative share of nontradeables is relatively large and is considerably higher than, e.g., for food, alcoholic beverages or clothing. As we expect that costs of arbitrage only play a role for tradeables but not for nontradeables it is not surprising that for categories such as recreation the relative importance of the distance variable is smaller. The second reason is the following: To obtain the results in table 2, all bilateral border variables are forced to be equal. As the detailed descriptive statistics for relative price volatilities in table C show, there are good reasons to assume that the size of the border effect differs considerably across country pairs. Forcing these effects to be equal across country pairs can lead to biases in the estimate of the distance variable. As we will see below, all distance coefficients turn positive when we allow for heterogeneity across estimated border effects.

To check the sensitivity of our results, we use two alternative measures of relative price volatility. Tables 3 and 4 report the regression results when these two measures are employed. The measure denoted as 'volatility measure 2' is computed as the spread between the 10th and the 90th percentile of the distribution of two-month changes in the relative price between two locations. The results for volatility measure 2 (see table 3) basically confirm the results obtained for volatility measure 1. For all goods categories the border coefficients are highly significant indicating strong segmentations across national markets. Distance coefficients are positive and significant for the more traded goods categories whereas they are partly insignificant or even negative for the categories with a relatively high share of nontradeable goods. Volatility measure 3 is obtained as the standard deviation of the two-period ahead in-sample forecast error of each relative price series. The forecast is based on an estimated AR(6) model. Results are presented in table 4 and are also similar to those obtained for volatility measure 1. Summarizing the results from all three measures shows, that European markets are segmented considerably.

What explains these border effects? Nominal exchange rate variability in conjunction with sticky national prices presents a prime candidate. Analogous to the construction of volatility measure 1, we compute nominal exchange rate dispersion as the variability of two-month nominal exchange rate changes, which of course is zero for all intra-national pairs. The results from including nominal exchange rate volatility as an explanatory variable are reported in table 5. For our overall sample, the coefficient on nominal exchange rate variability is 0.85. Including nominal exchange rate variability substantially weakens the effect of the border dummy, whose point estimate falls from 4.39 to 0.67. This suggests that a very large part of the border effect stems from variable nominal exchange rates under sticky prices. However, even with $V(\Delta s_{ij})$ in the regression, the border dummy remains positive and significant. These results are in line with our interpretation of the descriptive statistics where we found a close, but not perfect relationship between nominal and real exchange rates. A somewhat worrisome result in table 5 is that the distance estimate becomes significantly negative. The reason for this result is the same that we already mentioned above: The role that bilateral exchange rate volatility plays for bilateral real exchange rate volatility differs considerably across country pairs. When forcing the impact of nominal exchange rate volatility to be equal for all country pairs, biases in other variables result. This can be seen when we look at specification 2 of table 5 where we include an additional variable for all Italian bilateral nominal exchange rates. As the results from specification 2 the impact of the border and its significance increase considerably and the coefficient on distance turns significantly positive again when we control for Italian exchange rate volatility.

One way to capture the heterogeneity in border effects across European countries is to include individual border dummies for each included country pair. Table 6 reports results for individual border estimates when total index data are used. The table contains results from four different estimations. The first two columns present the results of regressing volatility measure 1 on log distance, 21 borders, and 81 individual location dummies (one for each of our cities, not reported for reasons of convenience). All coefficients have the anticipated sign and are significant at least at the five percent level. The coefficients on the border dummies range between 2.19 (t-stat 33.21) for the German-Austrian border to 27.78 (t-stat 160.21) for the Italian-Swiss border, which is more than ten times as large. The largest border effects are found for bilateral Swiss real exchange rates and for the 'Northern' European countries (Germany, Austria and Finland) relative to Italy. Summarizing, our findings confirm the results documented by Engel and Rogers (1996) and Engel and Rogers (2001): Crossing an international border adds considerable volatility to relative city prices, even after accounting for the effects of distance and city-specific characteristics.

Table 6 also displays the results obtained when the distance function is quadratic, rather than logarithmic. This is reported in columns three and four, which is interesting because it allows a test for our assumption of a concave distance relationship. We find that distance has a significantly positive effect on price variability, whilst the square of distance has a significantly negative effect, as is postulated by a concave distance relationship. Again border dummies are positive and significant and

prevail the same country pattern as described above. The results in columns five to eight of table 6 show that our main results are not affected when volatility measures 2 or 3 are employed. We again find that the coefficients on distance and the border dummy are highly significant and of the hypothesized sign.
In table 7, we report evidence on relative border sizes for all those countries for which we have available data on COICOP subcategories. In all cases, the border variables are highly significant. Distance variables are mostly positive, but are not always significant.
What impact is the EMU likely to have on the importance of borders? Will national borders still matter in the EMU? In order to analyze the impact of the elimination of exchange rate volatility on the significance of European borders we now turn to the analysis of the subperiods.

6 Pre-EMU and EMU Subperiod: Has the EMU Reduced the Size of European Borders?

To study the effects of the EMU on the size of the estimated border coefficients we basically repeat the above analysis for two subperiods: 1995.01-98.12 and 1999.01-2002.12. The subperiods correspond to the late ERM and the early EMU period. The first period includes the aftermath of the foreign exchange market turbulence during the ERM crisis (1992.09-1993.07), when major exchange rate movements took place and Italy temporarily withdrew from the ERM. Before we start with our analysis one issue should be noted. The availability of regional data for the EMU period allows us to perform the first study - at least to our best knowledge - that examines the existence and size of border effects in an international environment "without trade barriers or currency fluctuations" (see Parsley and Wei (1996). Thus, instead of relying on measures for nominal exchange rate volatilities to control for the importance of nominal exchange rates in conjunction with sticky nominal prices, we can directly assess the size of the real border effect.
Some graphical evidence on the impact of the EMU on relative price volatility is given in figure 1. In this graph, we plot mean values of the relative price dispersion across country groups for the pre-EMU subperiod (y-axis) versus their corresponding values for the EMU subperiod (x-axis). We notice several characteristics of the data: First, intra-national (within-country) relative price volatility is low prior to the EMU and does not decline significantly during the EMU. Secondly, international relative price volatility is high prior to the EMU and particularly pronounced between Northern European countries (Germany, Austria and Finland) and Italy and is relatively low between the North European countries (Germany, Austria and Finland). Bilateral combinations involving Switzerland always lie above the cor-

responding combinations with EMU countries. Thirdly, international relative price volatility falls drastically for all EMU cross-border city pairs in the second subperiod. The EMU effect has been particularly strong for formerly quite volatile Southern European exchange rates, whilst for relatively stable exchange rates there has only been a minor effect. Finally, in addition to the strong decrease of *international* relative price volatility within the EMU, the data also reveal a sizeable reduction in relative price volatility between Switzerland and the EMU. This convergence process may be due to a deliberate policy of shadow-targeting the euro exchange rate by the Swiss National Bank.[12]

Figure 2 takes a closer look at all 3240 cross-city volatility measures in both the pre-EMU (panel a) and EMU sample (panel b). The scale of both graphs is chosen to be the same, so that the reduction of relative price volatility for all cross-border city pairs is more directly visible. In panel (b) of Figure 2 it is impossible to discriminate visually between within-country and within-EMU relative price volatility, whereas the EMU-Swiss city pairs are still clearly identifiable as having higher volatility. As a first approximation one may therefore be tempted to conclude that the EMU has eliminated international differences in relative price volatility between EU cities. The formal analysis below will show that this conclusion is not valid and that national borders continue to matter for relative price volatility even in the EMU.

Table 8 displays individual border estimates for the log-linear and quadratic distance function specifications for our two subperiods when total index data are used. In both cases, the regression coefficients on distance have the correct hypothesized signs but are insignificant in the second subsample. This may be due to the problems with our distance measure, the large share of nontradeables in the total index (in conjunction with a declining importance of transaction costs in the EMU) or due to a strong decline in transaction costs of arbitrage in the EMU as a consequence of increased price transparency. A second interesting feature of table 8 is the significance of all border dummies in both the pre-EMU and EMU sample. We find sizeable and significant border effects for all country-pairs. However, estimated sizes have decreased dramatically in most cases. In the pre-EMU subperiod, we can observe large differences in estimated coefficients: The smallest coefficients are found for the Northern European countries Germany, Austria and Finland (3.08 for ge-au, 9.41 for ge-fi and 8.77 for au-fi) and between Spain and Portugal (7.66). The highest coefficients are found for all northern countries (Germany, Austria, Finland and Switzerland) via Italy. We already saw that for the Italian case, these results are due to a large 'nominal' share in the overall border estimate. Other border

[12] Engel and Rogers (2001) find a sizeable reduction in intra-national relative price volatility in an earlier sample, and attribute this decline to an increased economic integration within countries, which is likely to be caused by advancements in transportation, communication, etc.

estimates lie in between these extremes. In the EMU period, all border estimates (also relative to Switzerland) dropped. For EMU member countries all values now lie between a value of close to or even below 1 (au-it, it-po, au-po and ge-au) and somewhat above 3 (ge-fi, ge-sp, ge-po, fi-sp and fi-po). The most drastic reduction (by 80 percent and more) in the estimated border effects are found for the Southern European periphery (Italy, Spain, Portugal) relative to the Northern European core-countries (Germany, Austria and Finland). Although relative Swiss border effects have decreased as well, their relative reduction is markedly smaller (by around 50% for ge-ch, au-ch, fi-ch and sp-ch). On average, border estimates have decreased by around 75%. These results are basically unchanged when the quadratic distance specification is considered.

To summarize, the shift in the variance of the nominal exchange rate from the ERM to the EMU sample can be interpreted as an exogenous event, as part of a political process that ultimately led to European Monetary Unification. Relative price volatility has fallen as a result of the decline in the nominal exchange rate variance. We find that during the EMU distance is notably smaller and insignificant. The estimated border coefficients are still positive and significant, but are less than one-fifth their size in the pre-EMU sample. This suggests that a common monetary policy leads to greater market integration of regions within countries and between countries.

We now turn to the analysis of COICOP subcategories. The results for individual border estimates are given in tables 9 to 13, summary results are given in table 14. The general pattern of the results is the same as that for the total index data: In most cases the coefficient for distance is both positive and often significant in both subperiods. In almost all cases estimated border coefficients drop drastically across subperiods, but remain highly significant in both subperiods. In the second subsample (EMU subperiod) only the border dummy for Italy relative to Germany becomes insignificant for clothing. As for the total index, the dispersion across estimated border coefficients generally declines in the second subperiod. In table 14 we present average border and distance estimates for the COICOP subcategories. Several issues are noteworthy: First, the average distance estimates do not change largely across subperiods. This indicates are more or less stable relationship between distance and relative price volatility. Secondly, when comparing average border estimates, we see that there is a considerable decline for all borders. Looking at the variation of estimated border coefficient across country pairs one can see that - despite some ‘outliers’ such as ge-it and it-po - integration across European countries is relatively equally spread.

7 Regression Results for GEMU

The above results for the EMU were obtained for a relatively small sample (1995 - 2002) and the 48 months of data from the EMU subperiod reveal only a limited amount of information about the long-run effects of the monetary union on within-EMU relative price dispersion. In particular, we are unable to obtain formal evidence about the speed of relative price convergence amongst European cities because estimating dynamic time series models on such few data points is unlikely to yield reliable results about these long-run effects.
An interesting feature of our data set is that it contains West and East German data for the entire 12-year period of the GEMU. Can we learn anything about the long-run dynamic effects of the EMU on relative price dispersion from the German experience? In order to find out, we estimated all our regressions from above for this sample as well. As a control group we added Austrian and Swiss cities to the sample. We think that analyzing such a homogeneous sample of adjacent countries with a common language and long-standing political, cultural and economic linkages will provide an estimate for the upper bound of the speed of convergence which we may realistically expect from the EMU. To estimate the direct effects of the GEMU on East-West German integration we include a shadow-border in the form of a East-West German border dummy in all regressions.

7.1 Cross-Sectional Evidence on Intra-German Border Effects

Table 15 displays the estimates of equation (1) with volatility measure 1 for the overall GEMU period (1991.01-2002.12) and for three GEMU subperiods (1991.01-1994.12, 1995.01-1998.12, 1999.01-2002.12). In equation (1) the regression coefficient on log distance is significant in the overall period and the first and second subperiods, but insignificant in the third subsample. This result is in line with the evidence found for EMU where the distance coefficient also becomes insignificant for the EMU subperiod when total index data are used. As outlined there, this might be due to a decrease in transaction costs of arbitrage, our insufficient measure of these transaction costs or the large proportion of nontradeable goods in the total index (in conjunction with decreasing importance of transaction costs). A second feature of table 15 is the significance of the East-West border dummy in the overall period and both pre-EMU subperiods. Thus, even after the start of the monetary union in Germany there was a significantly different pattern of relative price changes amongst cities within each part of Germany in comparison with cities located across the former 'iron curtain'. This is most likely the result of slow price deregulations and a gradual unlocking of formerly administered prices for housing, rent and certain goods in East Germany. By 1999, much of this price deregulation between East and West Germany appears to have been completed and the shadow-border is no longer significant. An interesting characteristic of our results for the immediate post-unification period is that our measure of economic integration suggests that West Germany, Austria and Switzerland were by far more integrated with each other than with East Germany. Integration of East and West Germany proceeded speedily during 1995-1998, when the East-West German border effect fell by over 90 percent from 25.31 to 0.79, as compared to the minor reduction of the German-Austrian border effect from 3.60 to 3.06 for the same period. By this metric, the two parts of Germany became four times more integrated during the 1990s than Germany and Austria did in spite of a long history of virtually no exchange rate volatility between the two countries. Tables 16 and 17 provide analogous results when the degree of goods market integration is determined using measures 2 and 3. As one can easily see all the results cited above remain valid: The West-East German border variable is highly significant for the first and second pre-EMU subperiod and becomes insignificant in the EMU subperiod. Additionally, there is a drastic drop in its value between the first and second subperiod. Furthermore, results indicate that - at least in terms of relative price volatility - West and East Germany are relatively well integrated.

7.2 Time Series Evidence on Intra-German Goods Market Integration

7.2.1 Single-Equation Time Series Evidence

In the above analysis we have identified an EMU-effect that is equal to an 80 percent reduction in intra-EMU relative price volatility for core-Europe relative to the southern periphery and in our GEMU sample we even found a reduction of intra-German relative price volatility that exceeded 90 percent. Both monetary unions therefore resulted in impressive integration effects. Like in the convergence regressions popularized in the growth literature, a low initial degree of economic integration thereby appears to be succeeded by a more rapid convergence progress. In order to examine this proposition more formally, the following analysis makes explicit use of the time series dimension of the data.

Instead of running a cross-section regression with 666 (=37*(36/2)) city pairs for the German-Austrian-Swiss sample we constructed 666 time series of two-month relative price (real exchange rate) changes between our 666 city pairs. For each of these time series we then ran an Augmented Dickey Fuller (ADF) unit-root test for the overall sample period (1991.01- 2002.12) by regressing the change of the real exchange rate on its past level and six difference terms. Instead of reporting here the 666 AR(1) coefficients, figure 3 displays the kernel density estimates of these AR(1) coefficients for the various intra-national and international city pairs. Within-country AR(1) coefficients are typically quite dispersed and skewed towards unity. The lowest AR(1) coefficients and hence the highest convergence speeds are found for West Germany/Austria and Switzerland relative to East Germany. All AR(1) coefficient estimates relative to East Germany have a relatively narrow density distribution, around a mean value of less than or equal to 0.8. Table 18 summarizes these coefficient estimates. The half-lives implied by these coefficients are between 5 and 72 months when unadjusted coefficients are used and between 5 and 84 months when Kendall bias-adjusted coefficients are used. The estimated speed of convergence is very low within countries for West Germany, Austria and Switzerland. Since relative price volatility was found to be smallest within countries as well, we conclude that most price convergence within countries has already been achieved in the past and that a further convergence is unlikely. For East Germany, however, convergence speed is much higher. This is probably due to the fact that East Germany faced quite different initial relative price movements and hence displayed the largest speed of relative price convergence. Looking at international bilateral combinations we can see that again for all relative prices between West Germany, Austria and Switzerland observed adjustment speeds are relatively low. For combinations, however, that include East Germany convergence speeds are very fast.

7.2.2 Panel Evidence

A major problem with the evidence presented above is that averaging over a large number of independently estimated AR(1)-coefficients may only yield a very imprecise picture of the convergence properties of relative prices within a monetary union. Furthermore, with only 12 years of data the power of such ADF-based tests in discriminating an AR(1)-coefficient close to unity from a unit root is known to be low. Pooling the cross-section data and performing a panel unit root test has been shown by Levin and Lin (1992), Levin and Lin (1993), Levin et al. (2002), Oh (1996) and Wu (1996) to increase the power of such tests considerably. In this section we will briefly discuss the convergence properties of relative prices in our GEMU sample found by running panel unit root tests. In doing so we use two different approaches: We will start by presenting evidence for the Levin-Lin (LL) test and will then proceed to results from the Im-Pesaran-Shin (IPS) test.

When constructing relative prices we choose one the capital city for each country as base region/city. To conduct the LL panel unit root test, the raw data are first transformed by subtracting the time-specific mean for each panel of relative prices. Let $\tilde{q}_{i,t}$ denote the transformed relative prices.[13] In order to correct for possible se-

[13]Note that there is a switch in notation. This is supposed to indicate that the sample of relative prices underlying the panel unit root analysis is not identical to the sample used in previous sections

rial correlation we employ the ADF method, and the estimation equation employed is given by:[14]

$$\Delta\tilde{q}_{i,t} = \alpha_i + \rho\tilde{q}_{i,t-1} + \sum_{j=1}^{k_i} \phi_{i,j}\Delta\tilde{q}_{i,t-j} + \epsilon_{i,t}. \tag{2}$$

As our goal is to compare long-run deviations from PPP across various markets we construct several panels of relative price series that correspond to the regional grouping that we have presented above for the individual AR(1) coefficients. In doing so, we choose one base city - the respective capital - for each country. Results for the LL test are presented in table 19. The reported p-values for the adjustment coefficients were obtained by a nonparametric bootstrap as described in section D.2 of appendix D. Note that we include a shadow East-West German border to study the convergence speed of relative prices between cities located in the two formerly separated parts of Germany. The pattern of the results is similar to that for the individual AR(1) coefficients. However, panel-based estimates of convergence speeds turn out to be much higher. Again, we find no signs of mean reversion in relative prices within West Germany, whilst East-West relative prices converged with a half-life of around two years. Convergence within East Germany turns out to be very low (with a half-life of almost ten years) and not significant. Finally, between the West German cities and Austrian or Swiss cities we find no or only weak indications of mean reversions. These last numbers roughly correspond to the evidence about the slow speed of real exchange rate convergence between industrialized countries reported in the introduction of the present paper. Our estimates of a slow rate of relative price convergence within a country are consistent with similar estimates provided by **?** for the U.S. economy. They study price level convergence among U.S. cities and find that relative price levels mean revert, but do so at a surprisingly slow rate. In a panel of 19 cities they estimate the half-life of convergence to be approximately 9 years. We find very similar results for our sample. **?** conclude that their estimates for the U.S.A. provide an upper bound on speed of convergence that participants in the EMU are likely to experience. Our results support these conclusions.

To check robustness of these findings we also employ IPS panel unit root tests. The results are presented in table 20. As for the LL test, we choose one city - the capital - for each country as a base city when computing relative prices. Adjustment coefficients (ρ) are computed as averages over all individual adjustment coefficients of the respective sample. Bias adjustment (for ρ_{adj}) is done using the formula by Kendall (1954). Mean t-statistics (t^*) are reported, their respective p-values are obtained by

due to our choice of a base city/region for each country.

[14] See appendix D for a detailed description of the estimation and bootstrapping approach followed to obtain the estimation results.

employing a nonparametric bootstrap. Results basically mirror those of the LL test: Only for East German relative to West German/Austrian and Swiss relative prices the null hypothesis of a unit root can be rejected. Half-lives for these cases lie in the range between 1 1/2 and 2 1/2 years. All other relative prices exhibit much lower (or no) mean reversion, the null hypothesis of non-stationarity cannot be rejected for them.
To summarize, the results of this section show that there is strong mean reversion between East German and West German/Austrian and Swiss relative prices but there are no indications for mean reversion between West German, Austrian and Swiss relative prices. Does these results suggest that West German, Austrian and Swiss goods markets are very segmented although these countries (particularly West Germany and Austria) not only share close economic links but also a common language? Not necessarily. There is an much for favorable interpretation of the results. It is based on the cross-sectional evidence from above that showed that the West German-Austrian border is relatively close when compared to other countries. How can this evidence be put in concordance with the missing evidence of significant mean reversion from the long-run unit root tests? The answer is relatively simple: Assume that transaction costs of arbitrage create a band of inactivity across the equilibrium real exchange rate. As transaction costs between West Germany and Austria are supposedly small, the width of this band and thus the observed average relative price dispersion is also small. When additionally relative prices do not cross this band of inactivity, then our long-run test would indicate no mean reversion. This, however, would be not a sign of segmentation but of relative strong integration.

8 Conclusions

The major message of our empirical results is that the elimination of nominal exchange rate volatility in a monetary union will give a major boost to economic integration by significantly reducing cross-border relative price volatility. However, moving to a common currency neither immediately nor in the long-run completely eliminates cross-country relative price volatility. Even in a monetary union national borders and distance continue to be important determinants of relative price volatility. Looking at both the German unification experience and the early phase of EMU we are able to establish that relative price convergence is likely to occur relatively fast. The half-life of the East-West German price level convergence is estimated to be between 1.5 and 2 years. We expect somewhat higher convergence speeds for the remaining European countries. Our evidence for the first years of the EMU suggests that price level convergence has already occurred to a large extent and that roughly 80 to 90 percent of the initial relative price dispersion has been eliminated by now. The literature on pricing-to-market has emphasized that, when markets are segmented, price discrimination can occur. The finding that distance is important in explaining price differences between locations in Europe lends support to this literature. The EMU is found to have greatly reduced but not completely eliminated the importance of intra-EMU borders. Our width-of-the-border metric suggests that due to EMU European locations have grown closer together. However, the results of this paper also confirm the finding of Engel and Rogers (2001) that despite more price transparency under a common monetary policy and the complete absence of the intra-EMU trade barriers, European product markets are still segmented.

9 Tables

Table 1: All Items and Subcategories, Descriptive Statistics, Overall Period, Volatility Measure 1, Summary Results

Categories	Country Groups		
	all	intra-nat.	internat.
allit	14.19	3.46	16.58
	(7.62)	(1.27)	(6.26)
food	12.98	7.03	14.94
	(4.44)	(2.83)	(2.85)
alco	15.23	6.23	18.27
	(7.05)	(7.20)	(3.53)
clot	32.19	9.58	39.62
	(22.24)	(10.58)	(19.93)
hous	12.34	7.62	13.89
	(3.57)	(2.57)	(2.24)
furn	9.59	4.97	11.11
	(3.28)	(1.59)	(2.03)
heal	12.97	7.32	14.83
	(5.90)	(3.21)	(5.38)
tran	11.43	5.29	13.45
	(4.35)	(2.94)	(2.41)
comm	14.57	2.59	19.83
	(8.99)	(3.02)	(4.65)
recr	18.70	9.51	21.73
	(7.57)	(4.18)	(5.77)
hote	14.16	8.62	15.99
	(6.41)	(3.51)	(6.09)
dist	1218	365	1409

Notes:
1) Relative price dispersion between region i and region j, denoted as $V(q_{ij})$, is computed as the standard deviation of two-month relative price changes between the two regions, i.e.,

$$V(q_{ij}) = \sqrt{var(\Delta q_{ij,t})},$$

where $\Delta q_{ij,t}$ denotes the two-month change in regions' i and j relative price and $var(.)$ denotes the empirical variance of $\Delta q_{ij,t}$.
2) In the column termed 'intra-nat.', the mean relative price dispersion (and its standard deviation across regions) of all regional pairs where both regions are located in the same country is reported. In the column termed 'internat.', the mean relative price dispersion (and its cross-regional standard deviation) of all regional pairs where both regions are located in different countries is reported.
3) There are 3240 observations for allit (81 regions) and 1596 observations (57 regions) for the subcategories.
4) A description of the employed short names for the COICOP subcategories is given in section A of the appendix.

Table 2: All Items and Subcategories, Aggregate Border Estimates, Overall Period (1995.01 - 2002.12), Volatility Measure 1

Category	(ln)dist		border		R^2	R^2_{adj}	*s.e.r.*
	Coeff.	t-stat	Coeff.	t-stat			
Total Index							
allit	0.50	5.61	6.94	29.98	0.864	0.861	0.0028
COICOP Subcategories							
food	1.28	12.01	5.19	31.09	0.893	0.889	0.0015
alco	0.84	5.25	11.68	35.92	0.901	0.898	0.0023
clot	4.59	8.87	16.99	20.56	0.907	0.903	0.0069
hous	0.87	8.14	5.09	39.26	0.879	0.875	0.0013
furn	1.07	13.32	3.91	31.98	0.886	0.882	0.0011
heal	1.04	8.11	4.48	16.55	0.905	0.901	0.0019
tran	0.26	2.88	6.87	54.01	0.917	0.914	0.0013
comm	0.15	1.41	14.80	96.92	0.982	0.981	0.0012
recr	-0.52	-2.40	13.85	36.52	0.792	0.784	0.0035
hote	0.62	5.66	4.74	17.81	0.930	0.927	0.0017

Notes:

1) Table 2 reports results from estimating equation (1) in section 4 of the main text. The dependent variable is the standard deviation of two-month differences in relative prices (volatility measure 1). All regressions contain as explanatory variables a dummy for each of the included regions in addition to the variables listed in the table. The regression for 'allit' additionally includes an extra Italian border dummy. Coefficients on log distance and border are multiplied by 10^3.

2) In brackets, t-statistics are reported. In computing these statistics, White's heteroscedastic-consistent standard errors were used.

3) R^2 denotes the (unadjusted) coefficient of determination, R^2_{adj} denotes the adjusted coefficient of determination and the term *s.e.r.* denotes the standard error of regression.

4) There are 1540 observations in each regression (the regression for 'allit' contains 3240 observations).

Table 3: All Items and Subcategories, Aggregate Border Estimates, Overall Period (1995.01 - 2002.12), Volatility Measure 2

Category	(ln)dist		border		R^2	R^2_{adj}	$s.e.r.$
	Coeff.	t-stat	Coeff.	t-stat			
			Total Index				
allit	0.45	8.35	4.39	21.53	0.867	0.863	0.0014
			COICOP Subcategories				
food	1.74	9.96	5.26	17.99	0.783	0.775	0.0028
alco	0.41	3.09	8.33	37.74	0.853	0.847	0.0019
clot	6.24	6.25	6.43	4.42	0.866	0.861	0.0124
hous	0.43	4.87	5.41	39.07	0.800	0.792	0.0015
furn	0.06	0.92	3.57	34.94	0.869	0.864	0.0011
heal	0.52	4.86	4.75	32.42	0.812	0.805	0.0015
tran	-0.90	-8.44	10.58	66.69	0.881	0.876	0.0018
comm	1.13	4.24	10.80	57.63	0.950	0.948	0.0018
recr	-0.80	-2.86	18.31	35.61	0.799	0.792	0.0045
hote	0.10	0.50	6.29	12.22	0.835	0.828	0.0035

Notes:

1) Table 3 reports results from estimating equation (1) in section 4 of the main text. The dependent variable is the spread between the 10th and the 90th percentile of the distribution of two-month changes in relative prices. (volatility measure 2). All regressions contain as explanatory variables a dummy for each of the included regions in addition to the variables listed in the table. The regression for 'allit' additionally includes an extra Italian border dummy. Coefficients on log distance and border are multiplied by 10^3.

2) In brackets, t-statistics are reported. In computing these statistics, White's heteroscedastic-consistent standard errors were used.

3) R^2 denotes the (unadjusted) coefficient of determination, R^2_{adj} denotes the adjusted coefficient of determination and the term $s.e.r.$ denotes the standard error of regression.

4) There are 1540 observations in each regression (the regression for 'allit' contains 3240 observations).

Table 4: All Items and Subcategories, Aggregate Border Estimates, Overall Period (1995.01 - 2002.12), Volatility Measure 3

Category	(ln)dist		border		R^2	R^2_{adj}	*s.e.r.*
	Coeff.	t-stat	Coeff.	t-stat			
			Total Index				
allit	0.14	1.84	5.74	26.62	0.871	0.868	0.0024
			COICOP Subcategories				
food	1.82	9.92	6.52	19.89	0.772	0.763	0.0030
alco	0.16	1.05	12.48	38.66	0.918	0.915	0.0023
clot	8.69	14.27	11.55	11.94	0.830	0.823	0.0093
hous	0.56	5.05	4.22	33.09	0.826	0.819	0.0013
furn	0.75	7.71	2.50	23.13	0.835	0.829	0.0012
heal	0.29	1.99	4.96	15.37	0.883	0.879	0.0023
tran	-0.10	-1.09	5.73	43.90	0.899	0.895	0.0013
comm	0.81	6.71	15.20	81.29	0.977	0.976	0.0015
recr	-0.23	-0.80	15.67	31.37	0.754	0.745	0.0045
hote	-0.25	-1.63	5.57	14.47	0.861	0.856	0.0027

Notes:

1) Table 4 reports results from estimating equation (1) in section 4 of the main text. The dependent variable is the standard deviation of the two-month ahead in-sample forecast errors from an AR(6) process (including 12 seasonal dummies) estimated for each relative price series (volatility measure 3). All regressions contain as explanatory variables a dummy for each of the included regions in addition to the variables listed in the table. The regression for 'allit' additionally includes an extra Italian border dummy. Coefficients on log distance and border are multiplied by 10^3.

2) In brackets, t-statistics are reported. In computing these statistics, White's heteroscedastic-consistent standard errors were used.

3) R^2 denotes the (unadjusted) coefficient of determination, R^2_{adj} denotes the adjusted coefficient of determination and the term *s.e.r.* denotes the standard error of regression.

4) There are 1540 observations in each regression (the regression for 'allit' contains 3240 observations).

Table 5: All Items, The Role of Nominal Exchange Rate Volatility, Overall Period (1995.01 - 2002.12), Volatility Measure 1

Spec.	ln(dist)	border	n.e.r.vol.	Italian n.e.r.vol	R^2_{adj}
Spec. 1	-0.43 (-16.12)	0.67 (7.70)	0.85 (287.30)		0.993
Spec. 2	0.03 (1.26)	1.34 (21.45)	0.69 (114.57)	0.92 (257.30)	0.996

Notes:

1) Table 5 reports results from estimating equation (1) in section 4 of the main text. The dependent variable is the standard deviation of two-month differences in relative prices (volatility measure 1). All regressions contain as explanatory variables a dummy for each of the included regions and an extra Italian border dummy in addition to the variables listed in the table. The term '*n.e.r.vol.*' denotes the standard deviation of two-month differences in the nominal exchange rate between two regions. Coefficients on log distance and border are multiplied by 10^3.

2) In brackets, t-statistics are reported. In computing these statistics, White's heteroscedastic-consistent standard errors were used.

3) R^2_{adj} denotes the adjusted coefficient of determination.

4) There are 3240 observations in each regression.

Table 6: All Items, Regression Results for Individual Border Estimates, Overall Period (1995.01 - 2002.12), Volatility Measures 1, 2 and 3

	Measure 1				Measure 2		Measure 3	
Variable	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat
ln(dist)	0.14	4.51			0.42	3.75	0.05	1.10
dist			0.29	4.57				
dist^2			-0.12	-5.61				
ge-au	2.19	33.21	2.24	34.54	5.32	24.63	1.94	23.60
ge-fi	6.78	83.25	6.95	81.69	14.15	41.25	8.21	57.46
ge-it	21.85	399.11	21.91	403.61	29.48	158.48	21.51	175.47
ge-sp	8.61	121.43	8.75	119.48	16.52	68.24	6.06	48.82
ge-po	9.82	60.29	10.12	61.33	13.40	24.55	7.75	21.17
ge-ch	12.07	67.57	12.13	67.42	31.03	54.05	14.51	106.29
au-fi	6.04	73.70	6.27	73.60	11.84	38.15	6.82	51.70
au-it	21.19	511.95	21.26	571.53	28.76	178.66	17.02	230.28
au-sp	8.14	119.62	8.32	121.56	13.82	52.86	5.74	53.84
au-po	8.86	56.05	9.26	57.67	9.88	17.99	4.57	18.58
au-ch	10.54	55.76	10.64	56.50	27.62	43.57	12.42	79.51
fi-it	21.20	300.07	21.55	220.04	27.89	73.48	13.52	102.50
fi-sp	9.95	120.06	10.61	72.89	16.08	46.26	9.08	64.93
fi-po	10.47	65.45	11.49	48.25	14.46	23.91	7.64	31.59
fi-ch	13.23	69.24	13.53	69.19	34.22	52.61	13.52	69.78
it-sp	14.60	292.31	14.71	286.80	18.64	101.06	14.57	130.46
it-po	11.22	75.98	11.51	78.08	16.31	32.35	9.29	36.64
it-ch	27.78	160.21	27.86	160.34	51.93	83.41	19.07	131.72
sp-po	5.44	38.93	5.49	42.93	10.46	21.98	3.18	15.88
sp-ch	14.77	81.20	14.90	82.37	37.01	63.11	11.42	76.86
po-ch	16.67	72.51	16.94	74.65	34.20	35.59	11.81	39.17
R^2	0.997		0.997		0.974		0.978	
R^2_{adj}	0.997		0.997		0.973		0.977	
$s.e.e.$	0.0000		0.0004		0.0010		0.0000	

Notes:

1) Table 6 reports results from estimating equation (1) in section 4 of the main text. In the specifications named 'Measure 1', 'Measure 2' and 'Measure 3', the dependent variable is volatility measure 1, 2 and 3, respectively. A description of how the individual volatility measures are constructed is given in the footnotes of tables 2, 3 and 4, respectively. All regressions contain as explanatory variables a dummy for each of the included regions in addition to the variables listed in the table. Coefficients on log distance and borders are multiplied by 10^3, coefficients on distance and distance squared are multiplied by 10^6 and 10^9, respectively.

2) In brackets, t-statistics are reported. In computing these statistics, White's heteroscedastic-consistent standard errors were used.

3) R^2 denotes the (unadjusted) coefficient of determination, R^2_{adj} denotes the adjusted coefficient of determination and the term $s.e.r.$ denotes the standard error of regression.

4) There are 3240 observations in each regression.

Table 7: Subcategories, Regression Results for Individual Border Estimates, Overall Period (1995.01 - 2002.12), Volatility Measure 1

Var.	Categories									
	food	alco	clot	hous	furn	heal	tran	comm	recr	hote	bord/dist
ln(dist)	0.38	0.28	0.35	0.22	0.17	0.04	0.14	0.33	0.49	0.23	
	(3.12)	(1.2)	(0.74)	(2.6)	(3.15)	(0.58)	(1.68)	(3.78)	(2.32)	(2.75)	
ge-fi	6.07	7.66	38.32	6.56	5.29	12.28	6.82		11.39	11.82	45.01
	(16.49)	(17.07)	(38.45)	(28.34)	(34.69)	(43.91)	(30.79)		(18.24)	(24.32)	
ge-it	7.50	11.42	5.63	7.09	6.54	11.38	7.04		12.20	10.61	33.65
	(28.45)	(35.94)	(12.31)	(43.13)	(78.63)	(47.31)	(55.64)		(24.16)	(26.08)	
ge-sp	9.73	12.68	19.37	6.19	7.75	12.22	8.94		13.83	10.58	42.93
	(32.07)	(25.76)	(25.12)	(31.59)	(70.39)	(50.27)	(51.94)		(23.33)	(24.34)	
ge-po	7.61	8.29	31.92	8.93	8.57	13.10	10.06		12.57	11.23	47.58
	(17.42)	(15.17)	(11.08)	(35.81)	(44.8)	(45.49)	(35.23)		(19.71)	(22.33)	
fi-it	7.28	12.84	41.40	9.90	6.42	8.78	8.36	16.93	6.11	6.06	47.32
	(25.11)	(29)	(38.84)	(49.11)	(42.43)	(47.58)	(37.06)	(45.11)	(14.53)	(24.4)	
fi-sp	10.04	14.30	36.39	9.40	8.65	9.49	9.43	15.21	16.35	7.85	52.30
	(31.59)	(25.04)	(29.43)	(41.35)	(51.24)	(47.47)	(32.18)	(38.95)	(32.32)	(29.5)	
fi-po	8.71	9.02	32.78	11.02	8.40	10.60	10.79	18.42	7.05	7.47	47.41
	(20.09)	(15.53)	(8.71)	(42.67)	(29.33)	(39.53)	(30.88)	(39.07)	(12.86)	(24.09)	
it-sp	5.78	13.63	20.58	5.18	4.36	4.06	6.66	14.83	14.74	4.19	35.86
	(36.51)	(39.1)	(31.17)	(37.22)	(50.24)	(37.23)	(55.16)	(117.47)	(45.87)	(33.05)	
it-po	5.12	7.63	34.17	5.93	4.46	4.44	5.25	12.25	1.99	2.70	32.02
	(16.43)	(17.09)	(11.96)	(28.55)	(30.06)	(25.71)	(21.73)	(43.19)	(5.28)	(13.86)	
sp-po	4.77	8.19	27.68	5.22	3.95	3.32	8.45	14.24	12.81	3.24	35.05
	(18)	(22.17)	(9.84)	(28.41)	(32.22)	(25.21)	(38.35)	(57.89)	(42.37)	(19.94)	
R^2	0.943	0.933	0.961	0.943	0.970	0.983	0.967	0.992	0.935	0.979	
R^2_{adj}	0.941	0.930	0.959	0.941	0.969	0.983	0.965	0.992	0.932	0.978	
$s.e.r.$	0.0011	0.0019	0.0045	0.0009	0.0006	0.0008	0.0008	0.0008	0.0020	0.0010	

Notes:

1) Table 7 reports results from estimating equation (1) in section 4 of the main text. The column denoted 'bord/dist' reports the ratio of that row's average border estimate divided by the average distance estimate. Averages are computed over the COICOP categories. For more detailed notes, see the footnotes of table 6.

Table 8: All Items, Regression Results for Individual Border Estimates, pre-EMU and EMU Subperiod (1995.01 - 1998.12, 1999.01 - 2002.12), Volatility Measures 1

	Specification 1				Specification 2			
	Subperiod 1		Subperiod 2		Subperiod 1		Subperiod 2	
Var.	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat
ln(dist)	0.16	4.72	0.06	1.53				
dist					0.37	4.67	0.12	1.46
$dist^2$					-0.15	-5.67	-0.03	-0.97
ge-au	3.08	46.93	1.38	12.73	3.14	50.12	1.40	13.13
ge-fi	9.41	81.17	3.59	28.16	9.59	77.65	3.60	26.61
ge-it	32.13	528.53	1.83	21.57	32.20	540.39	1.84	21.67
ge-sp	12.42	154.19	3.49	33.57	12.56	145.84	3.50	32.32
ge-po	14.52	90.36	3.27	15.00	14.87	88.64	3.29	13.59
ge-ch	15.06	118.60	8.95	32.83	15.12	115.39	8.97	33.01
au-fi	8.77	82.37	2.64	20.71	9.01	75.98	2.66	19.74
au-it	31.61	653.38	0.78	15.38	31.68	715.99	0.80	17.75
au-sp	12.25	158.74	2.40	27.68	12.45	149.13	2.42	26.73
au-po	13.82	87.85	1.56	7.71	14.27	87.57	1.60	6.80
au-ch	13.97	101.79	6.67	22.98	14.08	101.03	6.71	23.33
fi-it	30.89	319.67	2.67	25.26	31.28	234.52	2.68	18.92
fi-sp	14.37	130.14	3.41	30.49	15.13	79.47	3.46	17.75
fi-po	15.33	88.86	3.66	17.64	16.52	60.63	3.76	10.98
fi-ch	18.07	113.20	6.96	24.05	18.40	104.47	6.99	23.72
it-sp	21.36	348.51	2.66	44.39	21.47	318.07	2.67	41.73
it-po	16.86	117.09	1.63	8.68	17.19	116.57	1.65	7.74
it-ch	39.76	325.96	7.12	27.14	39.84	317.16	7.14	27.39
sp-po	7.66	58.03	2.60	14.09	7.72	65.06	2.61	14.08
sp-ch	20.58	147.69	6.89	25.36	20.71	145.32	6.92	25.71
po-ch	24.08	110.46	5.93	15.47	24.37	111.93	5.96	15.31
R^2	0.998		0.934		0.998		0.934	
R^2_{adj}	0.998		0.932		0.998		0.932	
$s.e.r.$	0.0005		0.0005		0.0005		0.0005	

1) Table 8 reports results from estimating equation (1) in section 4 of the main text. The dependent variable is volatility measure 1. For further notes, see the footnotes of table 6.

Table 9: Food and Non-Alcoholic Beverages (food) and Alcoholic Beverages and Tobacco (alco), Regression Results for Individual Border Estimates, pre-EMU and EMU Subperiod (1995.01 - 1998.12, 1999.01 - 2002.12), Volatility Measure 1

	food				alco			
	Subperiod 1		Subperiod 2		Subperiod 1		Subperiod 2	
Var.	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat
ln(dist)	0.23	1.76	0.53	4.04	0.11	0.38	0.66	4.49
ge-fi	10.02	26.77	1.85	3.81	11.22	23.13	4.01	11.07
ge-it	12.69	61.73	3.16	8.41	14.81	55.91	8.22	27.33
ge-sp	13.08	46.00	6.67	16.89	16.82	32.73	7.57	17.35
ge-po	11.89	25.58	3.31	6.15	12.00	18.64	4.11	9.06
fi-it	11.79	33.19	3.19	9.61	17.21	30.47	7.41	26.24
fi-sp	12.94	32.27	7.00	20.32	19.24	26.82	6.66	17.58
fi-po	12.58	26.79	4.34	8.60	12.32	16.35	4.30	10.29
it-sp	8.84	48.21	2.15	12.76	18.01	40.68	8.60	37.50
it-po	8.84	26.58	2.00	5.26	11.34	20.30	3.59	11.08
sp-po	6.53	23.58	2.85	8.66	11.09	23.20	3.95	13.89
R^2	0.948		0.903		0.931		0.938	
R^2_{adj}	0.946		0.899		0.928		0.935	
$s.e.r.$	0.0014		0.0012		0.0024		0.0014	

Notes:

1) Table 9 reports results from estimating equation (1) in section 4 of the main text. The dependent variable is volatility measure 1. For further notes, see the footnotes of table 6.

Table 10: Clothing and Footwear (clot) and Housing, Water, Electricity, Gas and Other Fuels (hous), Regression Results for Individual Border Estimates, pre-EMU and EMU Subperiod (1995.01 - 1998.12, 1999.01 - 2002.12), Volatility Measure 1

	clot				furn			
	Subperiod 1		Subperiod 2		Subperiod 1		Subperiod 2	
Var.	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat
ln(dist)	0.42	1.41	0.42	0.66	0.27	1.91	0.21	2.43
ge-fi	31.15	40.71	43.33	33.21	9.77	27.96	2.72	9.72
ge-it	10.45	36.73	0.10	0.16	11.62	52.21	2.12	13.54
ge-sp	11.26	24.68	25.61	24.72	9.75	35.83	1.46	7.59
ge-po	25.13	14.86	35.70	9.23	14.22	40.34	1.44	5.84
fi-it	33.63	42.79	45.31	33.08	15.72	49.18	3.58	13.91
fi-sp	34.51	40.33	38.11	23.70	14.14	40.94	2.80	9.97
fi-po	25.83	10.15	39.78	8.50	16.28	40.24	4.38	14.89
it-sp	7.50	19.20	26.26	29.98	7.82	35.64	2.51	18.36
it-po	26.33	15.44	35.88	9.54	9.86	30.87	2.20	11.81
sp-po	21.22	12.76	31.68	8.43	7.89	31.82	2.54	16.06
R^2	0.978		0.951		0.944		0.845	
R^2_{adj}	0.977		0.948		0.941		0.838	
$s.e.r.$	0.0026		0.0060		0.0013		0.0010	

Notes:

1) Table 10 reports results from estimating equation (1) in section 4 of the main text. The dependent variable is volatility measure 1. For further notes, see the footnotes of table 6.

Table 11: Furnishings, Household Equip. and Routine Maint. of the House (furn) and Health (heal), Regression Results for Individual Border Estimates, pre-EMU and EMU Subperiod (1995.01 - 1998.12, 1999.01 - 2002.12), Volatility Measure 1

	hous				heal			
	Subperiod 1		Subperiod 2		Subperiod 1		Subperiod 2	
Var.	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat
ln(dist)	0.32	3.70	-0.03	-0.56	0.13	1.39	-0.06	-0.93
ge-fi	8.55	36.91	0.91	5.52	17.86	55.64	4.13	13.97
ge-it	11.18	106.93	0.61	7.61	20.53	77.79	1.49	8.76
ge-sp	12.06	79.08	1.27	12.30	18.92	69.38	0.74	4.24
ge-po	13.37	45.83	1.12	6.85	20.39	58.02	1.37	6.15
fi-it	10.88	44.20	0.84	5.22	14.21	66.96	3.99	19.19
fi-sp	13.41	49.78	1.10	6.53	13.91	55.39	3.95	18.20
fi-po	12.91	29.35	1.64	7.98	15.57	41.84	4.71	20.05
it-sp	6.80	48.00	1.03	13.18	7.46	50.93	1.34	13.33
it-po	7.29	32.60	1.21	9.04	8.05	32.82	1.90	11.65
sp-po	5.61	32.40	1.80	14.91	5.33	26.41	0.91	7.24
R^2	0.968		0.911		0.983		0.975	
R^2_{adj}	0.967		0.907		0.983		0.974	
$s.e.r.$	0.0009		0.0005		0.0010		0.0008	

Notes:

1) Table 11 reports results from estimating equation (1) in section 4 of the main text. The dependent variable is volatility measure 1. For further notes, see the footnotes of table 6.

Table 12: Transportation (tran) and Communication (comm), Regression Results for Individual Border Estimates, pre-EMU and EMU Subperiod (1995.01 - 1998.12, 1999.01 - 2002.12), Volatility Measure 1

	tran				comm			
	Subperiod 1		Subperiod 2		Subperiod 1		Subperiod 2	
Var.	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat
ln(dist)	0.16	1.58	0.11	1.23	0.44	3.42	0.16	2.07
ge-fi	8.95	32.55	4.68	13.98	0.0		0.0	
ge-it	9.42	61.78	4.57	25.55	0.0		0.0	
ge-sp	11.73	62.27	6.27	28.74	0.0		0.0	
ge-po	13.20	37.55	6.97	18.77	0.0		0.0	
fi-it	11.19	38.30	5.79	19.03	20.31	33.55	14.05	23.80
fi-sp	13.31	44.82	5.04	13.05	17.61	28.61	12.91	21.49
fi-po	12.54	28.28	9.41	23.03	23.82	30.48	13.53	19.21
it-sp	7.11	48.31	5.61	41.97	19.23	103.22	11.20	97.33
it-po	6.91	25.64	3.80	13.42	15.89	36.43	8.55	26.42
sp-po	7.86	36.46	9.42	35.39	17.15	49.05	11.47	38.80
R^2	0.969		0.948		0.986		0.989	
R^2_{adj}	0.967		0.946		0.986		0.988	
$s.e.r.$	0.0010		0.0010		0.0013		0.0008	

Notes:

1) Table 12 reports results from estimating equation (1) in section 4 of the main text. The dependent variable is volatility measure 1. For further notes, see the footnotes of table 6.

Table 13: Recreation and Culture (recr) and Hotels, Cafes and Restaurants (hote), Regression Results for Individual Border Estimates, pre-EMU and EMU Subperiod (1995.01 - 1998.12, 1999.01 - 2002.12), Volatility Measure 1

	recr				hote			
	Subperiod 1		Subperiod 2		Subperiod 1		Subperiod 2	
Var.	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat
ln(dist)	0.14	0.66	1.04	4.05	0.34	2.67	0.22	2.29
ge-fi	13.19	20.0	10.73	13.49	13.84	44.57	11.45	16.69
ge-it	16.62	31.12	10.21	15.69	12.88	55.50	9.84	17.12
ge-sp	18.44	31.11	10.42	13.49	13.94	52.24	8.69	14.25
ge-po	17.39	24.07	9.27	11.47	14.38	30.04	9.73	15.26
fi-it	10.08	21.76	1.54	3.03	11.46	37.88	0.81	2.51
fi-sp	16.02	30.54	16.76	26.67	12.57	39.75	2.45	6.97
fi-po	11.79	18.51	-0.42	-0.63	12.44	32.73	1.29	3.17
it-sp	12.87	41.85	17.21	43.83	6.53	32.50	1.47	10.06
it-po	4.31	8.89	-0.25	-0.55	5.15	17.12	0.07	0.34
sp-po	10.27	25.69	16.29	41.64	3.80	16.20	2.65	14.72
R^2	0.935		0.941		0.966		0.972	
R^2_{adj}	0.932		0.938		0.965		0.971	
$s.e.r.$	0.0020		0.0024		0.0013		0.0012	

Notes:

1) Table 13 reports results from estimating equation (1) in section 4 of the main text. The dependent variable is volatility measure 1. For further notes, see the footnotes of table 6.

Table 14: The Relative Importance of the Border: Estimated Average Distance and Border Coefficients for Subcategories, pre-EMU and EMU Subperiod (1995.01 - 1998.12, 1999.01 - 2002.12), Volatility Measure 1

	Subperiod 1		Subperiod 2	
Border	Avg. Coeff.	bord/dist	Avg. Coeff.	bord/dist
ln(dist)	0.26		0.33	
ge-fi	13.84	54.10	9.31	28.56
ge-it	13.35	52.20	4.48	13.74
ge-sp	14.00	54.73	7.63	23.41
ge-po	15.77	61.66	8.11	24.88
fi-it	15.65	61.17	8.65	26.53
fi-sp	16.77	65.54	9.68	29.68
fi-po	15.61	61.01	8.30	25.44
it-sp	10.22	39.93	7.74	23.73
it-po	10.40	40.64	5.89	18.08
sp-po	9.67	37.82	8.36	25.63

Notes:

1) Table 14 reports average estimated distance and border coefficients for subcategories from tables 9 to tables 13.

2) For further notes, see the footnotes of tables 9 to 13.

Table 15: German Economic and Monetary Union (GEMU), Regression Results for Individual Border Estimates, Overall Period (1991.01 - 2002.12) and pre-EMU- and EMU Subperiods (1991.01 - 1994.12, 1995.01 - 1998.12, 1999.01 - 2002.12), Volatility Measure 1

	Overall Sample		Subperiod 1		Subperiod 2		Subperiod 3	
Var.	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat
ln(dist)	0.17	4.49	0.31	7.08	0.16	3.81	0.08	1.35
we-ea	14.20	76.88	25.31	76.04	0.79	9.07	0.27	1.52
we-au	2.64	60.57	3.60	49.68	3.06	60.15	1.30	12.64
we-ch	12.24	38.08	13.24	25.88	15.01	136.00	8.90	31.90
ea-au	14.32	76.21	26.25	76.86	2.83	30.02	0.82	5.89
ea-ch	19.74	52.52	29.70	48.05	14.85	105.95	8.37	27.66
au-ch	11.44	35.04	13.92	27.45	13.98	108.76	6.64	20.90
R^2	0.996		0.997		0.992		0.935	
R^2_{adj}	0.995		0.997		0.992		0.931	
$s.e.r.$	0.0005		0.0007		0.0005		0.0008	

Notes:

1) Table 15 reports results from estimating equation (1) in section 4 of the main text. The dependent variable is volatility measure 1. There are 666 observations in each regression. For further notes, see the footnotes of table 6.

Table 16: German Economic and Monetary Union (GEMU), Regression Results for Individual Border Estimates, Overall Period (1991.01 - 2002.12) and pre-EMU and EMU Subperiods (1991.01 - 1994.12, 1995.01 - 1998.12, 1999.01 - 2002.12), Volatility Measure 2

	Overall Sample		Subperiod 1		Subperiod 2		Subperiod 1	
Var.	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat
ln(dist)	0.34	4.67	0.55	4.15	0.38	4.02	0.06	0.45
we-ea	1.13	7.77	5.11	12.19	0.48	3.02	0.28	1.09
we-au	4.38	43.82	6.63	40.21	3.75	28.27	2.78	14.72
we-ch	21.41	71.12	27.40	49.37	24.57	63.99	13.86	28.45
ea-au	4.50	31.62	9.04	29.17	3.86	20.99	2.34	9.40
ea-ch	21.43	63.52	28.75	40.66	23.58	58.51	13.79	24.02
au-ch	18.98	58.07	26.44	47.34	25.47	73.22	10.14	23.84
R^2	0.984		0.983		0.980		0.888	
R^2_{adj}	0.983		0.982		0.978		0.880	
$s.e.r.$	0.0010		0.0014		0.0014		0.0016	

Notes:

1) Table 16 reports results from estimating equation (1) in section 4 of the main text. The dependent variable is volatility measure 2. There are 666 observations in each regression. For further notes, see the footnotes of table 6.

Table 17: German Economic and Monetary Union (GEMU), Regression Results for Individual Border Estimates, Overall Period (1991.01 - 2002.12) and pre-EMU- and EMU Subperiods (1991.01 - 1994.12, 1995.01 - 1998.12, 1999.01 - 2002.12), Volatility Measure 3

	Overall Sample		Subperiod 1		Subperiod 2		Subperiod 1	
Var.	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat
ln(dist)	0.55	8.92	1.36	6.33	0.67	7.03	-0.10	-0.65
we-ea	19.06	60.92	40.89	39.59	0.23	0.96	0.20	0.29
we-au	1.73	22.19	-0.03	-0.09	1.46	10.17	3.65	9.51
we-ch	25.05	146.34	28.95	27.29	23.40	70.87	23.92	30.49
ea-au	19.63	59.92	43.45	39.05	2.42	9.95	4.59	8.68
ea-ch	36.49	93.85	70.35	45.66	24.34	70.74	25.25	32.14
au-ch	21.91	119.00	27.29	25.31	22.08	71.94	16.72	44.77
R^2	0.995		0.987		0.985		0.915	
R^2_{adj}	0.994		0.986		0.984		0.909	
$s.e.r.$	0.0008		0.0027		0.0011		0.0023	

Notes:

1) Table 17 reports results from estimating equation (1) in section 4 of the main text. The dependent variable is volatility measure 3. There are 666 observations in each regression. For further notes, see the footnotes of table 6.

Table 18: German Economic and Monetary Union (GEMU), Descriptive Statistics for Single-Equation ADF Tests, Overall Period (1991.01 - 2002.12)

Location	ρ	ρ_{adj}	Std.Dvt.	Half-Life	Half-Life (adj.)
all	0.85	0.88	0.14	1.1	1.3
intra-nat.	0.90	0.93	0.11	1.7	2.3
internat.	0.82	0.85	0.14	0.9	1.1
we-we	0.97	1.00	0.06	6.0	-
we-ea	0.63	0.65	0.06	0.4	0.4
we-au	0.94	0.96	0.10	2.6	4.6
we-ch	0.95	0.98	0.02	3.3	7.0
ea-ea	0.85	0.87	0.11	1.0	1.3
ea-au	0.69	0.71	0.07	0.5	0.5
ea-ch	0.80	0.83	0.07	0.8	0.9
au-au	0.90	0.92	0.12	1.6	2.2
au-ch	0.90	0.93	0.04	1.6	2.2
ch-ch	0.91	0.94	0.05	1.9	2.8

Notes:

1) Table 18 reports means and standard deviations of estimated AR(1) coefficients for individual relative price series. The Augmented Dickey Fuller (ADF) unit root results are obtained by regressing the change of the (log) real exchange rate on its past level and six lagged difference terms. The overall number of considered relative price series is 666.

2) The term 'intra-nat.' refers to estimated AR(1) coefficients for relative prices between regions that are located in the same country. The term 'internat.' refers to AR(1) coefficients for relative prices between regions that are located in different countries. The terms 'we', 'ea', 'au' and 'ch' refer to West-German, East-German, Austrian and Swiss locations.

3) Bias adjustment is done using the formula by Kendall (1954).

4) Half-lives are computed using the formula: $half-life = \frac{ln(0.5)}{ln(\hat{\rho})}$, where $\hat{\rho}$ denotes the average estimated AR(1) coefficient.

Table 19: German Economic and Monetary Union (GEMU), Levin-Lin Panel Unit Root Test of Real Exchange Rate Convergence, Overall Period (1991.01 - 2002.12)

Loc.	ρ	ρ_{adj}	$t^{\star}$	p-value	half-l.	half-l.(adj.)	N	T
all	0.90	0.94	-20.55	0.036	6.6	10.7	138	63
we-we	1.02	1.22	1.38	0.987	-	-	6	63
we-ea	0.68	0.71	-20.09	0.009	1.8	2.0	12	63
we-au	0.95	1.00	-2.94	0.701	14.7	-	26	63
we-ch	0.91	0.95	-4.45	0.546	7.5	13.5	10	63
ea-ea	0.90	0.93	-3.48	0.355	6.3	9.8	5	63
ea-au	0.74	0.77	-23.25	0.002	2.3	2.6	25	63
ea-ch	0.77	0.80	-14.12	0.002	2.6	3.0	9	63
au-au	0.96	1.00	-3.51	0.716	15.4	-	19	63
au-ch	0.85	0.89	-9.85	0.113	4.3	5.7	23	63
ch-ch	0.91	0.95	-3.12	0.359	7.4	13.2	3	63

Notes:

1) Table 19 reports results from Levin-Lin panel unit root tests of real exchange rate convergence. The real exchange rate between two regions is computed as the ratio of the respective regions' CPI (denoted in the same currency). When constructing relative prices, for each country one city was chosen as base city: For West Germany Berlin-West was chosen, for East Germany Berlin-East was chosen, for Austria Vienna was chosen and for Switzerland Bern was chosen. A more detailed description of our procedure is given in section D of the appendix.

2) Relative prices are grouped into various classes. 'all' refers to the group of all relative prices, 'intra-nat.' involves only intra-national relative prices, 'internat.' denotes all international relative prices. The other terms ('we-we', ...) refer to bilateral relative prices as indicated by the respective country short names. 'we' refers to West-German locations, 'ea' refers to East-German locations, 'au' refers to Austrian locations and 'ch' refers to Swiss locations.

3) Bias adjustment is done using the formula by Nickell (1981).

Table 20: German Economic and Monetary Union (GEMU), Im-Pesaran-Shin Panel Unit Root Test of Real Exchange Rate Convergence, Overall Period (1991.01 - 2002.12)

Loc.	ρ	ρ_{adj}	$t^\star$	p-value	half-l.	half-l.(adj.)	N	T
all	0.83	0.84	-2.19	0.012	3.6	3.8	138	63
we-we	1.01	1.03	0.39	0.989	-	-	6	63
we-ea	0.66	0.66	-5.67	0.009	1.7	1.7	12	63
we-au	0.88	0.89	-0.82	0.722	5.6	6.2	26	63
we-ch	0.91	0.92	-1.41	0.563	6.9	8.0	10	63
ea-ea	0.85	0.86	-1.68	0.334	4.4	4.7	5	63
ea-au	0.72	0.72	-4.52	0.004	2.1	2.1	25	63
ea-ch	0.77	0.77	-4.59	0.002	2.6	2.7	9	63
au-au	0.87	0.88	-1.29	0.432	5.1	5.5	19	63
au-ch	0.82	0.83	-2.15	0.119	3.5	3.6	23	63
ch-ch	0.85	0.86	-2.02	0.268	4.4	4.7	3	63

Notes:

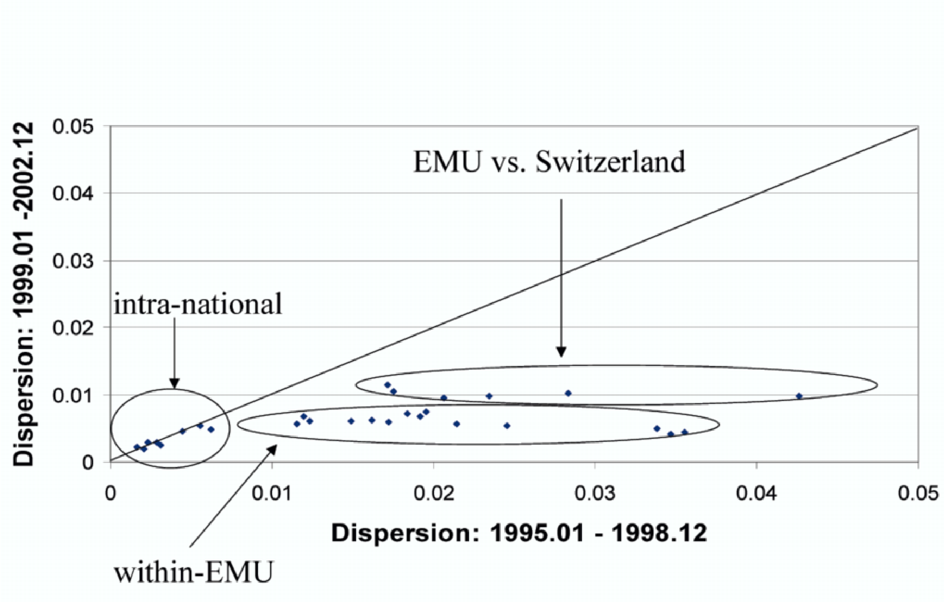
1) Table 20 reports results from Im-Pesaran-Shin panel unit root tests of real exchange rate convergence. The real exchange rate between two regions is computed as the ratio of the respective regions' CPI (denoted in the same currency). When constructing relative prices, for each country one city was chosen as base city: For West Germany Berlin-West was chosen, for East Germany Berlin-East was chosen, for Austria Vienna was chosen and for the Switzerland Bern was chosen. A more detailed description of our procedure is given in section E of the appendix.

2) Relative prices are grouped into various classes. 'all' refers to the group of all relative prices, 'intra-nat.' involves only intra-national relative prices, 'internat.' denotes all international relative prices. The other terms ('we-we', ...) refer to bilateral relative prices as indicated by the respective country short names. 'we' refers to West-German locations, 'ea' refers to East-German locations, 'au' refers to Austrian locations and 'ch' refers to Swiss locations.

3) Bias adjustment is done using the formula by Kendall (1954).

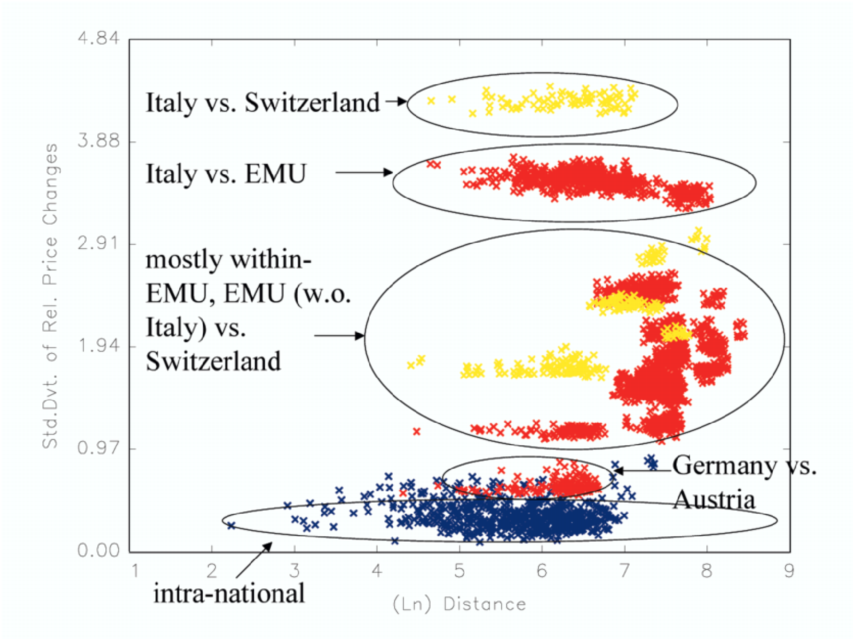
10 Figures

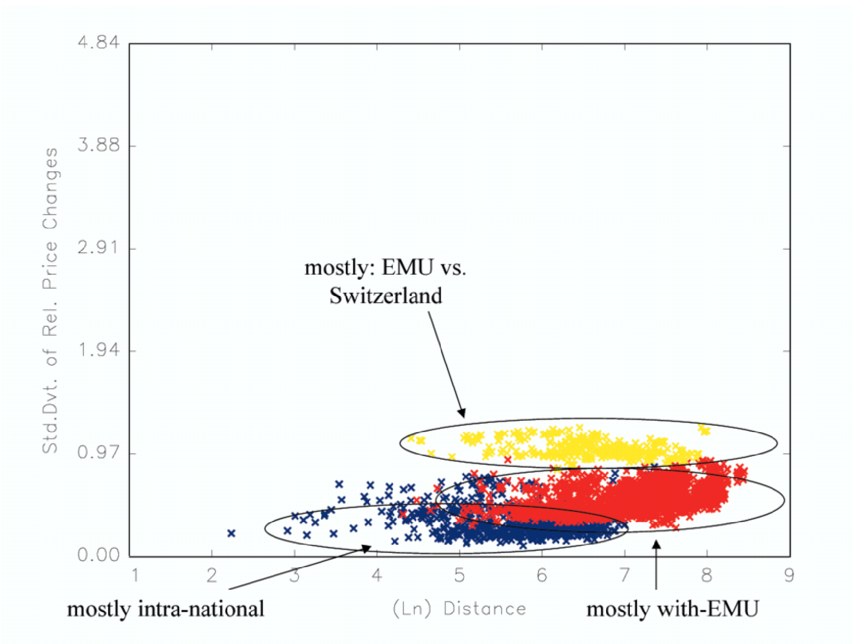
Figure 1: Relative Price Volatility in the pre-EMU- (1995.01-1998.12) and EMU- (1999.01-2002.12) Subperiod, Volatility Measure 1



Note: *Figure 1 plots mean values of the relative price dispersion across bilateral country groups (e.g., the mean of relative price volatilities of all German-Austrian locations) for the EMU-subperiod (1999:01-2002:12) on the vertical axis, and the pre-EMU-subperiod (1995.01-1998.12) on the horizontal axis. The solid line is the* 45° *line.*

Figure 2: Relative Price Volatility versus Distance, Pre-EMU- (1995.01 - 1998.12, Upper Panel) and EMU- (1999.01 - 2002.12, Lower Panel) Subperiod



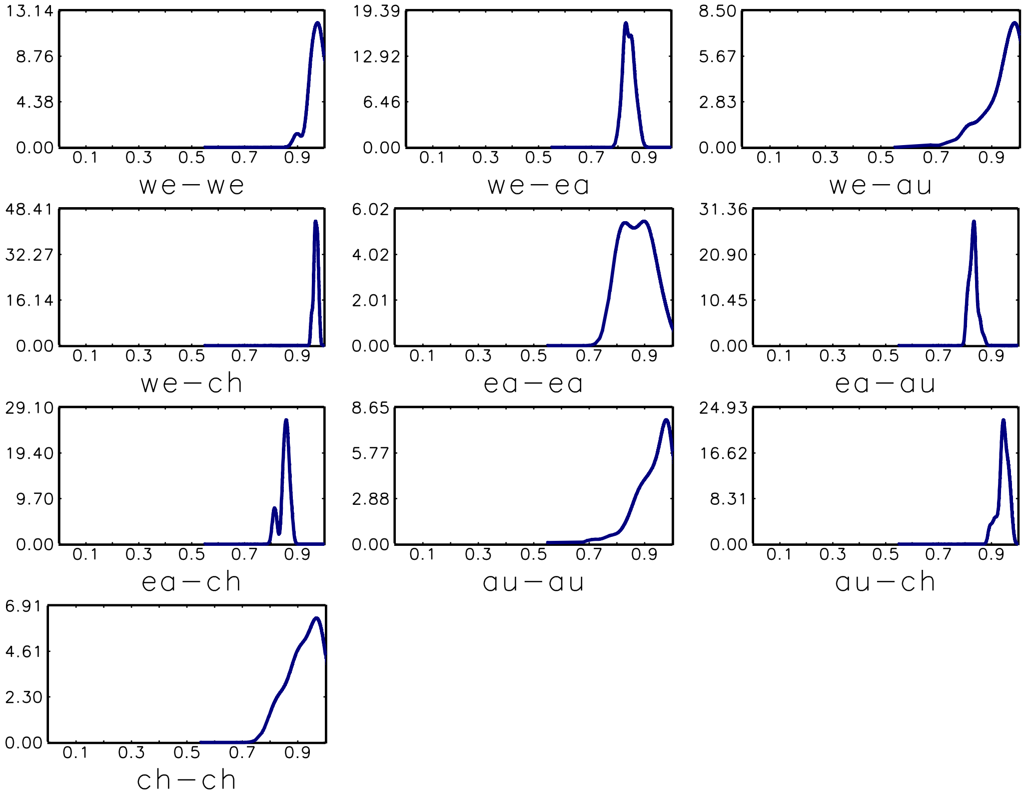


Note: *Figure 2 plots our measure for relative price dispersion across regions against the distance (in logs) between these regions for the pre-EMU-subperiod (upper panel) and the EMU-subperiod (lower panel). Relative price dispersion between region i and region j is computed as the standard deviation of two-month relative price changes between the two regions, i.e.,*

$$V(q_{ij}) = \sqrt{var(\Delta q_{ij,t})}$$

where $\Delta q_{ij,t}$ *denotes the two-month changes between region's* i *and region's* j *relative price and* $var(.)$ *denotes the empirical variance of* $\Delta q_{ij,t}$.

Figure 3: Kernel Estimates of the Empirical Distributions of the AR(1) Coefficients Obtained from Single-Equation ADF Tests, GEMU Sample



Note: *Figure 3 plots kernel estimates of the empirical distribution of the AR(1) coefficients obtained from single-equation ADF tests. AR(1) coefficients are retrieved by regressing the change of the real exchange rate on its past level and six lagged difference terms. Estimations were performed for all relative price series available in the GEMU sample (666).*

References

Campbell, J. and Perron, P. (1991). Pitfalls and opportunities: What macroeconomists should know about unit roots. In Blanchard, O. J. and Fischer, S., editors, NBER Macroeconomics Annual 1991, pages 141–201. MIT Press, Cambridge and London.

Cecchetti, S. G., Mark, N. C., and Sonora, R. (2000). Price level convergence among United States cities: Lessons for the European Central Bank. NBER Working Paper Series No. 7681.

Crucini, M. J., Telmer, C. I., and Zachariadis, M. (2000). Dispersion in real exchange rates. Working Paper 01-w20, Vanderbilt - Economic and Business Administration.

Devereux, M. B. and Engel, C. (1998). Fixed vs. floating exchange rates: How price setting affects the optimal choice of exchange-rate regime. NBER Working Paper No. 6867.

Engel, C., Hendrickson, M. K., and Rogers, J. H. (1997). Intra-national, intra-continental and intra-planetary PPP. Journal of the Japanese and International Economies, 11:480–501.

Engel, C. and Rogers, J. H. (1996). How wide is the border? American Economic Review, 86:1112–1125.

Engel, C. and Rogers, J. H. (2001). Deviations from purchasing power parity: Causes and welfare costs. Journal of International Economics, 55:29–57.

Hall, A. R. (1994). Testing for a unit root in time series with pretest data-based model selection. Journal of Business and Economic Statistics, 12:461–470.

Hufbauer, G. C., Rogers, J. H., and Wada, E. (2001). Prive level convergence and inflation in Europe. Institute for International Economics Working Paper No. 01-1, Washington, D.C.

Im, K. S., Pesaran, H. M., and Shin, Y. (2002). Testing for unit roots in heterogeneous panels. Journal of Econometrics.

Kendall, M. (1954). Note on bias in the estimation of autocorrelation. Biometrika, 41:403–404.

Levin, A. and Lin, C.-F. (1992). Unit root tests in panel data: Asymptotic and finite-sample properties. University of California, San Diego Discussion Papers No. 92-93.

Levin, A. and Lin, C.-F. (1993). Unit root tests in panel data: New results. University of California, San Diego Discussion Papers No. 93-56.

Levin, A., Lin, C.-F., and Chu, C. (2002). Unit roots in panel data: Asymptotic and finite-sample properties. Journal of Econometrics, 108:1–24.

Mussa, M. (1986). Nominal exchange rate regimes and the behavior of real exchange rate: Evidence and implications. Carnegie Rochester Conference Series on Public Policy, 25:117–214.

Nickell, S. J. (1981). Biases in dynamic models with fixed effects. Econometrica, 49:1417–1426.

O'Connell, P. G. and Wei, S.-J. (1997). The bigger they are, the harder they fall: How price differences across U.S. cities are arbitraged. NBER Working Paper Series No. 6089.

Oh, K.-Y. (1996). Purchasing power parity and unit root tests using panel data. Journal of International Money and Finance, 15:405–418.

Parsley, D. C. and Wei, S.-J. (1996). Convergence to the law of one price without trade barriers or currency fluctuations. Quarterly Journal of Economics, 111:1211–1236.

Parsley, D. D. and Wei, S.-J. (2001a). Explaining the border effect: The role of exchange rate variability, shipping costs, and geography. Journal of International Economics, 55(1):87–106.

Parsley, D. D. and Wei, S.-J. (2001b). Limiting currency volatility to stimulate goods market integration: A price based approach. NBER Working Paper Series No. 8468.

Wu, Y. (1996). Are real exchange rates nonstationary? Evidence from a panel-data set. Journal of Money, Credit and Banking, 28:54–63.

A Data

An overview of the countries and regions that are included in our study is given in table A of section B of the appendix. As one can see there we are using monthly price data for 13 German states ('Laender'), 20 Austrian cities, 5 Finnish regions, 20 Italian cities, 18 Spanish regions ('communidades'), 7 Portuguese regions and 4 Swiss cities. As table A indicates, all data were retrieved either from a country's national statistical office (Austria, Finland, Italy, Spain and Portugal) or from the respective region's statistical office (Germany and Switzerland).

For Germany, Finland, Italy, Spain and Portugal we have available both the aggregate price index and twelve sub-indices. The sub-indices follow the same - so-called COICOP - classification schemes across all countries. Out of these twelve sub-indices we use ten (Food and non alcoholic beverages (short name: food), alcoholic beverages and tobacco (alco), clothing and footwear (clot), housing, water, electricity, gas and other fuels (hous), furnishings, household equipment and routine maintenance of the house (furn), health (heal), transport (tran), communications (comm), recreation and culture (recr) and hotels, cafes and restaurants(hote)). We exclude the category 'miscellaneous goods and services' as it is too unspecific and the category 'education'. The reason for the exclusion of the latter category is best illustrated by a plot of the index. As figure A of section C shows educational systems across European countries seem to be strongly regulated. Thus, an economic analysis of the dynamics of relative prices is little promising. All of the price data (for all countries) are seasonally unadjusted. All of the used categories of goods are mutually exclusive. Together they comprise a large fraction of the overall CPI.

Consumer price data are closer to being monthly average data than to being point-in-time data. In order to compare prices internationally we use a monthly average exchange rates from the IMF (International Financial Statistics). We also use data on the distance between cities. Our distance measure is the great-circle distance computed from the latitude and longitude data of each city included in our sample.

B Tables

Table A: Included Regions/Cities

Germany (Short Name: ge, 12 regions)
Baden-Wuerttemberg (Stuttgart), Bayern (Muenchen), Western Berlin, Eastern-Berlin, Brandenburg (Potsdam), Hessen (Wiesbaden) Mecklenburg-Vorpommern (Schwerin), Niedersachsen (Hannover), Nordrhein-Westfalen (Duesseldorf), Saarland (Saarbruecken) Sachsen (Dresden), Sachsen-Anhalt (Magdeburg), Thueringen (Erfurt) **Source**: Statistical Offices of the German states; **Coverage**: 1991.01 - 2002.12
Austria (Short Name: au, 20 cities)
Amstetten, Baden, Bregenz, Dornbirn, Eisenstadt, Feldkirch, Graz, Innsbruck, Kapfenberg, Klagenfurt, Krems, Linz, Salzburg, Steyr, St.Poelten, Villach, Wels, Wien, Wiener Neustadt, Wolfsberg **Source**: Statistics Austria; **Coverage**: 1991.01 - 2002.12
Finland (Short Name: fi, 5 regions)
Uusimaa (Helsinki), Southern Finland (Tampere), Eastern Finland (Joensuu), Mid-Finland (Kokkola), Northern Finland (Oulu) **Source**: Statistics Finland; **Coverage**: 1995.01 - 2002.12
Italy (Short Name: it, 20 cities)
Ancona, Aosta, Bari, Bologna, Cagliari, Campobas, Firenze, Genova, L'Aquila, Milano, Napoli, Palermo, Perugia, Potenza, Reggio Calabria, Roma, Torino, Trento, Trieste, Venezia **Source**: Istituto Nazionale di Statisticia (ISTAT); **Coverage**: 1991.01 - 2002.12
Spain (Short Name: sp, 18 provinces)
Andalucia (Seville), Aragon (Saragossa), Principado de Asturias (Oviedo), Baleares (Palma de Mallorca), Canarias (La Palma), Caabria (Santander), Castilla y Leon (Valladolid), Castilla La Mancha (Albacete), Cataluna (Barcelona), Ceuta y Melilla (Ceuta), Extremadura (Badajoz), Galicia (LaCoruna), Communidad Madrid (Madrid), Cummunidad Murcia (Murcia), Navarra (Pamplona), Pais Vasco (San Sebastian), La Rioja (Logrona), Communidad Valenicana (Valencia) **Source**: Institutio Nacional de Estadistica (INE); **Coverage**: 1991.01 - 2002.12
Portugal (Short Name: po, 7 regions)
Acores (Ponta Delgada), Algarve (Faro), Altenejo (Evora), Centro (Coimbra), Lisbon (Lisbon), Madeira (Funchal), Norte (Vila Real) **Source**: Institutio Nacional de Estatistica (INE); **Coverage**: 1991.01 - 2002.12
Switzerland (Short Name: ch, 4 regions)
Basel, Bern, Genf, Zurich **Source**: Statistical Offices of the respective Cities; **Coverage**: 1991.01 - 2002.12

Notes:

1) Data for COICOP subcategories (Germany, Finland, Italy, Spain and Portugal) are available for the period from 1995.01 to 2002.12 only. For Austria and Switzerland no data for subcategories are used.

2) When data are available only for a larger region (such as a state in Germany), the city reported in brackets is taken to compute distances.

Table B: All Items, Descriptive Statistics, Overall Period, Volatility Measure 1, Individual Results

Loc.	mean (std.dvt.)	dist	n.e.r.vol.	Loc.	mean (std.dvt.)	dist	n.e.r.vol.
ge-ge	1.93 (0.44)	338	0.0	fi-it	23.90 (0.36)	2375	24.54
ge-au	5.53 (0.7)	513	0.42	fi-sp	12.70 (0.32)	3096	12.24
ge-fi	8.98 (0.27)	1686	8.23	fi-po	14.61 (0.77)	3731	13.74
ge-it	24.37 (0.36)	897	24.85	fi-ch	15.87 (0.35)	2073	15.85
ge-sp	11.22 (0.31)	1478	10.25	it-it	2.86 (0.37)	451	0.0
ge-po	13.84 (0.65)	2201	12.85	it-sp	17.65 (0.46)	1371	17.29
ge-ch	14.45 (0.24)	431	12.78	it-po	15.69 (0.83)	2155	14.77
au-au	4.54 (1.07)	214	0.0	it-ch	30.65 (0.42)	630	30.77
au-fi	9.61 (0.46)	1812	8.23	sp-sp	2.87 (0.76)	404	0.0
au-it	25.01 (0.45)	634	24.84	sp-po	9.78 (0.8)	891	6.54
au-sp	12.11 (0.47)	1607	10.16	sp-ch	17.75 (0.26)	1132	17.56
au-po	14.24 (0.64)	2387	12.83	po-po	5.66 (2.03)	773	0.0
au-ch	14.29 (0.53)	515	12.73	po-ch	21.09 (0.64)	1895	20.64
fi-fi	2.01 (0.25)	336	0.0	ch-ch	2.62 (1.21)	129	0.0

Notes:
1) Relative price dispersion between region i and region j, denoted as $V(q_{ij})$, is computed as the standard deviation of two-month relative price changes between the two regions, i.e.,

$$V(q_{ij}) = \sqrt{var(\Delta q_{ij,t})},$$

where $\Delta q_{ij,t}$ denotes the two-month change in regions' i and j relative price and $var(.)$ denotes the empirical variance of $\Delta q_{ij,t}$.

2) For an explanation of the employed country short names, see table A.

3) There are 3240 observations (81 regions).

Table C: Subcategories, Descriptive Statistics, Overall Period, Volatility Measure 1, Individual Results

Loc.	food	alco	clot	hous	furn	heal	tran	comm	recr	hote	dist	n.e.r.v.
intra-nat	7.03	6.23	9.58	7.62	4.97	7.32	5.29	2.59	9.51	8.62	365.00	
	(2.83)	(7.2)	(10.58)	(2.57)	(1.59)	(3.21)	(2.94)	(3.02)	(4.18)	(3.51)		
ge-fi	12.61	11.19	49.48	10.92	9.86	20.10	13.02		19.78	22.46	1686	8.23
	(0.66)	(1.09)	(3.13)	(0.8)	(0.39)	(5.26)	(1.11)		(2.5)	(7.6)		
ge-it	13.98	15.68	10.88	13.75	10.41	20.77	13.08		22.40	22.86	897	8.82
	(0.78)	(1.44)	(0.87)	(1.72)	(0.76)	(6.08)	(1.4)		(3.33)	(7.06)		
ge-sp	17.12	20.97	27.82	12.48	11.54	20.74	13.74		23.79	22.25	1478	10.25
	(1.72)	(2.27)	(4.28)	(1.79)	(0.86)	(5.41)	(1.45)		(3.08)	(7.19)		
ge-po	18.26	17.14	58.29	15.52	14.04	22.28	17.22		24.15	24.96	2201	12.85
	(2.55)	(1.99)	(20.1)	(1.41)	(1.17)	(5.26)	(1.13)		(2.07)	(6.8)		
fi-it	13.87	15.10	52.45	16.22	12.39	15.24	14.92	26.95	14.79	13.47	2375	9.87
	(0.78)	(0.4)	(2.67)	(1.54)	(0.8)	(1.27)	(0.87)	(3.49)	(2.73)	(1.13)		
fi-sp	17.42	20.50	50.54	15.29	14.49	15.06	14.71	25.15	24.66	14.63	3096	12.24
	(1.3)	(2.28)	(3.26)	(1.9)	(0.94)	(0.82)	(0.66)	(2.7)	(4.33)	(1.4)		
fi-po	19.28	15.73	64.78	17.16	15.90	16.83	18.40	30.45	16.88	16.26	3731.	13.74
	(3.32)	(1.34)	(10.02)	(1.43)	(1.06)	(0.76)	(1.11)	(3.55)	(0.89)	(1.58)		
it-sp	13.05	20.51	28.77	13.33	9.47	11.19	11.76	17.21	24.78	12.54	1371	7.77
	(1.14)	(3.01)	(4.58)	(1.96)	(0.93)	(1.63)	(2.22)	(0.51)	(5.29)	(1.71)		
it-po	15.68	15.10	60.30	14.40	11.26	12.23	12.72	16.80	13.68	13.13	2155	7.57
	(4.01)	(0.75)	(22.01)	(2.04)	(1.57)	(2.01)	(2.08)	(0.86)	(3.09)	(1.89)		
sp-po	15.62	19.23	56.44	12.95	10.41	10.18	14.44	18.30	23.49	12.73	891	6.54
	(3.42)	(3.23)	(16.41)	(1.92)	(1.42)	(1.52)	(2.27)	(1.44)	(4.73)	(2.03)		

Notes:

1) Relative price dispersion between region i and region j, denoted as $V(q_{ij})$, is computed as the standard deviation of two-month relative price changes between the two regions, i.e.,

$$V(q_{ij}) = \sqrt{var(\Delta q_{ij,t})},$$

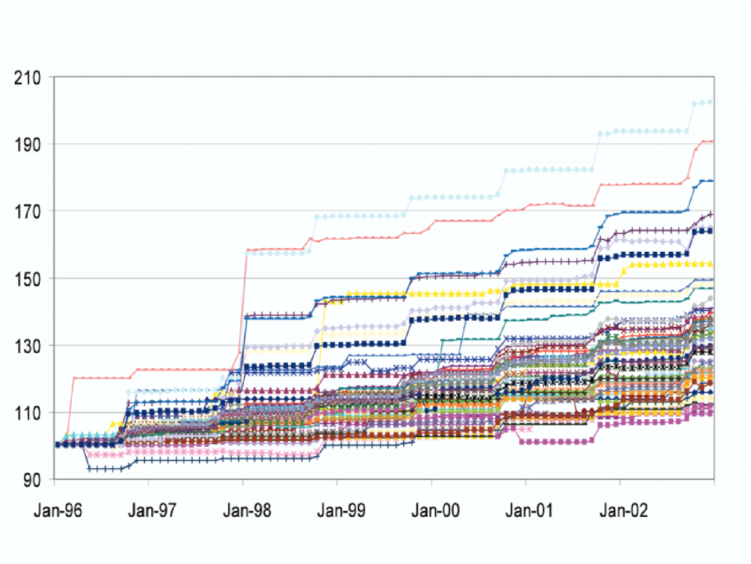
where $\Delta q_{ij,t}$ denotes the two-month change in regions' i and j relative price and $var(.)$ denotes the empirical variance of $\Delta q_{ij,t}$.

2) For an explanation of the employed country short names, see table A.

3) There are 1596 observations (57 regions).

C Figures

Figure A: Education (1996.01 = 100)



Note: *Figure A plots the indices for the COICOP subcategory 'Education' of our German, Finnish, Italian, Spanish and Portuguese regions. Data are normalized such that 1996.01 corresponds to 100 in all cases.*

D Levin-Lin Panel Unit Root Test

D.1 The Test Procedure

To obtain the Levin-Lin panel-unit root results in section 7.2.2, we proceed as follows: Let $q_{i,t}$ (with $i = 1, 2, \ldots, N$ and $t = 1, 2, \ldots, T$) be a balanced panel of real exchange rates consisting of N individuals with T observations, respectively. The starting point of our analysis is the following test equation:

$$\Delta q_{i,t} = \rho_i q_{i,t-1} + u_{i,t}, \tag{D.1}$$

where $-2 < \rho_i \leq 0$, and $u_{i,t}$ has the following error-components representation

$$u_{i,t} = \alpha_i + \theta_t + \epsilon_{i,t}. \tag{D.2}$$

In this specification, α_i represents an individual-specific effect, θ_t represents a common-time effect and $\epsilon_{i,t}$ is a (possibly serially correlated) stationary idiosyncratic shock. The Levin-Lin test procedure imposes (both for the null hypothesis of non-stationarity and for the alternative hypothesis of stationarity) the homogeneity restriction that all ρ_i are equal across individuals. Thus, the null hypothesis can be formulated as:

$$H_0 : \rho_1 = \rho_2 = \cdots = \rho_N = \rho = 0,$$

and the alternative hypothesis (that all series are stationary) is given by:

$$H_1 : \rho_1 = \rho_2 = \cdots = \rho_N = \rho < 0.$$

To test this null hypothesis we proceed as follows:

1. First, we control for the common-time effect by subtracting the cross-sectional means:

$$\tilde{q}_{i,t} = q_{i,t} - \frac{1}{N} \sum_{j=1}^{N} q_{j,t} \tag{D.3}$$

Having transformed the dependent variable we proceed with the following test equation:

$$\Delta\tilde{q}_{i,t} = \alpha_i + \rho\tilde{q}_{i,t-1} + \sum_{j=1}^{k_i} \phi_{i,j} \Delta\tilde{q}_{i,t-j} + \epsilon_{i,t}. \tag{D.4}$$

The lagged differences of $\tilde{q}_{i,t}$ are included to control for potential serial correlations in the idiosyncratic shocks $\epsilon_{i,t}$. Whereas we equalize the ρ_i across individuals we allow for different degrees of serial correlation, k_i (with $i = 1, \ldots, N$), across them. The number of lagged differences for each region is determined by the general-to-specific method of Hall (1994) which is recommended by Campbell and Perron (1991).

2. The next step in our testing procedure is to run the following two auxiliary regressions

$$\Delta\tilde{q}_{i,t} = \alpha_{1i} + \sum_{j=1}^{k_i} \phi_{1i,j}\Delta\tilde{q}_{i,t-j} + e_{i,t}. \tag{D.5}$$

$$\tilde{q}_{i,t-1} = \alpha_{2i} + \sum_{j=1}^{k_i} \phi_{2i,j}\Delta\tilde{q}_{i,t-j} + \nu_{i,t-1}. \tag{D.6}$$

and to retrieve the residuals $\hat{e}_{i,t}$ and $\hat{\nu}_{i,t-1}$ from these regressions.

3. These residuals are used to run the regression

$$\hat{e}_{i,t} = \rho_i\hat{\nu}_{i,t-1} + \eta_{i,t}. \tag{D.7}$$

The residuals of (D.7) are used to compute an estimate of the variance of $\eta_{i,t}$:

$$\hat{\sigma}^2_{\eta i} = \frac{1}{T-k_i-1}\sum_{t=k_i+2}^{T}\hat{\eta}^2_{i,t} \tag{D.8}$$

4. Normalizing the OLS residuals $\hat{e}_{i,t}$ and $\hat{\nu}_{i,t-1}$ by dividing them through $\hat{\sigma}_{\eta i}$ yields:

$$\tilde{e}_{i,t} = \frac{\hat{e}_{i,t}}{\hat{\sigma}_{\eta i}} \tag{D.9}$$

$$\tilde{\nu}_{i,t-1} = \frac{\hat{\nu}_{i,t-1}}{\hat{\sigma}_{\eta i}} \tag{D.10}$$

5. The normalized residuals are used to run the following pooled cross-section time-series regression:

$$\tilde{e}_{i,t} = \rho\tilde{\nu}_{i,t-1} + \tilde{\epsilon}_{i,t}. \tag{D.11}$$

Under the null hypothesis,- $\tilde{e}_{i,t}$ is independent of $\tilde{\nu}_{i,t-1}$, i.e., we can test the null hypothesis by testing whether $\rho = 0$. Unfortunately, the studentized coefficient

$$\tau = \frac{\hat{\rho}}{\hat{\sigma}_{\tilde{\epsilon}}\sum\limits_{i=1}^{N}\sum\limits_{t=2+k_i}^{T}\tilde{\nu}^2_{i,t-1}}$$

with

$$\hat{\sigma}_{\tilde{\epsilon}} = \frac{1}{NT}\sum_{i=1}^{N}\sum_{t=2+k_i}^{T}\tilde{\epsilon}_{i,t}$$

is not asymptotically normally distributed. Levin and Lin (1993) compute an adjusted test statistic based on τ that it is asymptotically normally distributed. However, we do not make use of their adjustment procedure but use bootstrap methods to compute critical values for the null hypothesis. This procedure is described in section D.2.

D.2 Bootstrap Procedure

Since the finite-sample properties of the adjusted τ statistics are unknown and since idiosyncratic shocks may be correlated across individual regions we rely on bootstrap methods to infer critical values for the τ statistics. More precisely, we employ a nonparametric bootstrap where we resample the estimated residuals from our model. The starting point of our bootstrap approach is given by the hypothesized data generating process (DGP) under the null hypothesis

$$\Delta q_{i,t} = \sum_{j=1}^{k_i} \phi_{i,j} \Delta q_{i,t-j} + \epsilon_{i,t}. \tag{D.12}$$

Our procedure is as follows:

1. We retrieve the OLS residuals from estimating the DGP under the null hypothesis. This yields the vectors $\hat{\epsilon}_1$, $\hat{\epsilon}_2$, ..., $\hat{\epsilon}_T$, where $\hat{\epsilon}_t$ is the $1xN$ residual vector for period t.

2. Then, we resample these residual vectors by drawing one of the possible T residual vectors with probability $\frac{1}{T}$ for each $t = 1, \ldots, T$.

3. These resampled residual vectors are used to recursively build up pseudo-observations $\Delta\hat{q}_{i,t}$ according to the DGP (using the estimated coefficients $\hat{\phi}_{i,j}$).

4. Next, we perform the Levin-Lin test (as described in section D.1) on these observations (without subtracting the cross-sectional mean). The resulting τ is saved.

5. Steps two to four are repeated 5,000 times. The collection of the τ statistics form the bootstrap distribution of these statistics under the null hypothesis.

E Im-Pesaran-Shin Panel Unit Root Test

E.1 The Test Procedure

To obtain the Im-Pesaran-Shin panel-unit root results in section 7.2.2, we proceed as follows: Let $q_{i,t}$ (with $i = 1, 2, \ldots, N$ and $t = 1, 2, \ldots, T$) be a balanced panel of real exchange rates consisting of N individuals with T observations, respectively. Following Im et al. (2002) we start our analysis by estimating the following ADF test equation

$$\Delta\tilde{q}_{i,t} = \alpha_i + \rho_i\tilde{q}_{i,t-1} + \sum_{j=1}^{k_i} \phi_{i,j}\Delta\tilde{q}_{i,t-j} + \epsilon_{i,t} \tag{E.1}$$

for each of the N individual real exchange rate series. In this equation the tilde above the variable q indicates that the cross-sectional mean has been subtracted from the real exchange rate series, i.e.,

$$\tilde{q}_{i,t} = q_{i,t} - \frac{1}{N}\sum_{j=1}^{N} q_{j,t}. \tag{E.2}$$

As the subindex i for the parameter k indicates we allow the number of included lagged differences to vary across individual series. For each series the number of included lags is determined according to the general-to-specific method by Hall (1994), recommended by Campbell and Perron (1991). The maximum number of lags is set to six.

The Im-Pesaran-Shin test procedure imposes for the null hypothesis of non-stationarity the homogeneity restriction that all ρ_i are equal across individuals. Thus, the null hypothesis can be formulated as:

$$H_0 : \rho_1 = \rho_2 = \cdots = \rho_N = \rho = 0.$$

Unlike the Levin-Lin test, however, there is no analogous homogeneity condition for the alternative hypothesis of stationarity which is given by:

$$H_1 : \rho_1 < 0 \cup \rho_2 < 0 \cup \cdots \cup \rho_N < 0.$$

To test this null hypothesis we individually estimate equation (E.1) for each relative price series and retrieve for each equation the studentized coefficient $\hat{\tau}_i$ which is given by $\frac{\hat{\rho}_i}{\hat{\sigma}_{\rho_i}}$ (where $\hat{\sigma}_{\rho_i}$ denotes the standard deviation of the estimated adjustment coefficient ρ_i). The panel unit root test statistics τ_{ips} is then obtained by averaging the t-values of the estimated ρ-coefficients, i.e.,

$$\tau_{ips} = \sum_{i=1}^{N} \hat{\tau}_i. \tag{E.3}$$

Im et al. (2002) show that this statistics is asymptotically standard normally distributed. However, we do not make use of this result (partly as it relies on the assumption that the errors ϵ_{it} are independent across individuals). The critical values reported in the main text are obtained via a non-parametric bootstrap procedure that is described in subsection E.2.

E.2 Bootstrap Procedure

Since the finite-sample properties of the Im-Pesaran-Shin test statistics τ_{ips} might differ considerably from their asymptotic properties and since idiosyncratic shocks may be correlated across individual regions we rely on bootstrap methods to infer critical values for the τ_{ips} statistics. As for the Levin-Lin test, we employ a non-parametric bootstrap where we resample the estimated residuals from our model. The starting point of our bootstrap approach is given by the hypothesized data generating process (DGP) under the null hypothesis

$$\Delta q_{i,t} = \sum_{j=1}^{k_i} \phi_{i,j} \Delta q_{i,t-j} + \epsilon_{i,t}. \tag{E.4}$$

Our procedure is as follows:

1. We retrieve the OLS residuals from estimating the DGP under the null hypothesis. This yields $\hat{\epsilon}_1$, $\hat{\epsilon}_2$, ..., $\hat{\epsilon}_T$, where $\hat{\epsilon}_t$ is the $1xN$ residual vector for observation t.

2. Then, we resample these residual vectors by drawing one of the possible T residual vectors with probability $\frac{1}{T}$ for each $t = 1, \ldots, T$.

3. These resampled residual vectors are used to recursively build up pseudo-observations $\Delta\hat{q}_{i,t}$ according to the DGP (using the estimated coefficients $\hat{\phi}_{i,j}$).

4. Next, we perform the Im-Pesaran-Shin test (as described in subsection E.1) on these observations (without subtracting the cross-sectional mean). The resulting test statistic $\hat{\tau}$ is saved.

5. Steps two to four are repeated 5,000 times. The collection of the $\hat{\tau}$ statistics form the bootstrap distribution of these statistics under the null hypothesis.

CFS Working Paper Series:

No.	Author(s)	Title
2001/02	**Roland Beck**	Do Country Fundamentals Explain Emerging Market Bond Spreads?
2001/03	**Markus Kern** **Bernd Rudolph**	Comparative Analysis of Alternative Credit Risk Models – An Application on German Middle Market Loan Portfolios –
2001/04	**Antje Brunner** **Jan Pieter Krahnen**	Multiple Lenders and Corporate Distress: Evidence on Debt Restructuring
2001/05	**Ralf Ewert** **Andrea Szczesny**	Countdown for the New Basle Capital Accord. Are German Banks Ready for the Internal Ratings-Based Approach?
2001/06	**Bernd Kaltenhäuser**	Explaining the Dollar-Euro Rate: Do Stock Market Returns Matter?
2001/07	**Guenter Beck** **Axel A. Weber**	How wide are European borders? On the Integration Effects of Monetary Unions
2001/08	**Yunus Aksoy** **Tomasz Piskorski**	Domestic Money and US Output and Inflation
2001/09	**Elke Hahn**	Core Inflation in the Euro Area: Evidence from the Structural VAR Approach
2001/10	**Olaf Ehrhardt** **Eric Nowak**	Private Benefits and Minority Shareholder Expropriation – Empirical Evidence from IPOs of German Family-Owned Firms
2001/11	**Daniel Gross**	Country-Specific and Global Shocks in the Business Cycle
2001/12	**Daniel Gross**	Trade Flows and the International Business Cycle
2001/13	**Guenter Beck** **Axel A. Weber**	Economic Integration and the Exchange Rate Regime: How Damaging are Currency Crises?

Copies of working papers are available at the Center for Financial Studies or can be downloaded (http://www.ifk-cfs.de).