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Endogenous Optimum Currency Areas and the Blend of Sectors – On the Determinants of Business Cycle Correlation across European Regions

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Abstract

This paper examines the degree of correlation of EU regional employment cycles and attempts to show whether these cycles reflect changing patterns of specialisation. By focusing on the regional level and by employing three different indicators of similarity of sectoral structure, it improves on existing studies. A dynamic panel data model is estimated for region pairs by within groups, i.e., by a standard fixed effects estimator. Special attention is paid to capture the rich dynamics which is typical of employment data. The key finding is that employment growth is more synchronised when regions look alike in their sectoral structure. The empirical results again highlight the problem of a common monetary policy for uncommon regions within the euro zone.

- **JEL classification:** E32, F15, R23
- **Keywords:** Regional Employment, European Union, Regional Business Cycles, Specialisation, Synchronicity

I. Introduction

The integration of nations within Europe has been an important issue for four decades now. With respect to the national level, there exists a variety of empirical studies focussing on the similarities and differences in sectoral structures (for

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example, the level of national specialisation or its inverse, the level of similarity), the sources of national business cycles and the importance of industry-specific shocks. Considerable theoretical and empirical research on this general issue has been conducted, but this line of research has usually emphasised differences across countries, assuming implicitly (instead of testing explicitly) that countries themselves are homogenous entities. Although the implicit assumption of homogenous entities is not warranted, still very few studies combine measures of regional instead of national specialisation with the co-movement of regional business cycles. In this paper we explore the relationship between regional employment cycles and the similarity of bilateral regional sectoral structures, and, secondly test the hypothesis that the similarity of regional industrial patterns enhances the bilateral correlation of regional employment cycles.

The structure of this paper is as follows. In section 2, we motivate the paper based on a survey of the empirical literature dealing with the determinants of the synchronicity of European national and regional employment cycles. In section 3, we explain the specification of the relevant variables in detail and give some information about data sources. Above all, we provide an in-depth-discussion of the construction of and the relative merits of several alternative bilateral indicators, measuring the degree of similarity between the sectoral structures of two regions. We also give some stylised facts about the potential correlation between regional business cycle synchronisation, the degree of regional similarity in the sectoral structure and about the results of panel unit root tests. In section 4, we introduce the empirical model, describe and give reasons for the chosen estimation procedure and discuss the empirical results of the panel regressions. The empirical method adopted is to estimate a dynamic panel data model for region pairs by within groups, i.e., by a standard fixed effects estimator. This model relates (rolling) correlations between regional employment cycles to a measure of sectoral similarity and to differences between regional incomes. Special attention is paid to capture the rich dynamics which is typical of employment data. We also test whether the results are robust to the inclusion of additional variables. In section 5, we summarise our findings and draw some policy conclusions. The key finding is that employment growth is indeed more synchronised in a statistically significant way when regions have a more identical sectoral structure. This evidence is robust with respect to the inclusion of additional variables such as differences in incomes between regions and to different specifications of specialisation indices.

II. Motivation

This paper takes uses the above literature as a starting-point and tries to trace back the synchronicity of regional employment cycles to idiosyncratic sectoral developments within different European regions. These regions are – as a stylised fact – characterised by the emergence of agglomeration. That is, certain areas experience a change in their employment figures and at the same time economic activity becomes more or less concentrated in certain industry branches. This in turn has on the one hand a sustained impact on the regions' economic performance and on the other hand the regional sectoral structure might differ in comparison to other regions. In order to examine the relationship between the similarities of regional sectoral structures and the fluctuations of the regional employment cycle, we intend to quantify changes of the regional sectoral structure caused by European integration over the last two decades. For this purpose, we construct a number of annual indices measuring the degree of similarity between the sectoral structures of two regions (in the following called "specialisation indices") for 30 European regions. It is generally acknowledged that a disaggregated representation of an employment cycle displays higher informational contents than its aggregated representation because different regional developments cancel out at the aggregated level. Moreover, the agglomeration phenomenon can more accurately be grasped by a higher resolution. Hence, we decided to follow Fatás (1997), Forni and Reichlin (1997) and Clark and Wincoop (2001) and focus on the regional instead of the national dimension.

Our analysis differs from the above cited papers however in several respects. Firstly, and in contrast to Artis and Zhang (1997), our procedure is not limited to the analysis of benchmark cycles. Instead, we correlate all possible region-pairs with each other. Second, we emphasise the regional dimension in the same way as Fatás (1997) by using NUTS1 regional employment data. However, in this paper we de-trend the latter with a more useful econometric filter technique. Third, we do not emphasise the role of nominal impact variables, such as the exchange rate regime, but focus on real variables, i.e. similarities in the sectoral structure, in determining the cyclical behaviour of the employment time series. Fourth, we dispense with an investigation the correlation between regional cycles and national or European cycles. Instead, we focus on the impacts of a change in the regional sectoral pattern of production on the degree of correlation of regional employment cycles.

Nearly all contributions in this field investigate the sources of national co-fluctuations empirically, but only a small number (Fatás, 1997, Barrios et al., 2001, Clark and Wincoop, 2001) deliver a detailed analysis at the regional level. However, studies which examine the relationship between regional sectoral patterns and the synchronicity of the employment cycles for a wide range of European regions are not as yet available[1]. Since industry-specific shocks usually play a more important role at the regional rather than at the cross-national level (Clark and Shin, 2000), such studies would fill an existing gap in the literature. Closest to this demand is the study by Clark and Wincoop (2001) which identifies correlations between regional cycles based on border effects and on a measure of the similarities in the bilateral national instead of regional production patterns. In other words, their study contributes to filling this gap but still applies an identical economic structure for each region of the respective nation. As an innovation, we start with their methodology but focus on the bilateral similarities or differences in regional instead of national production patterns determining the synchronicity of the employment cycles between these regions.

One branch of theory which tries to explain the choice of industries to locate in a certain region is the New Economic Geography literature, which explicitly conveys detailed theoretical information about the impact of increasing economic integration on the development of industrial structures. Moreover, this literature implicitly gives information regarding the resulting sectoral structures as well as the specialisation patterns. Under the scenario of increasing economic integration, there will be advantages for firms of one industry to cluster in a certain region – these agglomeration forces can further be reinforced by themselves ("cumulative causation") and will tend to encourage concentration of industrial activity by such cumulative causation. Centripetal or centrifugal forces are commonly regarded as the main reasons for such cumulative causation (for a survey see, e.g., Krugman, 1998). These forces have a significant impact on the decision of mobile production factors to agglomerate or deglomerate geographically. Changes in the sectoral structure of the respective economy will be the result. However, centripetal or centrifugal forces are themselves determined by the degree of integration or, to be more precise, by the magnitude of transportation costs.

In the typical NEG model, specialisation and agglomeration processes emerge within the manufacturing sector, depending on the transport costs and economies of

[1]One exception is Barrios et al. (2001) who investigate the impact of sectoral specialisation on the co-fluctuations among U.K. business cycles and on the co-fluctuations of the latter with EU country cycles.

scale in the respective industry. One of the sector investigated in this paper, i.e. ‘manufactured products’, is especially relevant from an NEG point of view. So the blend of sectors (manufacturing versus primary and tertiary) seems to offer an additional hint as to how far a region has come in terms of structural change or in terms of agglomeration rather than the specialisation of manufacturing industries. Some features of the New Economic Geography are used in the next section to construct several indices to quantify the degree of similarity in the sectoral structures of two regions.

III. Data and Stylised Facts

In order to test the conjectured impact of the degree of similarity (regional specialisation patterns) on the synchronicity of regional employment cycles empirically we employ a panel of 30 European regions from six countries, namely Belgium, France, (West) Germany, Ireland, the Netherlands and Spain. The sample for the final regressions runs from 1989 to 1996. This specific choice of regions is determined by the current limitations of data availability. However, it also makes sense from an economic point of view because these regions at least represent an area which has been called the core European Monetary Union (EMU) and for which the synchronicity of cycles as a precondition of a well-functioning currency union have been discussed exhaustively. From the data we were able to construct 435 region-pairs. These regions-pairs will later on represent the cross-sections for our panel regression analysis. The definition of regions used in this paper is based on the Nomenclature of Territorial Units for Statistics of EUROSTAT, namely the NUTS 1 level[2]. The NUTS was established by Eurostat to provide comparable regional breakdowns of the Member States of the European Union. In Table 1, all European regions included in our empirical analysis are listed.

For each of these 435 region-pairs, the degree of employment cycle synchronicity as well as the degree of bilateral similarity in the sectoral structure of both regions is determined as follows.

A. Synchronicity of regional employment cycles and bilateral regional specialisation patterns: specification and data

In order to determine the degree of employment cycle synchronicity, the regional

[2]We assume that Ireland is one region, and the reduction of the eleven West German regions to eight avoids using the so-called city-states in Germany.

Table 1. European NUTS 1 Regions: A Survey

Region	Nomenclature Eurostat	Nomenclature used in this paper
	BelgiUM	
RÈGION BRUXELLES-CAP.	BE 1	BE 1
VLAAMS GEWEST	BE 2	BE 2
RÈGION WALLONE	BE 3	BE 3
	France	
ÎLE DE FRANCE	FR 1	FR 1
BASSIN PARISIEN	FR 2	FR 2
NORD-PAS-DE-CALAIS	FR 3	FR 3
EST	FR 4	FR 4
OUEST	FR 5	FR 5
SUD-OUEST	FR6	FR6
CENTRE-EST	FR 7	FR 7
MÈDITERRANÈE	FR 8	FR 8
	Germany	
SCHLESWIG-HOLSTEIN	DE F	G 1
HAMBURG	DE 6	
NIEDERSACHSEN	DE 9	G 2
BREMEN	DE 5	
NORDRHEIN-WEST-FALEN	DE A	G 3
HESSEN	DE 7	G 4
RHEINLAND-PFALZ	DE B	G 5
SAARLAND	DE C	
BADEN-WÜRTEMBERG	DE 1	G 6
BAYERN	DE 2	G 7
BERLIN	DE 3	G 8
	Ireland	
IRLAND	IE	IRL
	Netherlands	
NOORD-NEDERLAND	NL 1	NL 1
OOST-NEDERLAND	NL 2	NL 2
WEST-NEDERLAND	NL 3	NL 3
ZUID-NEDERLAND	NL 4	NL 4
	Spain	
NOROESTE	ES 1	ESP 1
NORESTE	ES 2	ESP 2
COMUNIDAD DE MADRID	ES 3	ESP 3
CENTRO (E)	ES 4	ESP 4
ESTE	ES 5	ESP 5
SUR	ES 6	ESP 6
CANARIS	ES 7	ESP 7

Source: EUROSTAT (1999).

employment cycle for each European region is approximated by the de-trended time series of total regional employment. For the de-trending procedure the Hodrick-Prescott-Filter with parameter 10 is used in order to transform the non-stationary variable regional employment to a stationary one[3]. The advantages of this standard practice are, firstly, that it is easy to implement and, secondly, that the resulting cyclical residuals are similar to those of the band-pass filter introduced by Baxter and King (1999). We use the regional employment data provided by EUROSTAT. These data are available for the Belgian regions for a time span ranging from 1975 to 1997, for the French, West German and Dutch regions as well as for Ireland from 1970 to 1997 and for the Spanish regions from 1980 to 1997. To quantify the synchronicity of the bilateral regional employment cycles we use the well-known Bravais-Pearson correlation coefficient between both cycles. We base the correlation coefficient on a rolling window with a reasonable size of five periods[4].

In order to quantify the degree of similarity in bilateral regional production patterns and to use these results for our regressions in section 4, we employ popular indices used by many scholars in the empirical literature[5]. As an innovation, we do not focus on just one indicator. Instead and in order to conduct a "sensitivity analysis" our estimations are based on three alternative indices, each having their own merits. As a first indicator we use the index of conformity CON (Imbs, 1999). It is constructed analogously to the usual correlation coefficient, but without consideration of the statistical mean. Its realisations lie between zero (the sectoral structure of both regions observed is totally different) and one (both regions have the same sectoral structure). A disadvantage of this indicator is that the value of the results tends to be near one, even if there are significant differences in the sectoral structure between two regions. Hence, this index has to be interpreted rather cautiously.

[3]See Hodrick and Prescott (1997). Additionally, we conducted panel unit root tests which clearly rejected the non-stationarity for the generated series. See section IV.

[4]See Inklaar and DeHaan (2001) who use just five periods to calculate the correlation coefficients; for a 10 year window see, e.g., Caporale et al. (1999). In this paper, we dispense with using ten-years windows for constructing the rolling correlation coefficient (as, e.g., ECB, 1999) in order to avoid losing ten degrees of freedom. This would not be the case if monthly or higher frequency data were available. As an additional cross-check we also captured movements in synchronicity over time by recursive estimates of bilateral correlations in an earlier version of this paper. See Belke and Heine (2001).

[5]Surveys of the advantages and disadvantages of such indicators are widespread and can be found in Krieger-Boden (1999), Midelfart-Knarvik et al. (2000) or WIFO (1999).

Secondly, we use the Finger-Kreinin index FIN[6]. It is defined as the sum of the minima of the sectoral shares of two regions. The higher the empirical realisation of this index (maximum 1), the more similar the sectoral structures between the two regions are. Although this index compares two regions sector by sector, the strategy of summing up the minima might be a problem. For instance, the value of the index is unchanged when there are significant changes in the sectoral structure of one region but no changes in the minima.

The third and final index we employ is the specialisation coefficient SPEC. This was employed by e.g., by Krugman (1991), Molle (1997), Clark and Wincoop (2001), Kim (1999) and OECD (1999). This index allows for realisations between zero (both regions have the same sectoral structure) and two (the sectoral structure of both regions observed is totally different). In a strictly verbal sense CON and FIN are measures of similarity since their empirical realisations grow with increasing similarity between sectors and only SPEC is an index of specialisation / diversification since its realisation shrinks with increasing similarity. However, all the indices which we apply here try to measure the bilateral degree of similarity and quantify the regional specialisation patterns. The three indices employed are displayed and explained in Table 2.

A change in all of these specialisation indices can principally be put down to the fact that the relative shares of the sectors have changed. The more industrial sectors are available, the more meaningful the indicators are. However, in our case the number of available sectors might probably pose a slight problem but cannot be completely solved because of the limited data availability. In order to construct the different indicators we use the time series of the nominal regional gross value added for the 30 European regions, ranging from 1975 to 1996 for the Belgian regions, from 1975 to 1994 for the French and West German regions as well as for Ireland, from 1975 to 1993 for the Dutch regions and from 1980 to 1995 for the Spanish regions. Our choice of this variable significantly deviates from Imbs (1999) who estimates a specialisation index based on the sectoral total employment. We do not use employment data because there is lack of sectoral coverage of these data such that the specialisation index would not be very meaningful. In order to avoid such problems inherent in the use of an index of sectoral total employment, we

[6]See, e.g., Finger and Kreinin (1979), Greenaway and Hine (1991), and Amiti (1997).

Table 2. Indicators to quantify the degree of similarity in regional sectoral structures

Indicator	Formula	Range
Index of conformity	$\mathrm{CON} = \dfrac{\sum_{i=1}^{n} a_i \cdot b_i}{\sqrt{\left[\left[\sum_{i=1}^{n} a_i^2\right] \cdot \left[\sum_{i=1}^{n} b_i^2\right]\right]}}$	0 to 1
Finger-Kreinin index	$\mathrm{FIN} = \sum_{i=1}^{n} \min[a_i \cdot b_i]$	0 to 1
Coefficient of specialisation	$\mathrm{SPEC} = \sum_{i=1}^{n} \lvert a_i - b_i \rvert$	0 to 2

Note: The variable n corresponds to the amount of sectors, a_i is the share of sector i in country A and b_i is the share of sector *i* in country B.

use the regional gross value added for six different sectors (agricultural, forestry and fishery products, fuel and power products, manufactured products, building and construction, market services and non-market services). Unfortunately, European regional data do not yet allow a finer disaggregation of these data.

The lower the value of the index of conformity CON and the Finger-Kreinin index FIN, the higher is the degree of specialisation and, hence, the lower is the degree of similarity between the two examined regions and as a result the co-fluctuation between the employment cycles must be very low. With respect to the coefficient of specialisation SPEC the contrary holds. With respect to the coefficients of the "specialisation indices" in our regressions explaining correlations coefficients of regional employment cycles, we expect the following conditions to hold:

$$\frac{\partial KOR}{\partial CON} > 0; \ \frac{\partial KOR}{\partial FIN} > 0; \ \frac{\partial KOR}{\partial SPEC} < 0; \qquad (1)$$
$$0 \leq CON \leq 1 \ \ 0 \leq FIN \leq 1 \ \ 0 \leq SPEC \leq 2$$

with the variable KOR representing the Bravais-Pearson correlation coefficient measuring the 5-year moving correlation between the residuals of Hodrick-Prescott (HP)-filtered employment of two regions. In other words, employment growth should be more synchronised when regions look more alike in their sectoral structure – for example, when manufacturing plays an important role in both regions, or when agriculture dominates. More specifically, the higher the values of the index of conformity and the Finger-Kreinin index are, the more similar the sectoral structures are between the two examined regions and the higher the degree of synchronicity of the employment cycles should be. However, the contrary is valid for the coefficient of specialisation. The higher the value of SPEC, the less similar the sectoral structures are between the two respective regions and the lower the synchronicity of the employment cycles is.

In addition to these different specialisation indices we complement our analysis by including selected additional variables. Firstly, we experimented with a dummy variable to deal with the German unification in 1990 and dummies for common borders to capture geographic proximity. Only in a small number of cases were these variables included in our preferred specification however. Secondly, we include the difference in real income between two regions (see also Imbs, 1999) by taking the absolute value of the difference of the regional GDP deflated by the national consumer price index (annual data from 1975 to 1996). The time periods for the regional GDP data range from 1977 to 1996 for the Belgian regions, from 1980 to 1996 for the French and West German regions, from 1975 to 1996 for Ireland, from 1981 to 1996 for the Dutch regions and from 1980 to 1996 for the Spanish regions. We expect a negative sign on the estimated coefficient of the difference in real incomes, implying a higher correlation of employment cycles if incomes are more similar.

B. Stylised Facts

A variety of stylised facts for business cycles is accepted as valid at the national level by many scholars, but for the regional level such relatively indisputable facts are missing. In particular, this could be caused by the fact that adequate regional time series are currently not available. First steps in this direction were undertaken by Fatás (1997) and Clark and Wincoop (2001). In their investigations they find that the co-movements of economic variables between European regions are decreasing, but not to a statistically significant extent. On the contrary, the correlations of national cycles within Europe seem to be increasing significantly according to several studies. On a more general level, it seems fair to state that the following picture has emerged in the literature. Despite large regional policy expenditures, regional inequalities in Europe have not narrowed substantially over the last two decades, and according to some measures have even widened.

European States have developed increasingly common production structures. But European regions have become increasingly polarised in terms of their employment performance[7]. Seen on the whole, a puzzle emerges of a decreasing synchronicity of within-country regional business cycles alongside with increased synchronicity of national cycles. However, this puzzle can be at least partially solved by searching for possible impact factors (i.e., the several specialisation indices) behind this negative regional trend in synchronicity.

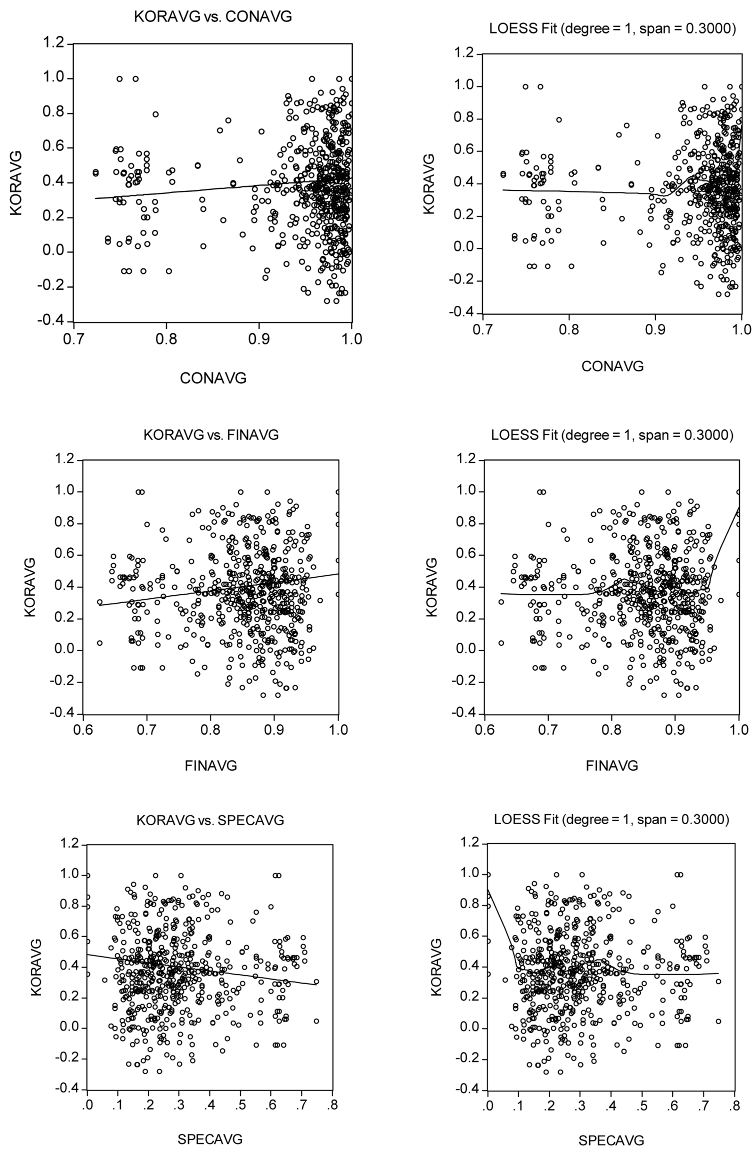
In order to check the empirical validity of the hypothesis that the degree of synchronicity of EU regional business cycles is determined by the degree of similarity we look at the bilateral synchronicity of regional employment cycle correlations. We start with some simple scatter plots, presented in Figure 1. This figure shows cross-plots of each of our bi-regional measures for the specialisation indices (CON, FIN and SPEC) and of the bi-regional correlation coefficients of residuals of the HP-filtered regional employment for all region-pairs. All variables are averaged over the period 1975 to 1997 (if available). In addition, we fit a tentative bi-variate regression of the correlation coefficient on the specialisation index and an intercept. The corresponding estimated regression line is represented by the straight line in each scatter plot. The least squares method, however, is very sensitive to the presence of a few outlying observations. For this reason we also carry out a form of weighted least squares where outlying observations are given less weight in estimating the regression coefficients (see Cleveland, 1993).

As expressed by the regression lines, the conjectured negative (positive) relationship between SPEC (CON and FIN) and correlation coefficient of regional employment growth cannot be rejected to exist at first glance. However, the correlations in the figures appear to be quite weak. Without taking into account additional variables, the extent of specialisation seems to have only weak explanatory power for the employment growth correlations. In the rest of this contribution we tackle this issue more formally.

We start our formal empirical analysis with tests of the non-stationarity of the levels and the first differences of the variables under consideration. One might argue, that we could run our regressions *in levels* since our variables are stationary

[7]See Imbs (1998), Clark and Wincoop (2001), Kalemi-Ozcan et al. (2000). A recent paper by Puga (2002) describes these trends, and discusses how recent location theories can help us to explain them and to reconsider the role of regional policies, especially transport infrastructure improvements, in such an environment.

Fig. 1. Regional employment growth and regional degree of similarity in the EU



Note: The plots display arithmetic averages (AVG) of the respective variables over the whole sample from 1975 to 1997, if available. LOESS indicates the popular Loess (or LOWESS) techniques for local polynomial regressions with bandwidth based on nearest neighbours described, e.g., in Cleveland (1993).

by definition. For example, realisations of the correlation coefficient KOR fall into a range of –1 to +1 by definition of a correlation coefficient. In addition, the latter measures correlations between two time series which are constructed as the residuals of the HP10 filter, hence, are stationary. Table 2 clearly demonstrates the existence of finite upper and lower bounds for the realisations of the specialisation indices. The same is valid with respect to the common border and unification dummies (DUM1, DUM2 and DUMGER lie between zero and one) and the relative income variable (RELINC) which essentially represents a share and therefore takes values between zero and one. However, it is unclear whether the bounded nature of the constructed variables is relevant within the sample, and empirically it is probably better to treat stationarity as a sample property, particularly given the short sample available.

The test we apply here is the first widely used panel data unit root test by Levin and Lin (1992)[8]. This test represents a direct extension of the univariate ADF test setting to panel data. The results by Levin and Lin indicate that panel data is particularly useful for distinguishing between unit roots and highly persistent stationarity in macroeconomic data and that their unit root test for panel data is appropriate in panels of moderate size (between 10 and 250 cross-sections) as encountered in our study.

Table 3 displays the results of applying this unit root test to our set of variables. As usual, we difference the data until it is stationary. In cases I to IV, this leads us to use also the *levels* (and not only the first differences) of the correlation coefficient KOR and the relative income RELINC in our estimations. As expected, the non-stationarity of the levels of all the variables under investigation can be rejected in spite of the rather high (in absolute values) critical values of the test-statistics[9]. We also conducted similar unit root tests for the three different kinds of specialisation indices explained in Table 2 and gained similar results which are available on request.

This leads us to focus solely on the *levels* of the variables of interest, namely the correlation coefficient KOR, the relative income RELINC and the three different specialisation indices explained in section 3.1.

IV. Empirical Evidence

In order to test for a significant relationship between the correlation of regional

[8]This test was augmented by Levin and Lin (1993) and critically surveyed by Higgins and Zakrajsek (1999).

[9]The unit root test results displayed above refer to those variables whose order of magnitude might be most disputable according to the anonymous referee. We would like to argue that one argument for taking first differences of the mentioned variables might pertain to the choice of the adequate estimation technique.

Table 3. Pattern of panel ADF-test statistics for labour-market variables

Variable	t-value (no lagged differences)	t-value (one lagged difference)	t-value (two lagged differences)
	I) ADF-test statistic (no constant, no trend)		
KOR	-38.18***	-41.25***	-34.47***
RELINC	-59.04***	-12.46***	-9.93***
D(KOR)	-72.16***	-63.61***	-51.20***
D(RELINC)	-37.01***	-32.83***	-25.30***
	II) ADF-test statistic (common constant, no trend)		
KOR	-27.97***	-30.57***	-23.84***
RELINC	-57.93***	-53.22***	-40.41***
D(KOR)	-72.18***	-63.70***	-51.11***
D(RELINC)	-36.94***	-32.52***	-25.26***
	III) ADF-test statistic (common constant and trend)		
KOR	-27.75***	-30.47***	-23.70***
RELINC	-58.48***	-53.03***	-40.22***
D(KOR)	-51.14***	-63.95***	-50.84***
D(RELINC)	-37.14***	-32.62***	-25.60***
	IV) ADF-test statistic (individual-specific constant and trend)		
KOR	-34.05***	-38.83***	-32.20***
RELINC	-80.24***	-33.09***	-30.99***
D(KOR)	-70.15***	-63.00***	-51.98***
D(RELINC)	-34.48***	-24.57***	-23.75***

Note: the t-value is the realization of the usual ADF-test statistic; */**/*** denotes significance of the lagged endogenous variable; the sample range is 1975-1997 with adjustments if necessary due to the lag structure and data availability. The number of used cross-sections is 435 (465 minus all region-pairs with region G8 (Berlin) involved for which not in all cases data were available).

Ad I) test equations correspond to model 1 in Levin and Lin (1992), p. 12. The relevant critical values are -1.31/-1.67/-2.34 (Levin and Lin (1992), Table 1, p. 45 (for N=300 cross-sections and t=25 periods).

Ad II) test equations correspond to model 2 in Levin and Lin (1992), p. 16. The relevant critical values are -1.35-1.71/-2.39 (Levin and Lin (1992), Table 2, p. 46 (for N=300 cross-sections and t=25 periods).

Ad III) test equations correspond to model 3 in Levin and Lin (1992), p. 19. The relevant critical values are -1.39/-1.75/-2.43 (Levin and Lin (1992), Table 3, p. 47 (for N=300 cross-sections and t=25 periods).

Ad IV) test equations correspond to model 5 in Levin and Lin (1992), p. 25. The relevant critical values are -23.04/-23.35/-23.89 (Levin and Lin (1992), Table 5, p. 49 (for N=300 cross-sections and t=25 periods).

EU employment cycles and the specialisation indices constructed for region-pairs, we run dynamic panel regressions of the correlation coefficient KOR between two regional employment cycles on its own past level, one of the measures of the degree of similarity, and a variable capturing relative differences with respect to

[10]If necessary, a dummy for German unification was included as well. We originally also experimented with dummies for common borders. However, these dummies did not enter the final specifications due to lack of significance and/or near singular matrix problems. In an earlier version of this paper, we also run re-gressions under the relatively rigid statistical assumption that all region-pairs in the pool react in the same manner to changes in the concentration measure. However, the results remained similar.

Table 4. Hausman tests for fixed versus random effects (1989-1996)

Measure of specialisation used	Test statistics	Empirical realisation
Index of conformity (CON)	chi-sqr(4) =	1085.459
	p-value =	0.00
Finger-Kreinin index (FIN)	chi-sqr(4) =	1066.93
	p-value =	0.00
Coefficient of specialisation (SPEC)	chi-sqr(4) =	1066.11
	p-value =	0.00

Numbers in brackets refer to the degrees of freedom of the Chi-square statistics. Estimations are in each case based on a least squares estimation of an equation which specifies the correlation coefficient KOR as a function of its own first and second lag, contemporaneous relative income and one of the three different measures of specialisation lagged one period.

regional income[10].

A. Empirical model and estimation procedure

Based on our theoretical arguments we conjecture that a decreasing degree of regional similarity tends to reduce the correlation between EU regional employment cycles in a cross-country panel analysis, after controlling for additional variables. To test for a significant negative impact of an increasing degree of specialisation on the correlation between regional employment cycles in the EU, we undertake fixed effects estimations, assuming region-pair specific intercepts and slope coefficients which are the same across units. Based on the general-to-specific approach, the final specification of the underlying regression equations is reached according to the usual diagnostics (see, e.g., Wooldridge, 2003).

We exclusively estimate fixed effects models[11]. However, in the literature random effects models are sometimes implemented instead of fixed effects models, mainly because the FE country-dummies are costly in terms of lost degrees of freedom. We decided to dispense with a random effects estimation because it would only be appropriate, if we really believed that our sampled cross-sectional units were drawn from a large common population as is not the case here. The key distinction between fixed and random effects models is whether the individual effects should be treated as correlated with the x's or not. Correlation would favour the estimation of a fixed effects model. No correlation would support a random effects estimation. However, in our case there is practically speaking little reason to

[11]The time dimension of the data available for the EU regions is too small to get reliable estimates from country-specific time-series regressions.

treat the individual effects as uncorrelated with the other regressors as assumed in the random effects model. Moreover, in our context it is rather clear that fixed effects are likely to be more rigorous as will be discussed later (Greene, 2003, pp. 293ff., and Hsiao, 2002, pp. 149ff.).

However, for reasons of methodological correctness, we also performed Hausman specification tests to check empirically whether fixed effects is the correct estimation procedure (against the possible alternative of random effects). For this purpose, we compute the Hausman test statistic for testing the null hypothesis of random effects against the alternative hypothesis of fixed effects (Hausman, 1978). Already with an eye on the final regression equations we are concerned with three specifications according to which the correlation coefficient KOR is a function of its own first and second lag, contemporaneous relative income and one of the three different measures of the bilateral similarity lagged one period. The results which reveal overwhelming evidence in favour of a fixed effects estimation are displayed in Table 4. The null hypothesis that random effects is the better option (i.e., unbiased) is rejected according to the chi-squared tables at the usual significance levels throughout our specifications[12]. Hence, we feel justified to dispense with random effects estimations and to focus on fixed effects models in the remaining part of the paper.

The empirical model we start with is the most common one and can be described as follows:

$$y_{it} = \alpha_i + x_{it}^{'}\beta + \varepsilon_{it} \tag{2}$$

with y_{it} as the dependent variable (synchronicity), x_{it} and β as k-vectors of non-constant regressors (e.g., one of the specialisation indices CON or FIN or SPEC, the relative income RELINC and a dummy for German unification (DUMGER) and parameters for $i = 1,2, \ldots, N$ cross-sectional units and $t = 1,2, \ldots, T$ as the periods for which each cross-section is observed. The parameter β has no subscript because our estimation procedure assumes that it is common to all units.

The corresponding setting with respect to a representative dynamic regression equation for one cross-section out of the whole system (described by the index i) is the following:

$$y_{it} = \alpha_i + x_{it}^{'} + \delta y_{i,t-1} + \varepsilon_{it} \tag{3}$$

[12]We also experimented with estimations based on White standard errors and covariances and got strikingly similar results.

However, when estimating our first-order model substantial complications might potentially have to be taken into account. This is due to the heterogeneity of the cross-sections analysed (Greene, 2000, pp. 582 ff.). The main problem to be treated in this context is the correlation of the lagged dependent variable (the correlation coefficient KOR lagged one or more periods) with the disturbance, even if the latter does not exhibit autocorrelation itself. However, the Nickell bias, i.e. the one found in the fixed effects model, is much less of a problem when the time dimension is large (Nickell, 1981). In fact we argue that the time dimension in our case is large enough to estimate by within groups, i.e. use the standard fixed effects estimator. We do not worry about instrumenting the lagged dependent variable, because in the present case the use of lags as instruments would look doubtful because measurement error and possibly the moving average nature of the variables tend to induce serial correlation in the error term. Hence, the standard FE estimator is not only simpler but also more rigorous in our case. Hence, we rely on the standard least squares FE estimates where all observations are given equal weight throughout our estimations. As an additional robustness check, we also use the Feasible Generalized Least Squares (FGLS) estimates of the empirical FE model assuming the presence of cross-sectional heteroscedasticity and autocorrelation but without correction for contemporaneous correlation. This neglect is no drawback since correlations across countries might become relevant only in the case of symmetric shocks to the regional labour markets and the probability of the latter is typically small in our large EU sample[13]. As usual, we call this kind of estimation procedure the cross-section weights case. It implies that each pool will have an unrestricted intercept and that each pool equation is down-weighted by an estimate of the cross-section residual standard deviation.

B. Results

The original sample has been chosen to range from the year of the fall of the iron curtain 1989 to the year 1997. The decision in favour of this sample which fits the data best was taken with an eye on robustness with respect to German unification (with economic anticipation effects felt already in 1989) and on the impact of this choice on the sample available for the so-called Wooldridge-test of serial correlation of residuals in dynamic panels which we use in our estimations. Due to the fact that this test implies a regression equation which contains lagged

[13]See the debate on optimum currency areas, e.g., Babetski et al. (2002).

Table 5. Impact of the degree of specialisation on the correlation between regional employment cycles in the EU

	(1)	(2)	(3)	(4)	(5)	(6)
Estimation method:						
Fixed effects + no weighting	X		X		X	
Fixed effects + cross-section weights		X		X		X
Regressors:						
Correlation coefficient KOR lagged one period	0.55***	0.52***	0.56***	0.53***	0.56***	0.53***
Correlation coefficient KOR lagged two periods	-0.30***	-0.23***	-0.30***	-0.23***	-0.30***	-0.23***
Coefficient of specialisation:						
CON	4.63***(-1)	4.37*** (-1)	/	/	/	/
FIN	/	/	2.25***(-1)	2.03***(-1)	/	/
SPEC	/	/	/	/	-1.10***(-1)	-1.00***(-1)
Relative income (RELINC)	-9.84E-05***	-0.0001***	-9.64E-05***	-0.0001***	-9.63E-05***	-0.0001***
(Weighted) statistics:						
R^2	0.66	0.97	0.66	0.97	0.66	0.97
F-statistics	11.05	188.97	10.99	196.94	10.99	197.32
Wooldridge-$\hat{t}$ (p-value)	-0.48 (0.63)	-0.12 (0.90)	-1.12 (0.26)	-0.70 (0.49)	-1.12 (0.26)	-0.67 (0.50)

Each estimation uses 2956 total panel observations and 435 cross-sections. The sample in each case is 1989-1996. Numbers in brackets indicate the lag of the implemented regressor. If not otherwise indicated the variable is inserted contemporaneously. / means that the variable is not included. ***, **, *indicate significance at the 1, 5 and 10 percent levels respectively.

residuals, the sample for the Wooldridge-test is diminished by one year and now comprises exactly the years in the wake of German economic and monetary union. Due to the fact that the indices used by us are available only up to 1996 the last year of the sample actually used is 1996. All of our representative final regression equation specifications which are displayed in Table 5 include two endogenous lagged correlation coefficient KOR, the respective specialisation index (the index of conformity CON, the Finger-Kreinin index FIN or the coefficient of specialisation SPEC), and the variables relative income RELINC and, if significant, also the dummy for German unification DUMGER. With an eye on the annual frequency of the data, we limit possible lags to a number from 0 to 2 (annual data) and then test down until we arrive at the best fitting specification. The number of lags of the relevant variables are determined by the usual goodness-of-fit criteria for panels. By this procedure, we try to ensure that the dynamics of the model are sufficiently rich that serial correlation of the residuals is eliminated.

The fit of each equation is checked by referring to the R-squared, the F-statistics and an AR(1) time series test for the autocorrelation of residuals[14]. The latter test for autocorrelation of the residuals of order one is highly recommended by Wooldridge (2002) for dynamic panels, even if the latter include lagged endogenous variables. It regresses the dependent variable on the independent variable and the lag of residuals from the original equation. Based on this regression for 1990 to 1996, a standard t-test on the significance of the coefficient of the lagged pool residuals is performed. Under the null hypothesis, this coefficient is zero and, thus, there is no autocorrelation of the residuals[15]. The empirical realization of the t-statistics is displayed in Table 5 jointly with the corresponding p-value (denoted as $\hat{t}$). Since the marginal significance level of the F-test of joint significance of all of the slope coefficients is in all cases clearly below one percent, the p-value is not explicitly tabulated by us. However, the degrees of freedom can be easily read off from the table[16]. We expect positive signs of the estimated coefficients of the index of conformity CON and the Finger-Kreinin index FIN and a negative sign of the

[14]See Wooldridge (2002), p. 176f. Serial correlation should not be present in a model which is supposed to be dynamically complete in the conditional mean.

[15]A nice feature of the statistics computed is that it works whether or not the regressors are strictly exogenous.

[16]The numerator degrees of freedom can be calculated as the number of explanatory variables less one and the denominator degrees of freedom corresponds to the numbers of observations minus the number of re-gressors.

coefficient estimate of the coefficient of specialisation SPEC. We now turn to a final discussion of our representative estimations in Table 5.

Table 5 contains a rather strong result. Independent of the specific measure of the bilateral degree of similarity (the index of conformity CON, the Finger-Kreinin index FIN or the coefficient of specialisation SPEC), the estimated coefficients measuring the impact of the degree of similarity on the correlation between regional employment cycles in the EU are always significant at the one percent level. Moreover, the coefficient estimates always display the expected sign. The economic impact of the relative income on the degree of correlation between regional employment cycles originally claimed by Imbs (1999) is small but economically non-negligible. If cross-section weights are used, the estimated coefficients are slightly smaller but nevertheless highly significant with the expected sign. The dummy for German unification is significant in only a small number of cases. The available test statistics in the lower part of both tables point towards correct specifications of the final regression equations. Most important, according to the Wooldridge test statistics the dynamics of the final models are sufficiently rich that serial correlation of the residuals is absent. The lags of the dependent and/or explanatory variables appear to be long enough in this respect. The broad result that emerges from Table 5 is that the exact specification of the specialisation index does not really affect the role of the regional specialisation pattern for business cycle synchronisation in general. Complementary investigations show that the use of co-variances which are robust to general heteroscedasticity do not change our results. Also from this point of view, the results displayed in Table 5 are quite representative.

Finally, we would like to stress a common feature of all estimation results, namely the fact that our specialisation indices enter most of the regression equations not contemporaneously but with a lag. This clear pattern underlines the validity of our prior that the measures of similarity are exogenous with respect to the correlation coefficient. However, a nice feature of the Wooldridge statistics computed to test for the absence of residual autocorrelation is anyway that it works whether or not the regressors are strictly exogenous.

V. Summary and Outlook

This paper examines the degree of correlation in EU regional employment cycles and attempts to show whether these reflect changing patterns of specialisation. The approach taken in this paper is closest in spirit to Clark and Wincoop (1999) although it looks at regional rather than national indicators of similarity. With its focus on the regional level and by using three different indicators of similarity of sectoral structure, it goes beyond existing studies. The empirical method adopted is to estimate a dynamic panel data model for region pairs by within groups, i.e. a standard fixed effects estimator. Special attention is paid to ensure that the dynamics of the model are sufficiently rich that serial correlation is eliminated. The key finding is that employment growth is more synchronised in a statistical and econometric way when regions look alike in their sectoral structure. This finding is robust in two dimensions. First, in contrast to most of the empirical literature, we used three different indicators to approximate the regional sectoral structure. Since the results are not sensitive to the choice of the indicator, the relationship between the sectoral structure and the co-movement of employment cycles seems to be stable. Second, our key finding closely corresponds with the results from Clark and Wincoop (2001) with the only exception, that they do not focus on and quantify the regional specialisation patterns. Instead, they used just a national index, which implicitly assumes, that regions are homogenous entities.

The section on stylised facts mentions that the degree of synchronicity of regional employment cycles has declined in the past for a majority of region pairs. The evidence in favour of a decreasing synchronicity of regional employment cycles is implicitly backed by other studies with a regional focus (Fatás, 1997, Clark and Wincoop, 2001). This result stands in contrast to many studies with a national focus (e.g., Barrios et al., 2001) which indicate an increasingly closer correlation of national business cycles between countries which are now members of EMU (e.g., Artis and Zhang, 1997 and Christodoulakis et al., 1995). In section 2 we called this a puzzle which is likely to be of fundamental importance in driving the results and is also important for the context that we view this contribution. According to our estimation results, the trend decline of synchronicity of business cycles at the regional level could be explained mainly through changes in the sectoral structure, given the small impact of relative incomes. However, this paper can only potentially solve part of the puzzle, because it cannot explain why synchronicity between EU national cycles has grown at the same time. However, well-founded speculations in this direction could for instance start from the insight that the industrial structure of the EU countries as a whole represents an average of its different regional sectoral structures. Hence, industrial structures do not differ so much from each other on a national level as they do on a regional level by construction.

However, one should be careful and not draw premature conclusions from the quite consistent empirical results gained in this paper. Above all, one should not open discussions which risk to imply that low synchronisation of regional cycles is a problem when it could be optimal. According to the New Economic Geography literature, low synchronisation might instead just be an expression of agglomeration tendencies on a regional level which take place according to an optimising calculus. Centripetal or centrifugal forces which determine the degree of agglomeration are determined by the degree of integration or, to be even more concrete, by the magnitude of transportation costs.

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References

Amiti, M., 1997, Specialisation Patterns in Europe, Weltwirtschaftliches Archiv 135, 573-593.

Artis, M. J. and W. Zhang, 1997, International Business Cycles and the ERM: Is there a European Business Cycle?, International Journal of Finance and Economics 2, 1-16.

Babetski, J., L. Boone and M. Maurel, 2002, Exchange Rate Regimes and Supply Shocks Asymmetry: The Case of the Accession Countries, CEPR Discussion Paper No. 3408, Centre for Economic Policy Research, London.

Barrios, S., M. Brülhart, R. Elliott and M. Sensier, 2001, A Tale of Two Cycles : Co-fluctuations between UK Regions and the Euro Zone, University of Nottingham research paper series, Nr. 02/2001.

Baxter, M. and R. G. King, 1999, Measuring Business Cycles: Approximate Band-Pass Filters for Economic Time Series, Review of Economics and Statistics 81, 575-593.

Belke, A. and J. M. Heine, 2001, On the Endogeneity of an Exogenous OCA-Criterion: The Impact of Specialisation on the Synchronisation of Regional Business Cycles, HWWA Discussion Paper No. 119, HWWA-Institut für Wirtschaftsforschung Hamburg.

Brülhart, M., 2002, Labour-Market Effects of Intra-Industry Trade: Evidence for the United Kingdom, Weltwirtschaftliches Archiv 138, 207-228.

Caporale, G. M., N. Pittis and K. Prodromidis, 1999, Is Europe an Optimum Currency Area? Business Cycles in Europe, Journal of Economic Integration 14, 169-202.

Christodoulakis, N., S. P. Dimelis and T. Kollintzas, 1995, Comparisons of Business Cycles in the EC: Idiosyncrasies and Regularities, Economica 62, 1-27.

Clark, T. E. and K. Shin, 2000, The Sources of Fluctuations within and across Countries, in: Hess, G. and E. van Wincoop, eds., Intranational Macroeconomics, Cambridge University Press, 189-217.

Clark, T. E. and E. van Wincoop, 2001, Borders and Business Cycles, Journal of International Economics 55, 59-85.

Cleveland, W. S., 1993, Visualizing Data, Hobart Press.

European Central Bank, 1999, Monthly Bulletin, July, Frankfurt, 35-58.

Fatás, A., 1997, Countries or Regions? Lessons from the EMS Experience, European Economic Review 41, 743-751.

Finger, J. M. and M. E. Kreinin, 1979, A Measure of 'Export Similarity' and its Possible Uses, Economic Journal 89, 905-912.

Forni, M. and L. Reichlin, 1997, National Policies and Local Economies: Europe and the United States, CEPR Discussion Paper No. 1632, Centre for Economic Policy Research, London.

Frankel, J. A. and A. K. Rose, 1998, The Endogeneity of the Optimum Currency Area Criteria, The Economic Journal 108, 1009-1025.

Greene W.H., 2000, Econometric Analysis, 4th ed., New Jersey.

Greene, W.H., 2003, Econometric Analysis, 5th ed., New Jersey.

Greenaway, D. and R. C. Hine, 1991, Intra-Industry Specialization, Trade Expansion and Adjustment in the European Economic Space, Journal of Common Market Studies 29, 603-622.

Hausman, J., 1978, Specification Tests in Econometrics, Econometrica 46, 1251-1271.

Higgins, M. and E. Zakrajsek, 1999, Purchasing Power Parity: Three Stakes Through the Heart of the Unit Root Null, Federal Reserve Bank of New York, mimeo.

Hodrick, R. J. and E. C. Prescott, 1997, Postwar U.S. Business Cycles: An Empirical Investigation, Jounral of Money, Credit and Banking 29, 1-16.

Hsiao, C., 2002, Analysis of Panel Data, Cambridge.

Imbs, J., 1999, Co-Fluctuations, CEPR Discussion Paper No. 2267, Centre for Economic Policy Research, London.

Inklaar, R. and J. DeHaan, 2001, Is There Really a European Business Cycle?, Oxford Economic Papers 53, 215-220.

Kim, S., 1999, Expansion of Markets and Geographic Distribution of Economic Activities: The Trends in US Regional Manufacturing Structure, 1860-1987, Quarterly Journal of Economics 110, 881-908.

Krieger-Boden, Ch., 1999, Nationale und regionale Spezialisierungsmuster im europäischen Integrationsprozess, Die Weltwirtschaft 2, 234-254.

Krugman, P., 1991, Increasing Returns and Economic Geography, Journal of Political Economy 99, 483-499.

Krugman, P., 1998, What's New about the New Economic Geography, Oxford Review of Economic Policy 14, 7-17.
Levin, A. and C.-F. Lin, 1992, Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties, University of California, San Diego, Department of Economics Discussion Paper 92-23, May.
Midelfart-Knarvik, K. H., H. G. Overman, S. J. Redding, and A. J. Venables, 2000, The Location of European Industry, CEPR Discussion Paper Nr. 2618, Centre for Economic Policy Research, London.
Molle, W., 1997, The Regional Economic Structure of the European Union: An Analysis of Long-Run Developments, in: Peschel, K., ed., Regional Growth and Regional Policy within the Framework of European Integration, Heidelberg, 66-86.
Nickell, S., 1981, Biases in Dynamic models with Fixed Effects, Econometrica 49, 1417-1426.
OECD, 1999, EMU: Facts, Challenges and Policies, Paris.
Puga, D., 2002, European Regional Policies in Light of Recent Location Theories, Journal of Economic Geography 2, 373-406.
Quantitative Micro Software, 2000, EViews – Command and Programming Reference, Irvine, California.
WIFO, 1999, Specialisation and (Geographic) Concentration of European Manufacturing, Austrian Institute for Economic Research, Report of a Study for the Enterprise DG of the European Commission, Brussels.
Wooldridge, J. M., 2002, Econometric Analysis of Cross-Section and Panel Data, MIT Press, Cambridge, Mass.

Variable Annex

KOR: Bravais-Pearson correlation coefficient measuring the 5-year moving correlation between the residuals of HP-filtered employment of two regions.

CON: Index of conformity as defined in Table 2.

SPEC: Finger-Kreinin index as defined in Table 2.

FIN: Coefficient of specialisation as defined in Table 2.

RELINC: Relative income for a region-pair, calculated as the absolute value of thedifference between the logarithms of real GDP of these two regions. Nominal regional GDP was deflated by the consumer price index of thecountry the respective regions belong to.

DUMGER: Dummy for German unification coded as 1 for the period from 1990 on, otherwise 0.

Data sources and more detailed specifications of each variable are described in section 3.