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Economic Integration in a Multicone World?

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Economic Integration in a Multicone World[•]

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Abstract

This paper examines whether economic integration favors countries' convergence into a common cone of diversification. We analyze the manufacturing specialization patterns for a sample of 19 current and potential European Union countries over the period 1963-1998, and assess the impact of integration on their evolution. We perform year-by-year threshold estimations of Rybczynski relationships to identify the diversification cones and then estimate discrete choice models to investigate whether membership in the European Union is associated with a higher probability of being in a same cone. Economic integration in Europe is found to have promoted convergence from lower to higher diversification cones.

Keywords: Multiple Diversification Cones, Convergence, European Integration, Threshold Estimation, Dynamic Probit Models

JEL Classification: F11, F15, C2

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Economic Integration in a Multiple Cone World

1 Introduction

In the standard Heckscher-Ohlin trade model two global equilibria may result: a "single cone" of production in which all countries have the same mix of tradables and offer the same factor rewards; or "multiple cones" where countries specialize in a subset of industries most suited to their relative endowments and factor prices are not necessarily equalized. While much work has been done attempting to econometrically verify the single product mix equilibrium, empirical studies addressing the multiplicity of cones are scarce.

In this paper, we treat the European Union as a Hecksher-Ohlin integrated world economy and examine whether members converge from lower to higher cones of diversification. In particular, we analyze the manufacturing specialization patterns for 19 current and potential European Union countries over the period 1963-1998, assessing the impact of trade integration on their evolution. In doing this, we relate two strands of economic literature: empirical tests of the Heckscher-Ohlin trade model and economic integration and convergence.

To our knowledge, only Debaere and Demiroglu (2003), Schott (2003), and Xiang (2004) have recently examined the multiple-cone question empirically. Debaere and Demiroglu (2003) use the lens conditions derived by Deardorff (1994) that compares country endowments with sectoral factor inputs. Considering a sample of 28 developed and developing countries for 1990, they find that, even though the endowments of countries are too dissimilar for all countries to be able to produce the same set of goods, rich OECD countries lie in the same diversification cone. Following Leamer (1984, 1987), Schott (2003) estimates Rybczynski relationships for a cross-section of 45 countries for the same year via maximum likelihood. He reports that the null hypothesis of a single-cone equilibrium is rejected in favor of an alternative hypothesis of two cones after correcting for cross-country intra-industry heterogeneity. Furthermore, most developed economies seem to be within a common cone. These results are thus compatible to those in Debaere and Demiroglu (2003). Comparing countries' distribution of capital-labor ratios over industries for 10 rich OECD countries on the basis of the Kolgomorov-Smirnov test statistics, i.e., non-parametric estimation techniques, Xiang (2004) uncovers, in turn, that there are non-uniform differences in the distribution functions of factor usage among these countries. In contrast to the two studies above, he identifies three diversification cones within the set of developed economies and shows that the United Kingdom, France, and the United States are in different groups.

This paper first differs from the previous investigations in that we use a different empirical methodology. In order to identify the number of cones of diversification, we perform year-by-year threshold estimations of Rybczynski relationships as described in Hansen (2000). Second, instead of considering a unique cross-sectional sample, we look at the question of multiple cones of diversification in a multi-period framework, which allows us to trace both the evolution of the number of cones and the country composition within each cone over time.

Having identified the path of the cross-sectional distribution of economies across different cones, we then go one step further in linking these developments to the process of economic integration. Venables (2003) has shown that trade agreements between high income countries are likely to cause a convergence of income and production structures. This theoretical result finds empirical support in Ben-David (1993, 1996). We contribute to this branch of the literature, providing additional evidence. We estimate discrete choice models to investigate whether membership in the European Union favors convergence of production patterns, i.e., whether this membership is associated with a higher probability of being in a same cone of diversification. Economic integration in Europe is found to have generally promoted convergence from lower to higher diversification cones. The exceptions seem to be Greece and Portugal.

Determining the existence of different cones is important for several reasons. Within a cone, definite predictions are generated about the factor content of trade and these have been tested (Bowen et al., 1987; Trefler, 1995; and Davis and Weinstein, 2001). The equalization of factor prices across countries and across regions has also been tested (Repetto and Ventura, 1997; Bernard and Schott, 2002; and Hanson and Slaughter, 2002). Integration of the world economy provides a force for income concentration such that the Stolper-Samuelson effect holds due to increased trade. Tests of this variety will be affected by the existence of multiple cones of product diversification and therefore whether countries lie in the same cone has significant implications for the link between trade and income inequality.

Within a cone, factor prices equalize and there is a lack of incentive to exploit cross-country factor price differences via outsourcing. Thus, the existence of multiple as opposed to the traditional single cone of diversification affects the international fragmentation of production (Feenstra and Hanson, 1997; Deardorff, 1998). Whether countries belong to the same cone is also important for the relationship between trade and growth (Deardorff, 2001) and economic development (Ventura, 1997).

In addition, establishing whether economic integration is a force for convergence or divergence is of interest because the way benefits and costs of trade arrangements are distributed between member countries plays a non-minor role in explaining their success or collapse (e.g., Venables, 1999). In the particular case of the European Union, promoting cohesion is a central concern explicitly expressed in the foundational Treaty of Rome (1957). Further, if as a result of declining trade barriers, integrating countries become more similar (dissimilar) in terms of their production structures, and thus less (more) sensitive to industry specific shocks, less (more) idiosyncratic business-cycles should be expected (e.g., Frankel and Rose, 1998). As it is well known, this is a key criterion in delimiting an optimum currency area and thus of fundamental relevance in assessing the current European Monetary Union as well as its future extensions.

The remainder of the paper is organized as follows: Section 2 presents the theoretical framework. Section 3 briefly describes our dataset, while Section 4 explains the econometric methodology. Section 5 discusses the estimation results and Section 6 concludes.

2 Theory

We begin with the multicone-version of the Heckscher-Ohlin model. In this world, each cone is a combination of input vectors. In the Lerner (1952) description of an open economy's accumulation of capital relative to fixed supplies of labor, there are two factors (capital and labor) and four industries. The four industries, I_1, I_2, I_3, and I_4, have unit-value isoquants which define three cones of diversification. Output levels are a linear function of the factor supplies.[1] The Lerner diagram, shown in Figure 1, illustrates that as capital is accumulated relative to labor, output in an industry evolves such that countries move into cones with progressively lower capital rental rates and higher wages. A GDP-maximizing country will produce two goods in the cone in which it resides. The capital-labor ratios mark the borders between the cones and are labeled γ_i where $i=[0,3]$. Thus, within a cone, countries produce a similar set of goods.

Countries abundant in capital produce the two capital-intensive goods while countries abundant in labor produce the two labor-intensive goods. As capital accumulates relative to labor, there is an evolution of output in industry i, in country c, per total workforce ($q_{ci}=Q_{ci}/L_c$) and the industry development path (à la Leamer, 1987) can

[1] Without a loss of generality, in Figure 1, each sector is displayed as having Leontief technology and factor intensity reversals are ruled out.

be mapped. Hence, as countries accumulate capital, they move to higher cones and the mix of commodities they produce shifts to be increasingly capital-intensive.[2]

The second issue we address in this paper is how this picture would change during a process of economic integration. We examine whether integration favors convergence of production structures and *per capita* income levels. In particular, we investigate whether membership in an economic agreement such as the European Union promotes convergence into a common cone of diversification.

Let us assume that comparative advantage is associated with *per capita* income via physical capital endowments, so that higher capital-income ratios are linked to higher *per capita* incomes. Now consider two rich economies, i.e., both with capital-labor ratios above the world average. Venables (2003) shows that membership in an economic union comprised of these two economies will lead to convergence. The country with an "intermediate" comparative advantage will do better from the union than the one with "extreme" comparative advantage because the former is likely to benefit from trade creation, while the latter is likely to experience trade diversion. The European Union would therefore promote convergence along the development path, i.e., as countries accumulate capital relative to labor, in such a way as to end up being more similar. We now turn to the empirical analysis.

3 Data

We consider a sample of 19 European countries: Austria, Belgium, Cyprus, Denmark, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Netherlands, Norway, Poland, Portugal, Spain, Sweden, Turkey, and the United Kingdom for which we have all the required data over a long time span. Some of these countries are founding members of the European Union, other countries became members during the 1970s, 1980s, and

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[2] We choose to follow the Leamer (1987) and Schott (2003) line of reasoning.

1990s and finally, the remaining ones only entered the union very recently or are still not members. Table 1 shows the evolution of EU country composition over the last four decades. The successive enlargements and thus the time varying membership provide us with a natural experiment to identify the impact of economic integration on specialization patterns.

In particular, we use value-added data for these countries across 28, 3-digit ISIC Rev. 2 industries over the period 1963-1998 from the United Nations Industrial Development Organization (UNIDO) database. As discussed in Schott (2003), focusing on manufacturing precludes testing whether endowment disparities are associated with specialization across broad economic sectors (agriculture, mining, manufacturing, and services), but has the advantages of including fewer non-tradables than other sectors and thus being closer to the theoretical assumptions.

We also use economy wide labor data, and arable, forest and woodland data from the Food and Agriculture Organization (FAO) as well as data on the skill level of working age population (i.e., population with at least complete secondary school) from the database prepared by Barro and Lee (2000) over the period 1960-2000. Finally, we apply the perpetual inventory method as indicated in Jacob et al. (1997) and Kamps (2004) with a depreciation rate of 13.3% (e.g., Schott, 2003) on gross fixed capital formation over the period 1970-2000 reported by the United Nation Commission for Trade and Development (UNCTAD).[3]

4 Econometric Methodology

The econometric methodology consists of a two step procedure. First, to determine the number of cones and the distribution of countries across them, we take a route that is

[3] In the presence of technological differences across countries, factor endowments should be adjusted by productivity (see Debaere and Demiroglu, 2003). We are currently working on deriving these adjusted endowments to check the robustness of our results. Similarly, we are aware of the problems that the intra-industry heterogeneity originating in the ISIC aggregation could create (see Leamer, 1987, and Schott, 2003).

different from those pursued in previous empirical investigations. We assess the existence of multiple diversification cones using a threshold regression model as described in Hansen (2000). Since we have annual data for the period 1963-1998, we can establish whether the number of cones has changed over time as well as to identify how countries re-group across these cones. We then investigate whether the probability of being in a same cone is affected by common membership in the European Union using binary choice models, including an accounting of dynamics in a probit model for panel data with unobserved heterogeneity, as explained in Wooldridge (2005).

We start by estimating the Rybczynski relationships for the cross-section of 19 countries for which we have the required data for each of the sample years. Formally:

$$q_{cit} = \theta_{1it} + \theta'_{2it} k_{cit-1} + \theta'_{3t} z_{ct-1} + \varepsilon_{cit} \tag{1}$$

where $q_{cit}=Q_{cit}/L_{ct}$ denotes output per worker in country *c*, in industry *i*, in year *t*; $k_{cit-1}=d_i k_{ct-1}$ with d_i being a row vector of industry dummy variables and $k_{ct-1}=K_{ct-1}/L_{ct-1}$, country *c 's* capital-labor ratio in year *t-1;* and z is vector of country endowments per worker included to control for the presence of production factors other than labor and capital: arable land, forest and woodland, and skill per worker.[4]

As discussed, economies shift out of labor-intensive manufacturing into capital-intensive manufacturing at higher levels of capital abundance. Hence, in a multicone world the derivative of output, with respect to endowments, changes along the development path as countries accumulate capital relative to labor. The regression coefficients therefore change across cones (see Schott, 2003). To assess whether this threshold effect is present in our sample, we employ a heteroskedasticity-consistent Lagrange Multiplier (LM) test for a threshold as described in Hansen (1996). Since the threshold is not identified under the null hypothesis of no threshold effect, the *p*-values are computed by a bootstrap analog, fixing the regressors from the right-hand side of Equation (1) and generating the bootstrapped dependent variable from the distribution

[4] We use one period-lagged explanatory variables to avoid simultaneity problems.

$N\left(0,\varepsilon_i^2\right)$, where $\hat{\varepsilon}_i$ is the OLS residual from the estimated threshold model. Let us rewrite Equation (1) in the following way:

$$q_{cit} = \theta_t' x_{ct-1} + \varepsilon_{cit} \tag{2}$$

where x is a vector including all endowments per worker. If we find evidence in favor of two diversification cones, the correct specification is the following:

$$q_{cit} = \theta_t' x_{ct-1} + \delta_{tn}' x_{ct-1}(\gamma) + \varepsilon_{cit} \tag{3}$$

where $n=CxI=532$ (number of countries times number of industries), $x_{ct-1}(\gamma) = x_{ct-1} d_{ct-1}(\gamma)$ with the indicator function $d_{ct-1}(\gamma) = \{k_{ct-1} \le \gamma\}$, so that the continuously distributed capital-labor ratio $k=K/L$ acts as a threshold variable, i.e., a variable that is used to split the sample into two groups. Let $x(\gamma) = x_\gamma$. In matrix notation, we therefore have:

$$Q_t = X\theta + X_\gamma \delta + \varepsilon \tag{4}$$

where we have suppressed the time index t to simplify the notation. To estimate our parameter of interest, namely γ, we follow the method proposed by Hansen (2000). This method consists of obtaining least squares estimates of the regression parameters $(\theta, \delta_n, \gamma)$ through concentration. Formally, let

$$S(\theta,\delta,\gamma) = \left(Q - X\theta - X_\gamma \delta\right)' \left(Q - X\theta - X_\gamma \delta\right) \tag{5}$$

be the sum of squared errors function. By definition, the least squares estimators $\left(\hat{\theta},\hat{\delta},\hat{\gamma}\right)$ jointly minimize Equation (5). For this minimization, γ is assumed to be restricted to a bounded set $\left[\underline{\gamma},\overline{\gamma}\right] = \Gamma$.[5] Conditional on γ, Equation (4) is linear in θ and δ. Thus, regressing Q on $X_\gamma^* = \left[X\ X_\gamma\right]$ we get the conditional OLS estimators $\hat{\theta}(\gamma)$ and $\hat{\delta}(\gamma)$. Let $S_n(\gamma)$ be the concentrated sum of squared errors function:

$$S_n(\gamma) = S_n\left(\hat{\theta}(\gamma),\hat{\delta}(\gamma),\gamma\right) = Q'Q - Q'X_\gamma^*\left(X_\gamma^{*\prime} X_\gamma^*\right)^{-1} X_\gamma^{*\prime} Q \tag{6}$$

[5] The least squares estimator is also the maximum likelihood estimator when ε_{ci} *is iid* $N(0,\sigma^2)$.

$\hat{\gamma}$ can then be obtained as the value that minimizes this function:

$$\hat{\gamma} = \underset{\gamma \in \Gamma_n}{argmin}\, S_n(\gamma) \tag{7}$$

where $\Gamma_n = \Gamma \cap \{k_{11},\ldots,k_n\}$ requires less than n function evaluations. Asymptotic confidence intervals are obtained by inverting the likelihood ratio statistics (see Hansen, 2000). The existence of additional breaks is evaluated by successively applying the same procedure on the sub-samples defined as indicated above. Once the thresholds have been identified, we group the countries into different diversification cones.

The second stage is to determine whether membership in the European Union favors convergence into a common cone of diversification. In particular, we test whether country pairs such that both are members of the European Union have a higher probability of being in a same cone than country pairs that are not.

Formally, we begin with a baseline specification given by the standard probit model:

$$P(sc_{it} = 1 \mid eu_{it}) = \Phi(\delta + \rho eu_{it}) \tag{8}$$

$$P(sc_{it} = 1 \mid eu_{it}) = \Phi(\eta_t + \rho eu_{it}) \tag{9}$$

where Φ is the standard normal cumulative distribution function, *sc* is a dummy variable that takes the value of 1 if both countries belong to the same diversification cone and 0 otherwise, *eu* is a dummy variable that takes the value of 1 if both countries are members of the European Union and 0 otherwise, and η_t are unrestricted years intercepts. Since there are no exact priors about the timing of the impact of EU membership, we also consider up to four lags for *eu*.

We then estimate a random effects probit model to account for unobserved heterogeneity:

$$P(sc_{it} = 1 \mid eu_{it}) = \Phi(\delta + \mu_i + \rho eu_{it}) \tag{10}$$

Finally, since the splitting up of countries displays inertia, we explicitly incorporate dynamics into the model following the approach proposed by Wooldridge

(2005). In particular, this approach consists of modeling the distribution of the unobserved effect conditional on the initial value and any exogenous explanatory variable and then performing conditional maximum likelihood estimation. We therefore estimate the following model:

$$P\left(sc_{it}=1 \mid sc_{it-1},..,sc_{i0},eu_{ii1},...,eu_{iT},\mu_i\right)=\Phi\left(\tau_t+\rho eu_{it}+\pi sc_{it-1}+\mu_i\right) \tag{11}$$

where t=1,…,T and 1 corresponds to 1964 and T corresponds to 1998. Since we have an unbalanced panel for the whole sample period and this methodology is basically appropriate for balanced panels, we replicate the exercise for the period for which we have for all country pairs in all years: 1967-1993.[6] The initial period, t=0, is 1963 (1966). The unobserved effect μ_i is assumed to satisfy $\mu_i \mid sc_{i0},eu_i \sim N\left(\alpha_0+\alpha_1 sc_{i0}+eu_i\alpha_2,\sigma_\upsilon^2\right)$ where $eu_i=\left[eu_{i1},...,eu_{iT}\right]$ is a row vector of indicators of EU membership in all time periods, which are included to allow for partial correlation between the unobserved effect and EU membership in all years.[7] The τ_t are time effects for the enlargement years: 1973, 1981, 1986, and 1995. We further consider the effect of lagged EU membership, in which case the model specification is as follows:

$$P\left(sc_{it}=1 \mid sc_{it-1},..,sc_{i0},eu_{i0},eu_{ii1},...,eu_{iT},\mu_i\right)=\Phi\left(\tau_t+\rho eu_{it-1}+\varphi sc_{it-1}+\mu_i\right) \tag{12}$$

Estimation results are presented in the next section.

5 Results

The first step of our empirical analysis consists of determining the number of diversification cones and of classifying the countries into the resulting groups. This is done using the sample splitting method proposed by Hansen (2000) as described above.

[6] This method assumes that the explanatory variables, in this case mainly *eu*, are strictly exogenous, which rules out feedback from unexpected movements in the outcome variable *sc* to future values of the explanatory variables. Since it might be argued that *eu* is endogenous we are working on the implementation of the GMM estimator proposed by Arellano and Carrasco (2003) to correct possible endogeneity biases.

[7] In particular, μ_i is specified as follows: $\mu_i=\alpha_0+\alpha_1 sc_{i0}+eu_i\alpha_2+\nu_i$, where ν_{ii} is independent of (sc_{i0},eu_i) and distributed as $Normal\left(0,\sigma_\nu^2\right)$. In this case, sc_{it} given $\left(sc_{it-1},...,sc_{i0},eu_i,\nu_i\right)$ follows a probit model with response probability: $\Phi\left(\rho eu_{it}+\pi sc_{it-1}+\alpha_0+\alpha_1 sc_{i0}+eu_i\alpha_2+\nu_i\right)$.

Table 2 reports annual threshold estimations over the period 1963-1998. According to these estimates, three diversification cones can be identified from the beginning of the 1960s to the end of the 1970s whereas there are essentially two cones in the last two decades.

Table 2 allocates the sample countries into the different cones according to their capital-labor ratios. During the 1960s and 1970s the lower cone included Cyprus, Greece, Hungary, Poland, Portugal, and Turkey. Austria, Finland, Italy, and the United Kingdom belong to the middle cone, while Spain and Ireland alternated between the lower and the middle cones. Finally, the upper cone consisted of Belgium, Denmark, France, Germany, Netherlands, Norway, and Sweden. In recent years, the middle cone vanishes. Specifically, we observe convergence of the countries in the middle cone to those in the upper cone. On the other hand, there is persistence in the lower cone, which is mainly integrated by two "latecomers" in the European Union, Greece and Portugal, and four countries that were not members by 1998: Cyprus, Hungary, Poland, and Turkey.

The next step is to assess whether economic integration in Europe favors convergence in production structures. In particular, we explore whether the probability of two countries being in a same cone is higher when both are members of the European Union. Operatively, we estimate discrete choice models where the dependent variable is an indicator of a common diversification cone. Specifically, from the information provided by the classification reported before we construct the dummy variable, *sc,* that takes a value of 1 if both countries belong to the same diversification cone and 0 otherwise and varies across all possible country pairs (19 x (18/2)=271) and over all sample years (1963 to 1998). The main independent variable, *eu*, is also an indicator of common membership in the European Union over time as shown in Table 1.

Table 4 presents probit estimates of Equations (8) and (9), i.e., with and without fixed time effects as well as comparable logit estimates.[8] The main results are robust across specifications and estimation methods: two countries that are both members of the European Union have a significantly higher probability of being in the same diversification cone. Since the impact of economic integration on specialization patterns can materialize with some lag and thus have varying intensities over time, we re-estimate Equation (8), this time allowing for up to a four year lag in the explanatory variable. The estimation periods are then successively redefined to ensure comparability. Estimation results are reported in Table 5. The estimated effect of membership in the European Union increases with the length of the lag, so that it seems to be increasing over time. Hence, the longer that countries share membership in this agreement, the more likely it is that both end up in the same diversification cone.

We next turn to estimations making use of the panel structure of our data. We first estimate a random effects probit model for panel data, i.e., Equation (10). This allows us to control for unobserved heterogeneity across country pairs. We consider both contemporaneous and lagged common membership as explanatory variables and unbalanced as well as balanced panels. Estimates are shown in Table 6. They confirm our original findings.

To account for the dynamic nature of the economic relationships under study and, in particular, for the fact that specialization patterns exhibit inertia, we estimate Equations (11) and (12).[9] The estimated effect of membership is again highly significant. As expected, the coefficient on the indicator of same cone is very statistically significant and seems to be factually large. The same is also true for the initial value of this variable.

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[8] The main difference between the normal and logit distribution are in the tails. Generally, the coefficients of the logit model are larger than those of the probit model. The reason is that the variance of a variable with a logistic distribution is $\pi^2/3$, while that of a variable with a standard normal distribution is 1. Amemiya (1981) suggest dividing the coefficients of the logit model by 1.6 to compare them with those of the probit one. In our case, estimated coefficients, using this division factor, are essentially the same.

[9] In an alternative specification, we also introduce a second lag of the dependent variable and accordingly modify the conditioning set as suggested in Wooldridge (2005). Estimations coincide with those reported here. Results of this alternative specification are not presented, but are available from the authors upon request.

This implies that there is substantial correlation between the unobserved heterogeneity and the initial condition (see Wooldridge, 2005). Finally, as suggested by the EU membership variables for the different sets of countries over time, the probability of being in a same cone is higher for those countries that have shared membership since 1973 and for those that have been members since 1995. As can be inferred from Tables 1 and 3, results for 1981 and 1986 are driven by Greece and Portugal.

6 Concluding Remarks

This paper aims at assessing whether economic integration favors convergence or divergence of countries specialization patterns. Examining the case of the Europe Union, we pursue a two-stage empirical methodology. In the first phase, we estimate Rybczynski relationships using the threshold model proposed by Hansen (2000) to identify the number of cones and how countries split across them on a year-by-year basis. Once countries are classified, we turn to the use of discrete choice models in the second phase in order to explore whether economic integration in Europe has favored convergence into a common cone of diversification. The main conclusion is that European integration has indeed lead to convergence from lower to higher cones. This confirms theoretical results in Venables (2003) suggesting that agreements between developed countries are a force for convergence.

If trade integration pushes towards convergence and all members eventually end up within the same diversification cone, there could be an implied effect on the decision to accept new members into the agreement. In particular, if the economic union desires to maintain production within several cones of diversification, when more recent members converge to pre-existing ones, an enlargement may be anticipated. Hence, this union would expand stepwise at discrete intervals as is in fact the case for the European Union.

Figure 1

Two-factor, Four industry Lerner Diagram

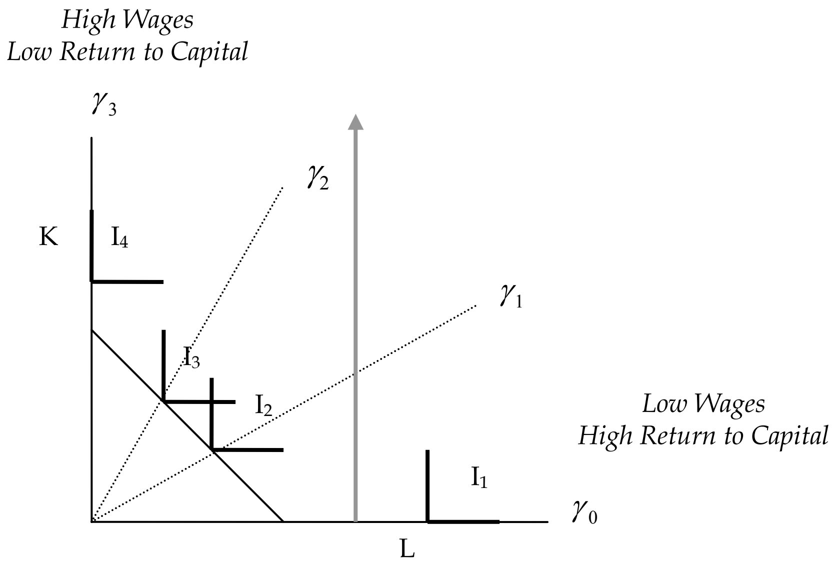

Table 1

Historical Widening of the European Union	
Year	**Members**
1958	Belgium, France, Germany, Italy, Luxembourg, Netherlands
1973	Denmark, Ireland, United Kingdom
1981	Greece
1986	Portugal, Spain
1995	Austria, Finland, Sweden

Source: Based on Baldwin (1994) and the European Union (2005)

Table 2

Threshold Estimation

Year		$\hat{\gamma}_1$	$\hat{\gamma}_2$	Year		$\hat{\gamma}_1$	$\hat{\gamma}_2$
1963	**K/L**	5.480	8.870	**1981**	**K/L**	11.200	
	LM-Test	101.275	83.101		**LM-Test**	106.026	
	B. P-Value	(0.000)	(0.007)		**B. P-Value**	(0.000)	
1964	**K/L**	5.430	8.880	**1982**	**K/L**	11.740	
	LM-Test	100.821	84.473		**LM-Test**	97.522	
	B. P-Value	(0.000)	(0.000)		**B. P-Value**	(0.000)	
1965	**K/L**	5.380	8.900	**1983**	**K/L**	12.060	28.220
	LM-Test	99.173	82.587		**LM-Test**	89.767	72.595
	B. P-Value	(0.000)	(0.008)		**B. P-Value**	(0.000)	(0.078)
1966	**K/L**	5.340	8.930	**1984**	**K/L**	12.160	31.280
	LM-Test	100.538	82.510		**LM-Test**	86.294	1214.930
	B. P-Value	(0.000)	(0.003)		**B. P-Value**	(0.004)	(0.010)
1967	**K/L**	5.300	8.960	**1985**	**K/L**	12.030	
	LM-Test	99.727	79.610		**LM-Test**	84.263	
	B. P-Value	(0.000)	(0.023)		**B. P-Value**	(0.005)	
1968	**K/L**	5.360	9.010	**1986**	**K/L**	12.160	
	LM-Test	91.591	96.640		**LM-Test**	85.908	
	B. P-Value	(0.000)	(0.000)		**B. P-Value**	(0.005)	
1969	**K/L**	5.370	9.060	**1987**	**K/L**	15.970	
	LM-Test	96.557	94.506		**LM-Test**	89.046	
	B. P-Value	(0.000)	(0.000)		**B. P-Value**	(0.001)	
1970	**K/L**	5.400	9.100	**1988**	**K/L**	13.760	
	LM-Test	107.496	96.590		**LM-Test**	86.854	
	B. P-Value	(0.000)	(0.000)		**B. P-Value**	(0.006)	
1971	**K/L**	5.390	9.140	**1989**	**K/L**	15.150	
	LM-Test	111.383	81.501		**LM-Test**	90.860	
	B. P-Value	(0.000)	(0.006)		**B. P-Value**	(0.004)	
1972	**K/L**	5.420	9.340	**1990**	**K/L**	23.040	
	LM-Test	102.284	77.117		**LM-Test**	100.178	
	B. P-Value	(0.000)	(0.041)		**B. P-Value**	(0.000)	
1973	**K/L**	5.700	9.770	**1991**	**K/L**	18.700	
	LM-Test	99.307	76.313		**LM-Test**	93.515	
	B. P-Value	(0.000)	(0.033)		**B. P-Value**	(0.000)	
1974	**K/L**	5.920	10.760	**1992**	**K/L**	20.160	
	LM-Test	104.653	73.343		**LM-Test**	89.410	
	B. P-Value	(0.000)	(0.078)		**B. P-Value**	(0.001)	
1975	**K/L**	7.020	12.340	**1993**	**K/L**	22.300	
	LM-Test	103.181	77.961		**LM-Test**	80.054	
	B. P-Value	(0.000)	(0.024)		**B. P-Value**	(0.032)	
1976	**K/L**	7.660	14.670	**1994**	**K/L**		
	LM-Test	103.663	77.624		**LM-Test**		
	B. P-Value	(0.000)	(0.041)		**B. P-Value**		
1977	**K/L**	8.280	17.240	**1995**	**K/L**		
	LM-Test	104.822	76.682		**LM-Test**		
	B. P-Value	(0.000)	(0.054)		**B. P-Value**		
1978	**K/L**	9.110	18.700	**1996**	**K/L**		
	LM-Test	117.508	78.735		**LM-Test**		
	B. P-Value	(0.000)	(0.025)		**B. P-Value**		
1979	**K/L**	10.700		**1997**	**K/L**		
	LM-Test	120.519			**LM-Test**		
	B. P-Value	(0.000)			**B. P-Value**		
1980	**K/L**	10.060		**1998**	**K/L**	41.620	
	LM-Test	108.053			**LM-Test**	83.686	
	B. P-Value	(0.000)			**B. P-Value**	(0.015)	

The above table reports the levels of capital-labor ratios at which a threshold effect is detected as well as the LM-test for no threshold with its associated p-bootstrapped value as calculated according to the method proposed by Hansen (2000).

Table 3

Diversification Cones in Europe (1963-1998)			
	Cone 1	**Cone 2**	**Cone 3**
1963-1966	Cyprus, Greece, Hungary, Ireland, Poland, Portugal, Spain, Turkey	Austria, Finland, United Kingdom	Belgium, Denmark, France, Germany, Netherlands, Norway, Sweden
1967	Cyprus, Greece, Hungary, Ireland, Poland, Portugal, Turkey	Austria, Finland, Italy, Spain, United Kingdom	Belgium, Denmark, France, Germany, Netherlands, Norway, Sweden
1968-1970	Cyprus, Greece, Hungary, Ireland, Poland, Portugal, Spain, Turkey	Austria, Finland, Italy, United Kingdom	Belgium, Denmark, France, Germany, Netherlands, Norway, Sweden
1971-1973	Cyprus, Greece, Hungary, Poland, Portugal, Spain, Turkey	Austria, Finland, Ireland, Italy, United Kingdom	Belgium, Denmark, France, Germany, Netherlands, Norway, Sweden
1974	Cyprus, Greece, Hungary, Poland, Portugal, Turkey	Austria, Finland, Ireland, Italy, Spain, United Kingdom	Belgium, Denmark, France, Germany, Netherlands, Norway, Sweden
1975-1976	Cyprus, Greece, Hungary, Ireland, Poland, Portugal, Turkey	Austria, Finland, Italy, Spain, United Kingdom	Belgium, Denmark, France, Germany, Netherlands, Norway, Sweden
1977-1978	Cyprus, Greece, Hungary, Ireland, Poland, Portugal, Turkey	Austria, Finland, Germany, Italy, Spain, United Kingdom	Belgium, Denmark, France, Netherlands, Norway, Sweden
1979		Cyprus, Greece, Hungary, Ireland, Poland, Portugal, Turkey	Austria, Belgium, Denmark, Finland, France, Germany, Italy, Netherlands, Norway, Spain, Sweden, United Kingdom
1980-1982		Cyprus, Greece, Hungary, Poland, Portugal, Turkey	Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Netherlands, Norway, Spain, Sweden, United Kingdom
1983	Cyprus, Greece, Hungary, Poland, Portugal, Turkey	Austria, Denmark, Finland, Ireland, Italy, Netherlands, Spain, Sweden, United Kingdom	Belgium, France, Germany, Norway
1984	Cyprus, Greece, Hungary, Poland, Portugal, Turkey	Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Netherlands, Spain, Sweden, United Kingdom	Norway
1985-1986		Cyprus, Greece, Hungary, Poland, Portugal, Turkey	Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Netherlands, Norway, Spain, Sweden, United Kingdom
1987		Cyprus, Greece, Hungary, Poland, Portugal, Spain, Turkey	Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Netherlands, Norway, Sweden, United Kingdom
1988-1989		Cyprus, Greece, Hungary, Poland, Portugal, Turkey	Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Netherlands, Norway, Spain, Sweden, United Kingdom
1990		Cyprus, Greece, Hungary, Poland, Portugal, Spain, Turkey	Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Netherlands, Norway, Sweden, United Kingdom
1991-1993		Cyprus, Greece, Hungary, Poland, Portugal, Turkey	Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Netherlands, Norway, Spain, Sweden, United Kingdom
1994-1997			Austria, Belgium, Cyprus, Denmark, Finland, France (*), Greece, Hungary, Ireland, Italy(**), Netherlands, Norway, Poland, Portugal, Spain, Sweden, Turkey, United Kingdom
1998		Cyprus, Greece, Hungary, Poland, Spain, Turkey	Austria, Denmark, Finland, Netherlands, Norway, Sweden, United Kingdom

The classification of countries in specialization cones is derived from threshold estimations performed using the method proposed by Hansen (2000).

(*) Data until 1996

(**) Data until 1994

Table 4

Diversification Cones and EU Membership				
	Probit		**Logit**	
	SC	**SC**	**SC**	**SC**
EU	0.754***	0.629***	1.215***	1.007***
	(0.042)	(0.044)	(0.070)	(0.073)
Observations	5852	5852	5852	5852
Time Effects	No	Yes	No	Yes
Log pseudo Likelihood	-3881.995	-3566.468	-3881.995	-3569.630
Pseudo R^2	0.040	0.118	0.040	0.118

Probit and Logit estimates of Equations (8) and (9).
The sample period is 1963-1998.
The dependent variable (SC) is a dummy variable taking a value of 1 if both countries are in the same diversification cone and 0 otherwise.
The explanatory variable (EU) is a dummy variable taking a value of 1 if both countries are members of the EU.
Robust standard errors are in parentheses

Table 5

Diversification Cones and EU Membership - Contemporaneous and Lagged Effects									
	1963-1998	1964-1998		1965-1998		1966-1998		1967-1998	
	SC	SC	SC	SC	SC	SC	SC	SC	SC
EU	0.754***	0.741***		0.728***		0.715***		0.700***	
	(0.042)	(0.043)		(0.043)		(0.043)		(0.043)	
EU(-1)			0.749***						
			(0.044)						
EU(-2)					0.767***				
					(0.046)				
EU(-3)							0.779***		
							(0.047)		
EU(-4)									0.831***
									(0.050)
Observations	5852	5699	5699	5546	5546	5393	5393	5240	5240
Log pseudo Likelihood	-3881.994	-3785.434	-3790.550	-3688.660	-3690.323	-3591.652	-3592.349	-3494.387	-3481.008
Pseudo R^2	0.040	0.040	0.038	0.039	0.039	0.038	0.038	0.037	0.041

Probit estimates of Equation (8).
The dependent variable (SC) is a dummy variable taking a value of 1 if both countries are in the same diversification cone and 0 otherwise.
The explanatory variable (EU) is a dummy variable taking a value of 1 if both countries are members of the EU (contemporaneous and lagged up to 4 periods, EU up to EU(-4), respectively).
Robust standard errors are in parentheses.

Table 6

Diversification Cones and EU Membership - Accounting for Unobserved Heterogeneity and Dynamics								
	Static Panel				**Dynamic Panel**			
	Unbalanced 1964-1998		**Balanced 1967-1993**		**Unbalanced 1964-1998**		**Balanced 1967-1993**	
	SC	**SC**	**SC**	**SC**	**SC**	**SC**	**SC**	**SC**
SC(-1)					2.540***	2.538***	2.246***	2.242***
					(0.061)	(0.061)	(0.088)	(0.090)
SC(0)					0.822***	0.823***	1.738***	1.742***
					(0.069)	(0.069)	(0.247)	(0.247)
EU	1.298***		1.134***		0.468***		0.348**	
	(0.062)		(0.081)		(0.108)		(0.141)	
EU(-1)		1.535***		0.846***		0.480***		0.380***
		(0.071)		(0.060)		(0.108)		(1.409)
EU(0)					-0.330	-0.345	-0.853***	-0.876**
					(0.213)	(0.214)	(0.404)	(0.405)
EU(1973)					0.803***	0.802***	2.250***	2.241***
					(0.180)	(0.180)	(0.514)	(0.513)
EU(1981)					-0.456***	-0.461***	-1.484***	-1.486***
					(0.173)	(0.173)	(0.495)	(0.494)
EU(1986)					-0.574***	-0.565***	-0.953***	-0.954***
					(0.106)	(0.105)	(0.219)	(0.218)
EU(1995)					0.480***	0.480***	1.160***	1.164***
					(0.074)	(0.074)	(0.192)	(0.192)
Observations	5699	5699	4617	4617	5196	5196	4131	4131
Log pseudo Likelihood	-2287.987	-2250.220	-1545.320	-1565.978	-1177.446	-1176.957	-811.848	-811.223

Estimates of Equations (10)-(12), i.e., static and dynamic Probit panels.
The dynamics panel specification follows that suggested by Wooldridge (2005).
The dependent variable (SC) is a dummy variable taking a value of 1 if both countries are in the same diversification cone and 0 otherwise.
SC(-1): One-period lagged/SC(0): Initial period.
The explanatory variable EU is a dummy variable taking a value of 1 if both countries are members of the EU (contemporaneous and lagged up to 4 periods).
EU(-1): One-period lagged/EU(0): Initial period.
EU (year) is a dummy variable taking a value of 1 if both countries are simultaneously members for the first time in that year.
The estimations include year-fixed effects for enlargement years (1973,1981,1986, and 1995) which are not reported.
Standard errors are in parentheses.

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