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Multi-dimensional regional inequality as an alternative allocation mechanism For EU Structural Funds remittances: The case of Spain and Hungary

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Abstract. We propose a multidimensional approach to regional inequality as an alternative allocation mechanism for EU Structural Funds remittances based on per capita GDP, particularly after EU enlargement. The indicators of regional inequality are combined to a composite index by means of Maasoumi's aggregator function. We propose Partial Common Principal Component Aanalysis as the estimator of the weights for the aggregator function. Application of the multidimensional approach to Spain shows that there are substantial differences between the rankings of the regions obtained by means of the multidimensional approach and the traditional approach based on per capita GDP. For Hungary, which is less developed than Spain, the rankings differ less.

JEL classification: R58, R15

1. Introduction

One of the challenges the European Union (EU) is facing in the context of enlargement is reform of its Common Structural Policy (CSP).[1] At present the main allocation mechanism is per capita regional income. Regions whose per capita income falls short off the treshold of 75% of EU average GDP are eligible for support from the Structural Funds. Spain, Portugal, Greece and Ireland have been the main beneficiaries up till now. After enlargement, the total number of regions eligible for support from the Structural Funds will substantially increase, if the present allocation mechanism is maintained. With a few exceptions, all regions in the new member states will be eligible.

Heijman (2001) presents some budgetary consequences of enlargement under the present allocation mechanism for Structural Funds remittances. For instance, if enlargement had been completed by 1994 total spending from the Structural Funds would have gone up from the present amount of

[1] Another policy area that will be substantially reformed in the future is the Common Agricultural Policy.

138.2 billion euro to 796.8 billion euro which is an increase of 475%. Furthermore, the new member states would have used up more than 83% of the total resources. Especially Poland, a country with a relatively large population and low income, would have benefited from the present allocation system: it would have received about 60% of the total remittances.[2]

Needless to say that such a scenario is strongly opposed by the present member states. First of all they face a drastic increase of their contributions to the financing of the Structural Funds. Moreover, the present beneficiaries expect substantial cuts in their remittances in a bid to reduce spending and to free funds for the new member states. On the basis of this, there is growing consensus that there is a need for reform of the CSP including the allocation mechanism.

Overhaul of the CSP should include the criteria applied to allocate Structural Funds remittances. As mentioned above, per capita regional GDP has been the most important criterion so far. This is in line with *traditional* theoretical approaches and empirical analyses of regional welfare and inequality.[3] But there is a growing recognition among economists and social scientists that economic indicators (particularly GDP) or income related measures by themselves are not sufficient to analyze the extent and distribution of welfare and inequality at country as well as the regional level (Ram, 1982). It has been shown that inequality is linked to many aspects of an individual's life, not just income (Sen 1973, 1985). The environment in which people live matters as much as income. Several aspects such as general economic, social, political, environmental, and cultural conditions rather than income alone affect welfare and inequality (McGranahan et al. 1985; Mazundar 1996; Cohen 2000).[4] Hence, measurement of inequality has to contend with the multidimensionality of the welfare concept.

The purpose of this paper is to present a brief description of the CSP. Moreover, we discuss a multi-dimensional approach to measure regional inequality as an alternative to an approach based on a single criterion, i.e., per capita GDP. In this context we present a generalization of conventional Principal Component Analysis to estimate the weights needed in a multidimensional analysis of regional inequality. Finally, we illustrate the difference between the traditional GDP based allocation mechanism and a multidimensional approach in a case study relating to Spain and Hungary in a bid to develop alternative mechanisms for the allocation of Structural Funds remittances.

The country selection was made with the aim to analyze regional disparities between a *Cohesion Country* (Spain) and one of the new Member

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[2] The proposed transfers would most likely have exceeded the absorption capacity of the new member states. To give an impression, the transfers would have amounted to 50% of the GDP of Estland. For Hungary this percentage would have been 37%, for Poland 57%, 5% for Slovania, 21% for the Czech Republic and 2 % for Cyprus. It is highly unlikely that a sufficient number of projects could have been identified and executed within the mandatory time span (Heijman, 2001).

[3] For a critical review of income as the dominant indicator see amongst others Dasgupta (1993), Adelman (1980), Shorrocks (1980), Kakwani (1981), Cowell (1995), Jenkins et al. (1998)

[4] Not all of the various dimensions are directly observable or measurable in monetary terms (Hansen 1995; Nijkamp 1998; Folmer 1986).

States (Hungary). The regions in the Cohesion Countries (Spain, Greece, Ireland, and Portugal) rank low in terms of per capita income. Per capita GDP in the richest EU countries (e.g., Germany, Austria, Belgium) amounted to approximately 25.000 US$ in 1997 while it ranged between 17.000 US$ (Ireland) and 10.000 US$ (Portugal) in the Cohesion Countries (United Nations, 1999). Moreover, (and related to this) regional inequality in terms of per capita income has been substantially larger in the Cohesion Countries than in the richest member states with a disproportionally large number of regions below the qualifying treshold for Structural Funds entilement.

The new member states in Central and Eastern Europe have extremely poor regions in terms of per capita GDP. As mentioned above, the completion of enlargement has resulted in a substantial increase of economic disparities in the EU. But, they have recently pursued social policies which may have improved people's standard of living. Actually, indicators of health (such as health expenditures as percentage of GDP) and education (school life expectancy) for these countries show figures similar to those in the Cohesion Countries.[5]

The remainder of this paper is organised as follows. In Sect. 2 we present a brief description of the SCP. In Sect. 3 we discuss the methodology applied in the case study while empirical results are presented in Sect. 4. The paper ends with a summary and conclusions.

2. The common structural policy [6]

Regional disparities in the European Union have been the rationale for the CSP. Its legal basis is the European Treaty (Art. 130A-E). The objectives are:

1. assisting regions with a per capita GDP smaller than 75% of the EU average;
2. assisting regions affected by the decline of traditional industries;
3. combating long-term unemployment;
4. helping workers adapt to technological change;
5. assisting rural areas in the process of structural reform of agriculture;
6. assistance to regions with extremely low population density (fewer than 8 per km^2) and below average per capita GDP.

The main instrument to achieve the objectives of the SCP is formed by the Structural Funds. They include: the European Social Fund, the European Regional Development Fund, the 'Guidance' section of the European Agricultural Guidance and Guarantee Fund, and the Fisheries Guidance Instruments. In addition, there is a Cohesion Fund aimed at financial contributions to projects in the fields of the environment and trans-European infrastructure.

[5] In 1996 the highest health expenditures in the Cohesion Countries was 8.2% of GDP (Portugal) while it was 12.8 % of the GDP (Czech Republic) in the group of new member states (World Health Organisation, 1999). School life expectancy in 1995 ranged between 13.5 and 14.5 in the Cohesion Countries, and 12.6 and 12.9 in the group of new member states (United Nations, 1999).

[6] See also Heijman (2001).

During the period 1994-1999 total expenditure was 138.2 billion euro (1994 prices). Objective 1 was by far the most important, absorbing more than 65% of total spending. In addition, the four poorest member states (Greece, Portugal, Spain and Ireland) received 14.5 billion euro in this period in the framework of the so-called Cohesion Fund. The total amount (152.7 billion Euro) of Structural Funds spent was about 0.33% of the member states' combined GDP. The target set by the European Council in Edinburgh in 1992 was 0.46% of total GDP for the period 1994–1999, which, apparently, was not reached.

In the main beneficiary countries (Spain, Ireland, Greece and Portugal) receipts from the Structural Funds amounted to about 5% of their GDP per year on average (Bainbridge 1997, p. 463; Senior Nello and Smith 1998, p. 35). In spite of these efforts, regional disparities have not been substantially reduced. On the basis of this Van Meerhaeghe (2000) concludes that structural policy has failed from this perspective. However, for some countries, for instance Ireland, there is evidence that structural assistance has substantially stimulated economic growth and that per capita GDP has increased.

In Agenda 2000, the European Commission proposed to maintain the percentage of 0.46 for the period 2000–2006. This concerns an estimated amount of 275 billion euro (1997 prices). A proposed 45 billion euro has been earmarked for the enlargement of the EU to be spent in countries from Central and East Europe.

Allocation of Structural Funds is based on the following principles:

- Concentration: funding should be strictly confined to the objectives of the CSP. Since several of the CSP objectives are related to the objectives of other policy areas, such as the Common Agricultural Policy, there is a risk of confounding policies. The concentration principle is aimed at restricting this.
- Partnership: carrying out structural policy should be done in close cooperation with national, regional and local authorities. This principle is related to the subsidiarity principle which states that policy measures should be determined by the lowest level of authority suited for a given problem so as to achieve compatibility with local preferences (Folmer and Howe, 1992). It is also a safety valve against the omnipotence of centralized bureaucracies.
- Programs: funding should be based on programs rather than individual projects. This principle is a safety valvet against fragmentation and inefficiency inherent to financing of individual projects. It is based on the assumption that regional development requires a comprehensive development policy. Individual projects cannot substitute for a development plan; they are constituting elements of it.
- Additionality: member states should co-finance programs that are funded in the context of Structural Funds. On average the EU finances 50%; the national government the other half (SER, 1997). The rationale of this principle is to guarantee member state commitment. Regional development is the prime responsibility of the member states whereas the EU supports, facilitates and complements member state policies.

An important aspect of enlargement is the so-called "phasing out" problem. It relates to the fact that regions currently eligible for remittances from the

Structural Funds would exceed the treshold of 75% of the average GDP per capita (the eligibility criterion) in an enlarged EU (Heijman 2001). This is a consequence of the fact that per capita GDP in the new member states is substantially lower than in the "old" member states including the Cohesion Countries. Consequently, average EU per capita GDP will go down after accession of the applicant countries. As mentioned in the previous section, phasing out is one of the main problems the Cohesion Countries are facing in the context of enlargement.

3. Methodological aspects of multidimensional analysis of inequality

Before going into detail, we make the following introductory remarks. Globally speaking, two main approaches to measuring inequality can be distinguished (Sen 1985). First, the so-called individualistic or non-paternalistic approach which takes its starting point in an individual's preferences and relates inequality to the basket of goods consumed. The second or "capability to function" approach seeks to evaluate inequality in terms of options available to a person for attaining her or his achievements. It is this latter approach that will be used in this paper. In the present context it consists of evaluating a set of indicators that reflect people's objective circumstances in a given cultural or geographic unit (Diener and Suh 1997). The social indicators used in this context are related to a wide range of elements of human life and living conditions.

As mentioned in the Introduction, we opt for a multidimensional rather than a single indicator approach in this paper. Several types of composite indices representing multidimensional inequality to compare regions or countries have been developed. The best-known composite index is perhaps the *Human Development Index* (HDI). It is published annually by the United Nations and consists of a set of three inter-country indicators (life expectancy, literacy rate and real per capita GDP). The term "human development" refers to a process of enlarging people's choices (UN 1999) in which income is considered a means of expanding choices and well-being. A similar index was suggested by Morris (1979), i.e., the *Physical Quality of Life Index* (PQLI) which is obtained by combining indicators relating to people's basic needs (life expectancy, infant mortality, literacy). Maasoumi and Nickelsburg (1988) constructed a composite index of well-being in a case study relating to the U.S. states using three indicators: per capita annual nominal income, net housing equity, and average schooling. It is based on the theoretical notions developed by Maasoumi (1986). Ram (1982) used additional indicators to those included in the PQLI index such as access to safe water, access to health care (i.e., number of physicians), calorie intake, amongst others. Hirschberg et al. (1991), Slottje (1991), Slottje et al. (1991) and Diener (1995) consider an even a wider range of social and economic variables. Finally we mention COR (1999) who considered variables relating to population, income and wealth, health conditions, housing, services availability, crime and social pathology, employment and labour conditions, environment, personal relationships and participation.

A methodological problem in the context of a multidimensional approach relates to the strong correlations between welfare indicators and economic measures. Several studies (McGranahan 1972; Larson and Wildfor 1979;

McGillivray 1991) have indicated that the former fail to provide additional information. Strong correlations between economic measures and other sets of indicators would lead to suggest that the former are sufficient. However, Diener and Suh (1997) demonstrate that social indicators provide information which is not contained in economic measures. They conclude that the measurement of nations' wealth should be considered a first approximation to measurement of welfare but that it is not enough to reflect the quality of life experienced by people.

Another problem relates to the accuracy of data used in international comparisons of inequality. To some extent cross-country data are deficient for comparative analyses (Dasgupta 1993; Hirschberg et al. 1991) because of large variations in definitions and in measurement of indicators among countries (Hicks and Streeten 1979). Among inequality indicators used in many studies, the literacy rate is considered one of the weakest. The standards for abilities in writing and reading differ much among countries so that it is difficult to find an appropriate indicator for international comparisons. For other indicators such as infant mortality, differences across nations may arise due to differences in data collection. It should also be observed that unevenness of country data does not only relate to inconsistencies in definitions of variables and data collection, but also to the geographical scale at which information is gathered. In most international surveys, country is the only unit of measurement while there is not much statistical information available for other territorial divisions such as regions (Cohen 2000). This may result in the necessity to substitute real data by estimates.

We now turn to methodological aspects to measure welfare and inequality. In recent years there have been many attempts to develop methodological procedures to handle the multidimensional nature of inequality (see amongst others Morris 1979; Ram 1982; Maasoumi and Nickelsburg 1988; Hirschberg et al. 1991; McGillivray 1991; UNDP 1990–1997; Boelhouwer and Stoop 1999; Ferriss 2000). These procedures come down to combining the indicators corresponding to a wide range of social and economic aspects (including life expectancy, adult literacy rate, GDP, unemployment, calorie intake, etc) into a composite index by assigning weights to the indicators. Several approaches have been proposed to obtain weighting schemes. One is to equally weight indicators. But this is an arbitrary approach and, moreover, there is no rationale for assigning the same weights to different indicators. An alternative approach consists of weighting indicators by market prices. However, some indicators relate to non-market goods, like air qualiity or many other environmental amenities for which no prices exist. Weights are also obtained by consulting experts (Giannas et al. 1999).

Another class of approaches applies statistical methods such Principal Components Analysis, Factor Analysis or Hedonic Models (Hirschberg et al. 1991) to obtain weights. Particularly Principal Component Analysis (PC) has been applied in a variety of welfare studies such as Maasoumi and Jeong (1985), Maasoumi and Nickelsburg (1988), Hirchberg et al. (1991), Boelhouwer and Stoop (1999). Typical for these studies is that the coefficients of the first unrotated component constitute the weights assigned to the indicators.

For reasons of objectivity, we opt for a statistical approach in this paper, particularly PC which is increasingly being applied in the present kind of case studies (see e.g., Quadrado et al. 2001). In this context a further methodological decision has to be concluded: whether or not to consider different countries or regions (in the present case study the Hungarian and Spanish regions) as belonging to one population. In the one population case the regions in both countries are pooled and the sample variance-covariance matrix of the pooled data is used to estimate a single set of weights applying to all regions. Alternatively, the regions in different countries are considered belonging to different populations and consequently different sets of weights are used.

It is common practice to conclude the pooling decision on ad hoc information or intuition. However, a generalization of conventional PC to several groups has been developed by Flury (1988) and Flury and Neuenschwander (1995). This approach allows for statistical tests of similarities between the different data sets before pooling them and proceeding on the basis of one common variance-covariance structure for the various groups. The generalization is denoted Partial Common Principal Components Analysis (PCPC).[7] We adopt PCPC in this paper and test whether or not the Hungarian and Spanish regions belong to two different populations and whether or not the differences between the variances and covariances of the indicators for these countries are merely due to sampling variability or also to structural differences. Finally we observe that PCPC is preferred to PC on the basis of the *principle of parsimony*.

The following observations apply. First, due to inter alia restrictions implied by comparability between Spain and Hungary, the set of common indicators turned out to be very limited. Particularly, the set of indicators is too small to construct a composite index of welfare. Rather, the composite index presented below merely represents some aspects of inequality.

We now turn to the composite index and estimation of the weights of the indicators. Before going into detail, we introduce the following notation. Let w represent the composite index of regional inequality of a group of N regions and let the subscript i denote the ith unit. The composite index for region i is a function of a set of indicators represented by the vector s_i with elements s_{ik} $(k = 1, \ldots, K)$ (in this paper the indicators in Table 1):

$$w_i = w(s_i) \tag{1}$$

Moreover, we define w_i^* as:

$$w_i^* = \frac{w_i}{\sum_{i=1}^{N} w_i} \tag{2}$$

We use Maasoumi's (1986) aggregator function to construct the composite index. It enables us to reproduce the maximum amount of information contained in the original data. Maasoumi's aggregator function reads:

$$w_i = \sum_{k=1}^{K} \delta_k s_{ik} \tag{3}$$

where $\delta_k = \alpha_k / \sum_{k=1}^{K} \alpha_k$ are the weights associated with s_{ik}.

[7] Empirical applications are mostly in biology (Flury 1988) and rare in the social sciences. (Quadrado 1999) and Quadrado et al. (2001) are recent applications in the social sciences.

Table 1. Indicators used to construct the composite index

Infant mortality rate: per 10,000 live-born.
Unemployment rate: percentage of the labour force.
Dwellings supplied with public sewerage: percentage of total number of dwellings.
Telephone lines: per 1000 inhabitants.
GDP: per capita.
Doctors (physicians and paediatricians): per 10,000 inhabitants.
Hospital beds: per 10,000 inhabitants.
Students at primary and secondary school: per 10,000 inhabitants.

(3) implies that different indicators are weighted using different weights. The weights α_k and hence δ_k are estimated by means of PCPC if the data support the hypothesis of a common eigenvector of the variance-covariance matrix for the pooled data set and PC otherwise.

The essence of PCPC can be summarized as follows. Consider T ($t = 1, \ldots, T$) groups, i.e. in the present case, groups of countries. The basic assumption of PCPC is that a subset (usually but not necessarily consisting of 1 element) of principal components is common for all groups whereas the remaining principal components are specific to each group. (In the present case study $T = 2$, i.e. the Spanish regions and Hungarian counties). In the case that data supports PCPC, the selected common principal component coincides with the one associated with the largest eigenvalue. The variance associated with that component may differ in both groups.

The null hypothesis of one common principal components H_{PCPC} (1) and the alternative of unrelated structures: H_{PC} read:

$$\begin{aligned} H_{PCPC}(1) &: \Psi_t = \beta\Lambda\beta \\ H_{PC} &: \Psi_t = \tilde{\beta}_t\Lambda_t\tilde{\beta}_t' \end{aligned} \tag{4}$$

where

Ψ_t is a positive symmetric (p × p) covariance matrix for the t-th group;

Λ is a (p × p) diagonal matrix with λ_1 the eigenvalue common to the T groups and $\lambda_2, \ldots, \lambda_p$ eigenvalues specific for the t-th group;

β is a (p × p) matrix $(\beta^c, \beta_t^s)$where β^c is the common eigenvector for all groups and β^s (with dimension $(p \times (p\text{-}1))$ is the matrix of eigenvectors that are specific for group t.

$\tilde{\beta}_t$ is the (p × p) matrix of specific eigenvectors for group t;

Λ_t = diag $(\lambda_{t1}, \ldots, \lambda_{tp})$ is the (p × p) matrix of eigenvalues specific fot the t-th group.

The test statistic for the exact maximum likelihood test of the null hypothesis of a common eigenvalue versus the alternative of unrelated covariance matrices is (Flury 1988):

$$\ell = -2\log\frac{L(\hat{\Psi}_1, \ldots, \hat{\Psi}_T)}{L(S_1, \ldots, S_T)} = \sum\nolimits_{t=1}^{T} n_t \log\frac{|\hat{\Psi}_t|}{|S_t|} \tag{5}$$

where S_t is the t-th sample variance-covariance matrix and $\hat{\Psi}_t$ is the estimate of the variance-covariance matrix for the th group under the null hypothesis. Under the null hypothesis the distribution of ℓ is asymptotically chi squared with $\frac{1}{2}\{(t-1)p(p-1)-(p-q)(p-q-1)\}$ degrees of freedom. (For further details see Flury 1988.)[8]

If the null hypothesis of a common eigenvector is not rejected, the weights attached to the variables (component coefficients) will be the same for the various groups, in this case study Hungary and Spain. If the hypothesis is rejected, PC is applied to each country separately. Then weights assigned to indicators are different for each country; they are the coefficients of the first component for each country.

We can investigate the stability of the coefficients of the first common component by means of the standard errors. Particularly, if the standard error of a given coefficient is smaller than 0.1, then that coefficient may be assumed to be stable (Flury 1988).

The use of a variance-covariance matrix to obtain the weights using PCPC or PC is based on the assumption that all indicators are comparable in terms of variances and covariances and units of measurement. In this paper we use indicators with negative (e.g., infant mortality) and positive (e.g., the number of doctors) associations with inequality. To achieve comparability, the indicators are re-scaled so that they range between 0 and 1, with 1 representing the highest level of welfare. Nijkamp (1988) suggests the following transformation:

$$\tilde{s}_{ij} = \frac{s_{ij}}{s_j^{\max}}, \text{ if j is a positive indicator} \tag{6}$$

$$\tilde{s}_{ij} = 1 - \frac{s_{ij}}{s_j^{\max}}, \text{ if j is a negative indicator} \tag{7}$$

where $s_j^{\max} = \max_i s_{ij}$

Division by s_j^{max} brings about that the coefficients range between 0 and 1. Equation (7) transforms an indicator with a negative association with welfare (e.g. infant mortality) into an indicator with positive association.

4. Empirical Results

In the case study we consider socio-economic indicators that are available at the NUTS-II level[9]. Hence, counties are considered for Hungary while regions are the corresponding units for Spain. The sources are the Hungarian Statistical Office and the Spanish Statistical Office and data is for 1996 or 1997. The list of indicators is presented in Table 1.

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[8] The test has the usual properties of a likelihood ratio kind of test, i.e., it is unbiased and consistent.

[9] Eurostat has defined a regional classification system that is applicable in all the EU member states. The 'nomenclature des unités territoriales statistiques (NUTS)' has four main levels (NUTS 0-III), with NUTS 0 being the member states and the higher levels indicating subnational regional units of decreasing order. In the UK, for example, NUTS III is equivalent to counties and local authority regions, NUTS II to county/local authority groupings and NUTS I to standard regions. Because of the availability of data on that level, most European regional analyses are based on NUTS II level of regional disaggregation (McDonald and Dearden 1999).

Table 2. First and second eigenvalues, standard errors and proportions of total variance recovered

	Hungary		Spain	
	1st PC	2nd PC	1st PC	2nd PC
Eigenvalues	0.1016	0.0524	0.0736	0.0631
Standard Errors	0.033	0.017	0.025	0.022
Proportion of Total Variance	0.49	0.25	0.40	0.34

As a first step, we apply PC[10]. The results obtained are given in Tables 2 and 3. In Table 2 we present information concerning eigenvalues for the first and second principal components for both Hungary and Spain. In both cases, the first principal component recovers less than 50% of the total variance which is substantially less than the 70% usually required (Flury 1988).

The goodness of fit can be improved by considering a greater number of principal components. Table 2 reveals that the accumulated proportion of variance accounted for by the first two components together is 74% so that the 2 component PC model fits the data well (according to this criterion). We now interpret these two components (see Table 3).

Two non-economic indicators (dwellings supplied with public sewerage and number of physicians) and two economic indicators (unemployment and GDP) are the dominant variables of the first component for Hungary (Table 3). Unemployment, GDP, telephone lines, hospital beds and physicians dominate the first component for Spain. The component coefficients for both countries are positive for all variables.

The dominant variables of the second component are sewerage (with a negative coefficient), physicians and education (with positive coefficients) for Hungary. For Spain sewerage is the only significant variable of the second component[11].

The robustness of the component coefficients is investigated by means of the standard errors. As argued in the preceding section, if the standard error of a coefficient is smaller than 0.1, then that coefficient may be assumed to be stable. On the basis of this rule of thumb, the standard errors of the first component reveal weak stability of the coefficients since several of them (e.g., infant mortality, sewerage and physicians) are larger than 0.1. This holds for both countries. Therefore, the first component cannot be adequately interpreted.

We do not attempt to interpret the second principal component because its coefficients are also unstable. Most of the standard errors of the second component are higher than those of the first component. Taking this into

[10] As outlined in the preceding section, the natural first step would be PCPC and to resort to PC in the hypothesis of a common principal component were rejected. We reverse the order here to illustrate some differences between PC and PCPC.

[11] The reason the sewerage variable is included in the analysis lies in the situation in Hungary, where in contrast to Spain, coverage is limited and varies over regions. For Spain the lowest percentage is 93.7% while most regions have percentages higher than 95%. The differences among the Spanish regions are so small that this indicator does not seem useful to discriminate among them.

Table 3. Coefficients of the first and second principal component[a)]

	Hungary				Spain			
	1st PC		2nd PC		1st PC		2nd PC	
Infant mortality	0,104	(0.131)	0,075	(0.260)	0,059	(0.252)	0,115	(0.223)
Unemployment	0,505	(0.071)	−0,090	(0.198)	0,596	(0.183)	−0,098	(0.951)
Sewerage	0,520	(0.244)	−0,709	(0.86)	0,032	(1.562)	0,985	(0.062)
Telephone lines	0,269	(0.064)	0,163	(0.105)	0,370	(0.070)	−0,029	(0.590)
GDP	0,411	(0.084)	0,158	(0.183)	0,526	(0.051)	0,018	(0.835)
Physicians	0,368	(0.162)	0,419	(0.190)	0,321	(0.145)	0,074	(0.517)
Hospital beds	0,176	(0.079)	0,180	(0.097)	0,345	(0.108)	−0,005	(0.560)
Education	0,237	(0.174)	0,473	(0.146)	0,071	(0.028)	−0,009	(0.115)

[a)] Standard errors in brackets.

Table 4. Partial common principal component test

Number of estimated parameters in the model	65	
Likelihood ratio test[1)]	4.221	
Degrees of freedom	7	
p-Value	0.7540	
Characteristic roots for the first common principal component.		
	Hungary	Spain
Eigenvalues	0.09471	0.06578
Proportion of total variance	0.46	0.36

[1)]Relates to the test for a common principal component.

account, interpretation of this component would be fallacious. In view of this, we conclude that the principal component model for each country separately based on two common components is not adequate for comparing inequality between Hungary and Spain.

Next we apply PCPC and test the hypothesis that Spain and Hungary share the first component. The results in Table 4 show that the χ^2 value obtained is 4.221 whereas the critical value at the 5% level of significance with 7 degrees of freedom is 14.07 (p-value 0.7540). Hence, we accept the assumption that Spain and Hungary have the first component in common.

Another feature of Table 4 is the low proportion of variance recovered. Compared to the PC analysis it has slightly decreased from 0.49 to 0.46 for Hungary whereas for Spain it has slightly increased from 0.34 to 0.36. The country-specific coefficients for the second principal component were deficient in the sense of wrong signs and standard errors substantially larger than 0.1.[12] Therefore the analysis was restricted to the common principal component. Because the proportion of total variance explained is below the usual requirement of approximately 0.75, the results presented below should be interpreted with caution.

The quality of the analysis could possibly be improved by expanding the set of indicators. For instance, there is only one indicator for education. A breakdown of the education variable into enrollment at primary and higher

[12] The results are available from the authors upon request.

Table 5. Approximate PCPC maximum likelihood estimates[a)]

	1st PCPC	
Infant mortality	0.0152	(0.1076)
Unemployment	0.4818	(0.0943)
Sewerage	0.7457	(0.1713)
Telephone lines	0.2011	(0.0601)
GDP	0.3294	(0.0697)
Physicians	0.2082	(0.0888)
Hospital beds	0.1353	(0.0554)
Education	0.0301	(0.0261)

[a)] Standard errors in brackets.

education might improve the results. Further improvement could possibly be achieved by including variables such as quality of social security, infrastructure, etc. Also the use of panel data., to analyze the developments over time, would be worthwhile considering in this regard.

In Table 5 the approximate maximum likelihood estimates of the coefficients of the first common component are displayed along with their standard errors. The dominant variables are dwellings supplied with public sewerage and unemployment, followed by GDP, telephone lines and physicians. We observe that the standard errors of sewerage and infant mortality exceed 0.1 substantially and slightly, respectively. This, in combination with the low proportion of variance explained, affects the reliability of the remainder of the analysis.

The weights obtained in Table 5 are used to rank the Spanish and Hungarian regions in terms of the indicators presented in Table 1. The results are presented in Table 6. Most of the Spanish regions rank higher than the Hungarian counties in terms of the composite index. The first 9 positions are taken by Spanish regions while there are only two Hungarian counties (Gyõr-Moson-Sopron and Vas) in the top 15. These are the most industrialized counties, and moreover, they benefit from their geographical location (both are on the Austrian border). The Spanish regions Extremadura and Andalucía, and especially Galacia rank low (23rd ,24th ,34, respectively) compared to the Hungarian counties.

Table 6 also contains the ranking for both countries in terms of per capita GDP. First of all we observe that all Spanish regions rank higher than the Hungarian regions. The difference between the poorest Spanish region (Extremadura) and the richest Hungarian region is 1260 Euro. Moreover, the difference between the poorest Spanish region and Hungarian county is 5772 Euro.

We observe substantial differences between both rankings for Spain. For instance, whereas Madrid ranks first in terms of per capita GDP it only ranks 5th in terms of the composite index. For Hungary both rankings closely correspond to each other. The largest differences are observed for Gyor-Moson-Sopron and Vas who rank substantially higher in terms of the composite index than per capita GDP.

The above leads to the tentative conclusion that the disparity between rankings on the basis of GDP and the composite index is larger for less

Table 6. Ranking of the Spanish and Hungarian regions in terms of GDP and the composite index

Spanish regions	GDP	Rank order	Composite index	Rank order	Hungarian counties	GDP	Rank order	Composite index	Rank order
Andalucia	8482	16	0.6087	24	Bács-Kiskun	2522	30	0.5034	30
Aragon	12447	7	0.7791	6	Baranya	2584	28	0.5794	26
Asturias	10036	12	0.7335	12	Békés	2473	32	0.5149	29
Baleares	14162	3	0.7791	6	Brosod-Abaúj-Zemplén	2395	34	0.3188	35
Canarias	10890	8	0.7335	13	Bud-Pest	6460	18	0.7182	16
Cantabria	10574	11	0.7689	7	Csongrád	3097	23	0.6937	20
Castilla La Mancha	9495	14	0.7152	18	Fejér	4040	19	0.6466	22
Casilla Leon	10875	9	0.7589	9	Gyõr-Moson-Sopron	3896	21	0.7570	10
Cataluna	13970	4	0.8387	3	Hajdú-Bihar	2531	29	0.5798	25
Comunidad Valenciana	10767	10	0.7593	8	Heves	2490	31	0.5495	28
Extremadura	7279	17	0.6088	23	Jász-Nagykum-Szolnok	2592	27	0.4613	32
Galicia	9232	15	0.3806	34	Komárom-Esztergom	2969	24	0.6580	21
Madrid	14991	1	0.8064	5	Nógrád	1948	36	0.2580	36
Murcia	9552	13	0.7231	15	Somogy	2420	33	0.4909	31
Navara	14503	2	0.8822	1	Szabolcs-Szatmár-Bereg	1997	35	0.4014	33
Pais Vasco	13519	5	0.8709	2	Tolna	2904	25	0.5750	27
Rioja	13213	6	0.8351	4	Vas	3937	20	0.7476	11
					Veszprém	2769	26	0.7179	17
					Zala	3145	22	0.6948	19

developed countries like Hungary than for more developed countries such as Spain. Needless to say, that this is an interesting topic for further research.

5. Conclusions

The enlargement of the European Union necessitates an overhaul of several of its policies. This definitely applies to the Common Structural Policy since expenditure would go up by more than 400% if the present allocation rules were maintained. Moreover, several of the present beneficiary regions in notably Spain, Portugal, Greece and Ireland would be phased out which will induce strong opposition. In the paper we have argued that reform should include the allocation mechanism. Allocation at present is primarily based on per capita GDP. However, there is a vast theoretical and empirical literature showing that a multivariate approach to welfare and inequality is superior to a univariate approach.

We present a composite index of welfare or inequality and propose Maasoumi's aggregator function to construct it. This aggregator function has several favorable characteristics to construct a composite index for interregional comparisons of welfare and welfare inequality. We also argue that Partial Common Principal Component Analysis is an appropriate methodology to estimate the weights of Maasoumi's aggregator function. We apply this approach in a case study relating to Spain and Hungary. Due to data limitations, the application is merely an illustration of the approach. Given this restriction, we find that the ranking of the regions obtained by means of the composite index differs substantially from the ranking obtained by the univariate index per capita GDP for Spain. For Hungary both rankings correspond rather closely. We conclude that these findings about the disparity and correspondence of the rankings for countries of different income levels could be an interesting topic for further research.

We repeat that the empirical analysis is subject to some caveats. Nevertheless, it shows that the multivariate approach is worthwhile considering as an alternative to the univariate approach to identify regions eligible for Structural Funds remittances. Moreover, it may be used in identifying the most important factors influencing regional disparity. Concentration on these factors could be instrumental in reducing the budgetary problems following enlargement in the sense that the Structural Funds could be targeted at the dominant factors in the first place.

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