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Social Exclusion and Economic Growth: An Empirical Investigation in European Economies

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Social Exclusion and Economic Growth: An Empirical Investigation in European Economies

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Abstract

The aims of this article are to propose an overall index of social exclusion and to analyze its relationship with economic growth in European countries. We approach social exclusion as a multidimensional phenomenon by a three-mode principal components analysis (Tucker3 model). This method is applied to estimate an indicator of social exclusion for 28 European countries between 1995 and 2010. The empirical evidence shows that in short run (a) Granger causality runs one way from social exclusion to economic growth and not the other way; (b) countries with a higher level of social exclusion have higher growth rates of real GDP per capita; and (c) social exclusion has a larger effect than the income inequality on the economic growth. The policy implications of our analysis is that social inclusion is not a source of economic growth in the short term.

Keywords: Social exclusion, Economic Growth, Multidimensional index, Three-mode principal components analysis.

JEL: D63, I3, O15, O52, C82.

Introduction

The aims of this paper are (i) to propose an aggregate overall index of social exclusion in order to assess the intensity and evolution of this phenomenon in European countries, and (ii) to analyze its relationship with economic growth in Europe over the period 1995–2010.

Social exclusion is a multidimensional general concept that refers not only to material, economic, or health deprivation, but also to deprivation from social relationships and participation in the society. A growing literature addresses the issue of how appropriately an index of overall social exclusion (e.g. Burchardt et al. 1999, Bradshaw et al. 2000, Whelan et al. 2002, Tsakloglou and Papadopoulos 2002, Chakravarty and D'Ambrosio 2006, Poggi 2007, Bossert et al. 2007) can be measured. The assessment of social exclusion by scientific debates has become an even more important task both for understanding and for policy evaluation. In fact, an aggregate index of social exclusion can serve as a useful indicator of both the overall level of social disadvantage and the overall effectiveness of governmental social policies (Micklewright, 2001).

With regard to the first issue, while proposing quite a new approach for the economic field (i.e. *Multiway Principal Component Analysis*), we present an overall (macro) index of social exclusion based on five main areas: employment, poverty, income inequality, education, and health. These meta-dimensions include many accepted indicators that have been selected according to the EU approach of social exclusion (e.g. European Commission, 2004, 2008).

With regard to the second issue - analyzing the relationship between social exclusion and growth in European countries - we attempt to broaden the perspective of existing literature. This is made possible by testing the relationship between economic growth and the previously estimated overall index of social exclusion. This analysis is carried out on a dataset by collecting observations for 28 European countries over the period 1995–2010.

The relationship between income inequality and growth has spawned a large theoretical and empirical body of literature (for a survey, see Barro 2000, Forbes 2000, and Kanbur 2000, among the others). In this research, we aim at extending this literature by examining the relationship between economic development and a wider concept of disparities/inequality, that is, social exclusion. The latter captures the multi-dimensional aspects of (relative) socioeconomic disadvantages more effectively than income inequality and poverty do.

Recent studies point out the role played by social exclusion in economic growth. The increasing scientific interest in this issue has been amplified by greater relevance that the fight of social exclusion has for European institutions. Since the Treaty of Amsterdam (1997) and the launch of *Lisbon's strategy* (2000), social inclusion has been considered one of the strategic objectives of the EU.¹ Again, in March 2010, the European Council re-launched social cohesion as one of the five key areas of the "Europe 2020 strategy." Due to the greater relevance that social inclusion has for European policies, the literature on social exclusion also has been mainly concerned with problems in European countries (Sen, 2000).² Following this strand, this article has its geographical focus on Europe. Thus, our results indirectly provide an empirical investigation of the theoretical background of the EU policies, which, in fact, assume a significant positive relationship between social inclusion and economic growth.

¹ The objective of Lisbon strategy is "to make European Union the most competitive and dynamic knowledge-based economy in the world, capable of sustainable economic growth with more and better jobs and greater social cohesion."

² See, among others, Silver and Wilkinson (1995), da Costa (1997), Gore and Figueiredo (1997), Walker and Walker (1997), Maxwell (1998), Tsakloglou and Papadopoulos (2002).

The empirical evidence shows that, in the short run, countries with a higher level of social exclusion have higher growth rates of GDP per capita.

This article is organized as follows. The definition and measurement issues related to social exclusion are discussed in Section 1. In Section 2, we outline our proposed framework for the measurement of the overall index of social exclusion and present the results. A summary of the various theories on the relationship between inequality and growth is presented in Section 3.1. Section 3.2 describes econometric model. Section 3.3 gives the results of the empirical analysis on the relationship between social exclusion and economic growth. Section 4 contains the conclusion. The data available and comparison among alternative approaches to estimate the index of social exclusion are briefly reviewed in the Appendixes A and B, respectively.

1. Social Exclusion: Definitions and measurement issues

1.1 Defining Social Exclusion

The history of the concept of social exclusion is relatively short, but the literature on the subject is already large and rapidly growing (Flotten, 2006). Several definitions are proposed in literature; however, in essence, they hold the concept that social exclusion includes a wide range of dimensions of marginalization and exclusion. These dimensions refer, not only to material, economic, or health deprivation, but also to deprivation from social relationships and participation in the society.³

Essentially four approaches to define social exclusion are proposed in the literature: (i) social exclusion as the lack of participation in social institutions (Duffy 1995, Rowntree Foundation 1998, Paugam and Russell 2000); (ii) social exclusion as the denial or non-realization of social, political, and civil rights of citizenship (Room 1995, de Haan, 2000); (iii) social exclusion as an increase in the distance among population groups (Akerlof 1997); and (iv) social exclusion as a process that leads to a state of functioning deprivations (Sen 1998, 2000).

Social exclusion borrows much from earlier literature on deprivation and poverty, and it is related, but not equivalent, to concepts such as inequality and poverty (Atkinson 1998, Robila 2006, Sen 1998, 2000, Atkinson et al. 2002, Atkinson and Marlier 2010).⁴ Opinions differ over the nature of this relationship. On one hand, poverty has been described as a cause of social exclusion, in as much as it prevents participation and access to the various networks that can help people into education, homes, jobs, and services (e.g. de Haan 2000). On the other hand, social exclusion has been considered a cause, or an element, of poverty (e.g. Jordan 1996). However, social exclusion differs from early “income” poverty, as the last is only concerned with lack of economic resources; whereas social exclusion refers to a broad range of dimensions of deprivation (Berghman 1995, Atkinson and Marlier 2010).⁵

³ For an overview of the definitions of social exclusion, see, among the others, de Haan (2000), Flotten (2006).

⁴ Notions of multidimensional inequalities and poverty can be found even in Aristotle's *Nicomachean Ethics*, and notions of exclusion and inclusion are an essential part of the concept of poverty defined by Adam Smith in the *Wealth of Nations* (Sen, 1998, 2000). So, “the helpfulness of the social exclusion approach does not lie in its conceptual newness, but in its practical influence in forcefully emphasizing - and focusing attention on - the role of relational features in deprivation” (Sen, 2000, p. 8).

⁵ However, since the work of Townsend (1979), even the income distribution approach has accepted the idea that inequalities and poverty are multidimensional phenomena (Tsakloglou and Papadopoulos, 2002).

The relationship between social exclusion, inequality, and poverty can be usefully considered in the perspective of Sen's *capability approach*.⁶ In this perspective, all these concepts deal with functioning disparities or failures, but social exclusion refers to functioning failures in terms of the inability to participate and emphasizes the role of relational features in the deprivation of capability (Sen, 2000). As Sen points out, social exclusion can be constitutively a part of capability deprivation. It happens when one is not able to relate to others and to take part in the life of the community; this can directly impoverish a person's life. For example, persistent exclusion from the labour market or credit access may impoverish a person and cumulatively lead to other deprivations, such as underconsumption, undereducation, or homelessness. However, social exclusion as well can be instrumentally a cause of diverse capability failures, inequalities, and poverty. It happens when even significant exclusions may not be impoverishing in themselves, but can lead to the impoverishment of human life through their causal consequences (Sen 2000).

This article refers to Atkinson and Marlier's (2010, p.1) definition of social exclusion as "the involuntary, not transitory exclusion of individuals and groups from society's political, economic and societal processes, which prevents their full participation in the society in which they live." According to this description, individuals or groups are excluded when (i) they suffer disadvantages in terms of education, training, employment, housing, financial resources, health, and so on; (b) their chances of gaining access to socioeconomic relations and to major social institutions that distribute these life chances are low if compared with the others; and (c) these disadvantages persist over time (Room 1990, Silver 1994, de Haan 1998, Byrne 1999).

A general consensus has formed around some key attributes of social exclusion. First, social exclusion is a purely relative and relational concept. "Exclusion" refers to a specific society at a specific point in time. It is concerned with comparisons of different individuals or groups and emphasizes the quality of their (socioeconomic) relationships. Individuals can be considered socially excluded only relatively to other members of the specific society in which they live, if there is a major discontinuity in their relationships with the rest of society, a lower degree of social participation, and lack of socioeconomic integration. So, unlike poverty, social exclusion can be only relative (Tsakloglou and Papadopoulos 2002, Bossert et al. 2007, Scutella et al. 2009). In this comparative context, if the focus is on disparities or deprivation of opportunities, we move in the direction of the idea of poverty as capability deprivation. In a similar way, social exclusion is concerned with the absence of interrelation between the opportunities enjoyed by different members of the community (Sen 2000).

Second, unlike inequality and poverty, social exclusion is a dynamic concept. Exclusion does not depend only on current situations, but also on "prospects for the future" (Atkinson 1998). Individuals or groups can be considered socially excluded when deprivation, inadequate social relations, and lack of participation persist or may worsen over time. Furthermore, according to Atkinson and Marlier (2010, p. 13), "social exclusion is not only a matter of ex post trajectories but also of ex ante expectations". Within this framework, social exclusion has to be analyzed as a process, and it requires the inclusion of time as an important variable. Therefore, forward-looking indicators are needed (Atkinson 1998, Tsakloglou and Papadopoulos 2002, D'Ambrosio and Peragine 2007).

⁶ "The capability perspective on (inequality) and poverty is inescapably multidimensional, since there are distinct capabilities and functionings that we have reason to value" (Sen, 2000, p. 4).

1.2 Measuring Social Exclusion

Social exclusion, as we have already seen, is a multidimensional concept that focuses on deprivation in different areas. As a consequence, a wide range of indicators of living standards is required to measure an overall index of this phenomenon. Taking this into account, most of the EU countries now produce, often at a micro level, several indicators that are related to the specific areas of individual deprivation and social exclusion: incomes, employment, health, education, and access to housing and other services (Atkinson et al. 2002, Marlier et al. 2007, Scutella et al. 2009).

Indices of overall social exclusion or deprivation are needed both for understanding such a multidimensional phenomenon and for policy evaluation. For example, an aggregate index makes it possible to compare social exclusion levels among different countries, to evaluate the impact of socioeconomic policies implemented by governments, and to measure changes over time and/or across countries. In addition, according to Micklewright (2001), there are at least two obvious arguments for the use of an aggregate measure of poverty and social exclusion: it would summarize the overall picture, thus avoiding the problem of fuzziness of multiple indicators, to allow a more clear understanding to emerge; it would guarantee a better communication, getting more attention on the field, as happened with the United Nations Development Program's Human Development Index.

In this regard, the multidimensional nature of social exclusion gives rise to a methodological problem of aggregation. It is about how to get a synthetic measure of overall social exclusion from a portfolio of indicators referring to several different areas of deprivation and exclusion, generally defined at the micro level. In some sense, defining a measure of overall social exclusion can be seen as a problem of multivariate analysis (Poggi 2003).

There is a very substantial body of literature related to the measurement of social exclusion and linked concepts (e.g. Burchardt et al. 1999, Bradshaw et al. 2000, Whelan et al. 2002, Tsakloglou and Papadopoulos 2002, Bourguignon and Chakravarty 2003, Atkinson 2003, Daly 2006, Chakravarty and D'Ambrosio 2006, Poggi 2007, Bossert et al. 2007, Hayes et al. 2008, Amendola and Dell'Anno 2008, Bossert et al. 2013).

According to Atkinson and Marlier (2010), it is possible to distinguish between two different forms of aggregation of the different dimensions (characteristics) of social exclusion. The first combines different characteristics at the individual level (for example, persons or households), which are aggregated over individuals to form an overall index.⁷ The second approach aggregates first across people and then across fields (characteristics) and, thus, utilizes a combination of aggregate indicators (e.g. Atkinson, 2003). These two approaches introduce two methods that measure social exclusion. In fact, the first one, that is predominant in literature, focuses on multiple deprivations at the individual level, which requires microdata sets containing information covering the different relevant domains (i.e. empirical analyses are performed using longitudinal data). The latter at the macro level compares the relative dimension of social exclusion across countries and over time (i.e. empirical analyses are performed using a panel dataset of aggregate variables). This article focuses on the latter approach.

⁷ In this strand of literature, there is a line of research that develops an axiomatic approach to the measurement of social exclusion. Among these, Chakravarty and D'Ambrosio (2006) and Bossert et al. (2007), similar to this research, estimated the social exclusion in EU member states, respectively, for the period 1994–98 and 1994–2001.

Differences between two approaches may exist as a consequence of different nature and methods of micro and macro data collection. In particular, the sizes of response bias⁸ change by aggregating individual-level data (e.g. based on household panel) instead of using data collected according to the methodology of national accounting. Considering that this kind of measurement bias usually occurs when respondents omit or distort - sometimes unconsciously - answers involving sensitive and private aspects of their life (e.g. mental health, material deprivation, social or political participation), thus differences in response bias may be relevant to measure a concept as social exclusion.⁹ In this sense, macro approach may reduce the respondent bias on these sensitive questions looking for these information in manifest variations of observable variables (e.g. crude death rate from suicide, inactivity rate, share of people with low educational attainment). However it may have the cost of a more rough approximation of the dimensions of social exclusion.

This argument strengthen authors' opinion that the two approaches are complementary and interdependent (e.g. macro data are usually based on the characteristics observed at the individual level).

According to UNDP (2011), in this area of research, the literature provides no straightforward algorithm for the construction of a multidimensional social exclusion index for a panel of countries. It rather provides some general guidance and suggestions of good practices in the area (Gordon and Pantazis 1997, Alkire and Foster 2007, Scutella et al. 2009, Atkinson and Marlier 2010). In particular, Atkinson and Marlier (2010) provided an overview of the challenges of measuring social exclusion and identified a minimum set of principles for social exclusion indicators. The first principle is that the set of indicators should be balanced across the different dimensions. It implies that the selection should ensure that all the main areas of concern are covered. The second principle is that the indicators should be mutually consistent and that the weights of single indicators in the portfolio should be proportionate.¹⁰ The third principle is that the portfolio of indicators should be as transparent and accessible as possible. This implies that indicators should be easy to read, examine, and understand. In this research, these guidelines are applied for combining indicators into an overall index of social exclusion. Once the indicators are selected, the key issue of the macro-level index of social exclusion is the method of aggregation.¹¹ An increasing number of studies use principal components analysis (PCA) to construct aggregate indices of social exclusion or relative deprivation. Among these studies, Cavassini et al. (2004) use PCA to construct an index of marginality for regions in Costa Rica; Mangan and Stephen (2007), for regions in Queensland; and Daly et al. (2008) and Lewis and Corliss (2009), for regions and states of Australia. We aim at contributing to this line of research by proposing an advancement of PCA (i.e. multi-way PCA) that is more appropriate for aggregating the (multi-) dimensions of social exclusion with panel data.

⁸ It is a systematic pattern in the difference between the respondents answers to a question and the true values.

⁹ Macro data make available

¹⁰ The term "proportionate" means that the interpretation of the set of indicators is greatly facilitated when the individual components have degrees of importance that are not grossly different.

¹¹ See Njong and Ningaye (2008) for an overview of the methods of aggregation of a multidimensional index. The study analyzes the index of poverty, but these methods are also appropriate for social exclusion.

2. Estimating an Overall Index of Social Exclusion in European Countries

2.1 The Dimensions of the Overall Index of Social Exclusion

There are no commonly shared definitions and measures on overall social exclusion. Therefore, the proposal of a synthetic index of such a complex phenomenon is not only a noteworthy but also a puzzling task. In order to reduce the arbitrariness of measuring social exclusion, we follow the minimum set of principles for social exclusion indicators proposed by Atkinson and Marlier (2010).

Concerning the first principle, we identify five main dimensions of the social exclusion: labour market, poverty, income inequality, education, and health. For each of these dimensions we define, among available data, a balanced number of observed variables for each dimension. As regards the second principle proposed by Atkinson and Marlier (2010), taking into account that method of dimensionality reduction (e.g. MPCA) estimates the loadings for each variables, the number of variables in each dimension reflects authors' idea about which are the fundamental determinants of the social exclusion.¹² With reference to the third principle, to obtain a transparent, reliable and accessible index of social exclusion we use only data collected by "institutional" database. Data published by institution as Eurostat assure that data are collected, elaborated and reported by standard and certified methods.

Although include more variables may increase the completeness of the estimated latent variable (i.e. social exclusion) a significant trade-off exists between inclusiveness of social exclusion definition and reliability of the estimated index (Atkinson and Marlier 2010). Thus, we consider as worthwhile to exclude variables with inadequate coverage of European countries or lacking time span.¹³ In sum, we fine-tune this choice balancing the dimensions of social exclusion between social and economic variables on the basis of the quality and availability of data.

With regard the data source, in the 2001 the Laeken European Council developed a system of social indicators to better measure and understand social exclusion in EU. It proposed a set of common statistical indicators for social inclusion, which allow monitoring in a comparable way to member states' progress toward the agreed EU objectives (Eurostat 2003).¹⁴ In May 2006, the Social Protection Committee endorsed new best practice criteria for indicator design, and

¹² That is because the relevance of the five dimensions is indirectly determined by the numerosness of variables associated to each dimension. In this sense, we include five variables related to (un)employment status because we consider that it is the most important cause of social exclusion. Similarly, poverty, education, and health (with three variables) are more relevant than income inequality (with two variables). Alternative approaches to fix the weights of dimensions of a multidimensional index are recently proposed in literature. For instance, Bellani (2012) proposes a multidimensional deprivation index where the weighs of each dimension are based on their perceived importance by members of alternative reference groups. However this approach is not suitable for macro data approach.

¹³ In particular add other variables (e.g. social life, political participation) may yield (1) problems to balance the five dimensions of social exclusion (in contrast to the first principle); (2) increase the number of missing values and (3) mixing data published by Eurostat with data extracted by different sources making harder to hold the third principle proposed by Atkinson and Marlier (2010).

¹⁴ According to Atkinson et al. (2004, p. 49), Laeken indicators "... are intended to support mutual learning and exchange of good practices in terms of policies. They can also prove useful for illustrating areas where more policy action is needed." They cover four important dimensions of social inclusion (financial poverty, employment, health, and education), which highlight the "multidimensionality" of the phenomenon of social exclusion (Eurostat 2003).

the list was updated in September 2009. The common indicators are currently divided into four strands attached to specific objectives with some indicators being a part of more than one portfolio. The four strands are as follows: (i) the *overarching indicators*; (ii) the *social inclusion indicators*; (iii) the *pensions indicators*; and the (iv) *health and long-term care indicators*. These strands are structured in three sections: *primary indicators* cover the most important elements, thereby leading to poverty and social exclusion; *secondary indicators* are intended to support the lead indicators and describe other important dimensions of the phenomena; and, finally, a set of *context indicators* that have been specified as providing “context” information which helps in interpreting trends in the primary and secondary indicators (Marlier 2007).

In this research, we consider the 2011 updating of *Laeken indicators* both as a theoretical background and as a reference data source for our analysis.¹⁵ From this set of indicators, we extract a subset of nine variables, in order to balance the number of indicators over the five EU's meta-dimensions of social exclusion and data availability. To have a better picture of the social exclusion phenomenon, we also include seven additional variables.

Conclusively, the dimensions of social exclusion are related to the labour market, poverty, income inequality, education, and health. These dimensions are defined by the following variables:¹⁶

- (1) *Employment*: inactivity rate,¹⁷ share of temporary employment, total unemployment rate, young unemployment, and long unemployment*;
- (2) *Poverty*: at-risk poverty before social transfer*, relative median at-risk-of-poverty gap*, and poverty rate elderly people*;
- (3) *Income inequality*: income inequality index of Gini*, income quintile share ratio*;
- (4) *Education*: adult non-participation in education and training*, early leavers from education and training*, and people with low educational attainment*;
- (5) *Health*: infant mortality, suicide rates in the ages 50–54, and suicide rates in the ages 15–19.¹⁸

The sixteen variables are then aggregated into an overall index based on weights that represent the relative importance of sub-indicators in the latent variable so-called “social exclusion.” The list of these variables, arranged according to the five dimensions of social exclusion, is summarized in Table 2. Further details on the data source and definitions are provided in Appendix A.

¹⁵ For methodology and definition details of Laeken indicators, see Eurostat (2010). Data are retrievable from the Eurostat online database. See the appendix A for details.

¹⁶ Laeken indicators are marked by asterisk.

¹⁷ We extract data on activity rate and further convert it into inactivity rate. It is easier to give an interpretable meaning to the overall index of social exclusion. In this way, an increase in inactivity rate means higher social exclusion. The same approach is followed for adjusting one of the two variables related to the education sub-indicators of social exclusion: “Non participant-Life-long learning.”

¹⁸ They can be considered important proxies for the prevalence of mental health disorders in a country. According to the United Nations (2007) mental health disorders, especially depression and substance abuse are associated with 90% of all suicides. These international organizations are a major impediment to the well-being of populations in developed and developing countries. People with these disorders are often subjected to social isolation, poor quality of life, and increased mortality. Moreover, Eckersley and Dear (2002) suggest that an increase in youth suicide reflects a failure of Western societies to provide appropriate sites or sources of social identity and attachment, and the promotion of unrealistic or inappropriate expectations of individual freedom and autonomy. According to this evidence, we also include the rate of youth suicide as a proxy of these factors.

2.2 Three-Mode Principal Components Analysis

To construct an aggregate measure of overall social exclusion, Multiple-way Principal Component Analysis (MPCA) is applied. As argued by Russell et al. (2000), MPCA is a dimensionality reduction technique that allows a much easier interpretation of the information present in the dataset, as it directly takes into account its three-way structure. In particular, the Tucker3 model is the most common model for performing three-way PCA (Pardo et al. 2004).¹⁹

In particular, we propose two models that have not been used earlier within this field of research: the Tucker3 and PARAFAC models.²⁰ The first one was originally applied by Tucker (1966) in psychometrics and currently, there are numerous examples within other disciplines such as chemometrics (e.g. Henrion et al. 1995, Pardo et al. 2004). The second one is the parallel factor analysis (PARAFAC), also named “canonical decomposition,” proposed simultaneously by Harshman (1970) and Carroll and Chang (1970). Kiers (1991) shows that PARAFAC can be considered a constrained version of Tucker3. The generality of the Tucker model made it an often used model for decomposition, compression, and interpretation in many applications.

The fundamental idea behind (multiple and two-way) PCA is to reduce the dimensionality of a data set consisting of a larger number of interrelated variables, while retaining as much as possible the variation present in the original data set. This is achieved by transforming the original data array in a more condensed form.²¹ In this sense, the Multi-way PCA framework (e.g. Tucker, PARAFAC) generalizes the classical (two-way) PCA solution.

The Tucker method is an extension of PCA to N -way data arrays, which preserve the original multiway structure of the data during model development. The Tucker3 method decomposes the three-way data arrays X into three orthonormal loading matrices, denoted as A ($I \times P$), B ($J \times Q$), C ($K \times R$) and the core matrix G ($P \times Q \times R$), which describes the interactions among A , B , and C .²² In sum notation, Tucker3 becomes

$$x_{ijk} = \sum_{p=1}^P \sum_{q=1}^Q \sum_{r=1}^R a_{ip} b_{jq} c_{kr} g_{pqr} + e_{ijk} \quad (1)$$

where the values p , q , and r are the number of components selected to describe the first, the second, and the third mode, respectively, of the data array. The number of factors in each mode is not necessarily the same (i.e. $P \neq Q \neq R$). Each of A , B , and C matrices can be interpreted as a loading matrix in the classical two-way PCA. g_{pqr} denotes the elements (p , q , r) of the core matrix G .

¹⁹ Although PCA could be also applied to a three-dimensional dataset (e.g. Countries x Time x Variables) by transforming data, results could be difficult to interpret, because the information of the three modes can be mixed (Pardo et al. 2004).

²⁰ As best we know, the first application in economics is by Mourao (2008).

²¹ In two-way PCA it is made by estimating the so-called principal components, which the first principal component has the largest possible variance and each succeeding component in turn has the highest variance possible under the constraint that it be uncorrelated with the preceding components. In computational terms, the principal components are found by calculating the eigenvectors (i.e. coefficients of PCs) and eigenvalues (i.e. variances of PCs) of the data covariance matrix.

²² For a comprehensive analysis of this approach, see Kroonenberg (1983, 2008) and Acar and Yener (2009). The empirical analysis of this research is performed by using N-ways toolbox for MATLAB (downloadable at <http://www.models.life.ku.dk/~pih/parafac/chap0contents.htm>) and described by Andersson and Bro (2000).

The largest squared elements of G indicate the most important factors that describe X . The core array is another relevant difference between two-way PCA and the Tucker3 model. In standard two-way PCA, there are no interactions among PCs; whereas the Tucker3 model allows such interactions. All loading vectors in one mode (can) interact with all loading vectors in the other modes, and the strengths of these interactions are provided in the core array g_{pqr} . Statistically, the squared element g_{pqr}^2 reflects the amount of variation explained by the factor p from the first mode; factor q from the second mode; and factor r from the third mode.²³ It is essential to choose a model with a reasonable number of PCs in all directions, because too many PCs result in over-fitting (modeling noise); whereas too few components lead to under-fitting (lack-of-fit) (Jørgensen et al. 2006). The number of factors chosen for the Tucker3 model determines the dimension of the core. We have an optimal dimensionality when the increase in the complexity of the model no longer increases the fit of the model significantly. One can use the development in explained variation as the model increases in dimensionality to indicate the best/simplest model of X . Therefore, the model with the last clear increment in explained variance is usually the one of interest. Table 1 shows the explained variation (sum of squares) of the core for three models (1,1,1), (2,1,2), and (2,2,2). It confirms that only the first component is sufficient to explain the variation of X . Therefore, our choice is definitely the dimension (1,1,1). It implies that we estimate a matrix of loading vector for “country” dimension A (28 x 1), for the “time” dimension B (16 x 1), and for variables dimension (16 x 1).

G (1, 1, 1)			G (2, 1, 2)			G (2, 2, 2)		
Index to elements	Explained var. of the core	Core entry	Index to elements	Explained var. of the core	Core entry	Index to elements	Explained var. of the core	Core entry
g_{pqr}			g_{pqr}			g_{pqr}		
1	1 1 1	100 %	2540.45	1 1 1	96.616%	-2539.15	1 1 1	96.361%
2				2 1 2	3.384%	475.17	2 1 2	3.254%
3				2 1 1	0.000%	0.000	2 2 2	0.078%
4				1 1 2	0.000%	0.000	1 2 2	0.008%
								-2537.66
								-486.68
								72.40
								22.43

Table 1 - Core matrix with dimensionality (1,1,1), (2,1,2) and (2,2,2)

With reference to the most parsimonious model (1,1,1), we see that the core element (1,1,1) indicates interaction among factors A_1 , B_1 , and C_1 . By definition, it explains the highest amount of variation. The second core entry multiplied by the vectors (2,1,2) is a rather small part of the total structural information. It supplies the models with an additional 3.3% of the total explained variation.

Furthermore, Table 1 shows a rough superdiagonality of core matrixes G (i.e. all the elements of the superdiagonal are null). It has two relevant implications. The first one is that it considers it unnecessary to apply orthogonal core rotations in order to estimate a new solution that can be easier interpreted. The second implication is that superdiagonality makes it possible to use the simplest three-way model: the PARAFAC model. With the same symbols of equation (1), the decomposition method in the PARAFAC model is as follows:

$$x_{ijk} = \sum_{p=1}^P a_{ip} b_{jp} c_{kp} + e_{ijk} \quad (2)$$

²³ For example, if the element g_{121} is high compared with other elements of G , then the combination of the first loading vector in mode one, the second in mode two, and the first in mode three is important.

Obviously, the down side of a more complex structure of the Tucker model, in addition to it being less parsimonious, is that the interpretation of findings not only involves the components themselves for all three modes, but also all interactions between these components. Therefore, we consider it worthwhile to also estimate a PARAFAC model that supports both loading interpretations and verifies the robustness of output.

A simple way of assessing whether the model structure is reasonable is by monitoring the distribution of superdiagonal and off-superdiagonal elements of G . The core consistency diagnostic indicates how well the model is in concert with the distribution of superdiagonal and off-superdiagonal elements of the Tucker3 core. According to Bro (1998), the PARAFAC model is appropriate if all the superdiagonal elements are close to one another and the off-superdiagonal elements are close to zero. Figure 1 reports the so-called “core consistency diagnostic” (CORCONDIA) proposed by Bro (1998). It is also a helpful test for determining the right number of components. The analysis using CORCONDIA indicates that only two factors are necessary, because the utilization of more factors leads to a great decrease in the core consistency. In the plot of Figure 1, the circles are the superdiagonal elements of G ; hence, these should ideally be one. The squares are the off-superdiagonal elements that should ideally be zero. Of course, for PARAFAC (1,1,1), the plot is trivial, because we have a scalar; therefore, there are not off-diagonal elements in the plot.

Figure 1 - Core consistency diagnostic PARAFAC

As shown in Figure 1, the core consistency plot of a two-component model is perfect.²⁴ Further, the superidentity is 99.93%. According to this explorative analysis, Table 2 reports the estimated loadings of the matrix C estimated by Tucker3 (1,1,1), (2,1,2), (2,2,2), and PARAFAC (2,2,2).²⁵

²⁴ The core elements on the superdiagonal are close to one, and all off-diagonal elements are zero.

²⁵ Estimated loading associated with variables (third mode in our sample – matrix C) of Tucker3 (1,1,1) is equal up to the fifth decimal of the estimated loadings of PARAFAC (1,1,1). As consequence, the PARAFAC estimates are not reported.

		Tucker3 (1,1,1)	Tucker3 (2,1,2)		Tucker3 (2,2,2)		Parafac (2,2,2)	
		C1	C1	C2	C1	C2	C1	C2
Employment								
EI	Inactivity rate	0.258	-0.258	-0.009	-0.258	-0.013	0.257	0.041
ET	Temporary Employment	0.090	-0.089	-0.093	-0.089	-0.089	0.085	-0.069
EUT	Total Unemployment rate	0.070	-0.070	0.036	-0.070	0.037	0.072	0.051
EUY	Young unemployment	0.153	-0.153	0.032	-0.153	0.035	0.154	0.067
EUL	Long unemployment	0.031	-0.031	0.018	-0.031	0.017	0.032	0.023
Poverty								
PR	At Risk poverty before social transfer	0.208	-0.210	0.051	-0.210	0.054	0.213	0.096
PM	Relative median at-risk-of-poverty gap	0.178	-0.178	0.010	-0.178	0.009	0.179	0.046
PE	Poverty rate elderly persons	0.165	-0.163	-0.089	-0.163	-0.090	0.158	-0.054
Income inequality								
IG	Index of Gini	0.247	-0.247	-0.001	-0.247	0.001	0.247	0.053
IQ	Income quintile share ratio	0.039	-0.039	-0.007	-0.039	-0.007	0.039	0.001
Education								
ENL	Non participant-Life-long learning	0.768	-0.771	0.112	-0.771	0.116	0.776	0.275
ELA	Person with low education attainment	0.271	-0.267	-0.666	-0.267	-0.670	0.233	-0.599
EE	Early leavers from education and training	0.140	-0.135	-0.310	-0.134	-0.317	0.118	-0.282
Health								
HAS	Crude death rate from suicide (50-54)	0.205	-0.202	0.634	-0.201	0.627	0.232	0.656
HSY	Crude death rate from suicide (15-19)	0.061	-0.058	0.152	-0.058	0.144	0.066	0.153
HI	Infant mortality	0.051	-0.051	0.039	-0.051	0.036	0.052	0.046
Explained Variation of X (%)		94.582	97.447		96.879		96.261	

Table 2 -Loadings of Matrix C (Variables)

The interpretation of Tucker3 and PARAFAC models is different. PARAFAC can be interpreted in a similar manner to two-way PCA, whereas the interpretation of Tucker3 results should also take into account the magnitude and sign of the non-zero elements of G. This means that for Tucker3, by comparing loadings from different modes, some attention needs to be paid to the core array, because the magnitude of a direction is given by the core array. For this reason, the signs of C1 loadings estimated by Tucker3 (2,1,2) and C2 of Tucker3 (2,2,2) are opposite to PARAFAC.²⁶ According to this analysis, we conclude that the results are robust with regard to different Tucker3 dimensions and model specifications (i.e. PARAFAC). Although the estimated coefficients and the fitting of the three models are very similar, we choose the Tucker3 (1,1,1) – or equivalently the PARAFAC(1) that yields identical results - to extract the factor loadings of the component matrix C (third mode). In particular, (a) Tucker3 explains a significantly higher variance with regard to PCA (94.6% instead of 80.6%); (b) the indexes based on Tucker3 (1,1,1) and Tucker3 (2,1,2) have very similar estimated values with a correlation coefficient of 99.8%.²⁷ Thus, as simple is better, Tucker3 (1,1,1) is regarded as

²⁶ In particular, the $c1_{ij}$ of Tucker3 (2,1,2) to be compared with PARAFAC coefficients should be multiplied for the (negative) element g_{111} . At the same time, the signs of $c2_{ij}$ of Tucker3 (2,2,2) should be multiplied for the (negative) element g_{212} .

²⁷ See appendix B for further details.

the preferred model in term of trade-off between share of explained variance and model complexity. As concern the relative weights of the dimensions of social exclusion - by aggregating the squared factor loadings estimated by Tucker3 (1,1,1) in Table 2 - we find that education variables account for 68.3% of the total explained variance, employment proxies for 10.4%, poverty measures for 10.2%, income inequality indexes for 6.3% and health indicators for 4.8%.

2.3 An Index of Social Exclusion in European Countries

In this section, the loadings of the vector C of Tucker3 (1,1,1) are employed to calculate a lower dimensional representation of social exclusion. Analogous to two-way PCA, the factor loadings are the correlation coefficients between the 16 proxies of social exclusion and the latent factor. Thus, we estimate the index of social exclusion according to Table 2.²⁸

$$SE_{ij} = 0.258 * EI_{ij} + 0.090 * ET_{ij} + \dots + 0.051 * HI_{ij} \quad (3)$$

where $i = 1, 2, \dots, 28$ (countries) and $j = 1995, 1996, \dots, 2010$.

Figure 2 depicts both trend and growth rate of the European social exclusion index calculated as unweighted and weighted for a population average over the 28 countries of our sample. According to our index, a higher score means a higher level of social exclusion.

Figure 2 - The dynamics of European Index of Social exclusion

Figure 2 reveals a decreasing trend (i.e. negative growth rates) of social exclusion with the exception of the years 2001, 2002 and 2009.

Figure 3 reports the ranking of 28 European countries calculated as normalized averages over the sample period.²⁹ Country rankings in terms of social exclusion are compared with the levels of income inequality (Gini index). The two indexes show a significant positive correlation ($r = 0.62$).

²⁸ In order to estimate SE_{ij} , we substitute the missing values in the 16 variables of eq. (3) with the observations available for the next year(s). Since the missing values are mainly from 1995 to 1997 (529 on 1344 observations), the estimates of the index for these years should be considered cautiously.

²⁹ $Norm.Index_i = [Index_i - \min(Index_i)] / [\max(Index_i) - \min(Index_i)]$, where $Index$ denotes the average of the index of social exclusion or the index of Gini for each country.

Figure 3 - Normalized indexes of Social Exclusion and Gini (averages 1995-2010)

Figure 3 clearly depicts the two measures of income inequality and social exclusion; although correlated for construction (see eq. 3), they quantify quite different phenomena.

Table 3 reports values of the social exclusion index for each of the 28 European countries for several years.³⁰

³⁰ We also perform a sensitivity analysis in order to verify how the overall index of social exclusion changes by reducing the number of the variables included in the dataset. We observe that the index is adequately robust. For instance, the correlation between the index based on the six variables with the highest estimated loadings (i.e. coefficients higher than 0.20 in Table 2) and the index based on the whole sample (16 variables) is equal to 0.98. Again, the correlation between the index based on the ten variables with coefficients lower than 0.20 and the index including all the variable is 0.78. Details on this analysis are available from the authors on request.

	1995	2000	2005	2006	2007	2008	2009	2010	Averages '95-'10	Average Gr. '95-'10
Austria	115.9	110.7	105.5	104.6	104.3	104.0	103.7	103.6	109.2	-0.74%
Belgium	131.2	124.5	121.1	121.3	120.1	119.5	119.7	119.1	123.8	-0.63%
Bulgaria	126.3	126.8	121.3	123.2	123.6	123.8	125.3	125.9	124.7	0.00%
Cyprus	124.8	122.9	120.0	116.5	114.3	112.7	114.1	113.8	119.5	-0.60%
Czech Rep.	107.2	107.3	109.1	108.1	106.1	104.7	106.6	106.0	107.5	-0.07%
Denmark	100.4	99.5	91.9	89.8	91.4	91.8	92.0	91.4	96.1	-0.61%
Estonia	131.8	121.2	118.6	113.3	112.9	113.6	118.0	116.4	121.2	-0.80%
Finland	112.8	107.2	103.7	104.0	101.6	102.3	102.9	100.2	106.4	-0.78%
France	126.9	123.9	120.0	119.8	118.6	117.6	119.1	119.9	122.0	-0.37%
Germany	114.8	110.1	109.7	109.9	110.3	109.0	108.6	108.3	110.7	-0.39%
Greece	137.9	132.7	125.6	126.3	125.0	123.3	123.3	124.4	129.8	-0.68%
Hungary	131.1	125.8	123.7	126.8	121.4	122.5	122.3	122.1	125.0	-0.46%
Ireland	130.7	125.8	120.2	117.4	117.6	116.9	118.4	119.2	123.4	-0.60%
Italy	136.2	128.3	127.2	125.7	124.9	124.2	124.8	124.6	128.9	-0.59%
Latvia	127.9	122.0	121.1	121.9	118.2	125.4	128.6	123.7	124.3	-0.18%
Lithuania	133.6	134.5	126.7	125.1	121.0	123.0	126.9	128.2	130.1	-0.26%
Luxembourg	127.9	115.3	112.1	116.5	117.7	114.4	105.7	110.3	116.9	-0.92%
Malta	136.6	140.1	133.5	133.0	132.4	131.4	130.7	130.5	135.6	-0.30%
Netherlands	112.3	104.2	103.2	101.9	100.5	99.5	100.1	100.7	104.2	-0.72%
Norway	100.7	101.2	98.2	104.1	100.3	98.0	100.3	100.4	100.4	0.00%
Poland	120.4	121.1	124.9	122.0	118.5	117.3	118.9	118.5	121.4	-0.10%
Portugal	142.6	139.2	137.7	136.5	136.5	134.5	133.2	133.4	138.3	-0.44%
Romania	123.7	123.3	130.6	130.6	130.8	128.2	127.4	126.3	126.0	0.15%
Slovakia	111.8	113.9	114.7	112.0	109.2	109.9	112.5	114.6	112.8	0.18%
Slovenia	120.4	118.3	110.1	108.6	107.0	105.9	107.4	106.6	113.8	-0.80%
Spain	142.4	132.4	126.1	125.4	124.0	124.9	127.2	128.3	132.1	-0.68%
Sweden	95.8	96.0	102.3	104.2	102.0	100.0	102.1	99.2	99.2	0.25%
UK	114.3	107.2	99.6	100.1	105.6	106.0	105.2	105.6	105.9	-0.50%
Unw. average	122.8	119.1	116.4	116.0	114.8	114.4	115.2	115.0	118.2	
Unw. Gr. aver.		-0.43%	-0.38%	-0.41%	-0.25%	-0.98%	-0.36%	0.67%	-0.12%	-0.41%

Table 3 - Index of Social Exclusion for each country

Notes: The remaining values ('94, '96, '97, '98, '99, '01, '02, '03, '04) are available upon request from the authors.

According to our index, Scandinavian countries with Netherlands and United Kingdom exhibit the lowest levels of social exclusion between 1995 and 2010. On the contrary, South European countries (Portugal, Malta, Spain, Greece, Italy) and Lithuania exhibit the highest values. In order to provide a rough evaluation of the effectiveness of the European policies that are aimed at reducing social exclusion, we find that since the declaration of Lisbon's Strategy in 2000, the countries with the best performances in promoting social inclusion are Sweden (-3.4%), Romania (-2.4%), Latvia (-1.4%); at the opposite side, we have Slovenia (+10.0%), Denmark (+8.1%) and Cyprus (+7.4%).

3. Social Exclusion and Economic Growth

This section analyses the relationship between social exclusion and economic growth in Europe over the period 1995–2010. Section 3.1 summarizes theoretical explanations that may explain how social exclusion affects growth. Section 3.2 describes the empirical model. Section 4.3 discusses the results.

3.1 What does Theoretical Literature Say?

The relationship between economic disparities and economic development has been explored in many empirical studies, which mainly focus on the relationship between income distribution and growth.³¹ Notwithstanding, it still remains a puzzle in terms of (i) the sign of correlation; (ii) the nature of this relation (short term or long term), and (iii) whether causality runs from economic growth to inequality and/or vice versa.

Various theoretical explanations have been suggested that explain how inequality could affect growth. Ehrhart (2009) classifies this literature into two main strands: (i) political economy explanations and (ii) purely economic explanations.³²

With reference to political economy explanations, the first group of models argues that a greater degree of inequality motivates more social demand for redistribution throughout the political process (e.g. Bertola 1993, Alesina and Rodrik 1994, Persson and Tabellini 1994, Perotti 1996). Typically, the transfer payments and the associated taxation will distort economic decisions, and, through this channel, the inequality would reduce the growth. The second group of models (a.k.a. “socio-political unrest theory” by Barro, 2000) hypothesizes that high economic disparities cause “political instability” (Alesina and Perotti 1994, 1996) and motivate the poor to engage in crime and disruptive activities (Bourguignon 1999). Through these dimensions of socio-political unrest, more inequality tends to reduce overall productivity and economic growth.

With reference to “purely economic” explanations, in the first approach, the hypothesis of a (negative) relationship between inequality and growth is due to the presence of imperfect capital markets (Galor and Zeira 1993, Aghion et al. 1999). This proposition assumes that a more unequal distribution of assets means that an increased number of individuals does not have access to credit and, thus, cannot carry out productive investments. Through this channel, inequality would reduce growth rate. According to the so-called “endogenous fertility approach,” income inequality noticeably reduces the future growth rate because of the positive effect of inequality on the overall rate of fertility (e.g. Becker *et al.* 1990, Galor and Zang, 1997). Thus, a worsening in the inequality jointly generates a rise in the fertility rate and a drop in the rate of investment in human capital, and this reduces the future growth rate of GDP per capita. The third approach claims that a more unequal distribution of incomes results in smaller domestic markets (Murphy et al. 1989). The size of home demand is, thus, too small to generate markets large enough to fully develop local industries or to attract foreign direct investments. Following this approach, inequality reduces growth rate as a consequence of a lower exploitation of the economies of scale and of incentives to foreign direct investment.

³¹ Aghion et al. (1999), Temple (1999), Kanbur (2000), Barro (2000), Banerjee and Duflo (2003), Knowles (2005), and Ehrhart (2009) for a survey of the literature on inequality and growth.

³² See also Tam (2008).

In our view, “political economy explanations” and “purely economic arguments,” originally proposed to explain the relationship between income inequality and economic growth, can also be applied to the wider concept of social exclusion.

In conclusion, our proposal is to analyze the relationship between inequality and growth by means of a broader measure of *social exclusion* referring to the distribution of income, economic, and social opportunities as well as access to several (social) citizenship rights (Marshall 1950).

This wider approach to inequality and growth analysis also finds support in Ehrhart’s (2009) conclusions. According to this survey, only the models of “political instability” and “endogenous fertility” find less controversial data validation. Thus, since the agents’ decision processes are significantly affected by social factors included in the notion of social inclusion, we assume that social exclusion may be much more relevant for affecting political instability and fertility decisions of economic agents than (just) income distribution.

3.2 Empirical Model

The empirical analysis of the relationship between the overall index of social exclusion and economic growth is pursued by different econometric approaches. It aims at verifying whether the theoretical arguments behind the relationship between (income) inequality and growth are empirically validated, irrespective of whatever a wider concept of inequality (i.e. social) is applied.

Two main limitations affect this analysis: (i) the small time dimension of the dataset (1995–2010) and (ii) the structural break of the recent economic crisis (2008–2009). Thus, the analysis on the long-run properties of the relationship between social exclusion and growth should be considered with caution.

The econometric analysis starts by testing logarithms of GDP per capita and the index of social exclusion for the presence of unit roots. We apply Levin et al. (2002) panel unit root tests, which also accounts for the structural break in 2008–2009. Accordingly, we find that both growth rates of GDP per capita (Ggdp) and the levels of social exclusion have not common unit root process.³³ It means that we cannot validate the existence of a systematic co-movement in the long run between social exclusion and economic growth. Thus, if a statistical link between social exclusion and GDP’s growth exists, it may have only short-run characteristics. Subsequently, Granger causality tests that are adapted to the panel structure of datasets are run to determine whether causality runs from social exclusion to economic growth and/or vice versa.

We carry out Granger (1969) causality tests in a context of panel data, to test whether previous changes in one variable help in explaining current changes in other variables.³⁴

³³ We also apply a battery of tests detecting for individual unit root processes so that unit root may vary across cross-sections: for growth rates of GDP per capita we reject the hypothesis of individual unit root processes while for the level of social exclusion index null hypothesis cannot be rejected. As a consequence, we also perform Pedroni’s and Westerlund’s test of cointegration on these two variables. The results have ambiguous findings with a prevalence to support no cointegration. Conclusively, as these diagnostics have low power when the time dimension is small as in our dataset ($T=16$), we opt to investigate on the short run properties of relationship between growth and social exclusion because it is definitely more suitable to our sample size.

³⁴ According to Granger’s test, a stationary variable x is said to Granger cause a stationary variable y if and only if y is predicted better by using the previous value of x , along with the previous values of y itself, rather than by using only the previous values of y .

However, a bi-variate framework without considering other relevant variables may lead to a spurious causality. Therefore, this study adopts a multivariate dynamic autoregressive model. Since the two variables employed in the Granger test should be stationary, we employ the growth rate of GDP per capita (Ggdp) and the level of social exclusion (SE). We also include period dummies for the structural break of the economic crisis (2008 and 2009) and a set of control variables to reduce potential omitted-variables bias. These are some potential causes of economic growth (i.e. foreign direct investment, population being more than 65 years of age, level of GDP, and index of openness at the international trade) and social exclusion (i.e. tertiary school enrolment).³⁵ However, when fixed effects are included in regressions (4) and (5), then several of these control variables become statistically insignificant. As a result, we omit to report these extended specifications later on. Finally, we run two regressions, including cross-country dummies, a dummy for economic recessions (Dummy '08/'09), an index of international trade openness (T_open), for economic growth regression (4), and the gross enrolment rate of tertiary school (Sch3), for social exclusion regression (5):

$$Ggdp_{ij} = \alpha_0 + \sum_{l=1}^m \alpha_l Ggdp_{ij-l} + \sum_{l=1}^m \beta_l SE_{ij-l} + \gamma_1 Dcrisis_i + \gamma_2 T_open_{ij} + c_i + u_{ij} \quad (4)$$

$$SE_{ij} = \alpha_0 + \sum_{l=1}^m \alpha_l SE_{ij-l} + \sum_{l=1}^m \beta_l Ggdp_{ij-l} + \gamma_1 Dcrisis_i + \gamma_2 Sch3_{ij} + c_i + u_{ij}; \quad (5)$$

where Ggdp is the first difference of the logarithm of GDP per capita. The χ^2 (Wald) statistics for the joint hypothesis: $H_0: \beta_1 = \dots = \beta_m = 0$ is the usual test that is used to investigate for the presence of Granger causality. In particular, we fix the length of lags (m) equals to two in order to save degrees of freedom.³⁶

According to redundant fixed effects and the Hausman test for correlated random effects, the best model specification for (4) and (5) is a fixed-effect model that includes cross-country and time dummies. Accordingly, we consider two specifications of the least-squares dummy variable (LSDV) estimator: an LSDV one way with a dummy for 2008 and 2009 and an LSDV with periods and countries fixed effects (LSDV two ways). Since Nickell (1981) demonstrates that the LSDV estimator is not consistent for a finite time dimension, three further econometric strategies are applied to check the robustness of econometric results: (a) Kiviet's (1995) approach to correct (downward) biased estimates in the LSDV estimator with a lagged dependent variable (LSDVC). We report bias-corrected LSDV estimators using the bias approximations in Bruno (2005a), who extends the results by Bun and Kiviet (2003) and Kiviet (1995) to unbalanced panels; (b) linear generalized method of moments (GMM) estimators proposed by Arellano and Bond (1991) - First Differences GMM; and (c) Arellano and Bover (1995) and Blundell and Bond (1998) - System GMM (2 steps)-. Following Roodman (2009a, b), one of the main drawbacks in implementing the dynamic panel GMM estimator is that too many instruments can overfit endogenous variables and fail to remove their endogenous components. Roodman (2009a) suggests two main techniques for overcoming this issue: first, to use only certain lags instead of all available lags for instruments; second, to combine instruments through addition into smaller sets (so-called "instruments collapsing"). We apply both of them: lags 2 (lag 1) through 7 (6) are included for the equation in differences, and lags 1 (lag 0) through 7 (6) are used for the level equation for endogenous (predetermined) variables. In particular, to implement the GMM approach in equations (4) and (5), lagged

³⁵ See appendix A for the sources of data.

³⁶ We start estimating with more lags, but these additional coefficients of lagged variables are statistically significant only in few cases of lagged Ggdp.

variables are treated as endogenous; control variables (trade openness and tertiary school enrolment) are handled as predetermined; and dummies are considered exogenous. To verify the reliability of the GMM estimates, the robust version of Sargan's (1958) test (Hansen J-statistic) is applied to check for the validity of instrumental variables. Since we have an unbalanced panel with gaps, the sample is maximized by using the forward orthogonal deviation (Arellano and Bover 1995) instead of the first-difference approach. Finally, we report standard error estimates computed by Windmeijer's (2005) finite-sample correction for the two-step covariance matrix.

Unfortunately, there is no general consensus in the literature on how to identify the best choice of panel estimator given the dimensions of the dataset used here. In particular, the most utilized dynamic panel estimator with endogenous regressors – the GMM approach – is designed for situations with few time periods (T) and many cross-sectional units (N), but our dataset does not hold this structure as it has only 28 cross-units and relatively many time periods. Thus, in spite of the applied techniques to reduce the instrument set (i.e. limiting the lag depth and “collapsing” the instrument set) dynamic GMM may be still inconsistent as the number of instruments becomes too large with $T=16$.³⁷ Additional source of bias for the estimates calculated by the first difference GMM occurs when the individual series shows a high degree of persistence.³⁸ It is the case of the regression 5 where the coefficient of the lagged index of social exclusion (α_1) is about 0.85.

Conclusively, given the dimensions of our dataset and the high persistence of the index of social exclusion a clear best choice of panel estimator is not available. However, some useful indications to address this issue are provided by Judson and Owen (1999) and Buddelmeyer et al. (2008). These studies run Monte Carlo simulations on a range of different estimators of panel data model with fixed effects in order to compare their biases. Matching the dimension of our dataset to Judson and Owen (1999) and Buddelmeyer et al. (2008) simulations, reveals that the best-performing estimator may be the Kiviet's bias-corrected LSDV estimator.³⁹ However, LSDVC assumes strictly exogenous regressors, thus to extend our analysis to endogenous regressors the first difference GMM and the system GMM can be regarded as the most reliable estimators for regressions 4 and 5, respectively.

³⁷ This concern is more relevant for system GMM because it uses more instruments than the difference GMM, thus estimates based on Arellano and Bover (1995) and Blundell and Bond (1998) should be considered cautiously.

³⁸ In this case the instruments available for the equations in first-differences are likely to be weak (Bond 2002).

³⁹ See Judson and Owen (1999, Table 2) and Buddelmeyer et al. (2008, Table 1).

3.3 Empirical Findings

This section shows the results of regressions (4) and (5) in Tables 4 and 5, respectively. In order to check robustness of findings for each estimator we report estimates based on two different techniques to aggregate the dimensions of social exclusion, i.e. PCA and MPCA.⁴⁰ In general, this test confirms that the findings are robust to the choice of the method to combine the dimensions of social exclusion.

Findings of Table 4 show that, with exclusion of the system GMM model, the Wald tests reject the null hypothesis that the coefficients of social exclusion (β_i) are jointly equal to zero both in MPCA and PCA index. Accordingly, social exclusion Granger causes real GDP growth. Empirical outcomes show a statistically not significant impact of the lagged social exclusion's value on the growth at the time t but a positive effect of social exclusion on the growth rate after two years (with exception of LSDVC and GMM-system). This result means that, in the short run, the hypotheses of negative relationship between level of social exclusion and economic growth - i.e. the above entitled "political instability" and "endogenous fertility" theories - are not empirically supported.

⁴⁰ See appendix B for details on PCA and a comparison of alternative measures of social exclusion. Estimated output based on MPCA (2,1,2) are not reported for the sake of brevity. These results are qualitatively very similar of MPCA (1,1,1).

	LSDV One way		LSDV Two-ways		LSDVC		GMM 1 st -diff.		GMM-sys (2-steps)	
	PCA	Tucker3	PCA	Tucker3	PCA	Tucker3	PCA	Tucker3	PCA	Tucker3
Ggdp_{t-1}	0.342 ^{***} (5.65)	0.383 ^{***} (6.62)	0.420 ^{***} (5.45)	0.454 ^{***} (5.88)	0.511 ^{***} (8.59)	0.553 ^{***} (9.17)	0.375 ^{***} (5.27)	0.456 ^{***} (7.77)	0.446 ^{***} (8.22)	0.491 ^{***} (9.83)
Ggdp_{t-2}	-0.150 [*] (-1.73)	-0.194 ^{**} (-2.22)	-0.143 [*] (-1.59)	-0.196 ^{**} (-2.19)	-0.163 ^{***} (-3.25)	-0.219 ^{***} (-4.53)	-0.076 [*] (-0.75)	-0.159 ^{**} (-2.10)	-0.227 ^{**} (-2.16)	-0.143 [*] (-1.47)
SE_{t-1}	-0.018 [*] (-0.12)	-0.035 [*] (-0.43)	0.073 [*] (0.51)	0.011 [*] (0.14)	0.106 [*] (0.86)	0.028 [*] (0.37)	-0.624 [*] (1.79)	0.118 [*] (0.77)	0.096 [*] (0.42)	0.051 [*] (0.41)
SE_{t-2}	0.364 [*] (2.54)	0.177 [*] (2.31)	0.190 [*] (1.90)	0.13 [*] (1.69)	0.159 [*] (1.34)	0.077 [*] (0.99)	0.297 [*] (1.71)	0.178 [*] (2.13)	0.223 [*] (1.14)	0.01 [*] (0.91)
Dummy ('08/'09)	-5.665 ^{***} (-14.77)	-5.803 ^{***} (-14.81)	--	--	--	--	-4.814 ^{***} (-5.56)	-5.640 ^{***} (-8.93)	-5.585 ^{***} (-8.09)	-6.050 ^{***} (-8.79)
T_{open}	0.095 [*] (8.21)	0.097 [*] (8.14)	0.062 ^{***} (4.16)	0.066 ^{***} (4.26)	0.058 ^{***} (3.90)	0.060 ^{***} (3.98)	0.195 [*] (4.79)	0.181 [*] (3.83)	0.090 [*] (3.12)	0.083 [*] (2.09)
Observ.(countries)	380 (28)		380 (28)		380 (28)		350 (28)		380 (28)	
Fixed Effects	Countries		Countries & Years		Years		--		--	
Robust St.Errors	Cross-section weights		Cross-section weights		Kiviet		Robust ^a		Windmeijer	
Adj-R ²	0.622	0.606	0.686	0.673	--	--	--	--	--	--
Durbin-Wats.	2.153	2.144	2.035	2.031	--	--	--	--	--	--
Wald Test β_i (p-v.)	0.000	0.001	0.001	0.022	0.001	0.090	0.006	0.024	0.131	0.817
GMM diagnostics										
AR(2) (p-value) ^b	--	--	--	--	--	--	0.805	0.368	0.206	0.246
# instruments	--	--	--	--	--	--	22	22	26	26
Hansen test	--	--	--	--	--	--	0.069	0.046	0.134	0.113
GMM instruments for levels (p-value)^c										
Exclud. Gr. ^d	--	--	--	--	--	--	--	--	0.066	0.048
Difference ^e	--	--	--	--	--	--	--	--	0.883	0.980
IV: (p-value)^c										
Exclud. Gr. ^d	--	--	--	--	--	--	0.065	0.039	0.113	0.111
Difference ^e	--	--	--	--	--	--	0.306	0.404	0.526	0.287

Table 4 - Dependent Variable: Growth rate of real GDP

Notes: ***, ** and * denote significant at 1%, 5% and 10% level respectively. Constant term and fixed effects are not reported. The numbers in parenthesis are the t-ratios.

Robust Covariance method known as Panel Corrected Standard Error (Beck and Katz, 1995), is applied to compute standard errors that are robust to heteroskedasticity across cross-sections but not to general correlation of residuals. For LSDVC is applied Kiviet (1995) correction estimator. The consistent estimator chosen to initialize the bias correction is Anderson and Hsiao estimator (1982). We apply xtlsdvc STATA's command (Bruno 2005b). ^a standard errors are robust to heteroskedasticity and arbitrary patterns of autocorrelation within countries. ^b Arellano-Bond test for AR(2) in first differences. ^c Difference-in-Hansen tests of exogeneity of instrument subsets. P-values greater than usual significance levels imply that we can not reject null that the instruments used in GMM and IV parts are valid. ^d Hansen test excluding group. The null hypothesis is that excluded instruments, as a group, are not correlated with independent variables (those which were assumed to be endogenous); ^e the null hypothesis is that the instruments are exogenous.

Taking into account the lack of robustness of these findings, any economic explanation should be interpreted with caution. Tentatively, (1) the statistically not significant impact of the lagged social exclusion's value on growth at the time t provides evidence that the interactions between these two phenomena have to be analysed in long run. It is reasonable to assume that the effects of social exclusion on growth (e.g. through political instability and changes in fertility rate) require a longer time span. (2) The statistically not significant impact of the social exclusion at $t-2$ on growth obtained by LSDVC and GMM-system suggests that not conclusive results are obtained by this analysis. However, if exists a correlation between these two variables, it is positive. In this circumstance two possible reasons may explain this finding. On the one hand, a higher level of social exclusion may have a positive impact upon growth (after two years) as a consequence of the expansion of government expenditures related to social welfare programs (e.g. welfare subsidies, education spending, public housing). Thus, higher level of social exclusion encourages expansionary fiscal policy and it stimulates the economy. On the other hand, a decrease in the level of social exclusion may have a negative impact on the growth because the costs to promote social integration (e.g. distortionary taxation of financing public expenditures) may outweigh the positive benefits of social inclusion in human capital accumulation. These arguments⁴¹ may explain the lack of a negative effect of promotion of social inclusion upon economic growth in the short run. However, the overall effect upon future growth is undetermined because it likely occurs out of the time horizon of the empirical model. According to Herzer and Vollmer (2012), panel cointegration techniques are the most suitable approach to address this issue. By means of this econometric approach, they find that income inequality has a negative long-run effect on income per capita (and thus long-run growth). However, this econometric technique cannot be applied in this study as cointegration method cannot be implemented with short data spans.

Findings of Wald's tests in Table 5 suggest that, with exclusion of the LSDV one way and the GMM system including the PCA index of social exclusion, the null hypothesis cannot be rejected. Accordingly, economic growth does not Granger-cause social exclusion at conventional significance levels.

Conclusively, there is empirical evidence that Granger causality runs one way from the social exclusion to the Ggdp and not the other way. This implies that, in the short run, the promotion of the social inclusion has a negative impact on economic growth instead of vice versa.

⁴¹ They are similar to the transition mechanisms proposed by Sylwester (2000) to explain a negative association between income inequality and economic growth.

	LSDV One way		LSDV Two-ways		LSDVC		GMM 1 st -diff.		GMM-sys (2-steps)	
	PCA	Tucker3	PCA	Tucker3	PCA	Tucker3	PCA	Tucker3	PCA	Tucker3
Ggdp_{t-1}	-0.097 ^{***} (-2.54)	-0.048 ^{***} (-0.83)	-0.095 ^{***} (-2.30)	-0.047 ^{***} (-0.76)	-0.092 ^{***} (-2.12)	-0.021 ^{***} (-0.28)	-0.088 ^{***} (-2.10)	-0.096 ^{***} (-1.51)	-0.146 ^{***} (-2.68)	-0.082 ^{***} (-1.10)
Ggdp_{t-2}	0.049 ^{***} (1.20)	0.167 ^{***} (2.63)	0.015 ^{***} (0.35)	0.107 ^{***} (1.65)	0.018 ^{***} (0.34)	0.100 ^{***} (1.17)	0.007 ^{***} (0.25)	0.092 ^{***} (1.40)	0.074 ^{***} (1.86)	0.072 ^{***} (0.99)
SE_{t-1}	0.921 ^{***} (11.56)	0.740 ^{***} (10.79)	0.907 ^{***} (12.10)	0.738 ^{***} (10.82)	0.925 ^{***} (13.06)	0.793 ^{***} (12.08)	0.890 ^{***} (6.40)	0.848 ^{***} (6.52)	1.000 ^{***} (7.67)	0.833 ^{***} (8.43)
SE_{t-2}	-0.194 ^{***} (-2.48)	-0.014 ^{***} (-0.22)	-0.196 ^{***} (-2.66)	-0.038 ^{***} (-0.58)	-0.162 ^{***} (-2.54)	-0.014 ^{***} (-0.18)	-0.234 ^{***} (-3.45)	-0.033 ^{***} (-0.62)	-0.041 ^{***} (-0.41)	0.042 ^{***} (0.60)
Dummy ('08/'09)	0.728 ^{***} (3.58)	0.229 ^{***} (0.65)	--	--	--	--	0.821 ^{***} (-3.31)	0.739 ^{***} (1.52)	0.952 ^{***} (2.94)	0.731 ^{***} (1.65)
School_3	-0.027 ^{***} (-3.17)	-0.050 ^{***} (-3.63)	-0.012 ^{***} (-1.07)	-0.018 ^{***} (-0.93)	-0.017 ^{***} (-1.00)	-0.009 ^{***} (-0.32)	-0.046 ^{***} (-1.98)	-0.050 ^{***} (-1.67)	0.007 ^{***} (0.61)	-0.024 ^{***} (-0.92)
Observ.(countries)	341 (28)		341 (28)		341 (28)		309 (27)		341 (28)	
Fixed Effects	Countries		Countries & Years		Years		--		--	
Robust St.Errors	Cross-section weights		Cross-section weights		Kiviet		Robust ^a		Windmeijer	
Adj-R ²	0.955	0.974	0.961	0.976	--	--	--	--	--	--
Durbin-Wats.	2.050	2.049	2.043	2.054	--	--	--	--	--	--
Wald Test β_i (p-v.)	0.039	0.030	0.053	0.258	0.100	0.505	0.129	0.198	0.012	0.440
GMM diagnostics										
AR(2) (p-value) ^b							0.764	0.501	0.425	0.388
# instruments							22	22	26	26
Hansen test							0.313	0.422	0.251	0.346
GMM instruments for levels (p-value)^c										
Exclud. Gr. ^d									0.515	0.367
Difference ^e									0.056	0.315
IV: (p-value)^c										
Exclud. Gr. ^d							0.257	0.413	0.257	0.306
Difference ^e							0.783	0.340	0.266	0.561

Table 5 - Dependent Variable: Index of Social Exclusion

Notes: See table 4

The last step of our empirical analysis involves a comparison of these results with the effect of income inequality on the growth rates. In this sense, we aim at investigating whether the EU institutions policy to promote economic growth while looking at the social disparities in preference to income inequality is empirically supported. To perform this analysis, we estimate regressions (4) and (5) substituting the MPCA index of social exclusion with the Gini index (Gini) and the income quintile share ratio (IQ). Table 6 summarizes the results for LSDV one way, LSDVC with period fixed effects and the GMM system.

Findings of Table 6 suggest not to reject the hypothesis of no-Granger causality between income inequality (e.g. Gini index and Income quintile share ratio) and the growth rate of GDP per capita. Thus, we do not find empirical validation of a statistically significant Granger relationship between the measures of inequality and economic growth in the short run. Accordingly, the assumption of this research to analyse the Granger-causality between growth and social exclusion instead of growth and income inequality is empirically supported.

With reference to the results for the inexistent relationship between income inequality and growth, it is not new in literature. As several surveys show the findings of this strand of empirical research are mixed. Many studies observe that inequality reduces economic growth (e.g. Perotti 1993, Alesina and Rodrik 1994, Persson and Tabellini 1994, Clarke 1995, Alesina and Perotti 1996, Deninger and Squire 1998, Kremer and Chen 2002, De la Croix and Doepke 2003, Josten 2003, Knowles 2005, Davis 2007, Pede et al. 2009). Other researches find a positive relationship between inequality and economic growth (e.g. Partridge 1997, Li and Zou 1998, Forbes 2000, Castello-Climent 2004, Nahum 2005). Other ones, similarly to Kuznets (1955), find evidence for a non-linear correlation (e.g. U-inverted shaped), (e.g. Barro 2000, Banerjee and Duflo 2003, Pagano 2004, Voitchovsky 2005, Bengoa and Sanchez-Robles 2005, Barro 2008, Castello-Climent 2010, Charles-Coll 2010), finally, there are studies who find not statistical significant or a non-conclusive nexus (e.g. Lee and Roemer, 1998; Panizza, 2002; Castelló and Doménech, 2002).

Banerjee and Duflo (2003) identify three main sources of bias that allow to explain the inconclusiveness of results in this strand of literature: (a) the measurement errors in data on inequality; (b) the choice of estimator approach;⁴² (c) the non-linearity of the relationship between the level of inequality and growth.⁴³

⁴² Banerjee and Duflo (2003) states that, in general, while the OLS regressions using one cross-section typically found a negative relationship between inequality and subsequent growth, the fixed effect approach yields a positive relationship between changes in inequality and changes in the growth rate.

⁴³ Among these studies, Barro (2000) finds a negative relationship between inequality and growth in the sample of poor countries and a positive relationship in the sample of rich countries.

Estimators	LSDV	LSDVC	GMM-sys (2-steps)	LSDV	LSDVC	GMM-sys (2-steps)	LSDV	LSDVC	GMM-sys (2-steps)	LSDV	LSDVC	GMM-sys (2-steps)
Dep. Variables:	Ggdp	Ggdp	Ggdp	Ggdp	Ggdp	Ggdp	Gini	Gini	Gini	IQ	IQ	IQ
Ggdp_{t-1}	0.422*** (4.99)	0.526*** (4.86)	0.552*** (5.62)	0.421*** (5.10)	0.540*** (5.52)	0.535*** (5.49)	0.046 (0.82)	0.088 (0.90)	-0.002 (-0.04)	0.003 (0.15)	0.012 (0.49)	-0.002 (-0.11)
Ggdp_{t-2}	-0.239 (-1.51)	-0.306** (-2.44)	-0.452 (-1.50)	-0.221 (-1.48)	-0.294*** (-2.81)	-0.421 (-1.51)	-0.003 (-0.03)	-0.034 (-0.23)	-0.030 (-0.40)	-0.001 (-0.04)	-0.017 (-0.45)	-0.005 (-0.19)
Gini_{t-1}	0.136 (0.88)	0.239 (1.46)	0.195 (0.40)	--	--	--	0.482*** (4.10)	0.688*** (4.42)	0.687*** (3.85)	--	--	--
Gini_{t-2}	-0.017 (-0.11)	-0.165 (-1.07)	-0.077 (-0.28)	--	--	--	0.041 (0.34)	-0.014 (-0.11)	0.091 (0.66)	--	--	--
IQ_{t-1}	--	--	--	0.729 (1.13)	0.920* (1.86)	0.774 (0.58)	--	--	--	0.521*** (4.13)	0.699*** (5.94)	0.642*** (4.25)
IQ_{t-2}	--	--	--	-0.372 (-0.58)	-0.837 (-1.62)	-0.714 (-0.65)	--	--	--	0.067 (0.51)	0.012 (0.13)	0.182 (1.28)
Dummy ('08/'09)	-6.323*** (-14.43)	--	-6.281*** (-7.95)	-6.291*** (-14.49)	--	-6.195*** (-9.50)	0.201 (0.69)	--	0.062 (0.20)	0.040 (0.53)	--	0.007 (0.09)
School_3	--	--	--	--	--	--	0.016 (1.18)	0.003 (0.06)	0.012 (0.74)	0.006 (1.26)	0.006 (0.44)	0.003 (0.46)
T_open	0.103*** (4.98)	0.081*** (2.73)	0.081*** (2.73)	0.104*** (5.08)	0.068** (2.07)	0.078** (2.66)	--	--	--	--	--	--
Observ. (countries)	235 (28)	235 (28)	235 (23)	244 (28)	244 (28)	244 (28)	189 (28)	189 (28)	189 (28)	199 (28)	199 (28)	199 (28)
Rob.St.Er.	Cr.sect.w.	Kiviet	Windm.	Cr.sect.w.	Kiviet	Windm.	White D.	Kiviet	Windm.	Cr.sect.w.	Kiviet	Windm.
Adj-R ²	0.604	--	--	0.605	--	--	0.872	--	--	0.879	--	--
Durb-Wats.	2.754	--	--	2.681	--	--	1.938	--	--	2.035	--	--
Wald T. β_i	0.590	0.309	0.887	0.514	0.140	0.756	0.599	0.666	0.904	0.989	0.858	0.960
GMM diagnostics												
AR(2)(p-v) ^b			0.225			0.204			0.728			0.242
# instrum.			26			26			26			26
Hans. Test			0.080			0.112			0.457			0.519
GMM instruments for levels (p-value)^c												
Excl. Gr. ^d			0.080			0.086			0.285			0.431
Difference ^e			0.537			0.469			0.960			0.630
IV: (p-value)^c												
Excl. Gr. ^d			0.114			0.117			0.492			0.626
Difference ^e			0.273			0.238			0.217			0.112

Table 6 - Regressions with indexes of income inequalities

Notes: See table 4.

Similarly, Ehrhart (2009) distinguishes two main reasons behind these controversial results. First of all, the statistical relationship between income inequality and growth may reflect the effect of omitted variables. For instance, Birdsall et al. (1995) sustain that the strong negative correlation is due to the omission of an educational variables (primary and secondary school enrolment rates). Perotti (1996) claims that the negative correlation between inequality and growth is not robust to the inclusion of the variable measuring the share of over sixty-five years of age. Birdsall et al. (1995) and Fishlow (1996) find that inclusion of regional dummy variables makes the income inequality variable insignificant in the growth regression.

The second issue of empirical analyses according to Ehrhart (2009) classification matches with Banerjee and Duflo's (2003) hypothesis that the vast majority of researches studying the impact of income inequality on economic growth does not measure inequality in a consistent manner as a consequence of lacking of comparable data. Knowles (2005) states that studies predating the release of the Deininger and Squire (1996) data set include data of dubious quality.

In this scenario, we check if our results of inconclusive findings can be explained by the same arguments proposed by this literature. In particular checking for Banerjee and Duflo's (2003) hypotheses, we control if the not statistically insignificant relationship between income inequality and growth is due to (a) measurement errors in data – by using two alternative indexes of inequality – (b) it depends on the choice of estimator approach – by using three alternative estimators – (c) it is due to non-linearity – by including a quadratic term of the measures of income inequality in the models. However, the finding of not statistical significance relationship is robust to all these checks.

Conclusively, on the question of whether income inequality is harmful or beneficial for growth, our data comply with evidence of irrelevant short run nexus. In spite of this, we guess that by using a long term perspective this finding may change. As recently shown by Herzer and Vollmer (2012), by using panel cointegration techniques to investigate the long-run effect of income inequality on growth, the effect of the inequality on per-capita income is statistically significant.

4. Conclusions

This article pursues a twofold objective. From a methodological point of view, the article proposes an aggregate overall index of social exclusion, estimated through a relatively new approach to this issue (i.e. MPCA). Atkinson et al. (2002) and Atkinson and Marlier (2010) outline the essential properties of the indicators of the social index (e.g. unambiguous, robust, responsive to policy without being subject to manipulation, consistent with international standards, balanced across the different dimensions, and readily understood by lay members of the community). We attempt to follow their recommendations by using macro data published by Eurostat for 28 European countries over the period 1995–2010.

The rationale behind the proposed index reflects the conviction that the quantification of an overall index of social exclusion is an essential step in assessing the economic relevance of this phenomenon. We consider as helpful a synthetic measure of social exclusion in terms of a communication strategy. In fact, a solo index allows us to overcome one frequent criticism of these non-monetary indicators. It is what they lack that has made GDP a success: the powerful attraction of a single headline figure allowing simple comparisons of socioeconomic performance over time or across countries (Stiglitz et al. 2009). Hopefully, this work will play a role in this scientific debate.

Looking at the country-by-country analysis, for the year 2010, Scandinavian countries (Denmark, Sweden, Finland, Norway) exhibit the lowest levels of social exclusion. On the contrary, some South European countries (Portugal, Malta, Spain) and Lithuania exhibit the

highest values. Looking at time trends, one can see that there has been a decrease in the overall index of social exclusion. This reveals the efficacy of national and European policies to promote social exclusion. In particular, these policies have been very effective in increasing the level of social exclusion between 2000 and 2010 in Slovenia, Austria, Cyprus, Finland. At the opposite side, Sweden, Romania, Latvia, Slovakia and Bulgaria are the only five European countries in which social exclusion increased in the last decade.

The second goal of this article is from a positive economic policy perspective. It consists of analysing the short run relationship between social exclusion and economic growth. In this sense, we provide a rough assessment on the effectiveness of Lisbon's strategy agenda. Initially, we test whether social exclusion causes economic growth or vice versa in the Granger sense. Results show that there is a (not very robust) statistically significant relationship of Granger causality between social exclusion and economic growth but not vice versa. Atkinson and Marlier (2010, p. iii) state that "Promoting social integration and inclusion will create a society that is safer, more stable and more just, which is an essential condition for sustainable economic growth and development"; thus, this research does not find empirical evidence that supports this view in the short run.

In general, this article has sought to extend the embryonic literature on the issues of measurement of the social exclusion and on the relevance of this phenomenon on the economic development in Europe. About the latter issue, it is worth noting that it is partially missing in theoretical literature on the possible mechanisms linking the social exclusion to economic growth. We make a first attempt to fill this gap, drawing on some strands of literature on income inequality. In this perspective, we also find that, in the short run, to promote economic growth looking at the social exclusion in preference to income inequality is empirically supported. However, further research needs to explicate the long run mechanisms through which the social exclusion matters for economic growth.

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Appendix A - Database

Var.	Description	Source [Code]	Mean	Max	Min	St.Dev	Obs
EI	Inactivity rate (100-Activity rate 15 to 64 years)	Eurostat [100-lfsi_act_a]	30.4	42.4	19.3	5.8	407
ET	Percentage of employees with temporary contracts	Eurostat [lfsi_emp_a]	10.8	35.2	1.0	6.7	399
EUT	Unemployment rate, annual average	Eurostat [une_rt_a]	8.2	20.1	1.9	3.9	407
EUY	Unemployment on population aged 15-24	Eurostat [une_rt_a]	17.8	42.5	5.0	8.1	407
EUL	Long-term unemployed (12 months and more) as a percentage of the total active population	Eurostat [tsdsc330]	3.6	12.2	0.3	2.5	405
PR	At risk of poverty rate before social transfers (pensions excluded from social transfers) (cut-off point: 60% of median equivalised income after social transfers)	Eurostat [ilc_li10]	24.9	40.4	0.0	4.0	328
PM	Relative median at-risk-of-poverty gap.	Eurostat [tsdsc250]	21.0	34.8	12.0	4.3	321
PE	At-risk-of-poverty rate of elderly people.	Eurostat [tsdde320]	19.4	52.0	4.0	9.6	320
IG	Gini index. Zero means perfect equality (everyone has the same income) and 100 means perfect inequality.	Eurostat [ilc_di12]	29.2	39.2	20.0	4.2	328
IQ	Income quintile share ratio. The ratio of total income received by the 20% of the population with the highest income (top quintile) to that received by the 20 % of the population with the lowest income (lowest quintile).	Eurostat [tsdsc260]	4.6	7.9	2.9	1.1	335
EE	Early leavers from education and training. Percentage of the population aged 18-24 with at most lower secondary education and not in further education or training	Eurostat [tsisc060]	16.1	54.4	4.1	9.6	368
ENL	(No) Life-long learning (No adult participation in education and training). 100 - the percentage of the population aged 25-64 participating in education and training over the four weeks prior to the survey	Eurostat [100- tsdsc440]	91.2	99.2	67.2	7.0	372
ELA	Persons with low educational attainment. It is defined the percentage of people aged 25 to 64 with an education level ISCED: pre-primary, primary and lower secondary education.	Eurostat[tsdsc430]	31.2	82.2	8	17.7	406
HSA	Crude death rate from suicide and intentional self-harm per 100,000 people, by age group (50-54)	Eurostat [tsdph240]	24.3	100.4	2.6	17.4	401
HSY	Crude death rate from suicide and intentional self-harm per 100,000 people, by age group (15-19)	Eurostat [tsdph240]	7.3	24.4	0.5	4.6	380
HI	Infant mortality rate. It represents the ratio of the number of deaths of children under one year of age during the year to the number of live births in that year. The value is express per 1,000 live births.	Eurostat [demo_minfind] for Luxemb. (OECD)	5.9	22.3	1.8	3.4	446
GDP percap	GDP per capita based on purchasing power parity (PPP). Data are in constant 2005 international dollars. (The growth rates are calculated as log differences and expressed in percentage points.)	WDI [NY.GDP.PCAP.PP.KD]	24787.2	74113	6182	11893	448
T_ope n	Trade openness Index. (Exports of goods and services in current US Dollars+Imports of goods and services in current US Dollars)/GDP at purchaser's prices in current US Dollars.	WDI [(BX.GSR.GNFS.CD+BM.GSR.GNFS.CD)/NY.GDP.MKTP.CD]	99.5	281.4	29.6	43.7	432
Sch3	School enrolment, tertiary (% gross)	WDI [SE.TER.ENRR]	52.2	95.1	9.8	18.2	396

Appendix B - Alternative Estimates of Social Exclusion Index

This appendix compares alternative indexes of social exclusion. They are estimated by Multiway Principal Component Analysis (Tucker, 1966) and Principal Component Approach (Hotelling, 1933). In table 7 we report the five Principal Components with eigenvalues greater than 1. The cumulative variance extracted is 80.6%. In line with missing values imputing method followed by Tucker3 index, we impute missing values to estimate PCA overall index of social exclusion with the observations available for the next year(s) in the country. The overall index is estimated by aggregating PCs following approach of eq. 3.

		PCA 1 st	PCA 2 nd	PCA 3 rd	PCA 4 th	PCA 5 th
Employment						
EI	Inactivity rate	0.273	0.268	-0.028	-0.255	0.309
ET	Temporary Employment	0.094	-0.146	-0.421	0.355	0.149
EUT	Total Unemployment rate	0.272	0.303	-0.126	0.386	-0.098
EUY	Young unemployment	0.286	0.266	-0.131	0.347	-0.122
EUL	Long unemployment	0.285	0.352	-0.135	0.128	-0.133
Poverty						
PR	At Risk poverty before social transfer	-0.010	-0.212	0.318	0.417	0.093
PM	Relative median at-risk-of-poverty gap	0.327	-0.037	0.199	-0.032	-0.315
PE	Poverty rate elderly persons	0.152	-0.292	0.306	0.046	0.059
Income inequality						
IG	Index of Gini	0.326	-0.196	0.312	0.098	-0.016
IQ	Income quintile share ratio	0.335	-0.190	0.308	0.111	-0.084
Education						
ENL	Non participant-Life-long learning	0.299	0.212	0.054	-0.283	0.236
ELA	Person with low education attainment	0.281	-0.262	-0.255	-0.074	0.422
EE	Early leavers from education and training	0.300	-0.254	-0.149	-0.139	0.327
Health						
HAS	Crude death rate from suicide (50-54)	-0.126	0.364	0.280	0.105	0.366
HSY	Crude death rate from suicide (15-19)	-0.169	0.239	0.351	0.239	0.452
HI	Infant mortality	0.169	0.213	0.228	-0.389	-0.205
Explained Variation of X (%)		32.60	18.92	13.63	9.90	5.49
Weights used in PCA		0.405	0.235	0.169	0.123	0.068

Table 7 - Principal Component Analysis

To make easier the comparison among these indexes, we normalise them to have a range [0, 100]. It is obtained by subtracting the minimum value and dividing by the range of the index values across the countries. In symbols we applied the following equation:

$$SE_{ij}^n = 100 * \frac{SE_{ij} - \min(SE_{ij})}{\max(SE_{ij}) - \min(SE_{ij})} \quad (6)$$

Figure 4 shows dynamics of the averages over time of 28 European countries of Tucker(1,1,1) Tucker3(2,1,2) and PCA(5).

Figure 4 - The dynamic of Normalised European Index of Social exclusion

Table 8 reports country average of normalised indexes. In parentheses are reported the countries' rankings.

	MPCA Tucker3 (1,1,1)		MPCA Tucker3 (2,1,2)		PCA (5 Princ. Compon.)	
Denmark	11.868	1	12.213	1	7.129	1
Sweden	17.787	2	18.553	2	17.276	4
Norway	20.049	3	20.904	3	13.497	3
Netherlands	27.196	4	27.486	4	13.319	2
United Kingdom	30.570	5	30.514	5	23.904	5
Finland	31.445	6	32.975	6	32.677	11
Czech Rep.	33.510	7	35.907	7	28.787	7
Austria	36.849	8	38.522	8	24.459	6
Germany	39.584	9	41.371	9	30.277	9
Slovakia	43.617	10	46.336	10	51.631	17
Slovenia	45.399	11	48.135	11	39.276	12
Luxembourg	51.249	12	52.383	12	31.804	10
Cyprus	56.215	13	56.793	13	29.165	8
Estonia	59.432	14	63.402	16	64.962	26
Poland	59.952	15	63.230	15	62.891	25
France	61.042	16	63.078	14	45.177	14
Ireland	63.625	17	65.083	17	44.262	13
Belgium	64.464	18	66.570	18	50.555	16
Latvia	65.267	19	69.294	20	71.661	27
Bulgaria	66.204	20	68.447	19	62.377	24
Hungary	66.728	21	70.329	21	58.451	23
Romania	68.586	22	70.972	22	56.747	20
Italy	73.974	23	74.532	23	54.050	19
Greece	75.694	24	76.761	24	57.036	22
Lithuania	76.426	25	82.271	26	85.765	28
Spain	80.080	26	80.299	25	56.843	21
Malta	86.740	27	85.879	27	50.486	15
Portugal	91.888	28	91.358	28	52.098	18
Average	53.77		55.49		43.45	
Cum.Var. Expl.	94.58%		97.45%		80.55%	

Table 8 - Countries' ranking for MPCA and PCA approaches

As first, figure 4 and table 8 show comparable results. As expected, similarities are significant between the two Tucker3 models while greater differences are with reference to the PCA approach. Correlations among the three indexes are 99.77% between Tucker3 (2,1,2) and Tucker3 (1,1,1); 85.11% between Tucker3 (2,1,2) and PCA(5) and 82.01% between Tucker3 (1,1,1) and PCA(5). The higher variance explained by Tucker3 models validates the benefit to apply a more complex approach to take into account the three-dimensional structure of dataset instead of two-way PCA.

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