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A Regional Perspective on Inequality and Growth in Portugal Using Panel Cointegration Analysis

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Abstract

This paper examines the relationship between inequality and economic growth for thirty Portuguese NUTS3 regions within a multivariate panel framework over the period 1995-2007, using panel cointegration techniques to test for the existence of a relationship between inequality and real GDP per capita. The results point to the existence of a relationship between the variables, where the effect of inequality, measured as the Gini index of the earnings distribution, on per capita output is negative. This negative influence seems to be determined by the behavior of the bottom of the earnings distribution, most likely by dampening investment in human capital, with the results pointing to the coexistence of a positive impact of inequality at the top of the distribution, supporting the incentives argument for the inequality-growth nexus. Additionally, the results confirm the predicted positive relationship between human capital and output, lending support to both the exogenous and endogenous growth models predictions on the importance of human capital for production both as inputs in the production of final goods and in the production of technology. Another interesting result concerns structural funds: we found a negative relation of this variable with regional output. We believe that EU structural funds were a source of Dutch disease for Portugal, which resulted mainly in a lack of external competitiveness of the Portuguese economy due to the excessive specialization in non-tradable goods made possible by the European funds

Keywords: inequality, economic growth, panel unit root tests, panel dynamic OLS

JEL classification: O12

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1. Introduction

In light of the recent concerns over the high income inequality indicators for Portugal, on the one hand, and its dismal growth performance, on the other hand, the discussion on the relationship between inequality and economic performance in the Portuguese economy seems a timely one. According to the European Commission's report Employment and Social Developments in Europe 2011, in 2010 Portugal presented the fourth highest Gini coefficient of income distribution, after Lithuania, Latvia and Spain. Eurostat's data on real GDP per capita growth rates show that Portugal grew faster than the EU15 average between 1996 and 2000, respectively, 3.84% and 2.56% on average, but since 2001 its growth rate has become increasingly lower, respectively, 0.59% and 1.57% for the EU15, on average, between 2001 and 2007, and its long-term growth prospects are quite poor. Additionally, the ongoing economic and public finances sustainability crisis with the associated cut in resources for social support increases the concerns with the long-term implications of potentially growing inequalities. We look at the inequality-output nexus from a regional perspective examining the thirty Portuguese NUTS3 regions over the period 1995-2007 using panel cointegration techniques.

The analysis of the relationship between inequality and economic growth has received a great deal of attention over the years resulting in a large body of theoretical and empirical literature on the subject. At the theoretical level, four views on the relationship between inequality and growth are typically identified (see e.g Duarte and Simões (2011)[1]; Ehrhart (2009); Barro (2000); Aghion *et al.* (1999)). The first view, named by Aghion *et al.* (1999) as "traditional theories", postulates that inequality is good for growth due to its beneficial effects on savings, investment and incentives, with redistribution distorting incentives to save and work. The second view, the fiscal or political economy approach, argues that greater inequality increases public support for redistribution, which leads to higher tax rates on capital accumulation and slower growth of the overall economy. The third view, called the credit markets imperfection approach suggests that barriers to financial markets access when, for instance, insufficient collateral leads some borrowers to forgo high-return projects, constitutes an impediment to capital accumulation and innovation so that higher inequality impedes

[1] A previous and longer version of this paper is available as Duarte and Simões, M. (2009), "Channels of transmission of inequality to growth: A survey of the theory and evidence from a Portuguese perspective", GEMF Working Paper No. 07/2009.

growth. Finally, the political instability approach, the fourth view, defends that inequality amplifies the risks of social and political crisis with the inherent instability reducing investment and growth. However, García-Peñalosa (2008) argues that since output growth, from a supply-side perspective, has four fundamental sources, physical capital, human capital, the labour supply and the level of technology, each of these represents a mechanism that relates the two variables, and so, depending on the source of growth, inequality and growth may be positively or negatively related.

Recent empirical studies have also failed to reach a consensus on the sign of the effect of inequality upon growth, arriving at varied and sometimes conflicting results (see e.g. Dominicis *et al.* (2008)). In general, cross-country studies suggest that there is a negative relationship between initial income inequality and subsequent economic growth, even after controlling for other important growth influences, but using panel evidence leads to different conclusions. Forbes (2000) argues that this can be interpreted as evidence that inequality is detrimental to growth in the long run but not over shorter time horizons. In any case, most studies are interested in identifying a sign for the relationship revealing little about why inequality might impede or be conducive to economic growth (exceptions include Perotti (1996)).

Although the empirical literature on the relationship between inequality and growth is well established the empirical studies on this nexus for the Portuguese economy are rather limited. To the best of our knowledge such an analysis has been performed only by Andrade *et al.* (2011) and Duarte and Simões (2012). This study extends the existing literature by adopting a regional perspective, examining the relationship within a multivariate panel cointegration framework which combines the cross-section and time series dimensions of the data. Individual country studies such as Andrade *et al.* (2011) are often confronted with the problem of a short data span that lowers the power of traditional time series unit root and cointegration tests so adding the cross-sectional dimension can be important to increase the power of such tests. As in Duarte and Simões (2012), we follow Dominicis *et al.* (2008) suggestion that empirical studies on the relationship between inequality and growth should focus on using single-country data at the regional level[2]. This is because the inability of recent empirical studies to reach a consensus is rooted in data quality problems, sample coverage variations and the

[2] "The analysis of the growth inequality linkage on a regional basis may therefore be more informative than the analysis based on worldwide cross-country data." Dominicis, Florax and De Groot (2008), p. 676.

use of different estimation methods that can be mitigated, at least the first two, by taking a regional perspective of a specific country. This study also extends the existing literature by estimating the relationship between inequality and output using the Dynamic Ordinary Least Squares (DOLS) estimator that allows for consistent and efficient estimates of equations with nonstationary variables, thus overcoming the spurious regressions problem. Additionally, the DOLS estimator also deals with the problem of the endogeneity of the regressors.

The remainder of the paper is organized as follows. Section 2 briefly overviews the hypotheses related to the inequality-growth nexus. Section 3 discusses the data, methodology, and empirical results. Section 4 offers some concluding remarks.

2. A brief overview of the literature on the inequality-growth nexus

Is inequality an impediment or a stimulus to economic growth? This is a recurrent debate in economic research. There is by now an extensive and rich literature on the relationship between inequality and economic growth that has been carefully and thoroughly surveyed by Benabou (1996), Perotti (1996), Aghion *et al.* (1999), Barro (2000), Bertola *et al.* (2006), García-Peñalosa (2008), and Ehrhart (2009), among others[3]. In the 1990's, a renewed interest on the subject took place that lead mostly to the development of theoretical models. Due to the scarcity of reliable and comparable data on inequality and due to the existence of poor measures of inequality, empirical literature on the subject lagged behind. The expected sign for the relationship between inequality and growth is a complex matter with early growth models on the subject predicting a positive sign, later ones arriving at a negative relationship and more recent growth models combining both effects. The evidence from empirical studies on the relationship between inequality and growth is also mixed. The cross-country studies usually find that higher inequality slows future growth, but the evidence from the studies that explore panel data is not compelling, pointing to an ambiguous effect of inequality on growth.

[3] Duarte and Simões (2011) provide a survey of the theoretical and empirical literature on two specific mechanisms of transmission from inequality to growth considered by the authors as the most likely to explain the relationship in the Portuguese economy.

According to some theories[4], also designated as traditional or classical theories on the impact of inequality upon economic growth, inequality promotes economic growth because: (i) growth depends positively on the accumulation of physical and human capital and richer individuals save more and thus invest more than poorer individuals and, in addition, there are indivisibilities and large sunk costs in physical capital investments implying a higher concentration of wealth in order to put into practice new investment projects; and (ii) it provides an incentive to the appearance of entrepeneurs/inventors expecting to belong to the wealthier part of the society, thus enhancing growth when innovation is the driving force of long run performance, as well as promoting higher effort by workers and thus efficiency (see e.g. Rodríguez-Pose and Tselios (2010)).

The negative effect of inequality on growth is justified on the basis of three main arguments or mechanisms of transmission. The fiscal or political economy approach literature is based on the interplay of two mechanisms, the political mechanism and the economic mechanism (see e.g. Bertola (1993); Alesina and Rodrik (1994), and Persson and Tabellini (1994)). The political mechanism states that in more unequal societies the median voter will vote for higher levels of taxation and government expenditure. These introduce distortions which will in turn discourage private investment, hindering in this way economic growth – the economic mechanism.

The credit markets imperfection approach, also known as the borrowing constraints in human capital investments channel explains the relationship between inequality and growth based on the analysis of investments in human capital, that foster growth, when there are imperfections in credit markets. Only those individuals that have a high enough initial level of wealth are able to invest in human capital because borrowing is costly and difficult. Thus, an economy with a less unequal wealth distribution will be growth enhancing because it invests more in human capital. For instance, Galor and Zeira (1993) show that in initially highly unequal societies, because of borrowing constraints, fewer individuals are able to invest in human capital, which in turn is detrimental to growth. Galor and Tsiddon (1994), on the other hand, develop a model in which the relationship between inequality and growth depends on the stage of development of an economy. At earlier stages of capitalism, the engine of growth is the accumulation of physical capital and a more unequal wealth distribution is beneficial to

[4] See e.g. Aghion *et al.* (1999) and Barro (2000).

growth because the owners of capital have a higher propensity to save. At later stages of capitalism the engine of growth is human capital accumulation. If there are borrowing constraints on individual's human capital investment due to the existence of capital market imperfections to the borrowers, at later stages of capitalism a less unequal wealth distribution enhances economic growth because individuals decide to invest more in human capital.

The social-political instability channel argues that in more unequal societies individuals are more likely to be involved in activities that act as a disincentive to private investment, such as violent protests against the regime, coups or criminal activities, which in turn hinders capital accumulation and thus growth (see e.g. Alesina and Perotti (1996) and Perotti (1996)). This channel is potentially more relevant to explain the inequality-growth nexus in less developed countries but less likely to occur in a country like Portugal[5].

García-Peñalosa (2008) conducts a survey of the literature on inequality and growth focusing on recent theories especially suitable to explain this relationship in industrialized economies. She argues that both the sign and the direction of causation of the relationship depend on the source of growth under consideration, technological change, human and physical capital accumulation, and changes in the labour supply. The author also points out that "Given the conflicting theoretical predictions, we would like to turn to the empirical evidence in order to assess the relative importance of these various mechanisms." (p. 75), although so far that does not seem to have been made possible.

At the empirical level, a set of studies (see e.g. Persson and Tabellini (1994), Alesina and Rodrik (1994), Alesina and Perotti (1996), Clarke (1995), Perotti (1996), Deininger and Squire (1998), Chen (2003), Balisacan and Fuwa (2003), Bleaney and Nishiyama (2004)) tried to assess the contribution of inequality to cross-country variation in growth rates, after controlling for a number of variables that have been found relevant in the explanation of cross-country growth performance in a large number of empirical growth studies. The estimates of the impact of inequality on economic performance of most of the cross-country studies reveal a negative long-run relationship between the two variables, a result that is usually robust to different sensitivity analyses such as

[5] Duarte and Simões (2009) review in more depth the fiscal policy and the borrowing constraints on human capital investments channels based on the argument that they are the relevant ones to explain the relationship between inequality and growth in an economy like Portugal.

controlling for different inequality measures, different samples and time periods, the presence of outliers, model uncertainty, measurement error, reverse causation, and heteroscedasticity.

However, following the release of the Deininger and Squire inequality dataset[6] that assembled more reliable data with time series information for a large enough group of countries, a number of studies estimated the inequality and growth relationship using panel data techniques (see e.g. Persson and Tabellini (1994), Forbes (2000), Barro (2000), Banerjee and Duflo (2003), Voitchovsky (2005)), trying to uncover in this way a short to medium-term relationship. Contrary to the common message conveyed by most cross-country studies, the panel data evidence is quite diverse, finding either a positive, negative or non-existent correlation between inequality and growth.

In face of the mixed evidence on the inequality-growth relationship provided by empirical studies, Dominicis *et al.* (2008) use the statistical methodology known as meta-analysis, that combines the results of several empirical studies on inequality and growth to give a quantitative summary of the main findings, to survey the empirical literature (Dominicis *et al.* (2008), p.661). The analysis is based on 37 studies that give a total of 407 estimates for the coefficient of the inequality measure[7], restricted to be the Gini coefficient of income distribution[8]. The results show that the variation in the estimates of the income inequality-growth relation are systematically associated with differences in estimation methods, sample coverage and data quality – fixed effects estimates are usually higher, the negative impact is stronger in poorer countries and when the growth period analyzed is longer, and poorer data quality leads to lower coefficient estimates. A final quite relevant suggestion by the authors is that, "(…) it is particularly promising if attention would shift towards samples of regions within one country, or a limited set of countries with similar characteristics, or alternatively with different characteristics to the extent that these can be controlled for in the specification of the regression model." Dominicis *et al.* (2008), p. 678.

Malinen (2012) argues that the appropriate econometric methods to assess the relationship between inequality and growth are panel data time series methods since these do not imply any loss of information and allow for the identification of a long-run

[6] See Deininger and Squire (1996).
[7] See the citations in the previous paragraphs for examples of many studies included in this meta-analysis.
[8] The authors point out that "A remarkable feature is the wide range of the different effect size estimates. Approximately 65% of the estimates are negative (263 out of a total of 407), and approximately 35% (144) are positive." Dominicis, Florax and De Groot (2008), p. 662.

relationship between the variables by taking into account the time series properties of the data. The author applies panel DOLS to a sample of 38 countries with at most 25 annual observations per country to test for a long-run relationship between real GDP per capita and the Gini index of income distribution (as well as investment). He finds evidence that the long-run relationship is negative in the sample as a whole and across different income groups, especially robust for middle-income and rich countries.

The regional dimension of the relationship between inequality and growth has been the focus of some studies. Partridge (1997), Partridge (2005), Panizza (2002) and Frank (2009) analyze evidence for the U.S. states[9]. Partridge (1997) and Panizza (2002) explore the panel structure of the data to uncover the sign of the relationship between inequality and growth. The first study considers ten-years growth episodes during the period 1960-90, finding a positive correlation between the Gini coefficient of income and growth, indicating a positive relationship between inequality and growth, but also a positive correlation between the income share of the middle quintile and growth, indicating a positive relationship between equality, not inequality, and growth. The second study uses state data from different sources covering a wider period (1940-80), Gini coefficients of income and the income share of the middle quintile, analyzes ten and twenty-years growth episodes, and uses different panel data techniques (pooled OLS, fixed effects, GMM), reaching the conclusion that the sign and significance of the estimated relationship changes when different measures of inequality are used and that controlling for outliers, serial correlation and structural breaks substantially changes the results. Partridge (2005) resumes the same issue for the period 1960-2000, estimating a short-term model with data averaged for 10 years periods and a long-run model in which 1960-2000 growth is regressed on the initial values of the explanatory variables, namely the initial Gini coefficient of income and the initial middle-quintile share. The results show a positive correlation between the middle-class share and overall inequality and long-run growth but the short-term results (associated with fixed effects estimations) are ambiguous with no definite conclusion possible. Frank (2009) uses state-level annual data over the period 1929-2000 to determine whether the income share of the top decile Granger causes income growth and to determine the sign of the relationship. The results confirm that inequality causes growth and give evidence that an

[9] Fallah and Partridge (2007) also analyze the inequality-growth relationship for the U.S. but using county data (not state data) for the 1990's. The authors find a positive correlation in urban areas but a negative one in nonmetropolitan areas.

increase in the income share of the top decile reduces future income growth but with regional differences – the relation is stronger for the more densely populated states.

Arbia *et al.* (2005), Ezcurra (2007) and Rodríguez-Pose and Tselios (2010) investigate the link between inequality and regional growth from the perspective of the EU. The first two studies apply spatial econometrics techniques. Arbia, Dominicis and Piras (2005) use regional data at NUTS II level for the period 1977-2002 dividing the sample into non-transition and transition European countries. For the sample without transition countries regions, the effect of income inequality on growth is positive and significant, but being located close to regions with a high level of income inequality is detrimental to a region's growth performance. The same conclusions apply when data for the transition countries regions are introduced indicating that while inequality within a country is growth enhancing, unequal realities in the neighbours have a detrimental impact. Ezcurra (2007) examining 63 NUTS I and II EU regions over the 1993-2002 period and using information from the European Community Household Panel (ECHP) to build income inequality measures, on the other hand, arrives at a negative relationship between a region's income dispersion and growth. Rodríguez-Pose and Tselios (2010) assess the contribution to regional growth of both income and education inequalities, again using data from the ECHP between 1994 and 2001 to compute dispersion measures for both distributions. The results indicate that both income and educational inequality matter for regional growth and have a positive impact upon it, although relatively small.

As far as the analysis of the relationship between inequality and regional growth is concerned, applied studies have not, to the best of our knowledge, derived any evidence from the use of panel cointegration techniques. In fact, the only study we know of that applies them is Malinen (2012) but from a cross country-level perspective. Examples of regional growth analysis that use panel cointegration methodologies can be found in Márquez *et al.* (2010) and Laurin (2012), but focusing on different growth determinants, respectively, factor accumulation and trade. If the series used in applied regional growth studies are non-stationary, the cross-section and panel econometric methodologies applied in earlier studies suffer from the spurious regression problem. The panel cointegration methodology allows us to deal with this problem by accounting for the integration and cointegration properties of the data.

3. Data, methodology, and results

To investigate the relationship between inequality and output we use the following empirical model:

$$ly_{it} = \alpha_i + \delta_i years + \beta_1 lineq_{it} + \beta_2 lh_{it} + \beta_3 lasf_{it} + \beta_4 e\_sector_{it} + \varepsilon_{it} \quad (1)$$

where i=1,…,N for each Portuguese NUTS3 region in the panel and t=1,…,T refers to the time period; *years* denotes an homogenous trend; *ly*, *lineq*, *lh*, and *lasf* are the natural logarithms of real output per capita, an earnings inequality measure (respectively, *lgini*, *lr1050*, and *lr9050*), the average number of years of education of the workforce, total or relative to a certain schooling level (respectively, *lh* – total; *lhs* - secondary, and *lhh* - tertiary), and the ratio of structural funds received relative to GDP, respectively, and *e_sector* is the employment share of either manufacturing or services (respectively, *e_Man* and *e_Ser*). The parameters α_i and δ_i allow for the possibility of region-specific effects and deterministic trends, respectively, and ε_{it} denotes the error term. In total we have thirty regional units[10] (N=30) and thirteen years (T=13) of data, for a total of 390 observations. The panel is balanced.

The choice of the explanatory variables, besides inequality[11], was determined by theoretical predictions, the convenience of a parsimonious specification and the availability of regional data for the Portuguese NUTS3 regions. Human capital, measured in many empirical growth studies by educational attainment, is expected to influence positively regional economic performance not only due to its role as an ordinary input into production but also as a major input in the production of new ideas (see e.g. Mankiw *et al.* (1992); Benhabib and Spiegel (1994; 2005); Sianesi and van Reenen (2003)). The influence of the sectoral composition of regional economic activity has also been considered has a relevant regional growth determinant in previous studies (see e.g. Applica and wiiw (2007), EC (2007), Ezcurra (2007), Cambridge Econometrics (2009)). Although investment ratios have been found to be one of the more robust growth determinants (see e.g. Levine and Renelt (1992), and Sala-i-Martin

[10] Portugal is composed of 30 NUTS3 regions: Minho-Lima, Cávado, Ave, Grande Porto, Tâmega, Entre Douro e Vouga, Douro, Alto Trás-os-Montes, Baixo Vouga, Baixo Mondego, Pinhal Litoral, Pinhal Interior Norte, Dão-Lafões, Pinhal Interior Sul, Serra da Estrela, Beira Interior Norte, Beira Interior Sul, Cova da Beira, Oeste, Médio Tejo, Grande Lisboa, Península de Setúbal, Alentejo Litoral, Alto Alentejo, Alentejo Central, Baixo Alentejo, Lezíria do Tejo, Algarve, Açores, and Madeira.

[11] Our main variable of interest is inequality. However, growth theories provide many compatible explanations for the behavior of output, an issue also known as 'open-endedness of theories' (see e.g. Durlauf *et al.* (2005)). We therefore include in our empirical analysis other explanatory variables selected on the basis of the criteria described in the text.

(1997)) this series is not available at NUTS3 level for Portugal. This problem may be to some extent overcome by the fact that we include the structural funds Portuguese regions received from the EU during the period under analysis as an explanatory variable, at least assuming that they have been used primarily for investment. In any case, structural funds have been found to be a robust regional growth determinant (see e.g. Cappelen *et al.* (2003), Soukiazis and Antunes (2006), Puigcerver-Peñalver (2007)).

3.1. Data

Data for this study refers to five variables: real GDP per capita, an earnings inequality measure, average years of schooling, a measure of the regional productive structure, and the ratio of structural funds received relative to GDP. Annual data from 1995 to 2007 were obtained mainly from the Regional Accounts of the Portuguese National Statistics Office (*INE*) and the Personnel Records (*QP*[12]) database[13], a rich Portuguese dataset with detailed and comprehensive information on workers and firms, which is the result of an annual compulsory survey conducted by the Ministry of Solidarity and Social Security *(*MSSS*)* where firms are required to provide information about their workers on items such as monthly compensation, highest schooling level attained, age, tenure and monthly hours worked[14]. The length of the period selected was dictated by data availability on regional accounts. Given the short time span of the data we compile it within a multivariate panel data framework.

Regional real output per capita corresponds to per capita GDP in year 2000 euros computed using data from the Regional Accounts of *INE*[15] and the Department of Prospective and Planning (DPP) of the Ministry of Environment[16]. The data concerning the regional productive structure, measured as the regional employment share in manufacturing or the services sectors was also obtained from *INE*. The ratio of structural funds received relative to GDP was computed with data from *INE* and the *Financial Office for Regional Development* (*IFDR*). Structural funds refer to the amount

[12] *Quadros de Pessoal*, in Portuguese.
[13] Data provided by GEP-MSSS.
[14] *QP* database does not include public sector workers and the self-employed.
[15] Regional accounts for the Portuguese economy at the NUTS3 level are available, on a comparable basis, from 1995 until 2010 (predicted values). We chose to limit our analysis to the 1995-2007 period in order not to capture potential effects of the 2007-08 financial crisis.
[16] We thank Natalino Martins from DPP for providing us data on Gross Value Added in year 2000 Euros that allowed us to compute regional price indexes. See also Martins and Barradas (2009).

received, per year, by each Portuguese NUTS3 region from FEDER, the European regional development fund, relative to regional GDP[17].

To measure earnings inequality and average years of schooling[18] of the workforce at the regional level in Portugal we use data from the *QP* database. Earnings inequality[19] measures are computed based on the analysis of two variables: average earnings[20] of the employees working full time and the number of employees distributed according to the economic activity of firms and the level of education of employees. The methodology used to assess earnings inequality at the regional level is the Lorenz dominance analysis (see e.g. (Fields (2001)). For each region we constructed weighted distributions of earnings from which we compute the inequality measures to be used in the estimations: the Gini index (Xu (2004)), and the ratios between percentiles 90% and 50% (*r9050*) and between percentiles 10% and 50% (*r5010*), that provide information on the impact of inequality in different parts of the distribution, top and bottom, respectively. Previous studies show that the relation between inequality and output might depend on inequality in different parts of the distribution, top or bottom, with inequality at the top fostering faster growth due for instance to the incentives argument, and inequality at the bottom reducing the growth rate because it leads to less investment in human capital (see e.g. Voitchovsky (2005) and García-Peñalosa (2008))[21].

Table 1 presents the mean values over the period 1995-2007 of the dependent and explanatory variables for each NUTS3 region. The regions appear in the table by descending order of mean real GDP per capita. The richest region on average is Grande Lisboa, the region where the capital city is located, and the top fifteen positions of the real GDP per capita ranking are occupied by regions located along the coast, on the

[17] Structural funds are given in 10^3 euros and GDP in 10^6 euros.
[18] See Fidalgo *et al.* (2010) for details on the construction of the variable average years of total schooling of the workforce at the regional level. Average years of schooling of the workforce by schooling level, secondary and tertiary, were taken from Cardoso and Pentecost (2011), who also use data from *QP* database.
[19] "Wage income is the main source of personal and household income, and hence its distribution has major implications for inequality. A large literature has hence examined the evolution of the distribution of labour earnings, and documented that in the last two decades of the 20th century a number of industrialised countries experienced a substantial widening in the earnings distribution." García-Peñalosa (2008), p. 61.
[20] This is a monthly average reported each year referring to the month of October. We consider average full earnings of the employees that performed complete working hours during the month of October.
[21] According to García-Peñalosa (2008), p. 75: "Inequality has two effects on the growth rate, a positive incentive effect, in line with the traditional literature, and a negative opportunity-creation effect operating through the constraints on human capital investment that it imposes on poor individuals. Greater inequality is hence conducive to growth if it occurs at the top of the distribution, and detrimental if it occurs at the bottom."

mainland, and also by the Portuguese islands, Madeira and Açores. These richer regions also tend to present higher values for the educational attainment variables, the manufacturing and services employment shares, and to some extent structural funds received. Grande Lisboa is almost three times richer than the poorest region, Tâmega, and slightly more than two times richer than the average value for Portugal. A quick inspection of the data indicates that the richest regions tend to register less earnings inequality, especially at the bottom of the earnings distribution. Considering the Gini index and the ratio *r1050*, only five out of the fifteen richest regions are in the first half of the earnings inequality ranking, where inequality is higher. With the ratio *r9050*, seven out of the fifteen richest regions are in the first half of the inequality ranking, and another three are in the top positions of the second half of the ranking. This evidence points also to possible differences concerning the link between inequality and output depending on the earnings groups affected.

[Insert Table 1 here]

3.2. Methodology

Our empirical investigation of the relationship between inequality and growth in the Portuguese NUTS3 regions is conducted in two steps. First, we test for the order of integration of the variables. Second, we estimate the relationship coefficients for the period 1995-2007 using the dynamic OLS (DOLS) panel cointegration estimation methodology[22].

Unit root testing in panels is quite recent (see e.g. Baltagi (2005)). In what follows, we provide a short description of some tests that assume cross-sectional independence and of a test proposed by Pesaran that assumes the presence of cross-sectional dependence. Afterwards, we also provide a short description of a stationarity test.

The simplest model for testing panel unit roots was a mere replication of a time series unit root test, based on the following equation:

$$y_{i,t} = \delta_0 + \alpha \cdot y_{i,t-1} + \mu_{1,t}, \quad i = 1, \ldots, N, \; t = 1, \ldots, T, \; \mu_{i,t} \sim iid(0, \sigma^2) \qquad (2)$$

[22] The analysis of cointegration does not allow us to identify the direction of causality, an issue that has always been associated with the discussion of the relationship between inequality and growth (see e.g. Gobbin and Rayp (2008)). However, since we only have thirteen annual observations the lag structure that would be necessary to carry out such an analysis would further restrain the degrees of freedom in an already small sample. As more data becomes available it will be possible to develop a causality analysis.

without or with a constant term (see Quah (1994) and Quah (1996)), where the null is H_0: $\alpha=1$ against the alternative H_a: $\alpha\leq 1$. Levin and Lin (1992) and Levin and Lin (1993) add individual and time specific components to equation (2), so that:

$$\mu_{i,t} = \delta_1 \cdot t + \eta_i + \nu_t + \varepsilon_{i,t},\ \varepsilon_{i,t} \sim iid(0,\sigma^2) \qquad (3)$$

where the correction for serial correlation is done in the same way as that employed in traditional ADF time series test. The authors also prove that the panel tests statistics have limiting Normal distributions.

The major limitation of these tests is the exclusion of heterogeneity by imposing a value of α common to all individuals ($i=1,...,N$). Im *et al.* (2003)[23] (henceforth IPS) propose a test that allows for heterogeneous individual behavior. Supposing that t_a ($i=1,...,N$) are the t-statistics associated with each cross-section estimation then $Z = \sqrt{N} \cdot \frac{t_\alpha - \mu}{\sigma} \sim N(0,1)$, $t_\alpha = \frac{1}{N} \cdot \sum_{i=1}^{N} t_{\alpha_i}$, $\mu = E[t_{\alpha_i}]$, and $\sigma^2 = Var(t_{\alpha_i})$. The null, H_0, is that all series contain a unit root against the alternative, null is H_a, that some of the individuals are stationary.

Harris and Tzavalis (1999) (henceforth HT) propose a unit root test in a fixed effects panel model, with or without time trend, taking T, the time dimension, as fixed. This test is appropriate for small T data samples, i.e. with few time series observations. The authors demonstrate the convergence property of the desirable statistic $\sqrt{N} \cdot (\alpha - 1 - B) \to N(0,C)$, with $B = \frac{-3}{T+1}$, $C = \frac{3 \cdot \left(17 \cdot T^2 - 20 \cdot T + 17\right)}{5 \cdot (T-1) \cdot \left(T+1\right)^3}$, and by appropriate standardization we have the usual $N(0,1)$ distribution. However this HT test assumes homogeneity, which might constitute a problem. The null, H_0, corresponds to the usual hypothesis of a unit root for all the data.

All these tests assume cross-section independence which goes against our perception that individuals, often regions and/or countries, are interdependent. Pesaran (2007) proposes a method to solve this problem of cross-section dependence that arises due to unobserved common factors. The equation corresponding to this Cross-Section Augmented Dickey-Fuller (CADF) for each individual is given by:

$$\Delta y_{i,t} = \delta_i + \alpha_i \cdot y_{i,t-1} + c_i \cdot \overline{y}_{t-1} + d_i \Delta \overline{y}_t + \grave{o}_{i,t}, \text{ where } \overline{y}_t = \frac{1}{N} \cdot \sum_{i=1}^{N} y_{i,t} \qquad (4)$$

[23] First published in June 1995.

Pesaran suggests to proxy the common factors with the individual mean ($\bar{y}_t$). We omitted lags of the dependent variable to solve eventual serial correlation problems for each individual. The cross-sectionally augmented IPS (CIPS) is given by:

$$CIPS = \frac{1}{N} \cdot \sum_{i=1}^{N} t_{\alpha i} \tag{5}$$

where the appropriate critical values are calculated in Pesaran (2007).

Hadri (2000) proposes an adaptation of the KPSS LM test for time series (Kwiatkowski *et al.* (1992)) to panel data. Suppose that the Data Generating Process can be represented by:

$$y_{i,t} = \delta_i \cdot t + r_{i,t} + \grave{o}_{i,t}, \ \grave{o}_{i,t} \sim iid(0, \sigma_{\grave{o}}^2) \tag{6}$$

where $r_{i,t}$ is a random walk, $r_{i,t} = r_{i,t-1} + \mu_{i,t}$, $\mu_{i,t} \sim iid(0, \sigma^2)$. Under the null hypothesis of stationarity around a level or a trend ($\delta_i \neq 0$), $\sigma_\mu^2 = 0$, or $\frac{\sigma_\mu^2}{\sigma_{\grave{o}}^2} = 0$, against the alternative $\frac{\sigma_\mu^2}{\sigma_{\grave{o}}^2} > 0$. The LM statistic is obtained by $LM = \frac{\sum_{i=1}^{N} \sum_{t=1}^{T} S_{i,t}^2}{\hat{\sigma}_{\grave{o}}}$, $S_{i,t} = \sum_{h=1}^{t} \hat{\grave{o}}_{i,t}$, assuming $\hat{\sigma}_{\grave{o}}^2$ is a consistent estimator of $\hat{\sigma}^2$.

In the presence of unit root variables Kao and Chiang (2001) demonstrate that the OLS estimator has a considerable bias in normal samples and that, in general, the FMOLS (fully modified OLS) estimator does not improve the OLS estimator. The authors conclude that the DOLS (dynamic OLS) estimator performs better than the other two estimators even if the limiting distribution of the regressors is the same for the DOLS and FMOLS estimators. Both the FMOLS and the DOLS estimators are supposed to correct for endogeneity and serial correlation. The Kao-DOLS approach is a derivation of Saikkonen (1991) for panel data where the observations for each unit are corrected using estimators of the long-run conditional variances and it is supposed that there is no cross-sectional dependence. Mark and Sul (2003) propose to weight the observations of each individual with an estimate of the asymptotic covariance matrix. Their approach requires a two steps DOLS. Both the Kao and Chiang (2001) and the Mark and Sul (2003) estimators consider fixed effects in the estimation of the cointegration regression, and so the latter authors argue that the expression "dynamic least squares dummy variable" is more accurate to designate these estimators although, for simplicity, they use the more commonly used designation DOLS.

For a homogeneous estimation, the β_{DOLS} coefficients are obtained from regression (7):

$$y_{i,t} = \alpha_i + x_{i,t}^{'} \cdot \beta + \sum_{j=-q}^{q} c_{i,j} \cdot \Delta x_{i,t+j} + \dot{v}_{i,t} \qquad (7)$$

with $\dot{v}_{i,t} = v_{i,t} + \sum_{|j|>q} c_{i,j} \cdot \varepsilon_{i,t+j}$, where ε has the usual i.i.d. properties.

Mark and Sul (2003) propose a homogeneous cointegrating vector where the presence of individual-specific fixed effects and individual-specific time trends is allowed for and a common time effect is used to control for cross-sectional dependency.

In summary, in this study we use new generation panel data methods, specifically panel time series econometric approaches, as in for instance Eberhardt and Teal (2011) and Eberhardt and Teal (forthcoming). These are especially suitable for empirical growth studies that use macroeconomic data that are usually non-stationary and thus growth regressions suffer from the spurious regression problem, as we have previously pointed out. If however the series are cointegrated, it is possible to establish long-run equilibrium relationships in the data by using the appropriate panel time series methods. We thus start by testing for the presence of unit roots. In all the panel unit root tests applied we demean the variables to take into account the possibility of cross-section dependence, following the suggestion of Pesaran (2007), and we also apply the Pesaran test (CADF) to account for this specific problem (see Pesaran (2007)). Since the results point to non-stationarity of the data we apply the DOLS panel time series method to estimate the different regressions, which assumes parameter homogeneity. Based on the fact that we are carrying out a study with regional data for Portugal, it is reasonable to assume homogeneity of the regression parameters as a more suitable hypothesis than that suggested by Brock and Durlauf (2001) concerning the adequacy of adopting the opposite assumption, that of parameter heterogeneity, which in our opinion would be completely justifiable only if we were using cross-country data. Nevertheless, we tested the Pedroni (2001) hypothesis of parameter heterogeneity. The homogeneity hypothesis was rejected only in the cases where the estimated regression includes one to two exogenous/explanatory variables. In all the other estimations with more than two exogenous/explanatory variables the homogeneity hypothesis was rejected. These results justify the estimation of homogeneous cointegrated DOLS parameters.

3.3. Results

Tables 2-5 present the results of the different panel unit root tests described in the methodology section for both the levels and the first differences of the variables, all in logs. For the IPS test we have removed the individual means in order to control for cross-sectional correlation. In this case, the two relevant statistics to determine whether the series have a unit root are the "t" and "Z" statistics without trend and with trend. The null hypothesis admits that all series contain a unit root and the alternative that some fraction of individuals/regions is stationary. The "Z" statistic of the HT test considers as the null that the entire sample has a unit root. The "t" statistic of the CADF test considers as null that all individual values have a unit root against the alternative that some fraction of individuals are stationary. The constant is present in the first, and in the second with the trend. The Hadri "Z" test considers as the null hypothesis the stationarity of all individuals against the alternative that some fraction of individuals has a unit root. We use a Bartlett kernel with one lag to control for serial correlation.

[Insert Tables 2-5 here]

As for the results concerning the variables in levels (see Tables 2-5), according to the IPS unit root test results we can conclude: (i) that all the inequality variables (*lgini*, *lr1050*, and *lr9050*) are stationary; (ii) that *lhh* is stationary around a trend; and (iii) for the presence of a unit root in all the other variables. These results are confirmed by the HT test, although in this case without a trend for *lhh*. The results for the CADF unit root test lead to the rejection of the null hypothesis of a unit root for all the inequality variables, as in the former tests, and also for the variables *ly* and *e_Ser* around a trend. The results for the Hadri panel unit root test allow us to reject the null hypothesis of stationarity for all the variables. Concerning the variables in first differences, the IPS and HT tests allow us to reject the presence of a unit root for all variables. The results with the CADF panel unit root test do not lead to the rejection of the presence of a unit root for the variables *dlhs*, and *dlhh*. The results with the Hadri panel unit root test reject the null hypothesis of stationarity for *dlh*, *dlhs* and *dlhh*. Based on the previously described results for the different panel unit root tests we consider that the hypothesis that all variables in levels have a unit root is a reasonable one, notwithstanding some of these results are somewhat contradictory.

Given that the panel unit root tests indicate that each variable is integrated of order one, the dynamic OLS (DOLS) technique is used to estimate the relationship between the variables in equation (1) over the period 1995-2007. We used two different methodologies in our DOLS estimations, that of Kao and Chiang (2001) and that of

Mark and Sul (2003). The latter considers four different behavioral hypotheses: "no time trend, no common time effect; "no time trend, with common time effect"; "heterogenous time trend, no common time effect"; and "heterogenous time trend, common time effect". In all four estimations using the Mark (2003) methodology that we carried out the results did not seem to make economic sense. For instance, we got negative and significant coefficients on human capital and substantial differences in coefficient values across regressions corresponding to the different behavior hypothesis. We thus present only the results obtained with the Kao (2001) DOLS estimation methodology. Tables 6-9 present the results of panel cointegration estimation of the regression equation using the DOLS estimation methodology and considering different inequality measures. To carry out the estimations we used the procedure *xtdolshm* for Stata 11.0 (see Amadou (2010)) assuming one lead and one lag.

[Insert Table 6 here]

Table 6 exhibits the results of panel cointegration estimation using the Gini index as the earnings inequality measure. Regressions 1-2 consider average years of total schooling, *lh*, as the human capital measure. Regression 3 consider that the relevant schooling level for production is secondary schooling and so the proxy used for human capital is average years of secondary schooling, *lhs*. In regression 4 it is considered in addition the contribution of tertiary schooling by introducing simultaneously average years of tertiary schooling, *lhh*, as an explanatory variable. In each case, we test for the relationship between output and the earnings inequality measure, the human capital variable(s), the homogeneous trend 'years', structural funds, *lasf*, and the influence of the regional productive structure by introducing the employment shares of the two major sectors of activity, combined or one at a time, depending on the statistical significance of the results.

The results in Table 6 point to a negative relationship between earnings inequality and output when measured as the Gini index of the employees' earnings distribution. The estimated coefficient of *lgini* is in all regressions negative and statistically significant so that an increase in inequality reduces output.

As far as the other explanatory variables are concerned, the human capital regressors are positive and statistically significant in all the regressions presented in Table 6, lending support to both the exogenous and endogenous growth models predictions on the importance of human capital for production both as inputs in the production of final goods and in the production of technology, although our empirical model does not allow

us to distinguish between the two mechanisms. An indication of the importance of human capital as an input in the production of technology is the fact that when we consider the separate influence of *lhs* and *lhh* the estimated coefficient on average years of tertiary schooling, the educational attainment level potentially more important for this task, is positive and statistically significant, although quantitatively lower than the *lhs* coefficient. The trend variable 'years' has an unambiguous negative and statistically significant sign, which is not surprising since we know from national/aggregate data that output growth stopped in 1998 and the stagnation in terms of trend output began in 2002.

Another interesting result is the negative and statistically significant coefficient of the structural funds variable in all regressions. Although we should expect a negative influence running from output to structural funds transfers (richer regions receive less funds), since some years after the beginning of these transfers it is natural to register a reduction in output differentials, the presence of a negative relation over the period 1995-2007 points to the null impact of these transfers in the reduction of output differentials[24]. As Cappelen *et al.* (2003), 641, point out: "to get the most out of the support [EU structural funds], this needs to be accompanied by policies that improve the competence of the receiving environments, for instance by facilitating structural change and increasing R&D capabilities in poorer regions". As far as the specific situation of Portugal is concerned, the negative sign for the coefficient of the structural funds variable might be related to a Dutch disease phenomenon. The Dutch disease concept was created by the The Economist (1977) and the seminal papers of Corden and Neary (1982) and Corden and Neary (1982), according to which the exploitation of natural resources leads to a decline in the manufacturing sector due to the resulting currency appreciation that makes manufacturing goods exports less competitive. Since then, other explanations for this decline, besides the exploitation of natural resources, have been envisaged, such as foreign aid funds proposed by Rajan and Subramanian (2011) and Verbeke (2007). The same arguments could be exploited for the Portuguese case in terms of the impact of European structural funds. Net transfers from the EU to Portugal

[24] Soukiazis and Antunes (2006) analyze regional convergence, absolute and conditional, in the 30 Portuguese NUTS3 regions over the period 1991-1999 using fixed effects panel techniques. Their results point to a positive and statistically significant growth impact of structural funds only in the coastal areas, although quantitatively very small. Puigcerver-Peñalver (2007) focus on regional growth in forty-one Objective 1 NUTS2 regions (GDP per capita below 75% of the Community average) from 1989 until 2000, dividing it according to the two programming periods defined for the transfer of structural funds. She finds a stronger and clearer impact on the growth rates of Objective 1 Regions during the first programming period, 1989-1993.

reached 2.6% of GDP in 1996-99, decreasing to 1.5% in 2004-07 and 1.3% in 2008-10. The majority of the European aid funds was used to build infrastructures, that is non-tradable goods. In general, we believe that public investment has a positive effect on long-run output, as evidenced in Aschauer (1989) and Easterly and Rebelo (1993). However, we must not forget that the estimated effect is often small or negligible and depends to a great extent on the empirical methodologies used (see e.g. Turrini (2004)). For the period 1960-2005, Afonso and St.Aubyn (2008) conclude that Portuguese public investment did not crowd out private investment but, at best, the effect of public investment was not much different from zero, with a negative total rate of return for Portugal (the same applied to Austria, Finland, Greece and Sweden). We thus believe it is possible to argue that EU structural funds were a source of Dutch disease for Portugal, which resulted mainly in a lack of external competitiveness of the Portuguese economy due to the excessive specialization in non-tradable goods made possible by the European funds.

Finally, in what concerns the regional productive structure, there is clear evidence of a positive relation between the importance of the services sector and output. As for manufacturing, although positive the estimated coefficient on the manufacturing employment share variable is not statistically significant.

[Insert Table 7 here]

The results of panel cointegration estimation using the ratio of percentile 10^{th} relative to percentile 50^{th} as the earnings inequality measure are presented in Table 7 since previous studies show that the relation between inequality and output might depend on inequality in different parts of the distribution, top or bottom. In this case we consider specifically the relation between inequality at the bottom of the distribution and output. An increase in the *r1050* variable corresponds now to a decrease in inequality, so that the confirmation of the previous results with the Gini index requires a positive estimated coefficient for this inequality measure. The consecutive regressions follow the same logic as those presented in Table 6. Regressions 5-6 consider average years of total schooling, *lh*, as the human capital measure, while regression 7 considers that the relevant schooling level for production is specifically secondary schooling measured as average years of secondary schooling, *lhs*. Regressions 8-9 consider the separate and additional contribution of average years of tertiary schooling, *lhh*.

From the inspection of the results presented in Table 7 we can see that the estimated coefficient of *lr1050* is in general positive and statistically significant. To be more

precise, it is positive in five regressions out of the seven estimated regressions in this case, pointing to, as was the case with the results with the Gini index, to a negative relation between inequality and output. The same conclusions also apply to the estimated coefficients for the remaining explanatory variables.

[Insert Table 8 here]

Table 8 presents the results of panel cointegration estimation using the ratio of percentile 90th relative to percentile 50th as the earnings inequality measure. In this case, we are controlling for a different sign of the relation between inequality and output at the top of the earnings distribution. An increase in the *r9050* variable corresponds once more to an increase in inequality, as in the Gini index case. Again the different regressions can be grouped in three main types according to the human capital variable(s) considered. Regressions 10 considers average years of total schooling, *lh*, as the human capital measure; regression 11-12 consider average years of secondary schooling, *lhs*; and, finally, regressions 13-14 consider in addition average years of tertiary schooling, *lhh*. In each of the three groups of regressions, we tested for the long-run relationship between output, the human capital variable(s), the inequality measure, *lasf*, the structural funds variable, the homogeneous trend, and the influence of the regional productive structure.

The results in Table 8 point to an interesting change in terms of the evidence on the relationship between earnings inequality and output relative to the results obtained with Gini index and the ratio of percentile 10th relative to percentile 50th. When the *lr9050* variable is used as the inequality measure, in only two out of the eighteen regressions the estimated coefficient is negative and statistically significant, the result that at first sight would confirm the evidence we got based on the other two inequality measures. In the vast majority of the regressions the estimated coefficient of this inequality measure is positive and statistically significant indicating that an increase in inequality is beneficial to growth. This result however does not invalidate the previous ones. In fact, they are compatible. As shown by Voitchovsky (2005), different mechanisms can be working at the same time so that an increase in inequality at the bottom of the earnings distribution will decrease output, while an increase in inequality at the top has the opposite effect. Inequality is good for growth since richer individuals have higher marginal propensities to save, thus leading to higher capital accumulation, and because it creates an incentive to belong to the wealthiest classes of society leading to higher working effort, both resulting in faster growth, but this only happens when inequality

increases at the top of the distribution. If inequality increases at the bottom of the distribution, it will prevent an increasing number of poor but talented individuals to invest in human capital, and in this way hurt growth. The conclusions regarding the remaining explanatory variables do not change when compared to the results presented in Tables 6 and 7.

[Insert Table 9 here]

Based on the evidence from the two previous tables, 7 and 8, in Table 9 we present the results of the estimations considering both inequality ratios at the same time (regressions 15-20) in order to confirm separate influences from different parts of the earnings distribution. Again the regressions are grouped according to the human capital variables considered. The results point to a negative influence of inequality at the bottom of the earnings distribution, together with a positive influence of inequality at the top. As for the other explanatory variables, the same conclusions as in Tables 6-8 apply.

4. Conclusions

Although Portugal ranks as one of the most unequal countries (in terms of income distribution) relative to other EU member states, inequality has hardly ever been considered as an explanatory factor in the studies that aim at explaining economic growth in the Portuguese economy, including studies devoted to the analysis of regional growth. To fill this gap, this paper investigated the relationship between inequality and economic growth, using panel cointegration estimation techniques. The empirical analysis exploits annual panel data for 30 Portuguese NUTS3 regions and the sample period runs from 1995 to 2007. Various panel unit root tests were conducted to demonstrate that the data variables are integrated processes with unit roots. The DOLS (dynamic OLS) panel cointegration estimation technique was then employed to estimate the regression equation.

Overall, the results do not reject the existence of a significant relationship between the earnings distribution and per capita output, as predicted by theoretical models, but the sign of the relationship varies according to the earnings groups affected by changes in inequality, so that different transmission mechanisms might be working simultaneously. Specifically, we found that an increase in inequality reduces per capita output most likely by dampening investment in human capital, but this negative sign

coexists with a positive relationship between inequality at the top of the earnings distribution and real GDP per capita, supporting the incentives argument. Additionally, the results confirm the predicted positive relationship between human capital and output. Another interesting result concerns the relationship between structural funds and output that point to a negative relationship. However, all these results are conditional on the sample period 1995-2007 so caution is required when trying to extend them to a longer time span. In fact, we may be capturing a specific short-term mechanism that relates the series during this period and so we should not try to infer the longer-term cointegration properties of the series.

The main policy implication emerging from this study is that policies aimed at raising the earnings and levels of education of the poorest are in order, even if there is evidence of a positive relationship between inequality and regional output due to inequality at the top end of the earnings distribution. When designing redistributive policy interventions it should be taken into account that the results in this study also have implications for the design of education programs since the negative output impact of inequality at the bottom of the earnings distribution is most likely explained by borrowing constraints to investments in human capital.

A better understanding of the overall effect of inequality on growth involves future research on the empirical assessment of the different channels through which inequality in different parts of the distribution may influence the growth process. These channels are surely important from a redistributive policy design perspective.

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Table 1: Mean values of the dependent and explanatory variables by NUTS3 region, 1995-2007

Nuts 3	*y* (euros)	*Gini*	*r1050*	*r9050*	*h* (years)	*hh* (years)	*hs* (years)	*asf*	*e_man*	*e_ser*
Grande Lisboa	19127.37	0.2497	0.630	1.989	9.332	2.693	5.532	13.40%	20.74%	78.61%
Alentejo Litoral	14817.01	0.2539	0.677	0.206	7.435	0.979	4.890	2.66%	24.89%	59.29%
Madeira	13142.20	0.1957	0.737	1.641	7.652	0.897	5.242	3.17%	26.55%	60.84%
Grande Porto	12255.04	0.2032	0.752	1.730	8.098	1.722	5.028	12.30%	33.16%	64.60%
Algarve	12148.51	0.1867	0.751	1.431	7.819	0.996	5.448	4.95%	19.61%	71.50%
Pinhal Litoral	12116.57	0.1465	0.732	1.366	7.644	1.196	5.024	4.28%	41.42%	43.34%
Baixo Mondego	11859.80	0.1846	0.703	1.622	8.005	1.543	5.156	7.16%	21.52%	58.76%
Baixo Vouga	11099.29	0.1876	0.746	1.572	7.622	1.272	4.885	10.10%	40.53%	42.50%
Lezíria do Tejo	10803.73	0.1976	0.675	1.532	7.605	1.049	4.958	2.54%	33.10%	54.90%
Médio Tejo	10239.54	0.1690	0.743	1.528	7.665	1.058	5.121	2.83%	30.36%	49.94%
Beira Interior S	10171.63	0.1843	0.706	1.503	7.363	1.025	4.708	1.89%	20.77%	42.02%
Oeste	10142.93	0.1811	0.722	1.530	7.512	0.942	5.009	3.00%	29.03%	46.33%
Entre Do. e Vo.	10022.86	0.1881	0.844	2.121	6.983	0.896	4.388	2.31%	61.80%	30.96%
Alentejo Central	9982.19	0.1912	0.658	1.475	7.353	0.959	4.736	6.04%	27.44%	61.04%
Açores	9884.40	0.2382	0.688	1.868	7.498	0.835	4.935	9.40%	23.01%	61.40%
Alto Alentejo	9847.82	0.2088	0.666	1.622	7.180	0.930	4.546	3.09%	24.30%	58.95%
Pen. de Setúbal	9423.67	0.2178	0.690	1.482	8.160	1.292	5.585	4.45%	31.69%	65.37%
Baixo Alentejo	9237.03	0.2295	0.661	1.795	7.354	1.042	4.695	2.96%	17.49%	63.62%
Ave	9156.87	0.1582	0.828	1.499	6.830	0.748	4.381	5.28%	62.37%	30.63%
Cávado	8974.61	0.1697	0.800	0.145	9.332	0.896	4.975	3.48%	46.76%	41.94%
Dão-Lafões	7964.62	0.1979	0.687	1.599	7.496	1.098	4.826	4.51%	25.62%	37.58%
Cova da Beira	7722.96	0.2641	0.570	1.927	7.730	1.030	4.551	1.85%	30.37%	38.44%
Beira Interior N	7601.45	0.2007	0.644	1.561	7.422	1.173	4.544	2.78%	17.11%	37.36%
Pinhal Interior	7558.24	0.3138	0.580	2.513	8.303	0.703	4.364	1.10%	20.37%	27.64%
Alto Trás-os-M.	7399.45	0.2816	0.514	0.186	7.944	1.119	4.919	2.95%	12.08%	38.14%
Douro	7336.89	0.2336	0.588	0.173	7.334	1.057	4.340	3.50%	14.31%	40.25%
Minho-Lima	7327.80	0.1824	0.687	1.468	7.565	0.888	5.255	3.32%	32.92%	40.29%
Pinhal Interior	6788.50	0.1884	0.749	1.610	6.904	0.767	4.220	2.95%	33.16%	35.58%
Serra da Estrela	6441.23	0.2072	0.751	1.744	6.792	0.901	3.799	0.91%	27.51%	38.48%
Tâmega	6396.96	0.1706	0.817	1.533	6.534	0.528	4.011	7.66%	54.02%	29.82%
Portugal	9899.71	0.2060	0.700	1.679	7.616	1.075	0.480	4.56%	30.13%	48.34%

Notes: The NUTS3 regions appear by descending order of mean real GDP per capita (*y*) values.

Table 2: IPS panel unit root test results

	Demean		Trend			Demean		Trend	
	T	Z	t	Z		t	Z	t	Z
ly	-1.38	0.60	-2.03	-2.37***	***dly***	-3.49***	-7.55***	-3.21***	-6.61***
lgini	-2,30***	-3,82***	-3.03***	-6.51***	***dlgini***	-4.32***	-9.08***	-4.45***	-9.40***
lr1050	-2.41***	-4.48***	-2.79***	-5.62***	***dlr1050***	-4.07***	-8.69***	-4.17***	-8.93***
lr9050	-2.67***	-5.43***	-2.61***	-5.54***	***dlr9050***	-4.31***	-9.14***	-3.89***	-8.61***
lh	-1.15	2.01	-1.57	-0.76	***dlh***	-3.54***	-7.93***	-4.30***	-9.86***
lasf	-1.54	0.32	-0.84	4.95	***dlasf***	-3,89***	-8,12***	-2,71***	-4,73***
e_Man	-1.30	1.02	-2.01	-2.74***	***de_Man***	-3,71***	-8,20***	-3,84***	-8,62***
e_Ser	-1.25	1.48	-2.22	-4.00***	***de_Ser***	-3,43***	-7,77***	-3,46***	-7,85***
lh	-1.00	3.24	-2.04	-2.74***	***dlh***	-3,86***	-8,28***	-4,21***	-8,98***
lhs	-0.71	5.26	-1.43	1.68	***dlhs***	-4,09***	-8,92***	-3,94***	-8,78***
lhh	-1.62	-0.12	-3.10***	-6.85***	***dlhh***	-4,60***	-9,41***	-4,75***	-9,84***

Notes: (***), (**) and (*) indicate statistical significance at the 1%, 5% and 10% levels, respectively.

Table 3: HT panel unit root test results

	Demean	Trend		Demean	Trend
	Z	Z		Z	Z
ly	0.30	0.64	***dly***	-22,28***	-7,46***
lgini	-7,69***	-26,97***	***dlgini***	-26,57***	-14,40***
lr1050	-10.60***	-9.28***	***dlr1050***	-26.82***	-14.76***
lr9050	-17.78***	-12.64***	***dlr9050***	-34.40***	-20.01***
lh	0,44	0.56	***dlh***	-20,57***	-10.55***
lasf	2.15	-0.77	***dlasf***	-24,97***	-10,11***
e_Man	1.63	0.28	***de_Man***	-21,81***	-8,82***
e_Ser	1.50	-1.50*	***de_Ser***	-20,13***	-8,33***
lh	1.61	-2.40***	***dlh***	-22,59***	-11,45***
lhs	3.93	-0.21	***dlhs***	-22,97***	-8,56***
lhh	-1.66**	-8.83***	***dlhh***	-27,59***	-14,38***

Notes: (***), (**) and (*) indicate statistical significance at the 1%, 5% and 10% levels, respectively.

Table 4: CADF panel unit root test results

	Demean	Trend		Demean	Trend
	t	t		t	t
ly	-2.02	-3,00***	***dly***	-3,03***	-3,21***
lgini	-2.25**	-2.03	***dlgini***	-2,21**	-2.4
lr1050	-2.13*	-2.24	***dlr1050***	-2.49***	-2.46
lr9050	-2.40***	-2.34	***dlr9050***	-2.36***	-2.52
lh	-1.90	-1.78	***dlh***	-2,14*	-2.22
lasf	-1.04	-2.42	***dlasf***	-2,33**	-2.19
e_Man	-1.49	-2.56	***de_Man***	-2,66***	-2.52
e_Ser	-2.04	-2,99***	***de_Ser***	-2,87***	-2,73**
lh	-1.29	-0.98	***dlh***	-1.45	-2.21
lhs	-1.00	-1.81	***dlhs***	-1.87	-2.16
lhh	-2,10*	-1.93	***dlhh***	-2,23**	-2.43

Notes: (***), (**) and (*) indicate statistical significance at the 1%, 5% and 10% levels, respectively.

Table 5: Hadri panel unit root test results

	Demean	Trend		Demean	Trend
	Z	Z		Z	Z
ly	14,99***	10,90***	***dly***	1,31*	6,17***
lgini	5,84***	4,39***	***dlgini***	-1.79	1,39****
lr1050	8.81***	4.05***	***dlr1050***	0.01	3.33***
lr9050	4.16***	3.63***	***dlr9050***	-2.27	2.47***
lh	9,87***	9,57***	***dlh***	2.57***	2.17**
lasf	16,90***	10,40***	***dlasf***	-0.48	6,59***
e_Man	17,24***	11,45***	***de_Man***	-0.24	4,67***
e_Ser	15,84***	8,78***	***de_Ser***	-0.61	2,46***
lh	13,61***	9,12***	***dlh***	2,73***	3,94***
lhs	16,67***	11,36***	***dlhs***	2,74***	2,73***
lhh	17,11***	5,08***	***dlhh***	0.75	2,72***

Notes: (***), (**) and (*) indicate statistical significance at the 1%, 5% and 10% levels, respectively.

Table 6: DOLS estimates with the Gini index as the inequality measure

Dependent variable: *ly*

	1	2	3	4
lh	0.501 (8.09)	0.480 (7.04)		
lhs			0.628 (6.47)	0.433 (4.25)
lhh				0.232 (4.19)
years	-0.012 (4.43)	-0.010 (3.17)	-0.014 (3.27)	-0.029 (4.67)
lgini	-0.058 (1.94)	-0.201 (5.95)	-0.126 (3.93)	-0.165 (5.20)
lasf	-0.014 (5.79)	-0.018 (6.39)	-0.019 (7.72)	-0.023 (9.20)
e_Man	0.499 (1.81)			
e_Ser	1.395 (4.61)	1.238 (3.80	1.051 (3.40	0.836 (2.75)
Wald(k)	411.68	399.22	208.18	366.88
Regions	30	30	30	30
Years	1995- 2007	1995- 2007	1995- 2007	1995- 2007
Observations	390	390	390	390

Notes: t-statistics values are reported in parentheses. All coefficients are statistically significant at least at the 10% level. The Wald(k) is a $\chi^2(k)$ statistic for the null of the regression coefficients. The DOLS methodology was applied in panel cointegration estimations using the procedure *xtdolshm* for STATA 11.0 (see Amadou (2010)) based on the GAUSS code from Kao, C. and Chiang, M. H.: 2002, "Nonstationary Panel Time Series Using NPT 1.3 A User Guide", Center for Policy Research, Syracuse University. Leads & lags=1.

Table 7: DOLS estimates with the ratio 10/50 as the inequality measure

Dependent variable: *ly*

	5	**6**	**7**	**8**	**9**
lh	1.151 (18.65)	0.545 (7.93)			
lhs			0.675 (6.93)	0.408 (4.49)	0.460 (4.50)
lhh				0.260 5.33	0.273 4.88
years	-0.039 (17.04)	-0.015 (4.83)	-0.018 (4.36)	-0.034 (6.25)	-0.038 (5.96)
lr1050	0.901 (14.57)	0.447 (6.27)	0.355 (5.27)	0.117 (2.01)	0.506 (7.55)
lasf	-0.027 (11.10)	-0.015 (5.47)	-0.017 (6.53)	-0.018 (8.12)	-0.020 (7.70)
e_Man	-0.576 (2.26			0.495 (1.95	
e_Ser		1.224 (3.74)	1.023 (3.30)	0.959 (3.36)	0.761 (2.49)
Wald(k)	1051.94	337.73	221.13	335.73	289.50
Regions	30	30	30	30	30
Years	1995- 2007	1995- 2007	1995- 2007	1995- 2007	1995- 2007
Observations	390	390	390	390	390

Notes: t-statistics values are reported in parentheses. All coefficients are statistically significant at least at the 10% level. The Wald(k) is a $\chi^2(k)$ statistic for the null of the regression coefficients. The DOLS methodology was applied in panel cointegration estimations using the procedure *xtdolshm* for STATA 11.0 (see Amadou (2010)) based on the GAUSS code from Kao, C. and Chiang, M. H.: 2002, "Nonstationary Panel Time Series Using NPT 1.3 A User Guide", Center for Policy Research, Syracuse University. Leads & lags=1.

Table 8: DOLS estimates with the ratio 90/50 as the inequality measure

Dependent variable: *ly*

	10	**11**	**12**	**13**	**14**
lh	0.455 (7.41)				
lhs		0.699 (8.17)	0.721 (7.38)	1.018 (11.22)	0.490 (5.40)
lhh				0.366 (7.56	0.246 (5.08
years	-0.012 (4.44)	-0.021 (5.74)	-0.019 (4.65)	-0.064 (12.14)	-0.036 (6.66)
lr9050	0.131 (3.93)	0.239 (7.54)	0.089 (2.43)	0.162 (5.18)	0.202 (6.64)
lasf	-0.014 (5.88)	-0.015 (6.65)	-0.021 (8.09)	-0.027 (12.34)	-0.018 (8.08)
e_Man	0.626 (2.31)	0.620 (2.4)		0.408 (1.74)	0.696 (2.77)
e_Ser	1.452 (4.80)	1.215 (4.20)	0.982 (3.13)		1.023 (3.57)
Wald(k)	394.19	315.01	267.67	592.11	412.20
Regions	30	30	30	30	30
Years	1995- 2007	1995- 2007	1995- 2007	1995- 2007	1995- 2007
Observations	390	390	390	390	390

Notes: t-statistics values are reported in parentheses. All coefficients are statistically significant at least at the 10% level. The Wald(k) is a $\chi^2(k)$ statistic for the null of the regression coefficients. The DOLS methodology was applied in panel cointegration estimations using the procedure *xtdolshm* for STATA 11.0 (see Amadou (2010)) based on the GAUSS code from Kao, C. and Chiang, M. H.: 2002, "Nonstationary Panel Time Series Using NPT 1.3 A User Guide", Center for Policy Research, Syracuse University. Leads & lags=1.

Table 9: DOLS estimates with the simultaneous influence of the ratios 10/50 and 90/50

Dependent variable: *ly*

	15	**16**	**17**	**18**	**19**	**20**
lh	0.479 (7.82)	0.539 (8.01)				
lhs			0.727 (8.51)	0.803 (8.41)	0.522 (5.77)	0.580 (5.77)
lhh					0.256 (5.25)	0.266 (4.82)
years	-0.013 (5.01)	-0.016 (5.40)	-0.023 (6.36)	-0.026 (6.27)	-0.040 (7.35)	-0.044 (7.07)
lr1050	0.111 (1.79)	0.541 (7.67)	0.138 (2.33)	0.542 (8.10)	0.279 (4.70)	0.673 (10.06)
lr9050	0.144 (4.25)	0.143 (3.73)	0.259 (8.07)	0.271 (7.44)	0.242 (7.60)	0.248 (6.86)
lasf	-0.014 (5.68)	-0.015 (5.49)	-0.014 (6.35)	-0.016 (6.17)	-0.017 (7.68)	-0.019 (7.36)
e_Man	0.542 (1.99)		0.510 (2.01)		0.498 (1.97)	
e_Ser	1.423 (4.71)	1.234 (3.81)	1.169 (4.06)	0.978 (3.19)	0.930 (3.26)	0.727 (2.40)
Wald(k)	385.15	362.32	308.39	323.19	395.73	385.03
Regions	30	30	30	30	30	30
Years	1995-2007	1995-2007	1995-2007	1995-2007	1995-2007	1995-2007
Observations	390	390	390	390	390	390

Notes: t-statistics values are reported in parentheses. All coefficients are statistically significant at least at the 10% level. The Wald(k) is a $\chi^2(k)$ statistic for the null of the regression coefficients. The DOLS methodology was applied in panel cointegration estimations using the procedure *xtdolshm* for STATA 11.0 (see Amadou (2010)) based on the GAUSS code from Kao, C. and Chiang, M. H.: 2002, "Nonstationary Panel Time Series Using NPT 1.3 A User Guide", Center for Policy Research, Syracuse University. Leads & lags=1.

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