The relationship between social transfers and poverty reduction: A nonparametric approach for the EU-27

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Abstract

Poverty reduction is one of the most important goals of developed welfare states. Among others, expenditures on social transfers consist in an important instrument for that goal. Using a nonparametric panel data model for the EU-27, for the period 2003-2013, this article studies the role of social transfers in combating poverty, allowing for the detection of saturation points and structural breaks between periods. Controlling for demographic and economic conditions specific to each country, we find a positive relationship between poverty gains and social transfers for values below 27% of GDP. Above that saturation point, expenditures on social transfers describe a situation of total inefficiency. The results suggest that Finland, Hungary and Luxembourg are the most efficient countries in reducing poverty via social transfers, whereas Greece, Portugal and Spain are the least efficient among the EU-27. We also conclude that social transfers are less efficient in the crisis period (2008-2013).

JEL Classification: C14, D61, H55 and I3.
Key Words: Poverty measure, social transfers, efficiency and nonparametric econometrics.

1 Introduction

Since its inception, the European project has gradually expanded its competences in the social ground. However, it wasn't until 2010 with the adoption of the Europe 2020 strategy, that the European Commission has set a quantitative target for poverty reduction and social exclusion, in line with the Lisbon Treaty, which marked the fight against poverty as an explicit goal of the European Union (EU). With more than 120 million people in the EU at risk of poverty and social exclusion, the Europe 2020 headline target is to reduce that number by at least 20 million by 2020[1]. Several measures can be implemented by the governments and related institutions to combat poverty. Among them, social transfers can be regarded as one of the most important instruments. Our objective is to assess the efficiency of such social benefits in reducing poverty among the European countries.

In the 2001 Leaken European council, the European Commission in partnership with the Members States developed a set of social indicators with the goal of measuring progress towards social inclusion objectives. The assessment of monetary poverty is one of the indicators in this list. Defining the poverty threshold at 60% below the median income of the population, before and after social transfers, this relative poverty measure was labeled as the *at risk of poverty rate*. Using a panel data for the EU-27 between the period 2003-2014, we construct an efficiency measure of poverty reduction (*poverty gains*), which is the difference between the risk of poverty rate before and after social transfers, to evaluate the role of each national social protection system in alleviating poverty levels. In order to do so, we propose a nonparametric model for the relationship between poverty gains and social expenditures that will allow us to detect structural breaks and implicitly inefficiency differences between countries and periods.

The remainder of the paper is organized as follows. Section 2 sets the stage by reviewing the findings of recent studies on the relationship between social expenditures and poverty rates. Section 3 describes the nonparametric methodology to be implemented. Section 4 presents the data, the relevant descriptive statistics and a nonparametric analysis between poverty gains, social expenditures and GDP per capita conditional upon year. Section 5 reports the results for the nonparametric panel data model. A systematic comparison between the nonparametric specifications and its parametric counterparts is also provided. Section 6 concludes.

2 Social transfers, poverty reduction and income distribution

Poverty reduction and income distribution are not the only goals of social transfers systems (Atkinson, 1997). Other benefits and tax allowances are also part of the social transfers system and their role is not reduced to poverty alleviation. However, despite the diversity of the social protection schemes among countries, and in particular among the European Union countries (Longford and Nicodemo, 2010), a significant amount of

[1] An overview of 2020 national targets can be accessed here: `http://ec.europa.eu/europe2020/pdf/annexii_en.pdf`

social transfers and benefits are designed for poverty reduction and income distribution purposes.

A commonly approach adopted to measure the efficiency in poverty reduction is based on the aggregate poverty gap (Beckerman, 1979). Defined as the sum of the individual income shortfalls from the poverty line, a system is called effective if the aggregate poverty gap is greatly reduced by the social transfers. Using the European Union *Statistics on Income and Living Conditions* in 2007, Longford and Nicodemo (2010) characterize the social transfer systems by their potential and effectiveness, and compare the countries for different definitions of the poverty threshold. The potential of a social transfers system is defined by comparing its total (or average) amount with the total (or average) shortfall of household income (poverty gap), and its effectiveness as the fraction of the total amount of social transfers that resulted in the reduction of the poverty gap. The authors conclude that the Scandinavian and central European countries are the most effective whereas the Mediterranean countries are the least effective.

According to Duffy (2010), there are significant differences in the amounts spent on social protection in the European countries. This means that establishing a comparison on their effectiveness in alleviating poverty is not an easy task. Using a descriptive analysis of the poverty rates, before and after social transfers, the author concludes that the impact of social benefits in the 2000's was weaker in Bulgaria, Spain, Greece, Italy and Cyprus. In turn, more than half of those at risk of poverty in Hungary, Sweden, Finland, Ireland and Denmark were removed from this risk as a result of social transfers. Our findings are in line with most of these conclusions.

A vast literature claims that high social efforts contribute for low poverty levels. Behrendt (2002) notes that quantitative studies suggest that there is a strong negative correlation between poverty and social expenditures across European countries. There might be however different patterns in the reduction of poverty rates, where countries like Germany stand out with a relatively low degree of effectiveness in relation to its level of spending, whereas the United Kingdom seems to be much more successful. As stressed by the European Commission (2008), there are clearly large differences among the European Union in social protection expenditure as a percentage of GDP and that countries with the lowest poverty rates are those who spend most on social benefits. Similar conclusions are drawn by Whiteford et al. (1997) for the OCDE countries.

One of the main limitations on the bivariate analysis between expenditure on social transfers and poverty is that the former may reflect differences in the living standards and in demographic, social and institutional structures specific to each country (Longford and Nicodemo, 2010). As referred by Kühner (2007), social expenditure ratios partly reflect demographic and economic conditions rather than social policy. For example, elderly population will probably have an influence between social spending and poverty levels across countries. Furthermore, expenditure ratios do not capture the impact of the tax system or social private arrangements (Caminada and Goudswaard, 2009) in combating poverty. In a cross-country analysis for the OCDE countries, Caminada et al. (2012) confirm the negative relationship between poverty and social expenditures even after controlling for socio-demographic factors such as an aging population, unemployment rates and GDP per capita. According to the authors, these seem to be

the most relevant control variables to be examined in a cross-country perspective.

The relationship between expenditures on social transfers and poverty alleviation should clearly be handled with caution. On the one hand, poverty is not a unidimensional problem (Nanak and Silber, 2008) having only to do with income. On the other hand, social transfers may reflect social, economic and institutional conditions specific to each country with different implications for poverty rates. Our approach seeks to establish the role of social transfers in reducing poverty taking into account this last issue. Using nonparametric techniques, we allow for saturation points in poverty gains, structural breaks and control for demographic and economic conditions between countries.

3 Methodology

The potential advantages of nonparametric econometrics methods lie on the ability to adapt many unknown features of the data (Pagan and Ullah, 1999). Relaxing the parametric assumptions commonly imposed on the data generating process, such techniques let the data determine an appropriate model, allow a smooth representation of the prevailing dynamics and facilitate the detection of structural changes. The most popular methods for nonparametric kernel regression fall into Watson (1964) and Nadaraya (1965) formulation of local-constant smoothing, Stone (1977) and Cleveland (1979) locally weighted estimation or Fan (1992) local polynomial smoothing, just to mention a few. In general, nonparametric models are specified without imposing any particular functional form.

An acceptable starting point for the analysis is to estimate the individual and joint densities of the relevant variables using the kernel method. In our panel data each variable is an observation on n countries for T consecutive years, yielding the form $\{x_{it}\}_{i=1,t=1}^{n,T}$ for each variable in the set. For those cases in which there is available data, $n = 1, ..., 27$ denotes the EU-27 countries and the time index ranges from 2003-2014. Following Racine (2008), a kernel density estimate, $\hat{f}(x)$, for the random variable x is given by

$$\hat{f}(x) = \frac{1}{nTh} \sum_{d=1}^{nT} K\left(\frac{X_d - x}{h}\right), \tag{1}$$

where $nT \equiv \sum_{i=1}^{n} T_i$ [2], h is a bandwidth, x the variable of interest and $K(.)$ the kernel function. To estimate the joint density between two variables, say $x(x_d^1, x_d^2)$, with $d = 1, ..., nT$, a bivariate kernel estimator can be constructed from (1) with a product kernel function of the form $K_h(x^q, x)$, $q = \{1, 2\}$, and a bandwidth $|h| = h^1 h^2$.

The presence of categorical variables requires nothing more than some additional notation and a generalization of the product kernel presented above. Suppose a mix of one continuous variable, x, and one ordered discrete variable, $\bar{x}$. Using the concept of generalized product kernel (Li and Racine, 2003), we can construct a kernel function

[2]For the case of a balanced panel the identity holds as an equality. We will assume throughout that this is the case.

$$K(.) = W(.) \times \overline{l}(.), \tag{2}$$

where $W(.)$ is a standard continuous kernel for x and $\overline{l}(.)$ the kernel for the ordered discrete variable as in Wang and Van Ryzin (1981) or Racine and Li (2004). In this case, the kernel estimator for the joint density between x and $\bar{x}$ is given by

$$\hat{f}(\bar{x}, x) = \frac{1}{nTh_x} \sum_{d=1}^{nT} \overline{l}(\bar{X}_d, \bar{x}) W\left(\frac{X_d - x}{h_x}\right), \tag{3}$$

where h_x is the bandwidth for the continuous variable.

For the sake of generalization, one can also consider an unordered discrete variable, $\hat{x}$, using a kernel function $\hat{l}(.)$ as in Aitchison and Aitken (1976) or Racine and Li (2004). For a given mix of discrete and continuous variables, say one ordered discrete, one unordered discrete and one continuous, build upon (2) a product kernel function yielding,

$$K(.) = W(.) \times \overline{l}(.) \times \hat{l}(.) \tag{4}$$

and derive from (3) a multivariate case,

$$\hat{f}(\hat{x}, \bar{x}, x) = \frac{1}{nTh_x} \sum_{d=1}^{nT} \hat{l}(\hat{X}_d, \hat{x}) \overline{l}(\bar{X}_d, \bar{x}) W\left(\frac{X_d - x}{h_x}\right). \tag{5}$$

From the results in (1) and (3) it is now possible to estimate the density of variable x conditional upon *year*, $g(x|year)$. Given its nature, we treat *year* as an ordered discrete variable and estimate,

$$\hat{g}(x|year) = \hat{f}(\overline{year}, x)/\hat{f}(x). \tag{6}$$

To explore the relationship between social transfers and poverty we propose to estimate a variant of Henderson et al. (2008) nonparametric panel data regression model

$$G_{it} = m(X_{it}) + \mu_i + \nu_{it} \qquad (i = 1, ..., n; t = 1, ..., T_i), \tag{7}$$

where G_{it} denotes poverty reduction at the t-th year for the i-th country. Here, the functional form $m(.)$ is not specified. The covariate $X_{it} = (x_{it}^1, x_{it}^2, ..., x_{it}^q)$ is of dimension q and the remaining variables are scalars. The errors are assumed to be *iid* with a zero mean, finite variance, and uncorrelated with X_{it} for all i and t. Further, μ_i has a zero mean and finite variance and it is allowed to be correlated with X_{it} with an unknown correlation structure. Hence, (7) is a fixed-effects model.

As suggested by Racine (2008), the fixed-effects component can be treated as an unordered discrete (cross-sectional dummy) variable, analogously to the least-squares dummy-variable approach in parametric panel data models. Therefore, instead of imposing in advance additive and separable individual effects, one can incorporate these into the unknown function. Nonparametrically estimating $E[G_{it}|X_{it}, \mu_i] = m(X_{it}, \mu_i)$, using the mixed discrete and continuous kernel technique described above, we test the possibility that the data is in fact potentially poolable.

Let $\hat{\lambda}$ denote the cross-validated smoothing parameter associated with μ_i. If $\hat{\lambda}$ is at its upper bound implies that the variable is irrelevant and is smoothed out. That is, the data is poolable and one gets $E[G_{it}|X_{it}, \mu_i] = m(X_{it})$. If, on the other hand, $\hat{\lambda}$=0 (or close to 0), the data is nonpoolable and this effectively estimates each $m_i(.)$ using only the time series for the ith country. The case $0 < \hat{\lambda} < 1$ suggests that the data is only partially poolable.

Time fixed affects can also be incorporated into the model in a similar manner. The only distinction is that time effects need to be treated as an ordered categorical variable. Here only the distance between points matters, while the magnitude of the data does not. With the inclusion of an ordered categorical variable that comprises several periods, we are implicitly testing for structural changes between points. As before, a cross-validated smoothing parameter associated with time fixed-effects close to zero suggests that these effects are relevant.

One of the most challenging tasks in nonparametric estimation relates to the choice of a proper bandwidth. The bandwidth parameter can be determined by a reference rule-of-thumb, plug-in (two stage) methods, cross-validated methods or bootstrap. For the case of joint density between each variable in the data panel and year, as specified in (3), we will use the normal reference bandwidth which takes the standard formula

$$h = 1.06\sigma n^{1/5}, \tag{8}$$

where σ is replaced by the sample standard deviation, $\hat{\sigma}$. Our choice for such method relies on a more clear noiseless graphical analysis, ensuring the robustness of the results using alternative methods.

For the nonparametric panel data model we use a local constant estimator as described in Racine (2008), with the data-driven AIC_c method (Hurvich et al., 2012) for bandwidth selection, which is based on minimizing a modified Akaike Information Criterion. Although the least-squares cross-validation procedure is similar to AIC_c, the latter is particular adequate for smaller samples. As stressed by Henderson and Parmeter (2015), little is known about the asymptotic performance of the panel-data estimator in the case of introducing fixed-effects as discrete regressors like it is proposed. We assume that the conditional mean is consistent in the presence of such regressors for cross-section and time.

Following Racine et al. (2006), the kernel regression significance test will be performed with a consistence test of significance of an explanatory variable in an nonparametric setting, that is similar to a simple t-test in a parametric environment. The nonparametric results will provide guidances on its possible parametric counterpart, which will be estimated, tested and compared with the nonparametric setting, regarding the respective r-squares and residuals' standard errors. Additionally, the nonparametric consistent model specification proposed by Hsiao et al. (2007) will be used to check whether the parametric models can be rejected against the nonparametric setting. The nonparametric analysis has been conducted using the np package in R software (Hayfield and Racine, 2008).

4 Data and preliminary analysis

The panel data for the 27 European Union countries (EU-27) is collected for the 12 year period 2003 to 2014. The definition of monetary poverty is adopted from EUROSTAT and its *Statistics of Income and Living Conditions* database that collects data from Members States regarding net family income after social transfers and taxes. As aforementioned, the poverty rate is defined as the share of population with an equivalized disposable income below the risk-of-poverty threshold, which is set at 60 percent of the national median disposable income. The EUROSTAT also provides information on that rate before social transfers. The at-risk-of-poverty rate before social transfers measures a hypothetical situation where social transfers are absent (pensions not being considered as a social transfer), directly providing insight about the contribution of those transfers to mitigate poverty.

Table 1 shows the average poverty rate before (*atrisk.bst*) and after (*atrisk.ast*) social transfers per country, between 2003 and 2014, plus some additional statistics. The highest poverty rate before social transfers is observed for Ireland in 2010, whereas the lowest in Czech Republic during 2013. As expected, poverty rates are lower once social transfers are accounted for. With 26.4% of total population below the poverty threshold, Latvia is the country that registered the highest risk-of-poverty rate after social transfers, in 2009. The lowest rate is again observed for the Czech Republic in 2009. On average, the lowest poverty rate is found in Czech Republic, followed by Slovakia and the Netherlands. At the bottom of this ranking we find Ireland, with a poverty rate of more than a third of the population, and the United Kingdom. Interestingly, both these countries climb the poverty rate ranking ladder once social transfers are considered. Romania and Belgium are the two countries with the higher poverty rates after social transfers. This suggests that, among the EU-27, there may be differences in the amount of social transfers directed to the most poor or that some countries are more efficient than others in reducing poverty.

(INSERT TABLE 1 HERE)

In order to better account for the role of social transfers in eradicating poverty, we define the concept of poverty gain (*gain*). This variable can be considered as an efficiency measure of poverty reduction via social transfers and corresponds to the difference between national poverty rates before and after social transfers. Social transfers (*soc.exp*) comprise expenditure on social protection in percentage of the GDP, containing social benefits to households or individuals, administration costs, which represent the costs charged to the scheme, for its management and administration, and other miscellaneous expenditure. Following Caminada et al. (2012), we also consider two additional variables: GDP per capita at constant 2010 prices in logs (*log.GDP*) and the percentage of population aged 65 years or above (*over65*). These control variables intend to capture the social and economic frameworks of the European countries, each with different implications for the social security systems. All variables were retrieved from the EUROSTAT.

As it is shown in Table 2, poverty gains seem to cover a wide range of values across countries (and years), with a coefficient of variation of 43%, while social transfers are much more stable. This may be an indicator of potential inefficiencies in the social protection systems towards poverty reduction, such as relatively high administration costs, inability to direct funds towards groups of population with higher needs or due to other structural characteristics like corruption, quality of the institutions or excessive income redistribution policies. There are, however, other aspects that may explain those differences. Demographic heterogeneity or changes in income distributions, including differences in the share of at-risk-of-poverty population clustered around the threshold, should also be accounted for. With a relatively low disparity in the percentage of population with 65 years or above, Table 2 suggests that differences in the structure of elderly population may not play a major role.

(INSERT TABLE 2 HERE)

In Figure 1 we depict the 0.25 (red dashed line), 0.5 (solid blue line) and 0.75 (green dashed line) conditional quantiles for the EU-27 GDP panel along with box plots of the raw data. The estimates are obtained after inverting the cumulative distribution function (CDF) at the desired quantile, using the nonparametric conditional CDF kernel estimator proposed by Li and Racine (2008). We treat *log.GDP* as continuous, *year* as an ordered discrete variable and use the standard formula (7) for bandwidth selection.

The results in Figure 1 suggest real convergence between countries in the period under analysis, since the estimated inter-quantile distance is narrow in 2014 when compared to that of 2003. Additionally, mean income has been decreasing since 2008, reflecting the adverse consequences of the financial crisis on economic activity. The non-smooth quantile estimates generated by the box plot are in agreement with those obtained via smooth CDF. It is then expectable that the underlying evolution had an impact on poverty rates and on the share of social transfers (and thus on poverty gains), in particular, before and after 2008. To make this more clear, we turn to the analysis of kernel density of poverty gain, social transfers and GDP.

(INSERT FIGURE 1 HERE)

Figure 2 plots the estimated density of poverty gain, social transfers, and GDP conditional upon *year*. We treat the first three variables as continuous and *year* as an ordered discrete. The conditional density uses second-order Gaussian kernel and a rule-of-thumb bandwidth selection. The results reported are robust to bandwidth choice, whether using the simple rule-of-thumb or data-driven methods such as least-squares cross-validation or likelihood cross-validation.

(INSERT FIGURE 2 HERE)

The distribution of poverty gains has evolved from a bimodal one in the early 2000s to a unimodal one after 2008 [Figure 2 (A)]. Some countries seem to have reduced

poverty gains while others increased those gains. Social expenses [Figure 2 (B)] set-off with a quasi-unimodal distribution and evolve to a markedly bimodal one. After 2008, one could arguably identify two distinct social protection policies among the EU-27 regarding the share of social transfers. With respect to income [Figure 2 (C)], the density suggests bimodal distribution but also real convergence between countries, confirming the previous analysis. This evolution is explained by economic growth in low-income countries between the period 2003-2014 but also, particularly after 2008, by a decreasing trend in the proportion of income in the left higher mode.

5 Results

We first restrict our analysis to the relationship between social transfers and poverty gains, which is the main focus of the present work. Therefore, in a first stage, other than categorical factors, only social transfers is included as a regressor in model (7). We then introduce control variables to assess the robustness of the bivariate evidence.

As mentioned in Section 3, the nonparametric poolability test specified by Racine (2008) is performed using local constant kernel method with AIC_c bandwidth selection. We treat each country as an unordered variable and include a categorical factor capturing the period before and after 2008, respectively *before crisis* and *crisis*, which we treat as an ordered discrete variable. Social transfers is assumed to be continuous. An examination of Table 3 reveals that the bandwidths for the individual country factor and for the ordered variable *crisis* are respectively 0.001 and 0.048. These results suggest that the model is not poolable across countries and time (before and after 2008). The residual standard error of the model is 0.79 and the r-squared equal to 97%. Additionally, as shown in Table 4, the significance test proposed by Racine et al. (2006) reject the null at the 1% level on the three variables of the model.

(INSERT TABLE 3 HERE)

(INSERT TABLE 4 HERE)

(INSERT FIGURE 3 HERE)

Figure 3 presents the partial response plots along with their bootstrapped error bounds. We disregard the interpretation of individual country effects for now. The evidence suggests that poverty gain increases with social transfers for values approximately below 23% and slightly decreases above that saturation point. This means that a parametric linear specification of the marginal effects of social transfers on poverty gains might be inappropriate. In our view, this is a sounding result with relevant policy implications. Assuming that social transfers' recipients are those with higher needs, and that benefits increase with poverty severity, a sufficient generous social protection system may cluster up the share of population around the poverty threshold in such a way that there is no incentive for wage earners to keep their jobs. In other words, increases in income redistribution can create substantial disincentives to labor market

participation. Moreover, the results in Figure 3 also suggest that the same share of social benefits, for the mean of EU-27, is less efficient in the crisis period, when compared to the period before 2008. One possible explanation is that for a country to maintain the same level of social expenditures, the share of such benefits in percentage of the GDP would have to rise, since income likely decreases. However, if the share of social benefits actually rose, one could hypothesize that poverty gains surpass the national saturation point, hence triggering the diminishing returns.

As implied by the previous model, a counterpart parametric regression can be inferred. In Table 5 we show the results of a least-squares dummy-variables (LSDV) regression including the country individual dummies, social expenditures and squared social expenditures (Model 1) and a time dummy variable (crisis) which takes the value 0 for the period before 2008 and 1 otherwise (Model 2). Country fixed-effects are omitted. In parenthesis we report the robust Arellano (1987) standard errors as there is indication of heteroskedasticity and serial correlation in the models.

Both parametric models are in accordance with the nonparametric specification. All variables included are significant at the 1% level. Poverty gain is increasing with social expenditures, *ceteris paribus*, while the parameter estimate for squared social expenditures is negative. The estimated saturation points are however higher than that of the nonparametric model. Decreasing gains are attained for social expenditures above approx. 32% and 42% for models (1) and (2), respectively. As expected, the dummy variable coefficient is negative, meaning that poverty gains are decreasing in the crisis period. We tested the inclusion of an interaction term between the dummy crisis and both social expenditures and squared social expenditures but the respective coefficients for the before-crisis period are no longer statistically significant.

Comparing the parametric models with the nonparametric one, we conclude that the latter specification performs better both in terms of the r-squared - approx. 97% against 88% for Model (1) and 90% for Model (2) - and residuals' standard errors - approx. 0.79 against 1.5 for Model (1) and 1.4 for Model (2). In addition, we implement the consistent test for correct specification of parametric regression models as described in Hsiao et al. (2007) which rejects the null of correct specification for models (1) and (2) - see Table 6. These results show that the nonparametric model is better than the parametric specifications for studying the systematic relationship between poverty gains and social expenditures.

(INSERT TABLE 5 HERE)

(INSERT TABLE 6 HERE)

We now proceed with the inclusion of control variables. The main objectives are as follows: 1) to take into account for other potential factors that may be relevant in explaining poverty gains and 2) to check the robustness of the previous results. Also, we will now be interested in interpreting the individual country fixed-effects, which indicate potential inefficiencies among countries. As aforementioned, the control variables to be included are a wealth dimension, the log of real GDP per capita (included in first-differences, d.log.GDP), and a demographic dimension, the percentage of population

with 65 years or above (over65)[3].

(INSERT TABLE 7 HERE)

As before, the nonparametric poolability test is performed using local constant kernel method with AIC_c bandwidth selection. We treat each country as an unordered factor, crisis is treated as an ordered categorical variable and the remaining regressors are considered as continuous variables. As shown in Table 7, the cross-validated smoothing parameters associated with individual country dummies and factor crisis are respectively 0.01 and 0.082. Again, these results suggest that the model is not poolable across countries and time. For this model, the residual standard error is approximately 0.72 and the r-squared 97.1%. There is then a slight improvement over the specification without controls. Moreover, the significance test proposed by Racine et al. (2006) shown in Table 8 suggests a strong rejection for all the regressors included.

(INSERT TABLE 8 HERE)

Figure 4 shows the partial regression plots and bootstrapped error bounds for social expenditures, factor crisis, country fixed-effects and controls. Social expenditures' saturation point is now higher, at approximately 27% in percentage of GDP. Also, for social transfers between approx. 22% and 23% there is, on average, a small decrease in poverty gains. The marginal effect with respect to factor crisis is similar. Other things equal, between 2008-2014, poverty gains are lower when compared to the before 2003-2007 crisis period. The percentage of population over 65 years old has an expected negative effect on poverty gain since the eldest population is related with the national needs for social transfers/pensions and with decreasing contributions to the social protection systems. Moreover, poverty gains increase with income. This result may reflect that the wealthiest nations have lower levels of poverty and, thus, less need for poverty reduction via social transfers. Another possible reason is that, as a country becomes richer, for a given share of social benefits, an efficiency factor is at stake in the allocation of such expenditures towards the poorest. We may however consider that the income effect is negligible as the bootstrapped error bounds indicate. An inspection of the lower-left plot of Figure 4, suggests that, *ceteris paribus*, poverty gains are decisively lower for three countries, namely, Greece, Portugal and Spain. In turn, Finland, Hungary and Luxembourg, followed by Sweden, the United Kingdom and Belgium, are particularly efficient in poverty reduction via social transfers. As mentioned in Section 4, one may argue that there are time-invariant unobserved institutional conditions related with the social protection system explaining such differences. Whether these arise from high administration costs, inability to direct funds towards groups of population with higher needs or due to other structural characteristics remains an open issue.

(INSERT FIGURE 4 HERE)

[3]Other control variables such as the percentage of population aged 20-24 year in tertiary education, trade openness, gini coefficient and remittances were also tested, however the results were unsatisfactory and thus unreported.

Table 9 shows the LSDV estimation including the control variables. As before, we compute heteroskedasticity and autocorrelation consistent standard errors, shown in parenthesis. The results are in line with the nonparametric model. Poverty gains depend on the level of social expenditures, decreasing for estimated values above 39%. Regarding GDP per capita, the estimated coefficient is positive and statistically significant. As expected, other things equal, a 1pp increase in the percentage of population with more than 65 years old is expected to reduce, on average, poverty gains by 0.78pp. Accordingly with the previous comparison, the nonparametric model performs better both in terms of the r-square and residual standard errors. In addition, as suggested by the Hsiao et al. (2007) consistent test show in Table 10, we reject the null of correct specification in the LSDV model. Lastly, in order to assess that our approach throughout has been correctly specified regarding individual (country-specific) effects, we performed a Hausman test on the parametric counterparts considered so far. In this sense, we tested the hypothesis that the country fixed-effects parameters μ_i are uncorrelated with the explanatory variables X_{it}. For the model presented in Table 11, the tests statistic, which under the null follows a Chi-squared distribution with 5 degrees of freedom, takes a value of 9, suggesting a rejection at the 10% level of significance.

(INSERT TABLE 9 HERE)

(INSERT TABLE 10 HERE)

(INSERT TABLE 11 HERE)

6 Conclusion

In the present work we analyzed the efficiency of social expenditures in reducing poverty levels among the EU-27 between the period 2003-2014. For this goal, we propose an efficiency measure which corresponds to the difference between national poverty rates before and after social transfers that we label as poverty gains. Using a nonparametric panel data fixed-effects model, we studied the relationship between poverty gains and social expenditures, and allowed for the detection of structural breaks and inefficiency differences between countries and periods. With this specification we assume that the time-invariant unobserved individual country effects are correlated with social expenditures and with the control variables to be included as regressors.

The results suggest that there is a significant relationship between poverty gains and social expenditures, unconditional after controlling for the percentage of population aged 65 or more years old and (the first-difference of log) real GDP per capita. The relationship is positive until some threshold, estimated for social expenditures in percentage of the GDP above 23%, without controls, and 27%, once these are included. This indicates that, other things equal, increases in social expenditures above that threshold do not contribute for poverty alleviation, for the mean of the EU-27, describing a situation of total inefficiency. As expected, poverty gains are decreasing with the percentage of population with 65 years or more, reflecting the higher needs for social transfers and pensions but also with decreasing funds directed to the social protection

system. Furthermore, poverty gains decrease in the crisis period suggesting that countries either can't cope with increasing social expenditures or that saturation points have been reached, triggering decreasing poverty gains. Poverty gains are increasing with income, a negligible effect once bootstrapped error bounds are considered. Wealthiest nations may have less need for poverty alleviation through social transfers or they may be more efficient allocating those transfers. Regarding individual country fixed-effects, the evidence suggests that, *ceteris paribus*, social expenditures are particularly efficient in reducing poverty for Finland, Hungary and Luxembourg, whereas Greece, Portugal and Spain are the least efficient among the EU-27. We hypothesize that this result may be related with the quality of the social protection system in the allocation of the funds. With this reasoning, it is possible that saturation points differ across EU-27. We highlight our intention to analyze such possibility, which we leave for further investigation.

The main findings of the nonparametric modulation are in line with the implied parametric counterparts. The coefficient estimates obtained from the LSDV specifications confirm the nonparametric results. Furthermore, we were able to show that the nonparametric models perform better than the parametric ones both in terms of the r-square and residual standard errors. Also, the nonparametric specification tests employed reject the parametric settings over the nonparametric ones. The Hausman test applied to the parametric model with control variables confirms that we correctly specified the country-specific effects. It is also our intention to construct an appropriate test for the nonparametric model.

As a final remark, we point out to some warning features regarding our poverty measures. The poverty rate at use in this study relies on a distributional measure with reference to a threshold, 60% of the median income. Thus, it assumes that an equally increase in the income, albeit significant, has no effect in poverty reduction, which we believe it's incorrect. Furthermore, that measure entirely neglects unemployed population that does not benefit from social assistance, hence disregarding the most poor. A correct assessment of poverty should be sensible to these features. We stress, however, that these issues were out of our scope.

Appendix

Table 1: Poverty rate before and after social transfers by country (2003-2014)

	atrisk.bst				atrisk.ast			
	min	max	mean	std.dev	min	max	mean	std.dev
AT	24.6	27.1	25.6	0.7	12.0	15.2	13.8	1.0
BE	26.3	29.1	27.4	0.8	14.3	15.5	15.0	0.4
BG	24.7	27.4	26.5	0.9	18.4	22.2	21.2	1.1
CY	21.0	24.6	23.0	1.2	14.4	16.1	15.4	0.6
CZ	16.6	21.6	18.8	1.7	8.6	10.4	9.4	0.6
DE	23.1	25.7	24.5	0.7	12.2	16.7	15.1	1.5
DK	26.8	31.5	28.8	1.5	10.9	13.3	12.2	0.8
EE	24.2	28.4	25.4	1.2	15.8	21.8	18.8	1.6
EL	22.5	28.0	24.3	1.8	19.6	23.1	20.9	1.3
ES	23.7	31.1	27.2	2.6	19.7	22.2	20.5	0.7
FI	26.2	28.9	27.5	0.9	11.0	13.8	12.8	0.9
FR	23.5	26.4	24.8	1.0	12.5	14.1	13.3	0.5
HU	26.3	30.4	28.5	1.4	12.3	15.9	13.6	1.2
IE	30.9	39.9	35.7	3.3	14.1	20.9	16.9	2.4
IT	23.3	24.7	23.9	0.6	18.4	19.8	19.2	0.4
LT	25.5	31.3	28.2	1.9	18.6	20.9	19.9	0.8
LU	22.4	29.4	25.8	2.7	11.9	16.4	14.1	1.3
LV	25.7	31.0	27.6	1.8	19.0	26.4	21.6	2.8
MT	20.1	24.0	22.6	1.3	14.2	15.9	15.2	0.6
NL	19.9	21.7	20.8	0.5	9.7	11.6	10.6	0.6
PL	22.9	29.8	25.1	2.4	16.9	20.5	17.8	1.2
PT	24.2	26.7	25.5	0.8	17.9	20.4	18.6	0.8
RO	27.5	30.9	28.9	1.3	21.1	25.4	23.0	1.4
SE	26.6	30.1	28.0	1.1	9.5	15.1	12.7	1.8
SL	22.0	25.9	24.2	1.2	11.3	14.5	12.9	1.2
SK	17.1	21.9	19.5	1.3	10.9	13.3	12.3	0.9
UK	28.9	31.0	30.0	0.6	15.9	19.0	17.3	1.3

List of countries included: Austria (AT), Belgium (BE), Bulgaria (BG), Cyprus (CY), Czech Republic (CZ), Germany (DE), Denmark (DK), Estonia (EE), Greece (EL), Spain (ES), Finland (FI), France (FR), Hungary (HU), Ireland (IE), Italy (IT), Lithuania (LT), Luxembourg (LU), Latvia (LV), Malta (MT), Netherlands (NL), Poland (PL), Portugal (PT), Romania (RO), Sweden (SE), Slovenia (SL), Slovakia (SK) and United Kingdom (UK).

Table 2: Poverty gains, social transfers and elderly population (EU-27 panel)

	min	max	mean	std.dev	coef.var
gain	2.6	24.7	9.95	4.33	0.43
soc.exp	10.5	33.7	22.78	5.81	0.26
over65	10.8	21.4	16.24	2.30	0.14

Figure 1: Nonparametric conditional quantile estimates for the EU-27 GDP panel

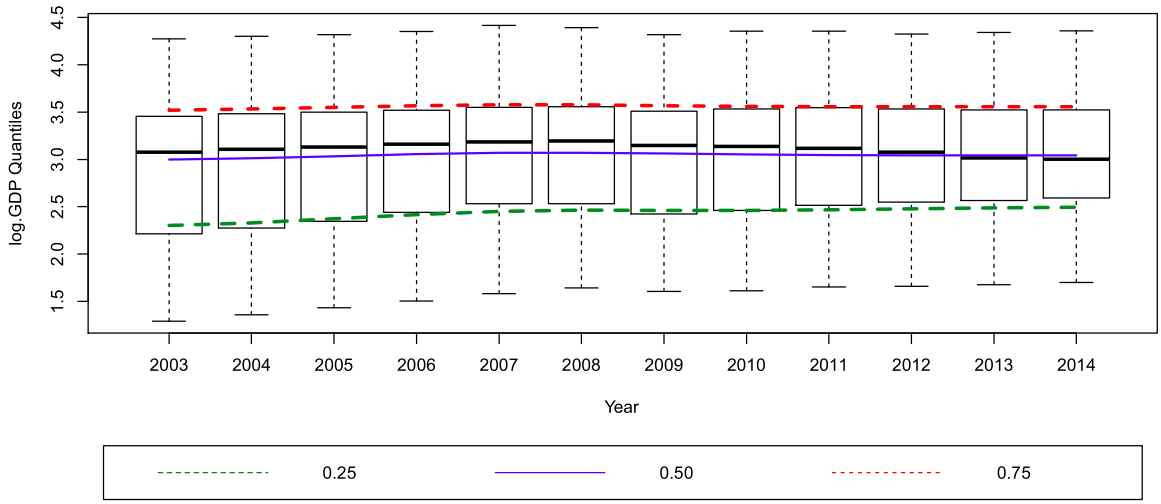

Figure 2: Nonparametric conditional density estimates for the EU-27 panel

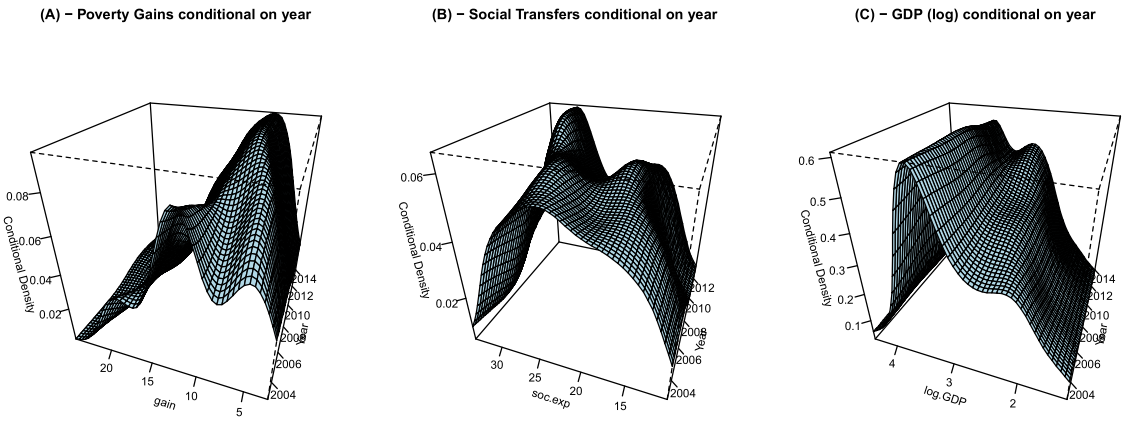

Table 3: Model summary for the nonparametric fixed-effects EU-27 panel

```
Regression Type: Local-Constant
Bandwidth Selection Method: Expected Kullback-Leibler Cross-Validation
Formula: gain ~ soc.exp + ordered(crisis) + factor(country)
Bandwidth Type: Fixed
Objective Function Value: 1.3 (achieved on multistart 1)

Exp. Var. Name: soc.exp          Bandwidth: 1.5      Scale Factor: 0.81
Exp. Var. Name: ordered(crisis)  Bandwidth: 0.048    Lambda Max: 1
Exp. Var. Name: factor(country)  Bandwidth: 0.0012   Lambda Max: 1

Residual standard error: 0.7936
R-squared: 0.9654

Continuous Kernel Type: Second-Order Gaussian
No. Continuous Explanatory Vars.: 1

Unordered Categorical Kernel Type: Li and Racine (normalized)
No. Unordered Categorical Explanatory Vars.: 1

Ordered Categorical Kernel Type: Li and Racine
No. Ordered Categorical Explanatory Vars.: 1
```

Table 4: Model summary for the nonparametric fixed-effects EU-27 panel

```
Kernel Regression Significance Test
Type I Test with IID Bootstrap (399 replications)
Explanatory variables tested for significance:
soc.exp (1), ordered(crisis) (2), factor(country) (3)

               soc.exp ordered(crisis) factor(country)
Bandwidth(s):      1.5           0.048          0.0012

Individual Significance Tests
P Value:
soc.exp         <2e-16 ***
ordered(crisis) <2e-16 ***
factor(country) <2e-16 ***
---
Signif. codes:  0 '***' 0.001 '**' 0.01 '*' 0.05 '.' 0.1 ' ' 1
```

Figure 3: Partial regression plot and bootstrapped error bounds for the EU-27 panel

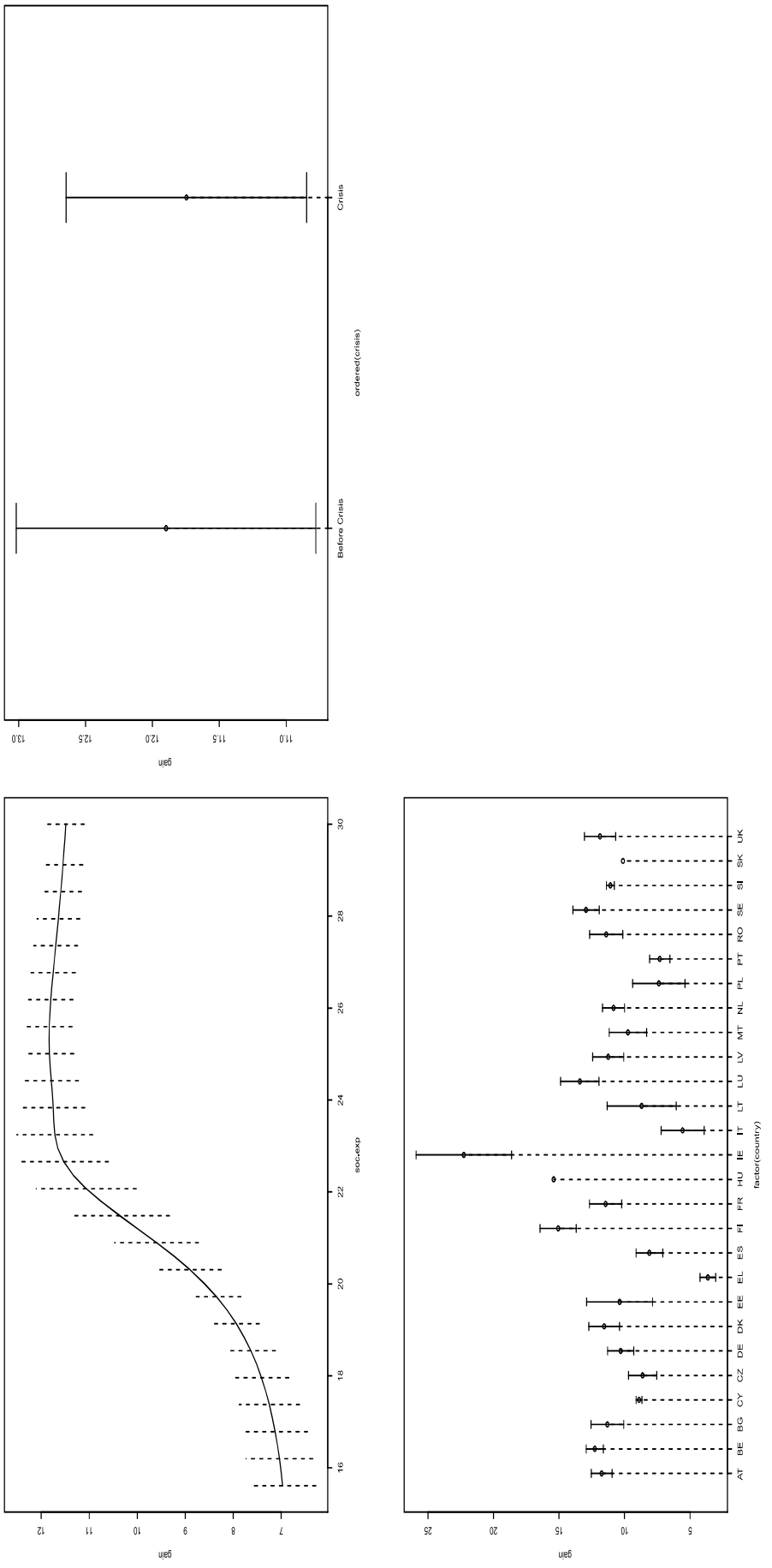

Table 5: LSDV fixed-effects results

	Dependent variable:	
	Poverty Gain	
	Model (1)	Model (2)
soc.exp	1.300***	1.600***
	(0.320)	(0.340)
$soc.exp^2$	−0.021***	−0.019***
	(0.006)	(0.006)
crisis		−1.800***
		(0.350)
Observations	252	252
R^2	0.880	0.900
Adjusted R^2	0.870	0.890
Residual Std. Error	1.500 (df = 223)	1.400 (df = 222)
F Statistic	61.000*** (df = 28; 223)	71.000*** (df = 29; 222)
Note:		*p<0.1; **p<0.05; ***p<0.01

Table 6: Summary of the model specification tests for the parametric LSDV fixed-effects models

```
## Model (1) ##
Consistent Model Specification Test
Parametric null model: lm(formula = gain ~ soc.exp + I(soc.exp^2) +
factor(region), data = data_panel, x = TRUE, y = TRUE)
Number of regressors: 2
IID Bootstrap (399 replications)

Test Statistic Jn: 3.8 P Value: <2e-16 ***
---
Null of correct specification is rejected at the 0.1% level

## Model (2) ##
Consistent Model Specification Test
Parametric null model: lm(formula = gain ~ soc.exp + I(soc.exp^2) +
factor(region) + crisis, data = data_panel, x = TRUE, y = TRUE)
Number of regressors: 3
IID Bootstrap (399 replications)

Test Statistic Jn: 6.4 P Value: <2e-16 ***
---
Null of correct specification is rejected at the 0.1% level
```

Table 7: Model summary for the nonparametric fixed-effects EU-27 panel including controls

```
Regression Type: Local-Constant
Bandwidth Selection Method: Expected Kullback-Leibler Cross-Validation
Formula: gain ~ soc.exp + over65 + d.log.gdp + ordered(crisis) + factor(country)
Bandwidth Type: Fixed
Objective Function Value: 1.2 (achieved on multistart 1)

  Var.: soc.exp         Bandwidth: 1.8     Scale Factor: 0.68
  Var.: over65          Bandwidth: 1.2     Scale Factor: 1.5
  Var.: d.log.gdp       Bandwidth: 0.1     Scale Factor: 7.7
  Var.: ordered(crisis) Bandwidth: 0.082   Lambda Max: 1
  Var.: factor(country) Bandwidth: 0.01    Lambda Max: 1

  Residual standard error: 0.7196
  R-squared: 0.9715

  Continuous Kernel Type: Second-Order Gaussian
  No. Continuous Explanatory Vars.: 3

  Unordered Categorical Kernel Type: Li and Racine (normalized)
  No. Unordered Categorical Explanatory Vars.: 1

  Ordered Categorical Kernel Type: Li and Racine
  No. Ordered Categorical Explanatory Vars.: 1
```

Table 8: Model summary for the nonparametric fixed-effects EU-27 panel including controls

```
Kernel Regression Significance Test
Type I Test with IID Bootstrap (399 replications, Pivot = TRUE, joint = FALSE)
Explanatory variables tested for significance:
soc.exp (1), over65 (2), d.log.gdp (3), ordered(crisis) (4), factor(country) (5)

                 soc.exp  over65  d.log.gdp ordered(crisis) factor(country)
Bandwidth(s):      1.79   1.214     0.1038         0.08201        0.01023

 Individual Significance Tests
 P Value:
 soc.exp         <2e-16 ***
 over65          <2e-16 ***
 d.log.gdp       <2e-16 ***
 ordered(crisis) <2e-16 ***
 factor(country) <2e-16 ***
 ---
 Signif. codes:  0 '***' 0.001 '**' 0.01 '*' 0.05 '.' 0.1 ' ' 1
```

Table 9: LSDV fixed-effects results with control variables included

	Dependent variable:
	Poverty Gain
soc.exp	2.000^{***}
	(0.270)
soc.exp^2	-0.026^{***}
	(0.005)
d.log.gdp	20.000^{***}
	(2.900)
crisis	-0.720^{**}
	(0.320)
over65	-0.730^{***}
	(0.200)
Observations	246
R^2	0.930
Adjusted R^2	0.920
Residual Std. Error	1.200 (df = 214)
F Statistic	94.000^{***} (df = 31; 214)
Note:	*p<0.1; **p<0.05; ***p<0.01

Figure 4: Partial regression plot and bootstrapped error bounds for the EU-27 panel including controls

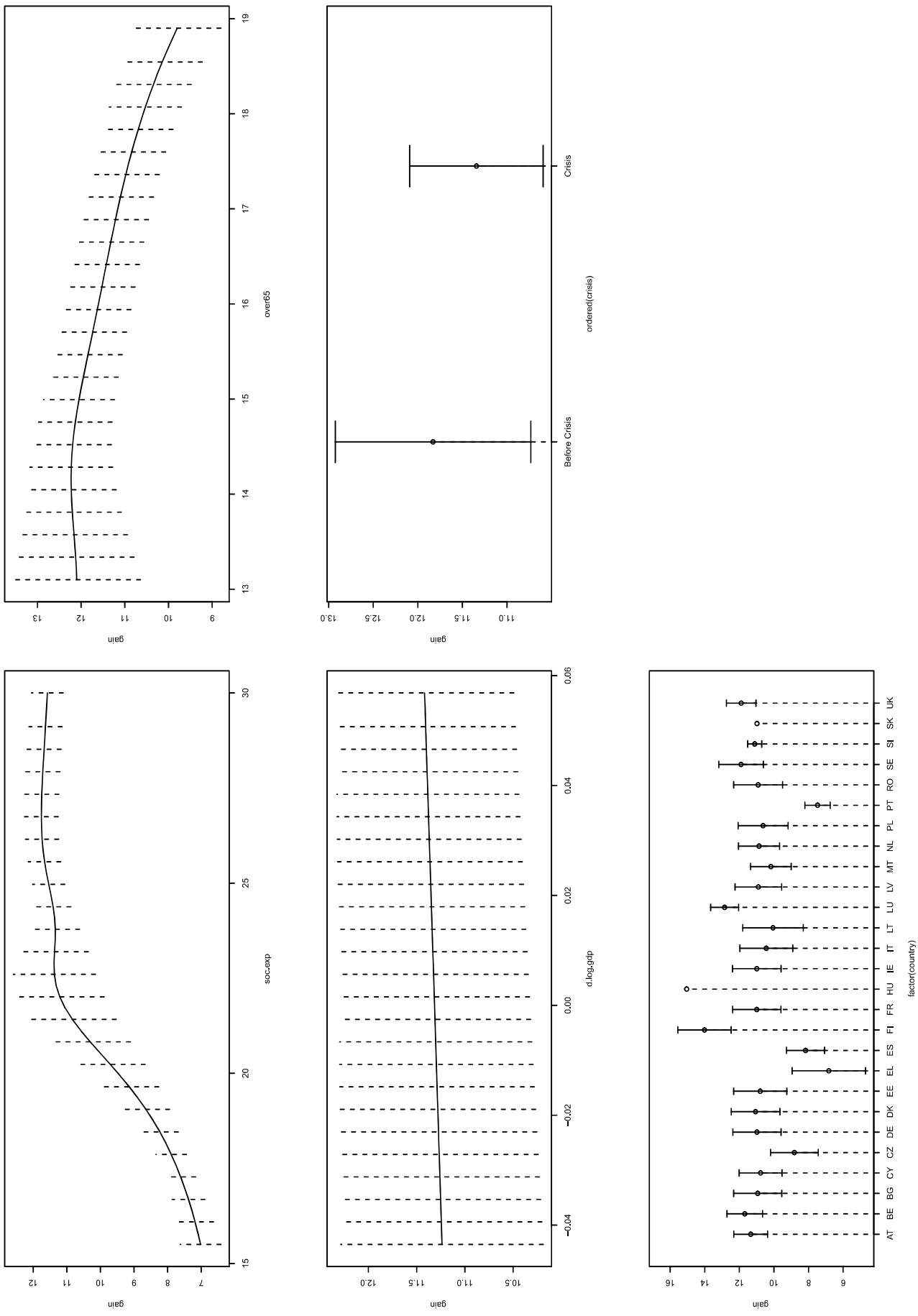

Table 10: Summary of the model specification tests for the parametric LSDV fixed-effects model with control variables included

```
Consistent Model Specification Test
Parametric null model: lm(formula = gain ~ soc.exp + I(soc.exp^2) +
d.log.gdp + over65 + factor(region) + crisis, data = data_panel,
x = TRUE, y = TRUE)
Number of regressors: 5
IID Bootstrap (399 replications)

Test Statistic Jn: 6.1   P Value: <2e-16 ***
---
Null of correct specification is rejected at the 0.1% level
```

Table 11: Hausman Test

```
data:  gain ~ soc.exp + I(soc.exp^2) + over65 +
d.log.gdp + crisis

chisq = 9, df = 5, p-value = 0.09
alternative hypothesis: one model is inconsistent
```

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