Revisiting the relationship between welfare spending and income inequality in OECD countries

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Abstract

The present paper estimates the effects of welfare interventions on income inequality. We propose a theoretical model showing that welfare policies follow the median voter constituency regardless of whether governments are center-left or center-right in the majority electoral system, whereas large differences exist between center-left and center- right coalitions in the proportional representation system. We exploit these differences in the mechanisms of welfare expenditure to estimate their elasticities on income inequality and find that a 1% increase in government spending reduces the Gini income index by half a percentage point. This result is robust under different compositions of expenditure, alternative imputation model specifications and falsification tests.

Keywords: Welfare policies; Electoral rules; Income inequality; Instrumental variable approach; OECD countries
JEL: H23; H53, E62; C26

1. Introduction

Income inequality has increased significantly in most countries over the past three decades. The rise was most marked in the United States, where the share of the richest 1% in all pre-tax income has more than doubled since 1980, reaching almost 20% in 2012. Top earners also fared very well in several other English-speaking countries including Australia, Canada, Ireland and the United Kingdom. A striking change is also observed in countries with a history of a more equal income distribution. Between 1980 and the late 2000s, the share of the top 1% increased by 70% in Finland, Norway and Sweden, reaching around 7-8%. By contrast, top earners saw their share grow much less in some of the continental European countries, including France, the Netherlands and Spain (Bergh and Nilsson, 2010; Organization for Economic Cooperation and Development, 2014).

Thus, an open question remains why rich OECD countries have significant long-term trends of unequal income distribution. One strand of literature points to inadequate or inefficient government intervention in the redistribution of resources (income) to the poor, which is typically measured as the difference between market and disposable income inequality at household level. While some early empirical investigations including government spending as an explanatory variable for income distribution found an equalizing effect (Buliu and Gulde, 1995; Gustafsson and Johansson, 1999; Li *et al.*, 2000; Schaltegger and Weder, 2014), the general trends in the 21st century were of cutbacks in welfare state generosity. This implies that benefit replacement rates have been lower than predicted, allowing existing income transfer programs to blunt the impact of rising earning inequality. These explanations can also be extended when we analyze the components of welfare state transfers (Smeeding and Grodner, 2000; Hacker *et al.*, 2005; Kenworthy, 2004; Smeeding, 2005), although when welfare state generosity includes the value of in-kind government and private services such as health care and public education it tends to limit the level of bottom-end inequality in the United States considerably (Adema and Ladaique, 2005; Garfinkel *et al.*, 2006; Hacker, 2002).

Another strand of literature suggests that further mechanisms may explain why differences in income distribution occur. Some authors have argued that some cultural and historical factors may play a role. In particular, (Bergh and Bjørnskov, 2014) have investigated the causality between trust and equality, and the potential role of welfare state policies as mediators of the causal associations. Instead, other authors have recognized the importance of electoral systems in determining large differences in welfare state policies and, in turn, in income inequality (see for example Alesina and Glaeser 2005; Milesi-Ferretti *et al.* 2002; Persson and Tabellini 2005). It is a fact that the European countries, which are regulated by a proportional representation system for most of the time, have been more prone to introducing redistributive policies, whereas English-speaking countries, which tend to be regulated by majority electoral systems, are less inclined to redistribute. Nevertheless, the electoral system interacts with the preferences of political parties emerging from the elections results. Alesina *et al.* (1997); Perotti and Kontopoulos (2002); Persson and Tabellini (2005) show that left-wing parties tend to spend and tax more than right-wing parties, and this difference can be particularly important in proportional representation systems in which it is easier for every group to find a voice in Parliament. The distinction between left and right-wing parties, at least in the proportional electoral formula, may become empirically relevant to explaining differences in welfare policies.

The present paper shows theoretically how majority and proportional electoral systems affect the redistributive policies and, in turn, income distribution. We extend the theoretical framework proposed by Iversen (2005) and Iversen and Soskice (2006),and suggest that the electoral system induces changes in welfare preferences when heterogeneous optimizing behaviors of "winning" coalition are included. We find that both left- and right-wing parties will follow the median voter constituency in the majority system, leading to similar redistributive policies. Conversely, in the proportional representation system there is a large difference between center-left and center-right coalitions in the targeted policies.

We propose a structural model for empirical analysis based on a panel of 21 OECD countries for the 1995-2010 period, and derived from the relevant inequality and welfare spending reduced forms, and test whether and to what extent more generous welfare transfers are related to more equality in the distribution of income. We exploit the outcome of the theoretical model suggesting that the interaction between the electoral systems and the parties or coalition winning the elections generates exogenous cross-country differences in allocating welfare transfers which serves as an instrument to estimate welfare-inequality relationship. Thus, an instrumental variable (IV) approach is applied in the empirical analysis.

The empirical results suggest that a larger budget for welfare state transfers is related to a lower income inequality. More interestingly, the estimated elasticities indicate that a one percentage increase in welfare transfers reduces the Gini index by about half a percentage point. When we extend the estimates for the components of welfare, we still find that the dimension of the impact of "not in kind" and "in kind" transfers and pensions is close to the aggregate result, and that these findings are robust even when the analysis is performed for countries with high and low levels of income inequality or for time-varying sample analysis.

The remainder of the paper is organized as follows: Section 2 summarizes the background and presents the theoretical model, while in Section 3 we show the identification strategy and the resulting structural equations. Section 4 then discusses the data set and the variables used. Section 5 presents the estimates, whereas various checks of the robustness of the results are discussed in Section 6. Concluding remarks are then presented in Section 7.

2. Theoretical model

2.1. Background

A society of individuals i is assumed to be classified in three classes. These classes characterize the low-income L, middle-income M and high-income H groups (Persson *et al.*, 2003). It is also assumed that the voting population is equally distributed between these three groups and that they have different preferences about the relevant policy choices[1]. We consider an indirect utility function $V^i : Q \to \mathbb{R}$ in which Q is a set of possible policy choices. Q includes $q^i \in Q$ representing the subset of policies that maximize the value of the indirect utility function of the group i, such that $V^i(q^i) \geq V^i(q)$, with the single-peaked preferences assumption $V^i(q)$ strictly concave. Thus, transferring a given quantity of income from one group to another, the indirect utility functions $V^i(q^i)$ depends on the disposable income of each group $(1-\tau)y^i$ and on the transfers received by the group i from the government (i.e., Tr^i). The indirect utility function reads:

$$V^i(q^i) = (1-\tau)y^i + Tr^i \tag{1}$$

where y^i is the gross income of group i and τ is a lump-sum tax on income ($\tau \leq 1$).

In order to include the progressive taxation in the model, which characterizes the redistributive policy, we follow Iversen (2005) in which transfers G represent a cost for the higher income groups M and H and a benefit for the lower income group L. We also assume that these transfers are costly and are paid for a non non-negligible share ϵ by people in M, and for the residual amount $1-\epsilon$ by people in H. We assume an upper limit G^* which is not modifiable without an agreement involving the financing groups. Furthermore, a constant cost α ($\alpha > 0$) on G is assumed to characterize the administrative costs, which includes red-type ones. This cost is not trivial, since it allows us to consider explicitly the measure of efficiency in the public goods provided by the government.

Next, we assume that the government finances the expenditure by revenues (Γ^i) for each group proportionally by leaving a flat-rate income tax (τ) and imposing the redistributive transfer policy discussed earlier. The budget constraint rules are:

$$\Gamma^L = \tau y^L \tag{2}$$
$$\Gamma^M = \tau y^M + (1+\alpha)\epsilon G \tag{3}$$
$$\Gamma^H = \tau y^H + (1+\alpha)(1-\epsilon)G. \tag{4}$$

Note that, the three income groups have different goals concerning G and τ. Since progressive taxation is assumed, the high income group will pay the largest amount of G without receiving any transfer (see equation 4). As a consequence, the optimal strategy for this group will be to reduce both G and τ to zero. On the other hand, the lower income group will receive the majority of transfers without paying G, and hence will prefer setting G to G^* and τ to 1. The middle income group, which is supposed to be equidistant between the other groups has conflictual strategies concerning τ and G. When τ is considered, the preferences of M are more in line with L, whereas its preferences are more closed to H when G is taken into account. In line with Meltzer and Richard (1981), there is a positive effect due to taxation which decreases by income and M may gain an higher utility from a positive τ. The optimality conditions for M will lead it to chose an intermediate level of taxation τ^{*m} (e.g., about 0.5) and to set G to zero. As it is clear by (3) and (4), when ϵ is negligible, H will pay the largest part of the cost of the redistributive policy and the preferences of M and L will converge perfectly if $\epsilon = 0$. Following these arguments, we can formalize the aims for each of the three income groups as a function of G and τ. For simplicity, we use the preferences v, expressed in terms

[1] Acemoglu and Robinson (2005) provides a version of the model in which there are only two groups of different size. In this case the chosen redistributive policy depends by the size of each group and the growth rate of inequality across groups.

of income shares, such that $g = G/y$ and we obtain:

$$
\begin{aligned}
v^L &= g + \tau && (5)\\
v^M &= -(\tau - \tau^{*m}) - g(1+\alpha)\epsilon && (6)\\
v^H &= -\tau - g(1+\alpha)(1-\epsilon) && (7)
\end{aligned}
$$

2.2. Modeling proportional representation

Intuitively, the basic model suggests that if one of the three income groups has the majority to be elected in the country elections, it will impose its preferences concerning τ and g upon the other groups. To define the proportional representation system, we introduce two different sets of preference functions for the LM and MH coalitions, respectively.

$$
\textit{Coalition } LM \quad \begin{cases} \hat{v}^L = \tau + g - \tau^{*m} \\ \hat{v}^M = (1-\tau^{*m}) - (\tau - \tau^{*m}) - g(1+\alpha)\epsilon \end{cases} \tag{8}
$$

$$
\textit{Coalition } MH \quad \begin{cases} \hat{v}^M = -(\tau - \tau^{*m}) - g(1+\alpha)\epsilon \\ \hat{v}^H = -\tau - g(1+\alpha)(1-\epsilon) \end{cases} \tag{9}
$$

Unlike equations 5, 6 and 7, we note that the preference functions in equations 8 and 9 depend on the preferences of the counterpart. For example, lets denote by $\hat{v}^L_{LM}$, the preferences of L when there is a coalition between LM; it includes $\tau + g$, which represents the preferences of the L and τ^{*m} which is the optimal tax rate for M. Since M prefers to set $g = 0$, τ^{*m} is the only variable into the M part. By aiming to find the solution of a multidimensional bargaining game, each group needs to satisfy the condition for each coalition which is symmetric for each player.

For example, when we consider the coalition LM and L is the "first player" the Rubinstein bargaining solution is obtained equalizing own preference on choice variables (g, τ) with those of M. Technically this implies substituting the key variables in the preference functions of each group with those of other group forming the coalition and maximizing under this constraint. Following this scheme, L is available to contract g and τ which makes L willing to accept a coalition and *vice versa*. Summarizing, the pay-offs for each coalition are as follow:

$$
\textit{Coalition } LM \quad \begin{cases} \textbf{L play} \Longrightarrow (1-\tau^{*m}) - (\tau^L - \tau^{*m}) - [g^* - g^L(1+\alpha)\epsilon = \\ \delta\left[(1-\tau^{*m}) - (\tau^M - \tau^{*m}) - [g^* - g^M(1+\alpha)\epsilon\right] \\ \textbf{M play} \Longrightarrow \tau^M + g^M - \tau^{*m} = \delta\left[\tau^L + g^L - \tau^{*m}\right] \end{cases} \tag{10}
$$

$$
\textit{Coalition } MH \quad \begin{cases} \textbf{M play} \Longrightarrow -\tau^M = -\delta\tau^H \\ \textbf{H play} \Longrightarrow -(\tau^H - \tau^{*m}) = -\delta(\tau^M - \tau^{*m}) \end{cases} \tag{11}
$$

where τ^L, τ^M and τ^H are the preferred tax rates for each group and where δ is a discount factor. Solving 10 and 11, we obtain a value of τ and g that allows the players to set-up a given coalition. When $\delta \rightsquigarrow 1$, these parameters read:

$$
\textit{Coalition } LM \quad \Big\{ \tau = \tfrac{1+\tau^{*m}}{2} - \tfrac{g}{2} + \tfrac{(1+\alpha)(g^*-g)\epsilon}{2} \quad \textit{and} \quad g = g^* \tag{12}
$$

$$
\textit{Coalition } MH \quad \Big\{ \tau = \tfrac{\tau^{*m}}{2} \quad \textit{and} \quad g = 0 \tag{13}
$$

where, (12) depends upon the cost of the redistributive policy α, the share of resources collected from the middle-income group ϵ and by the values of transfers costs g, whereas (13) depends only upon the optimal tax rate of the middle-income group. From the second condition, we find that in the coalition MH the tax rate is half of the optimal tax rate for M and the middle-income group will obtain a constant utility from this coalition that does not maximise its preferences. On the contrary, from the first coalition, we see that when

$g = g^*$, Ml obtains a higher utility from the redistributive policy. To understand when the LM coalition will be preferred to the MH, we run some comparative static on g, by introducing the optimal values of τ, obtained by (12) in the preference function of the player M. The comparative statics reads:

$$Coalition\ LM \quad \left\{ \frac{\partial \hat{V}^M}{\partial g} > 0 \quad if \quad \epsilon < \frac{1}{1+\alpha} \right. \tag{14}$$

From (14), we see that M will obtain a positive utility from the redistributive policy, but that the share of resources that M is willing to pay must not be higher than the inverse of the administrative costs for the redistributive policy. When there are no administrative costs $\alpha = 0$, M will always obtain a positive utility from the policy, but when these costs become relevant, individuals in group M will be willing to pay a lower amount of resources for policy. The progressiveness of the redistributive policy and the administrative costs represent the crucial parameters in determining the coalition between low and middle income groups. Table 1 summarizes the probabilities for a given coalition to win the elections, for different values of ϵ and

Table 1: Expected coalitions in a proportional representation system

Condition on ϵ	Condition on α	Probability of LM coalition	Tax system
If $\epsilon \rightsquigarrow 0$	$\alpha \rightsquigarrow \infty$	$Pr(LM) \rightsquigarrow 0$	Strong progressive tax rate
If $\epsilon \rightsquigarrow 0$	$\alpha \rightsquigarrow 0$	$Pr(LM) \rightsquigarrow 1$	
If $\epsilon \rightsquigarrow \frac{1}{2}$	$\alpha > 1$	$Pr(LM) \rightsquigarrow 0$	Progressive tax rate
If $\epsilon \rightsquigarrow \frac{1}{2}$	$\alpha < 1$	$Pr(LM) \rightsquigarrow 1$	
If $\epsilon \rightsquigarrow 1$	$\alpha > 0$	$Pr(LM) \rightsquigarrow 0$	Regressive tax rate
If $\epsilon \rightsquigarrow 1$	$\alpha \rightsquigarrow 0$	$Pr(LM) \rightsquigarrow 1$	

α. As shown by this Table, if the tax system is strongly progressive, we could expect a failure in the LM coalition. Moreover, the administrative costs can determine different coalition and hence different welfare policies. Following this scheme, we show that in a proportional representation system the coalition between center-left parties is more probable than a coalition between center-right ones; also for given countries with high administrative costs, a reverse probability can be found. Even more interesting, we show that targeted redistributive policies differ, depending on whether a center-left or a center-right coalition is in charge.

2.3. Modeling the majority representation

The majority electoral system is characterized by a winner-take-all approach for a restricted number of competing parties. We simplify the model by supposing that only a center-right CR and a center-left CL political party can take part in the election. That is, in the majority electoral system each party needs to attract the median income votes to win the elections, and, to do that, will converge to the policy preferences of the median voter constituency that is $\{g, \tau\} = \{0, \tau^{*m}\}$. As in proportional representation, the median constituent shares with H the same preferences on g, but is more similar to L when τ is considered. As a consequence, H and L will converge yo the preferences of M, with corresponding preferences as in equation (7).

When the government is in charge, the left- and right- wing parties have incentives to diverge by the median income preferences, adopting policies that reflect their own constituency needs (Persson *et al.*, 2003; Iversen, 2005). The median voter may reduce its own utility considerably but, at this stage, has no instruments to influence the redistributive policy promoted by the government. We can summarize the costs of a deviation from the electoral choice for the median voter as:

$$T_{CL} = g^* + \tau^{*m} \tag{15}$$
$$T_{CH} = -\tau^{*m} \tag{16}$$

In turn, this deviation may reduce the credibility of the party, favoring a convergence with its own constituency. The loss of reputation may be a matter for the government since it makes harder to deal with other agents. From another point of view, this credibility loss may reduce the possibility for the government to build other policies diverging by the willing of its own constituency. These costs, defined by c_{CL} and c_{CH}, constitute the pay-offs for a political party to diverge from its electoral promises. Using this framework, Table 2 summarizes the probabilities of a center-right and a center-left party winning the election. Under

Table 2: Expected probabilities for each party to be in charge in a majority system

Condition for CL	Condition for CH	Probability for CH win
If $T_{CL} < C_{CL}$	$T_{CH} < C_{CH}$	Indifference between CL and CH
If $T_{CL} < C_{CL}$	$T_{CH} > C_{CH}$	$Pr(CH) \rightsquigarrow 1$
If $T_{CL} > C_{CL}$	$T_{CH} > C_{CH}$	$Pr(CH) \rightsquigarrow 1$
If $T_{CL} < C_{CL}$	$T_{CH} > C_{CH}$	$Pr(CH) \rightsquigarrow 0$

this framework Iversen (2005) shows that the left party has a higher incentive to diverge from policies that are unattractive for the median income constituency. In the majority electoral system the median-voter will be less inclined to converge in a center-left party and, hence, the center-right party will win more often.

More interesting, the model shows that the policy promoted by the elected political party, following the median constituency preferences, will be less redistributive than in a proportional representation system. Also when a center-left party has the majority in a country, $\{g, \tau\} = \{0, \tau^{*m}\}$ target is chosen, since the deviation from it will produce a loss of credibility for the left party with a pay-off c_{CL}. Consequently, when a center-left or e center-right party wins the elections in a majority electoral system, the amount of welfare benefits will always be lower than the ones provided in a proportional system, since both the coalitions will converge to $\{g, \tau\} = \{0, \tau^{*m}\}$ target.

3. The identification strategy and the structural model

The theoretical model presented in the previous sections shows that the link between the electoral representation system and the share of resources devoted to welfare policies depends on the probability of a center-left or a center-right party or coalition winning the elections. More in detail, we found that in countries regulated by a majority electoral system (MS), from a welfare perspective, it is indifferent which coalition wins the elections whereas, in the proportional representation country, the center-left coalition (PS_{CL}) will redistribute more resources than the center-right coalition (PS_{CR}). This structure can be used to estimate the relation between welfare spending and income distribution, by controlling for potential endogeneity in the inequality regression. Focusing upon welfare transfers, the cross-country relation between the specific welfare policy and income inequality using the GINI indexes can be specified as:

$$Gini_{it} = \alpha_1 + \alpha_{2it} + \alpha_{3t} + \Psi_1 Wp^j_{it} + X'_{it}\Psi_2 + v_{it} \tag{17}$$

where $Gini_{it}$ measures the income distribution by the disposable income for the country i in year t and Wp^j_{it} is the specific welfare policy j for the country i in year t. α_{2it} and α_{3t} are trends at country level and year fixed effects, respectively, and X_{it} is a vector of controls variables that will be described in the next section.

Equation 17 may have some identification and estimation issues. Indeed, Wp^j_{it} may be influenced by feedback effects, as increased income inequality may lead to increase welfare expenditures, or expectations of the outcome of the process undertaken by the government to allocate expenditures may be correlated with the current level of income inequality. Since welfare policies might be viewed as a mechanism to reduce income inequality, the latter may induce policy-makers to increase welfare spending for high levels of income inequality, i.e., reverse causality (see, for example, Niehues 2010; Doerrenberg and Peichl 2014). Thus, welfare policy expenditure cannot simply be assumed to be exogenous, meaning that the empirical estimates are not easy to interpret because there are potentially other factors at play.

To deal with this, an instrumental variable (IV) approach can be used to identify suitable instruments. In the present context, we overcome this problem by using the main findings of the proposed theoretical model. The interaction between the electoral representation system and the political parties or coalition winning the elections is used as an instrument in this context, which characterizes the full set of preferences connected to the welfare policies and includes the progressive taxation[2]. The results defines an empirical ordinal variable ES_W, which increases with respect to the propensity to consider welfare transfers. This variable ranges from -1 to 1, where we encode: -1 countries with MS, 0 countries with (PS_{CR}) and 1 countries with (PS_{CL}).

Taking this IV approach, the inequality regression in the reduced form can be specified as:

$$Wp_{it}^{j} = \beta_1 + \beta_{2it} + \beta_{3t} + \delta_1 ES_{W_{it}} + X_{it}\delta_2 + u_{it} \tag{18}$$

$$Gini_{it} = \lambda_1 + \lambda_{2it} + \lambda_{3t} + \xi_1 ES_{W_{it}} + X_{it}\xi_2 + z_{it} \tag{19}$$

where welfare coefficient on inequality policy spending j in (17) is recovered by the ratio of the reduced form coefficients of ES_W on Wp_{it}^{j} and $Gini_{it}$, that is $\Psi_1 = \xi_1/\delta_1$.

4. Data and descriptive statistics

The present section provides a comprehensive description of data collected for 21 OECD countries in the1995-2010 period. In the section, we provide an overview of the link between welfare spending and income distribution, although special concern arises regard to the variable of income distribution, given the consistent missing values in these series. Below, we will deal with these shortcomings by using the imputation methods and show the consistency of these results by using the dataset with non-missing data.

4.1. Income inequality

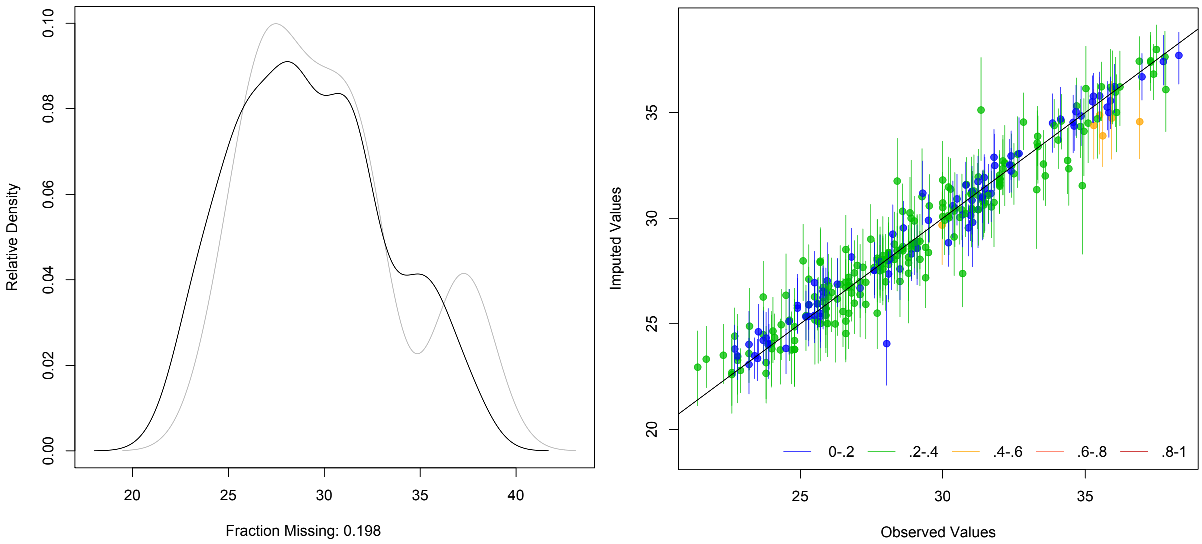
Our measure of inequality is $Gini_{it}$ index measured by the disposable income which accounts for the effect of progressive taxation within income percentiles. We collect data from different sources. The first index is extracted from the "National Accounts" of the OECD statistics ($GINI_{OECD}$); the second from the European Institute of Statistics ($GINI_{Eurostat}$); and third from the "GINI Growing Inequalities' Impacts" ($GINI_{GII}$), a project financed by the European Union, which integrates official information with different data sources derived from specific country statistics. Even if these three indexes are strongly correlated each other, they differ substantially in terms of the number of missing values. We carried out our analyses using $GINI_{GII}$, since it contains less missing data, while using the other indicators to implement imputation analysis and extract some robustness. Below, we will return to the data description.

We apply a multiple imputation method proposed by King *et al.* (2001); Honaker and King (2010) which uses the relevant information from the observed data to impute multiple values for each missing cell. This approach is based on two principal assumptions. The first regards the complete dataset D. Let $D = \{D^{OBS}, D^{MIS}\}$ a nxk matrix partitioned into its observed and missing elements that includes all dependent and explanatory variables and any other variable useful for the prediction of the missing values, with a multivariate Normal distribution $D \sim Nk(\mu, \Sigma)$. The second assumption regards the pattern of missingness that depends upon the observed dataset, D^{OBS}, so that the data are missing randomly (MAR). Let M be the missingness matrix, with the same dimension as D, with cells $m_{ij} = 1$ if the data in D is missing and $m_{ij} = 0$ if otherwise. M can be predicted by D^{OBS}. Formally:

$$p(M/D) = p(M/D^{OBS}). \tag{20}$$

[2]It is worth noting that this specification excludes the influence of omitted variable through the progressive taxation g, which is a direct channel in affecting welfare policies.

Figure 1: Statistic checks for the imputed GINI index ($Gini_{GII}$)



(a) Relative density (b) Falsification test

We briefly describe the iterative procedure that allows us to determine multiple imputations in the Appendix. An important point is that the missing values in the series of GINI indicators do not concern the same countries in the same period. Thus, a set of related macroeconomic variables which contain a few missing value as disposable income, growth rate of per-capita GDP and the share of total public spending in GPD (extracted by OECD) are included. The expected correlation with the GINI indicators should contribute to reproduce consistently missing data.

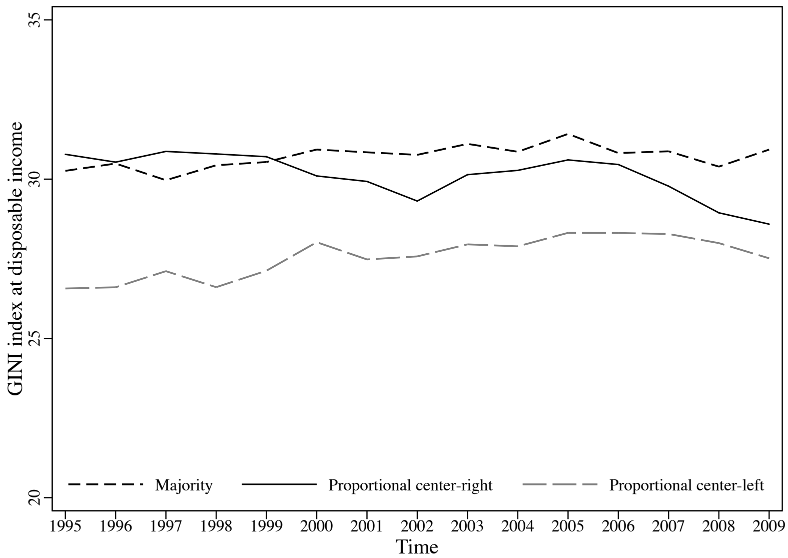
Although the multiple imputation involves imputing m values for each missing item and creating m completed data sets, it summarizes the matrix results in five independent series, obtained by a bootstrap replication of the missing values. Following King *et al.* (2001), we will use one of these series (e.g., the third), for the empirical analysis[3], whereas the remaining four series will be used for robustness.

Figure 1 inspects the goodness of the imputed GINI index ($Gini_{GII}$). Panel a) shows the relative density of the imputed and original series; panel b) shows a falsification test which uses the imputation method to replace each observation into the five series generated by the five models and investigates the distance with the linear prediction (Abayomi *et al.*, 2008). Both these checks confirm the robustness of the imputed $Gini_{GII}$, and specifically from the falsification test we find that variation across the imputations is small enough that the use of the median operator is full consistent.

Figure 2 propose a first look into the graphical link between ES_W and the temporal behavior of $Gini_{GII}$. The figure reports three different lines describing the mean value of the GINI index for the majority electoral system and for center-right and center-left coalitions in a proportional electoral system. The figure illustrates interesting regularities. Countries in a majority electoral system systematically have a higher value of the unequal distribution, with a significant difference with respect to the other electoral systems. In accordance with the theoretical model proposed above, there is clearly a greater propensity in the proportional electoral system - especially when the elections are won by a center-left government, to consider the detrimental effects of income inequality a priority.

[3]Following King *et al.* (2001), an analysis using 1000 imputed series, has been also used to ensure the constancy of imputed values across different bootstrap replications. The results of this analysis are omitted since no relevant differences are encountered among the imputed series.

Figure 2: Temporal behavior of the aggregate GINI index ($Gini_{GII}$)



4.2. Welfare expenditures

To characterize welfare policy interventions, we collected data from the National Accounts of the OECD, extracting aggregate information about social benefits and distinguishing between "in kind", and "not in kind" social spending. Not in kind transfers (Wp_{NIK}) are typically *cash* and their use is indistinguishable from income coming from other sources, whereas recipients of kind transfers (Wp_K) have no discretion over their use (Lequiller and Blades, 2007). Wp_S is the aggregate welfare transfers. As in Garfinkel *et al.*

Table 3: Classification of social expenditure

	Not in-kind	In-Kind
Pension	Early retirement pension, and other benefits	Residential care, home-help services and other benefits
Family	Family allowances, maternity and parental leave	Day care, home-help services and other benefits
Survivors	Pension and other benefits	Funeral expenses and other benefits
Incapacity related	Disability pensions, pensions, paid sick leave, other benefits	Residential care , home-help services, rehabilitation services and other benefits
Health	-	All services
Unemployment	Unemployment compensation, severance pay and early retirement for labour market reasons	Unspecified benefits
Housing	-	Housing assistance and other benefits
Other social policy areas	Income maintenance and other benefits	Social assistance and other benefits

(2006), we distinguish between in-kind and not in-kind transfers because it is of interest for policy-makers to understand if different returns exist in terms of income inequality. Indeed, while the impact of not in-kind transfers has been widely investigated (see for example Alderson and Nielsen 2002; Moller *et al.* 2003), the literature contains little evidence on the effects of in-kind transfers (exceptions are Garfinkel *et al.* (2006); Sefton (2002); Lampman (1984)). We also consider benefits for old-age pensions, which represent one of the main sources of not in-kind transfers in many countries. Table 3 labels welfare expenditures, whereas Table 4 lists the share of each expenditure in GDP by country. For example, we register that the Scandinavian countries (Denmark, Sweden), which have the highest shares of welfare benefits (first column of Table 4), spend considerably over an half of these transfers through in-kind benefits. Even if it is evidently a country-based variability, the mean over the time span and country shows a range between 46% and 54% between

”not in kind” and ”in kind” expenditures, irrespective of the type of the electoral system interacting with the coalition that won the elections.

Table 4: Share of welfare transfers by country

Country	Transfers			
	Total Wp_T	Not in-kind Wp_{NIK}	In-kind Wp_{IK}	Old-pension * Wp_P
Austria	30.592	18.815	11.778	10.607
Belgium	28.463	15.163	13.301	7.153
Canada	22.218	10.503	11.715	3.947
Czech Republic	21.784	11.873	9.911	6.613
Denmark	33.636	16.761	16.875	7.513
Finland	30.523	16.798	13.725	8.340
France	31.687	17.478	14.209	10.893
Germany	28.949	17.463	11.486	8.640
Greece	25.678	15.559	10.119	10.273
Hungary	25.357	13.904	11.454	7.791
Ireland	20.889	10.025	10.864	3.067
Italy	26.989	16.281	10.708	11.313
Luxembourg	23.498	14.058	9.440	6.220
Netherlands	25.117	11.159	13.959	5.447
Portugal	23.658	12.536	11.122	7.727
Slovak Republic	21.539	12.993	8.546	5.673
Slovenia	26.098	15.321	10.777	9.614
Spain	22.051	12.126	9.926	7.253
Sweden	33.159	15.178	17.981	9.487
United Kingdom	23.972	12.718	11.254	5.733
United States	.	11.684	.	5.320
MS	26.307	14.150	12.021	7.278
(PS_{CR})	25.414	13.843	11.295	7.224
(PS_{CL})	27.089	14.859	12.230	7.821

Notes: * MS is majoritarian electoral system; (PS_{CR}) is the proportional system with centre-right coalition (PS_{CR}); (PS_{CL}) centre-left coalition.

We also consider old age pension benefits which represent over a half of not in-kind transfers in the majority of the countries considered. In pay-as-you-go systems, pensions may have important redistributive effects in favor of poor elderly people, because it may reduce the inequality of distribution of income between the retirees (Heinrich, 2000) and the unequal distribution of earnings before retirement (Disney and Johnson, 2001; Alesina and Glaeser, 2005).

To provide a first impression of the relation between ES_W and the allocation of welfare transfers, at the bottom of the Table, we disentangle these policies between the two extremes of $ES_{W_{it}}$, that is, the center-right coalition in a majority electoral system (MS)and proportional representation with a center-left coalition (PS_{CL}). The Table shows that the second case also includes almost 2% of welfare transfers, shared by the components (about 1%) when each transfer is considered separately. Less evident differences can be described when we consider old-age pension benefits only.

4.3. Electoral systems, political variables and other controls

We extracted information characterizing the interaction between electoral system and political party orientation from the World Bank database of Political Institutions 2010 (WBPI2010) and considered the legislative and executive indexes of electoral competitiveness establishing whether electoral representation is ”proportional” or majority (Beck *et al.*, 2001). The first system is characterized by the condition following which the candidates are elected based on the percent of votes received by their political party. The latter is a system in which legislators are elected using a winner-take-all or first past the post rule. Using this definition, we include in the majority electoral system Canada, Czech Republic, France, Germany, Greece, Hungary, Italy, Korea, Poland, Spain, United Kingdom, United States and in the proportional representation system

Austria, Belgium, Denmark, Estonia, Finland, Ireland, Luxembourg, Netherlands, Poland, Portugal, Slovak Republic, Slovenia, Sweden.

Secondly, we considered the party orientation. This variable is coded through the following criteria: i) right, for parties that are defined as conservative, Christian Democratic, or right-wing; ii) left, for parties that are defined as communist, socialist, social democratic, or left-wing, and iii) center, for parties that are defined as centrist or when the party position can be described as centrist (e.g. party advocates strengthening private enterprise in a social-liberal context). Hence, to characterize center-left and center-right coalitions, we also collected information on the orientation of the parties, tha form the government majority.

Using this information, we set-up an ordered variable of the interaction variable between electoral system and political party orientation (ES_W), which we defined in paragraph 3, with the restriction suggested through the theoretical model of non-heterogeneous payoffs targeted redistributive policies when the center-left and center-right parties are in charge in a majority system.

From the same source of data we collected further information about the structure of government. As a first degree of approximation, we set up three different dummy variables to distinguish when a nationalist, regional-based and religious party supports the government. This variable allows us to improve the way in which we characterize different identities that can constrain the welfare policy and that are not ascribable to center-left and center-right coalitions (Alesina and Glaeser, 2005). In addition, we also consider how long the present government is due to remain in office and the share of votes obtained into the last round of voting.

To complete the dataset, we extracted the growth rate of pollution, the growth rate of per-capita GDP, the share of public spending in GDP and the World Bank government effectiveness index from the OECD National Accounts. The first series was used to account for differences in demographic pressure between countries while the growth rate of per-capita GDP accounts for the economic trends of each country and also acts as a control for the effect of the economic crisis that has taken place in many countries from 2006. Considering government spending as a share of GDP is useful to account for the size of the government and as a proxy of total taxation. Finally, the World Bank government effectiveness index is able to capture for differences between countries with regard to the services provided to citizens and is a proxy of the cost of government intervention.

5. Results

Table 5 presents the estimates in equation (18) of the correlations between electoral system interacted with the preferences of political parties (ES_W) and welfare expenditure, which includes time trends for each country, time fixed effects and the set of country covariates described in the previous section. We analyse the magnitude of this correlation for the share of welfare transfers (Wp_T), for the components of "not in-kind" (Wp_{NIK}) and "in-kind" (Wp_{IK}) expenditure and for the specific expenditure related to "old age pension benefits" (Wp_P).

The first column of Table 5 shows a significant and positive correlation between ES_W and Wp_T. Given the ES_W ordering, the significant positive effect of the parameter means that (PS_{CL}) provides a significant grater amount of public resources to redistributive policies with respect to the other electoral-government combinations. Disentangling this expenditure by considering in-kind and not in-kind transfers, we also find a significant correlation of ES_W in both transfers (columns II and III). This result is confirmed in the last column of the table. Fostered by the proportional system with center-left coalition we have a significant increase in old age pension benefits.

Table 6 reports the correlation between ES_W and income inequality indicator ($GINI_{GII}$) (equation 19). The three columns of the Table compare the *naive* model (column I), with the one that also includes fixed effects (column II) and models estimated with fixed effects and the country covariates (column III). The comparison shows that the linear fit of the model (Adjusted R^2) strongly increased with the inclusion of fixed effects and covariates. More interesting, the presence of a proportional system, with a center-left government significantly linked with a country-wide reduction of income inequalities, which supports our contention relative to this political party's greater attention to inequality.

Table 5: Reduced form estimates of the relationship between electoral representation system and parties winning the elections and welfare transfers (Equation 18)

	Transfers						Benefit	
	Total (Wp_t)		Not in-kind (Wp_{NIK})		In-kind (Wp_{IK})		Old-age pension (Wp_P)	
ES_W	3.358	***	1.852	***	1.506	***	0.724	***
	(0.374)		(0.265)		(0.209)		(0.195)	
Time trend	yes		yes		yes		yes	
Fixed effects	yes		yes		yes		yes	
Covariates	yes		yes		yes		yes	
Adjusted R^2	0.793		0.724		0.842		0.818	
N	282		282		282		275	

Notes: ES_W is the interaction variable between the electoral representation system and the political parties or coalition winning the elections. The robust standard error are shown in brackets. The asterisks give *p-value* significance levels: * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$.

Table 6: Reduced form estimates of the relationship between electoral representation system and parties winning the elections and GINI index (Equation 19)

	$Gini_{GII}$ I		$Gini_{GII}$ II		$Gini_{GII}$ III	
ES_W	-1.546	***	-1.535	***	-1.549	***
	(0.236)		(0.252)		(0.370)	
Time trend	no		yes		yes	
Fixed effects	no		yes		yes	
Covariates	no		no		yes	
Adjusted R^2	0.129		0.701		0.745	
N	282		282		282	

Notes: ES_W is the interaction variable between the electoral representation system and the political parties or coalition winning the elections. The robuts standard error are shown in brackets. The asterisks give *p-value* significance levels: * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$.

Table 7 lists the IV estimates of the structural welfare expenditure-income inequality model. Each specification includes time trends for each country, time fixed effects and the set of country covariates. From a first comparison across the columns of Table 7, we can see that the proposed specifications provide a fairly good approximation of the linear link between income distribution and the different welfare transfers considered. Note that the reduced-form estimates provide a check for the IV method in the form of the structural parameter estimate. For example, when social transfers in aggregate is accounted, the coefficient $\Psi_1 = \xi_1/\delta_1$ is equal to the ratio of the reduced form parameters in equation (19) and (18), respectively (i.e., $-1.546/3.358 = -0.461$).

For each specification, we test for weak instruments. Below the Table we report the first stage F-statistics and Wald statistics based on Cragg and Donald (1993) and the Kleibergen and Paap (2006) generalization to non-independently and non-identically distributed errors, along with the associated $p - values$ for weak-instruments hypothesis tests (Bazzi and Clemens, 2013). The tests support the choice concerning the instrument variable excluding that estimates may be somewhat biased toward ordinary least squares (OLS) estimates (Bound, Jaeger, and Baker 1995; Staiger and Stock 1997).

However, we are interested in comparing the impact of social expenditure with that of the components because we want to test if the composition of welfare transfers affects income inequality. Since all the expenditures are measured as shares with respect to GDP, we use the elasticities of these relationships which we list below the estimated parameters in bold character. A percentage increase of welfare transfers leads to a 0.43 percentage point reduction in income inequality. This estimated elasticity is close to the results obtained from in-kind (0.43) and not in-kind (0.42) transfers as also shown by the confidence intervals of the

Table 7: Estimates of structural parameters and elasticities of the relationship between welfare transfers and GINI index (IV estimates; instrument: ES_W)

	$Gini_{GII}$ I	$Gini_{GII}$ II	$Gini_{GII}$ III	$Gini_{GII}$ IV
Wp_{tot}	-0.461 ***			
	(0.094)			
	-0.430 ***			
	[0.087]			
Wp_{nik}		-0.836 ***		
		(0.181)		
		-0.425 ***		
		[0.092]		
Wp_{ik}			-1.028 ***	
			(0.233)	
			-0.438 ***	
			[0.099]	
Wp_p				-1.916 ***
				(0.590)
				-0.519 ***
				[0.161]
Time trend	yes	yes	yes	yes
Fixed effects	yes	yes	yes	yes
Covariates	yes	yes	yes	yes
Adjusted R^2	0.804	0.790	0.744	0.596
N	282	282	282	275
Kleibergen-Paap F test statistic†	67.967	41.370	43.873	11.665
	(0.000)	(0.000)	(0.000)	(0.001)
Cragg-Donald F statistic	114.610	61.838	90.398	18.495
	(0.000)	(0.000)	(0.000)	(0.001)

Notes: ES_W is the interaction variable between the electoral representation system and the political parties or coalition winning the elections. The elasticities are shown in bold. The robust standard error are listed in brackets. The asterisks give *p-value* significance levels: * $p <$ 0.1; ** $p < 0.05$; *** $p < 0.01$. We report first stage F-statistics and Wald statistics based on Cragg and Donald (1993) and the Kleibergen and Paap (2006) generalisation to non-independently and non-identically distributed errors, along with the associated $p - values$ for weak-instruments hypothesis tests (Bazzi and Clemens 2013). † Confidence intervals for the Kleibergen-Paap F test statistic follow Bazzi and Clemens (2013).

estimates. This may imply that while it is known that the components of government expenditures have a different productivity in growth rate and reduce their returns if the share of government expenditure increases (Barro, 1990, Devarajan et al. 1996), the effect of welfare expenditure shares on income inequality seems to be linked with the *proportional* propulsive mechanism generating disposable income. This statement seems to be confirmed in the old pension component. Although it is estimated a large effect by the parameter of this specific share on income inequality, the impact on GINI indicator is close to the aggregate effect when the elasticity is estimated.

5.1. Country's heterogeneous effects

As shown in the previous section, there is no differential impact within the components of welfare expenditure on the income inequality indicator, since all the estimated elasticities rely on the confidence interval of Wp_T. Table 9 below replicates the estimation results proposed in Table 7, where we compare countries with a GINI index above and below the mean of the OECD sample. As a general result, a comparison between the two sub-samples shows that the effect of welfare spending is roughly doubled when countries with high income distribution are analyzed, and that there is an elasticity of about -0.5 ($p - value < 1\%$) with respect to an elasticity of a lower income distribution of -0.27. A similar outcome is also found when we distinguish between in-kind and not in-kind transfers, while there is less difference between elasticities when old age pension benefits are analyzed. However, this variability is not sufficient to undermine our conclusions because confidence intervals at 95% significant level overcomes all the points estimate.

Table 8: Welfare transfers and GINI index: heterogeneous effects

	$Gini_{GII}$							
	High inequality sample	Low inequality sample	High inequality sample	Low inequality sample	High inequality sample	Low inequality sample	High inequality sample	Low inequality sample
Wp_T	-0.616 ***	-0.259 **						
	(0.129)	(0.118)						
	-0.503 ***	**-0.274** **						
	[0.105]	**[0.124]**						
Wp_{NIK}			-1.389 ***	-0.617 *				
			(0.327)	(0.345)				
			-0.606 ***	**-0.358** *				
			(0.143)	**(0.200)**				
Wp_{IK}					-1.107 ***	-0.446 **		
					(0.242)	(0.184)		
					-0.421 ***	**-0.214** **		
					(0.092)	**(0.088)**		
Wp_P							-1.507 ***	-1.411
							(0.257)	(0.981)
Wp_P							**-0.356** ***	**-0.440**
							(0.061)	**(0.306)**
Fixed effects	yes	yes	yes	yes	yes	yes	yes	yes
Covariates	yes	yes	yes	yes	yes	yes	yes	yes
Adjusted R^2	0.798	0.637	0.725	0.521	0.791	0.651	0.727	0.292
N	141	141	141	141	141	141	135	140
Kleibergen-Paap F test statistic†	13.089	35.072	7.658	10.648	19.933	30.515	16.185	6.093
	(0.000)	(0.000)	(0.004	(0.001	(0.000)	(0.000)	(0.000)	(0.014
Cragg-Donald F statistic	33.668	45.759	17.133	15.871	50.357	42.068	30.766	10.657
	(0.000)	(0.000)	(0.004	(0.001	(0.000)	(0.000)	(0.000)	(0.014

Notes: ES_W is the interaction variable between the electoral representation system and the political parties or coalition winning the elections. The elasticities are shown in bold. The robust standard error are listed in brackets. The asterisks give *p-value* significance levels: * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$. We report first stage F-statistics and Wald statistics based on Cragg and Donald (1993) and the Kleibergen and Paap (2006) generalisation to non-independently and non-identically distributed errors, along with the associated $p - values$ for weak-instruments hypothesis tests (Bazzi and Clemens 2013). † Confidence intervals for the Kleibergen-Paap F test statistic foolow Bazzi and Clemens (2013).

6. Robustness analysis

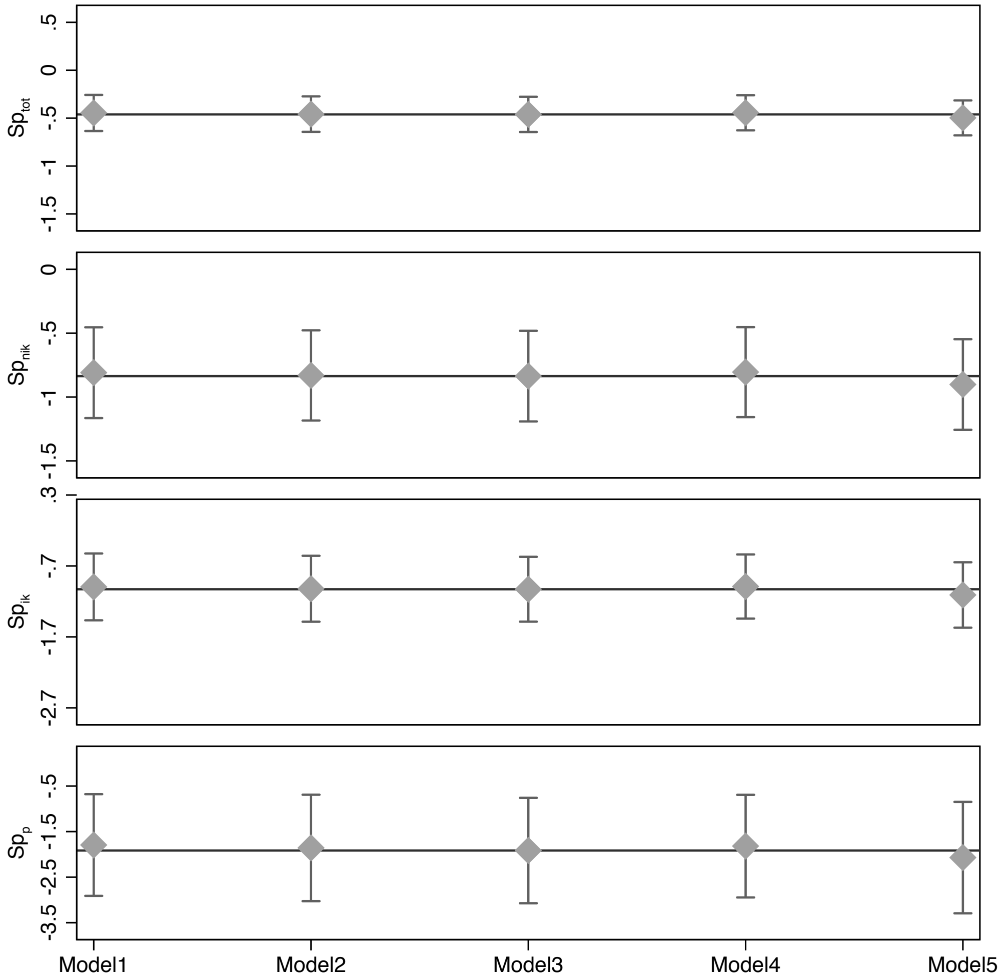
In this section we explore the robustness of the results. The initial check, we show how the estimates of the structural model may change when imputation models are used, since, as discussed early, imputation models use different independent bootstrap replication (King *et al.*, 2001). Figure 3 shows the estimated parameters for the $GINI_{GII}$ and the corresponding confidence intervals for the five imputed series using the welfare transfer variables presented in Table 7. We find small changes in the parameter magnitude and signs of the estimated relationship are in accordance with those in Table 7, showing the goodness of the proposed imputation method.

As a second robustness check, we show the stability of our results with respect to different time spans. This analysis has a two-fold aim. One the one hand, we test if our estimates are time-dependent and, on the other hand, we investigate whether the results are driven by the economic crisis that has taken place in many of the analyzed countries since 2008. For this purposes, Figure 4 reports the estimated parameter of equation two as in Table 7 for each considered welfare transfer. The time intervals considered are: i) 1997-2010, ii) 1996-2010, iii) 1995-2010, iv) 1995-2009, v) 1995-2008 and vi) 1995-2007. Again, the horizontal lines describe the coefficients in 7, whereas the vertical lines show the confidence interval of each estimated parameter. As in the previous case, the estimates are robust across different time intervals and always statistically significant. Note also that when the sample is restricted to 2007, thus excluding the possible effect of the economic crisis, we do not find any differences in the estimated parameters.

We are also aware that the imputation method may lead to a bias in the studied relationship. Although we cannot correct it because in principle we should include predictors which forecast the lack of an income inequality index perfectly, we can propose a restricted estimation which use non-imputed data. Table 10 lists the estimated parameters and elasticities. Even if the estimated effects of the welfare transfer elasticities are significant and pervasively lower, we note that they are not different to the ones found in the benchmark model.

Finally, Table 10 replicates the specifications reported in Table 7 by using the GINI index extracted from the National Accounts of the OECD ($GINI_{OECD}$). Unlike the ($GINI_{GII}$), this index has the advantage of usinga more homogeneous source of information, although the exclusion of supplementary surveys to obtain this index determines more missing, and in turn, more imputed values in the series.

Figure 3: Estimates of the relationship between welfare transfers and GINI index for different imputed models



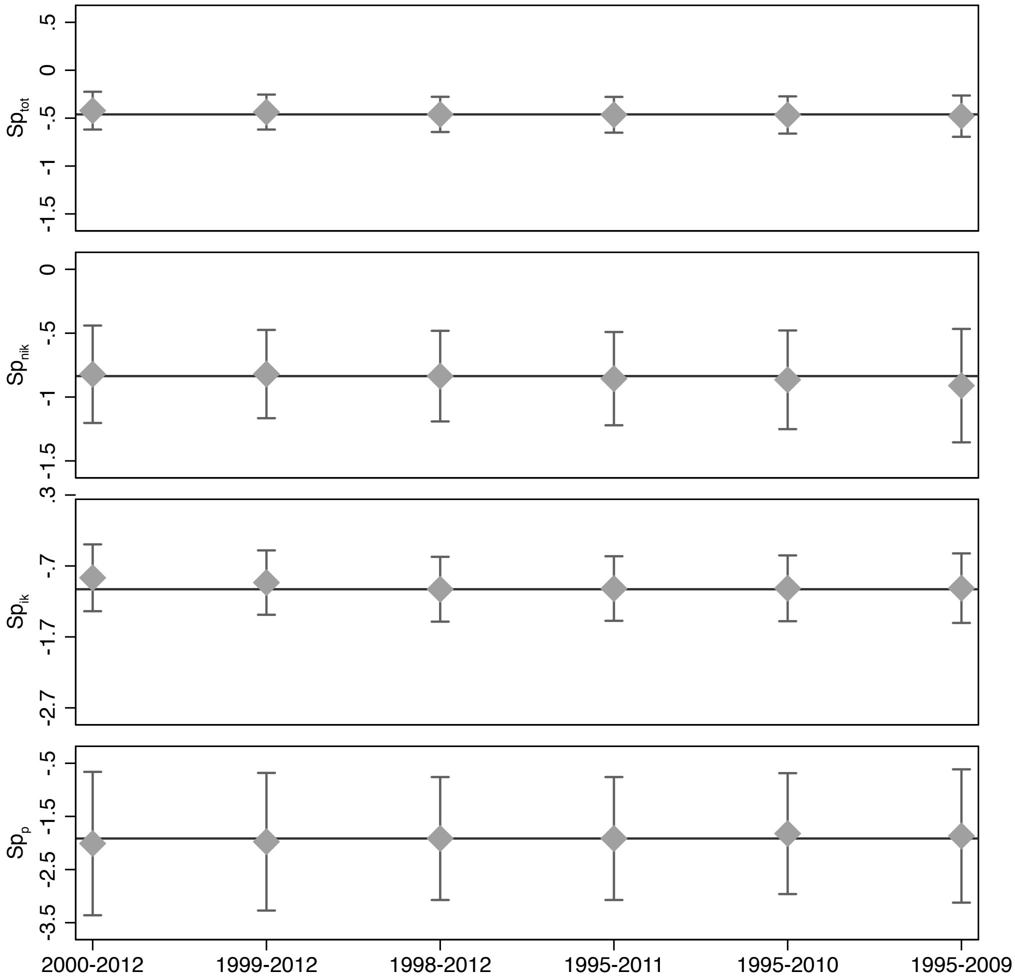
The estimates in Table 10 are in accordance with the ones presented above, even if the magnitude is higher than in the previous case. All the parameters and elasticities are significant at a 99% confidence interval. The more remarkable point estimates emerging from the Table derive from old age pension benefits. In this case, we find that a response of a 1% change in this component of not kind welfare leads to a 0.75 percent reduction in income inequality of, although once again it overlaps the confidence interval with that of the benchmark model.

7. Concluding remarks

The present paper contributes to the debate concerning the effects of the welfare spending on income inequalities. We develop a theoretical model to extract different mechanisms in which institutional (electoral) system and political representativeness lead the allocation of expenditure. Differences in the welfare policies are, hence, explainable through the probability of a center-left or a center-right coalition winning elections. Our model suggests that a majority electoral system produces similar redistributive strategies, whereas differences arising from center-left and center-right coalitions in the proportional representation system.

We exploit this exogenous variation in the allocation of welfare transfers across countries to estimate the

Figure 4: Estimates of the relationship between welfare transfers and GINI index for different time span



welfare-inequality relationship in a panel of OECD countries in causal terms. The empirical results suggest that a larger burden in transfers is related to lower levels of income inequality. Most interestingly, the estimated elasticities indicate that a one percentage increase in welfare transfers reduces the Gini index by about half percentage point, regardless of whether we use imputed or non-imputed data. When we perform the estimates for the components of welfare, we still find that the dimension of the impact of not in kind, in kind transfers and pensions is close to the aggregate result. This result is found to be robust also when the analysis is restricted in order to distinguish countries with high and low levels of income inequality. Also while we find that the effect of welfare transfers is roughly doubled in countries with high income inequality, the estimated elasticity measures for in-kind and not in-kind benefits, in both the samples considered, rely on the confidence interval of the aggregate spending.

While it is intuitively plausible that the political instruments is exogenous, it must also satisfy the exclusion restriction: shocks due to the center-left or center-right government in different electoral systems should affect income inequality only through welfare transfers. We acknowledge the possibility that transmission channels other than welfare transfers *per se* (i.e., interaction between center-left coalitions and unions) may be key underlying causes of income inequality in the aftermath of electoral systems and coalitions. The corporatist nature of labor market institutions and the close interaction between unions and left-wings parties in many OECD countries favor labor market agreements with business organisations, focused in particular

Table 9: Estimates of the relationship between welfare transfers and GINI index, non imputed data (IV estimates; instrument: ES_W)

	$GINI_{GII}$ I		$GINI_{GII}$ II		$GINI_{GII}$ III		$GINI_{GII}$ IV	
Wp_T	-0.447	***						
	(0.133)							
	-0.420	***						
	[0.125]							
Wp_{NIK}			-0.839	***				
			(0.256)					
			-0.423	***				
			[0.129]					
Wp_{IK}					-0.956	***		
					(0.313)			
					-0.418	***		
					[0.137]			
Wp_P							-1.194	**
							(0.511)	
							-0.324	**
							[0.139]	
Time trend	yes		yes		yes		yes	
Fixed effects	yes		yes		yes		yes	
Covariates	yes		yes		yes		yes	
Adjusted R^2	0.810		0.784		0.788		0.716	
N	228		228		228		225	
Kleibergen-Paap F test statistic†	32.587		18.436		31.026		10.759	
	(0.000)		(0.000)		(0.000)		(0.001)	
Cragg-Donald F statistic	61.090		32.805		63.939		26.808	
	(0.000)		(0.000)		(0.000)		(0.001)	

Notes: ES_W is the interaction variable between the electoral representation system and the political parties or coalition winning the elections. The elasticities are shown in bold. The robust standard error are listed in parenthesis. The asterisks give *p-value* significance levels: * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$. We report first stage F-statistics and Wald statistics based on Cragg and Donald (1993) and the Kleibergen and Paap (2006) generalisation to non-independently and non-identically distributed errors, along with the associated $p-values$ for weak-instruments hypothesis tests (Bazzi and Clemens 2013). † Confidence intervals for the Kleibergen-Paap F test statistic follow Bazzi and Clemens (2013).

to blue-collar workers, which are addressed to enhance redistributive policies. Unfortunately, we do not have reliable cross-country data on these other intermediate channels. There are, however, unionisation figures for approximately half of our sample period from the OECD, and we find that center-left government in different electoral systems is not significantly associated with unions, indicating that the strength of unions, when left-wings parties are at government, are unlikely to be driving our findings[4].

Another serious violation of the exclusion restriction is the possibility that the type of electoral system might directly affect income inequality independently of welfare conditions. For instance, net of the welfare magnitude and proportions, a large number of political parties in a proportional representation system may reduce the quality of the services provided by the government, given a likely growing inefficiency in the decision-making process. Note that this possibility is not a serious threat to our estimation strategy, since the proportional electoral system is empirically associated with significantly less income inequality in the reduced-form regressions. Thus to the extent that the hypothesised bias exists, our estimates would be lower bounds on the true impact of welfare spending on income inequality. Nonetheless, we acknowledge that we are unable to definitively rule out the possibility that electoral systems and coalitions could have some independent impact on income distribution beyond its impact working through welfare transfers, though we believe that these other effects are likely to be minor.

[4]Results are available upon request.

Table 10: Estimates of the relationship between welfare transfers and income inequality, $Gini_{OECD}$ (IV estimates; instrument: ES_W)

	$Gini_{OECD}$ I		$Gini_{OECD}$ II		$Gini_{OECD}$ III		$Gini_{OECD}$ IV	
Wp_T	-0.608	***						
	(0.100)							
	-0.657	*******						
	[0.078]							
Wp_{NIK}			-1.103	***				
			(0.210)					
			-0.728	*******				
			[0.127]					
Wp_{IK}					-1.356	***		
					(0.242)			
					-0.591	*******		
					[0.085]			
Wp_P							-3.061	***
							(0.875)	
							-0.754	*******
							[0.274]	
Time trend	yes		yes		yes		yes	
Fixed effects	yes		yes		yes		yes	
Covariates	yes		yes		yes		yes	
Adjusted R^2	0.791		0.760		0.720		0.283	
N	282		282		282		275	
Kleibergen-Paap F test statistic†	31.146		16.711		28.010		5.671	
	(0.000)		(0.000)		(0.000)		(0.001)	
Cragg-Donald F statistic	61.787		26.730		90.294		13.302	
	(0.000)		(0.000)		(0.000)		(0.001)	

Notes: ES_W is the interaction variable between the electoral representation system and the political parties or coalition winning the elections. The elasticities are shown in bold. The robust standard error are listed in parenthesis. The asterisks give *p-value* significance levels: * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$. We report first stage F-statistics and Wald statistics based on Cragg and Donald (1993) and the Kleibergen and Paap (2006) generalisation to non-independently and non-identically distributed errors, along with the associated $p-values$ for weak-instruments hypothesis tests (Bazzi and Clemens 2013). † Confidence intervals for the Kleibergen-Paap F test statistic follow Bazzi and Clemens (2013).

Appendix A: Multiple imputation.

Analyzing data with missing values is a serious problem in many fields especially in economic and social sciences, where the researchers deal with cross sectional, time series or time series cross sectional (TSCS) datasets.

Briefly we describe the iterative procedure that allows us to make multiple imputations in our paper following King *et al.* (2001); Honaker and King (2010). If the assumption of missing randomly (MAR) held and missingness matrix (M) doesn't depend on $\theta = (\mu, \Sigma)$ we can write the following expression:

$$p(D^{OBS}M/\theta) = p(D^{OBS}/\tau)p(M/\theta) \tag{21}$$

and derive the likelihood function of the observed data:

$$L(\theta/D^{OBS}) \propto p(D^{OBS}/\theta) \tag{22}$$

Moreover the multiple imputation allows us to incorporate knowledge via prior function on individual missing cell values. If, for example, a flat prior on θ is assumed, the posterior function is defined as $p(\theta/D^{OBS}) \propto p(D^{OBS}/\theta) = \int p(D/\theta)dD^{MIS}$. In order to draw from this posterior, Honeker et al. (2010) propose the use of a combination of a Expectation maximization (EM) algorithm, Dempster et al. (1977), with a bootstrap technique called EMB algorithm that works as follows. For each step m samples of size n with replacement from our dataset D are bootstrapped and then the EM algorithm is applied to find the posterior mode for bootstrapped data in order to obtain a punctual estimation for θ. Then imputations are made for each set of estimates by drawing values of D^{MIS} from its distribution conditioned on D^{OBS}. The result is m multiple datasets that can be used for subsequent analyses. If the use of bootstrapping methods occurs under regularity conditions, the bootstrapped estimation of θ has the correct properties.

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