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Women's voice on redistribution: From gender equality to equalizing taxation[☆]

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ABSTRACT

We investigate the role played by gender equality in redistributive policies through taxation and in preferences for redistribution. First, at the cross-country level, we study how the historical roots of gender equality –i.e., the time of women's enfranchisement and the role of women in the family– are related with the level of redistribution through taxation. We find that in countries that are historically more gender equal the tax system today is more redistributive. Second, at the individual level, in order to shed light on the cross-country evidence, we investigate whether gender equality is related with overall and/or gender-specific differences in preferences about redistribution. We find that in more gender-equal countries gender differences in redistributive preferences are significantly larger, while their average level does not vary: when gender equality is stronger, women are systematically more favorable to redistribution, whereas there are no significant changes for men. In turn, the component of gender equality driving this result is –not surprisingly– the equality of women and men within the political sphere.

1. Introduction

Redistribution is a major goal of public intervention, which largely takes place through a progressive tax system. In a representative democracy, policies are (or should be) a function of citizens' preferences, which differ across various dimensions, including gender. Both political institutions and preferences might have a strong inertial component, i.e., history could matter. In this paper, we examine the link between gender equality –and its historical roots– and redistribution.

The analysis we provide is two-fold. First, at the cross-country level, we study how the historical roots of gender equality –i.e., the time of women's enfranchisement and the role of women in the family (Todd, 1987)– are related with the level of redistribution through taxation. This aggregate analysis, which is based on 143 countries during the 2000–2016 period, shows that the historical roots of gender equality are significantly associated with the current redistributive features of tax systems: in countries with earlier women's enfranchisement and/or where women historically played a more relevant role within the family, the share of direct taxes

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(over indirect taxes or total tax revenue) is significantly higher than in countries with a later enfranchisement and with historical gender inequality within the family.

Second, we provide additional evidence at the individual level that might shed light on the black box of this cross-country partial correlation. From this point of view, we focus on the demand side of economic policies, i.e., redistributive preferences of citizens, as revealed by survey data. In turn, two different –but not mutually exclusive– mechanisms might play a relevant role. According to the first mechanism, a higher degree of gender equality might be linked to significantly stronger redistributive preferences, on average. This is because equality in the gender domain may ‘‘translate into’’ a drive for equality in the economic domain, thereby similarly influencing the preferences of both women and men. The second mechanism is that gender equality might be correlated in a disjoint fashion with the redistributive preferences of women and men, i.e., with a larger or smaller gender gap in redistributive preferences. Our individual level analysis, which is based on survey data from the European Social Survey (ESS) for 34 European countries during the 2002–2016 period, is consistent with the latter mechanism, while it rejects the former one. Our main finding is that in more gender-equal countries gender differences in redistributive preferences are significantly larger, whereas their average level does not vary. Furthermore, the component of gender equality driving this result is –not surprisingly– the equality of women and men within the political sphere: in other words, our evidence highlights that women are significantly more favorable to redistribution vis-à-vis men in *politically* more equal environments.

We strengthen the link between the cross-country and individual level analyses by adopting an Instrumental Variable specification, where we use the timing of women's enfranchisement and the historical role of women within the family as instruments for gender equality –and for gender equality in the political domain– nowadays.

Our individual-level results are consistent with the so called ‘‘resource hypothesis’’, as laid out by Falk and Hermle (2018), i.e., women are inherently inclined to prefer redistribution more than men –possibly because of a (social) insurance/risk aversion motive–, and that in more gender-equal societies this gender-based difference in redistributive preferences is ‘‘socially’’ allowed to emerge.[1] We find that the larger difference in redistributive preferences between women and men is driven by women being more favorable to redistribution when gender equality is higher, with no significant variation for men. This is consistent with women in more gender-equal environments becoming more vocal about redistribution which would benefit them, since on average they are poorer than men and/or are more altruistic (Croson and Gneezy, 2009).[2] On the other hand, our results do not support an alternative hypothesis, which refers to the role of men: when men feel ‘‘endangered’’ by more gender equality they react by being more adverse to redistribution. In other words, men are more focused on issues that can block the potential rise of women's voice.[3]

On the whole, our findings are in line with the predictions of standard political economy models. For example, within the median voter model, where the (one-dimensional) policy platforms proposed by candidates are about the overall amount of redistribution, citizens might differ in their preferences as a function of multiple factors, which could include their ideological position and their gender. If preferences for redistribution of women and men (differentially) change as a function of gender equality, then the preferences of the median voter could be affected, and thus induce candidates to alter their equilibrium policy platforms accordingly, i.e., with more gender equality the median voter could become more favorable to redistribution and thus the political/policy outcome would amount to a higher level of redistribution.[4]

This paper touches on three different, but strictly linked, strands of the literature.

First, it intends to add further elements to the vast literature on the political economy of taxation and the welfare state, since women's representation in the public sphere and in politics has, to a certain extent, revolutionized policy-making in the last century. More specifically, this literature has examined the impact of both the enlargement of the tax base and the gender composition of the electorate after the introduction of female suffrage, and shown an effect of significantly altering public spending choices, both in size and scope. This is the case, because women are more likely to request and vote for government spending in ‘‘gender-sensitive policies’’, i.e., health, child support, social protection, education, and welfare (Lindert, 1994; Lott and Kenny, 1999; Abrams and Settle, 1999; Aidt et al., 2006; Aidt and Dallal, 2008; Miller, 2008; Bertocchi, 2011; Carruthers and Wanamaker, 2015). Within this literature, the closest contribution to ours is the paper by Aidt and Jensen (2009), which explores the effect of the extension of the women's suffrage on tax composition in Western Europe during the 1860–1938 period.[5]

Second, we contribute to the literature on the determinants of women's empowerment and gender outcomes that has emphasized the role of history and its legacy through time (Alesina et al., 2013; Bertocchi and Bozzano, 2016; Hansen et al., 2015), of informal institutions, namely culture (Guiso et al., 2006; Fernandez and Fogli, 2009; Givati and Troiano, 2012) and religion, shaping the perception of the historical role of women within the family and in society in general (Guiso et al., 2003; Algan and Cahuc, 2006;

[1] According to the alternative ‘‘social role’’ hypothesis, gender-based differences in preferences might be attenuated in more gender-equal societies, since the differences between women and men in their social roles are reduced.

[2] Traditionally, women are found to prefer more redistribution than men (Alesina and Giuliano, 2011; Luttmer and Singhal, 2011). At the aggregate level, Nelson and Goel (2023) investigate the relationship between gender and income equality, by analyzing a panel of 150 countries during the 1985–2019 period.

[3] There is evidence on this mechanism in other spheres of inequality, e.g., ethnic diversity (Alesina and La Ferrara, 2005a), migration (Alesina et al., 2021), and pensions (Galasso and Profeta, 2002).

[4] More generally, political economy models that essentially deal with multidimensional policy spaces –such as the probabilistic voting model (Dixit and Londregan, 1996) and the citizen candidate model (Besley and Coate, 1997, 2003)– would naturally imply that changes in the redistributive preferences of all or some groups within society do affect the political and policy equilibrium: this would also happen when the policy space is one-dimensional, as it is implicit to our setting.

[5] Kose et al. (2021) also highlight a persistent, long run effect of women's enfranchisement on children's education. See Hessami and Lopes da Fonseca (2020) for a comprehensive review of this literature.

Bozzano, 2017). Within this field of studies, a central role is played by the influence of family ties and structures and their persistent effects on gender inequality (Alesina and Giuliano, 2010; Bertocchi and Bozzano, 2019).

Finally, our paper is also related to the growing literature on the determinants of the demand for redistribution and gender-specific preferences. There is overwhelming evidence about a widening gender gap in redistributive preferences, with women (i) being more favorable to redistributive policies (among others, Alesina and La Ferrara, 2005b; Alesina and Giuliano, 2010; Luttmer and Singhal, 2011; Ranehill and Weber, 2022, in an experimental setting) and (ii) leaning increasingly more to the left of the ideological spectrum, both in the U.S. and in Europe (Inglehart and Norris, 2000; Edlund and Pande, 2002).[6]

The paper is organized as follows: in Section 2 we describe the datasets we employ, while in Sections 3 and 4 we present our cross-country and individual-level results, respectively. Section 5 concludes.

2. Data description

Our empirical analysis is based on both aggregate cross-country data and individual-level survey data.[7] First, we build a new cross-country dataset spanning the 2000–2016 period for 143 countries in the world. As anticipated, we first investigate the redistributive features of the revenue side of the public budget as a function of the historical roots of gender equality. Our dependent variable is the ratio of direct taxes over indirect taxes, as computed from ICTD/UNU-WIDER Government Revenue Dataset.

Since we are interested to capture the equalizing effect of the tax system, this ratio conveniently embodies the level of redistribution created by the tax mix. Direct taxes are typically, if not intrinsically, progressive, because of the equalizing effects of tax brackets, fixed deductions and allowances, whereas indirect taxes are generally regressive. Since the equalization of disposable income is achieved through greater progressivity of the tax system, direct taxation should be preferred to indirect taxes to raise revenue and achieve redistribution (Sah, 1983; Saez, 2004). Thus, a higher level of redistribution is more likely to occur (ceteris paribus) when the balance between direct and indirect tax revenue tilts towards the former. In our case, the higher the ratio, the higher the equalizing effect of taxation on income.[8]

Historical roots of gender equality are captured by two different variables which reflect political and social factors: (i) the timing of women's electoral enfranchisement and (ii) the degree of feminism in historical family structures. The plausibly exogenous variation in both variables helps us reducing the recurring endogeneity problem that might afflict the link between gender equality and redistributive policies, both at cross-country and individual level.

The franchise extension to women is a central event in the democratization process: it has been adopted over a long period of time across different countries. We measure it by the date in which the suffrage has been extended to women for the first time, in some cases with restrictions on income, age, or education level.[9] In particular, we use the date of the partial recognition when it is granted to a reasonable representative share of the female population, i.e., educated women aged 30 or more.[10]

Regarding the degree of feminism in historical family structures, we construct a specific variable which builds on Todd's analysis on the status of women within a family system, dating back to the Middle Ages (Todd, 1987). Todd (1987) develops an anthropological classification of family types focusing on the status of women within a family system, i.e., the relationship between husband and wife. This status is in turn defined according to two dimensions: (i) the transmission of property via males only, females, or both, and (ii) the symmetric versus asymmetric importance of father and mother in the procreation of a child. Those dimensions are summarized into a trichotomic variable that measures the degree of feminism: maximum under matrilinearity, to varying degrees under bilaterality, and minimum under patrilinearity. Todd further qualifies those categories by identifying some different sub-groups within each of them. Bilaterality might be full and strongly pro-feminist, thus showing a very high status of women within the family, as it happens in the case of matrilinearity; alternatively, it might show a matrilineal or patrilineal bias, where the status of women is higher in the former than in the latter. Finally, patrilinearity might be weak, medium, or strong, which corresponds to a decreasing status of women within the family.

Accordingly, we code a categorical variable – degree of feminism - that ranges from 0 to 1, based on this taxonomy. A value of 0 corresponds to the lowest gender-equal family types, i.e., strong patrilinearity and polygyny, 0.2 to medium patrilinearity, 0.4 to weak patrilinearity, 0.6 to bilaterality with a paternal bias, 0.8 to bilaterality with a maternal bias, while a value of 1 corresponds

[6] However, results in laboratory experiments are mixed since in certain settings men proved to manifest stronger preferences for redistribution than women (e.g. (Checchi and Filippin, 2004; Assandri et al., 2018; Beraldo et al., 2022)). Mengel and Weidenholzer (2022) recently provided an extensive review of the literature on preferences for redistribution.

[7] While the analysis of country-specific policies is done at the world level, because of data availability the analysis of individual preferences is confined to countries participating to the European Social Survey.

[8] Data on direct and indirect taxes are the most consistent and complete across countries, and provide explicit information on the redistributive character of the tax systems. Among others, Martinez-Vazquez et al. (2011) use the ratio of direct to indirect taxes as a measure of the redistributive effect of tax systems. By applying both an OLS and an Instrumental Variable specification to a sample of 116 developed, developing, and transitional countries over the 1972–2005 period, they show that income inequality is significantly lower when the ratio between direct to indirect taxes is higher, especially so for developed countries. Moreover, see (Inchauste and Karver, 2018; Maier and Ricci, 2022; Hérault and Jenkins, 2022) for estimates of the regressive effects of indirect taxation.

[9] We retrieve this variable from the World Economic Forum Global Gender Gap Report 2006. Data are double-checked with http://womensuffrage.org/.

[10] We use the year of suffrage extension for state or national elections. In other terms, we disregard the franchise extension in local elections if it happened before the one for state elections.

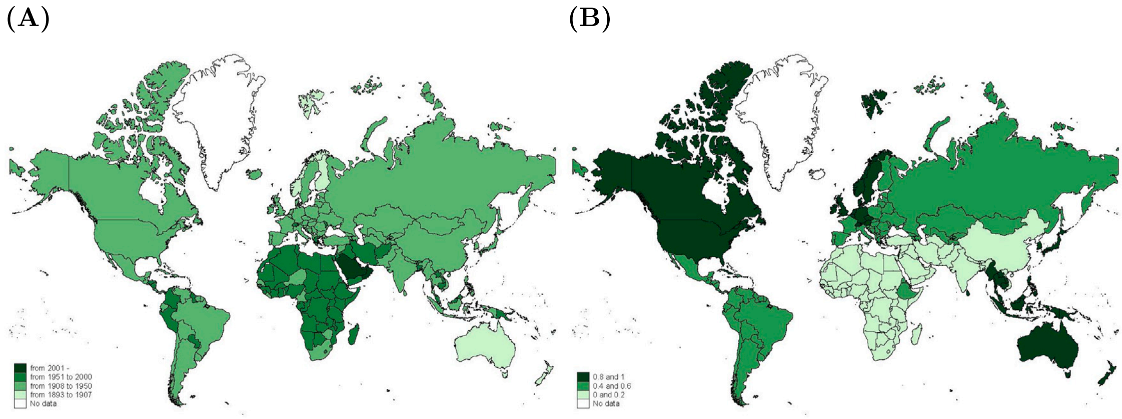

Fig. 1. Women's suffrage and degree of feminism around the world.
Panels (A) and (B) show world maps visualizing the timing of women's enfranchisement and the degree of feminism, respectively. The darker the shade of green for a country, the more recently voting rights to women were recognized in that country in panel (A), and the stronger the degree of feminism in family structure in panel (B). (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

to the family structures characterized by the highest status of women, i.e., full bilaterality and matrilinearity. Therefore, an increase in this variable has to be interpreted as a more relevant role of women within the family.[11]

Panel A of Fig. 1 shows the distribution in the timing of women's suffrage, while panel B displays the geographical variation in the variable degree of feminism around the world. The darker the shade of green for a country, the more recently voting rights to women were granted in that country. The first countries to recognize voting rights to women were New Zealand (1893), Australia (1902), Finland (1906), and Norway (1907). Many countries, mostly located in Europe, did it in the first half of the 1900s. Among the countries in which women's suffrage was guaranteed only in the second half of the 1900s, the first ones were Cote d'Ivoire, Greece, and Lebanon (1952), while the last ones were Iraq (1980), Central African Republic (1986), and Namibia (1989). Finally, the countries that most recently extended the right to vote to women were Bahrain (2002), Kuwait (2005), The United Arab Emirates (2006), and Saudi Arabia (2015).

Regarding the geographical variation in the degree of feminism, the map shows in darker green the countries in which women historically had the most relevant role within the family. More specifically, among the countries in which the feminism variable takes on a value of one we find Austria, Belgium, Cuba, Czech Republic, Germany, Ireland, Israel, Jamaica, Japan, the Republic of Korea, Norway, Sri Lanka, and Sweden. The status of women within the family was important –i.e., the feminism variable takes on a value of 0.8– in Australia, Brunei Darussalam, Cambodia, Canada, Fiji, Indonesia, Lao PDR, Malaysia, Myanmar, Netherlands, New Zealand, Papua New Guinea, Philippines, Singapore, Thailand, United Kingdom, and United States. On the contrary, many African, Middle-East and South-Asian countries are among the ones in which women historically had an inferior role within the family (i.e., the feminism variable takes on a value of zero).

We also collect data on standard economic, demographic, and political controls that are usually taken into account in empirical analyses about the revenue side of the public budget (Kenny and Winer, 2006). The economic variables (Eco) –taken from the World Development Indicators– are: per capita GDP, GDP growth rate, the share of oil rents over GDP, the share of trade over GDP. Demographic variables ($Demo$) are: population density (from UN World Population Prospects), the share of total population living in urban areas, the share of young population aged 0–14 and the share of the elderly aged 65 or more, and the labor force participation rate as a percentage of total population aged 15–64 (again from the World Development Indicators). Political variables (Pol) are: the Gastil index of democracy (from Freedom House), the ideological position on economic policy matters of the prime minister's or president's party (from the Database of Political Institutions), and a categorical variable that captures the regime type (taken from the Authoritarian Regimes Dataset).[12]

[11] In his analysis, Todd defines property transmission within the household and the relative importance of parents in the procreation of children as a first axis to categorize different family types. This is however complemented by a second axis which captures the degree of parental authority within the family, as identified by residential habits of different generations within the family itself. Families are classified as "vertical", or "authoritarian", when parents cohabit with their married children, so that three generations cohabit together. On the other hand, families are "non-vertical", or "nuclear", when new married couples automatically leave the parental household. In the robustness checks of our baseline analysis, we combine these two axes proposed by Todd and identify a set of nine family types (increasing in the status of women): patrilineal non-vertical and polygyny; patrilineal vertical (strong); patrilineal vertical (medium); patrilineal vertical (weak); bilateral non-vertical (patrilineal bias); bilateral non-vertical (matrilineal bias); matrilineal vertical; bilateral vertical. Indeed, as Todd suggests, maternal authority is maximum in maternal and bilateral vertical family structures and minimum in patrilineal non-vertical families and polygyny.

[12] The Gastil index is the average of two different indicators, i.e., civil liberties and political rights. Each country is graded on a 7-point rating scale from 1 (the highest degree of freedom) to 7 (the lowest degree of freedom) in both dimensions, according to several aspects, such as the freedom of expression and belief, rule of law, associational and organizational rights, personal autonomy and individual rights, political pluralism and participation, electoral process, and the functioning of the government. The leftist orientation with respect to economic policy of the party of the prime minister or president takes on a value of

Table 1
Summary statistics for cross-country analysis.

	Obs	Mean	Std. Dev	Min	Max
Direct taxes over Indirect taxes	143	0.68	0.48	0.04	3.24
Women's suffrage year	143	1947	22.39	1893	2015
Degree of feminism	143	0.38	0.35	0.00	1.00
Per capita GDP (thousand dollars)	143	11.99	17.75	0.23	102.68
GDP growth	143	4.08	2.70	−6.91	14.16
Oil rents over GDP	143	4.84	11.39	0.00	62.54
Trade over GDP	143	86.66	47.13	0.59	380.91
Population density	143	158.61	586.48	1.60	6803.67
Urban population	143	55.30	22.36	9.92	100.00
Young population 0–14	143	30.29	10.85	13.69	49.38
Elderly population 65 or more	143	7.59	5.37	0.89	21.21
Labor force participation rate	142	67.58	9.83	43.08	88.74
Gastil index	143	4.46	1.89	1.00	7.00
Government orientation	143	1.22	1.09	0.00	3.00
Regime type	143	52.42	45.30	1.00	100.00
English legal origins	140	0.30	0.46	0.00	1.00
French legal origins	140	0.54	0.50	0.00	1.00
Catholicism in 1900	141	0.25	0.38	0.00	1.00
Communism in 1970	141	0.20	0.40	0.00	1.00
Years of interstate conflicts	142	4.67	8.23	0.00	41.00
Plough	142	0.56	0.47	0.00	1.00
Latitude	142	20.00	24.58	−41.81	64.48
Longitude	142	19.98	58.10	−112.98	171.48
Tropical zones	142	35.99	43.52	0.00	100.00
Desert zones	142	4.22	12.52	0.00	77.28
Terrain ruggedness index	142	1.29	1.23	0.02	6.74
Land area	143	739.04	1645.30	0.34	9388.21
Arable land	143	15.08	13.63	0.35	60.97
Ethnic fractionalization	141	0.46	0.25	0.00	0.93
Religious fractionalization	142	0.44	0.23	0.00	0.86
Linguistic fractionalization	138	0.42	0.29	0.00	0.92

Finally, we complement our dataset with a series of potentially confounding historical, geographical, and cultural factors. Historical variables ($Hist$) are: English or French legal origins (Shleifer et al., 2008), the importance of Catholicism in each country in 1900 and the existence of a Communist regime in each country in 1970 (Barro, 2003), the total number of years of interstate conflicts in each country between 1816 and 2007 (Sarkees and Wayman, 2010), and the use of the plough in agriculture as a proxy of the origins of gender roles (Alesina et al., 2013).[13] Geographic variables (Geo) include: latitude, longitude, tropical zones, desert zones, terrain ruggedness index, all from Nunn and Puga (2012), and country's land area and percentage of arable land from the World Development Indicators.[14] Finally, cultural variables ($Cult$) are: ethnic, religious, and linguistic fractionalization indexes (Alesina et al., 2003).[15] Summary statistics for all the variables in the cross-country dataset are shown in Table 1.

The individual-level part of our dataset includes survey data from the European Social Survey (ESS), a biennial, repeated cross-section dataset gathering information on attitudes and behaviors across 34 European countries and over time. Data is drawn from 8 rounds between 2002 and 2016. As dependent variable, we rely on the question –repeated across all waves of the ESS– which most closely captures preferences for redistribution as follows: "To what extent do you agree or disagree with the statement: the government should reduce differences in income levels?". We recode this variable so that the higher the value, the more favorable to redistribution the individual is.[16] We also include a standard set of individual controls: sex, age, legal marital status, the main

1 for right, 2 for center, and 3 for left-wing party. Lastly, Regime type classifies authoritarian regimes around the world according to three modes of political power maintenance: hereditary succession (lineage), corresponding to monarchies; actual or threatened use of military force, corresponding to military regimes; and popular elections, designating electoral regimes. Non-authoritarian regimes are classified as democracies.

[13] The importance of Catholicism is measured by the share of total population that adhered to Catholicism in 1900 in each country, while the proxy of the origins of gender roles is an estimate of the fraction of the population currently living in a district (or country) with ancestors that traditionally engaged in plough agriculture as proposed by Alesina et al. (2013).

[14] Latitude and longitude are expressed in decimal degrees, for the geographical centroid of the country. Tropical zones is the percentage of the land surface of each country where any of the four Köppen–Geiger tropical climates is present, while desert zones is the percentage of the land surface area of each country covered by sandy desert, dunes, rocky or lava flows. The terrain ruggedness index is a measure of topographic heterogeneity which captures the amount of elevation difference between different areas of the same country. Lastly, we refer to country's total area (squared kilometers) and the fraction of land within a country which is arable (FAO definition).

[15] These variables capture the probability that two randomly selected individuals from the same country belong to the same ethnic, religious, or linguistic group respectively. Each index is arranged so that the higher its value, the more fractionalized the country.

[16] The variable is called *gincdif* and originally takes on values on the range from 1 (Agree strongly) to 5 (Disagree strongly). We use data from the ESS rather than from the World Values Survey (WVS) since the latter does not include a direct question on the role governments should play in addressing income inequality, but only a question on poverty reduction.

Table 2
Summary statistics for individual-level analysis.

	Obs	Mean	Std. Dev	Min	Max
Redistributive preferences	210 014	3.92	1.02	1.00	5.00
Female dummy	210 014	0.53	0.50	0.00	1.00
GGGI	210 014	0.15	0.05	0.00	0.29
Per capita GDP (thousand dollars)	210 014	36 607.98	20 958.67	2834.34	91 218.62
Left-wing dummy	210 014	0.10	0.31	0.00	1.00
Right-wing dummy	210 014	0.25	0.43	0.00	1.00
Household total net income	210 014	5.31	2.80	1.00	12.00
No children at home	210 014	1.62	0.49	1.00	2.00
Highest level of education	210 014	3.20	1.29	1.00	5.00
Religiosity	210 014	4.65	3.03	0.00	10.00
Age	210 014	49.07	18.07	14.00	114.00
Age squared	210 014	2734.71	1833.86	196.00	12 996.00
Marital status					
Married	210 014	0.53	0.50	0.00	1.00
Never married	210 014	0.26	0.44	0.00	1.00
Divorced or separated	210 014	0.11	0.31	0.00	1.00
Widowed	210 014	0.10	0.30	0.00	1.00
Main source of household income					
Wages or salaries	210 014	0.59	0.49	0.00	1.00
Self employment or farming	210 014	0.06	0.25	0.00	1.00
Pensions	210 014	0.27	0.45	0.00	1.00
Unemployment benefits	210 014	0.02	0.15	0.00	1.00
Social benefits	210 014	0.03	0.18	0.00	1.00
Investments	210 014	0.01	0.07	0.00	1.00
Other sources	210 014	0.01	0.12	0.00	1.00

source of household income, household total net income, the absence of children at home, and the levels of both education, and religiosity. Moreover, we include a set of 3 dummies which capture the political ideology of the respondent (i.e., left, center, right).[17] Finally, we match these individual-level variables with two country-level variables, i.e., current gender equality and per capita GDP. As a measure of current gender equality we employ the Global Gender Gap Index (hereafter GGGI), devised by Lopez-Claros and Zahidi and computed by the World Economic Forum every year since 2006 (World Economic Forum, 2006). This index is composed by 14 sub-indicators and is able to capture gender-based disparities in outcomes between women and men in four key dimensions: political leadership and empowerment (POLGGGI), economic participation and opportunity (ECOGGGI), educational attainment (EDUGGGI), and health and survival (HSGGGI). The GGGI ranges from 0 to 1, where lower values indicate gender inequality, while values approaching one mean higher gender equality, thus increasing in the degree of equality between the two genders.[18]

Summary statistics for all the variables in the individual-level dataset are shown in Table 2. There are 210,014 respondents, of whom 111,918 are women, i.e, 53.3 percent.[19]

3. Cross-country evidence: Historical roots of gender equality and redistribution

From a policy and aggregate perspective, we are interested in the link between our history-based measures of gender equality and the redistributive content of current tax systems.

We first look at some stylized facts. Panel A of Fig. 2 shows that in countries where the right to vote has been granted to women earlier, the ratio between direct and indirect taxes is higher. For example, New Zealand and Australia extended the suffrage to women in 1893 and 1902, respectively, then followed by Finland in 1906 and Norway in 1907. These countries show a high ratio of direct taxes over indirect ones. On the contrary, countries such as Bahrein, Kuwait, The United Arab Emirates, Saudi Arabia, and

[17] See the Appendix for additional details on these individual controls.

[18] The political leadership and empowerment sub-index (POLGGGI) is composed of 3 variables: the ratio of women with seats in parliament over male value, the ratio of women at ministerial level over male value, and the ratio of number of years of a female head of state (last 50 years) over male value. The economic participation and opportunity sub-index (ECOGGGI) is composed of 5 variables: the ratio of female labor force participation over male value, wage equality between women and men for similar work (converted to female-over-male ratio), the estimated female earned income over male value, the ratio of female legislators, senior officials, and managers over male value, and the ratio of female professional and technical workers over male value. The educational attainment sub-index (EDUGGGI) is composed of 4 variables: the ratio of female literacy rate over male value, the ratio of female net primary level enrollment over male value, the ratio of female net secondary level enrollment over male value, and the ratio of female gross tertiary level enrollment over male value. The health and survival sub-index (HSGGI) is composed of 2 variables: the ratio of female healthy life expectancy over male value and the sex ratio at birth (converted to female over male ratio). The four sub-index scores are calculated as weighted averages of the above variables within each sub-index. For more details on the technical framework and on the sub-indicators included in each dimension of the composite index, refer to WEF (2006). Notice that, in the econometric analysis and related figures, we normalize the GGGI by subtracting from its original value the minimum value it takes on within our sample, i.e., 0.5828.

[19] Regarding political ideology, i.e., the individual-level characteristic that should be most strongly related to preferences for redistribution, there are 11,452 left-leaning and 28,949 right-leaning women, while 71,517 women are centrist. In the case of men, 10,425 are left-leaning, 22,951 are right-leaning, while 64,720 are centrist.

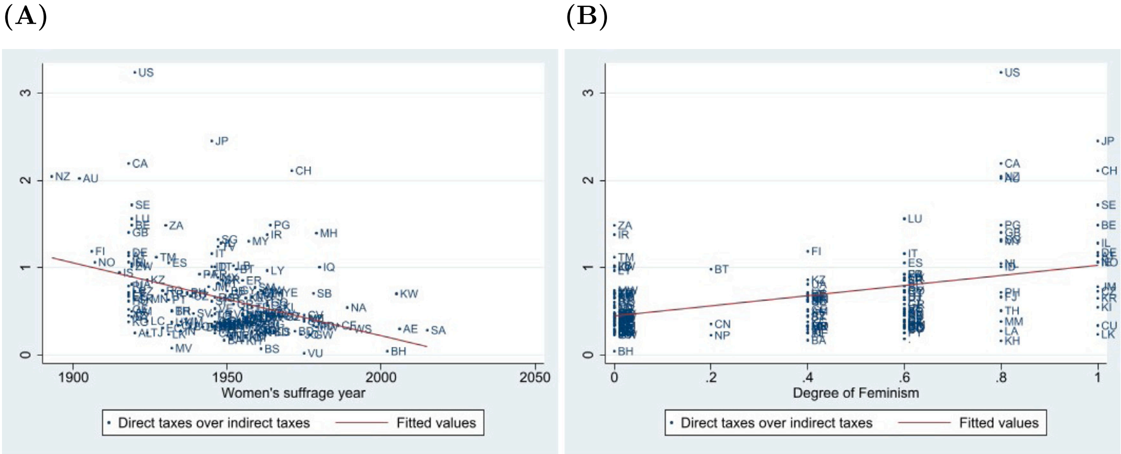

Fig. 2. Women's suffrage, degree of feminism, and direct taxes over indirect taxes.
Panels (A) and (B) show the unconditional correlations between the timing of women's suffrage, or the degree of feminism, and our measure of redistributive policy, i.e., the ratio of direct taxes over indirect taxes.

Brunei Darussalam granted the suffrage after year 2000 and show a low ratio between direct and indirect taxes. Similarly, panel B of Fig. 2 shows that in countries where the role of women within the family was more relevant in the past, i.e., Austria, Belgium, Cuba, Czech Republic, Germany, Ireland, Israel, Jamaica, Japan, the Republic of Korea, Norway, Sri Lanka, Sweden, the ratio between direct and indirect taxation is higher.[20]

In order to investigate the relationship between the average ratio of direct versus indirect taxes and the historical roots of gender equality, we estimate the following OLS model:

$$Y_c = \alpha + \beta_0 \text{Women's suffrage}_c + \beta_1 \text{Degree of Feminism}_c + \beta_2 \text{Eco}_c + \beta_3 \text{Demo}_c + \beta_4 \text{Pol}_c + \varepsilon_c \tag{1}$$

where Y_c is the average ratio between direct and indirect taxes in country c during the 2000–2016 period, Women's suffrage$_c$ and Degree of Feminism$_c$ are our historical explanatory variables, and ε_c is the error term. We apply White–Huber standard errors to deal with potential heteroscedasticity.

We consider two specifications displayed in Table 3. In a more parsimonious specification (column 1) we control for economic variables Eco_c (per capita GDP, GDP growth rate, the share of oil rents over GDP and the share of trade over GDP in country c), demographic variables $Demo_c$ (population density, the share of urban population, the share of young population aged 0–14, the share of the elderly aged 65 or more, and the labor force participation rate in country c), and political variables Pol_c (the Gastil index of democracy, the ideological orientation of the incumbent government, and the regime type in country c). Then, in a more demanding specification (column 2) we also add historical variables $Hist_c$ (English or French legal origins, the relevance of Catholicism in 1900, the existence of a Communist regime in 1970, the total number of years of interstate conflicts between 1816 and 2007, and the use of the plough in agriculture in country c), geographic (Geo_c) and cultural controls ($Cult_c$).[21] Columns 3–4 of Table 3 parallel the structure of columns 1–2 but with the addition of region fixed effects.

When not including region fixed effects, we find that both the timing of the extension of the suffrage to women and the historical role of women in the family have a statistically significant relationship with the ratio of direct over indirect taxes, in the expected direction: in countries where women have been enfranchised earlier and/or women played a relevant role within the family the ratio of direct over indirect taxes is significantly higher. On the other hand, when we add region fixed effects, only the timing of the extension of suffrage to women is significantly –and negatively– correlated with the ratio of direct to indirect taxation. This could be explained by the fact that there is very little variation in family structures within each region, as shown by panel B of Fig. 1.[22] In terms of magnitudes, based on the estimates from column 2 (our most demanding specification), one-standard-deviation increase in the timing of women's suffrage (22.013) is associated with a reduction of the ratio of direct to indirect taxes of 0.152 (22.013 * 0.007). Since the standard deviation of the dependent variable is 0.494, the implied variation is about 30.8 percent of that standard deviation. On the other hand, a one-standard-deviation increase in the relevance of the historical role of women in the family (0.352) is associated with an increase in the ratio between direct and indirect taxes of 0.123 (0.352 * 0.350), i.e., about 24.9 percent of the standard deviation of the dependent variable.

[20] In these countries we observe both bilateral transmission of property and symmetric importance of father and mother in the procreation of the child. Therefore, in these cases the variable degree of feminism takes on a value of one.

[21] More precisely, geographic controls are latitude, longitude, tropical zones, desert zones, terrain ruggedness index, country's land area, and percentage of arable land, while cultural controls refer to ethnic, religious, and linguistic fractionalization indexes respectively.

[22] Looking at the simple pairwise correlations, we notice that regions are strongly correlated with degree of feminism (72 percent), while they are only fairly correlated with women's suffrage year (44 percent).

Table 3
Direct over indirect taxes and historical determinants of gender equality, cross-country analysis.

Direct over Indirect taxes	(1) b/se	(2) b/se	(3) b/se	(4) b/se
Women's suffrage year	−0.007***	−0.007***	−0.010***	−0.008***
	(0.002)	(0.002)	(0.002)	(0.002)
Degree of feminism	0.283**	0.350***	0.075	0.152
	(0.139)	(0.128)	(0.170)	(0.139)
Per capita GDP	0.000***	0.000	0.000***	0.000
	(0.000)	(0.000)	(0.000)	(0.000)
GDP growth	−0.008	−0.024**	−0.002	−0.028**
	(0.011)	(0.011)	(0.010)	(0.011)
Oil rents over GDP	0.009**	0.011***	0.010***	0.012***
	(0.004)	(0.004)	(0.004)	(0.003)
Trade over GDP	−0.001	0.002*	−0.001	0.002*
	(0.001)	(0.001)	(0.001)	(0.001)
Demographic controls	Yes	Yes	Yes	Yes
Political controls	Yes	Yes	Yes	Yes
Historical controls	No	Yes	No	Yes
Geographic controls	No	Yes	No	Yes
Cultural controls	No	Yes	No	Yes
Region FE	No	No	Yes	Yes
Obs.	135	135	135	135
R-squared	0.492	0.711	0.585	0.751

OLS, cross country averages for the 2000–2016 period. The dependent variable is the ratio of direct over indirect taxation. In the first two columns we do not include region fixed effects, while we do so in columns (3) and (4). Demographic controls include population density, urban population, young people 0–14, elderly people over 65, and LFP rate. Political controls include Gastil index, left government, and regime type. Historical controls include English and French legal origins, Catholicism in 1900, Communist regime in 1970, years of interstate conflicts, and plough. Geographic controls include latitude, longitude, tropical zones, desert zones, terrain ruggedness index, country's land area, and percentage of arable land. Cultural controls include ethnic, religious, and linguistic fractionalization. Robust standard errors in parentheses.

*** $p < 0.01$.

** $p < 0.05$.

* $p < 0.1$.

We can also compute the proportion of the total variation in the direct to indirect taxes ratio that our historical gender-related variables are able to explain, controlling for confounding factors. The inclusion of our two historical gender measures increases the R-squared by 0.079 (from 0.7107 to 0.6316): they account for 7.9 percent of the total variation in the ratio of direct to indirect taxes and 21.4 percent of the residual variation that is left unexplained by the control variables.[23]

We also perform a series of additional robustness checks, starting from the specification of column 2 in Table 3, i.e., our most demanding specification without region fixed effects. Our results are reported in Table A.1 in the Appendix.

First, we use as separate dependent variables both direct taxes as a share of total tax revenue and indirect taxes as a share of total tax revenue. We find that our two historical variables are not significantly related to the share of indirect taxes over total tax revenue, while they are strongly related to the share of direct taxes over total tax revenue. This would suggest that in more gender-equal countries the "action" comes through a higher level of direct taxes. Second, we recode the degree of feminism within the family according to a slightly different criterion, thus obtaining a more polarized measure thereof: our results still hold. Third, we replace the variable degree of feminism with a full set of women status' dummies that –from a geographical point of view– mostly coincide with Todd's family types, as widely employed in the recent economic literature (Todd, 1985, 1990; Duranton et al., 2009; Bertocchi and Bozzano, 2015; Galasso and Profeta, 2018; Tur-Prats, 2019).[24] Our results confirm that the family types that are significantly related to a higher degree of feminism (bilateral and matrilineal family types) are in turn correlated with a higher level of direct to indirect tax ratio, as compared to patrilineal family types (both vertical and non-vertical) and polygyny.

Finally, in Appendix Table A.2 we also replicate our cross-country analysis by restricting the sample to the countries that are included in the individual-level analysis. For obvious lack of degrees of freedom, we run a more parsimonious specification than the one used in column (1) of Table 3. We confirm our previous results: both the women's suffrage and the feminism variables are significantly correlated with the ratio between direct and direct taxes in the expected direction.[25]

[23] The inclusion of the timing of women's suffrage increases the R-squared by $0.6924 - 0.6316 = 0.0608$, i.e., it accounts for 6.1 percent of the total variation in our dependent variable and 16.5 percent of the residual variation. On the other hand, the inclusion of the historical role of women in the family increases the R-squared by $0.6723 - 0.6316 = 0.0407$, i.e., 4.1 percent of the total variation in the dependent variable and 11 percent of the residual variation.

[24] See footnote 11 for details about the construction of those dummies for family types.

[25] (P-values are 0.068 and 0.1 respectively).

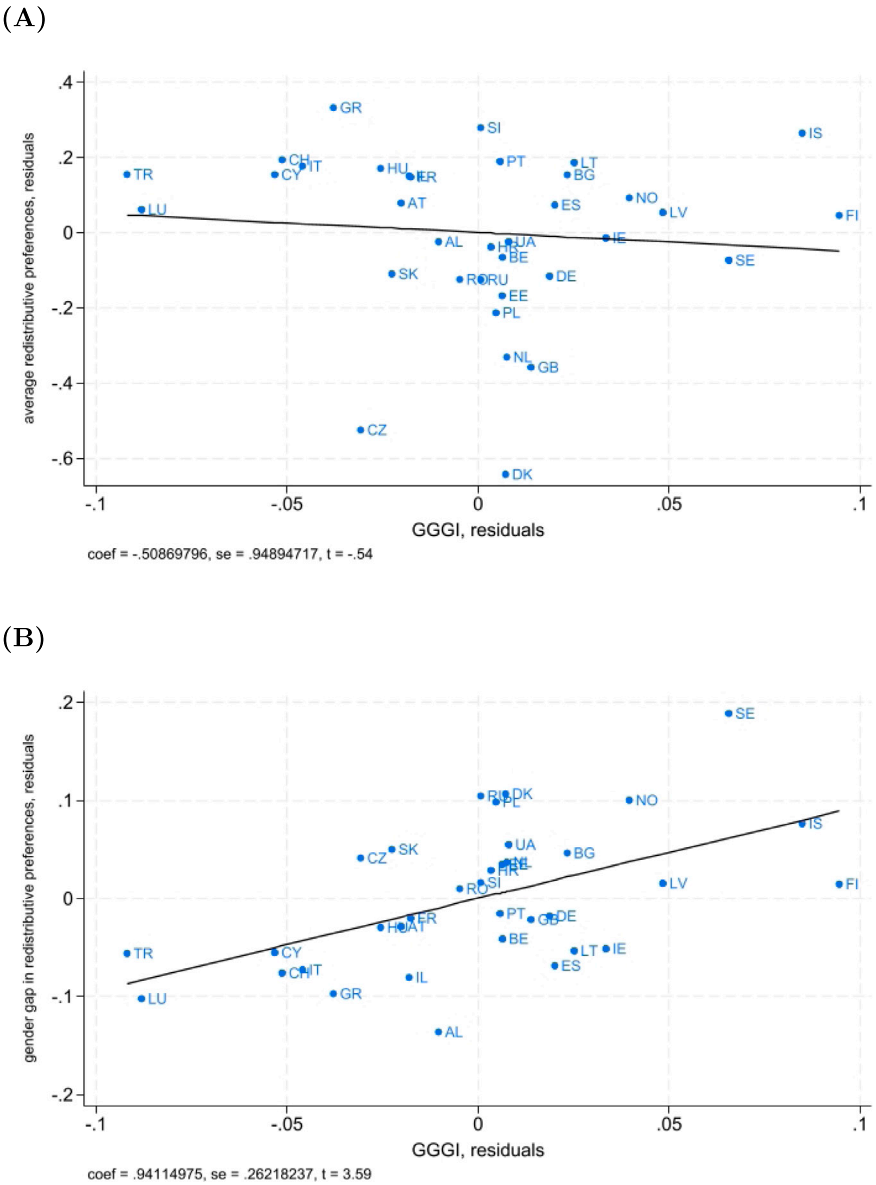

Fig. 3. The Global Gender Gap Index (GGGI) and average and gender-specific preferences for redistribution.
In Panel A we show a scatter plot of average redistributive preferences by country against the GGGI, having partialled out the correlation of both variables with per capita GDP. In Panel B we plot the gender gap in redistributive preferences against the GGGI, again having partialled out the correlation with per capita GDP. We focus on the countries that are included in the ESS survey.

4. Individual-level analysis: Gender equality and preferences for redistribution

We now turn to an individual-level analysis of redistributive preferences, with a focus on the role played by gender equality.

Our purpose is to shed light on the cross-country positive correlation between historically determined gender equality and redistributive taxation, by focusing on the demand side of economic policies, i.e., redistributive preferences of citizens. There are two mechanisms that might be at play. According to the first mechanism, in more gender-equal environments both women and men are more favorable to redistribution, which of course implies a positive correlation between gender equality and average redistributive preferences. According to the second mechanism, gender equality is differentially associated with the redistributive preferences of men and women, thus implying a larger or smaller gender gap in those preferences.

In panel A of Fig. 3, for the European countries that are included in ESS survey, we show a scatter plot of average redistributive preferences by country against the GGGI, having partialled out the correlation of both variables with per capita GDP. Per capita GDP is a relevant confounding factor because it is strongly and negatively correlated with average redistributive preferences, while it is strongly and positively correlated with the GGGI. This relationship between gender equality and average redistributive preference is not statistically significant at ordinary confidence levels (p-value of 0.59). On the other hand, in panel B of Fig. 3, we repeat the same exercise with the gender gap in redistributive preferences as the dependent variable. In this case the relationship is positive and strongly significant, i.e., the gender gap in redistributive preferences is significantly larger in more gender-equal countries (p-value of 0.001).

We check whether this explorative result is confirmed by the analysis of individual-level survey data. More precisely, we use the ESS data to investigate (i) whether redistributive preferences are on average related with the societal climate about gender, as captured by the overall degree of gender equality, (ii) to what extent those preferences are gender specific, and (iii) whether any gender-specific difference in preferences is significantly related to the overall degree of gender equality.

To deal with these empirical questions we use pooled OLS models, i.e., we merge the various years of the ESS survey and add year-of-interview fixed effects. However, similarly to Alesina and Giuliano (2011) and Luttmer and Singhal (2011), we do not include country fixed effects: the rationale behind this specification is that a relevant part of the variation in gender equality is *between* countries, with much less year-to-year variation *within* countries.

Formally, we estimate the following specifications:

$$y_{it} = \alpha_{it} + \beta_0 \text{Female}_{it} + \beta_1 \text{GGGI}_{ct} + \beta_2 \text{pc GDP}_{ct} + \beta_3 \text{Ideology}_{it} + \beta_4 X_{it} + \eta_t + \varepsilon_{ict} \tag{2}$$

$$\begin{aligned} y_{it} = {} & \alpha_{it} + \beta_0 \text{Female}_{it} + \beta_1 \text{GGGI}_{ct} + \beta_2 \text{pc GDP}_{ct} + \beta_3 \text{Ideology}_{it} + \beta_4 X_{it} + \beta_5 (\text{Female}_{it} * \text{GGGI}_{ct}) + \beta_6 (\text{Female}_{it} * \text{pc GDP}_{ct}) \\ & + xs\beta_7 (\text{Female}_{it} * \text{Ideology}_{it}) + \eta_t + \varepsilon_{ict} \end{aligned} \tag{3}$$

The dependent variable y_{it} is the answer to the survey question about whether – according to the respondent – the government should reduce income differences among citizens. Thus, it measures the preferences for redistribution of respondent i in interview-year t. Female_{it} is the female dummy for respondent i in year of interview t, while both GGGI_{ct} and per capita GDP_{ct} refer to the respondent's country of residence c in year of interview t. Notice that we normalize the GGGI – our measure of the gender environment at the country-year level– by subtracting from its original value the minimum value it takes on within our sample, i.e., 0.5828. Ideology_{it} stands for the Left-wing and Right-wing dummies for respondent i in year of interview t, with ideologically moderate respondents as the excluded category. X_{it} is a set of other individual-level features, i.e., household total net income, a dummy for no children at home, highest level of education, religiosity, age, age-squared, legal marital status, and the main source of household income. Finally, η_t stands for fixed effects that are year-of-interview specific. Standards errors are clustered at the respondent's country of residence level.

First, we are interested in the coefficients β_0 on the female dummy and β_1 on GGGI. Indeed, β_0 captures overall gender-specific differences in redistributive preferences, while β_1 captures whether average redistributive preferences are related to the GGGI. Second, and more importantly, when we include the interaction between the female dummy and the GGGI we are able to look at $\beta_0 + \beta_5 * \text{GGGI}_{ct}$: if this expression is significantly positive and increasing as a function of the GGGI, the gender gap in redistributive preferences would be significantly larger in more gender-equal environments. The opposite holds if this expression is decreasing. We perform a similar exercise with per capita GDP, properly interacted with the female dummy.

As a collateral analysis, we are also interested in the correlation between (self-reported) ideology and redistributive preferences, with the purpose of checking whether ideology differently affects women and men.

To perform this set of analyses, we step-wise include those interaction effects of the female dummy with the GGGI, per capita GDP, and ideology, first one by one and then in a joint fashion, as shown in Eq. (3).

The outcome of the baseline regression is shown in column 1 of Table 4: we find that women are significantly more favorable to redistribution, but there is no statistically significant relationship between the GGGI and average redistributive preferences, consistently with the aggregate pattern shown in panel A of Fig. 3.

However, when we add the interaction term between the female dummy and the GGGI (column 2), we find that gender-specific differences in redistributive preferences are larger when the GGGI is higher, i.e., in more gender-equal environments. Moreover, this variation is driven by women being more favorable to redistribution in those environments, while there are no significant changes for men. Following Brambor et al. (2006), we provide graphical summaries of the partial correlations computed from our regression results. Panel A of Fig. 4 shows that the partial correlation of the female dummy with pro-redistribution preferences is almost always positive and increasing with the GGGI, i.e., in more gender-equal environments the (positive) difference between women and men in their pro-redistribution attitudes is significantly larger. This difference is statistically insignificant in environments with the lowest level of gender equality.

In terms of magnitudes, being a woman is not significantly associated with stronger preferences for redistribution at the minimum level of the GGGI, while the difference between women and men in redistributive preferences is 0.106 at the average level of the GGGI, and 0.209 at the highest value of the GGGI: both differences are significantly different from zero. In relative terms, since the standard deviation of the dependent variable is 1.02, these differences for the average and the maximum level of the GGGI are about one-tenth and one-fifth of that standard deviation, respectively.

Overall, our findings are consistent with the "resource hypothesis", i.e., gender differences in preferences –in our case preferences for redistribution– are wider in more gender-equal contexts (Falk and Hermle, 2018). Digging further, we highlight that this increased gender cleavage is driven by women being more favorable about redistribution in more gender-equal environments, rather than men moving in the opposite direction and becoming less favorable to redistribution.

In what follows, we also explore the role of per capita GDP, by adding an interaction term between the female dummy and per capita GDP as shown in column 3 of Table 4. We find that the gender-based cleavage in redistributive preferences is positively

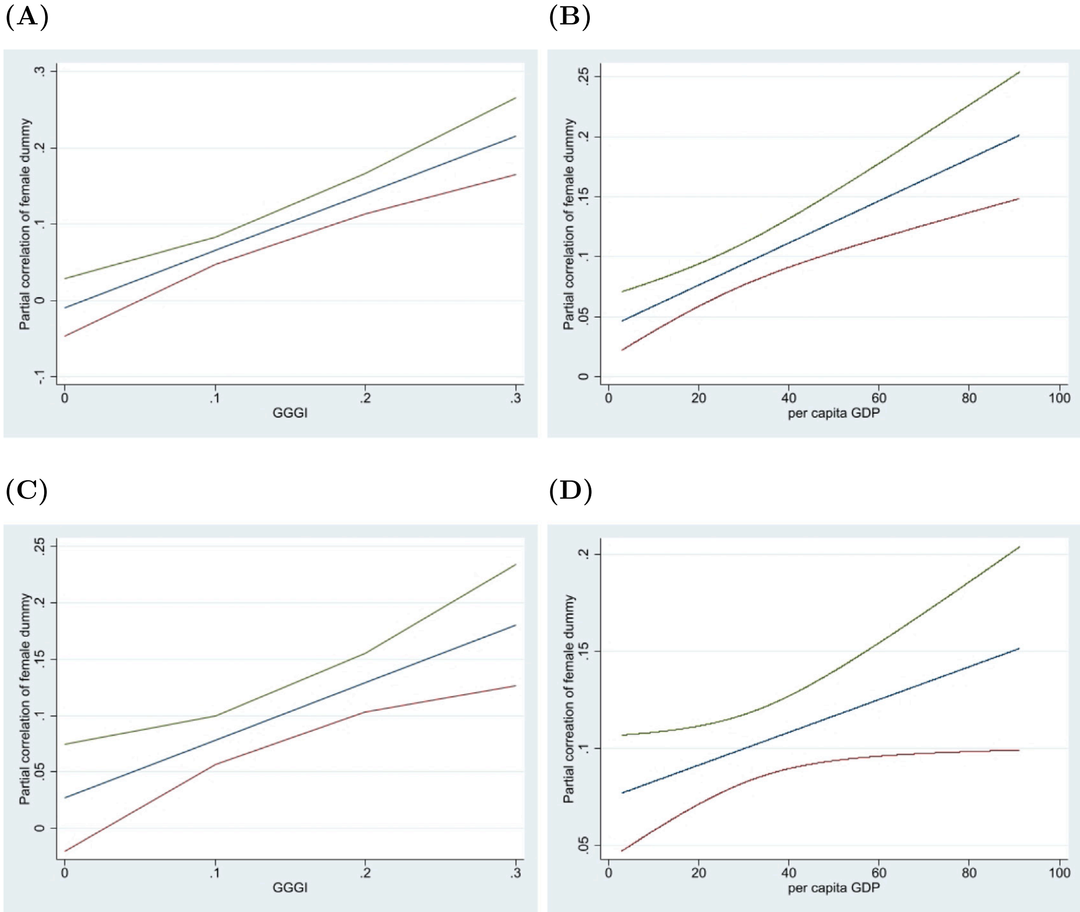

Fig. 4. Estimated partial correlations of the female dummy with redistributive preferences, as a function of GGGI and per capita GDP.
In panels (A) and (B) the blue line represents the estimated partial correlation of the female woman dummy in the regression for redistributive preferences –as computed from Table 4, columns (2) and (3) respectively–, while the red and green lines identify the upper and the lower bounds of the 95% confidence intervals. In panel (A) the partial correlation is a function of GGGI, while in panel (B) it is a function of per capita GDP. The estimated partial correlation of the female dummy that are shown in panels (C) and (D) are computed from column (4) in Table 4, and are functions of GGGI and of per capita GDP, keeping per capita GDP and GGGI at their mean values, respectively. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

and significantly associated with per capita GDP. Moreover, the partial correlation of the female dummy with pro-redistribution preferences is always positive and increasing with per capita GDP, as shown in panel B of Fig. 4.

Since the coefficients of the interaction terms of the female dummy with the GGGI and per capita GDP are both positive and significant, in column 4 of Table 4 we explore a more demanding specification which includes them both. One relevant concern here is that the GGGI would simply be a proxy for economic development, as measured by per capita GDP. Interestingly, the interaction between the female dummy and the GGGI displays a smaller coefficient, but still significant at the 5 percent confidence level, while the interaction with per capita GDP –albeit with a positive coefficient– is no longer statistically significant at ordinary confidence levels. However, panel C of Fig. 4 allows us to show that, when per capita GDP takes on its mean value, the estimated partial correlation of the female dummy is always positive and increasing with GGGI, still confirming that in more gender-equal environments the (positive) difference between women and men in their pro-redistribution preferences is significantly larger. At the same time, when the GGGI is at its mean value, the estimated partial correlation of the female dummy is almost always positive and increasing with per capita GDP (panel D of Fig. 4).

Columns from 5 to 8 of Table 4 parallel the structure of columns from 1 to 4 but with the addition of the interaction terms between the female dummy and both the left-wing dummy and the right-wing dummy.

Unsurprisingly, self-declared left-leaning respondents are more favorable to redistribution than moderate respondents, that in turn are more favorable to it than right-leaning respondents (Alesina et al., 2018; Jaeger, 2008). However, there is no gender-based difference in preferences for redistribution among left-leaning respondents, while ideologically moderate women are more favorable

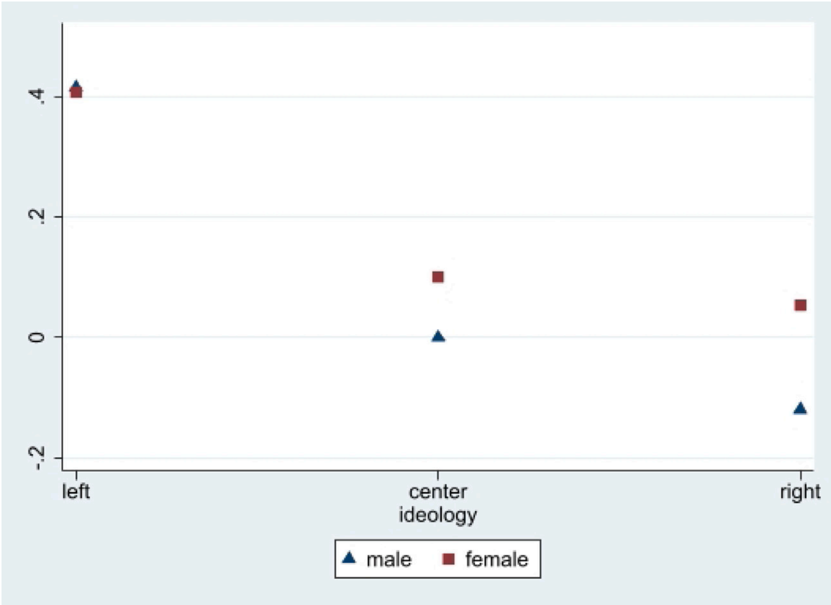

Fig. 5. Gender, ideology, and redistributive preferences.
On the basis of column (5) in Table 4, we show the estimated differences in the preferences for redistribution for women and men, conditional on their ideological leaning. Taking centrist men as the excluded category, the triangles show the estimated differences in those preferences for right-wing and left-wing men, while the squares represent the relative preferences about redistribution for women, again as a function of self-reported ideology.

to redistribution than moderate men, and this effect is even stronger among right-leaning individuals (see Fig. 5, based on column 5 of Table 4). Put in other terms, the ideology-led difference in redistributive preferences among women is much less pronounced than among men.[26]

Interestingly, when in column 6 of Table 4, we add the interaction terms between the female dummy and ideology together with the interaction term between the female dummy and the GGGI, the partial correlations show that left-leaning women are significantly *less* favorable to redistribution at the minimum level of GGGI, but significantly *more* favorable at its maximum level (see Figure A.1A in the Appendix). On the other hand, centrist women are significantly more favorable to redistribution especially at the maximum level of the GGGI (see Figure A.1B in the Appendix). Finally, for right-leaning respondents the partial correlation between being a woman and redistributive preferences is positive and significant at all levels of the GGGI, i.e., at the minimum level as well (see Figure A.1C in the Appendix).[27]

In column 8 we run our most-demanding specification in which we include all the above interactions. The partial correlations of the female dummy estimated from column 8 of Table 4 are reported in Fig. 6. In panels (A), (B), and (C) the estimated partial correlations of the female dummy are functions of GGGI and per capita GDP, keeping per capita GDP at its mean value, while in panels (D), (E), and (F) the estimated partial correlations of the female dummy are functions of GGGI and per capita GDP, keeping GGGI at its mean value. We show that left-leaning women are significantly *less* favorable to redistribution at the minimum level of the GGGI (a −0.08 difference with men, when per capita GDP takes on its mean value), while this does not hold for the minimum level of per capita GDP (when GGGI takes on its mean value). However, left-leaning women are not significantly different from men in their preferences at the maximum levels of both GGGI and per capita GDP. On the other hand, centrist and right-wing women are consistently *more* favorable to redistribution than centrist and right-wing men, and those differences are increasing with both GGGI and per capita GDP. More precisely, differences with men are 0.17 and 0.26 respectively at the maximum level of the GGGI when per capita GDP takes on its mean value, while they are 0.16 and 0.25 respectively at the maximum of per capita GDP when GGGI takes on its mean value).

In Table 5, we further investigate gender-specific differences in redistributive preferences by replacing the overall GGGI with its sub-indexes: political empowerment (POLGGGI), economic participation and opportunity (ECOGGGI), educational attainment (EDUGGGI), and health and survival (HSGGGI). While in columns 1–4 we separately include each of the GGGI components, in column 5 we include them jointly. We show that women are more favorable to redistribution vis-à-vis men in environments there are

[26] In Fig. 5 the value referring to left-(right-) wing women is computed as the sum of (i) the coefficient of female dummy + (ii) the coefficient of left-(right-) wing dummy + (iii) the coefficient on the interaction of female dummy with the left- (right-) wing dummy. On the other hand, the value that refers to left-(right-) wing men is simply the coefficient of the left- (right-) wing dummy.

[27] We find similar results when, in column 7, we include the interaction terms between the female dummy and ideology but we interact the female dummy with per capita GDP (see Figure A.2 in the Appendix).

Table 4
Redistributive preferences and the Global Gender Gap Index (GGGI).

Dep.var.: Redistributive pref	(1) b/se	(2) b/se	(3) b/se	(4) b/se	(5) b/se	(6) b/se	(7) b/se	(8) b/se
Female dummy	0.106***	−0.009	0.041**	−0.004	0.100***	−0.031	0.020	−0.025
	(0.014)	(0.023)	(0.016)	(0.023)	(0.015)	(0.026)	(0.017)	(0.026)
GGGI	0.252	−0.134	0.259	−0.008	0.247	−0.177	0.254	−0.008
	(0.824)	(0.843)	(0.822)	(0.830)	(0.824)	(0.848)	(0.823)	(0.839)
Per capita GDP	−0.008***	−0.008***	−0.009***	−0.009***	−0.008***	−0.008***	−0.009***	−0.009***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Left-wing dummy	0.358***	0.358***	0.358***	0.358***	0.415***	0.410***	0.411***	0.409***
	(0.037)	(0.037)	(0.037)	(0.037)	(0.042)	(0.042)	(0.042)	(0.042)
Right-wing dummy	−0.079**	−0.078**	−0.078**	−0.078**	−0.120***	−0.126***	−0.127***	−0.128***
	(0.034)	(0.034)	(0.034)	(0.034)	(0.041)	(0.040)	(0.040)	(0.040)
Household total net income	−0.047***	−0.047***	−0.048***	−0.047***	−0.047***	−0.047***	−0.047***	−0.047***
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
No children at home	−0.029*	−0.030**	−0.030**	−0.030**	−0.029*	−0.029*	−0.029*	−0.029*
	(0.015)	(0.015)	(0.015)	(0.015)	(0.014)	(0.014)	(0.014)	(0.014)
Highest level of education	−0.062***	−0.062***	−0.061***	−0.061***	−0.061***	−0.062***	−0.061***	−0.061***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
Religiosity	0.004	0.004	0.004	0.004	0.004	0.004	0.004	0.004
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
Female d. * GGGI	.	0.749***	.	0.511**	.	0.820***	.	0.501**
	.	(0.165)	.	(0.194)	.	(0.174)	.	(0.202)
Female d. * Per capita GDP	.	.	0.002***	0.001	.	.	0.002***	0.001**
	.	.	(0.000)	(0.001)	.	.	(0.000)	(0.001)
Female d. * Left-wing d.	.	.	.	.	−0.108***	−0.100***	−0.101***	−0.099***
	.	.	.	.	(0.022)	(0.022)	(0.022)	(0.022)
Female d. * Right-wing d.	.	.	.	.	0.073***	0.087***	0.090***	0.091***
	.	.	.	.	(0.020)	(0.020)	(0.020)	(0.020)
Interview-year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Obs.	210 014	210 014	210 014	210 014	210 014	210 014	210 014	210 014
Countries	34	34	34	34	34	34	34	34
R-squared	0.098	0.098	0.098	0.098	0.098	0.099	0.099	0.099

OLS estimates with the inclusion of interview-year fixed effects. For political ideology, the centrist dummy is the omitted reference category. In all regressions, we also control for age, age squared, legal marital status, and the main source of household income. For reasons of space, the estimates of these coefficients and of the constant terms are not reported.
Standard errors in parentheses are clustered at the country level.
*** $p < 0.01$.
** $p < 0.05$.
* $p < 0.1$.

politically, economically, and educationally more equal. We notice that no change in redistributive preferences occurs for men when gender equality goes up according to the economic and political components of GGGI, while women become more pro-redistribution. Interestingly, and perhaps surprisingly, in more gender-equal environments education-wise, men are significantly *less* favorable to redistribution, while the opposite is true for women. Finally, we do not find any significant correlation between the health-related component of GGGI and redistributive preferences, both unconditionally and conditionally to the respondent's gender. To note, when we include all four GGGI sub-indexes in the same specification (column 5), the only interaction term with the female dummy that remains statistically significant is the one with the political component.

Fig. 7 shows the estimated partial correlations for the female dummy for the different components of GGGI, computed from column 5 of Table 5, i.e., when including all GGGI components within the same specification.

4.1. An instrumental variable approach

The GGGI –as a measure of gender equality in country c in year t– might be significantly correlated with other unobserved factors at the country-year level that in turn are correlated with redistributive preferences, thus biasing our estimates. To deal with this potential instance of omitted variable bias, we use our country-specific historical variables, i.e., the year when women's suffrage was first introduced and the historical role of women in the family, as instruments within an Instrumental Variable (IV) specification. This empirical approach is valid to the extent that those historical variables (i) are significantly correlated with the GGGI and (ii) they have an effect on our dependent variable (redistributive preferences at the individual level) only through this channel of correlation, i.e., the current level of gender equality, as measured by the GGGI. Starting from the specification shown in column 2 of Table 4, the second stage of the IV specification is the following:

$$y_{it} = \alpha_{it} + \beta_0 \text{Female}_{it} + \beta_1 \text{GGGI}_{ct} + \beta_2 \text{per capita GDP}_{ct} + \beta_3 \text{Ideology}_{it} + \beta_4 X_{it} + \beta_5 (\text{Female}_{it} * \text{GGGI}_{ct}) + \eta_t + \varepsilon_{ict} \quad (4)$$

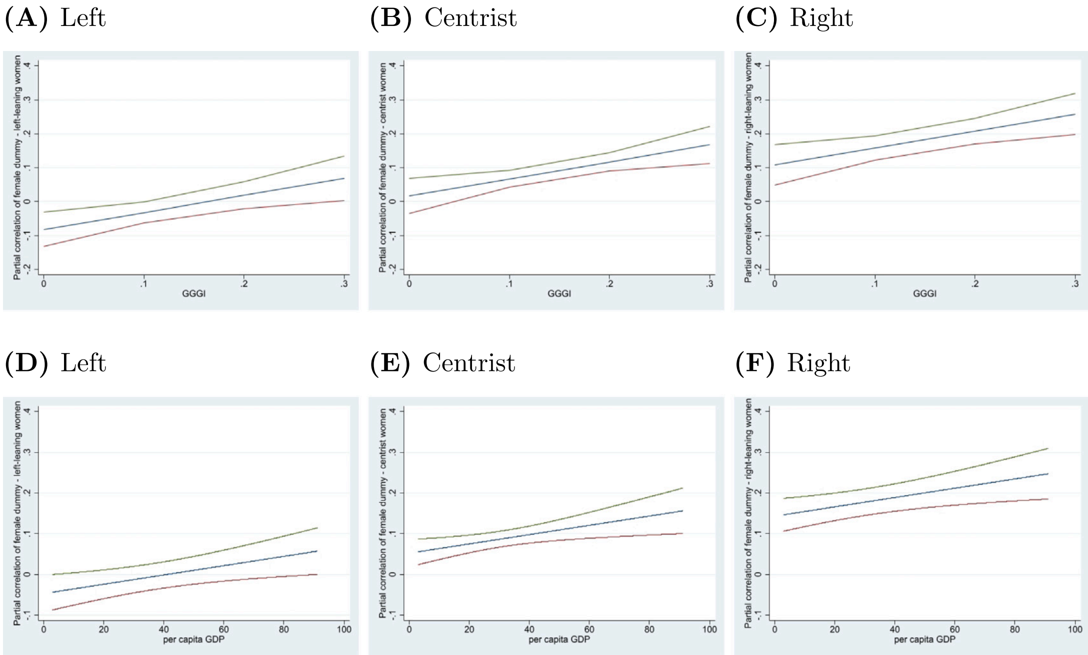

Fig. 6. Estimated partial correlations of the female dummy with redistribution preferences, as a function of political ideology, GGGI, and per capita GDP.
In panels (A), (B), (C), (D), (E), and (F) the blue line represents the estimated partial correlation of female woman dummy in the regression for redistributive preferences –as computed from Table 4, column (8)–, while the red and green lines identify the upper and the lower bounds of the 95% confidence intervals. In panels (A), (B), and (C) the estimated partial correlations of the female dummy are functions of GGGI and per capita GDP, keeping per capita GDP at its mean value, while in panels (D), (E), and (F) the estimated partial correlations of the female dummy are functions of GGGI and per capita GDP, keeping GGGI at its mean value. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

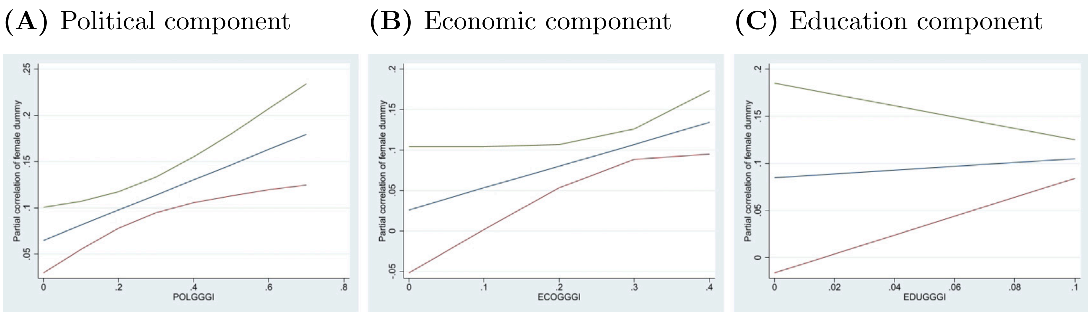

Fig. 7. Estimated partial correlations of the female dummy with redistributive preferences, as a function of specific GGGI components.
In panels (A), (B), and (C) the blue line corresponds to the estimated partial correlation of the female dummy in the regression for redistributive preferences –as computed from Table 5, column (5)–, while the red and green lines identify the upper and the lower bounds of 95% confidence intervals. The estimated partial correlation of the female dummy that are shown in panels (A), (B), and (C) are functions of the political, economic, and educational components of GGGI, each time keeping the other components of GGGI at their mean values, respectively. We do not show the figure referred to the health component of GGGI (HSGGGI) because it only takes on values between 0 and 0.05. However, the estimated partial correlation of the female dummy is 0.044 (not statistically significant) when HSGGGI is equal to zero, while it is 0.106 and 0.114 (both statistically significant) when HSGGGI is equal to 0.044 (its mean value) and to 0.05, respectively. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

In Eq. (4), we include all variables that appear in Eq. (2) above, plus the interaction between the GGGI and the female dummy. Then, we instrument (i) the GGGI and (ii) the interaction between the GGGI and the female dummy with women's suffrage year and the historical role of women in the family, by themselves and interacted with the female dummy, thus obtaining four instruments.[28]

We obtain the following first stages:

[28] For the rationale of this specification see Wooldridge (2010), Chapter 9.

Table 5
Redistributive preferences and the components of GGGI.

Dep.var.: Redistributive pref	(1) b/se	(2) b/se	(3) b/se	(4) b/se	(5) b/se
Female dummy	0.050***	−0.043	−0.038	0.053	−0.100
	(0.016)	(0.033)	(0.037)	(0.100)	(0.061)
POLGGGI	−0.004	.	.	.	−0.011
	(0.287)	.	.	.	(0.259)
ECOGGGI	.	−0.116	.	.	0.395
	.	(0.460)	.	.	(0.613)
EDUGGGI	.	.	−3.084**	.	−3.903*
	.	.	(1.272)	.	(1.952)
HSGGGI	.	.	.	3.904	4.731
	.	.	.	(5.057)	(4.932)
Per capita GDP	−0.009***	−0.008***	−0.008***	−0.008***	−0.009***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Left-wing dummy	0.358***	0.358***	0.354***	0.360***	0.360***
	(0.037)	(0.037)	(0.036)	(0.038)	(0.037)
Right-wing dummy	−0.078**	−0.080**	−0.080**	−0.078**	−0.080**
	(0.033)	(0.034)	(0.034)	(0.033)	(0.033)
Household total net income	−0.047***	−0.047***	−0.046***	−0.048***	−0.047***
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
No children at home	−0.030*	−0.030*	−0.030*	−0.032**	−0.034**
	(0.015)	(0.015)	(0.015)	(0.014)	(0.014)
Highest level of education	−0.061***	−0.063***	−0.058***	−0.062***	−0.062***
	(0.008)	(0.009)	(0.009)	(0.008)	(0.009)
Religiosity	0.004	0.004	0.003	0.005	0.004
	(0.006)	(0.007)	(0.006)	(0.006)	(0.006)
Female d. * POLGGGI	0.219***	.	.	.	0.164**
	(0.062)	.	.	.	(0.071)
Female d. * ECOGGGI	.	0.501***	.	.	0.270
	.	(0.123)	.	.	(0.169)
Female d. * EDUGGGI	.	.	1.336***	.	0.200
	.	.	(0.380)	.	(0.551)
Female d. * HSGGGI	.	.	.	1.195	1.417
	.	.	.	(2.113)	(1.499)
Interview-year FE	Yes	Yes	Yes	Yes	Yes
Obs.	210 014	210 014	210 014	210 014	210 014
Countries	34	34	34	34	34
R-squared	0.098	0.098	0.099	0.098	0.100

OLS estimates with the inclusion of interview-year fixed effects. In all regressions we also control for age, age squared, legal marital status, and the main source of household income. For reasons of space, the estimates of these coefficients and of the constant terms are not reported.
Standard errors in parentheses are clustered at the country level.
*** $p < 0.01$.
** $p < 0.05$.
* $p < 0.1$.

$$GGGI_{ct} = \alpha_{it} + \beta_0 \text{Female}_{it} + \beta_1 \text{Women's suffrage}_c + \beta_2 \text{Degree of Feminism}_c + \beta_3(\text{Women's suffrage}_c * \text{Female}_{it}) + \beta_4(\text{Degree of Feminism}_c * \text{Female}_{it}) + \beta_5 X_{it} + \eta_t + \varepsilon_{ict} \quad (5)$$

$$(GGGI_{ct} * Female_{it}) = \alpha_{it} + \beta_0 \text{Female}_{it} + \beta_1 \text{Women's suffrage}_c + \beta_2 \text{Degree of Feminism}_c + \beta_3(\text{Women's suffrage}_c * \text{Female}_{it}) + \beta_4(\text{Degree of Feminism}_c * \text{Female}_{it}) + \beta_5 X_{it} + \eta_t + \varepsilon_{ict} \quad (6)$$

The outcome of this exercise is shown in columns 1 and 2 of Table 6, with the output of the first stage regressions in the bottom panel. Women's suffrage year is negatively and significantly correlated with the GGGI. On the other hand, the interaction terms between the female dummy and the women's suffrage variable and the degree of feminism, respectively, are both statistically significant when the dependent variable is the interaction of the female dummy and the GGGI, and with the expected sign.[29]

The second stage of the IV specification is displayed in column 1 of the top part of Table 6, and is consistent with our core findings: the differences in redistributive preferences between women and men are positively and significantly associated with gender

[29] Since the dependent variable is the interaction between the GGGI and the female dummy, all its relevant variation happens in the case of women, because the dependent variable is –by definition– zero in the case of men. We check the *overall* estimated correlation of the degree of feminism, which is made of the coefficient on the degree of feminism by itself and the coefficient on its interaction with the female dummy. The former coefficient is negative and significant, while the latter is positive and significant, but larger in size. One can reject at ordinary confidence level the null hypothesis that the sum of those two effects is zero.

Table 6
Redistributive preference, GGGI, and POLGGGI: IV regression.

	(1) b/se	(2) b/se	(3) b/se	(4) b/se
Panel A: Two-Stage Least Squares				
Dep. var.	Redistributive pref.		Redistributive pref.	
Female dummy	−0.005	–	0.038	–
	(0.051)		(0.029)	
GGGI	−2.138	–	–	–
	(1.940)			
POLGGGI	–	–	−0.562	–
			(0.707)	
Female d. * GGGI	0.727**	–	–	–
	(0.354)			
Female d. * POLGGGI	–	–	0.267**	–
			(0.129)	
PC Female dummy for GGGI min	−.005	–	–	–
p-value	0.926			
PC Female dummy for GGGI average	.106	–	–	–
p-value	0.000			
PC Female dummy for GGGI max	.190	–	–	–
p-value	0.000			
PC Female dummy for POLGGGI min	–	–	.038	–
p-value			0.193	
PC Female dummy for POLGGGI average	–	–	.105	–
p-value			0.000	
PC Female dummy for POLGGGI max	–	–	.196	–
p-value			0.000	
Obs.	208 624	–	208 624	–
Countries	33	–	33	–
R-squared	0.094	–	0.096	–
Panel B: First Stage				
Dep. var.	GGGI	Female d. * GGGI	POLGGGI	Female d. *POLGGGI
Female dummy	2.450**	2.450**	−0.169	6.239*
	(0.048)	(1.162)	(0.159)	(3.452)
Women's suffrage year	−0.001***	0.000	−0.003***	0.000
	(0.000)	(0.000)	(0.001)	(0.001)
Degree of Feminism	−0.027	−0.057***	−0.106	−0.184***
	(0.034)	(0.008)	(0.090)	(0.027)
Female d. * Women's suffrage year	0.000	−0.001**	0.000	−0.003*
	(0.000)	(0.001)	(0.000)	(0.002)
Female d. * Degree of Feminism	−0.001	0.085**	0.010	0.263***
	(0.002)	(0.032)	(0.007)	(0.093)
Obs.	208 624	208 624	208 624	208 624
Countries	33	33	33	33
R-squared	0.717	0.908	0.731	0.782

IV estimates with the inclusion of year-of-interview fixed effects. In all the regressions, we also control for per capita GDP, political ideology (left-wing and right-wing dummies), household total net income, no children at home, highest level of education, religiosity, age, age squared, legal marital status, and the main source of household income. For space reasons, the estimates of these coefficients as well as of the constant terms are not reported. Standard errors in parentheses are clustered at the country level.
*** $p < 0.01$.
** $p < 0.05$.
* $p < 0.1$.

equality at the country level, and in turn this increased difference is driven by women being more favorable to redistribution in more gender-equal environments. The partial correlation between being a woman and redistributive preferences is indistinguishable from zero at the minimum level of gender equality, while it is significantly different from zero and positive at the average level of gender equality and even larger at their maximum level. The size of those estimated effects are almost identical to the ones derived from the OLS analysis (Table 3, column 2).[30]

[30] To check the strength and relevance of our instruments, we perform three different statistical tests to the outcome of the first stage regressions. Since we are in the presence of heteroscedasticity and clustering, we look at the Sanderson–Windmeijer (SW) multivariate F test of excluded instruments (Sanderson and Windmeijer, 2016). Following the standard rule of thumb, the F-statistics of excluded instruments should be greater than 10. In our case, we are able to reject the joint null hypothesis of weak instruments separately for each equation (GGGI, SW F = 40.15; interaction between GGGI and female dummy, SW F = 10.73, for both p-value < 0.0000). In both cases, our instruments are satisfactory also according to the Stock–Yogo weak ID F test critical values for the case of 2

To go even further with our analysis, we run our IV specification on each component of GGGI. Our results confirm that women are more favorable to redistribution vis-à-vis men in politically more equal environments (see columns 3 and 4 of Table 6). We also find a similar effect, but only mildly significant in more economically equal countries, while the health and education related components of gender equality are no longer statistically significant by themselves as well as when interacted with the female dummy.[31] Looking at the first stages, we notice that, as in the specification with the GGGI, the interaction terms between the female dummy and the women's suffrage variable and the degree of feminism, respectively, are statistically significant instruments when the dependent variable is the interaction of the female dummy and POLGGGI, with the expected sign. When we look at the interactions of the female dummy with ECOGGGI or EDUGGGI, only the interaction terms between the women's suffrage variable and the female dummy is significant, always with the expected sign. Finally, with HSGGGI no effect emerges, confirming our previous observations.

5. Concluding remarks

We show that historical roots of gender equality, as measured by the timing of women's enfranchisement and the historical role of women in the family, are significantly associated with the redistributive features of tax systems across countries: in countries with earlier women's enfranchisement and/or where women historically played a more relevant role within the family, the share of direct taxes (over indirect taxes and over total tax revenue) is higher than in countries with a later enfranchisement and with historical gender inequality within the family.

At the individual level we find that in more gender-equal environments gender differences in redistributive preferences are significantly larger, where the "action" comes from women being systematically more favorable to redistribution, and with no significant changes for men. Notably, the larger gender-based difference in redistributive preferences in more gender-equal environments is driven by the *political* component of gender equality. Thus, one could argue that gender equality in the political sphere is associated with women being more interested in income redistribution through government intervention.

We need additional analyses to identify the specific mechanisms through which stronger redistributive preferences for women in more gender-equal environments might influence economic policies. For example, it could be the case that on the demand side stronger redistributive preferences of women make redistribution more salient, since both politicians and media outlets might find it optimal to care about women as swing voters (Dixit and Londregan, 1996) and/or marginal consumers of media products (Hamilton, 2004). On the supply side, it would be interesting to investigate whether and how the presence and relevance of female politicians are conducive to more redistributive policy choices, particularly so in more gender-equal environments.

To the extent that income redistribution is normatively considered as a fundamental task for the government intervention in a market economy, our findings emphasize the opportunity to promote gender equality overall *and* in the political domain, in order to more effectively pursue this redistributive goal. From a predictive point of view, it would be interesting to explore whether in the long run the stronger pro-redistribution "voice" of women in more gender-equal environments would translate into faster reduction of gender differences in income and wealth (Goldin, 2006).

CRediT authorship contribution statement

Monica Bozzano: Conceptualization, Data curation, Formal analysis, Investigation, Methodology, Supervision, Validation, Visualization, Writing – original draft, Writing – review & editing. **Paola Profeta:** Conceptualization, Resources, Supervision, Validation, Writing – review & editing. **Riccardo Puglisi:** Conceptualization, Formal analysis, Investigation, Methodology, Supervision, Validation, Visualization, Writing – original draft, Writing – review & editing. **Simona Scabrosetti:** Conceptualization, Data curation, Formal analysis, Investigation, Methodology, Supervision, Validation, Visualization, Writing – original draft, Writing – review & editing.

Declaration of competing interest

The authors Monica Bozzano, Paola Profeta, Riccardo Puglisi, and Simona Scabrosetti declare none.

Data availability

Data will be made available on request.

Appendix A. Supplementary data

Supplementary material related to this article can be found online at https://doi.org/10.1016/j.ejpoleco.2023.102497.

instrumented endogenous variables and 4 excluded instruments (Stock and Yogo, 2005). Looking at the Kleibergen–Paap Wald rk F statistic, where the null hypothesis is that instruments are jointly weakly identified, we obtain analogous results (10.72). Finally, according to the SW Chi-sq tests, we also significantly reject the null of under-identification (respectively, SW Chi-sq = 124.25 p-value < 0.0000).

[31] Results for ECOGGGI, EDUGGGI, and HSGGGI are available upon request.

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