The Global Gender Gap in Labour Income

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Abstract

This paper introduces a new measure of economic gender inequality (EGI) based on the ratio of women's share of national labour income to men's. This measure captures both the principles of equal pay for equal work and non-discrimination. Importantly, it can be calculated from existing data and is comparable across countries and time. We show that EGI has only been improving slowly and that due to population growth aggregate EGI has been increasing and is equivalent to one billion working-age women working for nothing by 2017. Moreover, this gap is expected to continue to increase in coming decades. A decomposition shows that this is largely driven by the gender gap in labour market participation. Instrumental variable estimates suggest that while increases in income and democratisation reduce EGI, living standards would have to increase substantially for the most unequal societies, such as Tunisia or Iran, to approach those of the countries that are currently most equal such as Latvia or Norway.

Keywords: Economic Gender Inequality, Global Distribution of Income, Modernization Hypothesis

JEL Codes: J16, J71, D33, O15

This paper takes a modern view of what economic gender equality means. We argue that in a society with genuine economic gender equality there would be no difference in the total wages of men and women. Such equality of outcomes would reflect that men and women participated equally in the labour market, worked in a similar set of roles, occupations, and industries, and received equal pay for equal work.

We are thus claiming that when summing across men and women all national differences in total pay reflect inequality. That is, there is no justifiable difference in pay due to men or women preferring particular roles, having different preferences — on average — for risk or competition, or due to child-rearing and other forms of care giving.

This is a more expansive, and more demanding, measure of economic gender inequality (EGI) than the Gender Pay Gap (GPG). In support of our benchmark we note that the list of tasks that could *only* be done by men or women has shortened rapidly in recent years. Previous prohibitions on women in the armed (special) forces, certain sports, as astronauts, or firefighters have been overturned. While the list of roles, typically involving care, considered unsuited to men has likewise shortened.

Our benchmark incarnates our argument that higher-pay in male-dominated occupations and industries, and indeed the concentration of men and women in particular occupations and industries are not justifiable and constitute EGI. That is, as discussed in more detail in appendix A, in a gender equal society differences in preferences or other attributes will not lead to systematic differences in wages.

Economic Gender Inequality (EGI) has two components; discrimination between similarly qualified men and women, and differences in access to education, training, or particular sectors of the economy. Absent a few exceptions, there is a uniform commitment to eliminating both of these. Almost every country is a signatory to the Equal Remuneration Convention (1951), committing them to the *"principle of equal remuneration for men and women workers for work of equal value"*.[1] Similarly, almost all are signatories to The Convention on the Elimination of All Forms of Discrimination against Women (1979). The forms of discrimination addressed by this

[1]The USA is a prominent exception but has had a similar commitment since the 1963 Equal Pay Act.

convention lead to important aspects of EGI beyond equal pay for equal work, including differences in labour market participation, unemployment, and home production. The last century saw enormous progress, termed 'the grand gender convergence' by Goldin (2014), particularly in high-income countries. Nevertheless, the evidence suggests that despite such laws, a substantial pay gap remains. In the OECD, where we expect it to be smallest, the gender pay gap (GPG) (the difference between female and male median wages, divided by male median wages) remains over 15%, and it is as large as 37% in South Korea.[2] Outside the OECD, inequality is often even higher.

The measure of EGI this paper introduces captures both departures from equal pay for equal work and limits to women's labour market opportunities due to discrimination, which we term the labour share ratio. This is, the labour share of income of women — the compensation of female workers as a share of their value added, divided by the labour share of men. The idea is simple: one implication of 'equal pay for work of equal value' is that the ratio of compensation to value added should be the same for men and women. Our argument is that imperfect competition in labour and product markets mean that workers of both genders must bargain over their share of output. The extent to which male workers receive more, *ceteris paribus*, reflects differences in the relative bargaining strength of men and women.

Similarly, the elimination of discrimination against women implies equal access to education and training and no limitation in terms of occupation, sector, or rank. It thus also implies the elimination of most, if not all, differences in total value added (per hour) and the differences in labour market participation among men and women. Value added cannot normally be disaggregated by gender, but we do not need to to calculate our measure, all we need is the assumption that any systematic deviation from equal-value added per hour reflects a deviation from equality of opportunity. This assumption is empirically supported by the convergence documented by Goldin (2014) as well as the progressive elimination of explicit and implicit prohibitions on women working in roles previously restricted to men on the basis of presumed capability such as firefighters or front-line soldiers. Our measure therefore captures both equal pay for equal work and equality of opportunity.

[2]See, https://www.oecd.org/gender/data/genderwagegap.htm.

This approach has three key advantages. First, by focusing on the share of the value added we are able to abstract from cross-country variation in the determinants of value added that normally make meaningful cross-country and intertemporal comparison difficult. Importantly, for example, our approach does not require us to focus on full-time equivalent employees only, and allows for differences in self-employment and informality across place and over time. Second, this also makes aggregation meaningful and we are able to present estimates for total global EGI. Finally, our approach relies on well-understood data: the data that make up GDP statistics. Using these extant data we construct a panel for the period 1991-2018 covering over 90 countries accounting for (in 2018) nearly 4.5 billion working-age people.

Using these new data this paper studies how EGI varies across countries, its changes over time, and its composition. It also studies the evolution of aggregate global pay inequality. We find that whilst EGI has been slowly shrinking in most countries, that the relatively high birthrate in more unequal countries means that aggregate inequality has increased and will continue to increase until around 2050. The results imply that aggregate EGI in 2017 was equivalent to over one billion working-age women working for no compensation whatsoever. We then use the broad coverage provided by our new measure to analyse the causal impact of economic development and political emancipation on EGI. We find that while development reduces inequality, it does so only slowly. Our results suggest that Democracy also reduces inequality but only above a certain income level.

This paper contributes to three related literatures. Its first contribution is to the important literature on EGI, particularly research such as Blau and Kahn (1992, 2003) Olivetti and Petrongolo (2008) and Mulligan and Rubinstein (2008) that compares the GPG across countries or studies its evolution over time. The labour share ratio has the important quality that it is directly comparable across time and place, and as it captures both the principle of 'equal pay for equal work' and discrimination it implicitly controls for differences in the form of EGI across time and place. Moreover, our results quantify the relative importance of differences in pay and differences in labour force participation. Among the roughly 1.65 billion women currently working, 0.45 billion are unpaid equivalent due to low hourly pay and/or lower working hours compared with their men counterpart. The number of unpaid equivalents would have been 0.8 billion if labour force participation were the same among women as in men.

By introducing comparable new data on EGI, this paper's second contribution is to introduce gender to the prominent literature that studies trends in income inequality, such as Piketty and Saez (2003) and Atkinson et al. (2011). The broad coverage means that it also contributes to the related literature measuring the global income distribution, particularly the work of Jones (1997), Milanovic (2002) and, Sala-i Martin (2006) as well as the more recent work of Milanovic (2015). In both cases it contributes to these literatures by conducting similar analyses, but for EGI.[3] In doing so it is related to the surveys of Olivetti and Petrongolo (2016) who focus on the industrialized countries and Klasen (2020) who provides a global overview. Our results show that by ignoring gender differences in earnings the level of trends in actual inequality are understated. By documenting the considerable levels of EGI still routine in most of the world, our findings also provide a valuable counterpoint to the "Grand Gender Convergence" identified by Goldin (2014).

Our approach necessitates constructing the first gender-specific labour shares. Thus, this paper also contributes to the prominent literature that studies inequality through the lens of declining labour shares particularly the cross-country analyses of Karabarbounis and Neiman (2014) as well as the recent US focused papers by Acemoglu and Restrepo (2018), Autor et al. (2020).

The third contribution of this paper is to the growing literature on whether EGI is a symptom or a cause of underdevelopment. A case for the former is made by Fernández (2014) who shows theoretically and empirically that improvements in women's rights can be seen as an endogenous response to economic development. In a similar spirit, there is an increasing body of microeconometric evidence that documents how local changes in female political empowerment leads to changes in policy, such as Duflo (2004) or Bhalotra and Clots-Figueras (2014).[4] Given our focus

[3]Also related is Dorius and Firebaugh (2010) who study trends in aggregate gender inequality for a range of measures of literacy, life expectancy, and political representation.

[4]See Banerjee and Duflo (2005) for a review of the literature on gender inequality and development.

on EGI it is also related to work, notably Acemoglu et al. (2015), presenting evidence that a transition to democracy leads to a decrease in overall income inequality. A similar view, that (gender) inequality is a symptom of underdevelopment, is known in political sociology as the 'Modernization Hypothesis' (see, Inglehart and Baker, 2000), which states that rising living standards and political emancipation cause changes in values leading, *inter alia*, to improvements in gender equality. Taken together these different literatures suggest that EGI may be best understood as a symptom of underdevelopment rather than a separate pathology. This view is consistent with the Cubas (2016) who argues that women's labour force-participation increases with improvements in infrastructure and access to household appliances.

On the other hand, others such as Doepke et al. (2012) and Doepke and Tertilt (2019) have argued that gender inequality impedes economic development as it reduces investment in the human capital of children.[5] We provide the first causal analysis of the effects of improvements in income and political rights on EGI at the aggregate level. Our instrumental variables estimates suggest that an increase in per capita incomes from the bottom of the World Bank's lower-middle income category ($1,046) to the bottom its rich income category ($12,696) would be sufficient to reduce EGI by two-thirds – from a level similar to that of as in Tunisia, Iran, or Turkey, to one close to that of Norway, Sweden, or Latvia. We find that democratization reduces EGI, but only countries with per capita annual incomes of at least $10,000. Thus, while a substantial literature shows women's political empowerment leads to changes in policymaking in low income countries, such as Duflo (2004), we find no causal evidence that this is true for EGI at the national level. We also obtain results on the roles of other determinants, such as globalization as studied by (see, Oostendorp, 2009a) and Weichselbaumer and Winter-Ebmer (2007).

This paper proceeds as follows. The next section discusses how calculate EGI and the data we use. Section 3 describes patterns of gender equality around the world and provides estimates of the aggregate global gender gap. Section 4 provide estimates of the causal effects of prosperity and democratization on EGI. Section 5

[5] A strand of the literature advances the stronger alternative hypothesis that gender inequality may be a path, via a more competitive manufacturing sector, to growth. (see, Seguino, 2000, Schober and Winter-Ebmer, 2011, Seguino, 2011).

briefly concludes.

2 Measuring Gender Pay Inequality

Computing our measure of EGI, the labour share ratio, requires first computing the separate labour shares of men and women, λ_M, λ_F and then their ratio. In this section we first discuss the issues in computing labour shares and how we address them, their ratio, and the data we use to do so.

2.1 Measuring the labour share of income

The most straightforward measure of the labour share is simply the share of employees compensation in total value-added:

$$\lambda^U_{c,t,s} = \frac{\text{Compensation of Employees}_{c,t,s}}{\text{Total Value Added}_{c,t,s}} = \frac{\overline{W}^E_{c,t,s} N^E_{c,t,s}}{Y_{c,t,s}}, \tag{1}$$

where $\overline{W}^E_{c,t,s} = \overline{w}^E_{c,t,s}\overline{H}^E_{c,t,s}$ are average annual earnings per employee in country c in year t in sector s, and $N^E_{c,t,s}$ is the number of employees in that sector, country, and year. $\overline{w}^E_{c,t,s}$ are similarly average hourly earnings in that sector, country, and year. Normally we prefer to work with broad sectors such that $s \in (total\ economy, agriculture, industry, services, ...)$. But, we also work more disaggregated data for each of the 14 sub sectors. Using eq. (1) we can readily calculate the labour share of employees for the economy as a whole and each (sub-)sector of total economy.

However, as highlighted by Krueger (1999) and Gollin (2002), while conceptually simple measuring the compensation of employees is more complicated in practice. The so-called naive, or unadjusted, labour share captured by eq. (1) ignores mixed-income (income recorded in national accounts as due to multiple factors of production). Thus, it does not allow, *inter alia*, for the fact that the incomes of the self-employed will partially reflect returns to the capital employed and partly to the labour supplied. This will mean that the labour share is under-estimated.

This will be problematic in our context if the self-employed are disproportionately male or female. Gollin (2002) proposes three alternatives to eq. (1) designed to better capture self-employment income which we refer to as λ^{G1}, λ^{G2}, and λ^{G3} respectively. The first, as may be seen in eq. (2), attributes to the Compensation of Employees all mixed income, thus over-estimating the labour share by including returns properly attributable to land, capital, etc. The second, defined in eq. (3), assumes that the share of wages in mixed income is the same as in the wider economy, thus assuming that capital per worker etc., is the same. The third, given in eq. (4) assigns to self-employed workers the same average wage as employed workers. It thus assumes that wages are the same in self-employment and employment and thus that any additional returns are attributable to other factors of production.

$$\lambda^{G1}_{c,t,s} = \frac{\overline{W}^{E}_{c,t,s} N^{E}_{c,t,s} + M_{c,t,s}}{Y_{c,t,s}} \tag{2}$$

$$\lambda^{G2}_{c,t,s} = \frac{\overline{W}^{E}_{c,t,s} N^{E}_{c,t,s}}{Y_{c,t,s} - M_{c,t,s}} \tag{3}$$

$$\lambda^{G3}_{c,t,s} = \frac{\overline{W}^{E}_{c,t,s} N^{E}_{c,t,s} + \overline{W}^{E}_{c,t,s} N^{SE}_{c,t,s}}{Y_{c,t,s}} \tag{4}$$

Where, $N^{SE}_{c,t,s}$ is the number of self-employment in country c in year t in sector s and $M_{c,t,s}$ represents mixed income. Note that the last equation assumes that employees and the self-employed work an equal number of hours on average.

Self-Employment in LDCs

However, the nature of self-employment varies considerably between countries, particularly between high- and low-income countries. While self-employment in high-income countries is often a choice, in lower-income countries it often reflects the necessity of work in subsistence agriculture (Günther and Launov, 2012). This may mean that in lower-income countries the wages of the self-employed are lower than those of employed workers. An appealing solution is to use additional information such as sector-specific wages to make more refined adjustments. However, the necessary data is often available for only a subset of countries and for a limited period.[6] In an effort to produce reliable estimates while maximizing sample size Feenstra et al. (2015) employ different measures of the labour share depending on development status and data availability. For many low and middle-income countries this amounts to computing the labour share using λ^{G2} but proxying mixed-income with total value-added in the agricultural sector.

As van Treeck (2020) notes, however, this approach will overstate the labour share to the extent that other factors such as land and capital contribute to agricultural output. As well as double-counting agricultural employment (since it will be included in both as employment and again in total value-added in agriculture.). Using social-accounting matrices to inform and validate an alternative adjustment she proposes proxying self-employment in low- and middle-income countries and assuming the (average) wage of the self-employed is half that of the average employee in low-income countries and equal to the average wage in middle- and high-income countries.

$$\lambda^{VT}_{c,t,s} = \begin{cases} \lambda^{\dagger} = \frac{\overline{W}^{E}_{c,t,s} N^{E}_{c,t,s} + k\upsilon_{c,t,s} \overline{W}^{E}_{c,t,s} N_{c,t,A}}{Y_{c,t,s}} & \text{if } \lambda^{\dagger} \in c \notin \text{ high income group} \\ \lambda^{\dagger} = \frac{\overline{W}^{E}_{c,t,s} N^{E}_{c,t,s} + k\overline{W}^{E}_{c,t,s} N^{SE}_{c,t,s}}{Y_{c,t,s}} & \text{if } \lambda^{\dagger} \in c \in \text{ high income group} \end{cases} \tag{5}$$

Where, $N_{c,t,A}$ is the number of people in country c at time t working in the agricultural sector A, $\upsilon_{c,t,s}$ represents the employment share of sector s in the country total employment with $\upsilon = 1$ for the total economy and k captures the proportion of average wages assumed to be earned by self-employed workers. Following van Treeck (2020) we set $k = 1/2$ for low-income countries and $k = 1$ elsewhere. Again following van Treeck (2020) we treat λ^{G1} and λ^{G2} as lower and upper bounds on λ^{VT} and thus if $\lambda^{VT} \notin \left[\lambda^{G1}, \lambda^{G2}\right]$ then we adjust k such that $\lambda^{VT} \in \left[\lambda^{G1}, \lambda^{G2}\right]$.

[6] An alternative approach is to focus on the corporate (Karabarbounis and Neiman, 2013) or manufacturing sectors (Rodrik, 1999). But, this approach is better suited to high-income countries since these sectors often account for only a small fraction of total employment in low- and middle-income countries.

Hours Worked

One additional possibility is that average hours worked may differ between employment and self-employment. To the extent that hours worked are lower (higher) in self-employment than employment $\lambda^{\dagger}$ will over-estimate (under-estimate) the labour share. This is of particular concern here as a relatively subtle, but quantitatively important aspect of economic gender inequality is that in all countries women do more domestic work such as care or housework which limits the hours they can work in employment or self-employment (Hook, 2010, Ferrant et al., 2014).

To address this we modify $\lambda^{\dagger}$ to incorporate hours worked as follows:

$$\lambda^{\ddagger}_{c,t,s} = \frac{\overline{W}^{E}_{c,t,s} N^{E}_{c,t,s} + k \overline{w}^{E}_{c,t,s} \overline{H}^{SE}_{c,t,s} N^{SE}_{c,t,s}}{Y_{c,t,s}} \tag{6}$$

Where, $\overline{H}^{SE}_{c,t,s}$ are the average annual hours of self-employed. Following a similar logic to Feenstra et al. (2015) and van Treeck (2020) in the case that $\overline{H}^{SE}_{c,t,s}$ or $N^{SE}_{c,t,s}$ is not observed, as is the case for some lower income countries, then we proxy for it using total employment and average annual hours in agriculture sector $\overline{H}_{c,t,A} N_{c,t,A}$. As for λ^{VT} we adjust k, if necessary, such that $\lambda^{\ddagger} \in \left[\lambda^{G1}, \lambda^{G2}\right]$.

2.2 Measuring the labour Share Ratio

We will present results using both the *unadjusted* or naive labour shares $\lambda^{U}(F)$, $\lambda^{U}(M)$ and their ratio ρ^{U} as well as the *adjusted* labour shares $\lambda^{\ddagger}(F)$, $\lambda^{\ddagger}(M)$ and their ratio $\rho^{\ddagger}$.

To compute ρ requires calculating the labour share separately for men and women as follows:

$$\lambda^{U}_{c,t,s}(F) = \frac{\overline{W}^{E}_{c,t,s}(F) N^{E}_{c,t,s}(F)}{Y_{c,t,s}}, \lambda^{U}_{c,t,s}(M) = \frac{\overline{W}^{E}_{c,t,s}(M) N^{E}_{c,t,s}(M)}{Y_{c,t,s}}. \tag{7}$$

Which implies:

$$\rho^{U}_{c,t,s} = \frac{\lambda^{U}_{c,t,s}(F)}{\lambda^{U}_{c,t,s}(M)} = \frac{\overline{W}^{E}_{c,t,s}(F) N^{E}_{c,t,s}(F)}{\overline{W}^{E}_{c,t,s}(M) N^{E}_{c,t,s}(M)}. \tag{8}$$

Calculating $\lambda_{c,t,s}(M)$ and $\lambda_{c,t,s}(F)$ requires data on both value-added and compensation by gender. However, as eq. (8) makes clear, the value-added terms cancel by assumption, and so we need only gender-specific data on the compensation of workers.[7]

We can compute $\rho^{\ddagger}$ similarly:

$$\rho^{\ddagger} = \frac{\lambda^{\ddagger}(F)}{\lambda^{\ddagger}(M)} = \frac{\overline{W}^{E}_{c,t,s}(F)N^{E}_{c,t,s}(F) + k\overline{w}^{E}_{c,t,s}(F)\overline{H}^{SE}_{c,t,s}(F)N^{SE}_{c,t,s}(F)}{\overline{W}^{E}_{c,t,s}(M)N^{E}_{c,t,s}(M) + k\overline{w}^{E}_{c,t,s}(M)\overline{H}^{SE}_{c,t,s}(M)N^{SE}_{c,t,s}(M)}. \tag{9}$$

To isolate the role of hourly earning ($\overline{w}^{E}_{c,t,s}$) and employed hours ($\overline{H}_{c,t,s}N_{c,t,s}$) in $\rho^{\ddagger}$ we also calculate *employment share ratio* as below:

$$\varphi^{\ddagger} = \frac{\overline{H}^{E}_{c,t,s}(F)N^{E}_{c,t,s}(F) + \overline{H}^{SE}_{c,t,s}(F)N^{SE}_{c,t,s}(F)}{\overline{H}^{E}_{c,t,s}(M)N^{E}_{c,t,s}(M) + \overline{H}^{SE}_{c,t,s}(M)N^{SE}_{c,t,s}(M)}. \tag{10}$$

2.3 Data Used in the Construction of EGI

Our approach relies on well-understood data: the data that make up GDP statistics. Whilst, these data have been criticised, particularly for Sub-Saharan Africa (see, Jerven, 2013), they are compiled according to a well-defined standard designed to ensure comparability across countries and years.[8] This is a considerable advantage compared to the meta-analysis approach taken by Oostendorp (2009a). Perhaps most importantly, the ratios obtained by calculating (B.2) using (8) and (9) are dimensionless and thus do not suffer from an index-number problem.

The dataset is a country level panel taken from three major sources: International Labour Organization (ILO), United Nations System of National Accounts (SNA) and The World Bank World Development Indicators (WDI). To reproduce and extend the Gollin (2002) labour share calculations value added data on agriculture, industry, services, manufacturing and total economy's value added is taken from WDI. Annual value added of sub-sectors such as mining, education, construction, electricity, transport, storage and communications, wholesale and retail trade, etc. are taken from SNA.

To calculate the gender disaggregated labour shares at economy and sub-sectors levels we combine SNA and WDI data with ILO data on annual employment, mean

[7]Note, as computing λ^{G1} and λ^{G2} involves Mixed Income we cannot calculate them separately for men and women. Thus, when computing $\lambda^{\ddagger}(M), \lambda^{\ddagger}(F)$ we impose that the gender free $\lambda^{\ddagger} \in [\lambda^{G1}, \lambda^{G2}]$ and derive k value to be used in $\lambda^{\ddagger}(M), \lambda^{\ddagger}(F)$ calculations.

[8]Moreover, our estimates require sufficiently detailed GDP data that we are often forced to exclude those observations which Jerven (2013) argues should be taken least seriously.

Table 1: Summary Statistics of the Variables Used in EGI Calculation

	Observations	mean	sd	max
Male employees (millions)	3503	4.70	10.42	94.83
Female employees (millions)	3499	3.14	7.66	70.30
Male employees weekly hours worked	2089	43.33	6.46	68.00
Female employees weekly hours worked	2079	39.20	7.74	68.00
Male population ages 15-64 (millions)	7419	11.31	43.76	511.38
Female population ages 15-64 (millions)	7419	11.02	41.17	484.43
Male labour force participation rate	4585	0.74	0.08	0.96
Female labour force participation rate	4550	0.50	0.16	0.91
Male unemployment (%)	4538	7.60	5.77	36.96
Female unemployment (%)	4538	9.54	7.49	47.18
Male self-employed (millions)	2818	3.04	12.38	238.85
Female self-employed (millions)	2818	1.76	5.17	88.71
Male self-employed weekly hours worked	1508	42.87	4.87	68.00
Female self-employed weekly hours worked	1508	36.66	6.33	68.00
Gross value added (billions US dollar)	5501	244.23	1048.78	19838.00
Male employees mean nominal annual earnings	1713	18134.59	20444.12	131947.38
Female employees mean nominal annual earnings	1712	13808.04	16241.16	130554.45

Notes: The mean earnings are in current US dollar and only for countries for which reliable official exchange rate data was available from Penn World Table (PWT 10.0). Weekly hours greater than 68 are replaced with 68 hours per week.

weekly hours actually worked and mean nominal monthly earnings of employees (all disaggregated by gender and sectors). The weekly and monthly data are converted into annual numbers using ILO standard harmonization process: 5 days in a week, 4.33 weeks in a month and 52 weeks/12 months in a year. Earnings refer to gross earnings paid in cash and kind (including bonuses and gratuities and housing and family allowances paid by the employer directly to the employee) at regular intervals including pay during holiday/annual leave but excludes employers contributions to social security and pension schemes, and severance and termination pay.

For the global inequality calculations and inequality plots by income groups data on working age population, per capita gross domestic product and gross national income is collated from WDI, and labour force, labour force participation, unemployment, sectoral employment shares are taken from ILO. All the calculations are carried on data in national currencies except in Table 1 where the relevant variables are converted into current US dollar using official exchange rate data from Penn World Table (PWT 10.0). For some countries data for one or more variables are missing for one or more years subsequent to the first observation. We impute these missing observations using the Expectation Maximization procedure algorithm of Honaker and King (2010).

Our focus on compensation and value added has much in common with the literature on the overall labour share in that accurate measurement of compensation is essential. An important advantage of the labour-share approach is that ρ may be calculated using SNA and ILO data. That is the data used to calculate national accounts statistics.

3 Gender Inequality around the World

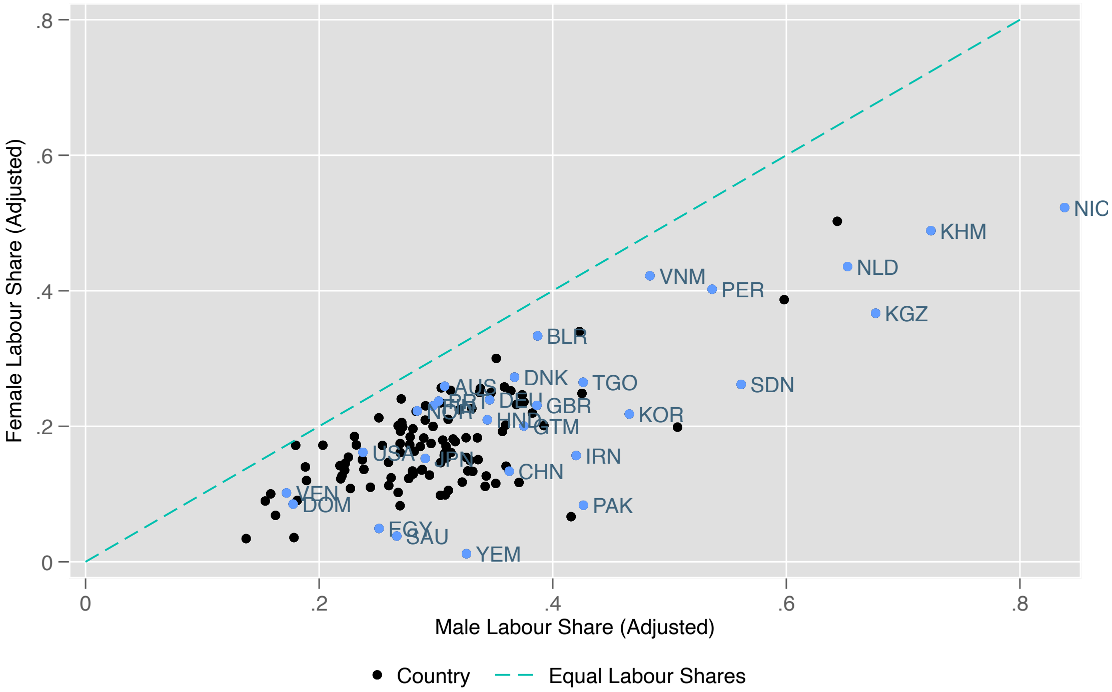
This section presents our new inequality data and establishes the existence of a large global gender gap. It begins by presenting the evidence that women do indeed have a lower labour share, the extent to which this varies across countries, and how this difference has tended to persist through time. It then moves on to document and discuss the aggregate extent of global EGI.

3.1 Labour Share Ratio

For the reasons discussed in section 2 we focus on the adjusted measure of EGI $\rho^{\ddagger}$, equivalent results for ρ^{U} are provided in appendix D. Figure 1 contains a scatter plot of the labour share of men in total value added, $\lambda^{\ddagger}(M)$ on the x-axis and the labour share of women $\lambda^{\ddagger}(F)$ on the y-axis by country for 2014. Hence, the dashed $45°$ line represents $\rho^{\ddagger} = 1$ or EGI (in means). It is immediately clear that in every country $\rho^{\ddagger} < 1$. Perhaps as expected, the countries closest to the line are Nordic countries such as Norway and Finland. Amongst large economies the labour share is highest in the Netherlands at 0.65. While high this figure is reasonably similar to the 0.6 for men and women estimated by Feenstra et al. (2015). The country with the absolute highest value of the female labour share is Nicaragua. The absolute value of the female labour share of total value added is also important as the relative shares of labour and capital share have important implications for inequality (see, Piketty and Saez, 2003). One, sometimes neglected, implication of this is that if gender differences in capital ownership mean that capital income disproportionately accrues to men, then a higher (female) labour share ratio will reduce the inequality of total (capital+labour) income.

There are a substantial number of countries where the labour share is low for both men and women. But Egypt, Saudi Arabia, and Yemen stand out given that the value added of men is close to the average but that of women is close to zero. This perhaps reflects a combination of both high inequality of opportunity and pay discrimination.

Figure 1: Scatter Plot of Female and Male Labour Share in 2014 (Adjusted Data)



Despite the substantial inequalities shown by figs. 1 and D.1 the mean country is more equal today than it has been in the past. Figure 2 presents box plots summarizing the cross-country distribution of $\rho^{\ddagger}$ between 1991 and 2018. Comparing the black horizontal lines in each we can see that there has been an increase of just under 10 percentage points in the cross-country median value of $\rho^{\ddagger}$ over the period from 0.52 to 0.62. The change in the mean was similar, increasing from 0.51 to 0.60.

The top and the bottom of each box depict the 75 and 25 percentiles of the distribution of $\rho^{\ddagger}$ in that year. These have seen similar increases to the median, with the 75^{th} perecentile increasing from 0.62 to 0.71 and the 25^{th} percentile increasing from 0.41 to 0.51. This suggests four conclusions.

First, the symmetry of the increase suggests that there has been neither rapid convergence by the most-unequal countries, nor particularly large improvements in the most-equal countries. Rather, there has been a relatively uniform increase.

Second, the exception to this broadly symmetric increase is the persistence of extremely low values of $\rho^\ddagger$ at the bottom of the distribution. This is captured by the lower adjacent values, depicted by the lower whisker in each year.[9] This highlights that the low female labour shares in countries such as Egypt or Saudi Arabia reported in fig. 1 are not a feature of that particular year. While, there is a large, unexplained, up-tick in 2018, otherwise the average of the first three years is similar to that of the 2015–2017.

Third, it highlights that overall progress has been slower than might be imagined given the increased awareness globally of gender inequality over the period, including initiatives such as the Millenium Development Goals. That is not to argue that the improvements are inconsequential, the 25^{th} percentile in 2018 is similar to the median in 1991, as is the 1991 75^{th} percentile and 2018 median. But, extrapolating from the average increase in $\rho^\ddagger$ of one third of a percentage point per year implies that it will be over a century before the median country (in our sample) has labour market gender equality.

Fourth, a further feature of the data is that the changes from year to year in the median, or other points on the distribution, are not always positive, reflecting that EGI sometimes decreases. One possibility is that this may be related to the economic cycle with gender equality suffering during recessions.

Fifth, and finally, looking at the blue and green boxes suggests that whilst there is evidence of a gentle upwards trend in the hourly wage ratio, the employment ratio has been almost constant. This suggests, in line with further results below, that while pay discrimination has been falling access to the formal labour market and other forms of gendered inequality of opportunity have not reducedin the same way.

To understand whether the distribution of EGI just reflects differences in economic development we consider the distribution by income group. Using the World Bank categorization, Figure 4 plots the distribution of the labour share ratio for High, and Upper and Lower Middle Income countries. Immediately, we can see that, as we expect, the High Income distribution is right-most, and the Lower distribution left-most. But, there is considerable overlap between, and heterogeneity within, categories.

[9] The whiskers report the upper and lower adjacent values which are, respectively, the values of x_i such that $x_i > 1.5 \times IQR + X_{75}$ and $x_i < X_{25} - 1.5 \times IQR$ respectively.

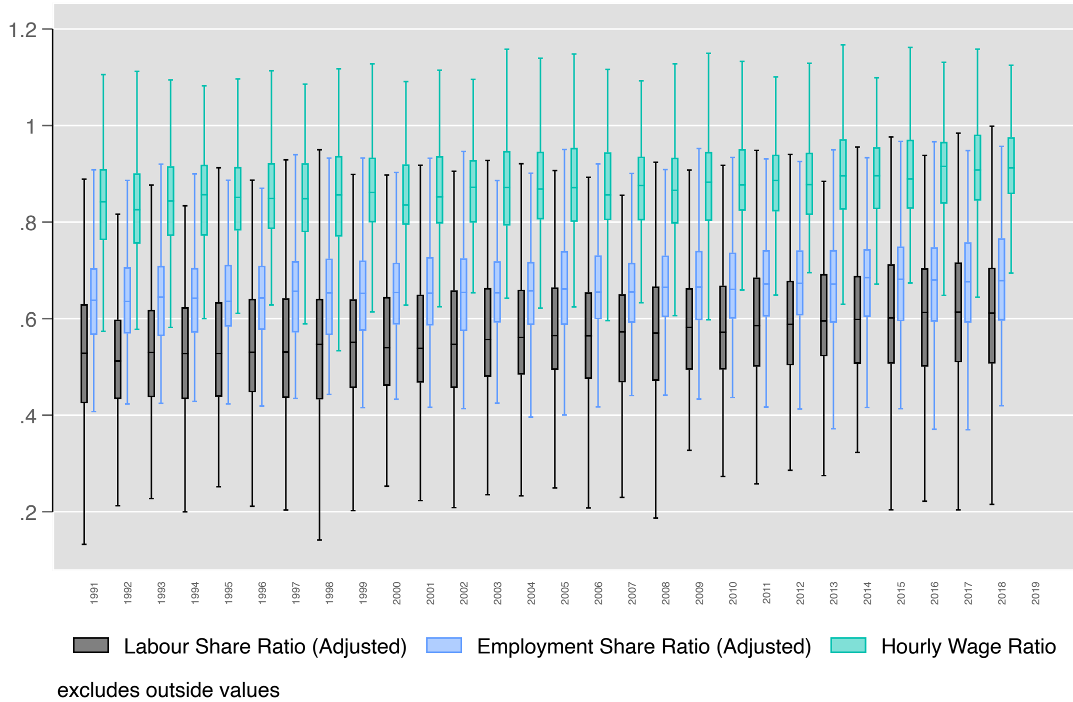
The difference between the High Income and the Upper Middle Income categories are relatively minor compared to between these and the Lower Middle category, but even this difference is second order compared to the within category variation. Thus, it would seem that Gender Inequality is not an automatic consequence of development. We return to this in Section 4.

Figure 2 treats countries as the unit of observation. This implicitly gives greater weight to women in smaller countries than those in big ones. This opens the possibility that changes in the cross-country distribution of $\rho^{\ddagger}$ fail to fully capture the change in the distribution of women's experience of EGI. To understand this fig. 3 reports kernel density estimates of the distribution of $\rho^{\ddagger}$ weighted by population. Comparing the distributions for 1994 (the solid blue line) and 2014 (the dashed blue line) we can see that our main conclusion is unaffected — there has been reduction in EGI across the distribution. However, comparing the weighted data with the unweighted data (the red lines) reveals that unweighted distribution will underestimate global EGI. In particular we can see considerable mass around 0.2 in 1994 for in the weighted data that is not apparent in the unweighted data. This is also reflected by a thicker left tail in the 2014 data.

3.2 EGI and Sectoral Composition

As a first step in understanding cross-country variation we now consider, as provided in fig. 5, the distribution of the ratio of annual wages for each sector separately. We focus on the ratio of annual wages $\rho^{\ddagger}\frac{N(M)}{N(F)}$ rather than $\rho^{\ddagger}$ as at the sector level the latter is hard to interpret as it will depend on both male and female employment in that sector. Looking at the data as a whole the first thing we note is that there is relatively little variation across sectors. The annual wage ratio is highest in construction and mining. This may reflect differences in the roles undertaken by men and women in those sectors. This seems particularly likely for construction where for a large share of observations female annual wages are higher than those of men. Looking at fig. D.3, which plots $\rho^{\ddagger}$ by sector we can see that the labour share ratio in those sectors is close to 0, implying that very few women work in those sectors but that those who do, do so in better paid managerial positions, etc.

Figure 2: Evolution of the Cross-Country Distribution of $\rho^{\ddagger}$.



Slightly more subtly we can see that, construction aside, the annual earnings ratio distribution is similar across sectors and there are very few cases where women's annual wages are larger than those of men. Our interpretation of this is that the absence of large differences across countries in the ratio re-emphasises that differences in labour market participation and hours worked are more important drivers of EGI than differences in wages.

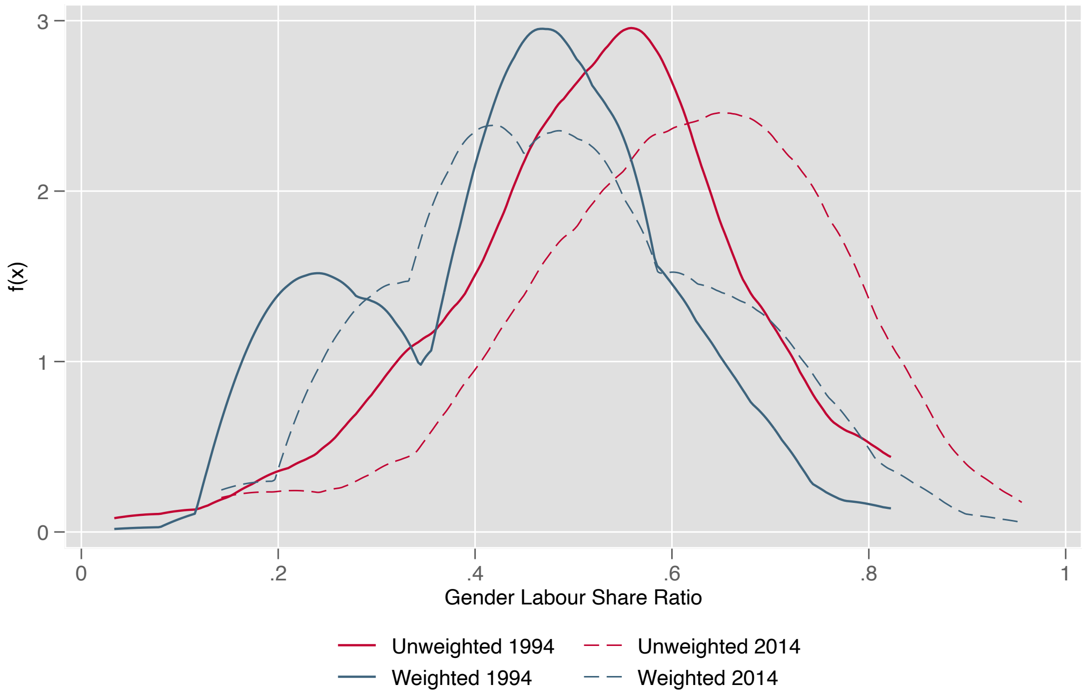
EGI is only one aspect of gender inequality, however, and in the next section we show that our measure is correlated with other dimensions of inequality suggesting that gender inequality, more broadly defined, may also be getting worse not better.

3.3 Aggregate Inequality

Having established the key features at the country level the remainder of this section focuses on the distribution of EGI at the population level. That is, we ignore the average differences between countries that were previously our focus, and now consider the total global extent of EGI ignoring national borders.

Differing population sizes and population growth rates mean that the rightwards

Figure 3: Population Weighted Distribution of $\rho^{\ddagger}$ in 1994 and 2014.



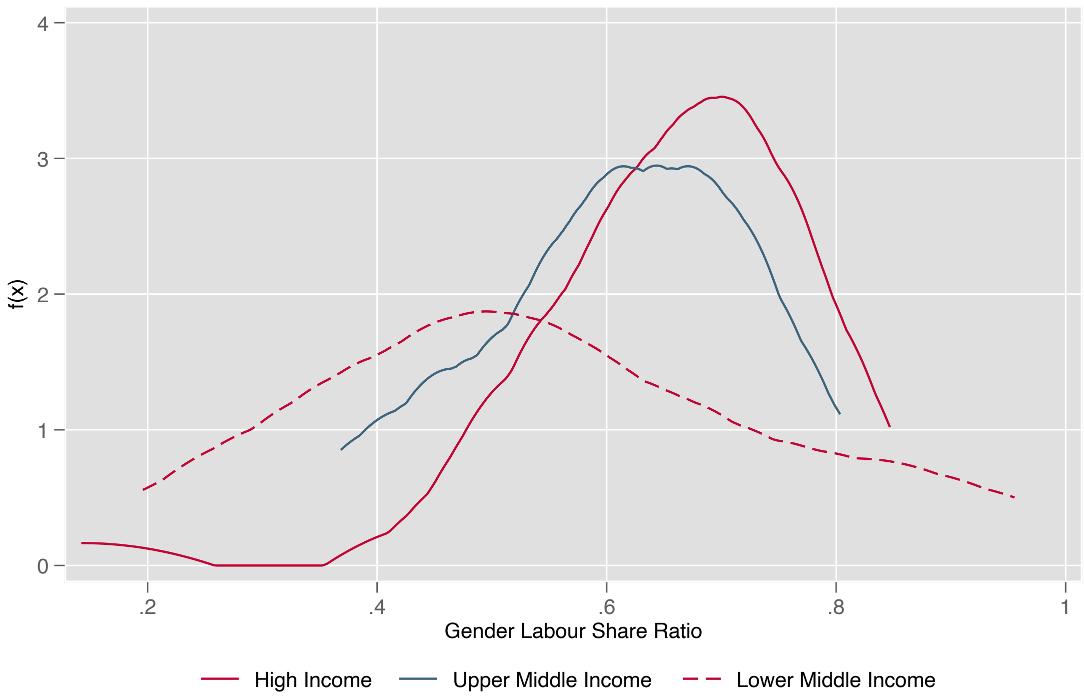
shift in the weighted and unweighted distributions of $\rho^{\ddagger}$ reported in fig. 3 may imply that while EGI has been reduced for the average woman, total EGI may have increased.[10] Measuring overall EGI requires calculating the total deviation from equality in each country and aggregating these across countries. It is convenient to interpret the resulting quantity as the number of unpaid equivalents.

Consider the following simple example — if for a total, evenly split, population of 200 men and women total compensation for men is 100 and 20 for the women, then this would imply a labour share ratio of $\rho = {}^{20}/_{100} = 0.2$. While one interpretation is that each woman receives one fifth of the income of each man, an equivalent is that one of the women earn the same as their male equivalents and the others are unpaid. The advantage of this is that it provides a straightforward summary statistic.

Thus, for each country c in our data the number of unpaid equivalent women is simply $(1 - \rho_c \frac{N_c(M)}{N_c(F)})P_c(F)$. The logic behind the multiplication of $\rho_{c,t}$ with the

[10]The literature on aggregate global income inequality shows that differences between nations are able to explain the majority of global inequalityMilanovic (2015). Thus, Jones (1997), Milanovic (2002) and Sala-i Martin (2006) show that, despite rises in within country inequality rapid growth in China and to a lesser extent India have reduced total world inequality.

Figure 4: Distribution of $\rho^{\ddagger}$ by Country Income Group.

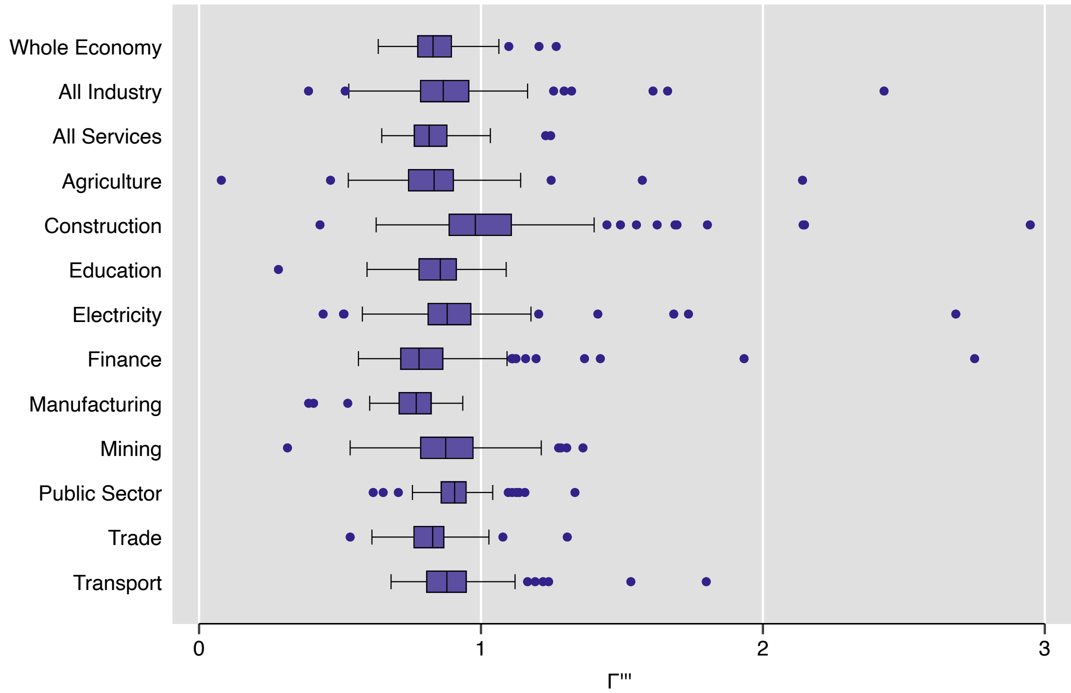


employment ratio is explained below. The total number of unpaid equivalent women in the world, aggregate world EGI, is then simply the sum over the set of all countries C. That is,

$$\Gamma_t = \sum_{c \in C} P_{c,t}(F) \left(1 - \rho_{c,t} \frac{N_{c,t}(M)}{N_{c,t}(F)}\right), \tag{11}$$

where $P_{c,t}(F)$ represents working-age women population in country c in time t, $N_{c,t}(M)$ and $N_{c,t}(F)$ denote total employment of men and women respectively. In a country with no inequality whatsoever $(1 - \rho_{c,t})$ must be zero. Gender bias against women would mean $(1 - \rho_{c,t}) > 0$ and vice versa. The multiplication of $\rho^{\ddagger}_{c,t}$ in eq. (11) by the ratio of male to female employment removes the effects that come from differences in the number of employed men and women. Thus, $\rho^{\ddagger}_{c,t} \frac{N_{c,t}(M)}{N_{c,t}(F)}$ shows the inequality that only arises from differences in working hours and hourly earnings. This, quantity is the ratio of female to male incomes and thus is similar in its interpretation to the GPG, as captured by η in eq. (B.4), albeit parameterized differently and our measure takes care of the differences in average working hours among self-employed; See,

Figure 5: Cross-Country Distribution of $\rho^{\ddagger}\frac{N(M)}{N(F)}$ by Sector



appendix B for a full discussion of this relationship.

Of course, as discussed above, EGI is driven not only by differences in earnings per hour and hours worked but also by differences in unemployment and labour market participation. Another advantage of our Γ_t is that it allows us to consider EGI amongst those out of the formal labour market. We can decompose Γ to assess the relative importance of each of these factors.

The number of unpaid equivalent amongst the employed due to per hour earnings differences is given by:

$$\Gamma_t^W = \sum_{c \in C} \left(N_{c,t}^E(F) + N_{c,t}^{SE}(F)\right) \left(1 - \frac{\overline{w}_{c,t}^E(F)}{\overline{w}_{c,t}^E(M)}\right) \tag{12}$$

Where $N_{c,t}(F) = \left(N_{c,t}^E(F) + N_{c,t}^{SE}(F)\right)$ is the total number of women in employment. Next we compute the number of unpaid equivalents amongst the employed due to differences in both per hour earnings and hours worked.

$$\Gamma_t^{WH} = \sum_{c\in C} N_{c,t}(F)\left(1 - \rho^{\ddagger}_{c,t}\frac{N_{c,t}(M)}{N_{c,t}(F)}\right). \tag{13}$$

We can analogously compute the number of unpaid equivalents in the female labour force as a whole by re-computing eq. (13) now including the unemployed:

$$\Gamma_t^{WHU} = \sum_{c\in C} \left(N_{c,t}(F) + U_{c,t}(F)\right)\left(1 - \rho^{\ddagger}_{c,t}\frac{N_{c,t}(M)}{N_{c,t}(F)}\right). \tag{14}$$

Where $U_{c,t}(F)$ is the number of women in the labour force but unemployed. This calculation thus assumes, conservatively, that were all unemployed women to be employed that they would face the same average level of wage inequality as faced by employed women since the actual inequality they would experience is unobservable.

A key form of EGI, as discussed previously, is gender difference in rates of participation in the formal labour market. Looking at table 1, we see that the cross-country average labour force participation rate is 74% among men but only 50% for women. As discussed in Appendix A, this is unlikely to mean that 24% of the women are economically inactive, they are much more likely to be working, but in unpaid home production, childcare and similar activities traditionally assigned to women. That is they are working, but not receiving wages.

We thus ask two related questions. First, what would be the number of additional unpaid equivalents recorded if female labour-market participation rates were the same as men's?

This is given by modifying eq. (11) to adjust for the male labour force participation rate:

$$\Gamma_t^{WHUP} = \sum_{c\in C} P_{c,t}(F)\Upsilon_{c,t}(M)\left(1 - \rho^{\ddagger}_{c,t}\frac{N_{c,t}(M)}{N_{c,t}(F)}\right), \tag{15}$$

where $\Upsilon_{c,t}(M)$ is the male labour market participation rate. Like eq. (14) this is predicated on the idea that were more women to participate in the formal labour market they would face similar rates of discrimination to those who currently do. Again this is conservative as one can imagine that women on the margin of labour market participation are more vulnerable to discrimination than the average women in employment. By imputing to these women the discrimination faced by employed women we are saying that women engaged in subsistence agriculture or home-production face average discrimination at least as great of the average employed women. Consider the the example of home production. We assume the running of a household in terms of cooking, cleaning, and care are equally valuable to men and women. Combining this assumption with the well documented fact that women do disproportionate amounts of home production in almost every society then their labour share in home production will be much lower. A similar argument applies to subsistence farming.

The second question is what would be the number of unpaid equivalents if we considered all working-age women? This quantity will capture the reality that there are very few working age women who are involved neither in market or non-market production. The latter may involve subsistence agriculture and or home-production such as cooking, cleaning, or care. There are of course some exceptions to this. Some women, like some men, live lives purely of leisure. But, these will be disproportionately concentrated in the richest countries where male labour market participation rates are very high and thus this final section of the female population will be small anyway. This quantity is given by $\Gamma_t - \Gamma_t^{HWUP}$.

We separate $\Gamma_t - \Gamma_t^{HWUP}$ and $\Gamma_t^{HWUP} - \Gamma_t - \Gamma_t^{HWU}$ as they involve slightly different assumptions. In the case of the latter, the counterfactual is clear. In a gender equal society rates of labour market participation would be the same. Thus, calculating what inequality would be if women participated at the same rate as men is relatively natural as it only involves the observed rate of labour market participation for men and women, and differences in women's observed employment income. The former quantity requires instead a comparison with men who are also not in the labour market. Now the assumption is that in which the labour market is unequal that home production is not more equal.

Figure 6 reports Γ_t and its composition for each year fron 1992 to 2017. The data suggest that in $\Gamma_{1991} = 800$ million, rising to a little over one billion by 2017. This suggests that global EGI increased over the period by around one quarter. The purple line plots (on the right-hand y-axis) the global working age female population. Over the same period this has increased from around 1.6 billion to 2.4 billion. This means that while in 1991 EGI was equivalent to one in two working age women being unpaid, by 2017 it had fallen to just over two in five. Put differently, in 1991 it took on average the earnings of two women to equal those of a single man, but by 2017 the earnings of five women were equal to those of just over two men.

This increase in the absolute number of unpaid equivalents and decrease in the proportion thereof reflects that the increase in Γ_t is because the rate of population growth has outstripped the slow increase in $\rho^{\ddagger}$ documented in fig. 2.

Given that the global working age population will continue to increase, albeit at a decreasing rate, for at least the next 20 years, with a total forecast increase of around 500 million working age women by 2042, if the historic growth of $\rho^{\ddagger}$ is maintained then it seems possible that Γ will not fall for quite some time to come.[11]

Considering the composition of Γ makes clear the relative roles of gender inequality amongst those in the labour market, and EGI due to differences in labour-market participation. The purple region describes the number of unpaid equivalents solely due to differences in average hourly wages, Γ_t^W. We can see that in 1992, that this alone accounted for around 164 million unpaid equivalent women. There has, however, been a small decline in this number over the period with it falling to 135 million by 2017. This perhaps reflects reductions in pay discrimination in both high income countries and elsewhere.

On the other hand, the number of unpaid equivalents associated with differences in overall earnings between employed men and women has grown. At the beginning of the period Γ_t^{WH} accounted for around half of, or 400 million, the total number of unpaid equivalents. While, it has increased by around 40 million over the subsequent 25 years Γ^{WH} has accounted for declining share of Γ. A corollary of this is that the role of hours worked $\Gamma^{WH} - \Gamma^{W}$ has increased over the same period.

A quantitatively small, but consistent, source of EGI is excess unemployment amongst women as depicted by the light blue seam in the picture. This shows that $\Gamma^{WHU} - \Gamma^{WH}$ has grown over the period from around 25 million unpaid equivalents to

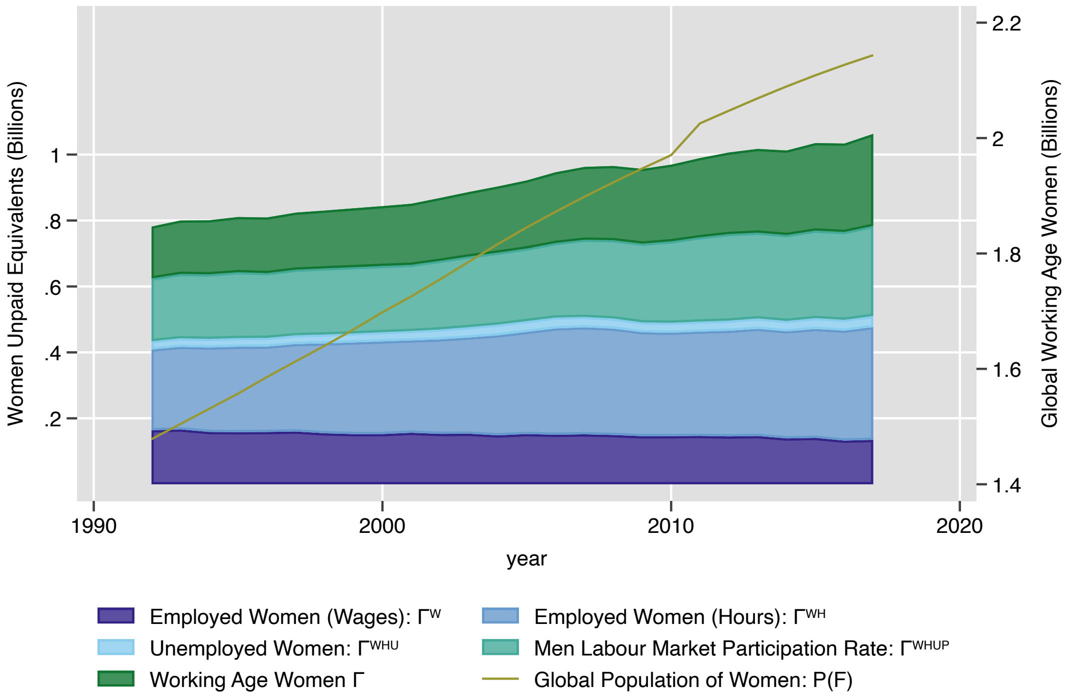
[11]UN World Population Prospects (2017) via ourworldindata.org, accessed Jun 2022.

34 million.

The two green areas report unpaid equivalents associated with women's work outside of the labour market. The light green area reports $\Gamma^{WHUP} - \Gamma^{WHU}$ or the number of unpaid equivalents associated with differences in men and women's labour market participation. That is, the number of additional unpaid equivalents we would record if women had the same rate of labour market participation as men, but faced the same rate of inequality as women already in the labour market. We can see that this is an important form of labour market inequality, equivalent to 190 million unpaid women at the beginning of our period and increasing to around 273 million by 2017.

Finally, the dark green area reports $\Gamma_t - \Gamma_t^{HWUP}$, the number of unpaid equivalents if we assume that all working age women not in employment face, on average, the same level of discrimination as those in the labour market. As discussed above this quantity requires the additional assumption that the population of global women who work either in employment or in some form of home production is small. On this basis the additional number of unpaid equivalents was 157 million in 1992 and had risen to just under 280 million by 2017. One way to interpret the scale of $\Gamma_t - \Gamma_t^{HWUP}$ is that it would require over 13% of working age women worldwide to be women of leisure which seems too high to be plausible. While, relaxing our assumption would no doubt reduce our estimate slightly there is good reason to believe that our estimate is in fact an underestimate given that the assumption that women not in employment face equal discrimination to those who are not is probably a conservative one.

Figure 6: Aggregate Global Economic Gender Inequality



4 Causes of Gender Inequality

This section studies the causal determinants of EGI. In particular, we ask whether increased incomes improve EGI, and whether democracy leads to reductions in discrimination. These questions are important for at least two reasons. Firstly, if it is the case that increased incomes rapidly lead to improvements in gender equality then this suggests that convergence in income per capita will lead to a rapid reduction in aggregate inequality. Equivalently, it also means that women in the LDCs will benefit substantially from growth. Alternatively, if improvements in living standards alone do not lead to reductions in inequality then this suggests that women will benefit comparatively little from development and that aggregate inequality may continue to rise for the foreseeable future.

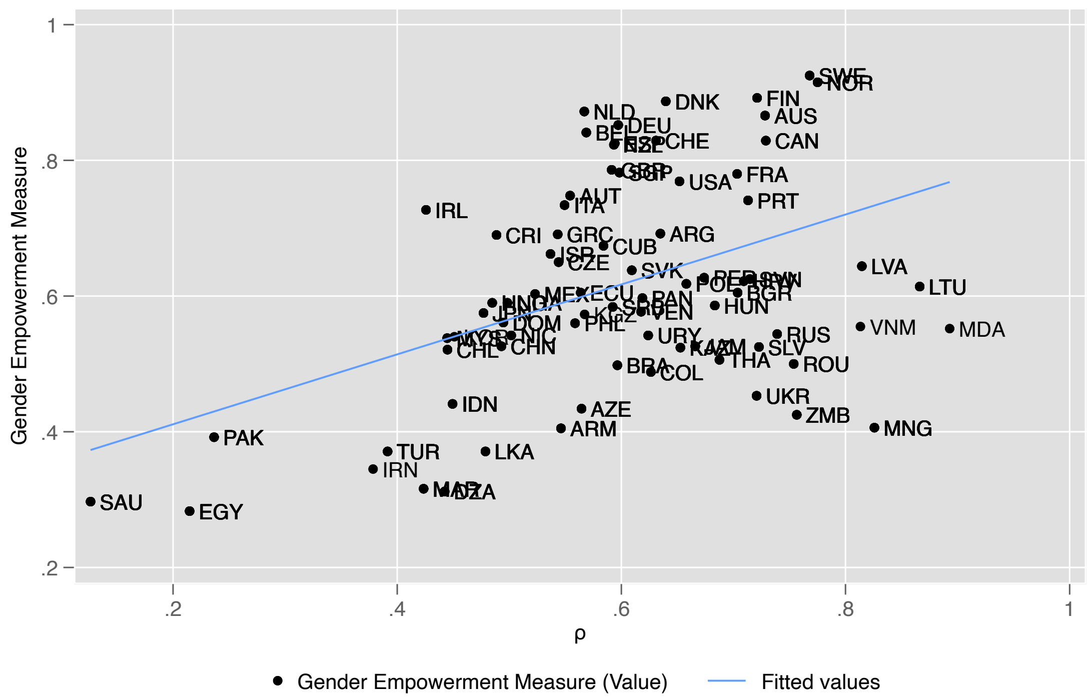
A prominent literature has considered how female empowerment may lead to economic development, which may in turn lead to further improvements in Gender Equality. A key issue in this literature is whether such feedback effects between gender empowerment and growth are sufficiently large to give rise to a virtuous circle of increasing women empowerment and increasing growth. Doepke et al. (2012) outline a model in which this takes place, and Fernández (2014) presents theory and evidence that as development takes place men become increasingly concerned about their daughters, leading to greater property rights for women. de la Croix and Vander Donckt (2010) consider how greater equality would lead to lower fertility thus hastening the demographic transition crucial for development. Cavalcanti and Tavares (2016) provide evidence that gender discrimination leads to lower per capita incomes. Relatedly, Doepke and Tertilt (2019) consider theoretically the effects of targeting transfers to women on development. Seguino (2000), Blackden et al. (2006) present cross-country evidence that there is a positive relationship between the two. But, Duflo (2012) cautions that the empirical evidence suggests that feedback effects may be insufficient for 'take-off' and that a 'continuous policy commitment to equality for its own sake may be needed'. Silva and Klasen (2021) provide a recent review of the theoretical literature. Klasen (2018), in their more empirically orientated survey, argues that many estimates of the effects of eliminating gender inequality are likely too large.[12]

A second dimension of development is the expansion of individual rights and political agency. Women in less developed countries often have even fewer rights and less political power than their peers in mature democracies as well as in some cases suffering limited physical integrity and various forms of social control. Duflo (2004) studies the random reservation of seats for women on Village Councils in India and finds that increased female political power leads to the greater provision of infrastructure targeted at women. Bhalotra and Clots-Figueras (2014) find that increased political power of women leads to improved health outcomes, while Bhalotra and Rawlings (2011) show that gender inequality in investments in the health of women and girls is a key margin for the inter-generational transmission of health.

[12] A largely separate literature studies the effects of gender equality on growth. Partly due to limited data availability, much of it has focused on the effects of educational inequality on growth Klasen (2002), Lorgelly (2010) but Dollar et al. (1999), Klasen and Lamanna (2009) do find that gender differences in labour market participation also retard growth. Others have studied the effect of trade-liberalization or globalization on gender equality. Oostendorp (2009a) finds that growth as well as trade and investment liberalization tend to be correlated with reductions in gender inequality – particularly in poorer countries. Additional evidence is provided by Neumayer and de Soysa (2011), Chen et al. (2013), Potrafke and Ursprung (2012), Cooray and Potrafke (2011), Richards and Gelleny (2007).

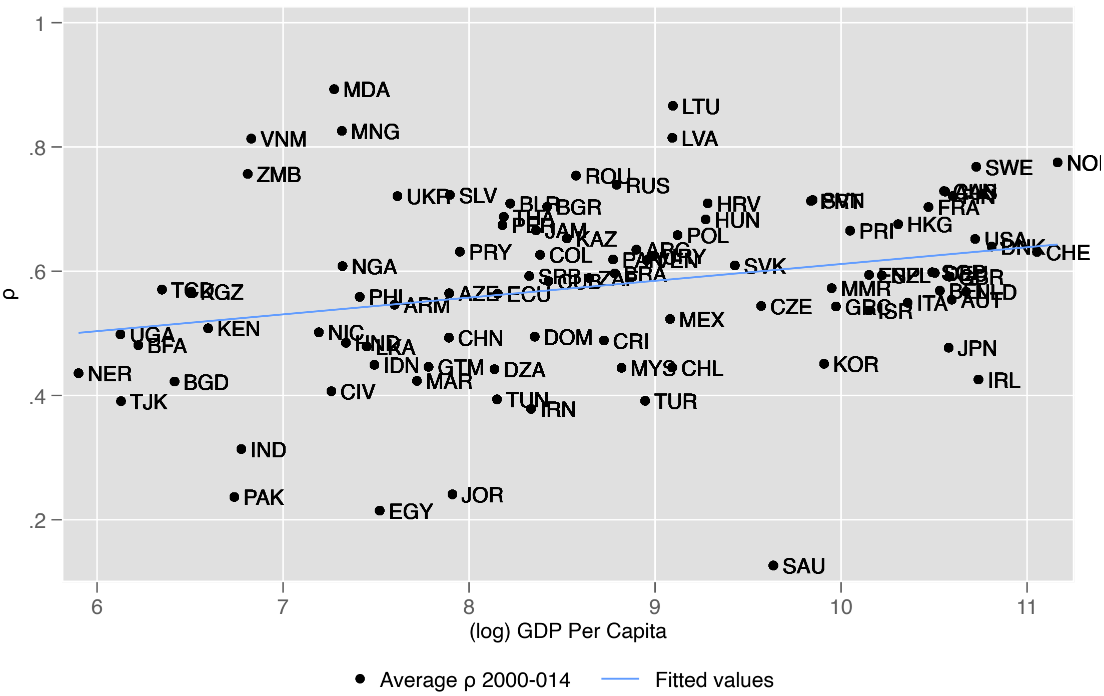
We do not study these dimensions directly but as can easily be seen in fig. 7, which compares $\rho^{\ddagger}$ with the Gender Empowerment Measure (GEM) of the UN employed by Doepke et al. (2012), there is a strong positive correlation between EGI and other dimensions of gender inequality. Doepke et al. (2012) show that whilst the slope describing the relationship between GEM (or several disaggregated measures) and per capita incomes is positive it 'is not very steep, in particular when moving from middle-income to low-income countries.' Considering fig. 8 we can see that this is similarly true for labour market discrimination.

Figure 7: $\rho^{\ddagger}$ is Positively Correlated with GEM



We build on this work by providing a causal analysis of the effects of income, $y_{c,t} = \frac{Y_{c,t}}{P_{c,t}}$, and democracy $D_{c,t}$, on labour-market gender-equality at the national level over the long-term. Both the previous literature and our intuition suggest that both income and democracy may be endogenous to EGI. But, these are two distinct causal relationships and thus, as such, we first consider them separately so that it is possible to interpret the results in a straightforward manner (Angrist and Pischke, 2009, Chapter 4). We do this by estimating two separate sets of 2SLS regressions, with separate instruments.

Figure 8: $\rho^{\ddagger}$ is Positively Correlated with Income



4.1 Construction of instruments and data

Our instruments for income capture external macroeconomic shocks which cannot be plausibly driven by domestic changes in gender equality. We measure these shocks with four different variables. The first is the gravity-weighted average of trading partners' GDP growth. To capture better the fact that many developing economies are particularly sensitive to changes in agricultural and mineral commodity prices we construct two instruments that capture terms of trade changes based on ex-ante shares of each commodity in trade as in Deaton and Miller (1996). Finally, we use the presence of IMF or WB emergency assistance, which tend to be a response to financial crisis to capture the effects of crises shocks. For Democracy, our instruments follow the strategy of Acemoglu et al. (2019). We code a country c as democratic in year t if Freedom House regards it as *Free* or *Partially Free* or Polity IV gives it a *non-negative* democracy score. Based on this and following Acemoglu et al. (2019) regional waves in transitions to and away from democracy are used as an instrument for country-level democracy. Our democracy instruments are thus exactly the same as in Acemoglu et al. (2019) but we update the data upto the year 2016. Appendix C describes the construction of each of these instruments in more detail.

For the causal analysis of income per capita and democracy on EGI, besides the instruments as highlighted above, we use a host of variables on democracy and its types, women representation in parliament, globalisation, financial liberalisation, women access to education and their employment share in different sectors of the economy. Table 2 gives a brief introduction of each of these variables and the data sources we rely upon.

Table 2: Variables used in Causal Analysis of EGI and their Data Sources

Variables	Type	Source
Gender empowerment measure	Continuous index (0–1)	Doepke et al. (2012)
Trade flows used in Gravity model	Current US $	IMF Direction of Trade
Gravity variables (trade partners GDPs, weighted distance, WTO/regional trade agreements memberships: Appendix C)	Different	Conte et al. (2021)
Gravity variables (indicators on common language, religion, legal systems, colonial history etc.: Appendix C)	Binary indicators (0,1)	Conte et al. (2021)
Global primary commodity prices (used in agriculture and mineral price indexes: Appendix C)	Current US $	IMF; World Bank
IMF programs and World Bank projects (used in construction of crisis instrument: Appendix C)	Binary indicators (0,1)	Dreher (2006a); Boockmann and Dreher (2003a)
Democracy measure ANRR (updated to 2016)	Binary indicator (0,1)	Acemoglu et al. (2019)
Regional democratization waves (updated to 2016: Appendix C)	Continuous index (0–1)	Acemoglu et al. (2019)
(log) GDP per capita	PPP (current international $)	WDI
Electoral democracy; Liberal democracy; Egalitarian democracy	Continuous indexes (0–1)	V-DEM project Coppedge et al. (2016)
Women's political empowerment; Women's civil liberties	Continuous indexes (0–1)	V-DEM project Coppedge et al. (2016)
Women's civil society participation; Political participation	Continuous indexes (0–1)	V-DEM project Coppedge et al. (2016)
Women lower house membership; Women upper house membership	Ratio	Inter-Parliamentary Union
KOF globalization (measures economic, social and political dimensions of globalisation)	Ranking (0–100)	Gygli et al. (2019); Dreher et al. (2008)
Chinn-Ito (measuring a country's degree of capital account openness)	Continuous index (0–1)	Chinn and Ito (2006)
Financial development (measuring access, depth and efficiency of financial institutions and markets)	Continuous index (0–1)	IMF
Ratio women employment industry; Ratio women employment services	Ratio	ILO
Ratio of girl primary enrollment to boys	Ratio	World Bank Education statistics

Notes: Inter-Parliamentary Union: https://www.ipu.org/our-impact/gender-equality/women-in-parliament

We assume the population relationship is linear. Then we have:

$$\rho^{\ddagger}_{c,t} = \beta y_{c,t} + \gamma D_{c,t} + \sum_{q=1}^{Q=5} \phi_q \rho^{\ddagger}_{c,t-q} + \mu_c + \tau_c t + \psi_{c,t} + \varepsilon_{c,t}. \tag{16}$$

Where we will assume, for now, that $\psi_{c,t} \sim N(0, \Sigma)$ with Σ clustered by country. The need to include fixed effects, μ_c, means that our estimates of ϕ will be biased and thus so will our other parameters. However, on average we have just under 20 observations per country and thus we expect the Nickell bias to be of the order $1/20$. We prefer this to the alternative of GMM estimation. However, in Table D.3 we employ the strategy of Acemoglu et al. (2019) and show a range of unbiased estimates obtained by fixing ϕ to a range of values.

We include fixed effects to allow for time-invariant country-specific factors that may determine EGI. Thus, our estimates will describe the impact of an increase in income on EGI *within* a given country, and not between variation. To capture any short run dynamics such as business cycle effects we include a set of lags of $\rho^{\ddagger}_{c,t}$. This will also help capture the reality that, given it represents in part a complex-nexus of legal, cultural and socio-economic factors gender inequality tends to change slowly.[13] We also, in our preferred specifications include country-specific linear trends $\tau_c t$ and year fixed effects. As discussed these allow for global changes in attitudes to gender inequality, the impact of pressure by external governments and NGOs, and so on.

One potential concern is that there may be a direct impact of the instruments on EGI beyond their effect via income, i.e. $Cov(Z, \rho^{\ddagger}|y) \neq 0$ where Z is the set of instruments. We test this by estimating the reduced form $\rho^{\ddagger}_{c,t} = \beta y_{c,t} + \gamma D_{c,t} + \sum_{q=1}^{Q=5} \phi_q \rho^{\ddagger}_{c,t-q} + \sum_{q=1}^{Q=5} \pi_q Z^q_{c,t} + \mu_c + \tau_c t + \psi_{c,t} + \varepsilon_{c,t}$. We are unable to reject the null joint hypothesis that the elements of π_q are equal to 0. This suggests that there is no such direct impact.

We begin by estimating a simplified specification in which we omit time-trends and year fixed-effects. That is, we set $\tau_c t$ to 0. We also ignore, for now, concerns about endogeneity and report simple OLS estimates. These results are reported in the first

[13] For example, Doepke et al. (2012) discuss the slow evolution of women's legal and political rights in the United States and the United Kingdom since the 17^{th} Century.

column of Table 4. We see that ϕ_1 is positive and precisely estimated, as is ϕ_2. ϕ_3–ϕ_5 are smaller and less precisely estimated. The sum of their coefficients is around 0.7 suggesting substantial persistence in EGI.

The coefficient on (log) GDP per Capita is also significant and precisely estimated, but is perhaps surprisingly small at 0.02. This coefficient implies that a tripling of income per capita will lead to only 12.6% increase in $\rho^{\ddagger}$ in the long run.[14] Column 2 reports the results now including year fixed effects. The estimated coefficient on β is now negative. This suggests, that increases in income are associated with reductions in EGI. This result is perhaps counter-intuitive, but it is robust to including country-specific trends (column 3) or allowing income to affect EGI with a lag (column 4). We attribute this to the combined effects of reverse-causality and the fact that growth is, other things equal, expected to be faster at lower levels of development.

To address this, Columns 5–8 of Table 3 relax the assumption that growth is exogenous. Columns 5 and 6 report specifications as in Columns 1 and 2 excluding year fixed effects and trends, and just country-specific trends respectively. Now the estimated coefficients are larger although, unsurprisingly, less precise. In our preferred specification in columns 7 the estimate is 0.073 implying a long run effect of $0.073/(1 - 0.242 - 0.172 - 0.042 - 0.069 + 0.002) = 0.153$, implying that an increase in per capita incomes from the bottom of the World Bank's lower-middle income category (\$1,046) to the bottom its rich income category (\$12,696) would be sufficient to raise $\rho^{\ddagger}$ from approximately 0.4 as in Tunisia, Iran, or Turkey, to close to 0.8 as in Norway, Sweden, or Latvia.[15] Such an increase in living standards is by no means automatic or swift but unlike the OLS coefficients these suggest that improvements in income alone could lead to wholesale improvements in women's lot.

We now turn to the effect of democracy on EGI, reported in table 4. Our instruments for Democracy follow the strategy of Acemoglu et al. (2019) and use the current democratic status of other countries in the same region having similar political history, that pre-sample, had the same democratic status. Full details are provided in Appendix C. Columns 1–3, report OLS estimates as in table 3 and columns 4–6

[14] The long-run effect is given by $\hat{\beta}/1-\sum_{q=1}^{Q=5}\hat{\phi}_q$.

[15] $(\ln 12,696 - ln 1,046) \times 0.153 = 0.382$

2SLS estimates. The results are surprising. Whilst, Electoral Democracy is associated with an increase in EGI in the simplest specifications, once we add in year effects or country-specific trends, the OLS and 2SLS estimates are consistently negative.

To understand this better, in table 5 we consider two possibilities. First, we augment our model to include both income and democracy. Column 1 reports OLS estimates of this specification, as above, the estimated coefficient on income is negative and now the effect of democracy is estimated imprecisely. Column 2 includes the interaction of Democracy and income. Now both income and democracy are negative and precisely estimated, while the interaction term is positive. This suggests that heterogeneous effects of democratization at different levels of income, or vice versa. Column 3 reports results for the same model, now estimated by 2SLS where we instrument for democracy and income simultaneously, as well as for their interaction, using the interaction of our two sets of instruments.

The interaction terms are hard to interpret directly, and so in figs. 9a and 9b we report the estimated marginal effects of democracy at different levels of income for each estimator. The results of both are similar, albeit the confidence interval on the OLS estimates is narrower as usual. The results suggest that at low incomes the effect of democracy are negative, but that at higher incomes they are positive. Looking at the x-axis more closely we can see a clear negative effect in countries with a GDP per capita of below $3,000 and a positive effect above $22,000 PPP.

The 2SLS estimates suggest that we cannot rule out a negative causal effect of democratization at any income level, but that there is some evidence of a positive effect above around $22,000. One possibility is that the contemporaneous effect of Democratization on EGI reflects short-run effects associated with political change rather than overall impact. To test this, in column 3 and fig. 9c we report the results of a specification in which $\rho^{\ddagger}_{c,t+5}$ is now the dependent variable, and to preserve sample size, we omit the lags of $\rho^{\ddagger}_{c,t}$. The estimated coefficients on Democracy and Income now tell us the conditional correlation of a change in income at time t with EGI in period $t+5$. We see that the overall pattern is similar. The direct effects are still negative, but the interaction is positive. Looking at fig. 9c we see that the estimated relationship is now stronger, and more precise. The qualitative result is similar, however, suggesting democratization at income per capita incomes below \$3,300 reduces $\rho^{\ddagger}$, but above around \$9,000 increases it.

Of course Democracy is a multifaceted concept and how it may be best quantified is contested.[16] We address this by repeating our analysis using several alternative measures of democratization. One debate in the literature has been whether Democracy should be seen as a binary or a continuous concept. Thus, we report in column 4 results using the binary measure of Acemoglu et al. (2019). The results are similar to those using the continuous measure. In columns 6–11 we study the effect of a range of alternative facets of democracy as measured by Coppedge et al. (2016). The first two are Liberal Democracy and Egalitarian Democracy both of which lead to similar estimates. Then we consider VDEM's measures of women's empowerment, their civil liberties, and their participation in civil society. These coefficients are similar in magnitude and precision to the others, again suggesting that democratization may have negative consequences at low per capita incomes. Finally, in columns 12 and 13 we report results using the share of women representatives in the lower and upper houses of the legislature respectively. The estimated coefficients are again similar although somewhat more noisily estimated. Table D.1 reproduces columns 5–13 for the specification reported in column 4 using $\rho^{\ddagger}_{c,t+5}$ as the dependent variable, and again we obtain similar results.

Our initial specification is deliberately parsimonious. Whilst, including additional controls might improve the precision of our estimates, if we include other variables that might also be driven by income or democracy on the right-hand side then we will again have an endogeneity problem.[17] This precludes including in our main regression other potential determinants of gender equality that have been discussed in the literature, such as Globalization and Trade or Financial Liberalization (see, Oostendorp, 2009a, Potrafke and Ursprung, 2012), as there are good reasons to imagine these may well be endogenous to income and democracy. However, having obtained consistent estimates of the effects of income and democracy we now, include measures of a number of other causes discussed in the literature and related outcomes, such as Globalization, in

[16]See Acemoglu et al. (2019) and Boese and Eberhardt (2021) and therein for discussion.
[17]This is the so-called *Bad Control* problem.

Table 6.

We begin by studying the role of Globalization, using the KOF index, as in (Potrafke and Ursprung, 2012) . The estimated coefficient is close to zero but not precisely estimated. Suggesting that Globalization is not associated with EGI beyond its effect on income or democracy, as captured by our measure.

One much debated policy available to governments, and historically encouraged by multilateral organizations is financial reform or liberalization. Columns 2 and 3 report specifications including the indices proposed by Chinn and Ito (2006) and Abiad et al. (2010). While we find no effect of capital account openness, as captured by the Chinn-Ito index, the Financial Development index of Abiad et al. (2010) suggests that a reduction in financial repression improves EGI beyond any impact on income itself. The estimated coefficient is indeed a large one, similar in magnitude to that of income estimated in table 3.

As discussed above, one argument that has been made in the literature is that EGI might promote the growth of the manufacturing sector and exports. Columns 6 and 7 thus include the gender imbalance in the manufacturing and service sectors (the ratio of female to male employees in these sectors). The ratio in industry has no effect, but more women working in services is associated with a higher value of $\rho^{\ddagger}_{c,t}$. One explanation for this is that greater employment of women in services represents greater female formal labour market participation.

The fraction of the working-age population captures both the relative scale of the burden of caring for children and the elderly which disproportionately fall upon women. It may also capture the relative scarcity of labour — a higher dependency ratio may reduce underemployment. The estimated coefficient is negative as expected, and precise. It is however, in common with many of our other estimates, relatively small. A decrease in the dependency rate of 10% is only associated with an increase in gender equality of 0.017.

In column 9 we consider another key dimension of gender equality, education. Taking the ratio of female to male primary enrolment as our measure we find that an increase in ratio of girls to boys attending primary school from one half to one, would reduce EGI by 0.05. This, is a relatively large effect although here we cannot identify any causal relationship.

Some variables such as the primary education enrolment ratio, or financial development might impact upon EGI with a lag. Table D.2 reports results for $\rho^{\ddagger}_{c,t+5}$ as above. The qualitative pattern in the results is similar, although the coefficients are now somewhat larger.

In sum, there would seem to be some other important correlates of EGI even conditional on income and democracy. In particular, demographic structure and access to education are important. As is financial development, suggesting that better functioning capital markets improve women's economic outcomes beyond any impact on overall living standards.

Table 3: Effects of Income on Economic Gender Inequality

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
(log) GDP, per capita (PPP $)	0.020***	−0.010**	−0.011**		0.040***	0.076*	0.073*	
	(0000)	(0042)	(0028)		(0000)	(0073)	(0091)	
$\rho^{\ddagger}_{c,t-1}$	0.296***	0.276***	0.276***	0.271***	0.245***	0.242***	0.242***	0.266***
	(0000)	(0000)	(0000)	(0000)	(0000)	(0000)	(0000)	(0000)
$\rho^{\ddagger}_{c,t-2}$	0.231***	0.220***	0.219***	0.216***	0.177***	0.172***	0.172***	0.204***
	(0000)	(0000)	(0000)	(0000)	(0001)	(0001)	(0002)	(0000)
$\rho^{\ddagger}_{c,t-3}$	0.063**	0.063***	0.063**	0.057**	0.041	0.042	0.042	0.045
	(0015)	(0009)	(0010)	(0024)	(0165)	(0159)	(0170)	(0126)
$\rho^{\ddagger}_{c,t-4}$	0.030	0.026	0.025	0.023	0.072**	0.068**	0.069**	0.002
	(0459)	(0541)	(0553)	(0598)	(0021)	(0025)	(0027)	(0962)
$\rho^{\ddagger}_{c,t-5}$	0.062*	0.068**	0.068**	0.074**	−0.014	−0.002	−0.002	0.041
	(0054)	(0034)	(0038)	(0020)	(0543)	(0925)	(0946)	(0279)
Lag (log) GDP, per capita (PPP $)				−0.012*				0.028
				(0053)				(0458)
Constant	−0.002	0.292***	0.302***	0.313***				
	(0942)	(0000)	(0000)	(0000)				
Estimator	OLS	OLS	OLS	OLS	2SLS	2SLS	2SLS	2SLS
Trends	C	CY	CST	CST	C	CY	CST	CST
Observations	3328	3328	3328	3229	2491	2491	2491	2493
Kleibergen-Paap p-value					0.00	0.00	0.00	0.00
Hansen J p-value					0.77	0.39	0.37	0.01

The dependent variable is $\rho^{\ddagger}_{c,t}$, the ratio of the female to male labour share ratio in country c and year t. *(log) GDP per capita* is the natural logarithm of per capita GDP (PPP). Columns $1-4$ report OLS estimates. Columns $5-8$ report IV estimates using the (lag) gravity weighted trade shocks Z_1, agricultural and mineral commodity price shocks, Z_2, Z_3, and IMF/WB interventions Z_4. $C =$ Country fixed effects, $CY =$ Country and year fixed effects, $CSY =$ Country, year fixed effects and country specific linear time trends.

* p < 0.1, ** p < 0.05, *** p < 0.01. Standard errors, clustered by country, in parentheses.

Table 4: Effects of Democratization on Economic Gender Inequality

	(1)	(2)	(3)	(4)	(5)	(6)
Electoral Democracy	0.016***	−0.017***	−0.017***	0.031***	−0.079***	−0.080***
	(0000)	(0002)	(0002)	(0000)	(0000)	(0000)
$\rho^{\ddagger}_{c,t-1}$	0.358***	0.325***	0.325***	0.356***	0.319***	0.319***
	(0000)	(0000)	(0000)	(0000)	(0000)	(0000)
$\rho^{\ddagger}_{c,t-2}$	0.272***	0.250***	0.250***	0.268***	0.243***	0.243***
	(0000)	(0000)	(0000)	(0000)	(0000)	(0000)
$\rho^{\ddagger}_{c,t-3}$	0.091***	0.083***	0.083***	0.091***	0.075***	0.075***
	(0001)	(0000)	(0000)	(0001)	(0001)	(0001)
$\rho^{\ddagger}_{c,t-4}$	0.056*	0.043	0.043	0.057*	0.035	0.035
	(0089)	(0225)	(0233)	(0093)	(0351)	(0363)
$\rho^{\ddagger}_{c,t-5}$	0.076**	0.070**	0.071**	0.077**	0.060*	0.060*
	(0024)	(0030)	(0032)	(0024)	(0067)	(0070)
Estimator	OLS	OLS	OLS	2SLS	2SLS	2SLS
Trends	C	CY	CST	None	CY	CST
Observations	4673	4673	4673	4510	4510	4510
Kleibergen-Paap p-value				0.00	0.00	0.00

The dependent variable is $\rho^{\ddagger}_{c,t}$, the ratio of the female to male labour share ratio in country c and year t. *Electoral Democracy Index* is taken from the V-DEM project Coppedge et al. (2016). It takes values in the interval 0 to 1 with higher values representing a greater degree of democracy. Columns $1-3$ report OLS estimates. Columns $4-6$ report IV estimates using the regional democratization waves Z_5 as instrument as defined in Appendix C. $C =$ Country fixed effects, $CY =$ Country and year fixed effects, $CSY =$ Country, year fixed effects and country specific linear time trends.

* p < 0.1, ** p < 0.05, *** p < 0.01. Standard errors, clustered by country, in parentheses.

Table 5: Effects of Income and Democratization on Economic Gender Inequality

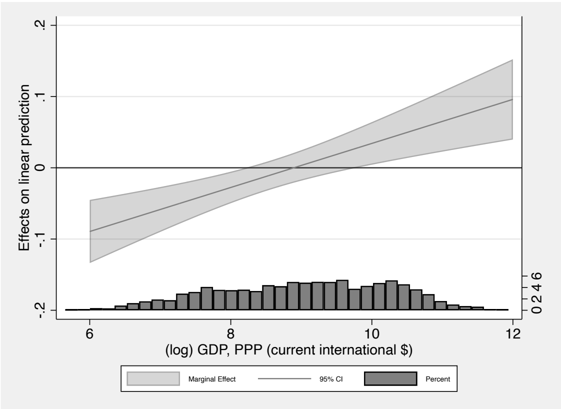
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
Electoral Democracy	−0.011	−0.275***	−0.501**	−0.519***									
	(0.207)	(0.000)	(0.043)	(0.000)									
(log) GDP, per capita (PPP $)	−0.012**	−0.022***	0.017	−0.029***	0.021	0.025	0.026	−0.026	−0.044	−0.022	0.011	0.062	0.056
	(0.022)	(0.000)	(0.634)	(0.004)	(0.546)	(0.460)	(0.437)	(0.465)	(0.239)	(0.542)	(0.749)	(0.124)	(0.162)
$\rho^{\ddagger}_{c,t-1}$	0.275***	0.268***	0.223***		0.227***	0.222***	0.222***	0.211***	0.211***	0.214***	0.219***	0.225***	0.227***
	(0.000)	(0.000)	(0.000)		(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$\rho^{\ddagger}_{c,t-2}$	0.220***	0.215***	0.158***		0.161***	0.157***	0.157***	0.154***	0.152***	0.152***	0.161***	0.157***	0.158***
	(0.000)	(0.000)	(0.006)		(0.002)	(0.004)	(0.004)	(0.006)	(0.006)	(0.006)	(0.004)	(0.003)	(0.003)
$\rho^{\ddagger}_{c,t-3}$	0.063**	0.058**	0.035		0.039	0.034	0.033	0.033	0.032	0.032	0.036	0.030	0.032
	(0.011)	(0.015)	(0.247)		(0.176)	(0.238)	(0.245)	(0.240)	(0.276)	(0.262)	(0.196)	(0.329)	(0.287)
$\rho^{\ddagger}_{c,t-4}$	0.025	0.021	0.062**		0.070**	0.060**	0.057**	0.053*	0.057*	0.055*	0.052*	0.047	0.054*
	(0.560)	(0.628)	(0.039)		(0.018)	(0.037)	(0.045)	(0.060)	(0.054)	(0.056)	(0.062)	(0.103)	(0.083)
$\rho^{\ddagger}_{c,t-5}$	0.068**	0.063*	−0.009		−0.002	−0.011	−0.013	−0.013	−0.012	−0.011	−0.015	−0.020	−0.025
	(0.041)	(0.052)	(0.721)		(0.927)	(0.636)	(0.555)	(0.537)	(0.594)	(0.632)	(0.505)	(0.472)	(0.372)
$\rho^{\ddagger}_{c,t}$				0.174***									
				(0.004)									
Interaction		0.031***	0.059**	0.060***	0.023**	0.059**	0.069**	0.131***	0.133***	0.124**	0.084**	0.091*	0.087*
		(0.000)	(0.029)	(0.000)	(0.037)	(0.021)	(0.025)	(0.003)	(0.004)	(0.017)	(0.016)	(0.084)	(0.061)
Democracy measure by ANRR					−0.186**								
					(0.037)								
Liberal democracy						−0.502**							
						(0.034)							
Egalitarian democracy							−0.573**						
							(0.048)						
Women's Political Empowerment								−1.069***					
								(0.006)					
Women's civil liberties									−1.106**				
									(0.017)				
Women's civil society participation										−1.049**			
										(0.012)			
Women's political participation											−0.715**		
											(0.027)		
Share Women Lower House												−0.703	
												(0.207)	
Share Women Upper House													−0.721
													(0.129)
Estimator	OLS	OLS			2SLS	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS
Trends	CST	CST	CST	CST	CST	CST	CST	CST	CST	CST	CST	CST	CST
Observations	3296	3296	2409	2846	2414	2414	2414	2403	2414	2414	2403	2294	2279

The dependent variable is $\rho^{\ddagger}_{c,t}$. The different indicators of democracy and women empowerment are taken from the V-DEM project Coppedge et al. (2016). Women shares in parliaments data is taken from Inter-Parliamentary Union. Columns $1-4$ report OLS estimates. Columns $5-13$ report IV estimates using the same instruments used in Tables 3 and 4. C = Country fixed effects, CY = Country and year fixed effects, CSY = Country, year fixed effects and country specific linear time trends.

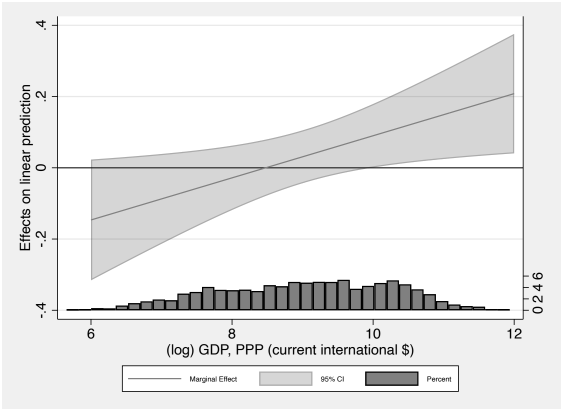
* p < 0.1, ** p < 0.05, *** p < 0.01. Standard errors, clustered by country, in parentheses.

Figure 9: Marginal Effects of Democracy on $\rho^{\ddagger}_{c,t}$ at Different Income levels

(a) OLS Estimates



(b) 2SLS Estimates



(c) 2SLS Estimates

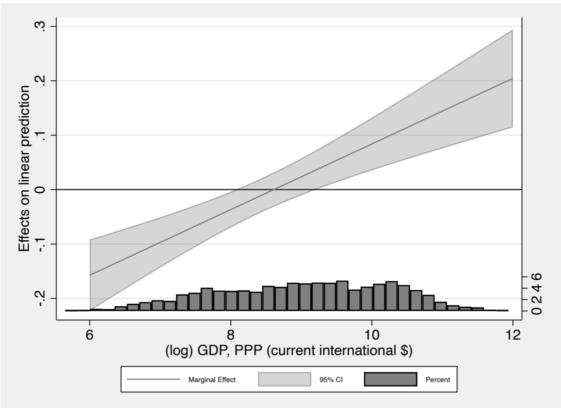


Table 6: Other Potential Determinants of Economic Gender Inequality

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
(log) GDP, per capita (PPP $)	−0012**	−0009	−0011*	−0015***	−0015**	−0014***	−0010*	−0018***
	(0022)	(0141)	(0062)	(0004)	(0011)	(0005)	(0063)	(0002)
Electoral Democracy	−0011	−0011	−0014	−0008	−0011	−0010	−0011	−0020**
	(0207)	(0270)	(0178)	(0350)	(0286)	(0373)	(0210)	(0045)
$\rho^{\ddagger}_{c,t-1}$	0.275***	0.271***	0.275***	0.273***	0.273***	0.270***	0.270***	0.269***
	(0000)	(0000)	(0000)	(0000)	(0000)	(0000)	(0000)	(0000)
$\rho^{\ddagger}_{c,t-2}$	0.220***	0.218***	0.224***	0.216***	0.220***	0.218***	0.217***	0.214***
	(0000)	(0000)	(0000)	(0000)	(0000)	(0000)	(0000)	(0000)
$\rho^{\ddagger}_{c,t-3}$	0.063**	0.060**	0.060**	0.057**	0.056**	0.055**	0.059**	0.050*
	(0011)	(0013)	(0022)	(0023)	(0040)	(0036)	(0014)	(0064)
$\rho^{\ddagger}_{c,t-4}$	0.025	0.022	0.032	0.025	0.019	0.018	0.021	0.013
	(0560)	(0607)	(0476)	(0571)	(0650)	(0680)	(0623)	(0781)
$\rho^{\ddagger}_{c,t-5}$	0.068**	0.065**	0.064*	0.067**	0.076**	0.074**	0.062*	0.070**
	(0041)	(0041)	(0065)	(0042)	(0032)	(0022)	(0058)	(0038)
KOF Globalization		−0000						
		(0234)						
Chinn-Ito			−0002					
			(0579)					
Financial Development				0.065***				
				(0000)				
Ratio Women Industry					−0004			
					(0925)			
Ratio Women Services						0.039**		
						(0019)		
Share Population 15-64							−0167***	
							(0000)	
Ratio Girl Primary Enrollment								−0101***
								(0000)
Estimator	IV	IV	IV	IV	IV	IV	IV	IV
Trends	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	3296	3232	3125	3242	3089	3089	3296	2860

The dependent variable is $\rho^{\ddagger}_{c,t}$, the ratio of the female to male labour share ratio in country c and year t. *KOF Globalization Index* is the overall measure compiled by Dreher et al. (2008). The *Chinn-Ito Index* measures capital account openness and is normalized to take values between 0 and 1 and taken from Chinn and Ito (2006). The *Financial-Reform Index* is from IMF and summarizes financial institutions and market access, efficiency and depth. *Ratio Women Industry* is the ratio of female to male employees in the Industrial Sector. *Ratio Women Services* is the equivalent for the Service Sector. *Share Population* $15-64$ is the percentage of the population aged $15-64$. *Ratio Girl Primary Enrollment* is the ratio of girls to boys enrolled in primary school.

5 Conclusion

This paper has presented a new approach to measuring EGI based on the ratio women's share of national labour income to men's. This approach corresponds precisely to the combination of the concepts of equal pay for equal work and equality of opportunity enshrined in international treaties. The resulting data can be readily compared across time and place, and we are thus able to provide new evidence about EGI varies across countries, its composition, and how it has evolved over time. We find that EGI, despite the progress documented by Goldin (2014), remains substantial at a global level. We present the first estimates of aggregate global EGI and suggest that this is equivalent to around one billion women working for no compensation whatsoever. Moreover, given demographic projections, this number can be expected to rise as population growth is projected to be concentrated on the poorest, and least gender-equal, countries over the next four decades. Our decomposition show that differences in hours worked and labour market participation are more important than differences in wages. Although, wage differences alone are equivalent to 135 million women working without compensation.

Using the labour share ratio approach we construct a panel covering around 120 countries for the period 1991-2016. We use these data to undertake a causal analysis of whether modernization leads to improvements in EGI or whether improvements in the treatment of women in the labour market are driven by a separate process. Our IV estimates support the modernization hypothesis: an increase in GDP per capita from \$1,046 to the bottom of the World Bank rich income category (\$12,696) would be sufficient to raise ρ from approximately 0.4 as in Tunisia, Iran, or Turkey, to close to 0.8 as in Norway, Sweden, or Latvia. At this level of income, and above, democratization also leads to reductions in EGI.

Compliance with Ethical Standards:

Conflict of Interest: The authors declare that they have no conflict of interest.

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A Defining Economic Gender Inequality

As discussed above, EGI is the product of two forms of discrimination. The first is differences in pay for the same value of work. The second is differences in value created due to inequality of opportunity. To fix ideas, we now discuss this more formally.

Consider a population of men and women $i \in \mathcal{I} = \mathcal{F} \cup \mathcal{M}$. Individuals have different types of occupation and some may have no occupation at all. Some are employed by firms in return for wages and other benefits. Others are self-employed, potentially in subsistence agriculture or other non-market activities. A given individual may indeed work in a combination of these activities. Some may undertake no productive activity at all.

In each productive activity an individual's labour creates value-added $v_i \in \mathbb{R}_+$. That is they produce something beyond the (market) cost of the other factors of production such as capital or land.[18] The value added an individual creates depends on their characteristics as well as their occupation. Their individual characteristics such as education, skills, experience, location, etc.) are denoted by the n-vector $\underline{\boldsymbol{\theta}}_{\boldsymbol{i}} \in \mathbb{R}^n$ and those of their occupation they perform (industry, occupation, equipment, hours etc.) by the m-vector $\overline{\boldsymbol{\theta}}_{\boldsymbol{i}} \in \mathbb{R}^m$. Then v_i is given by the scalar valued function of these two vectors:

$$v(\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}}) = v_i \tag{A.1}$$

[18]We ignore the possibility here that, Of course, some activities may have a negative value-added if costs such as environmental degradation are included.

We note that the lack of an i subscript on the value-added function, $v(\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}})$, reflects that this function is assumed to be identical for all individuals reflecting the idea that all relevant differences between individuals or occupations are captured by $\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}$ and $\overline{\boldsymbol{\theta}}_{\boldsymbol{i}}$ respectively.

Each individual receives total labour compensation c_i comprised of a wage $w_i > 0$ and benefits $b_i > 0$. (Note, some of the income of the self-employed, subsistence farmers, and other owners of capital will be due to their capital and not their labour. This is not included in c_i. While this is simple conceptually, it is more involved in practice as discussed in section 2.) b_i includes the costs of non-monetary compensation to an (self-) employer such as health-insurance, training, or food.

A worker's compensation c_i and their value added $v(\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}})$ are linked by their labour share. A worker whose compensation is equal to their value added would have a labour share of 1 – implying no compensation for capital, land, etc. Likewise, a worker who received half of the value-added they create in their role would have a labour share of 0.5. This labour share will vary depending on the technology of production and, to the extent that factor markets are not efficient, bargaining between the worker(s) and employers.

Individual i's labour share is then the ratio of their compensation c_i to their value-added v_i. That is:

$$\lambda_i = \frac{w_i + b_i}{v\left(\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}}\right)} = \frac{c_i}{v\left(\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}}\right)} \tag{A.2}$$

$$c_i = \lambda_i v\left(\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}}\right) \tag{A.3}$$

. Thus, the labour income of two individuals may vary because a) they have different characteristics or do different jobs (i.e. $\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}$ and or $\overline{\boldsymbol{\theta}}_{\boldsymbol{i}}$) are different) or because their compensation for the same-value added is lower. For example, if two lawyers charge their clients the same hourly rate and everything else about their work and benefits is identical, then any difference in their compensation will reflect differences in λ_i since $\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}}$ are identical and hence so is $v\left(\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}}\right)$.

This means that differences in λ reflect differences in pay for work of equal value. Of course, we may expect, even in a society without discrimination, some dispersion in θ due, for example, to frictions in wage-setting over the business-cycle. However, to the extent that the average labour share of women, $\overline{\lambda_F} = \frac{\sum_{i \in F} \lambda_i}{|F|}$, is different to that of men, $\overline{\lambda_M} = \frac{\sum_{i \in M} \lambda_i}{|M|}$ then there are deviations from the principle of equal pay for equal work.

These differences have two forms. The first is pure discrimination: women are sometimes paid less for the same work of equal value in the same job. The second is more subtle: roles that are predominantly filled by women may pay less than jobs creating the same value filled by men. Both of these, as explained immediately below, are captured by the ratio of the average labour share of all women, λ_F, with that of all men λ_M.

To see this, we assume that imperfect competition in labour and product markets means that workers of both genders must bargain over their share of output. The extent to which male workers received more, *ceteris paribus*, reflects differences in the relative bargaining strength of men and women. Our argument is similar to that of Rodrik (1999) who argues that, in part, differences in the labour share across countries reflect differences in the relative bargaining strength of workers rather than cross-national differences in production technologies. Similarly, we now discuss, why in our context, differences in the technology of production can not explain gender differences in factor shares.

The most immediate implication of (B.2) is that since it deals in factor *shares* any differences in output between men and women are allowed for. This in turn implies that any deviations from $\rho = 1$ in a competitive economy, if they do not represent discrimination, reflect differences in the technology of production. But, such an argument is hard to sustain. To see this consider the case where $v(\theta)$ is a simple Cobb-Douglas production function with inputs of labour, capital, and human capital, so $\theta_i = \{H_i, L_i, K_i\}$. Then $v(\theta_i) = A_i L_i^{\alpha_i} H_i^{\beta_i} K_i^{1-\alpha_i-\beta_i}$ such that there are constant returns. Then, the return to labour and the human capital embodied in it is $\lambda_i = \alpha_i + \beta_i$ and $w_i = (\alpha_i + \beta_i)v(\theta_i)$. Defining α_M and α_F as the average return to labour for men and women, and similarly β_F and β_M then we can write the labour share ratio as

$\rho = \frac{\alpha_F + \alpha_M}{\alpha_M + \alpha_F}$. Thus, if lower female labour shares are argued to be due to $\alpha_M + \beta_M > \alpha_F + \beta_F$ then we should expect women to have a correspondingly higher exponent on Capital and thus, in a competitive economy, to be concentrated in capital intensive roles. This is both at odds with casual empiricism and more importantly the micro-econometric evidence, such as that in Arai (2003).

In our data, human capital will be conflated with labour. But, again, any argument that the male labour share is higher because men for some reason make better use of human capital is contradicted by the empirical evidence (see, Pitt et al., 2012) and moreover would imply that capital intensity should be lower. If the argument is more plausibly that men often have higher levels of human capital due to parental decisions and institutional factors then this is adjusted for by our focus on the labour share ratio. This argument can easily be extended to other constant returns production functions.

Thus, an argument that differences in labour shares are argued to reflect differences production technology must now be an argument that the returns to scale in female production are lower than that from male production. Why this might be is not obvious, but given that estimates of the returns to scale in the overall economy are normally close to unity (see, Basu and Fernald, 1997) this implies that women have decreasing returns to scale and men increasing returns. Such a claim not only embodies extremely strong assumptions about differences in production between men and women, but also has odd implications for the economy as a whole. For example, it would imply that in the richest economies men are increasingly comparatively productive compared to women yet this is the opposite of what we find in our data.

A more subtle argument is that it is not that men's and women's production functions are different, but instead that differences in preferences mean that they tend to work in different sectors of the economy and differences in the labour share reflect this. However, were this to be true, it again would require that women to work in more capital intensive sectors. Moreover, we find a similar pattern of EGI across all sectors of the economy but we will also present estimates for the manufacturing sector only to address this concern.

A yet more subtle argument is that the economy lacks factors of production. But that scarce capital is unequally distributed again reflects differences in gender bargaining power, not efficiency. This is borne out by the evidence provided by Udry (1996) who shows that variable inputs are inefficiently allocated in favor of men when they controlled production within agricultural households. In particular, Udry (1996) provides evidence that differences in gender productivity are due to non-Pareto efficient allocations of fertilizer and labour. Finally, we provide an empirical argument that suggests that even when we compare service sectors, there still remain substantial differences in ρ across countries and almost always $\rho < 1$. However, allocation of capital on the basis of gender rather than efficiency is precisely labour market discrimination.

An alternative approach would be to consider a directed search model of the labour market in which there may be gender differences in preferences. For example, if women were assumed to be more risk-averse (on average) then this would imply that the average wage of women might be lower as they would be more likely take lower paying jobs rather than risk unemployment. However, were this to be true, other things equal, observed unemployment for women would be lower. In fact as documented by Azmat et al. (2006), Olivetti and Petrongolo (2008), unemployment rates are normally substantially higher for women.

B Measuring Gender Pay Inequality

Comparison with the Gender Pay Gap

EGI is also due to gender differences in occupational choice as women often disproportionately have occupations that create less value. The reasons include both differences in opportunity and differences in preference. Differences in opportunity vary from the obvious effects of social prohibitions on who can do which jobs, to more subtle requirements such as selection mechanisms that implicitly favor men (see, Goldin, 1990, England et al., 2020). There are also often differences in educational opportunity (see, Altonji and Blank, 1999), access to social-networks Beaman et al. (2018), Blackaby et al. (2005), glass-ceilings (see, Albrecht et al., 2003, Arulampalam et al., 2007, Bertrand, 2018, Duchini et al., 2020) and so forth. Gender differences in expected household production will also impact on hours worked: women also often engage in more (unmeasured) household production (see, Hook, 2010), and this impacts upon their pay and advancement.

Recent evidence also documents gender differences in preferences that may lead to differences in occupational choice. An important recent literature studies how differences in preferences for risk (see, Bertrand, 2011), working hours (see, Goldin, 2014), competition (see, Fershtman and Gneezy, 2001, Niederle and Vesterlund, 2007, Gneezy et al., 2009, Buser et al., 2014); the welfare of others and prestige affect occupational choice and earnings. Other studies consider the role of yet more subtle factors such as additional absenteeism due to the menstrual cycle (see, Ichino and Moretti, 2006) or the role of outside offers (see, Blackaby et al., 2005).

The benchmark in this paper is of no EGI, i.e. in a truly equal society the net impact of the gender differences discussed above would be small. Put differently, the premise of this alternative benchmark is that while gender differences are important in understanding persistent gender differences in otherwise equal settings, particularly in rich equal countries, they are relatively unimportant in understanding differences in EGI between countries or over time. Thus, this paper employs a different benchmark for what constitutes a society without EGI than that which is implicit in the GPG as well as much of the previous literature that uses Oxaca-Blinder decompositions and it is worthwhile to discuss how it differs.

Likewise, as people differ we may expect variation in $\underline{\theta}_i$ and $\overline{\theta}_i$. However, we do not expect systematic differences across groups. In particular while it may be that even in a completely equal society there are some systematic differences between men and women in relevant characteristics or occupations, we expect that these differences will be very small in aggregate. That is, if there are such characteristics there is no reason to believe that they disproportionately lead to higher male rather than female wages.

The implicit assumption in the Gender Pay Gap is that the correct counterfactual is the difference in compensation between equally qualified individuals doing the same job but differing only by gender. This same assumption is implicit in Oxaca-Blinder decompositions into group-differences (not discrimination) and the unexplained component (discrimination). Thus, these approaches, by design, neglect the two other dimensions of EGI; inequality of opportunity and differences in occupational choice.

The complication with including these other dimensions of EGI is that they potentially conflate discrimination with differences in preferences which should be taken as given. Thus, for example, this view would hold that if certain roles requiring, say, more education may be predominantly held by men, and this not imply EGI, if, for example, women rationally choose to invest less in the necessary education because they have different preferences in terms of labour-market attachment.

The alternative view advanced by this paper is that it is important to distinguish the role of genuine differences in preferences from the roles of societal institutions. For example, if a woman rationally chooses less education only because legislation requires that she will be obliged to take maternity leave rather than her partner take it then this represents discrimination not differences in preferences.

Similarly, in this alternative view one would argue that documented differences in preferences for risk taking between men and women can only justify EGI to the extent that such additional risk taking is productive in, as is often argued, the finance sector rather than as an arbitrary selection mechanism.[19] If a, protoypically male desire for prestige or an appetite for risks is rewarded in tournament style contests which sometimes characterise professional advancement then consequential inequality only represents differences in preferences and not descrimination to the extent that it is necessary or efficient for the mechanism employed to select on male characteristics.

[19] Sapienza et al. (2009) show that differences in risk taking are associated with testosterone levels rather than gender per se.

Again, as documented by Goldin (2014) salary differences between men and women are often concentrated , particularly after having had children, on those roles in which salary is a non-linear function of hours worked. The question is then, whether this reflects both the intrinsic nature of these roles and the (average) preferences of women. That is, are these roles indivisible and thus one person performing the role is creates more value-added than two equivalent people working half as much each. In terms of women's preferences it requires that they prefer to earn (much) less and work less rather than being constrained into doing so. If both of these assumptions are met then the salary differences Goldin (2014) identifies are not a source of EGI. On the other hand, it is equally easy to see that such an outcome could reflect something different than the natural consequence of the such preferences and technology. For example, consider roles which reward many hours worked as a signalling divice even if the number hours worked beyond a given amount is only loosely related to output if at all. This would be a system that rewarded presenteeism and if this selected out women then this would be a source of descrimination that is avoidable to the extent that there may be other, better, signalling technologies that do not have the same gendered characteristics.

Of course, none of the above is to argue that all observed differences in occupational choice reflect gendered institutions rather than fundamental differences in preferences over roles. The relevant question is are these differences quantitatively important and in which (if any) aggregate direction should we expect them to be.[20]

We argue that the consistent shrinkage, in many industrialised societies, of roles and occupations considered solely the domain of men or women, such as military service or childcare, suggests that caution is appropriate in ascribing too large a role to preferences. Notably, given that historically in many countries EGI considerably lagged development and technological progress and institutions governing accredition and training in fields such as relgion, medicine, or the law, will have developed without consideration of the consequences of differential impacts upon women. Thus, we see our benchmark of no-difference as reflecting the elimination of these. It will overstate EGI if gender differences in preferences play a large role independent of institutions that mediate them, but the evidence for this is, to the best of our knowledge, scant.

Of course, the discrimination and inequality of opportunity may interact – for example, women's educational choices will be distorted by pay discrimination. These differences will all be captured by $v(\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}})$.

A considerable literature, see for example Blau and Kahn (1992, 2003), Weichselbaumer and Winter-Ebmer (2007), Olivetti and Petrongolo (2008), has studied cross-country differences in EGI through the lens of the GPG. [21] In its simplest form and in our notation, denoting the median woman as $\tilde{F}$ and the median man as $\tilde{M}$, the definiton of pay equality embodied in the GPG requires that:

$$\text{GPG} = \frac{c_{\tilde{M}} - c_{\tilde{F}}}{c_{\tilde{M}}} = \frac{\lambda_{\tilde{M}} v(\underline{\boldsymbol{\theta}}_{\tilde{\boldsymbol{M}}}, \overline{\boldsymbol{\theta}}_{\tilde{\boldsymbol{M}}}) - \lambda_{\tilde{F}} v(\underline{\boldsymbol{\theta}}_{\tilde{\boldsymbol{F}}}, \overline{\boldsymbol{\theta}}_{\tilde{\boldsymbol{F}}})}{\lambda_{\tilde{M}} v(\underline{\boldsymbol{\theta}}_{\tilde{\boldsymbol{M}}}, \overline{\boldsymbol{\theta}}_{\tilde{\boldsymbol{M}}})} = 0 \tag{B.1}$$

Thus, while corresponding to a sensible and intuitive definition of EGI, the GPG conflates gender differences in earnings due to some forms of discrimination, such as education, but not others such as participation, with departures from equal pay for equal work. This makes cross country comparison difficult.

Amongst others Oostendorp (2009a) proposes the (log) occupational wage gap to address these concerns. This measure in the notation above is $\lambda_{\tilde{F}} v(\underline{\boldsymbol{\theta}}_{\tilde{\boldsymbol{F}}}, \overline{\boldsymbol{\theta}}_{\tilde{\boldsymbol{F}}})|Occupation) - \lambda_{\tilde{M}} v(\underline{\boldsymbol{\theta}}_{\tilde{\boldsymbol{M}}}, \overline{\boldsymbol{\theta}}_{\tilde{\boldsymbol{M}}})|Occupation)$ thus it captures both the occupation-specific labour share ratio as well as differences in value added within occupation due to differences in hours, rank, human capital, or other characteristics. Oostendorp argues that as occupations are measured

[20]It is entirely possible for there to be quantitatively important gender differences in average preferences that give rise to substantial differences in the numbers of men and women doing different roles, but for the aggregate consequences of these to be similar for men and women's average compensation.

[21]There is also a prominent related literature which studies pay gaps within countries, such as Mulligan and Rubinstein (2008), Manning and Swaffield (2008) or Black and Spitz-Oener (2010).

relatively precisely this should control for differences in human capital, but this is in contrast to the findings of Goldin (2014) who finds considerable pay inequality within even elite jobs due to differences in hours or child-rearing.[22] Furthermore, as Oostendorp (2009a) notes, this measure also does not control for gender differences in occupational choice. Whilst, in principle one could condition on additional person and job characteristics to further improve the ease of interpretation. However, such a data intensive approach is not normally feasible, and is certainly not for the purpose of this paper. Moreover, such measures would not correspond to total EGI that we focus on.

Instead, if given our assumption that differences in $\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}}$ reflect only discrimination and under our assumption that $v(\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}})$ is the same for men and women, then wages $w_i = v(\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}})\lambda_i$ will only be unequal if there is inequality of opportunity, i.e. differences in $\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}$ and $\overline{\boldsymbol{\theta}}_{\boldsymbol{i}}$, or departures from equal pay for equal work, i.e. differences in λ_i. Thus, we define the absence of EGI as requiring:

$$\rho \equiv \frac{c_F}{c_M} = \frac{v(\underline{\boldsymbol{\theta}}_{\boldsymbol{F}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{F}}))}{v(\underline{\boldsymbol{\theta}}_{\boldsymbol{M}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{M}}))} \frac{\lambda_M}{\lambda_F} = 1 \tag{B.2}$$

The assumption of no gender difference in $v(\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}})$ means that the ratio of the female to male labour share $\zeta = \frac{\lambda_F}{\lambda_M}$ is equivalent to departures from equal pay for equal work, ζ. Where:

$$\zeta \equiv \frac{\lambda_F}{\lambda_M} = \frac{c_F / v(\underline{\boldsymbol{\theta}}_{\boldsymbol{F}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{F}}))}{c_M / v(\underline{\boldsymbol{\theta}}_{\boldsymbol{M}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{M}}))} \tag{B.3}$$

Which means that:

$$\rho \equiv \frac{c_F}{c_M} = \frac{\lambda_F}{\lambda_M} \frac{v(\underline{\boldsymbol{\theta}}_{\boldsymbol{F}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{F}}))}{v(\underline{\boldsymbol{\theta}}_{\boldsymbol{M}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{M}})} = \underbrace{\zeta}_{\text{Equal Pay}} \underbrace{v_F / v_M}_{\text{Non-Discrimination}} \tag{B.4}$$

As discussed above, much of the literature analysing pay gaps within countries and some cross-country work such as Oostendorp (2009b), have attempted to condition on sector, occupation or other characteristics. One way to interpret ζ is as an idealised GPG in which all relevant characteristics have been controlled for. Indeed, were there no discrimination such that $v_F / v_M = 1$, then the GPG and departures from equal pay for equal work would be quivalent, specifically $GPG = \zeta - 1$.[23] Thus, our measure of EGI, the labour share ratio, is equivalent to an 'ideal' GPG calculated conditioning on all of the elements of $\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}}$, ζ multiplied by differences in value added due to discrimination.

Our parameterization has the important advantage for our purpose that it uses consistent data, and abstracts from $v(\underline{\boldsymbol{\theta}}, \overline{\boldsymbol{\theta}})$, $\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}}$, and λ_i. This eliminates the increased difficulties in interpretation when trying to make comparisons over time or place as both $v(\underline{\boldsymbol{\theta}}, \overline{\boldsymbol{\theta}})$, the technology of production, and the distribution of $\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}}$ will vary across time and place. This means that interpretation of changes in the wage gap will now conflate changes in $\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}}$, changes in λ, and changes in $v(\underline{\boldsymbol{\theta}}, \overline{\boldsymbol{\theta}})$. This makes clear the challenges faced by previous studies that have, for example, attempted to conduct meta-analyses of the GPG across countries (see, Weichselbaumer and Winter-Ebmer, 2007, Schober and Winter-Ebmer, 2011, Seguino, 2011). It may well be that there are particular applications where conditioning only on some elements of $\underline{\boldsymbol{\theta}}_{\boldsymbol{i}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{i}}$ provides a useful and interpretable quantitity but in general they will make comparison more difficult across time and place and so we do not pursue this issue here.

[22] Other partitions are sometimes used, for example, Wolszczak-Derlacz (2013) focus on the sectoral wage gap in their analysis of trade liberalization.

[23] In general, when $v_F / v_M \neq 1$:

$$GPG - 1 = \frac{\lambda_F v(\underline{\boldsymbol{\theta}}_{\boldsymbol{F}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{F}}))}{\lambda_M v(\underline{\boldsymbol{\theta}}_{\boldsymbol{M}}, \overline{\boldsymbol{\theta}}_{\boldsymbol{M}}))}$$

.

C Construction of Instrumental Variables

The analysis of whether economic growth and democratization lead to improvements in gender equality, and how rapidly, in Section 4 employs four instrumental variables for the GDP per capita and an instrument for the democracy. The first is a gravity-weighted trade shock measure. The second and third measure terms of trade shocks via changes in the prices of commodity imports and exports for agricultural and mineral commodities respectively. The final instrument proxies for financial crises using IMF or World-Bank crisis-interventions. The instrument for the democracy indicator is based on political science literature that democratization comes in regional waves Acemoglu et al. (2019). We now outline the construction of these variables in turn.

Gravity

We estimate a standard Trade-Gravity model of the form:

$$T_{c,c,'t} = \alpha_0 Y_{c,t}^{\alpha_1} Y_{c',t}^{\alpha_2} \aleph_{c,c',t}^{\alpha_3} e^{\sigma_c d_c + \sigma_{c'} d_{c'}} \tag{C.1}$$

Where $Y_{c,t}$ and $Y_{c',t}$ are the GDPs of countries c and c' in year t. $\aleph_{c,c',t}$ is a vector containing measures of the 'distance', broadly conceived, between c and c' in year t. In our case this includes whether the countries are contiguous, share a common language, colonial history, currently colonial relationship, common legal system, a common currency, common religion, are members of the same regional trade agreement, and whether the origin or destination country are members of GATT, and their respective GDP per capita. d_c and $d_{c'}$ are fixed-effects for the origin and destination countries respectively. These capture other, unmeasured, country characteristics that may cause them to export a particularly large or small amount.

Using the data used by Head et al. (2010) we estimate (C.1) using the Poisson pseudo Maximum Likelihood estimator proposed by Silva and Tenreyro (2006). We then obtain predicted flows for each pair of countries for each year. Our instrument is then:

$$Z_{c,t}^1 = \sum_{c'} \widehat{T_{c,c',t}} \times \Delta Y_{c',t} \tag{C.2}$$

Commodities

Our commodity price shock instruments, follow the approach of Deaton and Miller (1996), are given by the product of changes in the global price for each commodity in a given year multiplied by the share of that commodity in a country's trade in a fixed year. By fixing a year, we are able to rule out changes in the composition of the economy in response to price shocks. We use the year 2000 as our fixed year.

$$Z_{c,t}^{2,3} = \sum_{cm.} \Delta G_{t,cm.}^p \times X_{c,2000,cm.}, \tag{C.3}$$

where the summation is over commodities $cm.$ within agriculture and mineral sectors respectively, G^p denotes the global price of the respective commodity and X denotes trade share of the commodity. The data on global commodity prices are taken from IMF Primary Commodity Prices. In case price data was not available from IMF, it was then taken from the World Bank Commodity Price Data (The Pink Sheet). Data on commodities share in the country trade is taken from UN Comtrade. Commodities included within agriculture price index are; sugar, tea, barley, coffee, cotton, wheat, wool, olive oil, rice and rubber. Mineral price index includes petroleum, copper, gasoline, iron, natural gas, silver, uranium, coal, aluminium and zinc.

Crises

Our crisis instrument Z_4, is based on the data of Boockmann and Dreher (2003b) and Dreher (2006b), and is defined as the total number of World Bank projects and IMF Arrangements agreed or in effect in a particular year. The arrangements used in the main analysis include IMF Standby Arrangement agreed, IMF Standby Arrangement in effect for at least 5 months in a year, IMF Structural Adjustment Facility Arrangement agreed, IMF Structural Adjustment Facility Arrangement in effect for at least 5 months in a year, IMF Poverty Reduction and Growth Facility Arrangement agreed, and IMF Poverty Reduction and Growth Facility Arrangement in effect for at least 5 months in a particular year.

Regional Democratization Waves

The political science literature reports that transition to democracy mostly takes place in regional waves (for example, the Arab Spring). Based on this idea, Acemoglu et al. (2019) come up with instrumental variables that focus on regional waves of democratization for countries with common political histories during 1960. More formally, a regional democratization waves instrument is constructed as follow. For each country c, $D_{c,t0}$ denotes whether the country was a democracy or non-democracy in 1960 (or, depending on the country entry to the sample, at the start of the sample for that country) , and R_c denotes the geographic region of country c; seven regions as defined in Acemoglu et al. (2019). Then the countries are distributed into groups as:

$$I_c = \{c' : c' \neq c, R'_c = R_c, D_{c',t0} = D_{c,t0}\}. \tag{C.4}$$

That is, two countries can be in the same group only if they are within the same region and share similar political history at the start of the sample or in 1960. Using these grouping, the instrument is defined as:

$$Z^5_{c,t} = \frac{1}{I_c} \sum_{c' \in I_c} D_{c',t}. \tag{C.5}$$

This means that for country c, the democracy indicator $D_{c,t}$ is instrumented by the mean democracy within the country's c group, leaving out the country c observation. Note that, depending on the choice of initial year democracy $D_{c,t0}$, different instruments can be constructed. The current study relies on the four choices of $D_{c,t0}$ as in Acemoglu et al. (2019) and update the data upto the year 2016.

D Additional Results

Figure D.1: Scatter Plot of Female and Male labour Share in 2014 (Unadjusted Data)

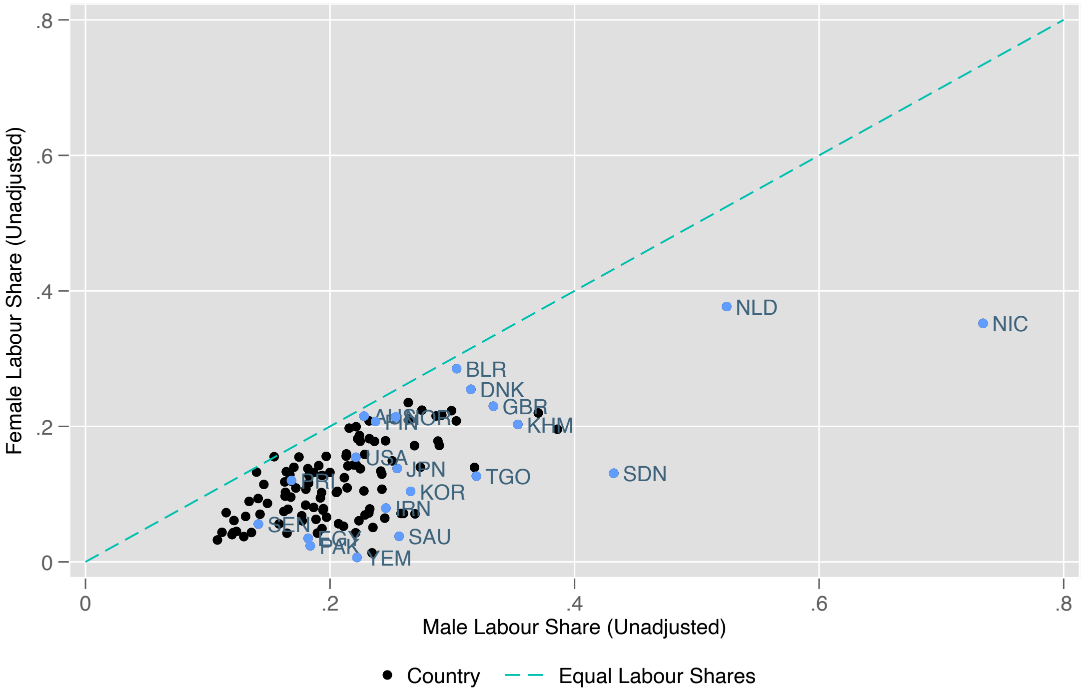

Figure D.2: Cross-Country Distribution of $\rho^{\ddagger}$ by Sector

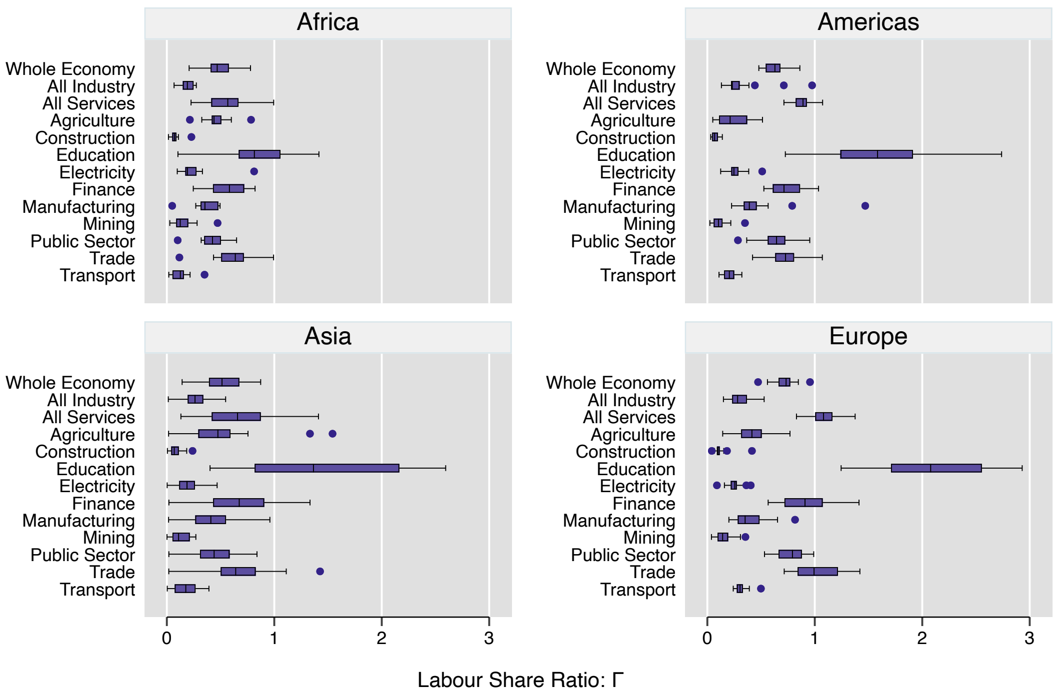

Figure D.3: Cross-Country Distribution of $\rho^{\ddagger}$ by Sector

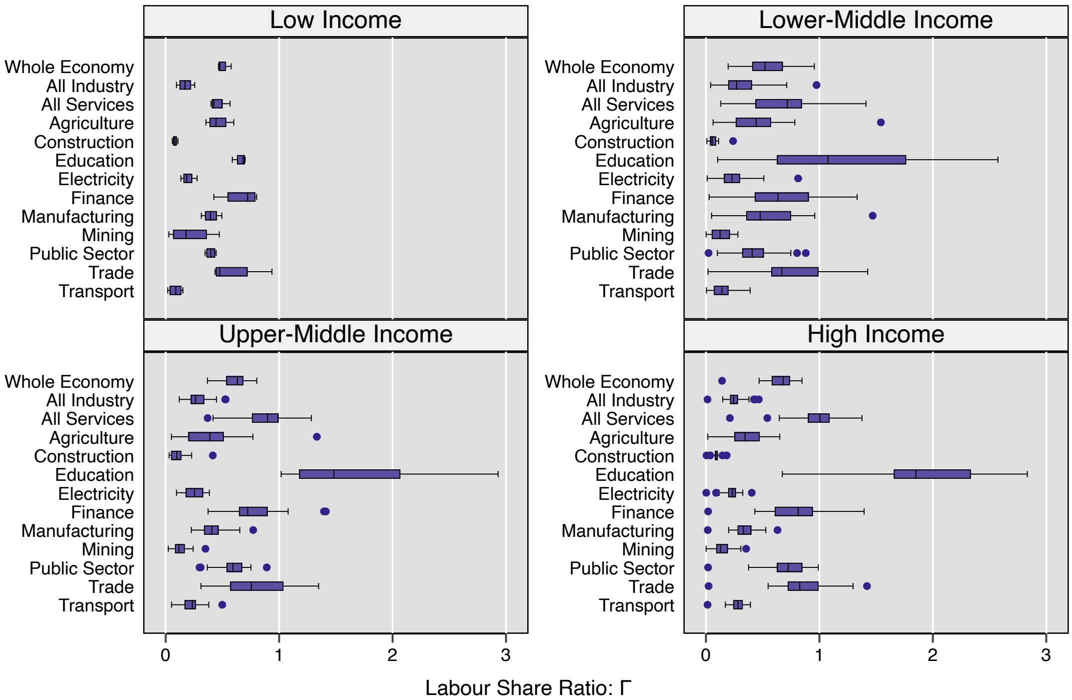

Table D.1: Effects of Income and Democratization on Economic Gender Inequality: Five Year Lag

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Electoral Democracy	−0.973***									
	(0.003)									
(log) GDP, per capita (PPP $)	0.023	0.024	0.040	0.040	0.019	−0.083	−0.003	0.079	0.082	0.086
	(0.723)	(0.707)	(0.535)	(0.511)	(0.796)	(0.214)	(0.968)	(0.292)	(0.439)	(0.322)
Interaction	0.121***	0.060***	0.112***	0.127***	0.145	0.218***	0.153	0.121	0.159*	0.127*
	(0.002)	(0.001)	(0.003)	(0.003)	(0.104)	(0.003)	(0.122)	(0.134)	(0.063)	(0.099)
Democracy measure by ANRR		−0.483***								
		(0.001)								
Liberal democracy			−0.883***							
			(0.008)							
Egalitarian democracy				−0.995**						
				(0.011)						
Women's Political Empowerment					−1.444**					
					(0.040)					
Women's civil liberties						−1.663***				
						(0.006)				
Women's civil society participation							−1.412*			
							(0.052)			
Women's political participation								−1.217*		
								(0.085)		
Share Women Lower House									−1.894**	
									(0.045)	
Share Women Upper House										−0.789
										(0.357)
Estimator	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS
Trends	CST	CST	CST	CST	CST	CST	CST	CST	CST	CST
Observations	2093	2093	2093	2093	2075	2093	2093	2075	1979	1964

The dependent variable is $\rho^{\ddagger}_{c,t+5}$, the ratio of the female to male labour share ratio in country c and year t. *(log) GDP per capita* is the natural logarithm of per capita GDP (PPP). Democracy and women empowerment vairables comes from the V-DEM project Coppedge et al. (2016). Women representation in parliament data comes from Inter-Parliamentary Union. C = Country fixed effects, CY = Country and year fixed effects, CSY = Country, year fixed effects and country specific linear time trends.

* p < 0.1, ** p < 0.05, *** p < 0.01. Standard errors, clustered by country, in parentheses.

Table D.2: Other Potential Determinants of Economic Gender Inequality: Five Year Lag

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\rho^{\ddagger}_{c,t}$	0.198***	0.185***	0.199***	0.181***	0.180***	0.176***	0.181***	0.190**
	(0001)	(0002)	(0001)	(0002)	(0000)	(0000)	(0001)	(0015)
Electoral Democracy	−0017	−0013	−0022	−0004	−0015	−0014	−0019	−0016
	(0271)	(0451)	(0258)	(0782)	(0376)	(0399)	(0233)	(0342)
(log) GDP, per capita (PPP $)	−0010	−0004	−0010	−0017	−0005	−0007	−0007	−0020*
	(0311)	(0714)	(0419)	(0108)	(0650)	(0466)	(0474)	(0077)
KOF Globalization		−0001						
		(0138)						
Chinn-Ito			0.003					
			(0683)					
Financial Development				0.142***				
				(0000)				
Ratio Women Industry					0.045			
					(0149)			
Ratio Women Services						0.070***		
						(0009)		
Share Population 15-64							−0328***	
							(0003)	
Ratio Girl Primary Enrollment								−0203***
								(0000)
Estimator	IV	IV	IV	IV	IV	IV	IV	IV
Trends	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	2846	2789	2668	2796	2643	2643	2846	2450

The dependent variable is $\rho^{\ddagger}_{c,t+5}$, the ratio of the female to male labour share ratio in country c and year t. *KOF Globalization Index* is the overall measure compiled by Dreher et al. (2008). The *Chinn-Ito Index* measures capital account openness and is normalized to take values between 0 and 1 and taken from Chinn and Ito (2006). The *Financial-Reform Index* is from IMF. *Ratio Women Industry* is the ratio of female to male employees in the Industrial Sector. *Ratio Women Services* is the equivalent for the Service Sector. *Share Population* $15-64$ is the percentage of the population aged $15-64$. *Ratio Girl Primary Enrollment* is the ratio of girls to boys enrolled in primary school. All other details as for Tables 5 and 6.

Table D.3: Effects of Income and Democratization on labour Market Gender Inequality, Assuming Different Values of ϕ

	$\phi = 0$	$\phi = 0.25$	$\phi = 0.5$	$\phi = 0.75$	$\phi = 1$	$\phi = 0$	$\phi = 0.25$	$\phi = 0.5$	$\phi = 0.75$	$\phi = 1$
(log) GDP, per capita (PPP $)	0.071***	0.055***	0.039***	0.022***	0.006	0.075**	0.066**	0.057**	0.047	0.038
	(0.000)	(0.000)	(0.000)	(0.000)	(0.293)	(0.015)	(0.019)	(0.041)	(0.107)	(0.250)
Estimator										
Trends	No	No	No	No	No	Yes	Yes	Yes	Yes	Yes
Observations	2546	2546	2546	2546	2546	2546	2546	2546	2546	2546
Kleibergen-Paap p-value	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Hansen J p-value	0.30	0.45	0.61	0.70	0.73	0.01	0.06	0.23	0.48	0.61

Each column reports a different assumed value of the AR(1) coefficient ϕ. Columns 1-5 do not include country specific trends. Columns 6-10 do. All other details as for Tables 5 and 6.