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Child Gender, Egalitarian Attitudes, and Economic Inequalities

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Abstract

We study the effects of gender composition of children on the sociopolitical attitudes of their parents. Having daughters significantly increases parental support for gender-egalitarian viewpoints, a result that holds across a range of different indicators. Since male/female birth outcomes are plausibly exogenous, the estimates are likely to be causal, and pass a number of diagnostics related to identification. The results are stronger for fathers, and coincide with subtle shifts in personality as characterized by the Big Five traits. We also show that attitudes to gender are strongly associated with other egalitarian views, and, when aggregated to the national level, are correlated with markers of income inequality and political representation. As redistributive policy in democratic countries is dependent upon public attitudes, the results have implications for addressing economic disparities.

JEL Classification: D63, D91, J62

Key Words: Discrimination, Gender Inequality, Income Inequality, Social Attitudes

1 Introduction

There is an implicit puzzle underpinning the current policy debate around socioeconomic inequality. On one hand, there is a widespread sense of agreement that highly disparate outcomes are undesirable, especially with respect to race, class, and gender.[1] On the other, inequalities such

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[1] The view that inequality represents a social ill is so widespread amongst economists that Clark and D'Ambrosio (2015, p1148) describe it as *"almost taken as an axiom"*. Public opinion is similarly strong - e.g. see the survey in Jenkins (2019). While distaste for inequality appears to be very general, some forms of seem to elicit stronger reactions than others. For example, inequalities due to factors beyond personal control (i.e. *Inequalities of Opportunity* (Roemer, 1998)) are usually regarded as especially socially corrosive. Negative associations between wellbeing and inequality show up regularly in empirical works (which regress life-satisfaction scores against inequality metrics such as the Gini), and experimental research, which elicit preferences over more and less egalitarian distributions of income (see Clark and D'Ambrosio, 2015).

as these exist in all developed countries, are persistent over time, and appear across a variety of variables, including earnings, education and health. Since policymakers have access to a number of powerful redistributive tools,[2] this pervasiveness raises an important question. Simply put, why is mitigating inequality so hard?

One plausible answer is that public attitudes may be fundamental in both generating and maintaining unequal outcomes. Discrimination based on immutable traits such as race or gender can generate fissures that widen over the life-cycle, resulting in large differentials in economic wellbeing (Lang and Spitzer, 2020). And policies designed to narrow these gaps (such as progressive taxation, labor market regulation, or anti-discrimination legislation) are constrained in part by political feasibility, as set by the general public (Pecoraro, 2017; Weinzierl, 2017). Since social attitudes are often relatively static, ingrained economic disparities may follow directly as a natural consequence.

Our goal in this paper is to study these social attitudes. We focus on one specific form - perspectives on gender inequality, and aim to determine their malleability with respect to an individual's external environment. Ordinarily, teasing out causal estimates in this context is difficult, because sociopolitical views are determined by an array of factors, many of which are both unobservable, and linked to a person's broader circumstances.[3] However, causal estimates can still be obtained under quasi-experimental conditions (Angrist and Pischke, 2011), where a randomly assigned treatment affects an individual's environment and their social attitudes adjust in response.

We present such analysis for parents, using the birth composition of their children as identifying variation. The central hypothesis is that daughters (relative to sons) influence familial conditions in ways that promote broadly *gender-egalitarian* perspectives. Using high-quality Australian panel data we show this is indeed the case, with parents of daughters expressing stronger gender-egalitarian sentiments *ceteris paribus* than those with sons. The results are particularly striking for fathers, older individuals, and persons with lower educational attainments. However significant estimates can be found for mothers and younger individuals as well, indicating that raising daughters predicts progressive attitudinal shifts across the population as a whole.

Understanding the ways that these social preferences are formed is important, as they can help to explain cross-national patterns in economic disparities. In the second part of the paper, we present evidence to this effect in two successive stages. Initially, we take two semi-global data sets on attitudes and values,[4] and show that beliefs related to gender are highly correlated with

[2] A detailed review of policy options for addressing income inequalities is provided by Atkinson (2015). These include minimum wages, capital endowments to be paid at adulthood, highly progressive personal income taxes, earned income tax credits, capital taxation and participation incomes.

[3] These factors include childhood experiences (Jennings *et al.*, 2009), social networks (Harmon *et al.*, 2019) and media exposure (Gavin., 2018), alongside deeper determinants such as genetics (Hatemi *et al.*, 2014).

[4] The World Values Survey and the European Values Survey.

other egalitarian views. This is consistent with the idea that egalitarian preferences represent a fundamental (and potentially partially innate) characteristic of the human psyche,[5] informing responses on a range of social and political issues. Then, when aggregated at the national level, we show that countries where egalitarian beliefs flourish are more equal on objective markers of economic disparity, including income inequality and female representation in politics. While these are only associations, we argue that causal chains whereby attitudes affect outcome distributions (both through influencing individual behavior, and by shaping policy) are an important part of this story.

Our work adds to an emerging body of research on birth composition and social attitudes/behavior. Beginning in the early 1990s, a series of papers by Warner (1991), Warner and Steel (1999), and Borrell-Porta *et al.* (2018) build an increasingly strong case that parenting girls increases gender-egalitarian sentiment.[6] Further, extensions of this research demonstrates that these attitudes feed through to affect political decision-making. For example, Washington (2008) links birth composition with voting patterns in United States congressmen, while Oswald and Powdthavee (2010) broaden this result to cover left-wing voting in general elections in the United Kingdom and Europe.

We contribute to this literature in two ways. Firstly, we further generalize these results, showing that links between birth outcomes and parental views hold (i), in a new country, and (ii) across an increased (but inexhaustive) range of attitudinal indicators. A small theoretical model is used to make sense of these findings, and explain why the results are stronger for men - a puzzle in the extant research. We also show that these effects run deeper than merely affecting sociopolitical views, and occur alongside other subtle shifts in personality. Counter-intuitively, parents with a greater fraction of daughters become slightly *less* conscientious relative to parents with more sons, a finding we attribute to compensatory behavior within households. Since conscientiousness is a driver of a variety of social behaviors (Jackson *et al.*, 2010), our results indicate that birth composition could have small but wide-ranging effects on a host of socioeconomic outcomes.

Secondly, (and more importantly) we go further than existing work in examining causality in these models. While child gender is plausibly exogenous, there are still some substantial threats to identification. Parents can affect the gender composition of their families deliberately, (e.g. through reproductive timing), or naively (e.g. through health behaviors that affect sex ratios). Or they may do so by basing future reproductive decisions on the mix of children they already

[5]Evidence to this effect comes from lab experiments in humans, and ethological studies of chimpanzees. For example, results from laboratory experiments suggests that a desire to mitigate inequalities is a fundamental factor (distinct from revenge) in motivating punishment decisions in ultimatum-style games (Bone and Raihani, 2015). Preferences for equal outcomes have also been seen in other primates (Leimgruber *et al.*, 2016), which the authors suggest is an evolved trait designed to facilitate cooperation.

[6]We note that this type of finding is not ubiquitous, for instance Healy and Malhotra (2013) find that men with sisters are more likely to be politically conservative than men with brothers. Whether this is due to differing effects for parents and siblings, or some other phenomenon, remains unclear.

have. Since these actions are plausibly related to pre-existing social attitudes,[7] they can open up non-causal correlations between egalitarian preferences and familial composition. However, using a variety of new diagnostic methods (Altonji *et al.*, 2005; Nunn and Wantchekon, 2011; Pei *et al.*, 2019; Oster, 2019) we study these potential threats, and show that our results are remarkably stable throughout. Biases that plague observational studies arising from unobserved heterogeneity or endogenous selection are likely to have only negligible impacts. As such, there is strong evidence that causality flows from birth outcomes to the sociopolitical views of parents.

We also study the mechanisms through which this causal flow operates, contrasting an *identity* (or *incentive*) effect with an *exposure* effect. The former occurs when parents take on their children's identities when calibrating their own views (Oswald and Powdthavee, 2010). Since girls are more likely to experience discrimination (Charles *et al.*, 2018), be disadvantaged in labor markets (Goldin, 2014), and face social constraints on professional achievement (Azmat *et al.*, 2020), we might expect parents to actively counter these views. Conversely, an exposure effect occurs when individuals encounter and engage with differing political views. Survey data show that women have greater preferences for redistribution (Alesina and La Ferrara, 2005; Tóth and Keller, 2011), and their wellbeing is more negatively correlated with perceptions of inequality than for men (Clark and D'Ambrosio, 2015).[8] Having more female children would therefore affect the composition of opinions shared within families, shifting the viewpoints of other members.

By examining when differences in parental opinions begin to emerge, we can gain some idea as to which effect accounts for our results. Since identity effects can emerge as soon as the child's gender is known (potentially even before birth), while exposure effects take time, patterns of gradual change are more indicative of the latter. Our estimates show that paternal preferences only diverge when their children reach early adulthood, coinciding with the onset of their own political maturity (Chan and Clayton, 2006). Therefore, it appears that attitudes to gender-inequality may spread through social interactions between adult children and their parents, rather than from parents taking on the identities of their children.[9]

The rest of the paper is structured as follows. Section 2 introduces the data sets, and Section 3 presents our econometric estimates. Section 4 provides some diagnostics bolstering evidence of

[7]For example, parental age has been linked with sex ratios (Jacobsen *et al.*, 1999; Mathews and Hamilton, 2005) and also political beliefs (Pew Research, 2018), such that older parents are more likely to have sons, *and* be more conservative. As we discuss later, maternal smoking and drinking increase the chances of female births (as male fetuses are more prone to miscarriage) and are correlated with politically salient factors like social class (Graham, 2012). Parents may also explicitly influence their child's gender via reproductive timing (James, 1987), medical intervention in IVF procedures, or gender-selective abortions (Kippen *et al.*, 2011; Edvardsson *et al.*, 2018).

[8]Similarly, women are also more positive towards welfare states (Goossen, 2020) and anti-discrimination legislation (Strolovitch, 1998).

[9]While our analysis shows that attitudinal shifts appear later, neurological scans provide evidence in support of more immediate change. Before-and-after neuroimaging from new parents (taken several months apart) reveal that both mothers and fathers see increased grey-matter volume occurring in areas of the brain associated with reward processing, emotion and hormone control (the striatum, amygdala and hypothalamus). See Kim *et al.* (2014) and references therein.

causality. Section 5 then shows the relationships between (i) various progressive social attitudes, and (ii) the links between these attitudes and economic disparities in democratic countries. Section 6 concludes, and supplemental material is presented in the appendix.

2 Data

HILDA Data

Our data are drawn from several different sources. Our main dataset is the Household Income and Labour Dynamics in Australia (HILDA) survey - an approximately nationally representative panel similar in scope and structure to the US PSID or German SOEP. HILDA began in 2001 with annual waves covering almost 20,000 individuals and 7,000 families, and is occasionally topped up to mitigate attrition. We use the 18th release which includes data sourced until 2018. The panel contains extensive questions on social attitudes and includes additional psychometric indicators related to personality.

Our key dependent variables are six statements on gender inequality. These take the form of Likert-type responses indicating agreement/disagreement with a given question using seven-point ordinal scales.[10] The variables appear only intermittently (approximately every four years) in our data, with four waves completed so far. The six questions are outlined below, and distributional plots of each are given in the appendix. Where appropriate we invert the scales, such that higher values always imply a greater preference for gender-neutral outcomes.

- *V1. On the whole, men make better political leaders than women.*
- *V2. It is not good for a relationship if the woman earns more than the man.*
- *V3. It is better for everyone involved if the man earns the money and the woman takes care of the home and children.*
- *V4. Mothers who don't really need the money shouldn't work.*
- *V5. If both partners in a couple work, they should share equally in the housework and care of children.*
- *V6. Children do just as well if the mother earns the money and the father cares for the home and the children.*

[10] Strongly agree, agree, somewhat agree, neutral, somewhat disagree, disagree, strongly disagree.

Controls

Since the attitudes presented above are likely to be related to individuals' economic and demographic backgrounds, we source variables related to these as exogenous controls. Two distinct sets are used throughout. The first set is a collection of background markers measured *before* parents have had their first child. By ensuring that our controls are recorded pre-treatment, we avoid identification problems where our covariate of interest affects other variables in the model,[11] which is desirable when fitting regression equations with causal interpretations (Angrist and Pischke, 2011). Employing this set allows us to account for cultural, demographic and socioeconomic factors potentially linked to perceived gender roles. We do this using parental age, gender, place of birth, and a number of markers of family background such as maternal and paternal educational attainments.[12] Characteristics of early family life are also obtained, such as whether an individual's parents were together, if their father was gainfully employed, and whether they grew up with other children.

Our second set of controls adds in standard contemporaneous variables, and is intended to be descriptive in nature. Here we include current income (log of real household equivalent[13]), educational attainment dummies {less than high school; high school; diploma; university degree; postgraduate}, area of residence {major cities; inner and outer regional; remote} and marital status {single; married; de facto; separated; divorced; never married}. Both control vectors are outlined below in Table 1.

Additional Data Sets

In the second part of the paper, we use data from the World Values Survey (WVS), European Values Survey (EVS) and World Bank to perform some cross-national comparisons. The WVS and EVS are repeated cross-sectional surveys on a wide variety of social attitudes jointly covering over 100 countries. Again, the surveys are designed to be approximately representative of within-country distributions such that national summary statistics can be obtained. The two data sets also share a common structure such that they can be combined to increase coverage.[14] Data from these surveys is then merged with data on income inequality (Gini coefficients) and female representation in politics (the proportion of women in elected office) sourced from the World Bank.

[11] For example it is possible that having girls may affect parental behaviors, such as labor supply decisions or marital stability. Including these variables as controls offers an additional path through which the treatment affects the dependent variable, resulting in biased estimates on child gender.

[12] The maternal and paternal variables refer to the mothers and fathers of the parents whose views we are examining.

[13] We account for economies of scale within households using the square-root equivalence scale, and drop zero and negative incomes from the analysis. Since the income variable is measures post-tax there are very few non-positive values.

[14] We note that the WVS asks a wider range of questions although there is substantial overlap.

Since these latter variables appear intermittently in the data and are sometimes volatile, we use averages over 10 years to obtain country-level summaries.

Descriptive Statistics - HILDA

Below we present descriptive statistics from our extract from HILDA, while information on WVS and EVS data is provided in Appendix A2. Dependent variables and pre-treatment controls appear in the left columns, and our contemporaneous controls are on the right (note that they partially overlap). Since decisions to have children are plausibly related to social values, we only look at families where there is at least one child, but start with the largest possible sample size satisfying this requirement. Children are then identified by matching via HILDA's cross-sample identifiers, which leaves ambiguous cases (such as with step-parents or non-biological parents) in the hands of the respondents.[15] To avoid heterogeneity associated with very young parents (>18) and the elderly (<75), we then drop these values, but note that including them has little impact upon the analysis. Our sample has a greater fraction of mothers (55.3%) than fathers (44.7%) (as there are more female single parents), and the average age is a little over 42. Most of our parents were born in Australia (76.3%), to English-speaking parents (90.8%), and only a small fraction (0.2%) identify as Indigenous or Torres Strait Islander. The average number of sons is a little higher than daughters (0.996 vs 0.935), resulting in almost two (1.931) children per parent.

[15] These are the parents/guardians in children under 15, and the children themselves after this age. Note that deceased children will drop out of our sample, however this should not affect our estimates as it is unlikely that child mortality is correlated with parental attitudes.

Table 1: Descriptive Statistics - HILDA Extract

Dep/Pre-Treatment	n	$\bar{x}$	$\hat{\sigma}$	Pre/Contemporaneous	n	$\bar{x}$	$\hat{\sigma}$
# Daughters	106,615	0.935	0.832	Age	106,615	42.83	10.73
# Sons	106,615	0.996	0.857	Age Squared	106,615	1949	963.7
# Children	106,615	1.931	0.936	Aboriginal or TS Is.	106,615	0.002	0.043
V1 Men Better Political Leaders	19,267	5.401	1.755	Married	100,222	0.713	0.452
V2 Better if Man Earns More	19,279	5.701	1.523	De Facto	100,222	0.139	0.346
V3 Man Should Be Breadwinner	24,492	4.759	1.869	Separated	100,222	0.033	0.178
V4 Mothers Shouldn't Work	24,505	4.608	1.883	Divorced	100,222	0.052	0.222
V5 Housework Should Be Shared	24,513	5.977	1.261	Never Married	100,222	0.016	0.125
V6 Ok if Parents Reverse Roles	24,475	5.346	1.528	Male	106,615	0.447	0.497
Age	106,615	42.83	10.73	Female	106,615	0.553	0.497
Age Squared	106,615	1949	963.7	Major City	106,608	0.631	0.482
Aboriginal or TS Is.	106,615	0.002	0.043	Inner Regional	106,608	0.245	0.430
Male	106,615	0.447	0.497	Outer Regional	106,608	0.106	0.308
Female	106,615	0.553	0.497	Remote	106,608	0.014	0.119
Born Australia	100,212	0.763	0.426	Very Remote	106,608	0.003	0.057
Born Non-Aust Anglophonic	100,212	0.092	0.289	Born Australia	100,212	0.763	0.426
Born Non-Anglophonic	100,212	0.145	0.353	Born English Speaking	100,212	0.092	0.289
Father Employed	93,374	0.865	0.342	Born Non-Eng Speaking	100,212	0.145	0.353
Raised By Both Parents	100,221	0.794	0.404	Employed	100,239	0.748	0.434
Had Siblings	100,042	0.966	0.182	Unemployed	100,239	0.029	0.169
Mother High School	106,615	0.222	0.415	Not in Labor Force	100,239	0.222	0.416
Father High School	106,615	0.226	0.419	Log Eq HH Income	106,391	10.86	0.617
Mother Non-U Tertiary	106,615	0.142	0.350	Post-Graduate Degree	100,239	0.118	0.323
Father Non-U Tertiary	106,615	0.177	0.382	Bachelor/Hons Degree	100,239	0.157	0.363
Mother University Ed	106,615	0.075	0.264	Diploma	100,239	0.345	0.475
Father University Ed	106,615	0.112	0.315	High School	100,239	0.127	0.333
				Less than High School	100,239	0.253	0.435

Note: The table presents sample sizes, means and standard deviations for the HILDA subsample. Contemporaneous covariates are given in the left panel and pre-treatment covariates are given in the right panel. Calculations are performed on a pooled sample from 2001-2018 with 106, 615 observations. Cardinality is assumed for the calculation of means and standard deviations of ordinal variables. The data set is trimmed to contain individuals with at least one child.

Histograms of the {1-7} attitudinal variables are presented in the Appendix. In all cases the variables are heavily left-skew, indicating that highly *in*egalitarian views are relatively rare. Modal responses are either six, or the maximal value of seven, and in most cases, these two outcomes account for more than 50% of total responses. Analogously, scores of 1 (indicating strong disagreement with gender-egalitarian views) were the rarest response for five from the six indicators, with the exception being V4 (mothers who don't need the money shouldn't work). Since our data are constrained by maxima and minima, some of our sample is not able to adjust in response to birth outcomes. This phenomenon is stronger in the right tail, where between 15% and 40% of the sample is already at its maximum value. Thus any distributional change is confined to the center and left tail of our attitudinal markers, and our models may understate true effects due to attitudinal change being missed within the top ordinal category.

3 Models and Estimates

Baseline Models

Here we present our baseline models, showing that parents of daughters report more gender-egalitarian values than parents of sons. If the sex of a child is random, effect sizes are straightforward to estimate and are equal to raw differences in average outcomes. However, (as outlined below) confounding in this context is still possible, and including controls will reduce predictive error and provide tighter confidence intervals, even if not required for unbiased estimates.

Our main models are linear regressions in the spirit of Ferrer-i-Carbonell and Frijters (2004) (ordinal models that relax cardinality assumptions are presented later in the paper). The basic specification is given in EQ (1), and we work with four different functional configurations.

$$y_{jit} = \alpha_l + \psi_m + \gamma_t + \phi D_{it} + \mathbf{x}_{it}^{'}\boldsymbol{\beta} + \varepsilon_{it} \qquad j = 1, ..., k \tag{1}$$

Here $y_{jit} \in \{1, 2, ..., 7\}$ is attitudinal indicator j of individual i in time period t. D_{it} is the number of daughters, and α_l, ψ_m and γ_t represent individual-specific effects on child numbers, households and years respectively. Vector $\mathbf{x}_{it}^{'}$ represents individual time-varying controls. The specification is therefore designed such that parameter ϕ captures the *ceteris paribus* difference in attitudes associated with a child being a daughter rather than a son.

We estimate EQ (1) by OLS, allowing for correlations between α_l, ψ_m, γ_t and $\mathbf{x}_{it}^{'}$. We do not report fixed-effect on individuals (which use only longitudinal variation to identify parameters) as there is insufficient intertemporal variation to generate useful estimates. Nonetheless, estimates from these models produce effect sizes in line with those below, albeit with larger standard errors.[16]

Our four specifications of EQ (1) are as follows. Under the assumption of random treatment assignment, estimates with no controls are presented in Column 1. To account for potential associations between cultural attitudes and behaviors that affect sex ratios), we add our pre-treatment controls in Column 2. We regard this as our preferred model for estimating treatment effects. Column 3 presents estimates adding in the contemporaneous set of controls. Lastly, Column 4 adds household fixed-effects to the pre-treatment controls, which remove all heterogeneity associated with inter-household variation, but increases the risk of attenuation bias (Pischke, 2007).

Results in Table 2 show estimates for V1 and V2 - our variables measuring whether (i) men are perceived to make better political leaders, and (ii) whether families do better if men earn more than their spouses. Values for $\hat{\phi}$ are always positive for both attitudinal variables, and are significant

[16] These estimates are available upon request.

in all models, except the household fixed-effects specifications. In terms of magnitudes, parenting a daughter rather than a son increases disagreement with the idea that men are superior political leaders by 0.04-0.08 points.[17] Similarly, daughters raise disagreement with the notion that of it being better for relationships if the man earns more by 0.03-0.06 units, with estimates clustered towards the higher end of this range. While significant, we note that these are fairly subtle effect sizes - at the upper end each corresponds to a shift of 0.03-0.04 standard deviations.

Table 2: Effects of Child Gender on Attitudes to Gender Equality - V1 & V2

	V1. Men Better Political Leaders†				V2. Better if Man Earns More†			
	M1	M2	M3	M4	M1	M2	M3	M4
$\hat{\phi}$	0.082***	0.077***	0.038	0.071***	0.058***	0.045***	0.034	0.046***
$SE\left(\hat{\phi}\right)$	(0.019)	(0.019)	(0.075)	(0.018)	(0.016)	(0.016)	(0.063)	(0.016)
Controls P-T	N	Y	N	N	N	Y	N	N
Controls Cont	N	N	Y	Y	N	N	Y	Y
Year FE	Y	Y	Y	Y	Y	Y	Y	Y
Child # FE	Y	Y	Y	Y	Y	Y	Y	Y
Household FE	N	N	Y	N	N	N	Y	N
R^2	0.010	0.086	0.621	0.095	0.007	0.043	0.606	0.058
F	19.09	74.55	34.27	77.26	13.70	24.49	2.627	33.63
N	19267	18038	15615	19231	19279	18050	15631	19243

Note: The table presents estimates of parameter ϕ from models derived from EQ (1). The LHS gives estimates for on male vs female political leaders (V1) and the RHS on whether it is detrimental for relationships if the woman earns more money (V2). M1 regresses outcomes on fixed effects for child number and year. M2 includes pre-treatment controls based upon race, gender, class and family background. M3 also includes household-specific fixed effects and M4 uses standard contemporaneous controls such as income, education and marital status. Standard errors are clustered at the household level. *, ** and *** denote significance at 10%, 5% and 1% respectively. † indicates that the variable has been inverted such that higher values indicate greater preferences for equality.

Estimates in Table 3 show the effects on V3 and V4 - (dis)agreement with the beliefs that (iii) it is better if men are breadwinners, and (iv) mothers who don't need the money shouldn't work. Again our estimates are always positive and are significant in seven from eight cases. The effect sizes are similar to those reported in Table 2 (0.04-0.05, or 0.03 standard deviations), although the household fixed-effects models produce much higher estimates (0.182 and 0.134 for V3 and V4 respectively). Notably these models also have greater standard errors - approximately four times larger than those from the other models. Thus, the point estimates in Table 2 would be significant at standard levels if the SEs remained the same. It therefore appears that the discrepancies in effect sizes are due to increased sampling error, where the household fixed-effects strip away most of the identifying variation in the data.

[17] In terms of our seven point linear scale.

Table 3: Effects of Child Gender on Attitudes to Gender Equality - V3 & V4

	V3. Men Should Be Breadwinners†				V4. Better if Mothers Don't Work†			
	M1	M2	M3	M4	M1	M2	M3	M4
$\hat{\phi}$	0.051***	0.037**	0.182**	0.025	0.057***	0.052***	0.134*	0.042**
$SE\left(\hat{\phi}\right)$	(0.019)	(0.018)	(0.073)	(0.017)	(0.019)	(0.019)	(0.069)	(0.018)
Controls P-T	N	Y	N	N	N	Y	N	N
Controls Cont	N	N	Y	Y	N	N	Y	Y
Year FE	Y	Y	Y	Y	Y	Y	Y	Y
Child # FE	Y	Y	Y	Y	Y	Y	Y	Y
Household FE	N	N	Y	N	N	N	Y	N
R^2	0.022	0.105	0.677	0.136	0.027	0.056	0.659	0.099
F	38.68	95.67	14.73	122.09	52.980	49.23	8.333	85.03
N	24492	22973	19936	24437	24505	22982	19942	24449

Note: The table presents estimates of parameter ϕ from models derived from EQ (1). The LHS gives estimates for views on gender norms for breadwinner roles (V3) and the RHS on whether financially secure mothers should work (V4). M1 regresses outcomes on fixed effects for child number and year. M2 includes pre-treatment controls based upon race, gender, class and family background. M3 also includes household-specific fixed effects and M4 uses standard contemporaneous controls such as income, education and marital status. Standard errors are clustered at the household level. *, ** and *** denote significance at 10%, 5% and 1% respectively. † indicates that the variable has been inverted such that higher values indicate greater preferences for equality.

Table 4: Effects of Child Gender on Attitudes to Gender Equality - V5 & V6

	V5. Share Equally in Housework				V6. Children OK with Reversed Parenting			
	M1	M2	M3	M4	M1	M2	M3	M4
$\hat{\phi}$	0.005	0.003	0.066	0.007	0.000	-0.012	0.017	-0.012
$SE\left(\hat{\phi}\right)$	(0.019)	(0.018)	(0.073)	(0.017)	(0.015)	(0.015)	(0.061)	(0.014)
Controls P-T	N	Y	N	N	N	Y	N	N
Controls Cont	N	N	Y	Y	N	N	Y	Y
Year FE	Y	Y	Y	Y	Y	Y	Y	Y
Child # FE	Y	Y	Y	Y	Y	Y	Y	Y
Household FE	N	N	Y	N	N	N	Y	N
R^2	0.009	0.045	0.596	0.048	0.006	0.058	0.631	0.063
F	19.21	45.68	25.65	44.87	10.93	39.53	8.69	40.56
N	24513	22989	19948	24457	24475	22961	19924	24419

Note: The table presents estimates of parameter ϕ from models derived from EQ (1). The LHS gives estimates for attitudes towards housework (V5) and the RHS for the efficacy of counter-traditional parenting roles (V6). M1 regresses outcomes on fixed effects for child number and year. M2 includes pre-treatment controls based upon race, gender, class and family background. M3 also includes household-specific fixed effects and M4 uses standard contemporaneous controls such as income, education and marital status. Standard errors are clustered at the household level. *, ** and *** denote significance at 10%, 5% and 1% respectively.

While we observed consistently positive and significant results in Tables 2 and 3, estimates from Table 4 show that not all our attitudinal variables are responsive to birth composition. Neither V5 (a preference for equally shared housework when both partners are employed) nor V6 (the notion that children do just as well when parenting roles are reversed) appear affected. Across all models, these estimates are close to zero, insignificant, and occasionally negative - precisely what we would expect if birth outcomes have no effect upon these variables.

What accounts for these null results, and why do they differ from those in Tables 2-3? A prosaic explanation is that these may be false negatives, occurring simply because the correlations are not strong enough to show up repeatedly in our data. If this is the case, we would continue to conclude that child gender has widespread effects on attitudes.

However, another possibility is that these variables are crucially different in nature to V1-V4, and that the phenomena they capture are genuinely unresponsive to child gender. Notably, V5-V6 may be sensitive to both pragmatic intra-household organizational factors, as well as views on the roles of women in society. This would have the effect of adding variation to the covariate of interest and therefore attenuate the coefficient towards zero.[18] For example, there is evidence that intensity in paid work differs between men and women (Bolotnyy and Emanuel, 2018), and women may have greater opportunity costs to employment (Cook *et al.*, 2018), which may lead them to take on greater household responsibilities. Thus, an individual may hold egalitarian social views (e.g., believe that men and women are of equal political skill - V1), but feel that women should take on more housework out of practical considerations. Similarly, attachment theory posits that children form a deep bond with one caregiver and require ongoing contact with that individual (Ainsworth, 1978). As this is disproportionately the mother (Fox *et al.*, 1991), parents may feel that maternal caregiving is better for children (i.e. V6), while still holding emancipatory political views.

Ordinal Models and Distributional Effects

In this section, we present two extensions of the results outlined above. Firstly, we relax the assumption of cardinality (which may be suspect for our 1-7 assessment scales) using an ordered probit model as an analog to EQ (1). Secondly, we use this model to study distributional effects, rather than simply estimate changes in conditional means. The equation we use is of the general form $y^*_{it} = \alpha_l + \gamma_t + \phi D_{it} + \mathbf{x}'_{it}\boldsymbol{\beta} + \varepsilon_{it}$, where y^*_{it} is a latent variable assigning y_{it} to values $\{1, ..., 7\}$ across intervals $-\infty < y^*_{it} < \pi_1$, $\pi_1 < y^*_{it} < \pi_2$,... $\pi_7 < y^*_{it} < \infty$. Parameters α_l, γ_t, ϕ, $\boldsymbol{\beta}$ and $\pi_1, \pi_2, ...\pi_7$ are estimated using Maximum Likelihood, and again our standard errors account for clustering at the household level.[19] For brevity, we report results in terms of changes in probability of parents reporting the most egalitarian response $\Delta P\,(y = 7)$ evaluated at the sample means.

[18] Such a phenomenon would need to be more apparent in V5 and V6 than the other covariates to have this effect in our data.
[19] Household fixed-effects estimations are excluded here due to convergence problems in these models.

Table 5: Marginal Effects - Ordered Probit Models

	V1	V2	V3	V4	V5	V6
$\Delta P\,(y=7\mid x=\bar{x})$	0.0203***	0.0182***	0.0083**	0.0087***	0.0019	-0.0004
$SE\,(\Delta P)$	0.0045	0.0047	0.0033	0.0028	0.0041	0.0027
Controls	N	N	N	N	N	N
Child #FE	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y
N	19,267	19,279	24,492	24,505	24,513	24,475
$\Delta P\,(y=7\mid x=\bar{x})$	0.0204***	0.0151***	0.0066**	0.0081***	0.0010	-0.0031
$SE\,(\Delta P)$	0.0044	0.0048	0.0033	0.0029	0.0043	0.0033
Controls Pre	N	N	N	N	N	N
Child #FE	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y
N	18,038	18,050	22,973	22,982	22,989	22,961
$\Delta P\,(y=7\mid x=\bar{x})$	0.0187***	0.0155***	0.0040	0.0064**	0.0022	-0.0031
$SE\,(\Delta P)$	0.0046	0.0046	0.0031	0.0027	0.0042	0.0032
Controls Cont	N	N	N	N	N	N
Child #FE	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y
N	19,231	19,243	24,437	24,449	24,457	24,419

Note: The table presents results from ordered probit regressions of all six attitudinal variables upon various sets of controls. The reported estimate is the change in probability for a parent to express the highest level of agreement/disagreement with the given statement, with all other covariates set at the sample means. Cluster-robust inference is used throughout. *, ** and *** denote significance at 10%, 5% and 1% respectively.

The results in Table 5 reinforce those from the linear model, indicating that the baseline results are not an artifact of specification error. The ordinal estimates are actually a little more significant, and show that for variables V1-V4, daughters increase the probability of the average parent holding strongly gender-egalitarian values by around 1-2%. As above, no significant results were found for attitudinal markers V5 and V6.

To assess the strength of the relationships at different values of our dependent variables, we contrast the estimates above with probability changes at the other outcome levels (i.e. $y = 1, ..., 6$). This allows us to determine whether female birth outcomes act (for example) at the center-right tail by strengthening existing egalitarian views, or at the left tail, by drawing relatively *in*egalitarian attitudes towards the center. We suppress the results for the sake of brevity,[20] but note that effect sizes tend to be small, but negative, for outcomes 1-5, indicating that daughters slightly reduce the rates of a series of less-egalitarian responses. Large positive effects then appear for outcome $y = 7$. The effects of female births are therefore not concentrated in only the right tail of the distribution. Rather, having daughters seems to have small effects over the center and left tail, which, when aggregated, translate into large increases in the probability of expressing a highly egalitarian viewpoint.

[20]These estimates are available upon request.

4 Heterogeneous Effects and Mechanisms of Social Change

Having established some general links between birth composition and attitudes, we now drill more deeply into the data to see if there are population subsets for whom the effects are stronger. Uncovering this heterogeneity is valuable as it offers some clues regarding the mechanisms through which the observed social changes occur. Initially, we proceed by estimating our preferred model from EQ (1) over different demographic strata. As other authors have noted, fathers tend to alter their social attitudes more strongly when parenting daughters (we examine this in more detail using the model in the appendix), and we look for similar patterns with respect to age and education.[21] We divide our sample into approximately equal-sized groups on the basis of each of these variables,[22] and present the results in Table 6.

Table 6: Heterogeneous Effects by Sex, Age and Education

	V1	V2	V3	V4	V5	V6
Male	0.096***	0.037	0.055**	0.086***	-0.006	0.010
	(0.029)	(0.023)	(0.025)	(0.025)	(0.019)	(0.021)
Female	0.062***	0.052**	0.023	0.027	0.011	-0.027
	(0.022)	(0.021)	(0.023)	(0.023)	(0.015)	(0.018)
Younger	0.065**	0.039*	0.039	0.037	0.015	0.021
	(0.027)	(0.023)	(0.025)	(0.026)	(0.017)	(0.021)
Older	0.085***	0.053**	0.025	0.059**	-0.009	-0.050**
	(0.026)	(0.024)	(0.026)	(0.027)	(0.018)	(0.021)
More Educated	0.068*	0.072*	0.015	0.078**	-0.037	-0.043
	(0.040)	(0.037)	(0.037)	(0.038)	(0.024)	(0.031)
Less Educated	0.082***	0.030	0.057***	0.056**	0.019	-0.011
	(0.023)	(0.020)	(0.022)	(0.022)	(0.016)	(0.018)

Note: The table shows estimates from our preferred specification from EQ (1) (pre-treatment controls, year and child # fixed effects) where the data are stratified by gender, age and education. All estimates are produced using OLS and standard errors are clustered at the household level. *, ** and *** denote significance at 10%, 5% and 1% respectively.

Obtaining significant estimates is harder in this context as our sample sizes are much smaller, however some interesting patterns still emerge. As expected, we do uncover more convincing estimates for men, although we also obtain significant estimates for women (of the expected sign) for our first two variables. Where significant effects are present, the parameters do not change much between men and women. Rather, it appears that for men, the impact of having daughters is wider, in that it manifests more robustly, across a greater range of indicators. Our estimates are also larger for parents in the lower educational bracket, where traditional values tend to be stronger (Feldman and Newcombe, 2020). Notably, we also see somewhat more significant effects for older individuals.

[21] The central insight is that gaps in social attitudes are likely to be larger between fathers and daughters than for mothers, since men hold more traditionally conservative views on gender issues. Reconciling these views across generations will therefore require men to moderate their views more so than women.

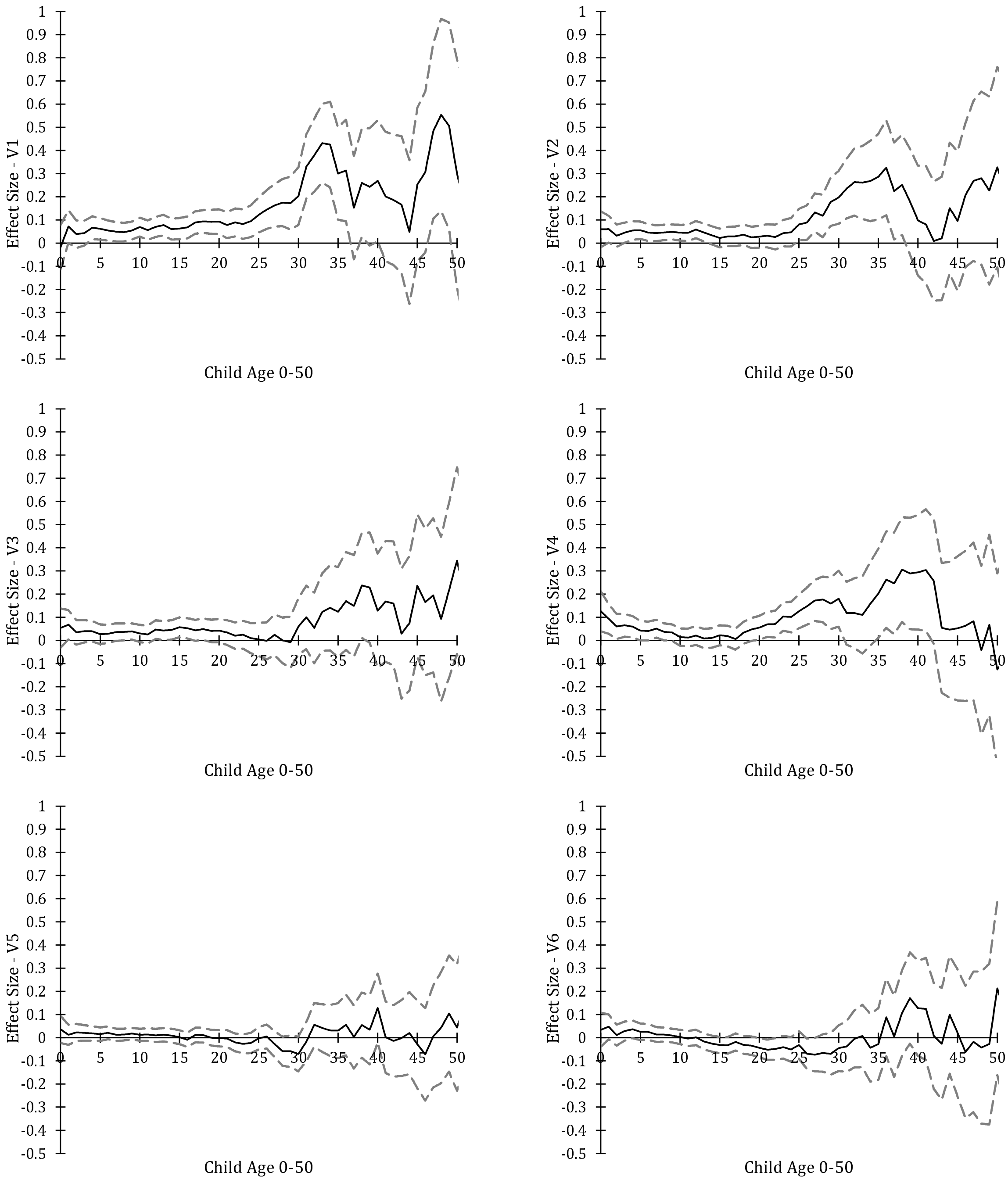
[22] Age is partitioned at 42 (41 and younger; 42 and older) while education is on the basis of at least some post-secondary attainment (including non-university education).

This heterogeneity associated with age sheds some light on the two theories of attitudinal change outlined in the introduction, as it is more consistent with the *exposure* hypothesis rather than the *identity* hypothesis. Since parents learn their children's gender very quickly, identity effects should appear rapidly based upon a newly acquired set of incentives. Conversely, change through exposure occurs slowly, when children are autonomous and express their own viewpoints. Exposure effects are therefore more likely to be present in older individuals.

To examine this issue in greater detail, we take the same specification as above, and produce rolling estimates, using the age of a parent's first child as the estimation frame. Again the idea here is to determine *when* any attitudinal change takes place. Figure 1 plots effect sizes (OLS estimates for ϕ) for variables V1-V6 (with 90% confidence intervals in grey), where the horizontal axes are based upon averages from 10-year rolling windows at the age of the firstborn child.

In all six panels, we see very little evidence of attitudinal change in parents in the first few years following childbirth. The point estimates are always close to zero, and are insignificant in all instances except V4 (note that the reduced sample sizes in rolling regressions make significance less likely here). However, across variables V1-V4, where we obtained significance in our full-sample models, there is no sign of structural change until children reach 20-25. Beyond this point the effect sizes are much larger, and much more likely to be significant at standard levels. Thus, the opinions of parents of sons and daughters do not begin to diverge until their offspring reach adulthood themselves. Since the transition to young adulthood involves substantial social development and expanded political rights, the correlations we observe are likely driven by changes in children rather than the parents themselves.

Figure 1: Heterogeneity in Effect Sizes Over First Born Child Age



Note: The figure depicts estimates of ϕ (vertical axes) for V1-V6 based upon rolling regressions for child age (horizontal axes). We use our preferred Specification 2 throughout and report estimates for each child age {0-50} +/- 4.5 years obtained by OLS. 90% confidence intervals are given using grey dashed lines. Note that stratifying the sample in this manner means that large standard errors (due to small sample sizes) will appear at higher child ages.

Do Daughters Affect Personality?

If the gender composition of children influences their parents' political attitudes, then the effects may run deeper and appear in other dimensions of human psychology. In such an instance, the changes in social values outlined above would reflect only a subset of a spectrum of potential psychological shifts. Here, we study whether these attitudinal shifts coincide with changes to broader notions of personality. While individuals' personalities are regarded as moderately stable during adulthood (Cobb-Clark and Schurer, 2012), they are also known to be responsive to major life changes (Specht *et al.*, 2011). We search for wider effects by taking our preferred model from EQ (1) and replacing the attitudinal markers with psychometric scores from the big five personality traits. These variables represent a five-dimensional taxonomy known to capture most of the interpersonal variation in psyche (John and Srivastava, 1999). The dimensions are captured with 1-7 scales depicting *Extraversion* (outgoing, talkative *vs.* quiet, passive), *Agreeableness* (friendly, pleasant *vs.* confrontational, challenging), *Conscientiousness* (disciplined, organized *vs.* careless, messy), *Openness* (inquisitive, adventurous *vs.* cautious, limited) and *Neuroticism* (sensitive, anxious *vs.* carefree, confident).

Does the gender of a child affect parents along these lines? The estimates in Table 7 below indicate that some subtle effects are indeed present. As above, we are testing many hypotheses (35 in total), and hence the occasional false positive is expected. We see occasional rejections of the null of no effect for extraversion, agreeableness, and emotional stability, however the effect sizes here are small, and appear only at lower levels of significance. Thus the evidence for child gender affecting these variables is relatively weak. However, the frequent rejections of the null hypothesis with respect to conscientiousness (and the uniformly negative signs) suggest a statistical link is present for this dimension. Here, our estimates are significant for the pooled sample, and are stronger for men, older individuals, and persons with lower educational attainments - precisely the same population subgroups that were more responsive in terms of the attitudinal indicators. The fact that these results coincide so closely strengthens the case that genuine sociocultural change is occurring within these sections of the population.

Table 7: Effects of Daughters on Big Five Personality Traits

	Extrav	Agree	Conc	Emot St	Open
Pooled	0.011	-0.011	-0.024**	-0.008	-0.012
	(0.011)	(0.009)	(0.011)	(0.011)	(0.011)
Male	-0.010	-0.009	-0.054***	-0.021	-0.003
	(0.016)	(0.014)	(0.015)	(0.016)	(0.016)
Female	0.027*	-0.012	0.001	0.004	-0.020
	(0.016)	(0.012)	(0.014)	(0.014)	(0.015)
Younger	0.031*	-0.015	-0.015	0.015	0.002
	(0.016)	(0.014)	(0.016)	(0.016)	(0.016)
Older	-0.009	-0.007	-0.033**	-0.029*	-0.022
	(0.016)	(0.013)	(0.015)	(0.015)	(0.015)
Less Educated	0.000	-0.034**	-0.031*	-0.001	-0.011
	(0.020)	(0.017)	(0.019)	(0.020)	(0.019)
More Educated	0.015	-0.001	-0.022*	-0.012	-0.018
	(0.014)	(0.011)	(0.013)	(0.013)	(0.013)

Note: The table gives estimates of parameter ϕ from our preferred specification for EQ (1) (pre-treatment controls, child # and year fixed effects) where the dependent variable is replaced with big five personality traits. The first row gives estimates for the pooled sample while stratifications based on gender, age and education appear in the rows below. Results for extraversion, agreeableness, conscientiousness, emotional stability and openness to experience appear in the columns. Standard errors are clustered at the household level. *, ** and *** denote significance at 10%, 5% and 1% respectively.

Why might daughters make their parents less conscientious? The result is curious, as females are substantially higher in this trait than males (Vecchione *et al.* 2012). One potential answer comes from conscientiousness requiring mental effort, and that this effort may produce positive spillover effects within families (Alvarez and Miles, 2003). Conscientious effort from one individual will create headroom for others, either to relax, or to substitute their mental efforts elsewhere. Such a process would generate negative correlations between effortful behaviors across individuals that shared common tasks, and explain the results obtained above. Interestingly, conscientiousness is also an input into a spectrum of human behaviors (Jackson *et al.*, 2010), and therefore may produce small but widespread flow-on effects. For instance, if reductions in conscientiousness appear across the board, we would expect to see contractions in a wide variety of mentally effortful behaviors. Conversely, if what we are observing is due to a substituting of mental effort, this implies that parents of daughters (and perhaps conscientious children in general) may achieve more highly in non-familial domains. In such an instance the effect would likely occur due to a diminished need to allocate effort to child-raising. We leave this matter to future research.

5 Causal Diagnostics

So far throughout the paper, we have been fitting our econometric estimates with causal interpretations, attributing variations in individuals' attitudes and psychology directly to birth outcomes.

In this section we consider the validity of these interpretations. While the apparently random nature of child gender would imply our estimates are indeed causal, there are some instances where this assumption of exogeneity may break down. If parents are able (either deliberately or subconsciously) to influence the gender of their children, and the tendency to do so is associated with political attitudes, this will open up non-causal correlations in our data. Equally, even if parents cannot affect child gender, they may have additional children as a way of manipulating their overall familial composition, which would have the same effects upon our estimates.

We consider several threats to identification along these lines. Using some new diagnostic methods, we study (i) the potential for unobservables to affect our estimates (by assessing treatment assignment, covariate balance and coefficient stability), and (ii) biases arising from endogenous stopping rules. In all cases, our results are consistent with causality flowing from birth outcomes to our attitudinal variables.

Potential Sources of Bias

Initially, we consider the evidence that parents could possess unobservable characteristics that jointly drive the key LHS and RHS variables in EQ (1). This argument implies a two-part evidentiary chain, where (i) child gender preferences are linked with social attitudes, and (ii) parents with child-preferences go on to influence the gender composition of their families. Consider a scenario where parents with traditional views engineer male births, while parents with progressive views do the reverse. This process would exert an upward bias on statistical estimates of ϕ in our models.

The idea that gender preference could be explicitly linked to social views can be empirically assessed. In addition to the variables already sourced, HILDA contains a question on whether parents would prefer a future child to be a boy, a girl, or are equally open to either. In Table 8 below, we stratify our attitudinal indicators by this indicator to see if child-gender preferences are related to sociopolitical views.

Our data show that most parents are equally happy with boys or girls (58%), and that these individuals were also the most egalitarian in their views. Approximately 22.1% of our sample preferred male births, and these parents expressed the lowest levels of egalitarian sentiment. Parents who favored girls (19.9%) were typically midway between the former two groups. In around half of all cases these differentials were significant. Thus, while son preference is correlated with traditional views on gender roles, the reverse (daughter preference predicting gender-egalitarian views) does not hold. Nonetheless, since the data show that child gender preference is at least *partially* linked to social attitudes, there is meaningful evidence supporting the first link of our two-part evidentiary chain.

Table 8: Parental Views by Preferred Gender of Next Child

Variable	No Prefer	Prefer Boy	Prefer Girl	No Pref-Boy	No Pref-Girl	Pref Girl-Boy
V1	5.539	5.113	5.416	0.426***	0.123	0.303***
V2	5.944	5.705	5.716	0.239***	0.228***	0.0107
V3	5.072	4.666	4.785	0.406***	0.287***	0.119
V4	4.824	4.552	4.590	0.272***	0.234***	0.038
V5	5.960	5.866	6.007	0.094	-0.047	0.141*
V6	5.475	5.398	5.368	0.077	0.107	-0.030
	58.0%	22.1%	19.9%	-	-	-

Note: The table presents averaged attitudinal variables V1-V6 stratified by response to parental preference for the gender of their next child. 58% of individuals expressed no preference and their attitudinal aggregates are given in the first column. Results for parents who prefer a boy (22.1%) or a girl (19.9%) are in the second and third column. Differences are presented on the right. *, ** and *** denote significance at 10%, 5% and 1% respectively.

Could these preferences for male or female births feed through to affect sex ratios? Child gender is plausibly affected by reproductive timing (James, 1987), and can be induced artificially via In Vitro Fertilization (IVF). However the former method is unreliable and provides minimal shifts in the likelihood of male/female births (Cagnacci *et al.*, 2003), while the latter is illegal in Australia, and therefore unlikely to be widespread. Gender-specific abortions are another possibility, where social attitudes may make marginal changes to the likelihood that a fetus is aborted, resulting in affected pregnancies being eliminated from our sample. Empirical evidence suggests that gender specific abortions do occur in developed countries, usually with a preference for boys (Edvardsson *et al.*, 2018). However, even in the most extreme instances, sex ratios vary only by a percentage point or so, which suggests again that this is unlikely to meaningfully affect our estimates. Deliberate engineering of the sex of a specific child therefore seems a negligible source of statistical bias.

However, sex ratios may also be affected incidentally, by other forms of behavior. There is some evidence that older parents are more likely to have girls (Jacobsen *et al.*, 1999; Mathews and Hamilton, 2005), and lifestyle factors such as smoking (Koshy *et al.* 2010) and diet are also known to affect child gender. Since male fetuses are more prone to miscarriage (the *fragile male hypothesis*),[23] biological (and socioeconomic) stressors tend to shift ratios towards female births (Kraemer, 2000). If these incidental factors differ strongly across birth outcomes, then it is plausible that other unobserved differences may also be present. We present such an analysis below.

Covariate Balance - Newly Born Daughters and Sons

We examine whether child gender is plausibly correlated with our two sets of covariates in Table

[23]Since male births are larger but gestate for the same length of time as female births, they grow faster, which leaves them more vulnerable to maternal health shocks. For this reason we expect to see more female births to mothers who experience health complications as males will be more likely to miscarry.

9, while equivalent results for a set of health behaviors appear in Table 10. In all cases, we take new births within the sample and stratify our data by outcomes from that year. The rationale for focusing only on new births is that it allows us to circumvent any reverse-causal flows, such that we can consider the effects of both background and contemporaneous variables. Here, random treatment assignment will result in the same distributions of covariates for parents of sons and daughters (Pei *et al.*, 2019). Means and standard deviations for both groups are given in the leftmost and central columns while differences (with t-tests) are presented on the right, where significant differences imply some form of selection.

Table 9: Covariate Balance - New Male/Female Births

	Daughters		Sons		Diff		Daughters		Sons		Diff
Contemp	$\bar{x}_D$	$\hat{\sigma}_D$	$\bar{x}_S$	$\hat{\sigma}_S$	$\bar{x}_D - \bar{x}_S$	Pre-Treat	$\bar{x}_D$	$\hat{\sigma}_D$	$\bar{x}_S$	$\hat{\sigma}_S$	$\bar{x}_D - \bar{x}_S$
Age	36.90	9.823	36.61	9.678	0.296	Age	36.90	9.823	36.61	9.678	0.296
Age Squared	1458	809.8	1434	796.0	24.56	Age Squared	1458	809.8	1434	796.0	24.56
Aborig/TS Is.	-	-	-	-	-	Aborig/TS Is.	-	-	-	-	-
Married	0.701	0.458	0.703	0.457	-0.002	Male	0.463	0.499	0.456	0.498	0.007
De Facto	0.212	0.409	0.196	0.397	0.016	Female	0.537	0.499	0.544	0.498	-0.007
Separated	0.020	0.141	0.022	0.148	-0.002	Born Aust	0.805	0.396	0.810	0.393	-0.005
Divorced	0.029	0.167	0.026	0.160	0.002	Born Eng C	0.080	0.271	0.082	0.275	-0.003
Never Married	0.005	0.068	0.007	0.084	-0.002	Born N-Eng	0.115	0.320	0.108	0.310	0.007
Male	0.463	0.499	0.456	0.498	0.007	Father Empl	0.828	0.378	0.843	0.364	-0.015
Female	0.537	0.499	0.544	0.498	-0.007	Both Parents	0.751	0.433	0.743	0.437	0.007
Major City	0.609	0.488	0.596	0.491	0.013	Had Siblings	0.973	0.161	0.968	0.175	0.005
Inner Region	0.261	0.439	0.244	0.429	0.018	Mother HS	0.288	0.453	0.291	0.454	-0.003
Outer Region	0.108	0.311	0.142	0.349	-0.033***	Father HS	0.276	0.447	0.279	0.449	-0.003
Remote	0.016	0.126	0.015	0.122	0.001	Mother Tert	0.185	0.388	0.175	0.380	0.010
Very Remote	0.005	0.069	0.003	0.058	0.001	Father Tert	0.197	0.398	0.206	0.404	-0.009
Born Aust	0.805	0.396	0.810	0.393	-0.005	Mother Univ	0.102	0.303	0.110	0.313	-0.008
Born Eng C	0.080	0.271	0.082	0.275	-0.003	Father Univ	0.144	0.351	0.129	0.335	0.015**
Born N-Eng	0.115	0.320	0.108	0.310	0.007						
Employed	0.633	0.482	0.642	0.479	-0.010						
Unemployed	0.031	0.174	0.032	0.177	-0.001						
Not in Labor F	0.336	0.473	0.325	0.469	0.011						
Log HH Inc	10.83	0.588	10.801	0.590	0.024*						
Post-Grad Ed	0.114	0.318	0.103	0.304	0.011						
Bachelors Ed	0.174	0.379	0.180	0.384	-0.007						
Diploma	0.349	0.477	0.339	0.473	0.010						
High School	0.140	0.347	0.152	0.359	-0.012						
Less than HS	0.223	0.417	0.226	0.418	-0.002						

Note: The table presents sample sizes, means and standard deviations for the HILDA subsample. Contemporaneous covariates are given in the left panel and pre-treatment covariates are given in the right panel. Calculations are performed on a pooled sample from 2001-2018 with 106, 615 observations. Cardinality is assumed for the calculation of means and standard deviations of ordinal variables. The data set is trimmed to contain individuals with at least one child.

The results in Table 9 show that there are virtually no differences in demographics or socioeconomics between parents of sons and daughters in the year of their birth. Across a host of contemporaneous and background variables, the only significant difference is in the sex ratios in very remote regional areas, which are more likely to report male births. However given the large number of hypothesis tests in Table 10, and the marginal significance of this variable (it is significant at 5% but not 1%), we argue that false positives as a result of multiple comparisons provide a plausible explanation. Nonetheless, if the probability of male birth does vary by geographical area, this would upwardly bias our parameter estimates, as per the example at the beginning of this section. Remote regional areas in Australia are typically socially conservative, and the local economies tend to be more dependent upon manual labor. Both these factors are likely to coincide with preferences for sons (e.g. see Dahl and Moretti (2008) and references therein). However, even if present, such an effect would affect less than 1% of our sample, and again likely be too small to produce meaningful statistical biases.

Equivalent results for health behaviors are given in Table 10. In line with the fragile male hypothesis, there is a higher rate of parental smoking for female births, although the difference is not significant at standard levels. We also see no significant differences for alcohol consumption (with a reverse sign to smoking), physical exercise, or diet, except the consumption of pasta/carbohydrates, which was slightly predictive of male births. In contrast to tobacco use, we are unaware of any cultural association between carbohydrate consumption and sociopolitical attitudes. Therefore, we again do not see evidence that male and female births are being drawn disproportionately from sub-populations with differing behavioral patterns.

Table 10: Covariate Balance - Health Behaviors

	Daughters		Sons		Diff
	$\bar{x}_D$	$\hat{\sigma}_D$	$\bar{x}_S$	$\hat{\sigma}_S$	$\bar{x}_D - \bar{x}_S$
Freq Vegetable	5.687	1.600	5.802	1.492	-0.115
Freq Pasta/Carbohydrate	4.740	1.183	4.866	1.147	-0.126**
Freq Snack Food	4.032	1.415	4.025	1.359	0.006
Freq Potato/Starch	3.674	1.159	3.607	1.180	0.067
Freq Red Meat	5.049	1.084	5.008	1.098	0.041
Freq Proc Meat	4.236	1.298	4.226	1.312	0.010
Freq Physical Activity	3.382	1.516	3.343	1.516	0.039
Cigarette Consumption	84.54	65.92	78.80	60.87	5.735
Freq Alcohol Consumption	5.473	2.331	5.555	2.272	-0.082

Note: The table presents averaged indicators of health behaviors for new births within our sample. The left two columns give averaged frequencies for parents who had a daughter while the middle two columns give the same values for parents who had a son. The difference in outcomes (with a t-test for significance) is presented in the rightmost columns. *, ** and *** denote significance at 10%, 5% and 1% respectively.

Coefficient Stability

An alternative method of assessing potential bias involves assessing the stability of causal parameters from EQ (1) over differing configurations of control variables. The key idea is that the degree of confounding by omitting observables offers some guidance as to the effects of excluding unobservables. If D is randomly assigned, then ϕ is unconfounded, and will be stable in the presence of any pre-treatment controls. Conversely, if unobservables affect y, then the inclusion of observed controls will reduce confounding, provided there is an association between observed and unobserved confounders. Thus, if estimates of ϕ are sensitive to permutations of variables from the observed set, they are unlikely to be robust to the inclusion of additional (hypothetical) controls from an unobserved set.

We measure this effect in two ways: a stability measure developed by Altonji *et al.* (2005) (popularized by Nunn and Wantchekon (2011)), and a related method derived by Oster (2019). The former involves comparing estimate $\hat{\phi}_R$ from a restricted model (with limited or no controls) to that from a fully specified model $\hat{\phi}_F$ containing the full set of observed controls. The ratio $U = \hat{\phi}_F / \left(\hat{\phi}_R - \hat{\phi}_F\right)$ then captures the effect of observables (given by $\hat{\phi}_R - \hat{\phi}_F$), and $\hat{\phi}_F$ is expressed relative to this quantity. In instances where the causal parameter is unconfounded, the denominator is zero (aside from sampling error) and U approaches infinity. The effect of unobservables can be measured using the value of U required to transform $\hat{\phi}_F \rightarrow 0$. In this instance, unobservables must be U times more important than observables to reduce the causal parameter to zero, where a common heuristic is that U should exceed 3.

Oster (2019) presents a refinement of this idea, where the effects of additional controls are scaled by the increase in model fit provided. Consider a regression where irrelevant (independent) random variables are included as controls, and stability is assessed using U. In such an instance $\phi_R = \phi_F$ and hence the estimates will incorrectly appear robust. The key to diagnosing this error is to note that controls that reduce confounding must also increase model fit. Thus, stability matters when $\hat{\phi}_R - \hat{\phi}_F$ is preserved, alongside positive changes in R^2. By expressing parameter differentials relative to changes in fit (allowing for a maximum explainable value R^2_{Max}, which can be less than one), we can ensure that stability is not an artifact of poor quality controls.

Results for both methods are given below. We restrict ourselves to our preferred model (with fixed effects on child counts and years, but not households) and present the results in Table 11. For the Oster (2019) method we posit a maximal fit $R^2_{Max} = \Pi R^2$ where $\Pi = 1.3$, and again calculate the required proportional selection on unobservables to completely remove our effect sizes. In this instance δ should exceed one. In both cases, we treat age and its square as essential controls, and examine robustness with respect to the pre-determined controls based upon family background.

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Table 11: Coefficient Stability and The Effects of Unobservables

	V1	V2	V3	V4	V5	V6
$\hat{\phi}_R$	0.0772***	0.0508***	0.0383**	0.0493***	0.0025	-0.0118
$\hat{\phi}_F$	0.0768***	0.0450***	0.0369**	0.0522***	0.0059	-0.0051
$U\left(\hat{\phi}_R, \hat{\phi}_F\right)$	192.2	7.759	26.35	-18.25	-1.735	0.761
R^2_R	0.018	0.026	0.053	0.039	0.009	0.014
R^2_F	0.086	0.043	0.105	0.056	0.045	0.058
$\delta\left(\hat{\phi}_R, \hat{\phi}_F\right)$	36.76	12.89	6.341	10.09	-3.536	-4.699

Note: The table presents estimates from our preferred model using (i) a restricted set of controls (year and child # fixed effects) and our full set of predetermined controls. ϕ_R and ϕ_F are the parameter estimates while $U(\phi_R, \phi_F)$ gives the ratio of unobservables to observables required to offset our estimates. R^2_R and R^2_F give the fit of the restricted and fully specified models, while $\delta(\phi_R, \phi_F)$ gives Oster's (2019) statistic using $\Pi = 1.3$.

The Altonji *et al.* (2005) results are considered first and appear in the top three rows. Over the first four indicators (that were significant in our original models), the estimates are nearly identical when the additional controls are included. Indeed the largest discrepancy is on the second variable, where the uncontrolled and controlled indicators were 0.0508 and 0.0450. Taking the ratios (presented in the third row) shows that unobservables have to be dramatically more important than observables to offset our effect sizes. These values range between 7.759 and 192.2, which comfortably exceed the heuristic benchmark of three. For our fourth indicator, the ratio is negative (-18.25), which occurs when the coefficient *moves away* from zero. In this case, selection on unobservables would need to *reverse* the direction of that arising from observables to dispense with the effect size. Estimates for variables V5 and V6 were not well identified in our original models, and hence we do not expect the results to be robust. This is the case, with none of the four coefficients significant, and both ratios close to zero.

To ensure that the stability results hold when using controls that increase the explanatory power of our models, we turn to the estimates in the lower three rows.[24] In all cases, the R^2 terms increase sharply with the inclusion of the additional controls, from factors of almost two to greater than four. For our first four indicators, the delta terms are again much larger than the threshold of one (from 6.341 to 36.76) and hence the results for these variables appear not to be vulnerable to confounding from unobservables.

Endogenous stopping rules

Lastly, we consider the possibility that parents' attitudes to equality affect the number of children

[24] Note that there are some small discrepancies between these results and those reported in Table 2. This is due to missing values appearing in regressions with included control variables. In the cases of V4 and V6 this is sufficient to reverse the direction of the (small) mediation effect induced by the additional controls.

they have. Consider an example where egalitarian parents prefer daughters (or traditional parents prefer sons), and couples revise their plans for additional children based upon the composition of their current family. Since having additional children will balance gender ratios (via regression to the mean), this behavior will also create correlations between or variables of interest.[25]

We can determine whether decisions to have additional children are correlated with parental attitudes using hazard models. Usually used in the analysis of duration, these models describe processes of accumulation before stoppage, or *failure*. Here, "failure" (i.e. stopping on a particular number of offspring) is expressed as a function (i) the existing gender composition, and (ii) a collection of socioeconomic variables including our attitudinal markers. We focus on a subset of our sample that is expected not to have further children (persons over 45) and observe whether stopping rules are correlated with gender-egalitarian viewpoints. Ideally, such a regression would use attitudinal indicators taken precisely when they decided *against* having further children. However as this is infeasible with our data, we proceed with our data as structured in HILDA (where attitudes will have likely been measured after stopping). Our results therefore rest on the assumption that the attitudinal markers have remained relatively stable since this time.

Two equations are estimated, a Cox proportional hazard model, where the underlying hazard function is not specified, and an exponential proportional model, where the underlying hazard is flat. For simplicity we use continuous models (note that all observations have "failed" which excludes discrete-time models like log-logistic regressions). The basic model is

$$h\left(c\right) = h_0\left(c\right) \exp\left(\beta_1 x_1 + \beta_2 x_2 + ... + \beta_k x_k\right) \tag{2}$$

where in the Cox model $h_0\left(c\right)$ is left unspecified while $h_0\left(c\right) = 1$ defines the exponential.

[25]If male/female births appear at a fixed probability then this behavior will not affect the societal composition of births, however it will affect the distribution of outcomes across families.

Table 12: Proportional Hazard Models - Cox and Exponential: Key Estimates

	Pre-Treatment		Contemporaneous	
Variable	Cox Prop H	Exponential	Cox Prop H	Exponential
V1 Men Better Political Leaders	1.0176	1.0091	1.0233*	1.0122
V2 Better if Man Earns More	0.9838	0.9916	0.9909	0.9952
V3 Man Should Be Breadwinner	1.0035	1.0015	1.0083	1.0044
V4 Mothers Shouldn't Work	1.0022	1.0007	1.0009	1.0003
V5 Housework Should Be Shared	1.0350**	1.0185*	1.0323**	1.0171*
V6 Ok if Parents Reverse Roles	1.0051	1.0024	1.0014	1.0002
Extraversion	1.0267	1.0147	1.0271	1.0147
Agreeableness	0.9634	0.9806	0.9654	0.9822
Conscientiousness	0.9893	0.9931	0.9998	1.0002
Emotional Stability	1.0071	1.0070	1.0051	1.0044
Openness to Experience	0.9999	0.9990	1.0047	1.0017
Child # FE	Y	Y	Y	Y
Year FE	Y	Y	Y	Y
N	1,411	1,411	1,470	1,470
# Clusters	1,028	1,028	1,056	1,056
$\ln L$	-9,208	-1,514	-9,652	-1,574

Note: The table gives parameter estimates from Cox proportional hazard models (where the hazard is undefined) and an exponential hazard model. The number of children is the dependent variable and stopping on a particular value is modeled using attitudinal variables and personality characteristics. Values higher than one indicate an increased hazard ratio. *, ** and *** denote significance at 10%, 5% and 1% respectively.

We report our estimates in terms of hazard ratios, where higher values indicate a greater likelihood of parents stopping on a given number of children. Estimates that use our pre-treatment controls appear on the left, while we use the full set including contemporaneous controls on the right. Our focus here however is on the attitudinal and personality traits, which are remarkably similar in their hazard ratios across (i) parametric specifications and (ii) configurations of controls. We see very little evidence that decisions to have further children are in any way linked to gender attitudes, or even to broader personality constructs. Of our 11 psychosocial variables (and 44 hypothesis tests), only V5 is linked to reproductive outcomes - and only at fairly low levels (5% and 10%) of statistical significance. Since attitudinal and personality factors do not appear to predict childbearing decisions, we conclude that endogenous stopping behavior is not apparent in our data.

6 Socioeconomic Correlates of Egalitarian Attitudes

Having established that socioeconomic views are malleable, and are causally affected by child gender composition, we now consider some wider economic implications of these results. Specifically, we are interested in whether attitudinal variables can help explain patterns in economic disparities across developed countries, and (indirectly), if social change represents an important tool for mitigating inequality. We do not link socioeconomic inequalities directly with population sex ratios *per se*, but rather, examine whether attitudinal variations line up with commonly observed cross-national patterns, such as high inequality in the US and lower disparities in the Nordic countries.

Two exercises are performed. The first shows that progressive attitudes towards gender attitudes are associated with other egalitarian sentiments, such that the former can be used as a proxy for the latter. The second exercise shows that these sentiments are highly correlated with socioeconomic disparities across countries, a pattern that we would expect if attitudes represent an important fundamental cause. We then review some of the empirical evidence suggesting that causality does indeed flow from attitudes to variations in outcomes.

Tables 13 and 14 present the results of the first exercise, showing the associations between our gender-egalitarian variables in HILDA, and the statistical links between similar gender-oriented questions in WVS/EVS data and attitudes towards other facets of inequality. The top panels in each table present correlations (which implicitly invoke our cardinality assumption) while the lower panels give Spearman rank correlations, which preserve the ordinal structure of the variables.

Results from Table 13 show that our six gender-specific variables share positive associations with each of the other markers. Across both correlation types, we obtain 30/30 positive and significant pairwise associations. This is to be expected if fundamental notions of gender-egalitarianism are a latent factor informing all such responses, such that a positive outcome on one question predicts similar responses to all others. Nonetheless, these associations are occasionally small (in the 0-0.2 range), indicating that views on this issue are genuinely multidimensional.

The estimates in Table 14 link the variables used in HILDA to the more general set of attitudinal markers obtained from the WVS/EVS.[26] We have access to two variables that are nearly identical to the questions asked in HILDA (on men being better business leaders, and being better suited to being primary breadwinners), but also four more variables that capture different dimensions of egalitarian sentiment.

The six variables we draw here are as follows:

- *W1. Men make better political leaders than women - {1-4}.*
- *W2. Men make better business leaders than women - {1-4}.*
- *W3. Don't like as neighbors: people of a different race - {0-1}.*
- *W4. Don't like as neighbors: homosexuals - {0-1}.*

[26] Details on the data set are given in the appendix.

- *W5. (A) man's job is to earn money, (A) women's is to look after her home and family - {1-4}.*
- *W6. Important: eliminating income inequality {1-4/1-10}.*

Again we see that all 30/30 pairwise correlations are positive and significant. For example, individuals who expressed gender-egalitarian views in the EVS/WVS data were much less likely to object to having sexual or racial minorities as neighbors, and much more likely to support policy action to reduce income inequality. This further generalizes the idea that socially progressive values account for clusters of correlated responses across social attitudes.

Table 13: Correlation Structure - Gender Equality Variables

Pearson Correlations	V1	V2	V3	V4	V5	V6
V1. Men Make Better Political Leaders†	1.000					
V2. Bad For Relationship if Woman Earns More†	0.404	1.000				
V3. Women Shouldn't Be Primary B-Winner†	0.418	0.415	1.000			
V4. Mothers Who Don't Need Money Shouldn't Work†	0.286	0.281	0.468	1.000		
V5. Partners Should Equally Share Housework	0.180	0.139	0.105	0.033	1.000	
V6. Children Ok Female B-Winner Male Carer	0.275	0.316	0.442	0.173	0.245	1.000
Spearman Rank Correlations	**V1**	**V2**	**V3**	**V4**	**V5**	**V6**
V1. Men Make Better Political Leaders†	1.000					
V2. Bad For Relationship if Woman Earns More†	0.477	1.000				
V3. Women Shouldn't Be Primary B-Winner†	0.449	0.451	1.000			
V4. Mothers Who Don't Need Money Shouldn't Work†	0.314	0.309	0.475	1.000		
V5. Partners Should Equally Share Housework	0.246	0.230	0.155	0.079	1.000	
V6. Children Ok Female B-Winner Male Carer	0.328	0.387	0.491	0.215	0.309	1.000

Note: The table presents pairwise associations between the six gender attitudinal variables obtained from HILDA. All correlations are based upon the pooled sample. All estimates are significant at 5% (asterisks are suppressed for brevity). The upper panel gives correlation coefficients while the lower panel gives Spearman rank correlations, avoiding the cardinality assumption employed in the former calculations. † indicates that the variable has been inverted such that higher values indicate greater preferences for equality.

Table 14: Correlation Structure - Socioeconomic Equality Variables

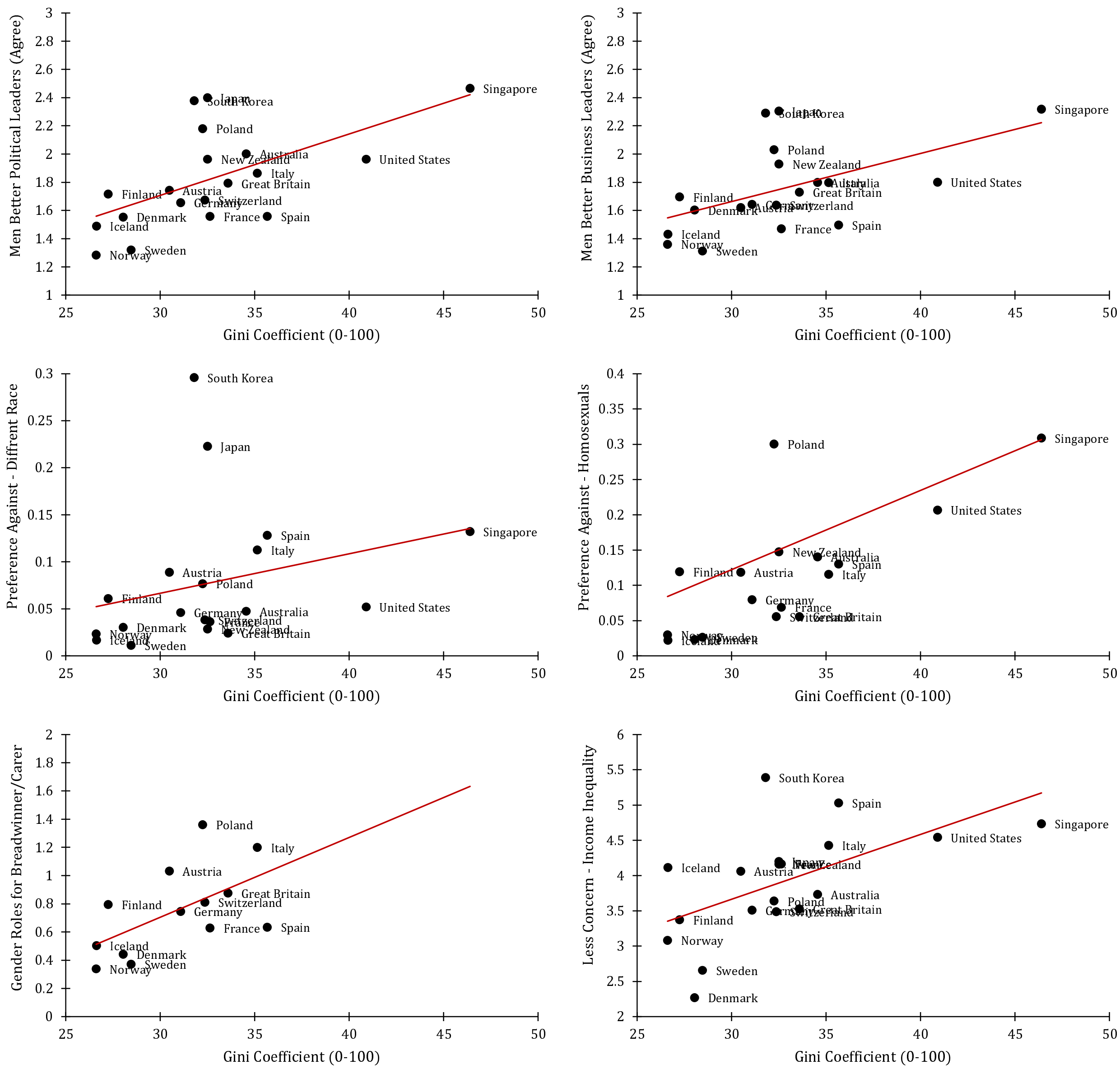
Pearson Correlations	V1	V2	V3	V4	V5	V6
W1. Men Make Better Political Leaders	1.000					
W2. Men Better Business Leaders	0.717	1.000				
W3. Racial Preference for Neighbors	0.172	0.169	1.000			
W4. Oppose Homosexuals for Neighbors	0.322	0.321	0.321	1.000		
W5. Men Should Be Breadwinners	0.562	0.530	0.182	0.358	1.000	
W6. See Combating Income Inequality as Important†	0.029	0.013	0.050	0.108	0.101	1.000
Spearman Rank Correlations	**V1**	**V2**	**V3**	**V4**	**V5**	**V6**
W1. Men Make Better Political Leaders	1.000					
W2. Men Better Business Leaders	0.724	1.000				
W3. Racial Preference for Neighbors	0.172	0.167	1.000			
W4. Oppose Homosexuals for Neighbors	0.316	0.314	0.321	1.000		
W5. Men Should Be Breadwinners	0.572	0.541	0.184	0.359	1.000	
W6. See Combating Income Inequality as Important†	0.025	0.009	0.055	0.115	0.101	1.000

Note: The table presents pairwise associations between the attitudinal variables taken from the WVS/EVS data. All correlations are based upon the pooled sample. All estimates are significant at 5% (asterisks are suppressed for brevity). The upper panel gives correlation coefficients while the lower panel gives Spearman rank correlations, avoiding the cardinality assumption employed in the former calculations. † indicates that the variable has been inverted such that higher values indicate greater preferences for equality.

Are these attitudes associated with objective economic disparities? To compare national-level aggregates on social attitudes with other markers of socioeconomic inequality, we also take data from the World Development Indicators (WDI) sourced from the World Bank. To measure income inequality we take WDI Gini Coefficients applied to annual incomes, and female political representation is used by the fraction of women in elected office. The idea here is that together, these two variables will act as barometers for economic and social disparities within a country. We then examine the associations between these objective outcomes and country-level averages of the six attitudinal markers presented in Table 14.

Figure 2 shows the relationships between the Gini coefficient and the mean response on each variable. The top two panels show links between two gender equality variables (preferences for male political and business leaders) and income inequality, while the middle two show the same for discriminatory views on the basis of race and sexual orientation. The final two panels give preferences for traditional breadwinner/carer roles and desires to lower income inequality. In all six cases, we see that countries with more (in)egalitarian attitudes also had more income inequality. E.g. Nordic and Western European countries have typically low scores on both metrics (i.e. they have both progressive outcomes on attitudes and lower income disparities) while Singapore and the United States had high corresponding values. The results seem to vary in intensity (e.g. the links between racial preferences and Gini coefficients are a little weaker than for gender) but the takeaway point here is their generality.

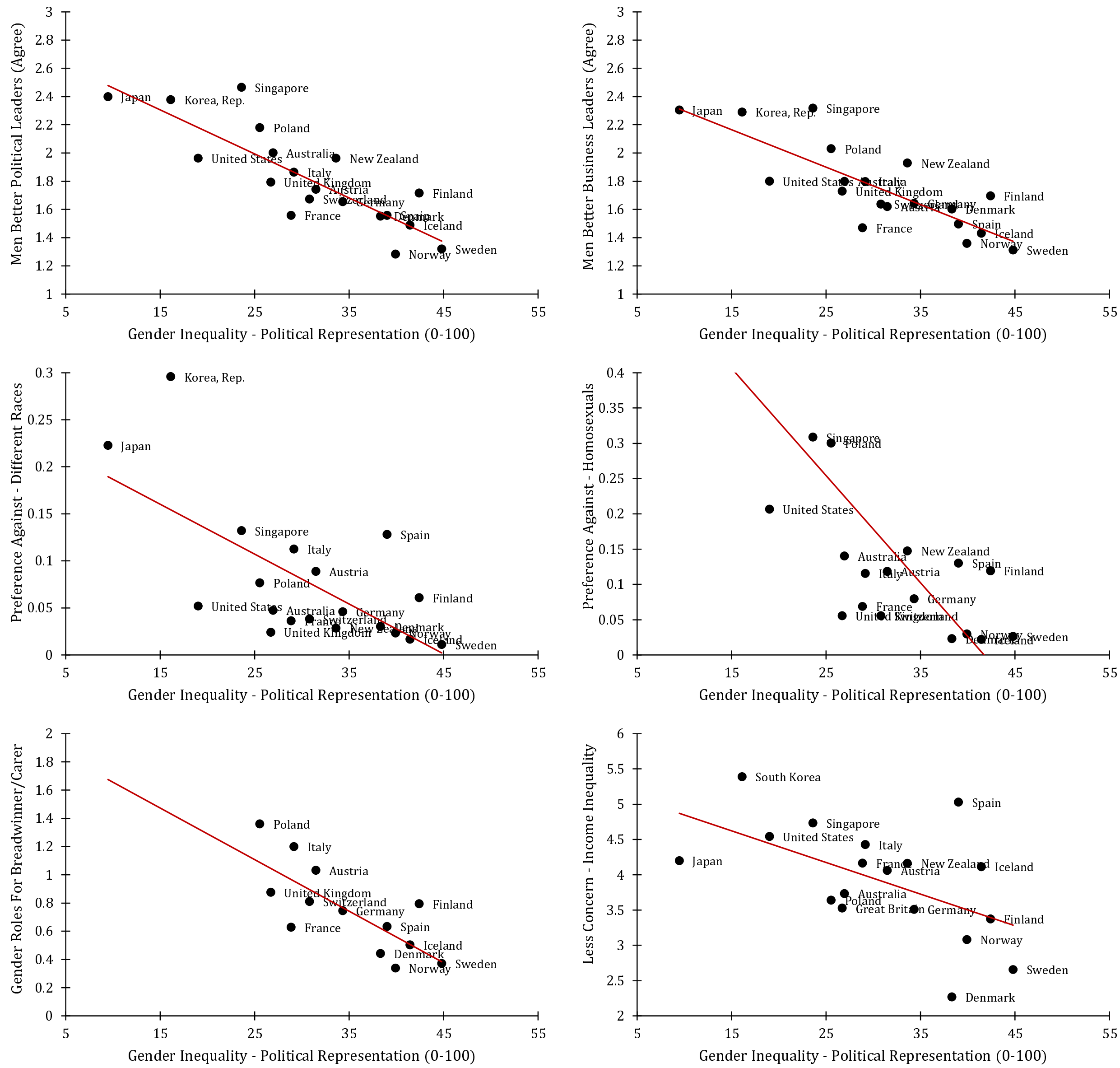
Figure 2: (In)Egalitarian Attitudes and Economic Inequalities - Country-Level Data



Note: The figure presents scatter plots between attitudinal variables from WVS/EVS averaged at the national level, and Gini coefficients sourced from the World Bank's WDI database and CIA Factbook. We restrict our data to developed countries that had observations in the WVS/EVS data since 2010 and also reported Gini coefficients and political representation data. Since the latter two variables tend to be volatile over time we use averaged values over a 10 year time horizon. A linear regression line fitted by OLS is also depicted in each plot.

Results for female representation in politics are given in Figure 3. The same ordering of attitudinal variables are used on the vertical axes, while the horizontal axes show the proportional representation (as a percentage) of elected officials. Here the associations are all negative, such that inegalitarian views predict lower female participation in politics (i.e. the same conceptual relationship in Figure 1). As above, there is some heterogeneity in the strength of the correlations, but overwhelmingly we see that attitudes are predictive of objective social inequalities.

Figure 3: (In)Egalitarian Attitudes and Economic Inequalities - Country-Level Data



Note: The figure presents scatter plots between attitudinal variables from WVS/EVS averaged at the national level, and political representation data sourced from the World Bank's WDI database. We restrict our data to developed countries that had observations in the WVS/EVS data since 2010 and also reported Gini coefficients and political representation data. Since the latter two variables tend to be volatile over time we use averaged values over a 10 year time horizon. A linear regression line fitted by OLS is also depicted in each plot.

Several bodies of literature suggest that the correlations depicted in Figures 2-3 at least partially reflect a causal flow from attitudes to outcomes. We consider the two possible causal paths outlined above - (i) that attitudes shape market outcomes in ways that directly exacerbate inequalities, and

(ii) in democratic countries (such as those in our sample), governments are accountable to a general public that may not support policy action.

The clearest example of path (i) is discrimination. If employers, educational institutions, legal systems and healthcare providers have preferences based on gender (or race or class), then this can be expected to impact upon a host of socioeconomic outcomes, including, but not limited to, income inequality. Empirical evidence of these effects includes Goldin and Rouse (2000) who examines gender-based labor market discrimination, and Lang and Lehmann (2012) who summarize a variety of findings related to race. While the effect sizes in these studies are sometimes small, a key point here is that different forms may be mutually self-reinforcing, in that inequalities in one dimension spill over to create inequalities in others. Further, these disparities may go on to be self-perpetuating, as inequalities may yield statistical discrimination, which in turn opens up additional future outcome gaps (Lang and Spitzer, 2020).

However, social norms can exert influence even when discrimination is not present. For example, beliefs about differing societal roles for men and women, racial minorities, and social classes, can shape individuals' life trajectories in ways that result in highly disparate outcomes. For instance, Akerlof and Kranton (2000) develop a theoretical framework showing that social expectations can augment utility functions and drive behavior. Similarly, Barigozzi *et al.* (2018) show empirically that social norms (e.g. expectations about career choice and childcare) have substantial influence over women's career decisions, and therefore contribute to gender wage inequalities. Perception of limited opportunity may also affect educational decisions and labor supply, and are correlated with factors such as race (Fouad and Byars-Winston, 2012) and social class (Zimmerman, 2020).[27] These effects are potentially large and may have ongoing ramifications. As an example, Chetty *et al.* (2016) show that children relocated to better neighborhoods have better long-term outcomes, an effect likely to be partially due to changing social expectations. And Guyon and Huillery (2020) show that (socially modifiable) aspirations are highly predictive of academic success, even once ability is controlled for, suggesting that "aspiration poverty traps" may be a source of ingrained disadvantage.[28]

Social attitudes may also affect inequalities through political channels (i.e. path (ii)). As outlined earlier, Washington (2008) and Oswald and Powdthavee (2010) have shown that familial birth composition affects voting patterns of both elected officials and the general public. Since left-wing policy prescriptions typically look to compress economic outcomes, these papers highlight direct links between the types of families individuals live in, and wider efforts to mitigate inequalities.

[27] An archetypal example is the "Wisconsin Model", where parental social class is shown to be predictive of intentions to pursue tertiary education while controlling for IQ (Sewell *et al.*, 1969).

[28] See Dalton *et al.* (2016) who note that disadvantage imposes extra costs upon individuals, which then lowers returns to efforts and thereby depresses outcomes. Since this effect results in further disadvantage, there is a feedback mechanism resulting in spiraling inequalities.

However, similar empirical links also exist with respect to more general attitudes. For example, cultural factors and social norms are a primary determinant of partisan affiliation (Leege *et al.*, 2002) which in turn affect social policy. Social attitudes can also influence policy (and therefore inequality) by moving the *Overton Window* - the range of political positions acceptable to the general public (Lehman, 2010). Here, redistributive proposals that fall outside this range are infeasible, and likely to be punished by voters if expressed (Caughey and Warshaw, 2019). Structural theories of political change therefore emphasize that the Overton Window must move as a precursor to substantive shifts in policy (Pierson, 2000).

7 Conclusion

This paper has studied the empirical links between the sociopolitical attitudes of parents and the gender of their children. We have added to a small body of literature showing that parents who have girls are more likely to hold gender-egalitarian attitudes than parents who raise boys. The effects we identify appear quite consistently across a range of indicators, such as whether men are viewed as superior political leaders, or are better suited to breadwinning roles. Nonetheless, there are some exceptions – we find no evidence of an effect in a couple of instances where the questions overlap with pragmatic issues concerning intra-household organization. Interestingly, our results only emerge when the children themselves reach young adulthood, which is consistent with close social interaction with politically mature individuals being the mechanism of change.

Stratifying our sample revealed stronger effects for men, older individuals, and those with lower educational attainments – groups that are usually known for holding relatively traditional social values. Further, we document that other small shifts in personality occur alongside these changes. Notably, parents of girls report lower levels of conscientiousness (a trait that that is more prevalent in women), which we attribute to compensatory behavior within households. It therefore appears that there is substantial interplay between the effects of raising daughters and the social and demographic characteristics of their families.

We focused in particular on issues related to causality. While the gender of a child could be thought of as random, which would allow for causal estimates, identification is more complex than it first appears. Since parents have a number of ways of consciously or unconsciously affecting the gender balance of their families, and may do so in ways related to their social views, these correlations have the capacity to bias (likely upwardly) econometric estimates. However, using a variety of recent diagnostic methods, we found little to no evidence that these factors could account for our results.

Lastly, we discussed the possibility that social attitudes may be a fundamental driver of some corrosive forms of inequality. For instance, we showed that views on gender equality are strongly associated with other egalitarian preferences, such as attitudes towards homosexuals and racial minorities. Further, these general egalitarian preferences are related to global patterns of tangible economic disparities, such as income inequality and female representation in politics. While our analysis in this section only presented correlations, it is likely that causal mechanisms where attitudes affect outcomes (through factors like discrimination or governmental policy) represent part of this story. Attitudinal change may therefore be a key input into efforts to redressing harmful economic inequalities.

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Appendix

A1. A Simple Model of Birth Composition and Parental Views

In this section, we provide a simple theoretical model outlining the fundamental ideas in the paper. The model is based loosely on Oswald and Powdthavee (2010) and has two working parts, (i) a trade-off between individualistic *vs.* egalitarian values that differs between males and females, and (ii) parents who take on some of the characteristics of their children when forming preferences. We use the model to explain why daughters can shift their parents leftward, and why these shifts are stronger for fathers.

Let $E \in R_+$ represent a society's preference for egalitarian vs individualistic values. $S(.)$ is the social value of public goods and $I(.)$ the return to individualistic endeavors. More egalitarian sentiment increases S (i.e. $S'(E) > 0$; $S''(E) < 0$) and reduces I (i.e. $I'(E) < 0$; $I''(E) < 0$). These terms are combined within a CES utility function. Since both are directly dependent upon E, we write $S(E) = E^\rho$ and $I(E) = (1-E)^\rho$, giving

$$V_M = \left[\lambda E^\rho + (1-\lambda)(1-E)^\rho\right]^{\frac{1}{\rho}} \qquad \lambda \in (0\,1)\,, \rho \in (0\,1) \tag{3}$$

Here, λ weights between social and individualistic outcomes, with higher values indicating greater egalitarianism. Relative to EQ (3), women have a greater preference for public goods, and experience discrimination that lowers their returns to individualistic endeavors. This is captured by $0 < \xi < 1$; $\xi + \lambda \leq 1$ such that $V_W = \left[(\lambda+\xi)E^\rho + (1-\lambda-\xi)E^\rho\right]^{\frac{1}{\rho}}$. Children also prefer higher values of E, which is captured directly through augmentation parameter $\delta > 0$. This generates four utility functions $\{$*Man* - M, *Woman* - W, *Son* - S, *Daughter* - $D\}$ and corresponding maximization rules:

$$V_M = [\lambda E^\rho + (1-\lambda)(1-E)^\rho]^{\frac{1}{\rho}} \to E_M^* = 1/\left[1 + (\lambda/(1-\lambda))^{1/(\rho-1)}\right] \tag{4}$$

$$V_W = [(\lambda+\xi)E^\rho + (1-\lambda-\xi)(1-E)^\rho]^{\frac{1}{\rho}} \to E_W^* = 1/\left[1 + ((\lambda+\xi)/(1-\lambda-\xi))^{1/(\rho-1)}\right] \tag{5}$$

$$V_S = [\lambda(E+\delta)^\rho + (1-\lambda)(1-E-\delta)^\rho]^{\frac{1}{\rho}} \to E_S^* = \delta + 1/\left[1 + (\lambda/(1-\lambda))^{1/(\rho-1)}\right] \tag{6}$$

$$V_D = [(\lambda+\xi)(E+\delta)^\rho + (1-\lambda-\xi)(1-E-\delta)^\rho]^{\frac{1}{\rho}} \to E_D^* = \delta + 1/\left[1 + ((\lambda+\xi)/(1-\lambda-\xi))^{1/(\rho-1)}\right] \tag{7}$$

We assume that parents maximize a convex combination of their own utility and that of their child. Let $\theta \in (0\,1)$ where $\theta > E_{min}$ denotes the degree to which a child's utility is incorporated. This gives four parental functions that can be constructed from the maximization equations above. We consider outcomes only for a single child, with EQ (8) to EQ (11) describing men and women, with sons and daughters $\{MD,\ WD,\ MS,\ WS\}$, respectively.

$$V_{MD} = (1-\theta)V_M(E;\lambda,\rho) + \theta V_D(E;\lambda,\rho,\delta,\xi) \tag{8}$$

$$V_{WD} = (1-\theta)V_W(E;\lambda,\rho,\xi) + \theta V_D(E;\lambda,\rho,\delta,\xi) \tag{9}$$

$$V_{MS} = (1-\theta)V_M(E;\lambda,\rho) + \theta V_S(E;\lambda,\rho,\delta,\xi) \tag{10}$$

$$V_{WS} = (1-\theta)V_W(E;\lambda,\rho,\xi) + \theta V_S(E;\lambda,\rho,\delta,\xi) \tag{11}$$

Result 1. Daughters Make Fathers More Egalitarian

Proof: Since $\theta \in (0\,1)$ this is established if $E_{MD}^*(\lambda,\rho,\delta,\xi) - E_M^*(\lambda,\rho) > 0$. These terms differ by ratios $\lambda/(1-\lambda)$ and $(\lambda+\xi)/(1-\lambda-\xi)$, and by $\delta > 0$. The effects are considered in turn. Since $(\lambda+\xi)/(1-\lambda-\xi) > \lambda/(1-\lambda)$ for $\lambda \in (0\,1)$, $\xi > 0; \xi + \lambda \leq 1$, $E_{MD}^*(\lambda,\rho,\delta,\xi) - E_M^*(\lambda,\rho) > 0$, holding λ,ρ,δ constant. Parameter $\delta > 0$ shifts $E_D^*(\lambda,\rho,\delta,\xi)$ but not $E_M^*(\lambda,\rho)$ such that $\delta + 1/\left[1 + ((\lambda)/(1-\lambda))^{1/(\rho-1)}\right] > 1/\left[1 + (\lambda/(1-\lambda))^{1/(\rho-1)}\right]$ holding λ,ρ constant.

Result 2. Sons Have Indeterminate Effects Upon Mothers

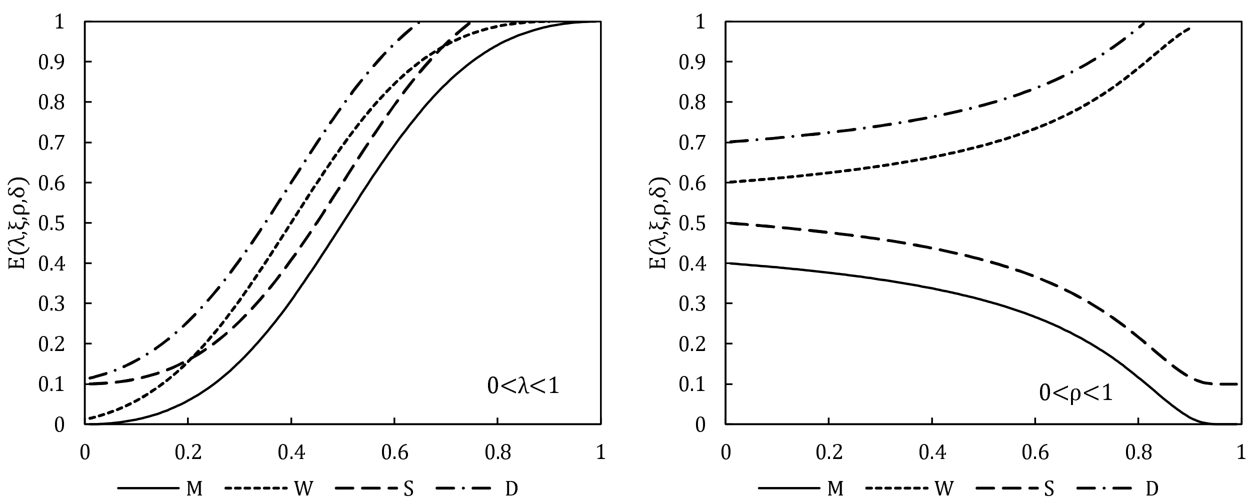
Proof: Since $\theta \in (0\,1)$ the *effect is captured by $E_W^*(\lambda,\rho,\xi) - E_S^*(\lambda,\rho,\delta)$ which can be positive or negative. As above, parameter $\xi > 0; \xi + \lambda \leq 1$ means that $(\lambda+\xi)/(1-\lambda-\xi) > \lambda/(1-\lambda)$ for $\lambda \in (0\,1)$, and therefore $E_W^*(\lambda,\rho,\xi) - E_S^*(\lambda,\rho,\delta) > 0$ holding λ,ρ constant and setting $\delta = 0$. Similarly $\delta + 1/\left[1 + ((\lambda)/(1-\lambda))^{1/(\rho-1)}\right] > 1/\left[1 + ((\lambda+\xi)/(1-\lambda-\xi))^{1/(\rho-1)}\right]$ holding λ,ρ constant and* *setting $\xi = 0$. The egalitarian effect dominates when $\delta\theta > \theta\left(1/\left[1+\left(\left(\lambda+\xi\right)/\left(1-\lambda-\xi\right)\right)^{1/(\rho-1)}\right]\right.$ $\left.-1/\left[1+\left(\left(\lambda\right)/\left(1-\lambda\right)\right)^{1/(\rho-1)}\right]\right)$.*

Result 3. Daughters Have Greater Egalitarian Effects on Fathers

Proof: This follows directly from R1 and R2. The incremental effect of a daughter for a father is proportional to $\theta\left[\delta+1/\left[1+\left(\left(\lambda+\xi\right)/\left(1-\lambda-\xi\right)\right)^{1/(\rho-1)}\right]-1/\left[1+\left(\lambda/\left(1-\lambda\right)\right)^{1/(\rho-1)}\right]\right]$. The corresponding effect for mothers is given by $\theta\left[\delta+1/\left[1+\left(\left(\lambda+\xi\right)/\left(1-\lambda-\xi\right)\right)^{1/(\rho-1)}\right]\right]$ $-\theta\left[1/\left[1+\left(\left(\lambda+\xi\right)/\left(1-\lambda-\xi\right)\right)^{1/(\rho-1)}\right]\right]=\theta\delta$, which is lower since $\xi > 0; \xi+\lambda \leq 1$.

Figure 4 illustrates the behavior of $E^*\left(\lambda,\xi,\delta,\rho\right)$ at some select values of these parameters. The left panel shows the four functions $\{$*Man*; *Woman*; *Son*; *Daughter*$\}$ over λ, defined with the restrictions $\xi=\delta=0$, $\xi=0$, $\delta=0$, $\xi;\delta\neq 0$. We see that $E^*_M < E^*_W, E^*_S, E^*_D$ for all λ, indicating that men without children hold the least egalitarian attitudes. Conversely $E^*_D > E^*_M, E^*_W, E^*_S$ such that daughters are always more egalitarian. Intersections occur between E^*_W and E^*_S as per R2, such that no unambiguous ordering can be established. The right panel presents the same exercise over ρ. Here there are no intersections and a complete hierarchy is obtained. Changes in substitutability therefore do not affect the ordinal relationships established on the left.

Figure 4: Behavior of Egalitarian Preferences in Families



Note: The left panel illustrates behavior of the functions E^*_M, E^*_W, E^*_S, and E^*_D over $\lambda > 0$ for fixed values of the other parameters $\rho = 0.5$, $\xi = 0.1$ and $\delta = 0.1$. Note that $E^*$is only defined on (0 1) and hence some values for λ are restricted. The right panel shows the same relationship over $\rho > 0$ using $\lambda = 0.4$, $\xi = 0.2$ and $\delta = 0.1$.

A2. Additional Data Description

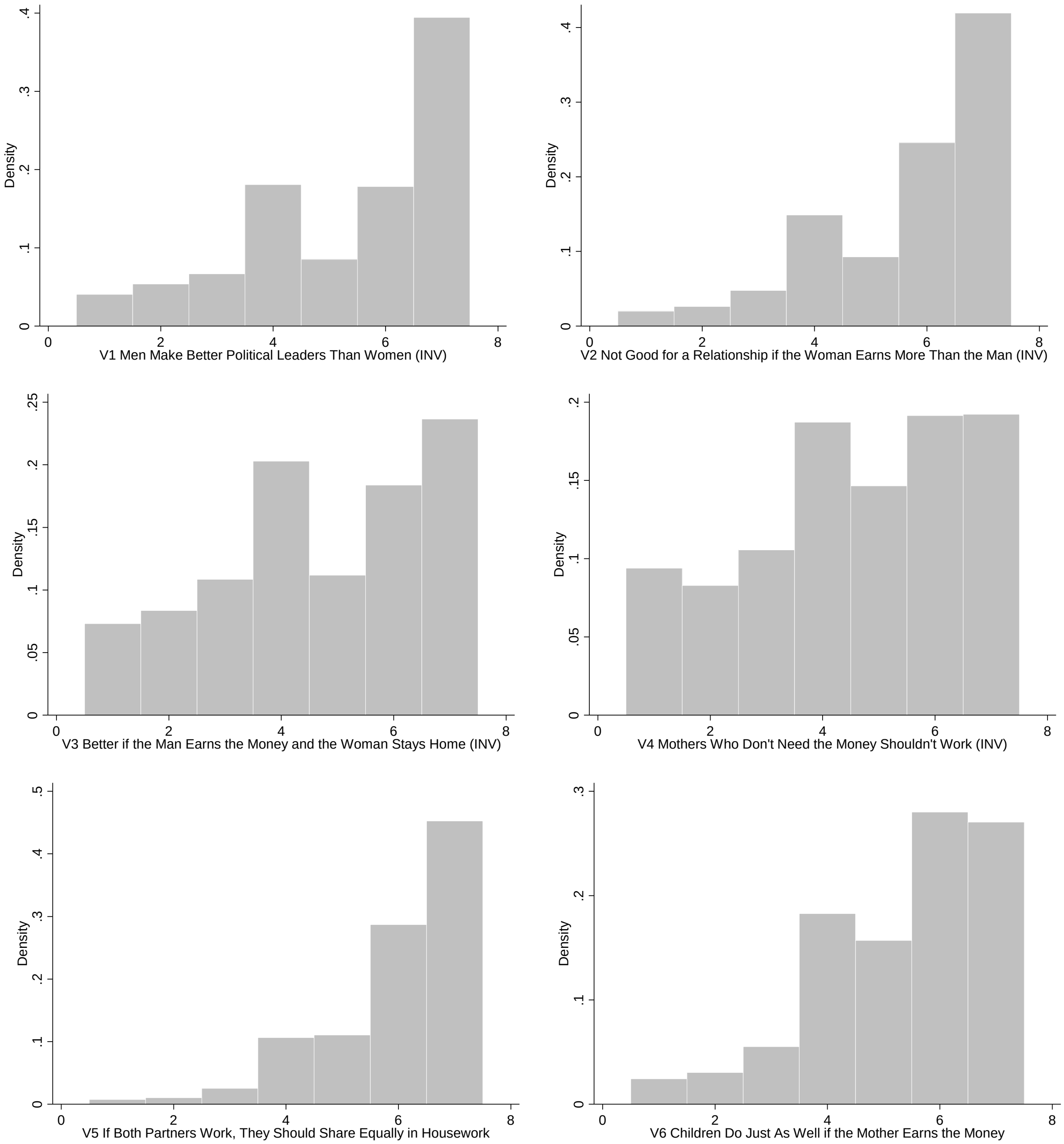
Our global data used in Section 6 are obtain from WVS/EVS (Inglehart *et al.*, 2017), which are approximately representative compatible national surveys on social attitudes. We draw additional data on income inequality and female representation in politics from the World Development Indicators and CIA World Factbook, and restrict our sample to developed countries that have observations since 2010. Consequently, data on certain countries (e.g. Canada) is unavailable, and occasionally there are missing values for variables W5 and W6. To handle instances where repeated observations occur over time (in either the WVS/EVS or WDI data) we produce longitudinal averages as country-level summaries. And since there is a change in methodology in sampling W6 across WVS and EVS (the EVS uses a 10 point scale while WVS uses 4 points) we take countries that appear in both surveys and calibrate the responses by matching the means and standard deviations in overlapping nations across survey formats. Descriptive statistics on average survey responses are presented below.

Table 15: Descriptive Statistics - EVS and WVS Samples

	Australia			Austria			Denmark			Finland		
	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$
W1	1,437	2.001	0.804	1,576	1.742	0.875	3,315	1.552	0.692	1,176	1.716	0.712
W2	1,457	1.798	0.720	1,584	1.621	0.805	3,307	1.603	0.722	1,176	1.696	0.708
W3	1,477	0.047	0.213	1,588	0.089	0.285	3,340	0.030	0.171	1,199	0.061	0.239
W4	1,477	0.140	0.347	1,588	0.118	0.323	3,340	0.023	0.150	1,199	0.119	0.324
W5	-	-	-	1,605	2.032	0.942	3,335	1.442	0.650	1,178	1.794	0.713
W6	1,447	3.733	2.708	1,604	4.062	2.386	2,939	2.268	2.655	1,167	3.374	2.447
	France			Germany			Great Britain			Iceland		
	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$
W1	1,786	1.558	0.769	2,111	1.655	0.698	1,748	1.792	0.658	1,609	1.489	0.633
W2	1,819	1.471	0.700	2,102	1.642	0.703	1,760	1.728	0.639	1,608	1.431	0.602
W3	1,848	0.036	0.187	2,085	0.046	0.210	1,782	0.024	0.153	1,551	0.017	0.128
W4	1,859	0.069	0.253	2,085	0.080	0.271	1,782	0.056	0.229	1,551	0.022	0.146
W5	1,853	1.629	0.892	2,138	1.745	0.778	1,771	1.875	0.764	1,603	1.503	0.668
W6	1,834	4.165	2.398	2,091	3.508	2.389	1,735	3.526	2.405	1,579	4.116	2.433
	Italy			Japan			New Zealand			Norway		
	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$
W1	2,200	1.863	0.788	1,587	2.399	0.724	750	1.963	0.717	1,116	1.283	0.652
W2	2,197	1.797	0.739	1,646	2.304	0.700	752	1.928	0.692	1,116	1.360	0.709
W3	2,149	0.113	0.316	2,443	0.223	0.416	841	0.029	0.167	1,115	0.023	0.151
W4	2,130	0.115	0.320	-	-	-	841	0.147	0.355	1,115	0.030	0.170
W5	2,226	2.199	0.885	-	-	-	-	-	-	1,122	1.339	0.723
W6	2,224	4.428	2.277	2,224	4.199	2.157	784	4.162	2.699	1,107	3.081	2.460
	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$
	Poland			Singapore			South Korea			Spain		
W1	1,229	2.180	0.745	1,972	2.465	0.723	1,186	2.377	0.832	1,175	1.558	0.708
W2	1,237	2.031	0.690	1,972	2.317	0.725	1,188	2.290	0.851	1,184	1.497	0.660
W3	1,239	0.077	0.266	1,970	0.132	0.339	1,200	0.296	0.457	1,194	0.128	0.334
W4	1,239	0.300	0.459	1,970	0.309	0.462	1,200	0.776	0.417	1,191	0.130	0.337
W5	1,309	2.361	0.869	-	-	-	-	-	-	1,198	1.634	0.806
W6	1,271	3.638	2.432	1,970	4.734	2.359	1,197	5.389	2.393	1,204	5.027	1.982
	Sweden			Switzerland			Taiwan			United States		
	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$	n	$\bar{x}$	$\hat{\sigma}^2$
W1	1,181	1.320	0.577	3,129	1.673	0.725	1,162	2.194	0.649	2,206	1.962	0.730
W2	1,178	1.312	0.558	3,131	1.638	0.715	1,188	2.185	0.684	2,191	1.799	0.680
W3	1,178	0.011	0.105	3,144	0.038	0.192	1,238	0.090	0.286	2,232	0.052	0.222
W4	1,178	0.026	0.160	3,141	0.055	0.229	1,238	0.426	0.495	2,232	0.207	0.405
W5	1,186	1.372	0.611	3,144	1.810	0.822	-	-	-	-	-	-
W6	1,167	2.657	2.649	3,118	3.486	2.476	1,185	4.872	2.659	2,192	4.541	2.540

Note: The table shows sample sizes, means and variances for the attitudinal variables W1-W6 obtained from the EVS/WVS data outlined in Section 6. Our sample is taken from 2010-2017 and only includes developed countries that also have Gini Coefficients and female political representation data. Variable W6 (importance in addressing income inequality) involves calibrating means and variances of overlapping countries across EVS/WVS. Estimates of $\bar{x}$ and $\hat{\sigma}^2$ assume cardinality in all cases.

Figure 5: Distributional Plots - Attitudinal Variables



Note: The figure depicts histograms of our six gender-attitudinal variables obtained from HILDA. The full variable descriptions are available in Section 2 and results are based upon the extract presented in Table 1. Where appropriate the variables are inverted such that higher values imply greater egalitarian sentiment.