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The 'mighty girl' effect: does parenting daughters alter attitudes towards gender norms?

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THE ‘MIGHTY GIRL’ EFFECT: DOES PARENTING DAUGHTERS ALTER ATTITUDES TOWARDS GENDER NORMS?

Abstract

Understanding the malleability of gender norms is crucial to address gender differences in the labour market and within households. We study the effect of parenting daughters on attitudes towards gender norms in the UK; more specifically, attitudes towards the traditional male breadwinner norm in which it is the husband’s role to work and the wife’s role to stay at home. We find robust evidence that parenting daughters decreases fathers’ likelihood to agree with a traditional gender division of work. This is especially the case for fathers of school-age daughters, for whom the effects are robust to a number of alternative specifications, and in particular the inclusion of individual fixed effects models. Our preferred estimates suggest that fathers’ probability to support traditional gender norms declines by approximately three percentage points (eight percent change) when parenting primary school-age daughters and by four percentage points (eleven percent change) when parenting secondary-school age daughters. In contrast, the effect of rearing daughters on mothers’ attitudes is smaller and generally not statistically significant. These findings are consistent with theories of exposure as well as (social) identity theories. We conclude that attitudes towards gender norms are not stable throughout the life-course and can significantly be shaped by experiences during adulthood.

Keywords. Gender norms, gender division of work, gender role attitudes, attitude formation, daughters, child gender.
JEL: J7, Z1.

1. Introduction

In the last decades, concerns about gender equality have been increasingly prominent in both the political and the social spheres, prompting governments to embark on the task of alleviating gender differences inside and outside the labour market. Nevertheless, progress towards achieving gender equality appears to have gradually slowed down (Gender Equality Index, 2017; Eagly and Wood, 2012). Against this background, a growing body of research has established the importance of traditional gender norms in explaining the persistence of gender inequalities in wages (Burda et al., 2007), in labour force participation (Fernandez et al., 2004; Fortin, 2005; Fernandez and Fogli, 2009; Farre and Vella, 2013; Johnston et al. 2014), and in the division of domestic work (DeMaris and Longmore, 1996; Greenstein, 1996, and see Davis and Greenstein 2009 for a review). However, there is limited evidence on how susceptible to change such norms are. This paper addresses this question.

Changing individual attitudes towards societal gender norms may be critical for further progress towards a less gendered division of work and towards gender equality more generally, since it may legitimize a wider range of social norms for both men and women (Eagly and Wood, 2012). So far, the literature has focused mostly on long-term changes in norms across cohorts (Baxter et al., 2015), although research on individual changes in attitudes towards gender norms across the life cycle is gradually increasing. This latter approach has studied the role played by the family environment, including that of marriage, parenthood, and women's labour patterns inside and outside the household (see interalia Cunningham, 2008; Baxter et al., 2015, Schober and Scott, 2012; Clarkberg, 2002; Corrigall and Konrad, 2007).

Our paper contributes to this literature by analysing one life course event that has received limited attention so far, namely the effect of parenting daughters. Using a British nationally representative longitudinal survey spanning two decades, we examine whether rearing daughters changes parental attitudes towards gender norms, and more specifically, attitudes towards the traditional male breadwinner norm in which it is the husband's role to work and the wife's role to stay at home. Given that a child's gender cannot be anticipated, we assume that rearing a daughter – as opposed to a son – is an approximately random event (Washington, 2008).[1]

In examining the individual change in attitudes[2] towards gender norms, we borrow the definition of gender norms from Pearse and Connell (2016), who define them as 'collective definitions of socially approved conduct in relation to groups constituted in the gender order – mainly distinctions between men and women'. Hence, norms are defined as 'features of a collective life' (Pearse and Connell, 2016, p. 34) that signal to other members of a group or

[1] However, some authors such as Hamoudi and Nobles (2014) found that relationship conflict between husband and wife predicted the sex of subsequent children, and hence, separate analyses need to be done looking at the effect of the first child only.

[2] By attitudes we understand *evaluations* of objects, behaviour, events, or people as good or bad; they vary on a positive/negative scale and can be expressed by statements such as "I like/dislike" or "I agree with/disagree with" (Schwartz, 2012; Bicchieri, 2017). Therefore, attitudes towards gender norms are individual evaluations of these gender norms; and a positive attitude would reflect an endorsement of the collective gender norm in the society.

society how they *should* behave' (Schwartz, 2012, p. 16, emphasis added), and closely follows definitions in other social science disciplines[3].

We find that parenting daughters significantly decreases the likelihood that fathers will agree with a traditional male breadwinner norm. This is especially the case for fathers of school-age daughters, for whom the effects are robust to a number of alternative specifications, and especially individual fixed effects models that control for time-invariant unobserved heterogeneity. Our fixed effects estimates suggest that fathers' probability to support traditional gender norms declines by approximately three percentage points (eight percent change) when parenting primary school-aged daughters and by four percentage points (eleven percent change) when parenting secondary school-aged daughters. In contrast, the effects on mothers are smaller and not statistically significant. While it is not possible to discern the exact mechanisms through which daughters affect parental attitudes, the heterogeneity of results between fathers and mothers combined with the finding that attitudinal change occurs only when daughters reach school age is in line with theories of exposure as well as with identity theories. Furthermore, given that attitudes towards gender norms are shaped by experiences during adulthood, our results also indicate some evidence of intra-cohort change in attitudes. Consistent with our findings on attitudinal change, we find that parenting school-age daughters is also associated with a lower likelihood that couples follow a traditional gender division of work.

To our knowledge, this is the first paper to explore the impact of child gender *across daughters' age* on individual changes in attitudes towards gender norms. This is important because our findings suggest that it is when daughters are of school-age – and not before – that father's attitudes become less traditional, thus coinciding with the period in which children experience a stronger social pressure to conform to gender norms (Lane et al., 2017). The paper also contributes to expanding the evidence beyond the United States, being the first paper to explore the impact of the gender of the child on attitudes towards gender norms in the UK. Finally, and unlike previous studies, we draw on data that covers very recent years – up to 2011 – which is important given the sharp change in patterns of gender inequalities during the past decades.

The structure of the paper is as follows. The next section reviews the relevant literature, and section three describes the data and empirical strategy. Section four contains the main results, section five robustness checks, and a final section concludes.

2. Related literature

On the malleability of attitudes towards gender norms

There are two main approaches in social science on the evolution of attitudes towards social norms, including gender norms. One approach suggests that attitudes are formed *before*

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[3] For example, feminist economist Stephanie Seguino uses a definition which is very close to that of Pearse and Connell, describing them as 'acceptable behavioural boundaries for men and women, congruent with the gender division of labour and male power' (Seguino, 2007, p. 2). Similarly, psychologist Alice Eagly uses the term *injunctive norms* to speak about 'consensual expectations about what a group of people ought to do or ideally would do' (Eagly and Karau, 2002, p. 574). The concept of gender norms is related but different to that of stereotypes, understood as 'consensual expectations about what members of a group actually do' (Eagly and Karau, 2002), or gender ideologies, used to 'justify the gender imbalance in power and resources' (Seguino, 2007).

reaching adulthood and remain stable thereafter[4]. Societal change in norms then occurs through processes of cohort succession, when older cohorts are replaced by younger ones who systematically differ in its social and historical early years' experiences (Mannheim, 1952; Brooks and Bolzendahl, 2004). An alternative approach – the one embraced by this paper - questions the stability of norms and embraces the viewpoint that *attitudes can change over the life course*[5], either due to social structural changes, or due to changes in individual circumstances (Baxter et al., 2015; Brooks and Bolzendahl, 2004; Hogg and Vaughan, 2008).

Empirical evidence concerning these two approaches remains inconclusive. Two early papers analysing US data from the 1970s and the 1980s point at different directions, one suggesting that attitudes towards familial roles in the US occur mainly within cohorts (Mason and Lu, 1988) and another providing evidence for cohort replacement-based explanations (Wilkie, 1993). Further analyses with data from the 1990s and early 2000s have not resolved the debate. While some research supports cohort-replacement theories (Brewster and Padavic, 2000), there is also evidence which confirms the importance of intra-cohort change (Danigelis et al., 2007). Another paper finds that while cohort-replacement theories have a strong explanatory power, ideological influences and conflicts experienced by individuals in adulthood mediate their attitudes (Brooks and Bolzendahl, 2004).

Attention to intra-cohort change has recently shifted the focus of research towards the potential factors underpinning change, with a particular emphasis on family environment. To this purpose, longitudinal data has been increasingly used to study the impact on attitudes towards gender norms of women's decision to work (Cunningham, 2008), parenthood (Baxter et al., 2015; Evertsson, 2013), the interaction between work and childbirth (Schober and Scott, 2012; Berrington et al., 2008), and marriage and cohabitation (Corrigall and Konrad, 2007; Moors, 2003). Nonetheless, the gender of the child has received limited attention. In what follows, we will focus on the specific effect of child gender.

On the relevance of child gender for attitudes towards gender norms

Evidence on the effect of child gender on attitudes towards gender norms remains inconclusive.[6] Warner's pioneer study (1991) showed that daughters led Canadian and US parents to hold more modern attitudes towards gender norms (with the exception of American men). She explains her finding with the inclusion of children's wellbeing into parents' own utility function, consistent with (social) identity theory. The logic behind (social) identity theory is that individuals derive utility from behaving in line with the social roles and the social categories they identify with (see e.g. Hogg et al., 1995; Akerlof and Kranton, 2000). Consistently, mothers with sons are more likely to hold traditional views on gender norms (Downey et al., 1994)[7].

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[4] See for example the increasing persistence hypothesis (Glenn, 1980; Inglehart and Baker, 2000) or the impressionable years hypothesis (Carlsson and Karlsson, 1970; Ryder, 1965; Krosnick and Alwin, 1989).

[5] See for example the life-stages hypothesis (Sears, 1983; Visser and Krosnick, 1998) or the lifelong openness hypothesis (Brim and Kagan, 1980).

[6] In this paper, we focus on attitudes towards traditional gender norms as a dependent variable. A few papers from various disciplines have studied the effect of child gender on a range of other outcomes, among them parents' political party identification (Oswald an Powdthavee, 2010; Lee and Conley, 2016), voting behaviour on women's issues (Washington, 2008; Glynn and Senn, 2015) labour supply (Lundberg and Rose, 2002; Lundberg, 2005; Pabilonia and Ward-Butts, 2007; Choi et al., 2008), take-up of parental leave (Bartel et al., 2015), support for gender equality policies (Warner and Steel, 1999), and CEO's wage setting (Dahl et al., 2012). For an overview, see Lundberg (2005) and Raley and Bianchi (2006).

[7] Their explanation, however, is that sons are more valued than daughters and are in need for more protection, inducing mothers to think that working would be a disservice to them.

Nonetheless, the external validity of those studies is limited due to small and unrepresentative samples sizes. In contrast, Shafer and Malhotra (2011) use a large sample (the National Longitudinal Study of Youth 1979 from the US) and find that only fathers reduce their support for traditional gender roles when having a daughter, while mothers' attitudes remain unchanged. The intuition behind Shafer and Malhotra's finding is that mothers, unlike fathers, may have already crossed the 'threshold of exposure' by experiencing situations in their own lives that render them more sympathetic to a modern attitude towards gender norms. Having a daughter would expose fathers to a larger extent than mothers to new worldviews, leading to a more significant shift in their attitudes. This evidence is in line with theories of exposure, which posit that individuals - parents, in this case - 'develop or change their understanding of women's place in society (...) when they encounter ideas and situations that resonate with feminist ideals (Bolzendahl and Myers, 2004). However, their study cannot test the plausibility of exposure-related explanations, since they analyse the effect of childbirth, as opposed to the effect of parenting daughters of different age groups. This is an important issue we address in this paper.

The significant effect of child gender on attitudes is challenged by two papers. Katzev et al. (1994) use the National Survey of Families and Households and find that, against their expectations, mothers with boys are more likely to hold modern attitudes towards marriage and family life. Similarly, Conley and Rauscher (2013), using the 1994 General Social Survey, find no evidence (for any parent) of daughters promoting non-traditional views and gender norms. The null effect would not necessarily contradict exposure-related theories given that if fathers exhibit a son preference[8], they might spend less time with daughters and as a result be less exposed to their worldviews (Lee and Conley, 2016). However, both studies are limited by either the older nature or the cross-section dimension of the data, resulting in their findings potentially being affected by time-invariant unobservables. The longitudinal nature of our data should help overcome some of these limitations.

An additional advantage of large longitudinal evidence is that it allows for the examination of age specific interactions, which may play an important role. According to (social) identity theories, the event of a daughter's birth could be enough to trigger a re-adjustment of the parental utility to include their daughter's interests. However, it is possible that such re-adjustment only takes place once daughters are older and the effect of gender norms on the child's behaviour becomes more visible. Research in psychology suggests that children are aware of gender stereotypes already at the age of six (Bian et al., 2017), with social pressure to conform to existing gender norms mounting around the early-adolescence period (Lane et al., 2017). In line with theories of exposure (see Bolzendahl and Myers, 2004; Glynn and Sen, 2015), parents may thus start to become more aware of what is at stake for their daughters when they reach school-age, prompting a re-adjustment of their gender norm attitudes around this time (and not earlier). The rest of the paper will be devoted to understand empirically how the gender of the child influences parental attitudes across child age.

[8] Dahl and Moretti (2008) found a son preference in the US using data from the 1960s to the 2000s. This son preference is less evident in more recent US data (see Blau et al., 2017).

3. Data and empirical strategy

We use rich longitudinal data from the British Household Panel Survey (BHPS) together with the BHPS sample of the UK Household Longitudinal Study (UKHLS).[9] The data are a nationally representative sample of British households collected annually since 1991. Attitudes towards the gender division of work – our outcome variable of interest – were collected biennially, hence our data comprise eleven distinct waves covering the period from 1991 to 2012.

3.1 Sample

We restrict the sample to individuals who we observe with a child under the age of 21 in the household in at least one wave, thus excluding individuals who never have children in the household as well as individuals who are only observed after their children in the household are already in their twenties or older. Since we follow individuals over time and are interested in the variation from children entering the household and not the variation from children leaving the household, we follow individuals only as long as the number of children in the household is increasing or constant.[10] After dropping observations for which the main outcome variable or any of the covariates are missing, the final sample is an unbalanced panel including 48,822 observations of 11,405 individuals (5,073 men with 20,851 observations and 6,332 women with 27,971 observations).

3.2 Measures

Our dependent variable measures attitudes towards traditional gender norms, specifically towards a traditional gender division of work. Respondents are asked to rate agreement with the statement 'a husband's job is to earn money; a wife's job is to look after the home and family' on a five-point scale ranging from 'strongly agree' to 'strongly disagree'. We recoded the answers so that higher values mean stronger agreement with the statement, hence more traditional attitudes.[11] We also binarize the dependent variable into non-traditional (taking a value of zero for 'disagree' or 'strongly disagree') and traditional (taking a value of one for 'neither agree nor disagree', 'agree', and 'strongly agree') and report results alongside those for the ordinal dependent variable.[12]

Our main regressor of interest is the binary variable 'at least one daughter' that takes on a value of one if the respondent has at least one daughter living in the household, and zero otherwise.[13] To account for the potential interactive effect between child gender and child age on attitudes to gender norms, we also distinguish different age groups: 'daughter 0 to 5', 'daughter 6 to 10', and 'daughter 11+' are dummy variables indicating whether there is at least one daughter of the respective age group living in the household. If there is more than one daughter, we consider the age of the youngest daughter, but we also run a robustness check using age of the oldest daughter. We also refer to these age groups as pre-school-aged daughters, primary school-aged daughters, and secondary school-aged daughters, and to refer to the latter two at the same time we say school-aged daughters.

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[9] The BHPS was collected annually from 1999 to 2008 and became part of the new UKHLS, which started in 2009 and is still ongoing.

[10] We replicated the main results of section four without this further sample restriction and results remained similar.

[11] We focus on this attitudinal question as it captures the essence of the traditional male breadwinner norm, by which it is the husband's role to work and the wife's role to stay at home.

[12] We also ran the main results (Tables 1 and 2) when considering "neither agree nor disagree" as non-traditional and results remained similar, albeit with generally smaller effect sizes.

[13] In a robustness check in Table A.6, we use alternative ways to represent child gender.

Our data contain information on children living in the household (as opposed to fertility records) which we use to construct our daughters measures, and we equally consider natural children, adopted children, foster children, and partners'/step-children.[14] Since we are interested in the longitudinal effect of *parenting* daughters via exposure or identity changes rather than the one-time event of a birth of a daughter, we consider information from co-residence preferable over that from fertility histories.

3.3 Estimation approach and identification assumptions

We are interested in identifying the effect of parenting daughters on parental attitudes towards gender norms, and start by estimating the following pooled OLS baseline specification:

$$y_{it} = \alpha + \beta_1 d_{it} + \beta_2 (d_{it} * fem_{it}) + \beta_3 fem_{it} + \beta_4 c_{it} + \beta_5 X_{it} + \beta_6 T_t + \varepsilon_{it} \quad (1)$$

where *y* stands for the level of agreement with traditional gender norms of individual *i* at time *t*, fem is a dummy variable indicating the individual is female, *c* are dummy variables controlling for the total number of children in the household, *X* are a set of individual characteristics, and *T* are wave fixed effects. *d* is our key regressor for parenting daughters and takes two forms. Once a dummy with a value of one if the individual has at least one daughter in the household, and zero otherwise. In a second specification testing age-of-daughter specific effects, *d* is a categorical variable with the four categories 'no daughter', 'at least one daughter with the youngest aged zero to five years', 'at least one daughter with the youngest aged six to ten years', and 'at least one daughter with the youngest aged at least eleven years'. Standard errors are clustered by individual.

Thus, we estimate the effect of parenting at least one daughter, while holding family size constant. When the number of children is controlled for, the coefficient on the daughter dummy captures the additional effect of parenting daughters as opposed to parenting only sons. In all models, we interact the effect of parenting daughters with a dummy for females, thus allowing the effect of parenting daughters to vary by gender of the parent. We then start adding control variables in a stepwise fashion: first, by adding the 'basic controls' age, square of age, as well as region and wave fixed effects. Then, by additionally adding the 'additional controls', which consist of two educational level dummies, five marital status dummies, eight employment status dummies, and ln household income (see Table A.1 for summary statistics and definitions of all variables).

While estimates from pooled OLS regressions are comparable to previous literature that uses cross-sectional data, we also estimate individual fixed effects (FE) models with robust standard errors, in order to eliminate bias arising from unobserved individual heterogeneity and to capture the effect of daughters on changes in attitudes within individuals longitudinally rather than comparing those with to those without daughters.[15] Hence, the error term in equation (1) takes the form:

$$\varepsilon_{it} = \mu_i + \rho_{it}. \quad (2)$$

[14] 94.4 percent of our children-wave pairs are on natural children, 4.8% on step/partners' children, 0.7% on adopted children, and 0.2% on foster children. If anything, we expect our results to be downward biased compared to considering only natural children.

[15] For ease of interpretation we estimate linear models, but our results remain very similar when we re-estimate the main results for the binarized dependent variable with logit and fixed effects logit models (see Table A.9).

We follow previous literature (e.g. Washington, 2008; Oswald and Powdthavee, 2010) in considering the gender of any given child entering the household as approximately random[16]. However, we still face potential endogeneity because of three reasons: First, fertility stopping rules may depend on the sex mix of children already in the household, hence our key regressor of having at least one daughter conditional on the total number of children may be selected. Indeed, evidence suggests that parents in Western countries including the UK are more likely to have a third child if their first two children are of the same sex, as they prefer a 'balanced' sex mix of children (Lundberg, 2005; Iacovou, 2001). There is also some evidence suggesting that a first-born boy increases the probability of further children (Ichino et al., 2014). This implies that only the gender of a firstborn child is truly random, and if more (less) traditional parents have a son (daughter) preference, this would introduce an upward bias to our OLS estimates. To account for these possibilities, we perform robustness checks in which we test for endogenous fertility stopping rules. We also estimate the effect of only first daughters on attitudes, and we re-estimate the main results on the subsample of observations with two or less children.

Second, there is potential selection into co-residence with daughters because a father's or mother's decision about co-residence may depend upon whether they have daughters or sons. For example, after couples split, we typically observe the resident parent with marital status 'divorced', while the non-resident parent drops out, i.e. attrition from our estimation sample[17]. Then, if the likelihood to divorce or to get child custody depends on child gender, this may bias our results.[18] Therefore, this essentially becomes a problem of attrition and we perform robustness checks testing whether having daughters is related to attrition and whether attrition affects our results.

Finally, we also check that our results are not driven by reverse causality, that is, that initial attitudes predict the probability of having a daughter. In addition to performing these robustness checks, we note that fixed effects account for the bias arising from time-invariant unobservable characteristics that are correlated with both the probability to live with daughters and with attitudes. For our fixed effects estimates to still be biased, it would be necessary that the timing of daughters entering the household is systematically correlated with shocks causing attitudinal change. Or alternatively, that individuals with higher malleability in attitudes are more likely to live with daughters versus sons. We argue that this is unlikely.[19]

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[16] This is based on the argument that parents cannot choose the sex of any given child, absent sex-selective abortion. Previous literature has regarded the possibility of sex-selective abortion as empirically unimportant in Western countries (e.g. Choi et al., 2008).

[17] This happens due to one of two reasons, either because they attrit from the panel due to the inability to track or retain fathers in the study after the family splits, or because we drop observations on non-resident fathers as we follow individuals only as long as the number of children in the household is increasing or constant. We replicated the main results reported in section four, keeping the observations for which variation came from children leaving the household, and results remained similar.

[18] However, there seems to be no evidence of an association between child gender and divorce or custody arrangements in Western countries other than the US (Lundberg, 2005) and we are not aware of any UK study investigating this.

[19] In line with this, Amato and Booth (1991) find that individuals in the US who were divorced at baseline held less traditional gender role attitudes, while getting divorced during the duration of the panel was not associated with changes in attitudes. We are, however, not aware of any UK evidence on this issue.

3.4 Sample descriptive statistics and randomization checks

Table A.1 reports summary statistics for the analysis sample. The average age of observations for men and women in our sample is 37.5 and 35.4 years, respectively. The average number of children is 1.62 and 1.66, respectively. The dummy of having at least one daughter in the household takes a value of one for 57 percent of observations on men and 58 percent of observations on women. None of the various variables on daughters in the household (at least one daughter, only daughters, first daughter) suggest that there are substantial differences between fathers and mothers in their probability to co-reside with daughters.

Figure 1 examines how attitudes towards the gender division of work differ by respondent sex and by the sex composition of offspring for individual-wave pairs with at least one child in the household. It shows the mean value of the ordinal (panel a) and binarized (panel b) outcome variable. On average, men have higher levels of agreement with the traditional male breadwinner norm than women. Men with at least one daughter or only daughters are less traditional than men with only sons. Women with daughters only are less traditional than women with sons only, but those with at least one daughter actually appear more traditional compared to those with only sons. To understand whether these differences are explained by other covariates requires further analysis.

- Figure 1 about here -

In Table A.2 we check whether our key regressor of parenting at least one daughter is related to any socio-demographic characteristics. In panel *a*, each row shows the coefficient of a separate regression of the probability to have at least one daughter on the respective socio-demographic characteristic, while controlling for the total number of children. Most of the individual characteristics are unrelated to the probability of parenting daughters, with a few exceptions. For females, having a higher degree is positively associated with parenting daughters, while being widowed/other marital status, being retired, and being in government training/having another employment status is negatively associated with parenting daughters. For men, being divorced is associated with a lower probability to be parenting daughters. To account for these differences, we control for marital status, employment status, and educational level. Panel *b* tests whether all socio-demographic characteristics shown in panel a, plus region and wave fixed effects, can predict having at least one daughter while holding family size constant. The F-test rejects joint significance for both the male and the female subsample. This supports our assumption of no selection into having at least one daughter, conditional on family size.

4. Main results

Our results are presented in Tables 1 to 3. Table 1 estimates the effect of parenting daughters on attitudes towards the traditional gender division of work. Table 2 subdivides this effect into parenting daughters of three different age groups. Finally, Table 3 studies whether parenting daughters also affects couple's probability of following a traditional gender division of work.

- Table 1 about here –

4.1 Baseline Results

Table 1 presents the effect of parenting daughters on levels of agreement with the traditional male breadwinner norm, using the continuous attitudinal scale of the survey as well as a binarized version to capture potential 'attitudinal shifts'[20]. In all specifications we include dummy variables for the total number of children in the household, hence we identify the effect of parenting at least one daughter while holding family size constant.[21] We allow the effect to vary by gender of the parent as we interact the daughter dummy with a female dummy, and we report the direct effect of parenting daughters for mothers via linear combination of estimates.

We find that having daughters is associated with lower levels of agreement with traditional gender norms among men, but not among women as reflected in column (1). Next, we add age, square of age, wave, and region fixed effects in column (2) and we observe that results are robust to such controls. In column (3) we additionally control for education, marital status, employment status, and the natural logarithm of household income. Importantly, the negative association between parenting daughters and traditional attitudes remains after controlling for these observable characteristics. Results hold too when attitudes are binarized in column (4), and we show that men with daughters are approximately three percentage points less likely to hold traditional attitudes compared to men without daughters, while holding family size constant. This coefficient reflects an eight percent reduction in the probability to hold traditional attitudes (approximately 37.1 percent of male observations without daughters in our sample hold traditional attitudes).

Given that these results could be explained by individual specific and time invariant unobservables that affect both co-residence with daughters and attitudes, columns (5) to (8) exhibit the individual fixed effects results and show that the negative association between parenting daughters and traditional attitudes persists in FE models. Comparing the FE columns with their respective OLS ones reveals that the size of the coefficient is approximately halved when accounting for time-invariant unobserved heterogeneity. Once we introduce the 'additional controls' for education, marital status, employment status, and ln household income in columns (7) and (8), the coefficient on the daughter dummy becomes statistically insignificant although the sign remains.

We then turn to examine attitudes among women. Women hold on average less traditional attitudes compared to men. The interaction effects between parenting daughters and being female show that the effect of parenting daughters is significantly different for mothers compared to fathers. The linear combinations of estimates capture the overall effect of parenting daughters among women. OLS estimates (columns (1) to (4)) show that parenting daughters is also negatively associated with traditional attitudes among women, the effect is smaller compared to the one for men and statistically significant only in two of the columns. When individual fixed effects are included, the coefficient on the daughter dummy is statistically insignificant in all specifications and the sign even turns positive for the specifications with ordinal dependent variable (columns (5) to (7)).

Taken together, the results suggest that having daughters is associated with lower levels of agreement with traditional gender norms among men. For women, the association is ambiguous. Once controlling for the full set of observable characteristics and time-invariant

[20] In columns (1) to (3) and (5) to (7) the dependent attitude variable is ordinal, while in columns (4) and (8) it is binarized.

[21] We only report the coefficients on the variables of key interest. For full results showing coefficients on all control variables see Table A.3.

unobservable characteristics in columns (7) and (8), effects are not statistically significant. This suggests that, on average, attitudes do not change with the birth of a daughter.

- Table 2 about here –

4.2 Age-of-daughter specific effects

Given that rearing a daughter might not exert an immediate effect on attitudes upon birth but instead may emerge after a certain exposure, we next turn to examining age-of-daughter specific effect of rearing daughters on parental gender norm attitudes (Table 2). We separate our previous dummy on parenting daughters into three different dummy variables: at least one daughter of age group 0 to 5, at least one daughter of age group 6 to 10, and at least one age group of 11 or older.[22] We refer to them in our paper as pre-school, primary school, and secondary school age daughters. Consistently with Table 1, the omitted category is thus 'not parenting any daughters'.

The OLS estimates suggest that parenting daughters of each age group is associated with less traditional gender attitudes among fathers, but the effect is larger for daughters at school-age (age groups 6 to 10 and 11 plus) rather than pre-school, which is consistent with exposure-based explanations. As expected, when we control for individual fixed effects (columns (5) to (8)), the coefficients on pre-school daughters and primary school daughters shrink. In the full specification in column (7), the coefficient on daughters aged 0 to 5 becomes insignificant, suggesting that on average, attitudes do not change at birth. In contrast, results suggest that parenting daughters of primary and secondary school age makes fathers on average less traditional, with a larger effect size for secondary school age daughters. Consistently, results for the binarized outcome variable in column (8) indicate that parenting pre-school daughters does not affect fathers' attitudes, while parenting primary school age daughters reduces the probability to hold traditional attitudes by about three percentage points, which constitutes an eight percent change compared to the baseline probability of holding traditional attitudes of 37.1 percent among men without daughters. Parenting secondary school age daughters reduces the probability to hold traditional attitudes by approximately four percentage points, which amounts to an eleven percent change compared to men without daughters.

As before, we the turn to examining the effects for mothers. The interactions between the different daughter dummies and the female dummy show that in many cases, the effect of parenting daughters is significantly different for mothers. When we test the joint significance of the linear combination of estimates, OLS estimates in columns (1) to (4) suggest that parenting pre-school and primary school daughters is associated with less traditional attitudes among mothers. However, once we introduce individual fixed effects the effect of parenting daughters on attitudes becomes insignificant for all age groups, while the sign of the coefficients turns positive for the younger two age groups.

Taken together, results from Table 2 suggest that after accounting for time-invariant unobserved heterogeneity, there is no robust effect among mothers. For fathers, attitudes to gender norms do not change with the birth of a daughter but instead, fathers' attitudes become significantly less traditional when parenting school-age daughters.

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[22] If there is more than one daughter, we define the age group based on the youngest daughter in the household, but we perform robustness checks in which we define the age group based on the oldest daughter in the household (see Table A.6).

- Table 3 about here –

4.3 Effects on behaviours

Given these interesting results for men, we next explore whether parenting daughters also changes behaviour reflecting the male breadwinner norm as it would be expected if attitudes inform behaviours (Table 3). To do so, we focus on the subsample of males living in a couple, and we define a variable which takes the value of one if the individual is employed or self-employed and the partner is neither of the two, and a value of zero otherwise.

We find that on average, having daughters does not change the probability to follow a male breadwinner norm (column (1) for OLS and (3) for FE). Studying different age groups again yields a more differentiated picture though. Parenting pre-school daughters is associated with a higher probability to behave traditionally. However, parenting primary and secondary school age daughters is associated with a lower likelihood to follow a traditional male breadwinner norm in which the man works and the woman does not work, and this results holds both cross-sectionally and longitudinally. In terms of effect size, fixed effects estimates shown in column (4) indicate that parenting daughters aged 6 to 10 reduces the probability of a traditional gender division of work by seven percentage points, and parenting daughters aged eleven or older reduces that probability by five percentage points. Compared to the baseline probability of following a traditional norm for those without daughters of 20.3 percent, this is a sizeable reduction of 36 and 25 percent, respectively. Therefore, *our finding that parenting daughters changes attitudes to gender norms is also reflected in behavioural changes concerning gender norms, for the subsample of individuals living with a partner.*

5. Robustness checks

Given the results in section 4, we next examine the robustness of our main findings in a number of different ways, presented in Tables A.4 to A.10. We generally include (but do not report) the full set of covariates, so the estimates are comparable to those in columns (3) (OLS) and (7) (FE) of Tables 1 and 2.

5.1 Endogenous fertility and reverse causality

We start by exploring in how far our results are affected by endogenous fertility stopping rules. In panel a of Table A.4, we examine whether the sex of the first child (columns (1) and (3)) and the sex mix of the first two children (columns (2) and (4)) in the household predict the total number of children an individual has in the last wave he or she is interviewed. To do so, we construct a collapsed cross-sectional dataset in which we only keep one observation per individual and summarize information from different survey waves. We analyse the male and female subsamples separately. We find that there is a negative but not statistically significant association between having a first daughter and the total number of children for both men and women.[23] However, we do find a negative and statistically significant correlation between having at least one daughter among the two oldest children, and the number of total children, for the subsample of individuals who have at least two children by

[23] The key regressor labelled ‘ever1stdaughter’ takes a value of one if in at least one survey wave the oldest child is female, and zero otherwise. The dependent variable is the total number of children in the wave with the most children, which coincides with the last wave the individual was interviewed.

the last interview wave.[24] This implies that those who have two sons are more likely to have further children compared to those with at least one daughter and hence there is positive selection of having at least one daughter. This is in line with previous research which shows that parents have a preference for having offspring with a balanced sex mix (Lundberg, 2005; Iacovou, 2001). To make sure this is not driving our results, we examine the effect of the gender of the oldest child only (Table A.6) and we analyse the subsample of individual-wave pairs with two or less children only (Table A.7), and we find that our results are robust to these additional checks.

Specifically, in Table A.6, we examine the robustness of our findings from Table 2 to three alternative specifications of the key regressors. These are dummy variables taking a value of one if there are only daughters of the respective age group in the household (columns (1) and (2)), if the oldest child in the household is female and in the respective age group (columns (3) and (4)), and if there is at least one daughter with the age group based on the oldest daughter (columns (5) and (6)), respectively, and zero otherwise. We find that the results from these alternative specifications are very similar to those obtained in Table 2. Overall, having daughters is negatively and significantly associated with less traditional attitudes among men, and the effect for secondary school-age daughters is robust to the inclusion of individual fixed effects. For women, the negative and mostly insignificant association between having daughters and traditional attitudes becomes positive for the two younger age groups and generally insignificant with FEs. The coefficients indicating the effect among fathers are largest for parenting only daughters and smallest when we only consider the gender of the first child, consistently with an exposure-based explanation. In sum, the robustness of our main results to considering only the gender of the oldest child suggests that they are not driven by endogenous fertility stopping rules. Moreover, we argue that it would be difficult for an individual to choose having only daughters, having at least one daughter, and having a firstborn daughter, all while holding family size constant. Therefore, results from Table A.6, when taken together, provide further support that our findings are not driven by endogenous fertility stopping rules.

Another way to test that our results are not driven by endogenous fertility stopping rules based on the sex mix of the first two children is to examine results for the subsample of individual-wave pairs with two or less children (columns (1) to (4) of Table A.7). Results for the subsample are very similar to the main results, providing a further indication that endogenous fertility stopping rules are not the driver behind our findings.

In panel *b* of Table A.4, we test for the possibility of reverse causality. We run regressions of ‘ever having at least one daughter’ on initial gender attitudes in the first wave the individual was interviewed, controlling for the full set of covariates including the total number of children. We find that initial attitudes are not associated with the probability of ever having a daughter while holding family size constant and hence, there is no evidence of reverse causality.

5.2 Selection into co-residence after divorce and attrition

As detailed in section 3, when parents divorce, the resident parent will typically be observed as divorced while the non-resident parent will drop out of our estimation sample, i.e. there is attrition from the sample. Hence, potential selection problems at family dissolution are

[24] The key regressor is a binary variable taking a value of one if there is at least one daughter among the two oldest children in the household when the individual was last interviewed, and zero if they are both sons.

essentially problems of potentially non-random attrition in our analysis. In Table A.5, we again use the collapsed cross-sectional dataset, to test for bias arising from attrition. Panel *a* shows that having a daughter in at least one interview wave is not associated with attrition from the estimation sample, that is, the probability of dropping out of the sample before the last wave. In panel *b*, we test for attrition bias in our results by performing a BGLW test.[25] To do so, we estimate the effect of ever having daughters on initial attitudes to gender norms. By comparing results using the total estimation sample to results using the sample of individuals who do not drop out of the sample before the last interview wave (i.e. the non-attrition sample), we examine how different coefficients would be if only the non-attrition sample was used. We find that ever having a daughter is generally unrelated to initial attitudes. P-values from tests of differences in coefficients between the two samples (columns (3) and (6)) suggest that there are no significant differences in coefficients between the total and the non-attrition samples. This also holds when we look at the daughter dummy differentiated by age groups. Hence, we can conclude that our coefficients are unaffected by attrition.

As a final check for selection into co-residence after divorce, and related attrition, in Table A.7 we re-estimate the main results for two additional subsamples: individuals who are never observed as divorced (columns (5) to (8)), and individuals who are not dropped from the estimation sample before the last survey wave (columns (9) to (12)). For both subsamples, results are very similar to the main results. The negative effect of parenting daughters (independent of age group) even becomes statistically significant at the 10 percent level with fixed effects (columns (7) and (11)). In line with the results from the BGLW test in Table A.5, there is no indication that family separation through divorce and related disappearance from our estimation sample drives our results.

5.3 Further robustness checks

Table A.8 checks the robustness of our results to an alternative dependent variable, which measures levels of agreement with the statement that "both the husband and wife should contribute to the household income". While this variable relates less directly to attitudes towards the gender division of work, it still captures attitudes towards gender norms more generally. We interpret higher levels of agreement with the statement as more gender-equal and hence less traditional attitudes. In contrast to our main outcome variable, the pooled OLS estimates indicate that both men and women with pre-school daughters are more traditional compared to those without daughters. Further, parenting secondary school-age daughters makes women less traditional. However, our main findings are reflected with this alternative outcome variable: parenting daughters makes men less traditional, the effect is strongest for fathers of school-age daughters, and there are no statistically significant effects for mothers that are robust to the introduction of individual fixed effects. Furthermore, we find that the main results with the binarized dependent variable (see columns (4) and (8) of Tables 1 and 2) are robust to estimating logit and fixed effects logit models (Table A.9). Finally, we conduct a placebo test in which we create a random variable of a 'fake' daughter to check that our results do not just pick up some other trend that occurs in people's lives around the time when they have school-age children.[26] We then re-run our OLS and individual FE models, and report results in Table A.10. As expected, there is no statistically significant

[25] Such a test is described in Fitzgerald et al. (1998) and was first used by and named after Becketti et al. (1988).
[26] We randomly assign a value of zero or one to create the placebo variable of a 'fake' daughter; if a value of one was assigned, then it is changed to one in all subsequent waves of the same individual. For those individuals with a placebo daughter, we randomly assign the age group in the first wave the daughter appears. After we observe that individual five times, we switch to the next oldest age group.

effect of the placebo variables on attitudes to the gender division of work. This holds for both males and females, the OLS and FE estimates, and the overall placebos as well as the 'age group' placebos.

6. Discussion

Understanding the malleability of gender norm attitudes is key to tackling gender inequality at its roots. Against the backdrop that attitudes start to form early in life, we show that parenting daughters decreases fathers' likelihood to agree with a traditional gender division of work while there is no robust effect on mothers. This finding is in line with previous US findings (Shafer and Malhotra, 2011) supporting the idea that having a daughter affects men and women differently. More importantly, we find that the effect on fathers' attitudes occurs when daughters who are of school age, which is a novel finding. We carefully check that our results are not driven by unobserved individual heterogeneity, endogenous fertility stopping rules, reverse causality, or attrition from the estimation sample. Finally, our findings on attitudes are also reflected in behavioural changes concerning gender norms, for the subsample of individuals living with a partner.

We build on the few papers which have examined the effect of child gender on attitudes towards gender norms in several ways. Not only are we one of few studies to provide longitudinal evidence and the first study to show UK evidence and to cover recent years, but we also enrich the literature by showing that there is an age-specific effect of daughters on fathers' gender norms attitudes.

While we are unable to test underlying mechanisms, the differential effect of parenting a daughter between fathers and mothers together with the finding of an age-specific daughter effect on attitudes is consistent with theories of exposure as well as with identity theories. Consistent with these theories, both father and mother may incorporate their children's wellbeing into their own utility function. Through parenting, fathers of daughters develop a better understanding of women's and girls' disadvantages in society, resulting in a significant shift in their attitudes towards gender norms. Conversely, mothers have already been exposed to situations of disadvantage first-hand, and as a consequence, parenting a daughter has a negligible effect on their attitudes towards gender norms, which are already less traditional than that of men. The timing of exposure - when daughters are of school-age - is in line with research in psychology which suggests that children become aware of gender stereotypes and social pressures to conform to gender norms around this age (Bian et al., 2017; Lane et al., 2017). Hence, fathers are likely to gradually become aware of the gender norms affecting their daughters' actions after that age, prompting the change in their gender norms attitudes. Our findings also provide evidence for theories of intra-cohort change in attitudes. Attitudes towards gender norms seem to be malleable to experiences during adulthood such as parenting a daughter, thus suggesting that indirect exposure to disadvantage has the potential to change people's attitudes.

Subsequent research could expand our findings in several ways. First, it could explore the mechanisms underlying our results and systematically test the plausibility of exposure theory with the help of time use data on father-daughter interactions. Second, while our paper focuses mainly on attitudinal change, we also found evidence of behavioural change, with couples parenting daughters being less likely to conform to a male breadwinner model. Were this finding to be confirmed by further research, the gender of the child should be considered as a potential source of heterogeneity for household division of labour. Third, our finding raises questions related to the consequence of indirect exposure to disadvantage. While having a daughter seems to make the father more likely to turn against gender norms that may disadvantage her, the same may not be necessarily true of other family members such as the spouse, let alone of non-family members. For example, Sands (2017) shows that exposure to socioeconomic inequality by racial and economic outgroups decreases support for redistribution. Further research should focus on the role of social identity in mediating the effect of indirect exposure to disadvantage or inequality.

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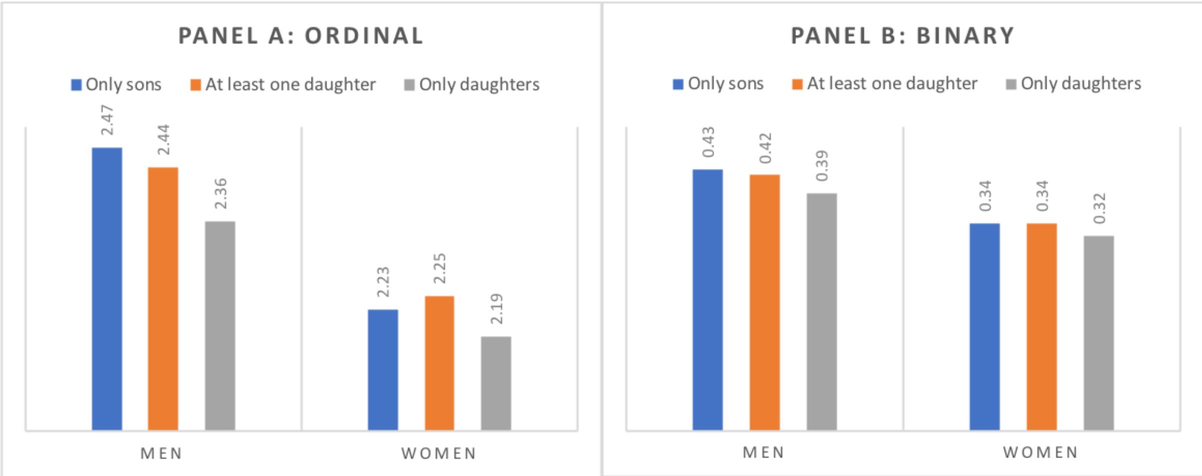
Tables

Figure 1: Mean attitudes to traditional gender division of work



Notes: sample includes individual-wave pairs with at least one child in the household, by gender of offspring. Panel a shows mean levels of agreement with the statement 'husband should work and wife stay at home' (scale 1 to 5). Panel b shows the share of observations with 'traditional attitudes' when the dependent variable is binarized.

Table 1

The effect of parenting daughters on attitudes to gender norms

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
		Pooled OLS				Fixed Effects		
		ordinal dep. variable		binary		ordinal dep. variable		binary
At least one daughter	-0.113***	-0.119***	-0.0930***	-0.0310***	-0.0547**	-0.0571**	-0.0415	-0.0102
	(0.0255)	(0.0249)	(0.0242)	(0.0115)	(0.0270)	(0.0271)	(0.0271)	(0.0140)
Female	-0.276***	-0.263***	-0.295***	-0.117***				
	(0.0235)	(0.0231)	(0.0223)	(0.0102)				
At least one daughter x female	0.0803***	0.0963***	0.0495*	0.00629	0.0919***	0.0905***	0.0626**	0.00912
	(0.0294)	(0.0290)	(0.0280)	(0.0130)	(0.0300)	(0.0300)	(0.0302)	(0.0156)
Linear combination of estimates: effect for females								
At least one daughter	-0.0330	-0.0228	-0.0435**	-0.0247**	0.0373	0.0334	0.0212	-0.00111
	(0.0227)	(0.0223)	(0.0210)	(0.00976)	(0.0257)	(0.0257)	(0.0256)	(0.0133)
Total number of children controls	yes	yes	yes	yes	yes	yes	yes	yes
Basic controls		yes	yes	yes		yes	yes	yes
Additional controls			yes	yes			yes	yes
Observations	48,822	48,822	48,822	48,822	48,822	48,822	48,822	48,822
R-squared	0.041	0.061	0.106	0.075	0.006	0.009	0.013	0.010
Number of individuals					11,405	11,405	11,405	11,405

Notes: Robust standard errors in parentheses (clustered by individual for columns (1) to (4)). *** p<0.01, ** p<0.05, * p<0.1. Total number of children controls: 4 dummies for number of total children. Basic controls: age, age squared, wave dummies, region dummies. Additional controls: 2 education dummies, 5 marital status dummies, 8 employment status dummies, ln household income.

Table 2
The effect of parenting daughters of different age groups on attitudes to gender norms

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Pooled OLS				Fixed Effects			
	ordinal dep. variable			binary	ordinal dep. variable			binary
At least one daughter (youngest 0 to 5)	-0.158***	-0.107***	-0.0659**	-0.0165	-0.0468*	-0.0474*	-0.0295	-0.00313
	(0.0277)	(0.0276)	(0.0271)	(0.0129)	(0.0274)	(0.0274)	(0.0275)	(0.0142)
At least one daughter (youngest 6 to 10)	-0.162***	-0.144***	-0.120***	-0.0480***	-0.0703**	-0.0790**	-0.0625*	-0.0281*
	(0.0308)	(0.0301)	(0.0296)	(0.0144)	(0.0322)	(0.0324)	(0.0324)	(0.0170)
At least one daughter (youngest 11 plus)	-0.0375	-0.116***	-0.103***	-0.0348**	-0.106***	-0.124***	-0.110***	-0.0410**
	(0.0307)	(0.0307)	(0.0297)	(0.0141)	(0.0344)	(0.0359)	(0.0359)	(0.0187)
Female	-0.276***	-0.263***	-0.296***	-0.118***				
	(0.0235)	(0.0232)	(0.0223)	(0.0103)				
Daughter 0 to 5 x female	0.0926***	0.0818**	-0.00739	-0.0222	0.0900***	0.0887***	0.0519*	0.00299
	(0.0336)	(0.0332)	(0.0322)	(0.0150)	(0.0307)	(0.0307)	(0.0309)	(0.0159)
Daughter 6 to 10 x female	0.0986***	0.116***	0.0811**	0.0216	0.0952**	0.0920**	0.0821**	0.0218
	(0.0368)	(0.0364)	(0.0353)	(0.0168)	(0.0370)	(0.0371)	(0.0372)	(0.0195)
Daughter 11 plus x female	0.0540	0.0976***	0.0832**	0.0235	0.0905**	0.0862**	0.0851**	0.0205
	(0.0366)	(0.0361)	(0.0347)	(0.0164)	(0.0418)	(0.0420)	(0.0421)	(0.0216)
Linear combination of estimates: effect for females								
At least one daughter (youngest 0 to 5)	-0.0653**	-0.0252	-0.0733***	-0.0387***	0.0432*	0.0413	0.0225	-0.000139
	(0.0255)	(0.0256)	(0.0242)	(0.0110)	(0.0259)	(0.0260)	(0.0259)	(0.0134)
At least one daughter (youngest 6 to 10)	-0.0632**	-0.0276	-0.0389	-0.0264**	0.0248	0.0130	0.0195	-0.00634
	(0.0270)	(0.0267)	(0.0253)	(0.0119)	(0.0299)	(0.0302)	(0.0301)	(0.0156)
At least one daughter (youngest 11 plus)	0.0166	-0.0187	-0.0195	-0.0113	-0.0151	-0.0374	-0.0249	-0.0205
	(0.0264)	(0.0268)	(0.0254)	(0.0120)	(0.0332)	(0.0351)	(0.0350)	(0.0177)
Total number of children controls	yes	yes	yes	yes	yes	yes	yes	yes
Basic controls		yes	yes	yes		yes	yes	yes
Additional controls			yes	yes			yes	yes
Observations	48,822	48,822	48,822	48,822	48,822	48,822	48,822	48,822
R-squared	0.042	0.061	0.106	0.076	0.006	0.010	0.014	0.010
Number of individuals					11,405	11,405	11,405	11,405

Notes: Robust standard errors in parentheses (clustered by individual for columns (1) to (4)). *** p<0.01, ** p<0.05, * p<0.1. Total number of children controls: 4 dummies for number of total children. Basic controls: age, age squared, wave dummies, region dummies. Additional controls: 2 education dummies, 5 marital status dummies, 8 employment status dummies, ln household income.

Table 3

The effect of parenting daughters on the traditional gender division of work (male subsample)

	(1)	(2)	(3)	(4)
	Pooled OLS		Fixed Effects	
	Traditional gender division of work			
At least one daughter	-0.0172		0.00003	
	(0.0118)		(0.0194)	
At least one daughter (youngest aged 0 to 5)		0.0699***		0.0122
		(0.0137)		(0.0196)
At least one daughter (youngest 6 to 10)		-0.0653***		-0.0724***
		(0.0141)		(0.0218)
At least one daughter (youngest 11 plus)		-0.0849***		-0.0507**
		(0.0139)		(0.0248)
Observations	18,144	18,144	18,144	18,144
R-squared	0.064	0.078	0.077	0.082
Number of individuals			4,889	4,889

Notes: Robust standard errors in parentheses (clustered by individual for columns (1) to (2)). *** p<0.01, ** p<0.05, * p<0.1. Control variables in all columns: 4 dummies for number of total children, age, age squared, 2 education dummies, 5 marital status dummies, ln household income, wave dummies, region dummies. Outcome variable: Dummy individual (self-)employed and partner not (self-)employed.