Family Norms and the Gender Turnout Gap

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Abstract

Gender differences in participation endure in many democracies even though socio-economic differences between men and women have narrowed and the number of female political figures has increased. We contribute to explaining this persistence by showing the highly gendered effects of family norms on the opportunity costs and social benefits of voting. To this end, we estimate the impact of changes in family composition–i.e., marital status and childrearing–on turnout among male and female voters using a unique administrative panel dataset from Italy. Our difference-in-difference impact estimates reveal that marriage equalizes turnout among spouses, increasing men's participation, and leaving women's unchanged. By the same token, young children reduce maternal turnout, leaving paternal turnout unchanged. Exploring potential mechanisms, we suggest that increased well-being among men and peer pressure within couples are the likely sources of the marriage effect, whereas gender imbalance in family chores allocation is the source of the effect of parenthood.

In the decades following the extension of suffrage, women were initially less willing to participate in elections than men (Teele, 2018). In recent years, though, levels of female turnout have increased and female voters now outnumber male voters in several advanced democracies including Norway and the U.S. (Kittilson, 2016; Quaranta and Dotti Sani, 2018). These trends have important implications for the political representation of women, who hold policy preferences that often diverge from men's and are more likely to elect female politicians (Alvarez and McCaffery, 2003; Teele, Kalla and Rosenbluth, 2018). Yet, a significant gender turnout gap–i.e., a lower propensity to vote among women compared to men–endures in some segments of the electorate. Specifically, while in most western democracies young women are more likely to vote than men of the same age, this difference tends to narrow over the life-course, and it often reverses among citizens aged 55 or older (see Figure 1 below; Quaranta and Dotti Sani, 2018; Bratsberg et al., 2019). What explains the persistence of gender differences in voter turnout?

The extant literature highlights the role of two different processes. A first strand of work argues that older generations of women may not have benefited from recent improvements to education and employment opportunities that have increased female turnout among younger generations of voters (Leighley and Nagler, 2013; Bratsberg et al., 2019). This suggests that the current gender gap may be driven by a cohort effect. A second strand of work contends that women's political engagement is higher when there are more female candidates and political figures and indicates that the socializing effects of political context are strongest among young women (Wolbrecht and Campbell, 2007; Dassonneville and McAllister, 2018). This also suggests that more time needs to elapse for the gender gap to evanesce. Yet, the literature documents similar patterns of persistence across democracies with different levels of female education, labor force participation, and descriptive representation (Leighley and Nagler, 2013; Bratsberg et al., 2019). Thus, it appears that gender differences in turnout may endure *even though* socio-economic differences between men and women have narrowed and the number of female representatives has increased.

We address this puzzle by proposing a third perspective. Our approach is related to recent work emphasizing the role of social context or "culture" (Dassonneville and Kostelka, 2019; Robinson

and Gottlieb, 2019). For example, the gender turnout gap correlates cross-nationally with the math achievement gap between boys and girls (Dassonneville and Kostelka, 2019). However, the extant literature is unclear about whether gendered social norms have a *causal* effect on political participation. To begin studying this important question, we focus here on how family socialization affects the opportunity costs and the social benefits of voting over the life-cycle.

We discuss theoretical reasons to expect that changes in family status–marriage, divorce, widowhood, and children of various ages–have differential turnout effects across genders. First, family norms can structure access to participatory resources to the disadvantage of women (Verba, Burns and Schlozman, 1997; Iversen and Rosenbluth, 2006). In particular, we contend that time constraints on females, on account of socially mandated household work, may increase the opportunity costs of political participation. Second, life-cycle transitions tend to improve men's relative physical and emotional well-being compared to women's and affect a variety of social indicators that are related to voter turnout (Wolfinger and Wolfinger, 2008; Burden et al., 2017). Third, changes in family status may alter how the social norm to vote is perceived and enforced (Stoker and Jennings, 1995; Dellavigna et al., 2017). We hypothesize that peer pressure may have positive effects on men's turnout, whereas the potential conflict between the civic duty to vote and other social expectations related to family duties may decrease electoral participation among women.

Our research design allows us to test these theoretical predictions and provides a rare opportunity to clearly discern gendered patterns of participation from other correlated factors like education and income. To this end, we leverage a unique individual-level panel dataset covering four elections in a large Italian municipality, which we build by merging three sources of administrative data: voter rolls, the civil register, and income tax files. Importantly, our data track the same individuals over a period of nine years (2004-2013) and contain complete information about turnout, gender, and family status. This enables us to identify within-individual effects through a series of differences-in-differences (DD) designs. Moreover, we exploit the longitudinal dimension of the data to examine the persistence of turnout effects and to (indirectly) test the parallel-trend assumption underlying our DD strategy. Specifically, we use event-study graphs to check whether "causes

happen before consequences"(Granger, 1969); that is, we verify that changes in voter turnout follow changes in family structure, and not vice versa.

We find significant and sizeable turnout effects of marital status and childrearing, and document considerable heterogeneity across genders. Our models show that the transition from never-married to married has a positive effect on voter turnout ($+1$ p.p.), whereas the transition from marriage to widowhood reduces political participation (-2 p.p.), and divorce has no effect. While these results are broadly in line with prior empirical studies (e.g., Wolfinger and Wolfinger, 2008; Hobbs, Christakis and Fowler, 2014), we contribute to the literature by showing that these effects are highly gendered: the positive effect of marriage is fully driven by male voters, whereas the widowhood effect is entirely attributable to female voters and appears to pre-date the spouse's death–potentially as a consequence of the deterioration of the spouse's health. We also examine the effects of children of different ages on parental voter turnout and uncover important gender differences. In particular, young infants (0-5 years old) induce a sizable drop in maternal turnout (-2 p.p., on average) that persists for several years, leaving paternal turnout unchanged. Interestingly, we observe that among unmarried parents–who plausibly hold progressive gender attitudes–the birth of a child has no significant effect on maternal turnout but has a positive effect on paternal turnout, which mirrors the mobilizing effect of marriage. Together, these results provide credible evidence that transitions in family status skew turnout patterns to the disadvantage of women.

In the last part of the paper, we collect additional evidence on possible mechanisms by leveraging nationally-representative survey data. First, we find no association between marriage and political knowledge. Heightened political interest, therefore, does not explain the higher electoral engagement of married men found in our estimates. This comports with the hypothesis that marriage equalizes men's turnout to the higher pre-marriage voter participation of their spouses (Stoker and Jennings, 1995). Second, while men with kids are more politically knowledgeable than their childless counterparts–which is consistent with our finding that turnout increases among fathers of school-aged children–this is not the case for women. Third, women hold more left-leaning policy stances than men *irrespectively* of their life-course stage, which raises the concern that women's

children-induced political disengagement undermines the representativeness of the voting electorate. Finally, our survey data confirm that the presence of young children sharply increases the number of total hours worked by women (both in absolute terms and relative to men), boosting the time they dedicate to household work and only partly reducing time of paid labor, which is consistent with our finding that marriage and young children correlate with men's income increasing relative to women's. This suggests that equal participation of women in politics would require men and women to share household responsibilities more equally (Silbermann, 2015).

While our results comport with prior work showing that traditional family norms in Italy prescribe strong gender imbalance in time devoted to household work (Alesina and Ichino, 2009; Anxo et al., 2011), they also indicate that these family obligations may interfere with the civic duty to vote, which in northern Italy is rooted in longstanding civic traditions (Putnam, Leonardi and Nanetti, 1994). Thus, studying gender turnout differences in this context may on its own provide interesting theoretical insights. However, we also discuss reasons to believe that our findings hold more generally. Specifically, barriers to voting are low in northern Italy and we find no effect heterogeneity by distance to the polls, thus suggesting that the convenience of voting can only take us so far in explaining the gender turnout gap. Moreover, levels of female labor force participation and pre-kindergarden childcare are higher in our setting than in other advanced democracies, on average, which should make it more difficult to find gendered turnout effects.

Conceptual Framework

Political scientists have long theorized that family acts as a source of interpersonal mobilization and that these effects vary between men and women (e.g., Wolfinger and Rosenstone, 1980). However, most prior studies focus only on one particular aspect of family status–like marriage, widowhood, or having kids (Wolfinger and Wolfinger, 2008; Stoker and Jennings, 1995; Hobbs, Christakis and Fowler, 2014; Dahlgaard, 2018). As a result, little is known about the overall consequences for the gender composition of the electorate.

Here we provide an explicit conceptual framework combining both resource-centered and moti-

vational theories of how gender roles affect voter turnout. We argue that changes in family composition can create different situational constraints for female and male voters and also differentially affect their level of civic engagement. Taken together, these heterogenous effects contribute to the enduring gender turnout gap.

Our framework draws, first, on the resource model of political participation (Schlozman, Burns and Verba, 1994; Verba, Burns and Schlozman, 1997). Gender differences in the stock of participatory resources–i.e., money, time, and civic skills–may have declined in recent years due to increasing female education and labor force participation. Nevertheless, family norms still structure individuals' control over these resource and the ability to devote them to politics. Studies of traditional societies in developing countries show that women are expected to play a greater role in the public sphere when they hold more control over material resources (Brule and Gaikwad, 2017; Robinson and Gottlieb, 2019). In advanced industrial societies, better economic opportunities and the possibility of divorce have improved women's bargaining power within families. Yet, marriage and the presence of children may still affect the gender division of household and paid labor to the disadvantage of women (Iversen and Rosenbluth, 2006; Alesina and Ichino, 2009). Recent work documents that work-life balance considerations related to raising a family often deter females from running for office (Silbermann, 2015). Here we argue that, if women are expected to do the greater share of household work, this may also increase the opportunity cost of voting. In particular, the time cost of getting informed about the election and turning out to vote may be prohibitive for the mothers (but not the fathers) of young children.

Whereas raising children puts greater pressure on women's resources compared to men's due to their social role as "care-givers", marriage tends to significantly improve men's well-being without having the same large positive effects on women. A second strand of related work documents these gendered effects of wedlock on psychological and emotional health. Although a happy marriage is a source emotional and material support for both men and women–reducing the incidence of depression and mental illness (Kiecolt-Glaser and Newton, 2001)–it has significantly stronger behavioral effects on males. Married men often adopt the social role of a "guardian", and en-

gage less in risky activities and more in healthy ones–perhaps for the sake of their partner (Wilson and Oswald, 2005). As these behaviors tend to be positively correlated with political engagement (Wolfinger and Wolfinger, 2008; Burden et al., 2017), this may in turn have positive downstream effects on men's levels of turnout, but may, however, leave women's participation unchanged.

Given the positive behavioral consequences of marriage, it may seem plausible that, conversely, divorce is a stressful experience that could depress voter turnout. Yet, Wolfinger and Wolfinger (2008) suggest that the effects documented in the U.S. literature may be driven by the fact that divorcees tend to move, which, in the American context, can result in loss of voter registration. Thus, the negative impact of divorce on voting may depend on the institutional barriers to voting. Prior work on widowhood is more clear-cut about the negative consequences of deteriorating mental and physical health due to loss of a partner (Hobbs, Christakis and Fowler, 2014). However, the gendered effects of widowhood on voter turnout are ambiguous. Whereas the sociological literature documents that men's well-being is more adversely affected than women's, especially in the period surrounding spousal death (Lillard and Waite, 1995), Hobbs, Christakis and Fowler (2014) find that the negative effect of widowhood on turnout is smaller among men than among women in the year surrounding spousal death. Here we suggest that married women's social role as "care-givers" may explain this paradox. That is, the spill-over effects of deteriorating health of the partner prior to spousal death may be stronger among women than among men.

A third line of theorizing suggests that marriage and having children may affect the social incentives to vote through socialization and peer pressure within the family. Spouses are likely to discuss politics and may encourage each other to participate in elections (Stoker and Jennings, 1995). They may also blame their partners for failing to comply with their civic duty to vote (Dellavigna et al., 2017). While these effects are well-documented in the literature, here we argue that they may have gendered consequences in two ways. First, marriage tends to equalize turnout among spouses (Stoker and Jennings, 1995). Given that unmarried men tend to vote less than unmarried women in advanced democracies, we may expect that the initial effect of marriage will increase male relative to female participation. Second, while having children may increase both

parents' feeling that voting is a civic duty, this social norm may conflict with other, family-related duties for women but not for men. As a result, having infant children may decrease female turnout, leaving male turnout unchanged, whereas having school-age children may increase turnout for men but not for women.

Prior Evidence

Distinguishing gendered turnout patterns from other correlated factors is challenging. A recent meta-analysis shows that most prior studies fail to find significant gender differences in electoral participation after controlling for other theoretically-important variables like education, income, and age (Smets and van Ham, 2013). Here we suggest that this decidedly mixed evidence likely follows from the biases of conventional approaches in the literature using conditional-on-observables estimation and survey measures of turnout. Namely, comparisons between men and women controlling for observable covariates may be biased because gender likely correlates with usually unobserved factors that also affect turnout, such as individual pro-social attitudes and civic skills. Moreover, validation studies show that surveys tend to under-estimate the gender turnout gap because non-voting women are less likely to (truthfully) respond to questions about participation compared to non-voting men (Dahlgaard et al., 2019). Importantly, the direction of the overall skew due to omitted variables and non-response bias is difficult to predict ex-ante. This shows that the literature needs studies combining credible research designs with administrative data.

Our study focuses on transitions in family status to explore whether gendered social norms can causally affect voter turnout. Yet, the prior empirical evidence about the effects of marital status is ambiguous. Wolfinger and Wolfinger (2008) estimate a strong, positive effect of marriage on turnout; though their cross-sectional analysis lends to limited causal interpretation, as the authors cannot rule out that their impact estimates are confounded by unobservable correlates of voter participation that differ across married and unmarried voters. Using panel survey data from the U.S., Stoker and Jennings (1995) find that, if anything, marriage depresses voter participation. But, consistent with the interpersonal voter mobilization hypothesis, they also find that, after marriage,

spouses adjust voter turnout to become more like each other. To our knowledge, Bhatti, Danckert and Hansen (2017) is the only prior work that uses longitudinal register data to estimate the effect of marital status on voter turnout. Although their paper focuses on the turnout effect of neighborhood ethnic diversity, they find that marriage substantially increases voter participation. Other recent work includes Hobbs, Christakis and Fowler (2014), who employ a matching and an event-study approach to show that voter participation in California declines after the death of a spouse. However, comparatively little is known about gender differences in the effects of marital status.

Evidence about the turnout effect of child-rearing is also scarce. Wolfinger and Wolfinger (2008)'s cross-sectional survey data suggest a positive effect of children on turnout. But the lack of longitudinal evidence limits the causal interpretation of their findings. A similar caveat applies to Arnold (2013), who finds a negative effect of children aged 5 or younger on parental turnout using the pooled American National Election Studies (ANES). In recent work, however, Dahlgaard and Hansen (2017) leverage a twin study design to show that, in the Danish context, having an additional child depresses turnout among both parents, but considerably more so among mothers than among fathers. Moreover, Dahlgaard (2018) utilizes a Regression Discontinuity Design based on children's age on Election Day to show the positive effect of children's voting eligibility on parental turnout in Denmark among both men and women.

To summarize, the main contribution of this study is to examine whether family norms are consequential for the gender turnout gap. To this end, we provide credible and comparable estimates of the gendered effects of various transitions in family status using large-scale administrative data and document the persistence of these effects over time. We also leverage different sources of survey data to explore potential mechanisms.

Research Setting

We study the 2004, 2009 municipal and European elections and the 2008, 2013 national parliamentary elections in the city of Bologna, a municipality of about 370,000 inhabitants in the Center-North of Italy. Voter turnout, though declining over time, has been historically very high in

Bologna. It was above 79% in the four elections we consider and was even slightly higher in the 2008 and 2013 national elections.

In 2004 and 2009, Bologna's mayoral race coincided with the election of Italy's Members of the European Parliament (MEPs). That is, voters who turned out on the 2004 and 2009 Election Days in Bologna received two separate ballots: one for municipal and one for European elections. Typically, European elections are less salient than municipal and national elections (Cantoni and Gazzè, 2018), which could have gendered consequences, given that women are less likely to turn out in second-order elections (Dassonneville and Kostelka, 2019). As we discuss in the Online Appendix, however, the average turnout difference between national elections and same-day municipal/European elections is small (1-1.5 p.p.). Thus, in the main analysis we pool all available data and control for election fixed effects; though we report models estimating effects separately by election-type in the Online Appendix.

Based on residential address, voters in Bologna are allocated to 436 voter precincts encompassing geographically close and contiguous areas. In turn, precincts determine assignment to pre-designated polling locations (typically public schools). Voter registration is automatic for Italian citizens. Except for Italians living abroad, there is no absentee or early voting in Italy. In the elections covered by our data, Bologna residents were allowed to vote on Sunday and the following Monday. In Table A13 in the Online Appendix, we show that there is no heterogneity in the effects of interest by distance to the polls, thus corroborating that barriers to voting are low.

Bologna and the region of Emilia-Romagna are known for their strong civic traditions and high levels of social capital (Putnam, Leonardi and Nanetti, 1994), and for being among the most progressive regions of northern Italy (in contrast to the considerably more conservative and traditional regions of southern Italy). Accordingly, the proportion of female city-councilors (36% in Bologna) is among the highest in the country. Moreover, levels of labor force participation among working-age women (67.4%) are above the average of Organisation for Economic Co-operation and Development (OECD) countries (62.6%). The coverage of public or publicly-funded pre-kindergarden schools is also high by national and international standards. Nevertheless, Italian

family norms–which prescribe strong gender imbalance in time devoted to household work–may still be strong in Bologna (Anxo et al., 2011). Thus, studying the effects of family status on voter turnout in this context is interesting on its own, but we should be less likely to find gendered patterns of voter turnout than in other parts of Italy and other OECD countries, on average, which enhances the generalizability of our findings.

Data

This project relies on three sources of data: administrative socio-demographic and voter turnout data at the individual level from the city of Bologna, survey data from the Italian National Election Studies (ITANES) to explore cross-sectional relationships between family structure and information acquisition, and survey data from the Italian National Statistical Agency (ISTAT) to test whether changes in family composition are reflected in changes in worked hours.

Our voter-level turnout data cover the universe of the voting-eligible population of the city of Bologna, in northern Italy. The data contain an anonymous, time-invariant voter identifier, which effectively gives us an unbalanced individual-level panel with up to four observations per voter. The retention rate is 84.2%–with deceased voters and out-of-town movers dropping out–over the nine years of our study. The data also feature a household identifier, which may vary if individuals change households.

The turnout data are complemented by detailed administrative socio-demographic information covering *every* resident of Bologna (i.e., including non-voting-eligible residents) updated as of, approximately, the four Election Days in the sample. Among others, these data contain: age in years, gender, marital status (i.e., never-married, married, divorced, or widowed), neighborhood, immigration status, position within the household, as well as income and income taxes paid in the year of the election. To construct the matched panel dataset with turnout and socio-demographic information, we digitized all Bologna's voter attendance sheets from the 2004, 2008, 2009, and 2013 elections. We then sent the turnout data to the municipal statistical office, which matched them against administrative socio-demographic records of the resident population. After anonymizing

and de-identifying the matched data, the municipality of Bologna sent us four files (i.e., one per election) with the turnout and socio-demographic information.

The demographic data also contain a variable for counts of household members. In some cases, that variable differs from the number of family members imputable by counting individuals with the same household identifier;[1] we exclude these cases from all samples. We also exclude 4,999 observations matched to no demographic data, 25 individuals who appear to have changed gender across elections, and six individuals with unknown marital status. Finally, to maintain a consistent sample across elections, we exclude voters who are not Italian citizens; citizens of other EU countries could, in fact, vote in the 2004 and 2009 European and municipal elections, but not in the 2008 and 2013 parliamentary elections.

Although the data do not say explicitly if an individual has children, we impute this information based on household structure. Specifically, one of the possible categories of the variable "position within household" is "Son/daughter of head of household." Because the demographic data cover the universe of the resident population (i.e., including children of any age), counting the number of individuals in that position gives the head of household's exact number of cohabiting children. Notice, however, that this imputation only makes sense for heads of households and their spouses; because the variable "position within household" is specified relatively to heads of households, it is complicated or even impossible to accurately determine whether individuals in other positions have children. For this reason, when we examine the effect of children on turnout, we limit the sample to heads of households and their spouses. Table A1 in the Online Appendix reports summary statistics for our administrative data.

Our second data source is the ITANES, which we use to explore (partial) correlations between family status, political knowledge, and information acquisition. By content, structure, and survey sampling strategy, the ITANES data follow the ANES. Professional polling companies are in

[1]Most of these inconsistencies are inmates and seniors living in retirement communities. All people living in the same community share the same household ID; but, according to the relevant variable in the demographic data, their households typically consist of one or two individuals.

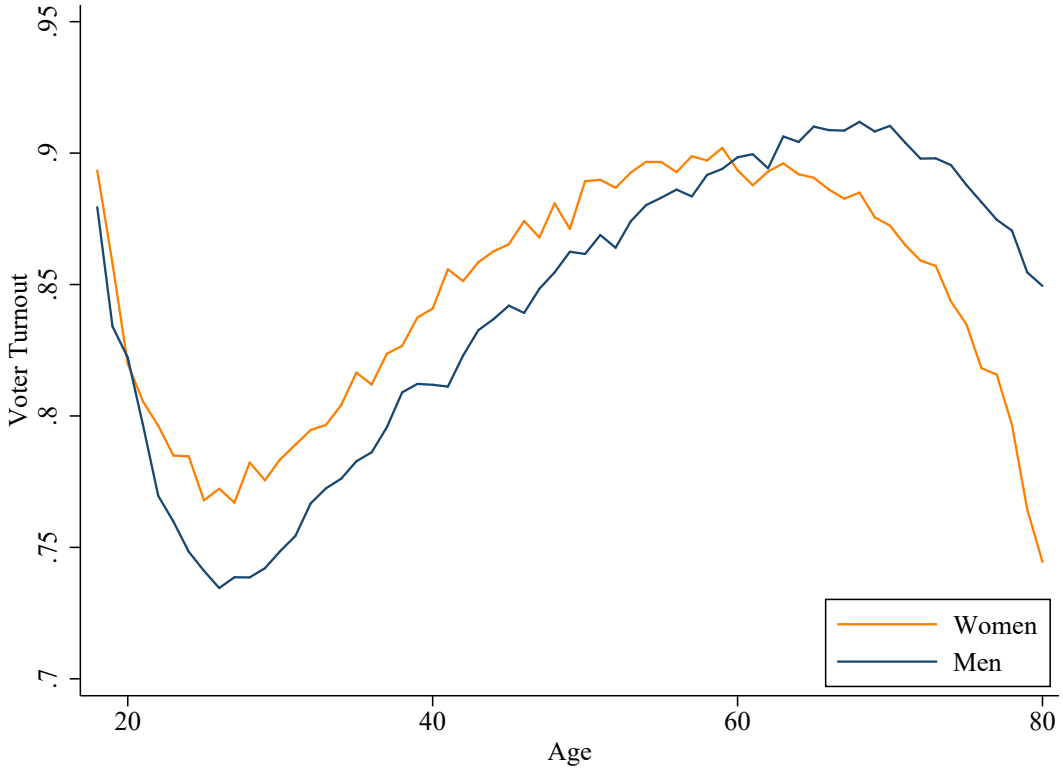
charge of administering the surveys to representative samples of the Italian voting-eligible population in the weeks following national elections. A typical ITANES survey asks a battery of socio-demographic questions (gender, age, education, marital status, presence of cohabiting children, etc.), questions on political opinions (e.g., which party the interviewee voted for in the most recent election, opinions on political leaders and policy-relevant issues), questions on information acquisition (e.g., frequency watching TV, reading newspapers), and questions testing factual political knowledge. For analysis based on ITANES data, we pool information from the 2001, 2006, 2008, and 2013 post-electoral surveys. Unfortunately, unlike the matched turnout and socio-demographic data from Bologna, the ITANES data do not track the same respondents over time; that is, we cannot exploit within-voter variation to estimate the effects of family status on voter information.

Our third and last source of data is the annual survey administered by ISTAT to collect information on salient aspects of Italian households' daily lives and behaviors. To construct the so-called AVQ data (from the Italian acronym for *Aspects of Daily Life*), each year ISTAT interviews a nationally representative sample of approximately 20,000 households and 50,000 people. We construct a pooled cross-sectional dataset using the 2005 through 2012 waves of the ISTAT AVQ data. Next to the respondents' basic socio-demographic information, we are interested in the number of hours of domestic and paid work, which we use to explore relations between family status and the allocation of labor across the two genders.

Results

Figure 1 shows average levels of turnout in Bologna among male and female voters. Consistent with patterns observed in other advanced democracies (Leighley and Nagler, 2013; Quaranta and Dotti Sani, 2018; Bratsberg et al., 2019), we find that young women are more likely to vote than men of the same age. This difference is largest among voters aged 25-30, though it narrows over the life-course, and reverses among citizens aged 60 or older. To contribute to explaining these patterns, we next explore the turnout effects of gendered social norms in the context of transitions in family status.

Figure 1: Voter Turnout by Age and Gender



Marriage Increases Voter Turnout

We begin the analysis by examining whether changes in marital status have an effect on turnout, above and beyond other correlated factors. To this end, Table 1 presents estimates from DD regressions of the following form:

$$voted_{it} = \beta^m married_{it} + \beta^d divorced_{it} + \beta^w widowed_{it} + \alpha_i + \delta_t + age_{it}^{sex} + X_{it}^{'}\gamma + \varepsilon_{it}, \qquad (1)$$

where $vote_{it}$ is a dummy for whether voter i turned out to vote in election t; $married_{it}$, $divorced_{it}$, and $widowed_{it}$ are mutually exclusive dummies for whether voter i's marital status as of Election Day t was, respectively, married, divorced, or widowed;[2] α_i, δ_t, and age_{it}^{sex} denote full sets of voter, election, and age in years-by-gender fixed effects, respectively; X_{it} is a set of controlling

[2]That is, the omitted category of marital status is "never married".

covariates.[3] Standard errors are two-way clustered by voter and household. Voter-level clustering accounts for potential serial correlation of regression residuals within voters (Bertrand, Duflo and Mullainathan, 2004); household-level clustering accounts for the marriage treatment simultaneously affecting couples of voters within the same household.

We find that marriage increases voter turnout by .7 to 1 p.p. relative to never-married voters. This estimate is significant at conventional levels and is virtually unaffected by controlling for neighborhood (column 2)[4] and family characteristics (column 3), earned income and income taxes paid during the year of the election (column 4), and separate gender-specific dummies for the presence of cohabiting children of the following ages: 0–5, 6–11, 12–17, 18 or more (column 5).

Due to the lower number of voters switching to or out of these marital statuses, turnout effects of divorce and widowhood are estimated less precisely. With this caveat in mind, divorced voters appear .8-to-1.1 p.p. more likely to vote than their never-married counterparts. Since that is also the magnitude of the marriage effect, this finding is consistent with the marriage-to-divorce transition inducing no change in voter participation. By contrast, widowhood reduces voter participation by .9 to 1.3 p.p. relative to never-married voters. Consequently, the marriage-to-widowhood transition

[3]We interact gender with age-in-years fixed effects for two reasons. First, transitions in marital status and birth of children typically occur at different ages for men and women. For example, female (resp. male) voters in our sample who have just switched from "never married" to "married" are 36.5 (resp. 38.5) years old, on average. Second, women and men of the same age may have different turnout rates, even in absence of any treatment. For example, due to their longer life expectancy, elderly women may be in better health than men of the same age. If better health translates to higher turnout, we may expect old women to turn out at higher rates than same-aged men. Accounting for these differences seems important to avoid omitted variable bias as we explore heterogeneous effects by gender, which we do in Subsection .

[4]Neighborhood controls are: precinct-year average age, income, and income taxes paid, as well as shares of female and Italian residents, and neighborhood-by-year fixed effects.

Table 1: Turnout Effect of Marital Status

	Outcome: Voter-Level Turnout									
	(1)		(2)		(3)		(4)		(5)	
1(Married)	.010	**	.010	**	.010	**	.010	**	.012	**
	(.003)		(.003)		(.003)		(.003)		(.003)	
1(Divorced)	.011	~	.011	~	.011	~	.010	~	.013	*
	(.006)		(.006)		(.006)		(.006)		(.006)	
1(Widowed)	-.009	~	-.009	~	-.009	*	-.014	**	-.012	*
	(.005)		(.005)		(.005)		(.005)		(.005)	
Voter FEs	✓		✓		✓		✓		✓	
Age×Gender FEs	✓		✓		✓		✓		✓	
Election FEs	✓		✓		✓		✓		✓	
Neighborhood controls			✓		✓		✓		✓	
Household controls					✓		✓		✓	
Income and taxes paid							✓		✓	
Children×Gender FEs									✓	
Never married $\bar{Y}$	.800		.800		.800		.800		.800	
N	1,084,202		1,084,202		1,084,202		1,084,202		1,084,202	

Notes: Neighborhood controls are: precinct-year average age and OECD modifed gross household income, as well as shares of female and Italian residence, and city neighborhood-by-year fixed effects. Household controls are the share of household members who are Italian citizens and the OECD modified gross household income. Children FEs are four dummies indicating presence of one or more children of the following ages: 0-5, 6-11, 12,-17, 18+. Standard errors are two-way clustered by voter and household.
** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

appears to induce a 2-p.p. drop in voter participation (i.e., $\beta^w - \beta^m \approx -.02$).

Marriage Increases Men's Voter Turnout, Leaves Women's Unchanged

As discussed above, turnout effects of changes in marital status could differ across genders for at least two reasons. First, marital status could create different situational constraints for female and male voters. Specifically, transitions across marital statuses could induce gender-specific changes in the time voters have to cast their ballots or to follow the political discourse. Second, marital status could differentially affect the level of civic engagement of voters of the two genders. For example, if single women and men tend to vote at different rates, marriage could equalize turnout by inducing the two spouses to vote (or abstain) together.

We explore heterogeneous effects by gender using DD regressions of the following form:

$$\begin{aligned} voted_{it} = & female_i \times \left(\beta^{m,female} married_{it} + \beta^{d,female} divorced_{it} + \beta^{w,female} widowed_{it}\right) + \\ & male_i \times \left(\beta^{m,male} married_{it} + \beta^{d,male} divorced_{it} + \beta^{w,male} widowed_{it}\right) + \\ & \alpha_i + \delta_t + age_{it}^{sex} + X_{it}^{'}\gamma + \varepsilon_{it}. \end{aligned} \tag{2}$$

Equation 2 augments regression 1 with gender-specific treatments. Table 2 reports estimated effects for female and male voters, along with female-minus-male differences in impact estimates.

The positive effect of marriage on turnout is concentrated entirely on male voters, ranging from 1.9 to 2 p.p.. By contrast, estimated effects on women are tightly centered around zero and insignificant across all specifications. Gender differences in impact estimates (i.e., $\beta^{m,female} - \beta^{m,male}$) range from -1.5 to -2.1 p.p. and are significant at the 1-percent level. That is, the never married-to-married transition increases men's turnout by 1.5 to 2.1 p.p. relative to women undergoing the same change in marital status.

Gender heterogeneity in marriage effects exactly offsets differences in turnout between never-married men and women–i.e., in our preferred specification in column 5, the differential impact estimate, $\beta^{m,female} - \beta^{m,male}$, is the same as the average turnout difference between men and women

Table 2: Turnout Effect of Marital Status by Voter's Gender

	Outcome: Voter-Level Turnout									
	(1)		(2)		(3)		(4)		(5)	
1(Married female)	-.001		-.001		-.001		-.001		.004	
	(.005)		(.005)		(.005)		(.005)		(.005)	
1(Married male)	.020	**	.020	**	.020	**	.019	**	.019	**
	(.004)		(.004)		(.004)		(.004)		(.004)	
1(Divorced female)	.008		.008		.008		.008		.013	~
	(.008)		(.008)		(.008)		(.008)		(.008)	
1(Divorced male)	.011		.011		.011		.010		.011	
	.008		.008		(.008)		(.008)		(.008)	
1(Widowed female)	-.021	**	-.021	**	-.021	**	-.026	**	-.021	**
	(.006)		(.006)		(.006)		(.006)		(.006)	
1(Widowed male)	.003		.003		.002		.001		.000	
	(.007)		(.007)		(.007)		(.007)		(.007)	
$\beta^{\text{married female}} - \beta^{\text{married male}}$	-.021	**	-.021	**	-.021	**	-.020	**	-.015	*
	(.006)		(.006)		(.006)		(.006)		(.006)	
$\beta^{\text{divorced female}} - \beta^{\text{divorced male}}$	-.003		-.003		-.003		-.002		.003	
	(.011)		(.011)		(.011)		(.011)		(.011)	
$\beta^{\text{widowed female}} - \beta^{\text{widowed male}}$	-.024	**	-.024	**	-.023	*	-.027	**	-.021	*
	(.009)		(.009)		(.009)		(.009)		(.009)	
Voter FEs	✓		✓		✓		✓		✓	
Age×Gender FEs	✓		✓		✓		✓		✓	
Election FEs	✓		✓		✓		✓		✓	
Neighborhood controls			✓		✓		✓		✓	
Household controls					✓		✓		✓	
Income and taxes paid							✓		✓	
Children×Gender FEs									✓	
Never-married female $\bar{Y}$	.813		.813		.813		.813		.813	
Never-married male $\bar{Y}$	.787		.787		.787		.787		.787	
N	1,084,202		1,084,202		1,084,202		1,084,202		1,084,202	

Notes: Neighborhood controls are: precinct-year average age and OECD modifed gross household income, as well as shares of female and Italian residence, and city neighborhood-by-year fixed effects. Household controls are the share of household members who are Italian citizens and the OECD modified gross household income. Children FEs are four dummies indicating presence of one or more children of the following ages: 0-5, 6-11, 12,-17, 18+. Standard errors are two-way clustered by voter and household.
** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

reported at the bottom of the table. Although not definitive, this is consistent with the notion that marriage equalizes voter participation across spouses by lifting men's turnout to the higher pre-marriage level as their wives'.[5] While divorce induces indistinguishable positive effects on male and female voters, transitions into widowhood appear to significantly reduce women's turnout with no effect on men's.

Since models 1 and 2 control for voter fixed effects, resulting estimates are free from omitted variable bias due to time-invariant individual characteristics that potentially correlate with family status. However, causal identification of family structure parameters hinges on a "parallel-trend" assumption. That is, voters who get married in the sample period (i.e., treated voters) would, absent changes in marital status, experience identical over-time changes in turnout as voters who do not change marital status (i.e., control voters). As the parallel-trend assumption is a statement on counterfactual outcomes—that is, unobservable changes in treated voters' turnout in absence of treatment—, it cannot be tested directly. Yet, it can be tested indirectly in the spirit of Granger (1969). The idea is to check that causes happen before consequences, and not vice versa. To this end, Figure 2a plots estimates of $\beta_{\tau}^{m,female}$'s (in orange), $\beta_{\tau}^{m,male}$'s (in blue), and corresponding 95-percent confidence intervals from the following event-study specification:

$$voted_{it} = \sum_{\tau} married_{it\tau} \times \left(\beta_{\tau}^{m,female} female_i + \beta_{\tau}^{m,male} male_i\right) + \alpha_i + \delta_t + age_{it}^{sex} + X_{it}^{'}\gamma + \varepsilon_{it}, \quad (3)$$

where $married_{it\tau}$ is a dummy equal to 1 if election t occurs τ elections since the first election voter i's marital status was "married".[6] Because our data span four elections, τ ranges between -3 and

[5]We further illustrate the powerful equalizing effect of marriage in Figure A1 in the Online Appendix: whereas single women are significantly more likely to vote than men with similar levels of income, this difference is small among married voters, with perhaps a slight advantage for men.

[6]Recall that we observe voters' marital status as of Election Day, but we do not observe the exact date of marriage.

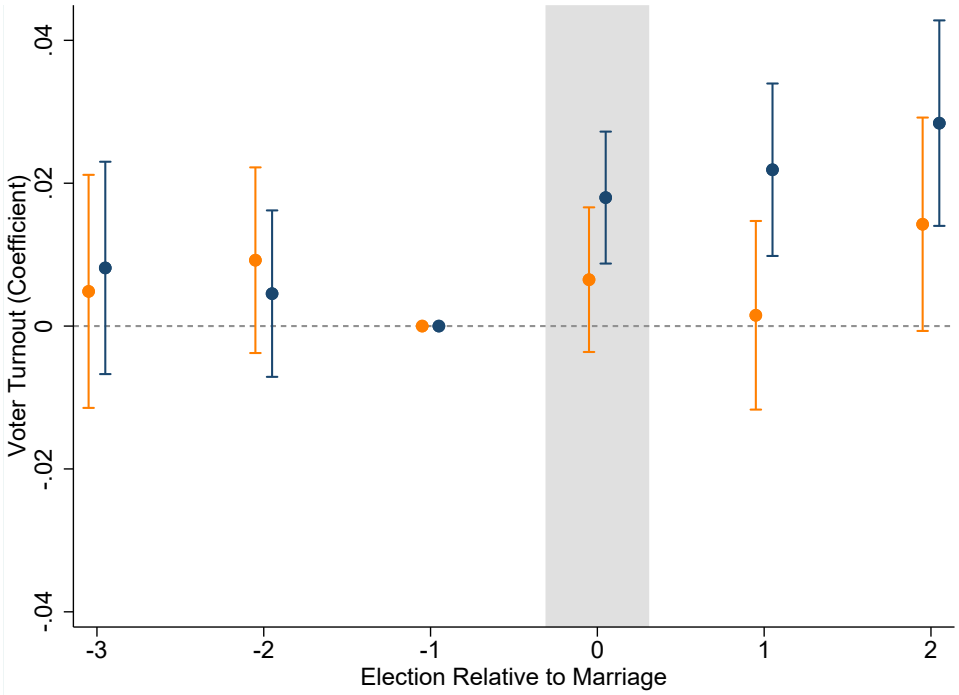
2. The coefficients of interest are the $\beta_{\tau}^{m,sex}$'s, measuring the turnout difference, conditional on controls, between married and control voters τ elections before ($\tau < 0$) or after ($\tau \geq 0$) marriage. All coefficients are normalized relative to $\tau = -1$; that is, the last election before marriage. The vector X_{it} includes all controls from the most demanding specification in Table 2 (i.e., column 5).

Reassuringly for our DD identification assumption, treated and control men share statistically indistinguishable voter turnout in pre-marriage elections (i.e., $\tau < 0$). In contrast, married men's turnout increases after marriage (i.e., $\tau \geq 0$) by a significant 2 p.p., a magnitude which is consistent with the findings from Table 2. Similarly, there are no obvious pre-trends in married women's voter turnout. At the same time, there is no noticeable change in turnout after marriage, with the possible exception of a marginally significant increase after three elections (i.e., at $\tau = 2$). Again, this finding supports the zero effect of marriage on women's turnout shown in Table 2.

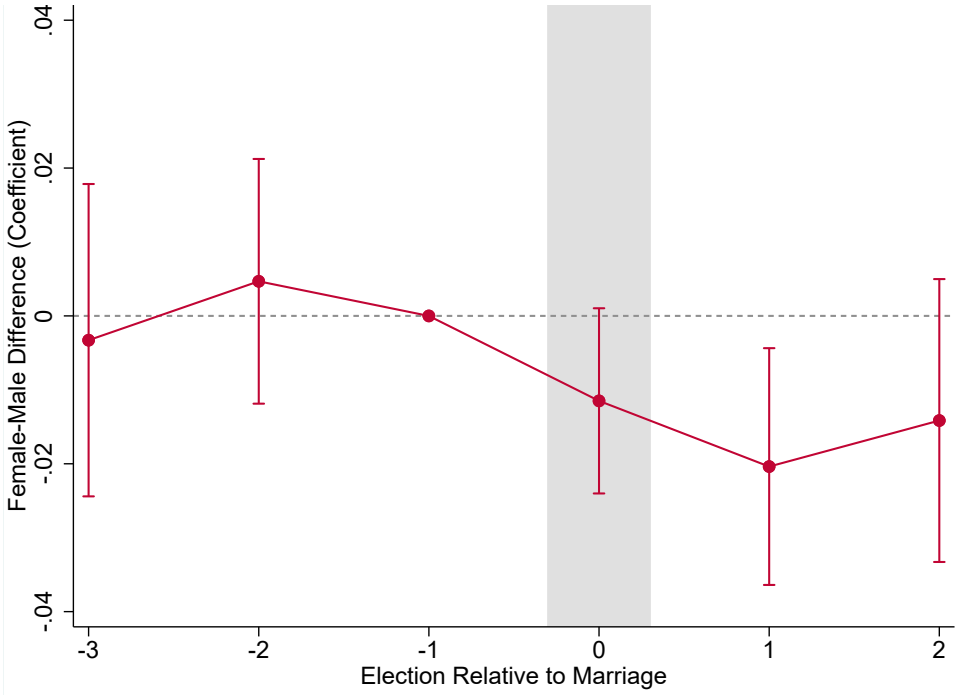
Figure 2b plots female-minus-male differences in marriage-induced turnout effects (i.e., $\beta_{\tau}^{m,female} - \beta_{\tau}^{m,male}$), along with corresponding 95-percent confidence intervals. Upholding the validity of the parallel-trends assumption, pre-marriage gender differences are centered around zero and insignificant. By contrast, in post-marriage elections, married women's turnout decreases by 1-to-2 p.p. relative to married men's, which is again in line with estimates of gender differences from Table 2. Event-study estimates of the turnout effects of divorce and widowhood (resp., on gender differences in impact estimates) are plotted, respectively, in Appendix Figures A2a and A3a (resp., A2b and A3b). Event-study plots for divorce support the findings from Table 2. For example, Figure A2a shows no change in turnout after divorce, which is consistent with the marriage-to-divorce transition inducing no change in voter participation. As for widowhood, Figure A3a reveals a significant decline in turnout that predates widowhood. A possible explanation is that the gradual deterioration of a spouse's health—which likely precedes widowhood—progressively reduces the surviving spouse's turnout (e.g., because the growing attention required by the dying spouse reduces the time available to follow politics and vote). Interestingly, gender differences in widowhood effects are driven by differences in pre-widowhood turnout (rather than by differential changes in men's and women's participation after widowhood), particularly so at $\tau = -3$.

Figure 2: Marriage Event Study

(a) Orange = Wife, Blue = Husband



(b) Marriage-Induced Gender Difference in Voter Turnout



Notes: Panel A plots event-study estimates of the effect of marriage on women's (orange) vs. men's (blue) turnout, along with 95-percent confidence intervals. All estimates are from a unique regression controlling for the same covariates included in Table 1, column 5. The x-axis denotes the election relative to the first election in which a voter's marital status is married. Panel B plots differences between female- and male-specific effects.

Children 0–5 Decrease Women's Turnout, but not Men's; Children 5–15 Increase Men's Turnout, but not Women's

In Table 2, the gender difference in the effect of marriage on turnout shrinks by a third from column 2 ($\beta^{m,female} - \beta^{m,male} = -.021$) to column 5 ($\beta^{m,female} - \beta^{m,male} = -.015$). Unlike other specifications, column 5 controls for interactions between gender and dummies for the presence of kids in the household, thus suggesting that children affect at least one of their parents' turnout. We explore this possibility in two steps. First, we estimate the average effect of children on parental turnout. Second, we test for differences in the effect of children on maternal vs. paternal turnout. Formally, we start with the following DD specification:

$$\begin{aligned} voted_{it} = & \beta^{0to5} kids0to5_{it} + \beta^{6to11} kids6to11_{it} + \beta^{12to17} kids12to17_{it} + \beta^{18+} kids18+ \\ & \alpha_i + \delta_t + age_{it}^{sex} + X_{it}^{'}\gamma + \varepsilon_{it}, \end{aligned} \tag{4}$$

which, relative to equation 1, replaces dummies for marital status with controls for the presence of kids aged 0–5, 6–11, 12–17, and 18 or more. Table 3 reports the results.[7]

Children aged 0 to 5 reduce parental turnout by .6 to .8 p.p., an effect that is significant at least at the 10-percent level across all specifications. Conversely, children aged 6–11, 12–17, and 18 or older *increase* voter turnout by .7–.8, .9–1, and 1.2–1.2 p.p., respectively. This pattern of effects—that increase with children's age—is consistent with younger children posing a situational constraint to parental electoral participation; for instance, due to the logistical difficulty of reaching one's polling location in presence of young children or to the limited time available to acquire in-

[7]Remember that we can accurately determine whether a voter has children only if that person is the head of household or her/his spouse. Thus, unlike Tables 1 and 2, whose estimation sample include all eligible voters, the sample for children regressions is limited to the subset of voters whose position within the household is "head of household" or "spouse/partner of head of household."

Table 3: Effect of Children on Turnout by Children's Age

	Outcome: Voter-Level Turnout									
	(1)		(2)		(3)		(4)		(5)	
1(Children aged 0-5)	-.007	**	-.006	**	-.006	*	-.006	*	-.008	**
	(.002)		(.002)		(.002)		(.002)		(.003)	
1(Children aged 6-11)	.008	**	.008	**	.008	**	.008	**	.007	**
	(.002)		(.002)		(.002)		(.002)		(.002)	
1(Children aged 12-17)	.009	**	.009	**	.010	**	.009	**	.009	**
	(.002)		(.002)		(.002)		(.002)		(.002)	
1(Children aged 18+)	.012	**	.012	**	.012	**	.012	**	.012	**
	(.002)		(.002)		(.002)		(.002)		(.002)	
Voter FEs	✓		✓		✓		✓		✓	
Age×Gender FEs	✓		✓		✓		✓		✓	
Election FEs	✓		✓		✓		✓		✓	
Neighborhood controls			✓		✓		✓		✓	
Household controls					✓		✓		✓	
Income and taxes paid							✓		✓	
Marital status×Gender FEs									✓	
No kids $\bar{Y}$	.820		.820		.820		.820		.820	
N	902,153		902,153		902,153		902,153		902,153	

Notes: Neighborhood controls are: precinct-year average age and OECD modifed gross household income, as well as shares of female and Italian residence, and city neighborhood-by-year fixed effects. Household controls are the share of household members who are Italian citizens and the OECD modified gross household income. Marital status FEs are three, mutually exclusive dummies indicating married, divorced, and widowed voters. Standard errors are two-way clustered by voter and household.
** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

Table 4: Effect of Children on Turnout by Children's Age and Voter's Gender

	Outcome: Voter-Level Turnout									
	(1)		(2)		(3)		(4)		(5)	
1(Female w/ children aged 0-5)	-.017	**	-.017	**	-.016	**	-.016	**	-.017	**
	(.003)		(.003)		(.003)		(.003)		(.003)	
1(Male w/ children aged 0-5)	.004		.004		.005		.004		.001	
	(.003)		(.003)		(.003)		(.003)		(.003)	
1(Female w/ children aged 6-11)	.004		.003		.004		.004		.004	
	(.003)		(.003)		(.003)		(.003)		(.003)	
1(Male w/ children aged 6-11)	.011	**	.011	**	.011	**	.011	**	.010	**
	(.003)		(.003)		(.003)		(.003)		(.003)	
1(Female w/ children aged 12-17)	.005	*	.005	*	.006	*	.006	*	.006	*
	(.003)		(.003)		(.003)		(.003)		(.003)	
1(Male w/ children aged 12-17)	.013	**	.013	**	.014	**	.013	**	.013	**
	(.003)		(.003)		(.003)		(.003)		(.003)	
1(Female w/ children aged 18+)	.013	**	.013	**	.013	**	.013	**	.013	**
	(.002)		(.002)		(.002)		(.002)		(.002)	
1(Male w/ children aged 18+)	.011	**	.011	**	.012	**	.012	**	.011	**
	(.002)		(.002)		(.002)		(.002)		(.002)	
$\beta^{0\text{-}5\ female} - \beta^{0\text{-}5\ male}$	-.021	**	-.021	**	-.021	**	-.020	**	-.018	**
	(.004)		(.004)		(.004)		(.004)		(.004)	
$\beta^{6\text{-}11\ female} - \beta^{6\text{-}11\ male}$	-.007	*	-.007	*	-.008	*	-.007	*	-.006	~
	(.004)		(.004)		(.004)		(.004)		(.004)	
$\beta^{12\text{-}17\ female} - \beta^{12\text{-}17\ male}$	-.008	*	-.008	*	-.008	*	-.008	*	-.007	*
	(.003)		(.003)		(.003)		(.003)		(.003)	
$\beta^{18+\ female} - \beta^{18+\ male}$	.002		.002		.001		.001		.001	
	(.003)		(.003)		(.003)		(.003)		(.003)	
Voter FEs	✓		✓		✓		✓		✓	
Age×Gender FEs	✓		✓		✓		✓		✓	
Election FEs	✓		✓		✓		✓		✓	
Neighborhood controls			✓		✓		✓		✓	
Household controls					✓		✓		✓	
Income and taxes paid							✓		✓	
Marital status×Gender FEs									✓	
Female w/o kids $\bar{Y}$	.809		.809		.809		.809		.809	
Male w/o kids $\bar{Y}$	.833		.833		.833		.833		.833	
N	902,153		902,153		902,153		902,153		902,153	

Notes: Neighborhood controls are: precinct-year average age and OECD modifed gross household income, as well as shares of female and Italian residence, and city neighborhood-by-year fixed effects. Household controls are the share of household members who are Italian citizens and the OECD modified gross household income. Marital status FEs are three, mutually exclusive dummies indicating married, divorced, and widowed voters. Standard errors are two-way clustered by voter and household.

** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

formation about the election. As children grow, this situational constraint dissipates or is offset by increasing political engagement, so that the net effect of children on political participation reverses sign and becomes positive. By the time children reach voting age, this positive effect becomes substantial (+1.2 p.p.), possibly because parents receive positive turnout spillovers from their children's voting eligibility (e.g., because parents accompany their children to vote for the first time; Dahlgaard, 2018).

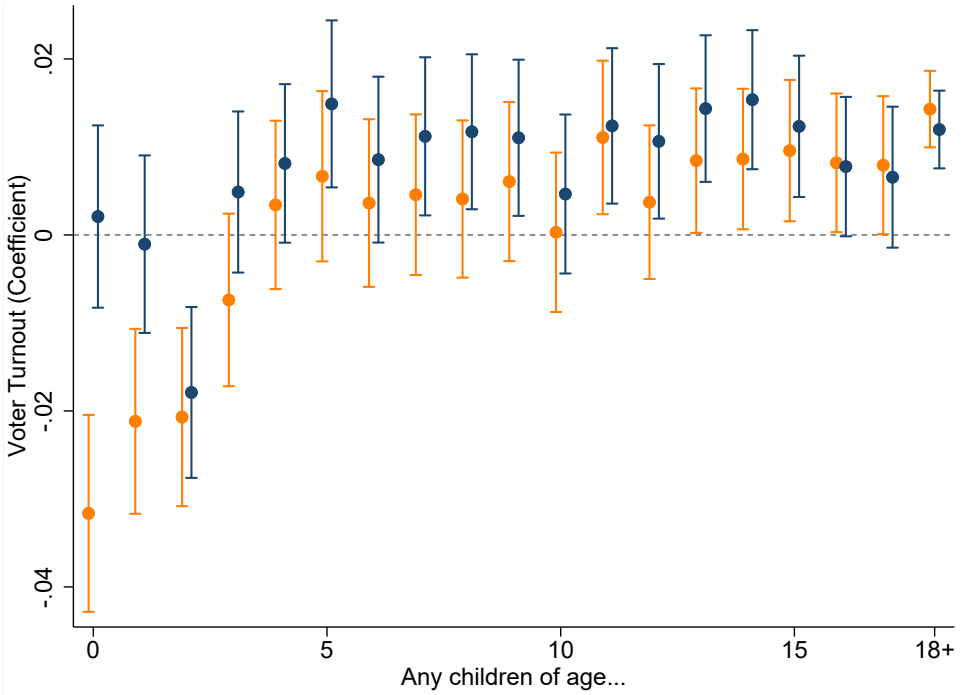
Table 4 reports estimates of the effect of children on maternal vs. paternal turnout and reveals that the negative effect of young children on political participation reported in Table 3 was driven by females voters. In fact, the presence of children aged 0 to 5 significantly reduces women's turnout, leaving men's participation unaffected. The difference between the negative effect on women and the zero effect on men is around 2 p.p. and is significant at the 1-percent level. While older children (aged 6 to 11 and 12 to 17) do not depress maternal turnout, they do increase paternal participation by approximately 1 p.p.. Only when children reach voting age (18 years), this heterogeneity dissipates, and both men's and women's turnout increases by around 1 p.p. relative to childless voters.[8] In other words, after a long break induced by motherhood, women resume their involvement in politics only around the time their children are themselves called to vote.

We explore whether the effects of childrearing vary depending on the parents' gender attitudes in the Online Appendix (Table A12). Our results suggest that the demobilizing effects of young children on mothers are stronger in families where both parents originally come from the (more conservative) South of Italy. Yet, we find that the negative effects of childrearing on maternal turnout are muted among unmarried parents–who likely hold relatively progressive views towards gender issues. Interestingly, the arrival of children also appears to boost the political participation of unmarried males to the level of their female partners, a process that mirrors the equalizing effects of marriage.

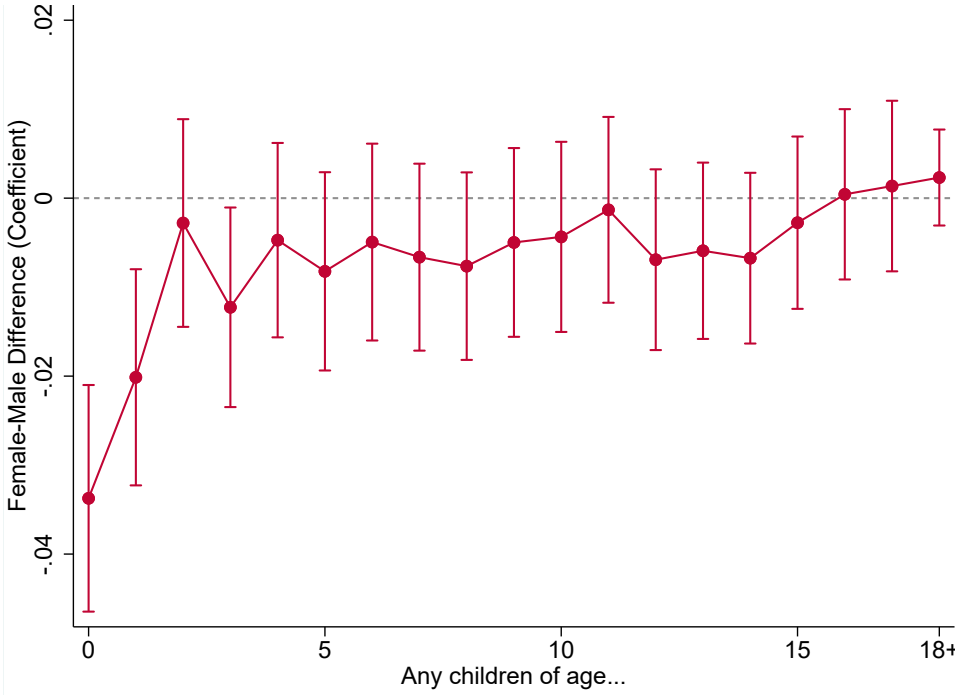
[8]In Appendix Tables A10 and A12, we show that estimates of the turnout effects of children are virtually unchanged (and, if anything, are slightly more precise) when we further restrict the sample to married voters.

Figure 3: Effects of Children on Turnout by Children's Age

(a) Orange = Mother, Blue = Father



(b) Children-Induced Gender Difference in Voter Turnout



Notes: Panel A plots estimates of the effect of children of different ages on maternal (orange) vs. paternal (blue) turnout, along with 95-percent confidence intervals. All estimates are from a unique regression controlling for the same covariates included in Table 3, column 5. Panel B plots differences between female- and male-specific effects.

To further investigate the pattern of children's effects, Figure 3a plots estimates of the turnout effects of children of specific ages. That is, the underlying regression controls for the same covariates as Table 4, column 5, but uses dummies based on 1-year age intervals for children's age (i.e., 0-year-olds, 1-year-olds, etc.) instead of four, broader intervals (i.e., 0–5, 6–11, 12–17, and 18+). The plot reveals that while the negative effect of children on maternal turnout is strongest in the months following childbirth, it persists for several years, and then vanishes when children turn four. By contrast, 0-to-5-year-olds have no effect on paternal turnout (with the possible exception of 2-year-olds), while older kids increase fathers' turnout. The ensuing gender difference is sizable (about 1 p.p.) and stable for children aged 5 through 15 (Figure 3b). When kids turn 16, this heterogeneity disappears.

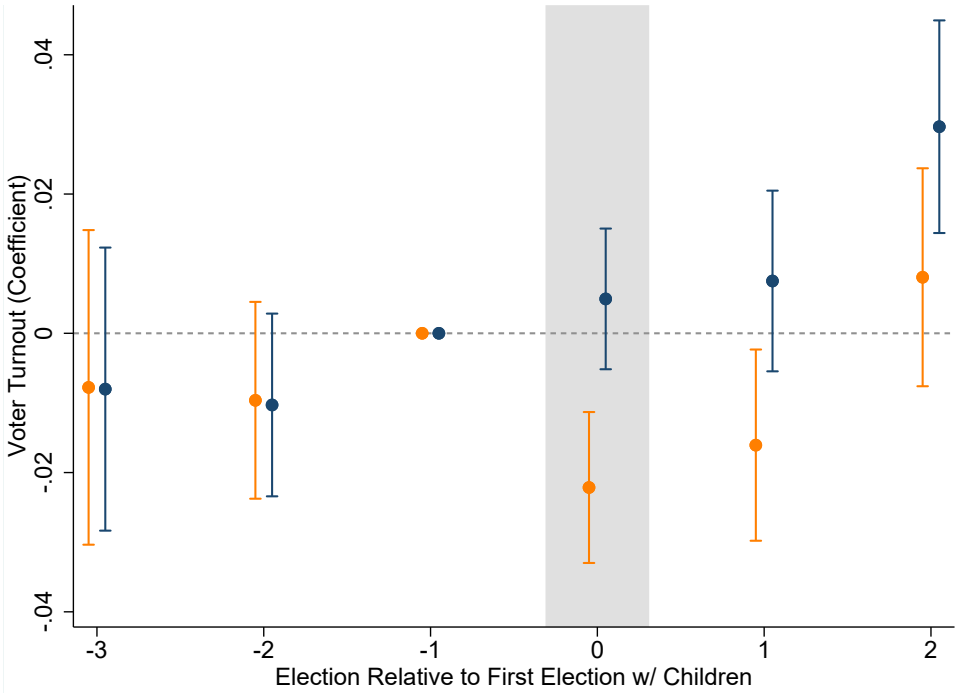
Figure 4a plots event-study coefficients of the effect of children on maternal (in orange) and paternal turnout (in blue). The underlying regression controls for the full set of covariates used for Table 3, column 5. Like in prior event-study plots, $\tau = 0$ denotes the first election a treated voter (i.e., a voter who switches from having no kids to having kids) is observed having at least one child. Analogously, $\tau = -1$ denotes the last election without kids, $\tau = +1$ is the second election with kids, etc..

Figure 4a also shows that there are no pre-trends in voter turnout; that is, treated and control voters have identical (conditional) turnout in pre-children elections. Corroborating the gender-heterogeneity documented in Table 4, treated women's turnout falls sharply (relative to control women) in the first election with children ($\tau = 0$), while treated men witness no drop in turnout. By the third election with children ($\tau = 2$), women's turnout recovers to pre-treatment levels and men's increases by approximately 2 p.p..

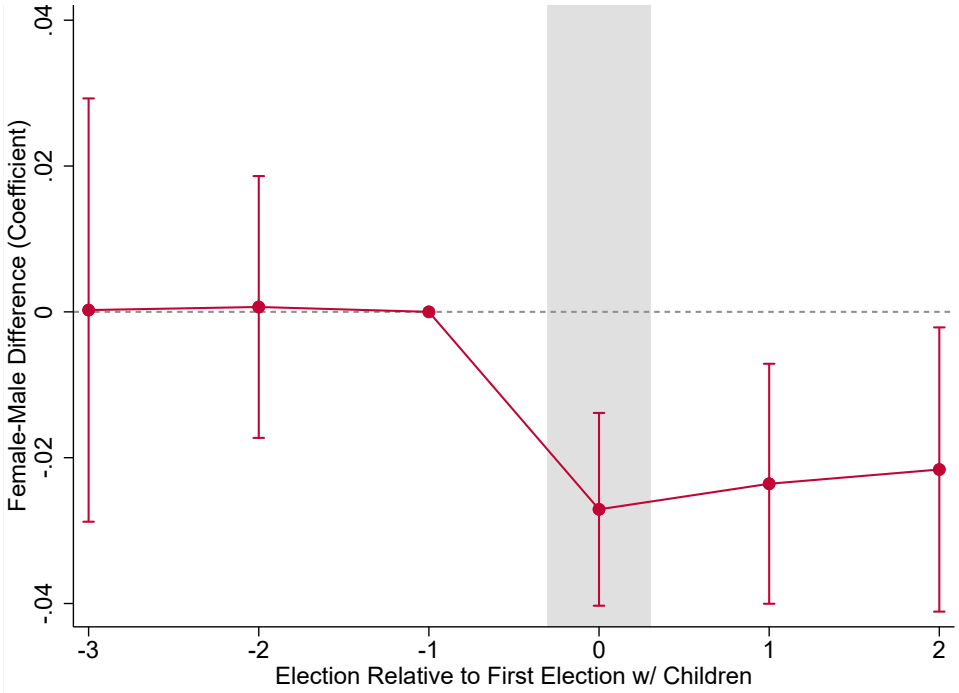
Figure 4b plots female-minus-male differences in event-study coefficients (i.e., $\beta_\tau^{m,female} - \beta_\tau^{m,male}$). Bolstering the causal interpretation of our findings, *treated* men and women share identical (conditional) turnout in pre-children years; that is, there is no pre-trend in the effect of children on the turnout gender gap, and all variation in said gap materializes suddenly and persistently in post-children elections.

Figure 4: Children Event Study

(a) Orange = Mother, Blue = Father



(b) Children-Induced Gender Difference in Voter Turnout



Notes: Panel A plots event-study estimates of the effect of children on maternal (orange) vs. paternal (blue) turnout, along with 95-percent confidence intervals. All estimates are from a unique regression controlling for the same covariates included in Table 3, column 5. The x-axis denotes the election relative to the first election in which a voter is first observed having children. Panel B plots differences between female- and male-specific effects.

Exploring Possible Mechanisms

We now explore the drivers of gender differences in turnout effects, using pooled cross-sectional survey data from ITANES and ISTAT as well as additional evidence from our administrative data. Specifically, we focus on the following three facts documented earlier in the paper: (i) the negative (resp., zero) effect of children aged 0–5 on maternal (resp., paternal) turnout, (ii) the positive (resp., zero) effect of children aged 6+ on paternal (resp., maternal) turnout, (iii) the positive (resp., zero) effect of marriage on men's (resp., women's) turnout. We report the results of our additional analyses in the Online Appendix and discuss our main findings here.

One mechanism by which having children may increase voter turnout is by stimulating parental civic involvement (Wolfinger and Wolfinger, 2008). To test for this mechanism, we check whether turnout effects are paralleled by similar patterns of correlations between the presence of children and political knowledge. For example, if higher civic engagement underlies the positive effect of children on paternal turnout, men with children should be more knowledgeable about politics than men without kids. Alternatively, children may raise paternal turnout through peer pressure (e.g., increasing the probability that, say, other parents will ask about one's turnout; Dellavigna et al., 2017). If peer pressure alone explains the positive effect of children on paternal turnout, then men with and without kids may be expected to display indistinguishable levels of political information.

Table A2 shows that men with kids are in fact more politically knowledgeable than their childless counterparts. This comports with the hypothesis that childrearing stimulates men's civic engagement. However, the presence of children does not correlate with mothers' political knowledge. Table A2 also shows that children induce women acquire more information on political campaigns. Yet, unlike men, women with children appear to rely on uninformative channels like TV ads, campaign leaflets, and campaign posters.

Another way family structure may differentially affect men's and women's voter turnout is by inducing heterogeneous effects on the quantity and type of time available to follow politics. For example, young children may require a disproportionate amount of maternal attention, leaving

new moms with little time for political participation (i.e., a "maternal time-constraint" effect). But children may also induce a "specialization effect", whereby fathers increase paid labor while mothers specialize in housework. If social interactions in the workplace spur greater interest in politics than those occurring during housework, then the specialization effect could help explain the heterogeneous impact of children on parental turnout.

Table A6 directly relates to this mechanism and shows that women with children report more hours of domestic work and less paid work than women without children. Young children also appear to increase male domestic workload, but to a much lesser extent than they increase women's. Furthermore, men with children of any age report more hours of paid work than their childless counterparts. Table A7 provides additional evidence from our administrative data and shows that the presence of children increases married men's income relative to their wives. This corroborates the notion that childrearing increases men's participatory resources compared to women's.

Interestingly, Table A8 documents that the transition from never-married to married increases men's income, leaving women's income unchanged. This finding is consistent with the notion that marriage improves men's relative well-being compared to women, which may have downstream consequences for gendered patterns of political participation. Furthermore, in Figure A4a, we show that cohabitation increases both men and women's turnout–unlike marriage, which increases only men's turnout (see Figure 2a). This suggests that moving in together has positive effects on the well-being of both male and female partners, whereas the effects of wedlock are gendered. This provides additional evidence of how gender roles shaped by family norms contribute to the enduring gender turnout gap.

Finally, we explore to what extent do gender differences in turnout effects matter for electoral outcomes. Do children or marital status also impact political preferences? After all, if Italian men and women share similar political preferences and family status does not affect political leanings, the differential effect of children on maternal vs. paternal political participation does not pose concerns about the representativeness of the voting electorate. Table A9 shows that, mirroring gender differences that have been documented for other democracies (e.g., Kittilson, 2016), Italian

women appear more leftist than men. However, neither the presence of children nor marriage seem systematically related to political leanings. Thus, even if changes in family status may not affect voters' political preferences directly, they may still be consequential by shifting the gender composition of the electorate.

Conclusion

In this article, we explore whether gendered social norms can causally affect political participation. To this end, we provide an explicit conceptual framework suggesting that the socializing effects of life-cycle transitions (changes in marital status and arrival of children) play an important role in gendered differences in voter turnout. In the empirical analysis, we use a unique panel dataset that covers four elections in a large Italian municipality and merges information on about 370,000 individuals coming from administrative voter rolls, the civil register, and income tax files.

Our findings show that life-cycle transitions lead to sizable and persistent changes in voter turnout that tend to disadvantage women. Using a voter-level DD strategy that controls for individual and age-by-gender fixed effects, we estimate a positive effect of marriage on male turnout, and a negative effect of widowhood on female turnout. We also show that the presence of young infants (0-5 years old) decreases maternal turnout, while children of voting age seem to increase both parents' turnout. We then use a variety of survey data to show that the estimated turnout effects are paralleled by similar patterns of gender-heterogeneous correlations between family structure, political knowledge, and time spent doing family chores. Taken together, our findings show that family norms contribute to the persistence of gender differences in voter turnout.

Our heterogeneous turnout effects have potentially important policy implications. Although our survey data reveal that political leanings are (partially) uncorrelated with family status, changes in family structure may still alter the political composition of the active electorate; for example, by mobilizing relatively right-leaning men upon marriage and by demobilizing relatively left-leaning women when they give birth to children. Although we cannot investigate such implications in our existing data, the resulting imbalance in political representation could in turn affect implemented

policies. For example, under-representation of mothers of young children may reduce support for public expenditure on child-care, with possibly self-reinforcing negative effects on women's political participation.

Finally, we believe that many of our insights travel to other advanced democracies. The northern Italian setting seems suitable to generalize about the consequences of gendered social norms for the composition of the electorate given that civic traditions are strong (Putnam, Leonardi and Nanetti, 1994) and average levels of turnout are high. Moreover, our findings demonstrate the potential of combining credible identification strategies with large-scale administrative data for further research on the enduring gender gap in political participation.

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Online Appendix

Descriptive Statistics

Table A1 reports summary statistics for the long version of the Bologna data. Columns 1 through 3 refer to the full sample, which consists of all eligible voters independently of their position within the household. Columns 4 through 6 refer to the household head samples; that is, the subset of eligible voters who are either heads of households or spouses thereof. Consistent with its high level of social capital, Bologna has high but declining voter participation: voter turnout in European elections decreased from 84.3 percent to 79.1 percent from 2004 to 2009, and a similar decline affected higher-salience political elections (from 85.2 percent in 2008 to 80.8 percent in 2013). On average, men are more likely to vote than women by 1-to-3 p.p., depending on the sample and election year. A majority of eligible voters are married (52.6 percent), while slightly less than a third (31.4 percent) have never been married. The longer life expectancy of women is reflected in their higher mean age (55.8 vs. 51.9 for men) as well as in the noticeable share of widows (18.7 percent). Bologna is a relatively affluent city, which shows in an average income of 25,483 euros and 6,004 euros of income taxes paid across the four elections in the sample.

Figure A1 describes the income-turnout relationship among non-married (top panel) and married voters (bottom panel). Specifically, each plot compares average turnout by 50 quantiles of OECD-modified household income for female (in orange) and male voters (in blue). Despite its descriptive nature, the pattern in the plots is consistent with the main marriage findings; that is, in each income quantile, non-married females are more likely to vote than males. By contrast, this turnout gender gap disappears (or even reverses sign) among married voters.

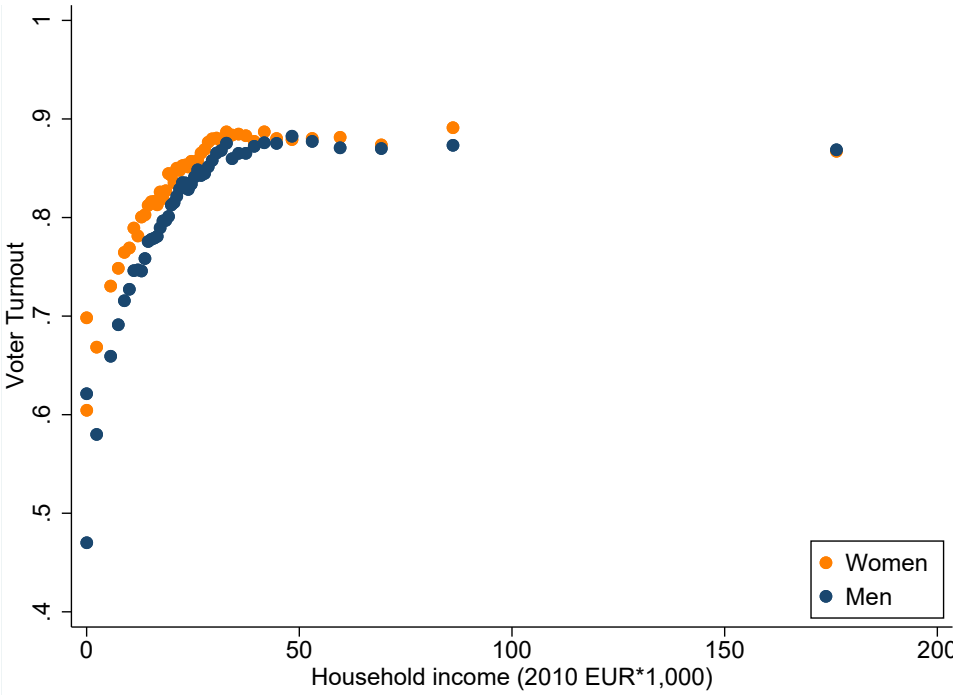
Table A1: Summary Statistics of Bologna Socio-Demographic and Turnout Data **(For Online Publication)**

	Full Sample			Children Sample		
	All	Women	Men	All	Women	Men
	(1)	(2)	(3)	(4)	(5)	(6)
Voted in year…						
2004	.843	.831	.856	.847	.834	.864
2008	.852	.846	.860	.858	.849	.868
2009	.791	.783	.800	.797	.788	.809
2013	.808	.794	.823	.813	.798	.833
Marital status:						
Never married	.314	.278	.357	.225	.205	.249
Married	.526	.489	.569	.599	.544	.668
Divorced	.039	.046	.031	.042	.049	.033
Widowed	.121	.187	.043	.135	.202	.050
Cohabiting kids aged…						
aged 0-5	-	-	-	.069	.066	.072
6-11	-	-	-	.071	.069	.073
12-17	-	-	-	.070	.070	.070
18+	-	-	-	.199	.211	.184
Age	54.0	55.8	51.9	57.2	58.4	55.7
	(19.1)	(19.5)	(18.4)	(17.6)	(17.9)	(17.0)
Income (2010€)	22	17	28	24	18	32
	(38)	(21)	(50)	(40)	(22)	(53)
Income taxes (2010€)	5,257	3,434	7,395	5,739	3,646	8,345
	(14,906)	(7,515)	(20,203)	(15,791)	(7,694)	(21,769)
N	1,163,355	628,043	535,312	978,722	542,795	435,927
N voters	381,257	202,345	178,912	328,025	177,652	150,373

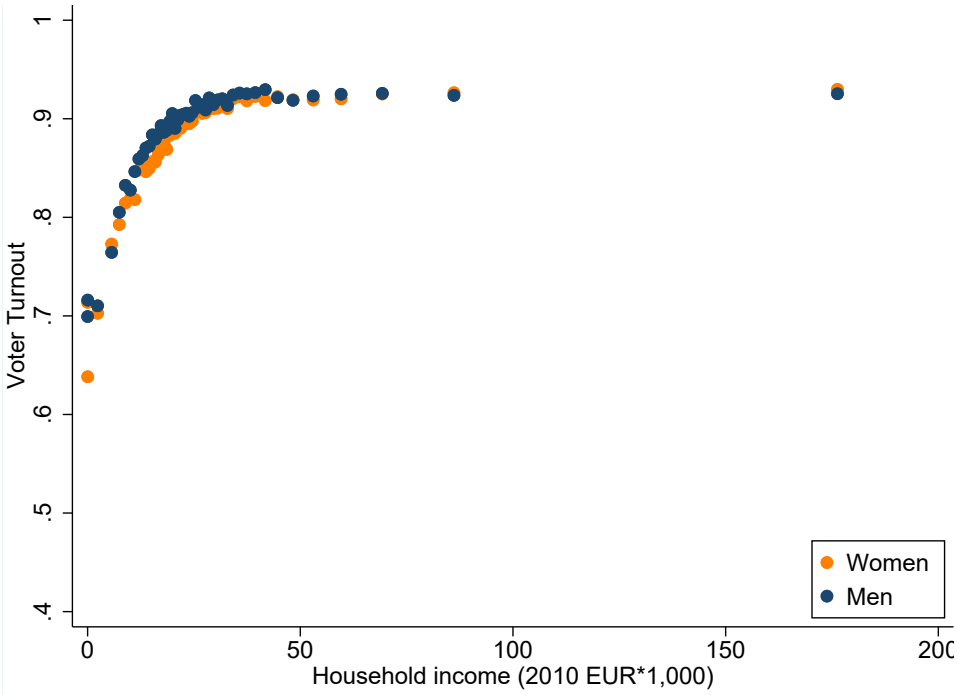
Notes: The table reports sample means and standard deviations in parentheses. Each children sample is a subsample of the corresponding full sample; specifically, children samples are limited to voters whose position within the household is either "Head of Household" or "Spouse/Cohabiting Partner of Head of Household".

Figure A1: Marriage Equalizes Turnout Between Men and Women **(For Online Publication)**

(a) Average Turnout by Income among Non-Married Voters



(b) Average Turnout by Income among Married Voters



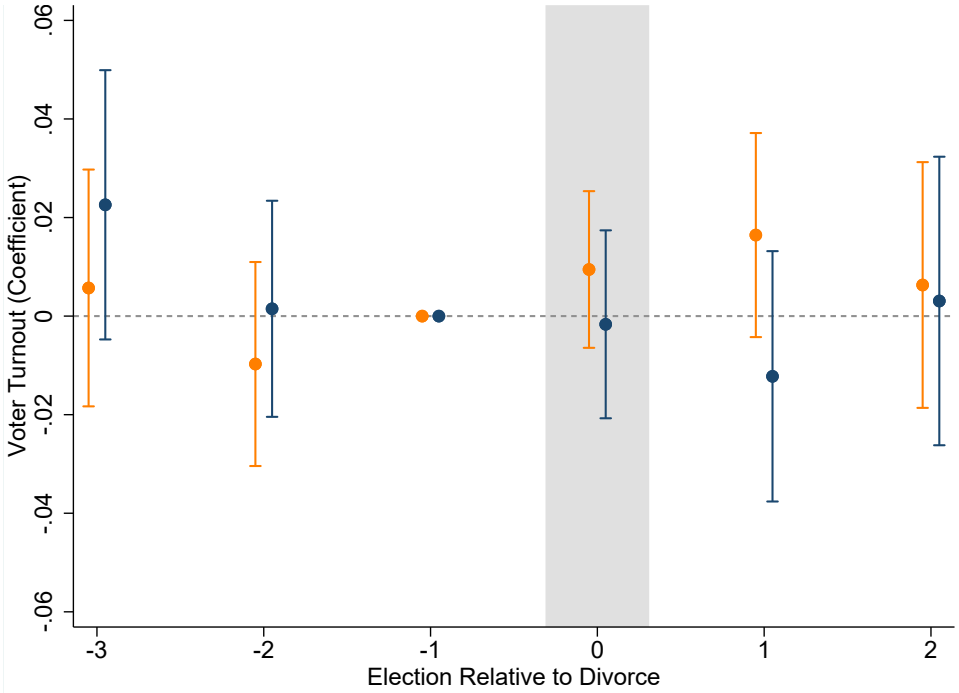
Notes: The Figure plots average turnout by 50 income quantiles (of OECD-modified household income, in 2010 euros) among female (in orange) and male voters (in blue). The top and bottom panels plot the gender-specific income-turnout relationship among non-married and married voters, respectively.

Divorce and Widowhood Event Studies

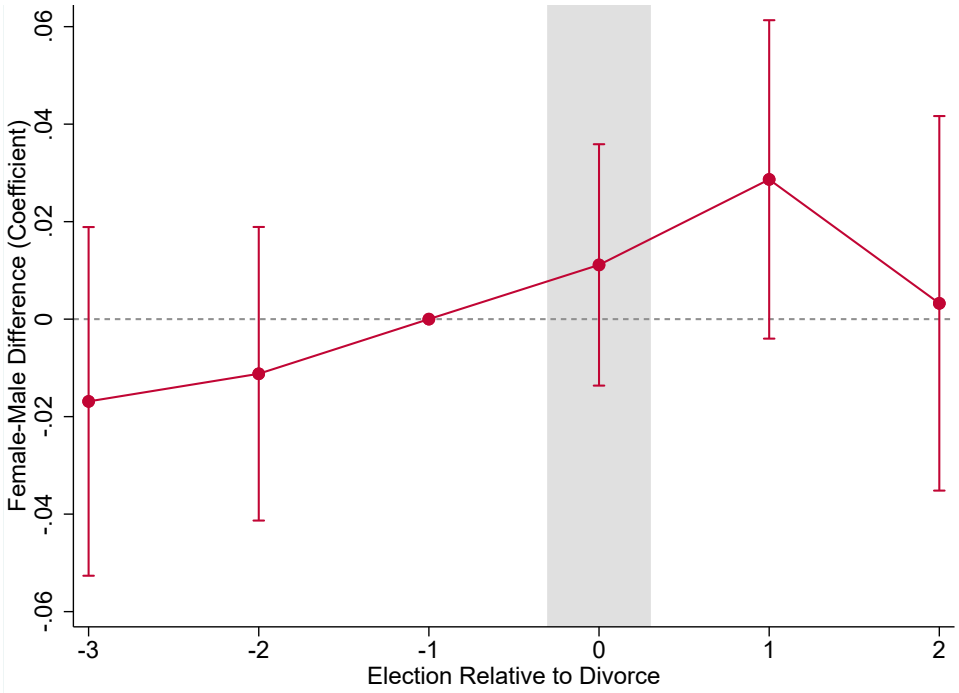
In this section, we plot event-study estimates for divorce and widowhood from regressions like specification 3. Figures A2a and A3a report gender-specific impact estimates on female (in orange) and male (in blue) voters for divorce and widowhood, respectively, while Figures A2b and A3b plot estimated differences in female- minus male-specific effects.

Figure A2: Divorce Event Study **(For Online Publication)**

(a) Orange = Women, Blue = Men



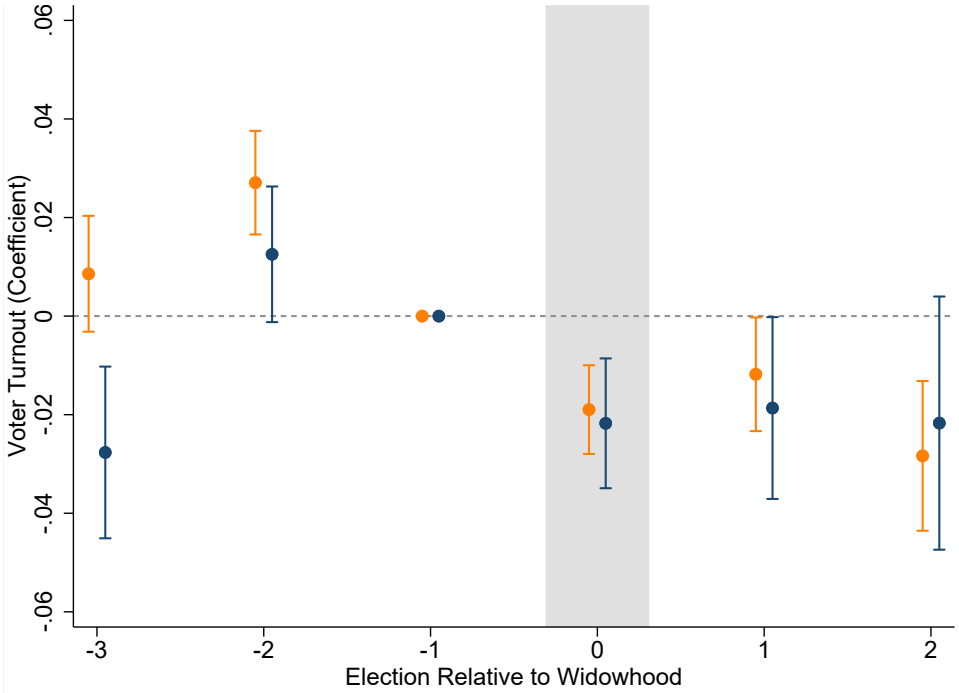
(b) Divorce-Induced Gender Difference in Voter Turnout



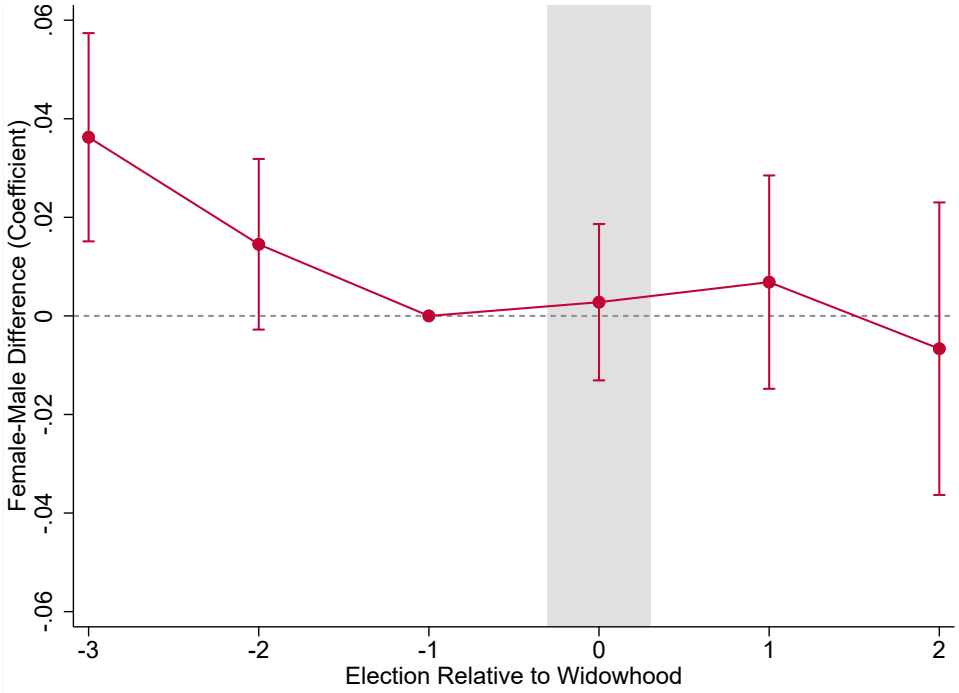
Notes: Panel A plots event-study estimates of the effect of divorce on women's (orange) vs. men's (blue) turnout, along with 95-percent confidence intervals. All estimates are from a unique regression controlling for the same covariates included in Table 1, column 5. The x-axis denotes the election relative to the first election in which a voter's marital status is divorced. Panel B plots differences between female- and male-specific effects.

Figure A3: Widowhood Event Study **(For Online Publication)**

(a) Orange = Widow, Blue = Widower



(b) Widowhood-Induced Gender Difference in Voter Turnout



Notes: Panel A plots event-study estimates of the effect of divorce on women's (orange) vs. men's (blue) turnout, along with 95-percent confidence intervals. All estimates are from a unique regression controlling for the same covariates included in Table 1, column 5. The x-axis denotes the election relative to the first election in which a voter's marital status is divorced. Panel B plots differences between female- and male-specific effects.

Mechanism: Civic Involvement and Political Knowledge

Here we explore whether having children may increase voter turnout is by stimulating parental civic involvement (Wolfinger and Wolfinger, 2008). To do so, we use 2001, 2006, 2008, and 2013 ITANES pooled cross-sectional survey data. On one hand, these data lack of a longitudinal dimension, which rules out the possibility of exploiting within-individual, across-time variation in family structure. On the other hand, the data offer a rich set of socio-demographic controls (e.g., education, employment status, religiosity), which may attenuate the omitted variable bias that likely affects cross-sectional estimates of the effect of family structure on political knowledge. On the whole, estimates from ITANES data should be interpreted as suggestive, rather than definitive, and they do not allow us to estimate effects of children of different ages.

The first six columns of Table A2 report estimates from regressions of dummy variables identifying correct responses to survey questions on factual political knowledge.[9] Following Kling, Liebman and Katz (2007), the outcome for column 7 is a summary index of political knowledge, defined as the equally weighted average of the z-scores of the outcomes from columns 1–6. All regressions use survey weights and control linearly for age (alone and interacted by gender),[10]

[9]The six questions are: "Do you know who elects the President of the Republic?", "Can you tell me the name of the Minister of External Affairs in charge during the last elections?", "Do you know the name of the President of the Chamber of Deputies in charge during the last elections?", "Do you know, approximately, how many representatives sit in the Chamber of Deputies?", "Do you know the name of the Prime Minister in charge during the last elections?", "How many years does the President of the Republic stay in office?". The exact wording of the original questions (in Italian) features minor differences across survey years.

[10]We demean age by gender, so the coefficient on female should be interpreted as the (conditional) difference in outcomes between average-aged women and men.

gender-specific dummies for divorce and widowhood, survey year and wave, as well as indicators for size of the city of residence, region of residence, education, father's education, employment status, and intensity of religious beliefs.

Men with kids are indeed more politically knowledgeable than their childless counterparts. They are significantly more likely to know who elects the President of the Republic and to recall the names of the Minister of Foreign Affairs, of the President of the Chamber of Deputies, and of the Prime Minister. The higher z-score of men with children also confirms their better political knowledge. Conversely, and perhaps unsurprising in view of the zero effect of children aged 5–16 on mothers' turnout (and the negative effect of younger children), the presence of children does not correlate with mothers' political knowledge.

Because of the positive effect on fathers and the zero effect on mothers, children seem to differentially affect men's and women's political information. For two out of eight outcomes reported in Table A2, impact estimates on men are significantly larger than corresponding estimates on women, and a third difference is marginally significant. Conversely, marital status does not correlate with better political knowledge, which is possibly consistent with marriage "leveling up" men's turnout to the higher level of participation of their wives without affecting either spouse's civic involvement.[11]

What information channels drive gender differences in political knowledge? To answer this question, Table A3 explores correlations between family structure and the self-reported use of different media channels to obtain information on the most recent political election. We distinguish between channels that plausibly require more ("hard info") vs. less ("easy info") time and effort by voters. Hard-info channels are: internet, radio, TV programs and news, newspapers, and participation in campaign meetings. We classify the following as easy-info channels: TV ads, campaign

[11] Yet, it is striking that, even controlling for the rich set of covariates included in these regressions, (average-aged) women are less knowledgeable than men.

Table A2: Factual Political Knowledge by Gender and Family Status(**For Online Publication**)

	Correctly Names…						
	How President Is Elected	Minister of Foreign Affairs	President of Chamber of Deputies	Number of Deputies	Prime Minister	President's Term Length	Sum of z-scores
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
1(Female)	-.115 **	-.136 **	-.141 **	-.103 **	-.135 **	-.023	-1.080 **
	(.021)	(.023)	(.023)	(.021)	(.024)	(.060)	(.134)
1(Has kids & female)	.006	-.036 ~	.011	-.005	.003	-.026	-.079
	(.017)	(.020)	(.020)	(.014)	(.022)	(.038)	(.099)
1(Has kids & male)	.047 **	.053 *	.047 *	-.022	.047 *	.028	.290 **
	(.016)	(.022)	(.021)	(.017)	(.019)	(.035)	(.099)
1(Married & female)	-.035	.015	-.021	.000	.022	.016	-.035
	(.022)	(.024)	(.024)	(.018)	(.027)	(.049)	(.129)
1(Married & male)	-.031	.009	.015	.004	-.020	.057	-.023
	(.021)	(.026)	(.025)	(.023)	(.022)	(.048)	(.133)
$\beta^{\text{female w/ kids}} - \beta^{\text{male w/ kids}}$	-.041 ~	-.088 **	-.036	.016	-.044	-.055	-.369 **
	(.024)	(.030)	(.029)	(.022)	(.029)	(.052)	(.139)
$\beta^{\text{married female}} - \beta^{\text{married male}}$	-.004	.006	-.036	-.004	.042	-.040	-.012
	(.030)	(.035)	(.035)	(.029)	(.035)	(.067)	(.184)
$\bar{Y}$	.616	.474	.520	.169	.769	.718	-.000
N	11,701	10,209	10,209	8,709	7,217	1,492	11,701

Notes: All regressions also control for age (alone and interacted with gender) and dummies for size of city of residence, region of residence, education, father's education, intensity of religious beliefs, survey year and wave, as well as gender-specific dummies for divorce and widowhood. Heteroskedasticity-robust standard errors are reported in parentheses.
** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

leaflets, and campaign posters/signs.

Men with children are significantly more likely than men without children to acquire political information from hard-info channels like TV programs and newspapers, while women with and without children are equally likely to seek hard info. Though at the limit of statistical significance, the ensuing gender differences (i.e., $\beta^{\text{female w/ kids}} - \beta^{\text{male w/ kids}}$) suggest that having children induces men, but not women, to spend more time listening to political TV programs and reading newspapers, a fact that possibly explains the heterogeneous effects in political knowledge documented in Table A2.

Overall, children seem to induce women to acquire more information on political campaigns. However, unlike men, women with children appear to rely on uninformative channels like TV ads, campaign leaflets, and campaign posters. Again supporting the hypothesis that marriage merely equalizes turnout across the two spouses without affecting their levels of political interest, virtually all correlations between marital status and information acquisition fall short of statistical significance.

Above we showed that children aged 0–5 reduce maternal but not paternal voter turnout. Unfortunately, the ITANES data do not contain information on the age of children, so we cannot test directly if younger children also depress maternal political knowledge. We can, however, limit the sample to relatively young respondents, who are more likely to be new parents. We take this approach in Appendix Tables A4 and A5, which replicate prior results restricting the sample to respondents aged 50 or younger. Although low statistical power rules out clear-cut conclusions (and the evidence from Table A5 is nuanced), estimates from Table A4 are broadly in line with the gender heterogeneity documented in the case of turnout. That is, younger kids seem to reduce their mothers' political knowledge, but not their fathers'.

Table A3: Acquisition of Political Information by Gender and Family Status(**For Online Publication**)

	Info from Internet	Info from Radio	Info from TV	Info from News-papers	Info from Campaign Meetings	Sum of z-scores (Hard Info)
	(1)	(2)	(3)	(4)	(5)	(6)
1(Female)	-.024	-.044 *	-.069 **	-.081 **	-.059 **	-.675 **
	(.015)	(.022)	(.022)	(.024)	(.017)	(.149)
1(Has kids & female)	-.001	-.012	.037	.028	-.003	.101
	(.012)	(.019)	(.023)	(.022)	(.014)	(.132)
1(Has kids & male)	.022	-.005	.066 **	.090 **	.025	.471 **
	(.016)	(.024)	(.021)	(.023)	(.019)	(.155)
1(Married & female)	-.042 **	-.020	-.018	-.050 ~	-.021	-.393 *
	(.016)	(.023)	(.027)	(.027)	(.017)	(.162)
1(Married & male)	-.015	.051 ~	-.011	-.013	-.006	.000
	(.019)	(.027)	(.024)	(.027)	(.023)	(.183)
$\beta^{\text{female w/ kids}} - \beta^{\text{male w/ kids}}$	-.024	-.007	-.029	-.062 ~	-.029	-.370 ~
	(.020)	(.030)	(.031)	(.032)	(.024)	(.203)
$\beta^{\text{married female}} - \beta^{\text{married male}}$	-.028	-.072 *	-.007	-.037	-.015	-.393
	(.025)	(.035)	(.035)	(.038)	(.028)	(.244)
$\bar{Y}$	.100	.245	.765	.541	.142	.000
N	9,581	9,581	9,581	9,581	9,581	9,581

	Info from TV ads	Info from Campaign Leaflets	Info from Campaign Posters	Sum of z-scores (Easy Info)		Sum of z-scores (All Info)
	(7)	(8)	(9)	(10)		(11)
1(Female)	-.088 *	-.089 *	-.115 **	-.616 **		-.892 **
	(.039)	(.042)	(.039)	(.193)		(.181)
1(Has kids & female)	.069 ~	.083 *	.110 **	.552 **		.375 *
	(.039)	(.038)	(.040)	(.190)		(.181)
1(Has kids & male)	.024	.044	.044	.233		.559 **
	(.038)	(.041)	(.039)	(.183)		(.185)
1(Married & female)	-.038	.022	-.086 ~	-.217		-.516 *
	(.044)	(.046)	(.047)	(.222)		(.211)
1(Married & male)	-.035	.013	-.066	-.186		-.082
	(.043)	(.048)	(.043)	(.208)		(.216)
$\beta^{\text{female w/ kids}} - \beta^{\text{male w/ kids}}$	.046	.039	.066	.318		-.184
	(.054)	(.056)	(.056)	(.262)		(.258)
$\beta^{\text{married female}} - \beta^{\text{married male}}$	-.003	.008	-.020	-.031		-.434
	(.062)	(.066)	(.063)	(.303)		(.300)
$\bar{Y}$	.683	.628	.654	.000		.000
N	3,201	3,201	3,201	3,201		9,581

Notes: All regressions also control for age (alone and interacted with gender) and dummies for size of city of residence, region of residence, education, employment status, father's education, intensity of religious beliefs, survey year and wave, as well as gender-specific dummies for divorce and widowhood. Heteroskedasticity-robust standard errors are reported in parentheses.
** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

Table A4: Factual Political Knowledge by Gender and Family Status: Respondents 50 or Younger **(For Online Publication)**

	Correctly Names…						
	How President Is Elected	Minister of Foreign Affairs	President of Chamber of Deputies	Number of Deputies	Prime Minister	President's Term Length	Sum of z-scores
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
1(Female)	-.060	-.155 *	-.144 *	-.003	.016	-.057	-.755 ~
	(.067)	(.072)	(.069)	(.065)	(.065)	(.190)	(.401)
1(Has kids & female)	-.043	-.067 ~	-.068 ~	-.052 *	-.087 **	-.129 ~	-.386 *
	(.030)	(.036)	(.035)	(.025)	(.034)	(.066)	(.172)
1(Has kids & male)	-.002	-.007	-.024	-.049 ~	.004	.027	-.211
	(.028)	(.040)	(.036)	(.028)	(.033)	(.063)	(.172)
1(Married & female)	-.059 ~	-.039	-.041	-.023	.008	.067	-.437 *
	(.033)	(.037)	(.035)	(.027)	(.034)	(.078)	(.186)
1(Married & male)	-.040	-.013	.010	.028	.009	.019	-.025
	(.030)	(.040)	(.037)	(.032)	(.033)	(.075)	(.200)
$\beta^{\text{female w/ kids}} - \beta^{\text{male w/ kids}}$	-.041	-.060	-.044	-.002	-.091 ~	-.155 ~	-.175
	(.038)	(.053)	(.049)	(.035)	(.046)	(.090)	(.228)
$\beta^{\text{married female}} - \beta^{\text{married male}}$	-.019	-.026	-.052	-.050	-.001	.048	-.412
	(.044)	(.054)	(.051)	(.041)	(.047)	(.105)	(.266)
$\bar{Y}$	.622	.428	.479	.182	.747	.688	-.157
N	4,294	3,800	3,800	3,323	2,829	494	4,294

Notes: All regressions also control for age (alone and interacted with gender) and dummies for size of city of residence, region of residence, education, employment status, father's education, intensity of religious beliefs, survey year and wave, as well as gender-specific dummies for divorce and widowhood. Heteroskedasticity-robust standard errors are reported in parentheses.

** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

Table A5: Acquisition of Political Information by Gender and Family Status: Respondents 50 or Younger (**For Online Publication**)

	Info from Internet	Info from Radio	Info from TV	Info from News-papers	Info from Campaign Meetings	Sum of z-scores (Hard Info)
	(1)	(2)	(3)	(4)	(5)	(6)
1(Female)	-.047	-.138 *	-.129 ~	-.153 ~	-.018	-1.141 *
	(.056)	(.068)	(.070)	(.079)	(.062)	(.524)
1(Has kids & female)	-.018	-.012	-.020	-.063	-.057 *	-.424
	(.027)	(.033)	(.038)	(.041)	(.027)	(.262)
1(Has kids & male)	.066 *	-.093 *	.009	.021	-.029	-.016
	(.033)	(.043)	(.038)	(.042)	(.037)	(.313)
1(Married & female)	-.042	-.029	.012	-.015	-.037	-.316
	(.031)	(.034)	(.039)	(.043)	(.030)	(.284)
1(Married & male)	-.054	.063	-.001	-.024	.011	-.050
	(.033)	(.043)	(.037)	(.042)	(.037)	(.319)
$\beta^{\text{female w/ kids}} - \beta^{\text{male w/ kids}}$	-.084 *	.081	-.029	-.085	-.028	-.409
	(.042)	(.054)	(.054)	(.059)	(.046)	(.405)
$\beta^{\text{married female}} - \beta^{\text{married male}}$	.011	-.092 ~	.013	.009	-.048	-.266
	(.045)	(.055)	(.054)	(.060)	(.048)	(.428)
$\bar{Y}$	.159	.247	.768	.574	.164	.335
N	3,561	3,561	3,561	3,561	3,561	3,561

	Info from TV ads	Info from Campaign Leaflets	Info from Campaign Posters	Sum of z-scores (Easy Info)		Sum of z-scores (All Info)
	(7)	(8)	(9)	(10)		(11)
1(Female)	-.153	.053	-.097	-.421		-1.332 *
	(.126)	(.132)	(.124)	(.594)		(.622)
1(Has kids & female)	-.024	.002	.076	.111		-.387
	(.058)	(.063)	(.063)	(.298)		(.329)
1(Has kids & male)	-.029	.023	-.061	-.143		-.090
	(.062)	(.068)	(.061)	(.284)		(.363)
1(Married & female)	.006	.039	-.101	-.119		-.358
	(.058)	(.064)	(.065)	(.296)		(.350)
1(Married & male)	-.030	.014	.010	-.015		-.056
	(.060)	(.068)	(.060)	(.281)		(.370)
$\beta^{\text{female w/ kids}} - \beta^{\text{male w/ kids}}$	.005	-.021	.136	.254		-.297
	(.084)	(.092)	(.087)	(.409)		(.485)
$\beta^{\text{married female}} - \beta^{\text{married male}}$	.036	.025	-.111	-.104		-.302
	(.084)	(.093)	(.088)	(.409)		(.508)
$\bar{Y}$	.730	.653	.705	.260		.431
N	1,316	1,316	1,316	1,316		3,561

Notes: All regressions also control for age (alone and interacted with gender) and dummies for size of city of residence, region of residence, education, employment status, father's education, intensity of religious beliefs, survey year and wave, as well as gender-specific dummies for divorce and widowhood. Heteroskedasticity-robust standard errors are reported in parentheses.
** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

Mechanism: Hours Worked

We now investigate how family composition affects the quantity and type of time available to follow politics. To this end, we explore correlations between family status and hours worked using 2005–2012 pooled cross-sectional data from the Italian National Statistical Agency. The so-called ISTAT AVQ data have similar shortcomings and strengths as the ITANES data. That is, they lack a panel dimension—which rules out DD specifications—but they contain detailed socio-demographic controls—which mitigate concerns about omitted variable bias. For this reason, the caveat remains the same, in that we treat evidence from ISTAT AVQ data as suggestive. Relative to the ITANES data, however, the ISTAT data have the advantage of providing the approximate age of cohabiting children.[12] We can therefore directly compare impact estimates on worked time by age of children with the corresponding turnout effects.

Table A6 reports estimates from three separate regressions. The outcomes for columns 1 and 2 are hours worked at home and hours of paid work, respectively. The outcome for column 3 is total worked hours; that is, hours worked at home plus hours of paid work. All regressions control for education, region of residence, year of interview, as well as gender-specific dummies for age,[13] divorce, and widowhood.[14]

Women with children report more hours of domestic work and less paid work than women without children. The net impact on female total hours of work is positive, particularly in presence

[12]Specifically, the raw data contain counts of children in the following age ranges: 0–5, 6–13, 14–17, and 18+.

[13]The exact respondent's age is not reported in the data. It is instead binned in the following intervals: 18–19, 20–24, 25–29, ..., 60–64, 65–74, 75+.

[14]Though the data contain employment status, we opted to exclude that control from these regressions. It seems indeed very likely that changes in worked hours are largely driven by changes in employment status.

Table A6: Children, Marital Status and Hours Worked **(For Online Publication)**

	Hours Worked at Home (1)		Hours of Paid Work (2)		Total Hours Worked (3)	
1(Female)	8.14	**	-1.69	**	6.46	**
	(.26)		(.37)		(.43)	
1(Female w/ children aged 0-5)	11.58	**	-4.58	**	7.01	**
	(.50)		(.38)		(.50)	
1(Male w/ children aged 0-5)	3.67	**	1.13	**	4.80	**
	(.25)		(.34)		(.38)	
1(Female w/ children aged 6-13)	5.97	**	-3.79	**	2.18	**
	(.34)		(.29)		(.36)	
1(Male w/ children aged 6-13)	.68	**	1.27	**	1.95	**
	(.16)		(.27)		(.29)	
1(Female w/ children aged 14-17)	4.45	**	-2.01	**	2.43	**
	(.68)		(.59)		(.72)	
1(Male w/ children aged 14-17)	-.37		.91		.54	
	(.33)		(.61)		(.64)	
1(Female w/ children aged 18+)	3.91	**	-1.00	**	2.90	**
	(.23)		(.20)		(.24)	
1(Male w/ children aged 18+)	-1.43	**	2.75	**	1.32	**
	(.12)		(.24)		(.24)	
1(Married & female)	14.01	**	-4.59	**	9.42	**
	(.15)		(.16)		(.19)	
1(Married & male)	-.37	**	5.36	**	4.99	**
	(.09)		(.16)		(.18)	
$\beta^{0\text{-}5\ female} - \beta^{0\text{-}5\ male}$	7.92	**	-5.71	**	2.21	**
	(.53)		(.51)		(.55)	
$\beta^{6\text{-}13\ female} - \beta^{6\text{-}13\ male}$	5.29	**	-5.06	**	.23	
	(.36)		(.39)		(.42)	
$\beta^{14\text{-}17\ female} - \beta^{14\text{-}17\ male}$	4.82	**	-2.93	**	1.89	*
	(.74)		(.82)		(.84)	
$\beta^{18+\ female} - \beta^{18+\ male}$	5.34	**	-3.75	**	1.58	**
	(.25)		(.30)		(.32)	
$\beta^{married\ female} - \beta^{married\ male}$	14.37	**	-9.95	**	4.42	**
	(.17)		(.22)		(.24)	
$\bar{Y}$	16.12		19.88		36.00	
N	269,030		269,030		269,030	

Notes: All regressions control for education, region of residence, year of interview, as well as gender-specific dummies for age, divorce, and widowhood. Standard errors clustered by household are reported in parentheses.
** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

of kids aged 0–5 (+7.01 hours). Children 0–13 also appear to increase male domestic workload, but to a much lesser extent than they increase women's. The presence of older children (ages 14–17 and 18+) does not correlate, or even correlates negatively, with male domestic work. Finally, men with children of any age report more hours of paid work than their childless counterparts.

Gender differences in correlational estimates point to children inducing a disproportionate increase in the total hours worked by women, an increase that is particularly marked for children aged 0–5 ($\beta^{0to5,female} - \beta^{0to5,male} = 2.21$ hours). This finding corroborates, at least partly, the hypotheses that younger children limit the amount of time their mothers have to follow politics and to vote.

Neither the "maternal time-constraint" nor the "specialization" effects can, however, fully account for gender differences in turnout effects. In fact, like children, marriage also appears to disproportionately affect women, by increasing the number of hours worked at home and reducing those of paid labor. But Table 2 showed that marriage does not affect female turnout; if anything, it merely increased men's voter participation to the pre-marriage voter turnout of their wives.

Mechanism: Income Changes

The Bologna data do not contain information on hours worked at home or hours of paid work. Yet, the data contain information on individual-level income, which is arguably a good proxy for employment. Therefore, we now report DD impact estimates of changes in family structure on individual income to complement the cross-sectional analyses of hours worked based on ISTAT data.

Table A7 reports gender-specific impact estimates of children on income. Children aged 5 or younger sharply and significantly reduce maternal income: the estimated effect is around -1,800 euros, which is approximately 10% of the average gross income of women in the sample. The effect remains statistically and economically significant in presence of children aged 6–11 (between -891 and -865 euros) and of children aged 12–17 (between -439 and -397), thus suggesting that some women never return to work after becoming mothers. The effect becomes positive and significant in presence of children aged 18 or older (between 250 and 300 euros). However, this positive effect is most likely driven by older women who experience income increases (e.g., start receiving disability benefits) at the same time as their adult children move back with them, rather than by women who resume working after their children move out.

By contrast, there is no significant relation between the presence of children aged 5 or younger on paternal income, while the relation becomes positive and significant in presence of older children. Consistent with the large children-induced gaps in hours of paid work documented using the ISTAT data, children generate startling gender gaps in income. Specifically, the income gender gap widens by around 2,200 euros in presence of children aged 0–5 (i.e., about 1/5 of the average income gender gap in the sample), by around 2,000 euros with children aged 6–11, and 1,700 euros with children aged 12–17. The effect of children aged 18+ on the gender income gap is insignificant, though this null effect is likely driven by old voters whose children move back with them (i.e., by voters who switch from having no cohabiting children to having cohabiting children 18+).

Table A7: Effect of Children on Income by Kids' Age and Voter's Gender **(For Online Publication)**

	Outcome: Voter-Level Income							
	(1)		(2)		(3)		(4)	
1(Female w/ children aged 0-5)	-1,853	**	-1,879	**	-1,879		-1,867	**
	(129)		(130)		(130)		(130)	
1(Male w/ children aged 0-5)	350		335		339		297	
	(348)		(347)		(347)		(375)	
1(Female w/ children aged 6-11)	-865	**	-890	**	-891		-891	**
	(128)		(128)		(128)		(128)	
1(Male w/ children aged 6-11)	1,099	**	1,072	**	1,073		1,066	**
	(405)		(405)		(405)		(410)	
1(Female w/ children aged 12-17)	-397	**	-416	**	-417		-439	**
	(121)		(122)		(122)		(122)	
1(Male w/ children aged 12-17)	1,275	*	1,269	*	1,270		1,269	*
	(532)		(531)		(531)		(535)	
1(Female w/ children aged 18+)	288	**	295	**	296		254	**
	(91)		(91)		(91)		(90)	
1(Male w/ children aged 18+)	684	*	695	*	696		691	*
	(342)		(342)		(342)		(341)	
$\beta^{0\text{-}5\ female} - \beta^{0\text{-}5\ male}$	-2,204	**	-2,213	**	-2,218		-2,163	**
	(369)		(369)		(369)		(395)	
$\beta^{6\text{-}11\ female} - \beta^{6\text{-}11\ male}$	-1,964	**	-1,962	**	-1,964		-1,957	**
	(421)		(421)		(421)		(426)	
$\beta^{12\text{-}17\ female} - \beta^{12\text{-}17\ male}$	-1,672	**	-1,685	**	-1,687		-1,708	**
	(544)		(544)		(544)		(548)	
$\beta^{18+\ female} - \beta^{18+\ male}$	-396		-400		-400		-438	
	(351)		(351)		(351)		(350)	
Voter FEs	✓		✓		✓		✓	
Age×Gender FEs	✓		✓		✓		✓	
Election FEs	✓		✓		✓		✓	
Neighborhood controls			✓		✓		✓	
Household controls					✓		✓	
Marital status×Gender FEs							✓	
Female w/o kids $\bar{Y}$	17,901		17,901		17,901		17,901	
Male w/o kids $\bar{Y}$	28,852		28,852		28,852		28,852	
N	883,208		883,208		883,208		883,208	

Notes: The sample is limited to married individuals. Neighborhood controls are: precinct-year average age and OECD modifed gross household income, as well as shares of female and Italian residence, and city neighborhood-by-year fixed effects. Household controls are the share of household members who are Italian citizens. Standard errors are two-way clustered by voter and household.

** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

Table A8: Effect of Marital Status on Income **(For Online Publication)**

	Outcome: Voter-Level Income							
	(1)		(2)		(3)		(4)	
1(Married female)	-0.315	~	-0.291	~	-0.286		0.230	
	(.174)		(.175)		(.175)		(.177)	
1(Married male)	1.437	*	1.466	*	1.485	*	1.323	*
	(.611)		(.610)		(.615)		(.650)	
1(Divorced female)	-0.028		0.013		0.017		0.498	~
	(.285)		(.285)		(.286)		(.286)	
1(Divorced male)	1.986	*	2.011	*	2.025	*	1.977	*
	(.898)		(.897)		(.904)		(.924)	
1(Widowed female)	10.720	**	10.712	**	10.717	**	11.225	**
	(.217)		(.218)		(.218)		(.220)	
1(Widowed male)	5.447	**	5.443	**	5.459	**	5.290	**
	(.762)		(.762)		(.765)		(.791)	
$\beta^{\text{married female}} - \beta^{\text{married male}}$	-1.752	**	-1.758	**	-1.770	**	-1.093	~
	(.605)		(.606)		(.608)		(.642)	
$\beta^{\text{divorced female}} - \beta^{\text{divorced male}}$	-2.014	*	-1.998	*	-2.008	*	-1.479	
	(.923)		(.923)		(.927)		(.945)	
$\beta^{\text{widowed female}} - \beta^{\text{widowed male}}$	5.273	**	5.269	**	5.258	**	5.935	**
	(.770)		(.770)		(.771)		(.796)	
Voter FEs	✓		✓		✓		✓	
Age×Gender FEs	✓		✓		✓		✓	
Election FEs	✓		✓		✓		✓	
Neighborhood controls			✓		✓		✓	
Household controls					✓		✓	
Children×Gender FEs							✓	
Never-married female $\bar{Y}$	16		16		16		16	
Never-married male $\bar{Y}$	18		18		18		18	
N	1,084,202		1,084,202		1,084,202		1,084,202	

Notes: Neighborhood controls are: precinct-year average age, income, and income taxes paid, as well as shares of female and Italian residence, and city neighborhood-by-year fixed effects. Household controls are the share of household members who are Italian citizens, average income across adult household members, and average income taxes paid. Children FEs are four dummies indicating presence of one or more children of the following ages: 0-5, 6-11, 12,-17, 18+. Standard errors are two-way clustered by voter and household.
** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

Mechanism: Cohabitation

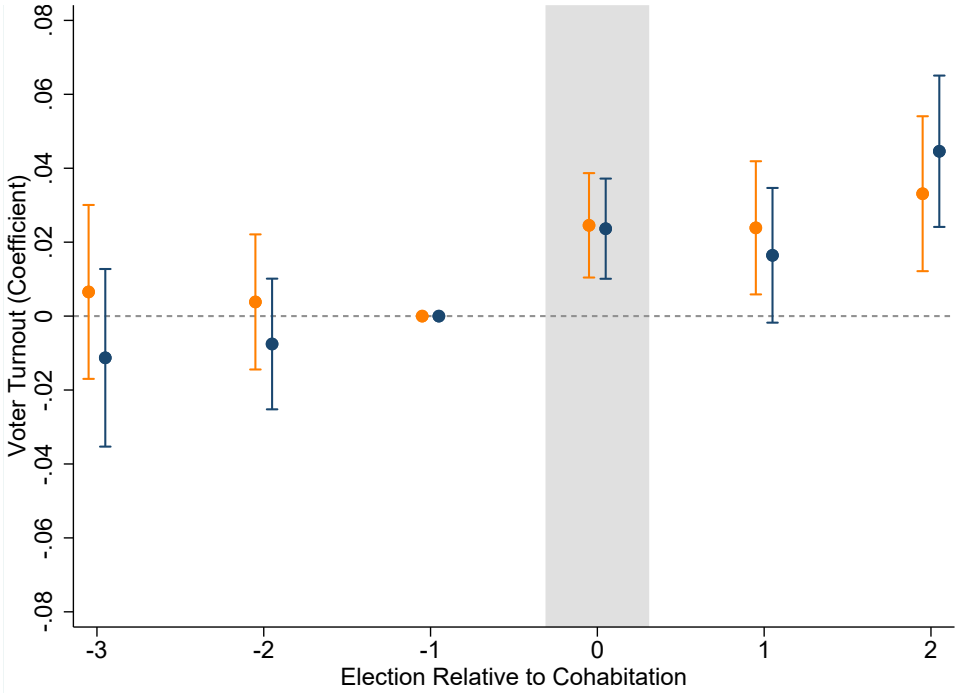
We now examine the turnout effect of cohabitation using event-study graphs. Because the Bologna data record cohabitation as a type of relationship to the head of the household (and not as a possible category of marital status), identifying couples of cohabiting partners is complicated. For example, the civil register may record cohabiting partners as heads of two separate households. In other cases, the civil register may record roommates who are simply sharing an apartment as "cohabiting" (i.e., assigning them a unique household identifier and making one of them the head of the household).[15] With this important limitation in mind, we define cohabiting couples as consisting of unmarried heads of households cohabiting with an unmarried individual of the opposite gender (where "cohabitation" status is determined from the relationship to the head of the household). The resulting sample consists of approximately 10,000 voters per election.

Figure A4a reports gender-specific estimates and 95-percent confidence intervals from an event-study regression (equation 3). As evidenced by the jump in voter participation occurring at $\tau = 0$, cohabitation increases both men's and women's turnout by about 2 percentage points. Unsurprisingly, the female-minus-male differences in event-study coefficients (i.e., $\beta_{\tau}^{cohabit,female} - \beta_{\tau}^{cohabit,male}$, plotted in Figure A4b) are centered around 0 and insignificant. Corroborating a causal interpretation of the results, neither figure shows significant pre-trends. Again, caution is needed in interpreting the results, as our definition of cohabiting couples (1) excludes cohabiting couples whose members are assigned different household identifiers and (2) includes cohabiting roommates of different genders sharing identical household identifiers.

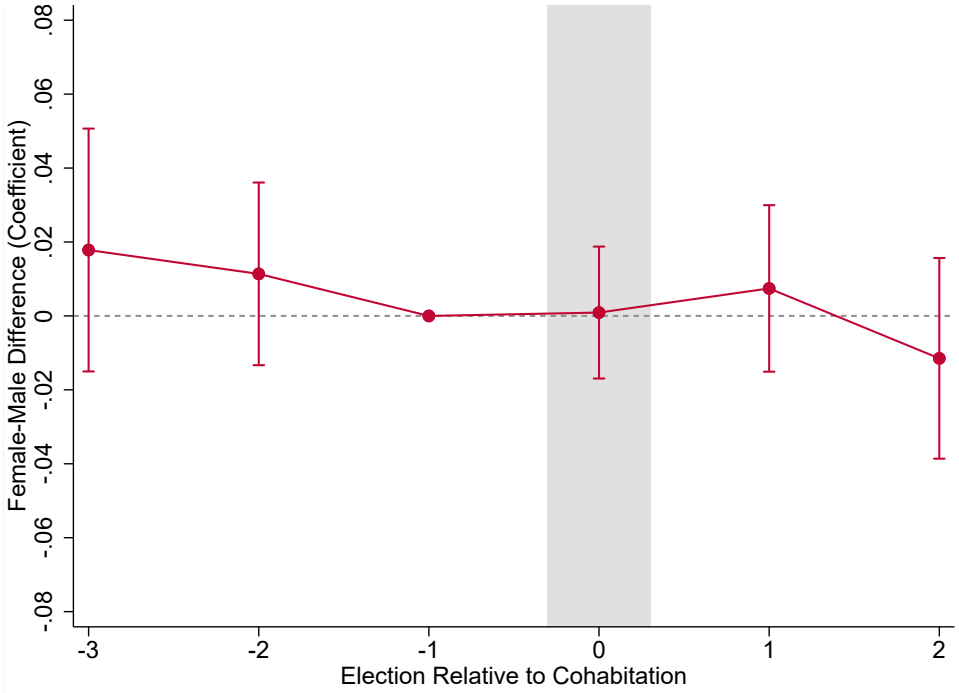
[15] Unfortunately, we could not determine when and why civil register officers record cohabiting individuals as belonging to the same or to different households.

Figure A4: Cohabitation Event Study **(For Online Publication)**

(a) Orange = Women, Blue = Men



(b) Cohabitation-Induced Gender Difference in Voter Turnout



Notes: Panel A plots event-study estimates of the effect of cohabitation on women's (orange) vs. men's (blue) turnout, along with 95-percent confidence intervals. All estimates are from a unique regression controlling for the same covariates included in Table 1, column 5. The x-axis denotes the election relative to the first election in which a voter's household composition status is cohabiting. Panel B plots differences between female- and male-specific effects.

Mechanism: Gender Differences in Political Preferences

To shed light on the effects of family composition on political preferences, we return to the ITANES survey data. Our outcomes are seven measures of political preferences, which we regress on the usual variables representing family status and the same controls used in previous regressions based on ITANES data. Specifically, the independent variables are an index of political ideology (ranging from 1–left–to 10–right), a dummy for having voted for a party in Silvio Berlusconi's coalition in the most recent political election, and the level of agreement with the following five statements: "Abortion should be harder to get," "When jobs are scarce, men should have more right to a job than women," "Drug users should not be punished," "Firms should be freer to hire and fire," and "Immigrants threaten natives' employment." Table A9 reports the results.

Mirroring gender differences that have been documented for other democracies (e.g., Kittilson, 2016), Italian women appear more leftist than men. They are significantly more likely to self-identify as left-leaning, 8.2 p.p. less likely to have voted for Berlusconi, and marginally more likely to disagree with the notion that firms should be freer to hire and fire.[16] Finally, women are less likely to agree with the statement that "When jobs are scarce, men should have more right to a job than women."[17]

Neither the presence of children nor marriage seem systematically related to political lean-

[16]Recall that all regressions control for (demeaned) age interacted with gender. Thus, the coefficient on $1(Female)$ should be interpreted as the (conditional) difference in political leaning between an average-aged woman and an average-aged man. In practice, though, we obtain substantively identical results controlling for age instead of age interacted with gender (results available upon request).

[17]Although, surprisingly, average-aged women are as likely as average-aged men to agree with the statement that "Abortion should be harder to get."

Table A9: Political Views by Gender and Family Status(**For Online Publication**)

	1 = Completely Disagree; 4 = Completely Agree						
	Abortion Should Be Harder to Get	If Jobs Are Scarce Men Should Have Priority	Drug Users Shouldn't Be Punished	Firms Should Be Freer to Hire, Fire	Immigrants Threaten Natives' Employment	1-to-10 Left-Right Index	Voted Berlusconi in Last Election
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
1(Female)	-.018	-.159 **	-.073	-.099 ~	.021	-.286 *	-.082 **
	(.054)	(.061)	(.063)	(.053)	(.051)	(.132)	(.021)
1(Has kids & female)	.014	.027	.013	-.009	.032	.072	.013
	(.045)	(.056)	(.059)	(.052)	(.048)	(.109)	(.015)
1(Has kids & male)	.074	.151 *	.025	.011	.062	.088	.029 ~
	(.045)	(.061)	(.067)	(.057)	(.053)	(.107)	(.015)
1(Married & female)	-.073	.092	-.030	.097	.064	.011	.035 ~
	(.055)	(.066)	(.068)	(.061)	(.057)	(.133)	(.020)
1(Married & male)	-.049	-.063	-.105	-.015	-.038	-.129	-.047 *
	(.058)	(.070)	(.078)	(.064)	(.060)	(.134)	(.020)
$\beta^{\text{female w/ kids}} - \beta^{\text{male w/ kids}}$	-.060	-.124	-.012	-.020	-.030	-.016	-.016
	(.064)	(.083)	(.088)	(.077)	(.071)	(.153)	(.021)
$\beta^{\text{married female}} - \beta^{\text{married male}}$	-.024	.155	.075	.111	.102	.140	.082 **
	(.079)	(.095)	(.103)	(.087)	(.082)	(.187)	(.028)
$\bar{Y}$	2.309	2.636	1.871	2.278	2.418	5.306	.283
N	8,891	6,501	6,376	9,004	9,371	9,522	11,701

Notes: All regressions also control for age (alone and interacted with gender) and dummies for size of city of residence, region of residence, education, employment status, father's education, intensity of religious beliefs, survey year and wave, as well as gender-specific dummies for divorce and widowhood. Heteroskedasticity-robust standard errors are reported in parentheses.

** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

ings.[18] This suggests that, even if changes in family status may not affect voters' political preferences directly, they may still shift the political composition of the *active* electorate (e.g., by mobilizing relatively right-leaning men upon marriage or by demobilizing relatively left-leaning women when they give birth to children).

[18]With the possible exception of voting for Berlusconi, which is positively correlated with marriage in the case of women, and negatively for men.

Robustness: Effects of Children on Turnout Are Unaffected By Further Restricting the Sample to Married Couples

Tables A10 and A11 report estimated effects of children on turnout among married couples. That is, we further restrict the children sample to *married* heads of households and their spouses (if any). Resulting estimates are substantively identical to those of Tables 3 and 4, thus suggesting that the effect of children on turnout does not differ substantively across married and cohabiting couples.

Table A10: Effect of Children on Turnout by Children's Age: Married Couples Only **(For Online Publication)**

	Outcome: Voter-Level Turnout							
	(1)		(2)		(3)		(4)	
1(Children aged 0-5)	-.012	**	-.012	**	-.011	**	-.011	**
	(.003)		(.003)		(.003)		(.003)	
1(Children aged 6-11)	.004		.004		.004		.004	
	(.003)		(.003)		(.003)		(.003)	
1(Children aged 12-17)	.007	**	.007	**	.008	**	.007	**
	(.002)		(.002)		(.002)		(.002)	
1(Children aged 18+)	.013	**	.013	**	.013	**	.013	**
	(.002)		(.002)		(.002)		(.002)	
Voter FEs	✓		✓		✓		✓	
Age×Gender FEs	✓		✓		✓		✓	
Election FEs	✓		✓		✓		✓	
Neighborhood controls			✓		✓		✓	
Household controls					✓		✓	
Income and taxes paid							✓	
No kids $\bar{Y}$	.872		.872		.872		.872	
N	543,705		543,705		543,705		543,705	

Notes: The sample is limited to married individuals. Neighborhood controls are: precinct-year average age, income, and income taxes paid, as well as shares of female and Italian residence, and city neighborhood-by-year fixed effects. Household controls are the share of household members who are Italian citizens and the OECD modified gross household income. Standard errors are two-way clustered by voter and household.

** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

Table A11: Effect of Children on Turnout by Children's Age and Voter's Gender **(For Online Publication)**

	Full Sample		Ever Married		Never Married		Ever Married Parents Born North		Ever Married Parents Born South	
	(1)		(2)		(3)		(4)		(5)	
1(Female w/ children aged 0-5)	-.017	**	-.018	**	-.008		-.016	**	-.029	**
	(.003)		(.004)		(.009)		(.004)		(.006)	
1(Male w/ children aged 0-5)	.001		-.001		.021	*	-.004		.004	
	(.003)		(.003)		(.010)		(.004)		(.006)	
1(Female w/ children aged 6-11)	.004		.002		.019	~	-.002		-.001	
	(.003)		(.003)		(.011)		(.003)		(.006)	
1(Male w/ children aged 6-11)	.010	**	.009	**	.015		.005		.014	*
	(.003)		(.003)		(.015)		(.003)		(.005)	
1(Female w/ children aged 12-17)	.006	*	.004		.021		.000		.008	~
	(.003)		(.003)		(.013)		(.003)		(.005)	
1(Male w/ children aged 12-17)	.013	**	.013	**	.015		.009	**	.014	**
	(.003)		(.003)		(.020)		(.003)		(.005)	
1(Female w/ children aged 18+)	.013	**	.012	**	.035	*	.010	**	.015	**
	(.002)		(.002)		(.015)		(.002)		(.004)	
1(Male w/ children aged 18+)	.011	**	.012	**	.022		.009	**	.020	**
	(.002)		(.002)		(.030)		(.002)		(.004)	
$\beta^{0\text{-}5\ female} - \beta^{0\text{-}5\ male}$	-.018	**	-.017	**	-.029	*	-.011	**	-.032	**
	(.004)		(.004)		(.013)		(.004)		(.007)	
$\beta^{6\text{-}11\ female} - \beta^{6\text{-}11\ male}$	-.006	~	-.008	*	.005		-.007	~	-.015	*
	(.004)		(.004)		(.019)		(.004)		(.006)	
$\beta^{12\text{-}17\ female} - \beta^{12\text{-}17\ male}$	-.007	*	-.008	**	.006		-.009	**	-.006	
	(.003)		(.003)		(.024)		(.003)		(.006)	
$\beta^{18+\ female} - \beta^{18+\ male}$	.001		.0003		.014		.0003		-.006	
	(.003)		(.0027)		(.033)		(.0028)		(.005)	
Voter FEs	✓		✓		✓		✓		✓	
Age×Gender FEs	✓		✓		✓		✓		✓	
Election FEs	✓		✓		✓		✓		✓	
Neighborhood controls	✓		✓		✓		✓		✓	
Household controls	✓		✓		✓		✓		✓	
Income and taxes paid	✓		✓		✓		✓		✓	
Marital status×Gender FEs	✓		✓		✓		✓		✓	
Female w/o kids $\bar{Y}$	.809		.813		.797		.824		.793	
Male w/o kids $\bar{Y}$	.833		.863		.762		.879		.823	
N	902,153		735,398		166,751		616,028		206,174	

Notes: Neighborhood controls are: precinct-year average age, and OECD modified gross household income, as well as shares of female and Italian residence, and city neighborhood-by-year fixed effects. Household controls are the share of household members who are Italian citizens and the OECD modifed gross household income. Marital status FEs are three, mutually exclusive dummies indicating married, divorced, and widowed voters. Standard errors are two-way clustered by voter and household.

** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

Heterogeneity in the Effects of Childrearing

In this appendix section, we explore heterogeneity of childrearing effects along several margins. Table A12 focuses on voters' marital status and geographic origin. Specifically, column 2 reports estimated children effects on ever-married voters;[19] column 3 reports effects on never-married voters; columns 4 and 5 restrict attention to married voters born in Northern and Center-South regions (as defined by ISTAT), respectively; column 1 reports estimates from the full children sample for comparison.

The negative effect of children 5 or younger on maternal turnout is larger on ever-married voters than on never-married ones and on married parents from Center-South regions than on married parents from Northern regions. Since conservative values – particularly on the role of mothers – are arguably stronger among married individuals and among voters from Center-South regions, this pattern of effects corroborates the importance of gendered social norms in explaining our children effects. Comparisons between female- and male-specific impact estimates largely support the same conclusion, as differences between effects on maternal and paternal turnout are typically larger among married voters and voters from the Center-South.

Table A13 explores heterogeneity by voting costs – proxied by distance to the polling place – and in the total number of children a voter had over the sample years. Specifically, column 2 restricts the sample to voters living within 400 meters from their polling place (measured in a straight line), corresponding to approximately the median distance in the sample. Column 3 looks at voters living farther than 400 meters from their polling place. Column 4 restricts the sample to voters who have exactly one kid in at least one election and never have more than one kid. Column 1 reports estimates from the full sample for comparison.

[19]Remember that, in our data, marital status can be: never married, currently married, divorced and not currently married, or widowed and not currently married. Ever-married voters are therefore defined as individuals whose marital status is different from "never married" in at least one election in our data.

All point estimates are substantively in line with those from the full sample, thus suggesting that gender differences in the cost of reaching one's polling location in presence of children are not the main drivers of our gender-heterogeneous effects. Yet, these results should be interpreted with caution, as distance to the polling place could correlate with other voter characteristics (e.g., income) that may attenuate the negative effect of children on the gender turnout gap.

Finally, Table A14 explores heterogeneity by political context. Column 2 restricts the sample to the two European/municipal elections covered by our data (2004 and 2009), while column 3 looks at the two national political elections (2008 and 2013). Columns 4 (respectively, column 5) focuses on residents of precincts where, in the 2008 national election, the center-left coalition received a vote share higher than (respectively, at or lower) than in the median precinct (57.2%).

There is no obvious pattern of heterogeneity by precinct-level results in the 2008 election. By contrast, effects on the gender turnout gap seem stronger in (slightly) lower-salience European/municipal elections than in national political elections. However, two important caveats are in order here. First, these regressions only use two data points for each voter. Second, the differential timing of the two types of elections could mask other margins of heterogeneity. For example, the 2008 and 2013 national elections span the years of the Great Recession, which could drive the relatively smaller effect of children on the gender turnout gap in several ways (e.g., making the allocation of family chores/parental duties more equal across partners).

For completeness, in Table A15 we report estimated gender-specific effects of marital status by political context. As in Table A14 (and thus subject to the same important caveats), national political elections seem associated with more tenuous differences in gender-specific effects, particularly for marriage.

Table A12: Effect of Children on Turnout by Children's Age and Voter's Gender: By Marital Status and Parents' Origins **(For Online Publication)**

	Full Sample		Ever Married		Never Married		Ever Married Parents Born North		Ever Married Parents Born South	
	(1)		(2)		(3)		(4)		(5)	
1(Female w/ children aged 0-5)	-.017	**	-.018	**	-.008		-.016	**	-.029	**
	(.003)		(.004)		(.009)		(.004)		(.006)	
1(Male w/ children aged 0-5)	.001		-.001		.021	*	-.004		.004	
	(.003)		(.003)		(.010)		(.004)		(.006)	
1(Female w/ children aged 6-11)	.004		.002		.019	~	-.002		-.001	
	(.003)		(.003)		(.011)		(.003)		(.006)	
1(Male w/ children aged 6-11)	.010	**	.009	**	.015		.005		.014	*
	(.003)		(.003)		(.015)		(.003)		(.005)	
1(Female w/ children aged 12-17)	.006	*	.004		.021		.000		.008	~
	(.003)		(.003)		(.013)		(.003)		(.005)	
1(Male w/ children aged 12-17)	.013	**	.013	**	.015		.009	**	.014	**
	(.003)		(.003)		(.020)		(.003)		(.005)	
1(Female w/ children aged 18+)	.013	**	.012	**	.035	*	.010	**	.015	**
	(.002)		(.002)		(.015)		(.002)		(.004)	
1(Male w/ children aged 18+)	.011	**	.012	**	.022		.009	**	.020	**
	(.002)		(.002)		(.030)		(.002)		(.004)	
$\beta^{0\text{-}5\ female} - \beta^{0\text{-}5\ male}$	-.018	**	-.017	**	-.029	*	-.011	**	-.032	**
	(.004)		(.004)		(.013)		(.004)		(.007)	
$\beta^{6\text{-}11\ female} - \beta^{6\text{-}11\ male}$	-.006	~	-.008	*	.005		-.007	~	-.015	*
	(.004)		(.004)		(.019)		(.004)		(.006)	
$\beta^{12\text{-}17\ female} - \beta^{12\text{-}17\ male}$	-.007	*	-.008	**	.006		-.009	**	-.006	
	(.003)		(.003)		(.024)		(.003)		(.006)	
$\beta^{18+\ female} - \beta^{18+\ male}$	.001		.0003		.014		.0003		-.006	
	(.003)		(.0027)		(.033)		(.0028)		(.005)	
Voter FEs	✓		✓		✓		✓		✓	
Age×Gender FEs	✓		✓		✓		✓		✓	
Election FEs	✓		✓		✓		✓		✓	
Neighborhood controls	✓		✓		✓		✓		✓	
Household controls	✓		✓		✓		✓		✓	
Income and taxes paid	✓		✓		✓		✓		✓	
Marital status×Gender FEs	✓		✓		✓		✓		✓	
Female w/o kids $\bar{Y}$	.809		.813		.797		.824		.793	
Male w/o kids $\bar{Y}$	.833		.863		.762		.879		.823	
N	902,153		735,398		166,751		616,028		206,174	

Notes: Neighborhood controls are: precinct-year average age, and OECD modified gross household income, as well as shares of female and Italian residence, and city neighborhood-by-year fixed effects. Household controls are the share of household members who are Italian citizens and the OECD modifed gross household income. Marital status FEs are three, mutually exclusive dummies indicating married, divorced, and widowed voters. Standard errors are two-way clustered by voter and household.

** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

Table A13: Effect of Children on Turnout by Children's Age and Voter's Gender: Distance to the Polls and Number of Children **(For Online Publication)**

	Full Sample		Distance 0-400m		Distance 400m+		Max 1 Kid	
	(1)		(2)		(3)		(4)	
1(Female w/ children aged 0-5)	-.017	**	-.015	**	-.017	**	-.019	**
	(.003)		(.005)		(.005)		(.004)	
1(Male w/ children aged 0-5)	.001		-.003		.003		.001	
	(.003)		(.005)		(.005)		(.004)	
1(Female w/ children aged 6-11)	.004		.006		-.000		.005	
	(.003)		(.004)		(.004)		(.004)	
1(Male w/ children aged 6-11)	.010	**	.008	~	.008	*	.012	**
	(.003)		(.004)		(.004)		(.004)	
1(Female w/ children aged 12-17)	.006	*	.006	~	.004		.007	~
	(.003)		(.004)		(.004)		(.004)	
1(Male w/ children aged 12-17)	.013	**	.010	*	.011	**	.015	**
	(.003)		(.004)		(.004)		(.004)	
1(Female w/ children aged 18+)	.013	**	.009	**	.016	**	.020	**
	(.002)		(.003)		(.003)		(.004)	
1(Male w/ children aged 18+)	.011	**	.005		.015	**	.025	**
	(.002)		(.003)		(.003)		(.004)	
$\beta^{0\text{-}5\ female} - \beta^{0\text{-}5\ male}$	-.018	**	-.012	*	-.019	**	-.020	**
	(.004)		(.006)		(.006)		(.005)	
$\beta^{6\text{-}11\ female} - \beta^{6\text{-}11\ male}$	-.006	~	-.002		-.008	~	-.006	
	(.004)		(.005)		(.005)		(.005)	
$\beta^{12\text{-}17\ female} - \beta^{12\text{-}17\ male}$	-.007	*	-.003		-.007		-.008	~
	(.003)		(.005)		(.005)		(.005)	
$\beta^{18+\ female} - \beta^{18+\ male}$	.001		.004		.000		-.005	
	(.003)		(.004)		(.004)		(.005)	
Voter FEs	✓		✓		✓		✓	
Age×Gender FEs	✓		✓		✓		✓	
Election FEs	✓		✓		✓		✓	
Neighborhood controls	✓		✓		✓		✓	
Household controls	✓		✓		✓		✓	
Income and taxes paid	✓		✓		✓		✓	
Marital status×Gender FEs	✓		✓		✓		✓	
Female w/o kids $\bar{Y}$	.809		.813		.807		.855	
Male w/o kids $\bar{Y}$	.833		.840		.831		.844	
N	902,153		455,662		425,311		176,866	

Notes: Neighborhood controls are: precinct-year average age, and OECD modified gross household income, as well as shares of female and Italian residence, and city neighborhood-by-year fixed effects. Household controls are the share of household members who are Italian citizens and the OECD modifed gross household income. Marital status FEs are three, mutually exclusive dummies indicating married, divorced, and widowed voters. Standard errors are two-way clustered by voter and household.
** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

Table A14: Effect of Children on Turnout by Children's Age and Voter's Gender: Political Context **(For Online Publication)**

	Full Sample		European/ Municipal Elections		National Political Elections		High Left Vote Share		Low Left Vote Share	
	(1)		(2)		(3)		(4)		(5)	
1(Female w/ children aged 0-5)	-.017	**	-.020	**	-.005		-.017	**	-.015	**
	(.003)		(.006)		(.006)		(.005)		(.005)	
1(Male w/ children aged 0-5)	.001		-.0003		.005		-.001		.004	
	(.003)		(.0054)		(.005)		(.005)		(.004)	
1(Female w/ children aged 6-11)	.004		.008		.003		.015	**	-.006	
	(.003)		(.005)		(.005)		(.004)		(.004)	
1(Male w/ children aged 6-11)	.010	**	.011	*	.009	~	.012	**	.006	
	(.003)		(.005)		(.005)		(.004)		(.004)	
1(Female w/ children aged 12-17)	.006	*	.019	**	-.004		.011	**	.001	
	(.003)		(.005)		(.004)		(.004)		(.004)	
1(Male w/ children aged 12-17)	.013	**	.015	**	.005		.013	**	.011	**
	(.003)		(.005)		(.004)		(.004)		(.004)	
1(Female w/ children aged 18+)	.013	**	.017	**	.011	**	.008	**	.018	**
	(.002)		(.003)		(.004)		(.003)		(.003)	
1(Male w/ children aged 18+)	.011	**	.012	**	.011	**	.011	**	.011	**
	(.002)		(.004)		(.004)		(.003)		(.003)	
$\beta^{0\text{-}5\ female} - \beta^{0\text{-}5\ male}$	-.018	**	-.020	**	-.010		-.016	**	-.018	**
	(.004)		(.007)		(.007)		(.006)		(.006)	
$\beta^{6\text{-}11\ female} - \beta^{6\text{-}11\ male}$	-.006	~	-.002		-.006		.003		-.012	*
	(.004)		(.006)		(.006)		(.005)		(.005)	
$\beta^{12\text{-}17\ female} - \beta^{12\text{-}17\ male}$	-.007	*	.004		-.009		-.002		-.010	*
	(.003)		(.005)		(.006)		(.005)		(.005)	
$\beta^{18+\ female} - \beta^{18+\ male}$	.001		.005		-.001		-.002		.006	~
	(.003)		(.004)		(.005)		(.004)		(.004)	
Voter FEs	✓		✓		✓		✓		✓	
Age×Gender FEs	✓		✓		✓		✓		✓	
Election FEs	✓		✓		✓		✓		✓	
Neighborhood controls	✓		✓		✓		✓		✓	
Household controls	✓		✓		✓		✓		✓	
Income and taxes paid	✓		✓		✓		✓		✓	
Marital status×Gender FEs	✓		✓		✓		✓		✓	
Female w/o kids $\bar{Y}$	.809		.821		.826		.818		.802	
Male w/o kids $\bar{Y}$	.833		.848		.857		.843		.827	
N	902,153		400,602		393,748		444,198		442,002	

Notes: Neighborhood controls are: precinct-year average age, and OECD modified gross household income, as well as shares of female and Italian residence, and city neighborhood-by-year fixed effects. Household controls are the share of household members who are Italian citizens and the OECD modifed gross household income. Marital status FEs are three, mutually exclusive dummies indicating married, divorced, and widowed voters. Standard errors are two-way clustered by voter and household.

** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$

Table A15: Effect of Marital Status by Voter's Gender: Political Context **(For Online Publication)**

	Full Sample		European/ Municipal Elections		National Political Elections		High Left Vote Share		Low Left Vote Share	
	(1)		(2)		(3)		(4)		(5)	
1(Married female)	.004		-.008		.011		.008		-.003	
	(.005)		(.008)		(.008)		(.007)		(.007)	
1(Married male)	.019	**	.017	*	.021	**	.024	**	.015	*
	(.004)		(.007)		(.007)		(.007)		(.006)	
1(Divorced female)	.013	~	-.006		.023	~	.016		.009	
	(.008)		(.013)		(.013)		(.011)		(.011)	
1(Divorced male)	.011		-.006		.017		.035	**	-.005	
	(.008)		(.014)		(.013)		(.013)		(.011)	
1(Widowed female)	-.021	**	-.032	**	-.023	*	-.014		-.031	**
	(.006)		(.010)		(.010)		(.009)		(.009)	
1(Widowed male)	.000		-.011		.001		.009		-.011	
	(.007)		(.012)		(.012)		(.011)		(.011)	
$\beta^{\text{married female}} - \beta^{\text{married male}}$	-.015	*	-.025	*	-.010		-.016	~	-.018	*
	(.006)		(.010)		(.010)		(.009)		(.009)	
$\beta^{\text{divorced female}} - \beta^{\text{divorced male}}$	.003		-.0001		.006		-.019		.014	
	(.011)		(.0190)		(.018)		(.017)		(.015)	
$\beta^{\text{widowed female}} - \beta^{\text{widowed male}}$	-.021	*	-.021		-.025		-.024	~	-.021	
	(.009)		(.015)		(.015)		(.013)		(.014)	
Voter FEs	✓		✓		✓		✓		✓	
Age×Gender FEs	✓		✓		✓		✓		✓	
Election FEs	✓		✓		✓		✓		✓	
Neighborhood controls	✓		✓		✓		✓		✓	
Household controls	✓		✓		✓		✓		✓	
Income and taxes paid	✓		✓		✓		✓		✓	
Children×Gender FEs	✓		✓		✓		✓		✓	
Never-married female $\bar{Y}$	.813		.821		.832		.818		.811	
Never-married male $\bar{Y}$	.787		.797		.807		.788		.789	
N	1,084,202		490,054		473,894		532,062		530,147	

Notes: Neighborhood controls are: precinct-year average age, and OECD modified gross household income, as well as shares of female and Italian residence, and city neighborhood-by-year fixed effects. Household controls are the share of household members who are Italian citizens and the OECD modifed gross household income. Children FEs are four dummies indicating presence of one or more children of the following ages: 0-5, 6-11, 12,-17, 18+. Standard errors are two-way clustered by voter and household.

** $p < 0.01$, * $p < 0.05$, ~ $p < 0.10$