The Daughter Effect in Politics: Evidence from Japanese National Legislators*

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Abstract

We examine how politicians' family structure shapes policy preferences, focusing on the influence of daughters on attitudes toward gender- and family-related policies. Using original data on the number and gender of children of candidates elected to Japan's House of Representatives in 2021, combined with candidate survey responses collected during the election campaign, we analyze whether exposure to daughters affects positions on optional separate surnames for married couples, same-sex marriage, and legislation promoting understanding of sexual minorities. Exploiting plausibly exogenous variation in the number of daughters conditional on the number of children, we find that daughters are associated with more progressive policy positions among male candidates. Our findings highlight the role of family-based socialization in shaping elite political preferences and extend evidence on the "daughter effect" to a non-Western context.

Keywords: D72, J16, D19

JEL Codes: Daughter effect; political socialization; gender attitudes

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1 Introduction

> "One of the first things we discover in these groups is that personal problems are political problems."
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> — Carol Hanisch (1970).

> "At the end of the day, I have to look my daughters in the eye and tell them, 'I believe I did what was right,' and I am doing that today."
>
> — Chris Madel.

Political preferences among political elites are shaped not only by ideology and constituency pressures, but also by personal life. Yet research on elites often treats personal life experience as statistical noise rather than a source of meaningful political variation. This omission can be consequential, as legislators make policy about marriage, work, and equality while simultaneously inhabiting these institutions as spouses, parents, and caregivers. A large literature shows that intimate experiences, such as parenthood, can shape how citizens think about politics (e.g., Oswald and Powdthavee (2010); Yang et al. (2025)). It is therefore plausible that similar experiences also influence how politicians evaluate policies that regulate family and gender relations.

This paper studies whether politicians' exposure to daughters is associated with their policy positions on gender- and family-related issues. The central idea is straightforward. Daughters can change what political actors—especially male actors—see, hear, and feel through new information, altered social networks, and the salience of gendered disadvantages and constraints. In many accounts, daughters are expected to make parents more supportive of gender equality. However, the mechanisms are not unidirectional. For some politicians, especially those already embedded in traditional gender norms or party organizations, having daughters may instead increase concern about social backlash, electoral vulnerability, or the perceived costs of rapid change. Whether daughters liberalize, polarize, or condition

preferences thus remains an empirical question, and one that is particularly informative when studied among elites.

Japan provides a demanding setting for testing family-based influences on elite preferences. Gender equality indicators remain comparatively low among OECD democracies, party organizations exert substantial discipline, and politicians operate in an electoral environment in which publicly articulated positions may be strategically constrained. If personal exposure to daughters is associated with elite policy preferences even in this context, it strengthens the case that family experiences can shape political attitudes under conditions that should mute idiosyncratic influences. Conversely, heterogeneous or null effects would clarify the limits of "daughter effects" in settings where social norms and organizational incentives may dominate.

We examine candidates elected to Japan's House of Representatives in 2021. We link information on legislators' number of children and daughters to original candidate survey responses collected during the election campaign. Our outcomes capture salient contemporary debates over family and sexual-minority policy: support for allowing married couples to retain separate premarital surnames, support for legal recognition of same-sex marriage, and support for early enactment of legislation promoting understanding of sexual minorities. Focusing on stated preferences during the campaign allows us to assess how private family exposure correlates with positions that candidates are willing to express publicly.

Our identification strategy exploits the fact that, conditional on total number of children, variation in the number of daughters is plausibly as-good-as random. We therefore compare candidates with the same parity but different sex composition of children, flexibly controlling for family size and observable covariates. This design addresses key concerns about selection into parenthood and helps separate exposure to daughters from having more children. We further probe the plausibility of the design using descriptive diagnostics based on the joint distribution of parity and daughters/ We also show robustness across alternative functional forms and specifications.

The results show that the association between daughters and policy preferences is sharply heterogeneous by candidate gender. For male legislators, additional daughters are consistently associated with more supportive positions on gender- and sexuality-related policies. For female legislators, exposure to daughters does not appear to be liberalizing. If anything, having daughters is associated with less supportive positions among female legislators. This pattern is difficult to reconcile with a single "daughters always liberalize" account and instead points to more nuanced mechanisms that depend on politicians' baseline constraints, strategic incentives, and the social meaning of gendered family roles.

Our main results are largely robust to alternative assumptions. First, although our main analysis uses linear probability models, the results are qualitatively unchanged when we estimate binary logit models. Second, we dichotomize ordered responses in the main specifications to aid interpretation, and the estimates remain qualitatively similar when we instead use ordered logit models. Third, the findings are largely unchanged when we restrict the sample to legislators aged 50 or older, reducing concerns that differences in fertility timing among younger legislators drive the results.

This paper contributes to three debates. First, it advances research on personal experience and political representation by testing whether family composition is linked to policy positions among elected elites, not only among mass publics. Second, it adds evidence from a least-likely case in which organizational discipline and conservative norms should attenuate idiosyncratic family influences. Third, by documenting gender-heterogeneous associations, it clarifies that exposure to daughters does not map mechanically onto more progressive attitudes; its political implications depend on who the politician is and how gendered experiences are processed within elite and partisan environments.

The remainder of the paper proceeds as follows. Section 2 reviews the related literature. Section 3 describes the institutional background and data. Section 4 explains the empirical strategy and identification assumptions. Section 5 presents the main results, and Section 6 reports robustness checks. Section 7 concludes.

2 Literature Review

Empirical work on how legislators' exposure to daughters relates to their political attitudes remains scarce, plausibly because data on politicians' children are typically difficult to compile. Existing evidence is shaped largely by a small set of pioneering studies of the U.S. Congress.[1] The canonical contribution is Washington (2008), who shows that, conditional on total number of children, U.S. legislators with more daughters are more likely to cast liberal roll call votes, with the clearest and most consistent patterns emerging on reproductive rights votes in the 105th to 108th Congresses. Building on this framework, Costa et al. (2019) extend the analysis to the 110th to 114th Congresses and find no statistically distinguishable effect of fathering daughters on roll call voting or cosponsorship support for women's issues once party is accounted for. They interpret these null results as consistent with a view that the Congress is becoming more polarized and party-disciplined, in which members have less latitude for private experiences to translate into legislative behavior. Later, Green et al. (2023) revisit these studies and extend the analysis to earlier Congresses, when partisan polarization was lower. They again find null effects, concluding that the daughter effect is confined to the specific cohorts analyzed by Washington (2008).

The mixed evidence from the U.S. Congress makes external validation essential. Green et al. (2023) argue that the pattern in Washington (2008) is cohort-specific and that striking natural-experiment findings can reflect statistical variation that does not replicate out of sample. Their critique shifts the burden of proof toward testing whether daughter effects generalize beyond the specific congressional cohorts and institutional setting that first produced a positive result. This motivates examining other polities where party discipline, electoral incentives, and gender norms differ. We answer that call by evaluating the daughter-effect hypothesis for Japanese national legislators, who face different partisan constraints, electoral rules, and gendered social expectations than their U.S. counterparts.

[1] One exception is Van Effenterre (2020), who provides a cross-national test by examining whether the number of daughters predicts male legislators' roll-call voting on abortion-related legislation in both the U.S. Congress and the French legislature.

While Green et al. (2023) conclude that daughters do not appear to have any appreciable effect on U.S. members' legislative behavior, it may be premature to generalize from this institutional setting to daughter effects more broadly, particularly given robust cross-national evidence from the mass public. Across countries, having daughters is associated with meaningful shifts in attitudes on gender and family-related issues (e.g., Oswald and Powdthavee (2010); Yu and Kuo (2018); Lee and Conley (2016); Perales et al. (2018); Chiba and Ono (2025)), although the direction and magnitude of these effects vary across contexts.[2] If exposure to daughters can reshape political preferences among ordinary citizens, it is reasonable to hypothesize that similar mechanisms may influence political elites, potentially in contexts where institutional and social conditions differ sharply from the U.S.[3]

Indeed, a growing body of research documents daughter effects beyond the political arena, linking children's gender composition to managerial decision-making, including CEOs' CSR choices, partners' hiring decisions in venture capital, and managers' workplace practices toward female employees (Cronqvist and Yu, 2017; Calder-Wang and Gompers, 2021; Wu et al., 2024; Ronchi and Smith, 2025). Together, these findings suggest that exposure to daughters can shape the preferences and actions of organizational decision-makers in consequential ways, reinforcing the plausibility that similar mechanisms may operate among elected officials.

The foregoing literature motivates a test of the daughter-effect hypothesis in Japan, where electoral institutions, party constraints, and gender norms differ from those in the U.S. Given the negative evidence from the U.S. Congress, examining a distinct polity is

[2]For an extensive survey, see Table SI 1.1 in the online appendix provided by Green et al. (2023). A recent meta-analysis of the European Social Survey across 39 countries reports that having daughters is associated with small but statistically significant increases in support for gender equality and homosexual rights, with little cross-national heterogeneity, and no detectable effects on left-right ideology or redistribution preferences (Yang et al., 2025).

[3]While Green et al. (2023) suggest that any daughter effect in the U.S. Congress is limited to the specific cohorts studied by Washington (2008), studies of the U.S. general public often find non-negligible daughter effects on gender-related beliefs and political attitudes (e.g., Conley and Rauscher (2013); Sharrow et al. (2018); Borrell-Porta et al. (2019); Van Effenterre (2020)). In addition, daughter effects have been documented in other elite political settings, including U.S. appellate courts and the Constitutional Convention (Glynn and Sen, 2015; Pope and Schmidt, 2021).

useful for distinguishing between two possibilities: that any daughter effects are limited to the particular cohorts studied by Washington (2008), or that the U.S. Congress is an institutional setting in which such effects are less likely to surface. Japan thus offers an informative out-of-sample setting for assessing the broader relevance of the daughter-effect hypothesis.

3 Institutional Context and Data

This study combines original data on politicians' family composition with candidate-level survey data from the 2021 Japanese House of Representatives election.

3.1 Why Japan Matters

Japan provides a particularly demanding context in which to examine whether politicians' family structure shapes policy preferences. First, among OECD countries, Japan ranks near the bottom on most measures of gender equality, including women's labor force participation, political representation, and legal protections related to family and work. These features make Japan a least-likely case for detecting a daughter effect: if exposure to daughters influences legislators' attitudes even in a highly gender-unequal setting, the underlying mechanism is likely to be substantively important.

Second, legislators' behavior in Japan is strongly shaped by party organizations. Party discipline is high, candidate selection is centralized, and policy positions are often coordinated within parties, particularly on socially salient issues. These institutional features limit the scope for individual legislators' private experiences to translate into publicly expressed policy positions, further biasing against finding effects of family background.

Third, our analysis relies on candidates' stated policy preferences elicited during the election campaign, rather than revealed behavior such as roll-call voting. Campaign periods are characterized by heightened strategic considerations, electoral incentives, and pressure to

adhere to party platforms. As a result, candidates' expressed positions during campaigns are likely to be more constrained by electoral calculas than by private beliefs. This measurement choice therefore provides a conservative test of whether exposure to daughters influences legislators' policy preferences.

Overall, Japan's gender norms, party-centered political institutions, and the strategic context in which preferences are measured all work against detecting a daughter effect. Finding systematic associations in this setting would therefore suggest that the influence of family structure on elite policy preferences is robust and not limited to more permissive political environments.

3.2 Election Survey Data

Data on politicians' policy positions come from the joint candidate survey conducted by the *Asahi Shimbun* and the University of Tokyo's Taniguchi Laboratory in conjunction with the 49th House of Representatives election.[4] The survey asks candidates to publicly state their positions on a wide range of political and social issues during the election campaign. Responses are recorded on a five-point scale ranging from strong opposition to strong support and capture candidates' campaign-period policy positions rather than post-election behavior. This survey has been widely used in the literature on Japanese electoral politics and is regarded as a reliable measure of candidates' stated preferences.

Survey responses are published during the election on the campaign special website of the *Asahi Shimbun*, one of Japan's major newspapers. The results are publicly available to anyone (not only subscribers) and presented in a graphically intuitive format. This high visibility makes responses consequential because constituents can use this information when deciding how to vote, and the seat is at stake. This feature of the survey strengthens both the credibility and the political relevance of the responses.

We focus on three question items from the survey. The outcomes of interest are can-

[4]The dataset is available at `https://www.masaki.j.u-tokyo.ac.jp/utas/utasindex.html#2021`.

didates' attitudes toward gender- and family-related policies, most notably support for the introduction of an optional dual-surname system for married couples.

3.2.1 Survey Question Items

Politicians' policy preferences are measured using responses from the Taniguchi Laboratory-*Asahi Shimbun* joint candidate survey conducted in advance of the 2021 House of Representatives election. The survey asks candidates to report their positions on a broad set of political issues spanning economic policy, national security, social welfare, constitutional reform, and social values.

Our primary dependent variables are drawn from Question 6 of the survey, which asks candidates whether they agree or disagree with a series of concrete policy proposals. The policy questions are measured on five-point Likert scales indicating the degree of agreement or opposition to specific policy statements. Response categories range from strong agreement to strong opposition, with an explicit neutral midpoint. Respondents may also skip individual items; such non-responses are coded separately and excluded from the analysis. Specifically, respondents may choose among the following six options: "Strongly agree"; "Agree"; "Neither agree nor disagree"; "Disagree"; "Strongly disagree"; "No response".

We focus on items related to gender and family policy. The main outcome variable captures candidates' attitudes toward the introduction of an optional dual-surname system for married couples, based on responses to the item: "If a married couple so desires, the law should allow spouses to retain their premarital surnames" (Q6_15).

To examine whether the effect of daughters is specific to women's issues or reflects a broader shift in social attitudes, we also analyze responses to related but distinct items concerning sexual minorities. These include attitudes toward the legalization of same-sex marriage (Q6_16) and support for the early passage of legislation to promote understanding of LGBT and other sexual minorities (Q6_17).

In our main empirical exercise, we recode each item into a binary indicator capturing

whether a candidate expresses support for the policy proposal. Candidates selecting either of the two support categories are coded as 1, and all other responses (including neutral and opposition categories) are coded as 0. This coding facilitates interpretation by distinguishing supporters from non-supporters.

3.3 Candidates' Children Data

The key explanatory variables measure the gender composition of politicians' children. Because no existing administrative, survey, or archival dataset systematically records family composition for Japanese legislators, we constructed an original dataset covering candidates elected to the House of Representatives in 2021. To our knowledge, this dataset is the first to provide comprehensive, candidate-level information on the number and gender of children for Japanese national legislators.

Information on the number and gender of children was hand-collected in 2021 from politicians' official websites, publicly available biographies, and verified media sources. When these sources were insufficient or ambiguous, we contacted legislators' offices directly via email and telephone to obtain the relevant information. We recorded only information that could be clearly verified and excluded cases in which family composition could not be determined with confidence.

Using this procedure, we obtained data on children for 298 of the 465 candidates elected in 2021. Legislators included in the analytical sample are broadly representative of the full population of elected members in terms of observable characteristics such as age, party affiliation, and gender, mitigating concerns about systematic selection into data availability. This original dataset enables a novel examination of how the gender composition of politicians' children relates to their policy preferences in the 2021 election.

3.4 Sample Characteristics

To assess potential selection into our estimation sample, Table 1 compares observable characteristics of all candidates elected to the House of Representatives in 2021 (panel (a)) with the subset of those candidates whose information on the number and gender of children is available (panel (b)). Panel (c) reports the corresponding differences in means ((b) less (a)).

Overall, the two samples are highly similar across key demographic and political characteristics. The proportion of male legislators and the age distribution are nearly identical across panels, with an average age difference of less than half a year. Party affiliation is also well balanced between the full sample and the analytical subsample. With the exception of a modestly lower share of Japan Innovation Party legislators in the subsample, differences in party composition are small and generally statistically indistinguishable from zero.

Measures of political experience and electoral performance show similarly limited differences. Legislators in the estimation sample have marginally higher single-member district vote shares, but these differences are economically small. Incumbency status, proportional-representation-only candidacy, and electoral outcomes are closely aligned across the two groups.

Crucially, the subsample does not differ meaningfully from the full population in policy attitudes toward gender- and sexuality-related issues. Support for optional separate surnames, same-sex marriage, and policies to promote understanding of sexual minorities is nearly identical across panels, suggesting that sample selection based on the availability of children information is unlikely to bias estimates of policy preferences.

Overall, these patterns indicate that restricting the analysis to legislators with available children information does not materially alter the demographic, political, or attitudinal composition of the sample. This similarity supports the internal validity of the empirical analysis that follows.

4 Empirical Strategy

To estimate the effect of daughters on political attitude, we begin with the following regression equation:

$$\begin{aligned} y_i =& \beta_0 + \beta_1 N_i^{\mathrm{D}} + \beta_2 \left(\mathsf{Male}_i \times N_i^{\mathrm{D}}\right) + \beta_3 \mathsf{Male}_i \\ &+ \sum_{k=1}^{7} \delta_k \mathbb{1}\{N_i^{\mathrm{C}} = k\} + \sum_{k=1}^{7} \theta_k \left(\mathsf{Male}_i \times \mathbb{1}\{N_i^{\mathrm{C}} = k\}\right) \\ &+ X_i'\gamma + \varepsilon_i, \end{aligned} \tag{1}$$

where y_i is each question item of interest; N_i^{D} is the number of daughters representative i reported to have; N_i^{C} is the number of children representative i reported to have; Male_i takes 1 if representative i is male and 0 otherwise; and X_i is a vector of candidate characteristics, including party fixed effects, an indicator for proportional-representation-only candidates, proportional-representation block fixed effects, the number of terms served in the House of Representatives fixed effects, age and its square, and candidates' residing prefecture fixed effects.

The parameters of interest in equation (1) are β_1 and β_2. Because the number of daughters is interacted with a male dummy, β_1 captures the marginal effect of an additional daughter for female elected candidates. The coefficient β_2 represents the incremental effect of daughters for male candidates relative to female candidates. Accordingly, the marginal effect of an additional daughter for male candidates is given by $\beta_1 + \beta_2$. We cluster standard errors at the political party and candidates' residing prefecture levels.

Our identification strategy follows the approach in Washington (2008). The identifying assumption underlying equation (1) is that, conditional on the total number of children, the gender composition of children is as good as randomly assigned. Under this assumption, variation in the number of daughters is orthogonal to unobserved determinants of candidates' political attitudes. By controlling flexibly for family size, we address potential concerns

related to fertility stopping behavior and ensure that identification comes from differences in the gender composition of children among candidates with the same number of children. The intuition is that, conditional on the total number of children, exposure to daughters—and thus to female family members—is plausibly exogenous.

In addition to the baseline specification (1), we also estimate two variants to explore potentially non-linear effects of daughters. First, we use a dummy indicating whether the candidate has at least one daughter. The estimating equation is

$$\begin{aligned} y_i =& \beta_0 + \beta_1 \mathbb{1}\{N_i^{\mathrm{D}} \geq 1\} + \beta_2 \left(\mathsf{Male}_i \times \mathbb{1}\{N_i^{\mathrm{D}} \geq 1\}\right) + \beta_3 \mathsf{Male}_i \\ &+ \sum_{k=1}^{7} \delta_k \mathbb{1}\{N_i^{\mathrm{C}} = k\} + \sum_{k=1}^{7} \theta_k \left(\mathsf{Male}_i \times \mathbb{1}\{N_i^{\mathrm{C}} = k\}\right) \\ &+ X_i'\gamma + \varepsilon_i, \end{aligned} \tag{2}$$

where $\mathbb{1}\{N_i^{\mathrm{D}} \geq 1\}$ is an indicator function that takes 1 if candidate i has more than or equal to one children and 0 otherwise. This specification allows the influence of daughters to operate through the presence of any daughter rather than through the number of daughters, abstracting from potential nonlinearities in family size beyond the first daughter. In this specification, β_1 captures the difference in political attitudes between female candidates with at least one daughter and female candidates with no daughters, holding constant the total number of children and other covariates. Thus, β_1 reflects the extensive-margin effect of exposure to daughters for female candidates. The interaction term coefficient β_2 measures how this effect differs for male candidates relative to female candidates. Accordingly, the effect of having at least one daughter for male candidates is given by $\beta_1 + \beta_2$.

Second, we use the ratio of the number of daughters to the number of sons.

$$\begin{aligned} y_i =& \beta_0 + \beta_1 R_i^{\mathrm{D}} + \beta_2 \left(\mathsf{Male}_i \times R_i^{\mathrm{D}}\right) + \beta_3 \mathsf{Male}_i \\ &+ \sum_{k=1}^{7} \delta_k \mathbb{1}\{N_i^{\mathrm{C}} = k\} + \sum_{k=1}^{7} \theta_k \left(\mathsf{Male}_i \times \mathbb{1}\{N_i^{\mathrm{C}} = k\}\right) \\ &+ X_i'\gamma + \varepsilon_i, \end{aligned} \tag{3}$$

where

$$R_i^{\mathrm{D}} = \begin{cases} \frac{N_i^{\mathrm{D}}}{N_i^{\mathrm{C}}} & \text{if} \quad N_i^{\mathrm{C}} \geq 1, \\ 0 & \text{if} \quad N_i^{\mathrm{C}} = 0. \end{cases}$$

This specification captures relative exposure to daughters within the family and allows the effect of daughters to scale with the proportion of female children rather than their absolute number. In this specification, β_1 captures the effect for female candidates of a unit increase in the proportion of daughters, holding the total number of children fixed. Because the number of children is controlled for flexibly, this variation reflects changes in the gender composition of children rather than changes in fertility. The interaction coefficient β_2 again measures the differential effect for male candidates relative to their female counterparts. The marginal effect of a higher share of daughters for male candidates is therefore $\beta_1 + \beta_2$.

The identifying assumption for equations (2) and (3) is the same as for the baseline specification, i.e., conditional on the total number of children, the gender composition of children is as good as randomly assigned.

4.1 Gender-Based Fertility Stopping

Washington (2008) points out that fertility stopping rules pose a potential threat to identification in designs that exploit variation in the gender composition of children. The core identifying assumption in such designs is that, conditional on parity (i.e., the total number of children), the gender composition of children is as good as randomly assigned. Stopping

behavior can violate this assumption if parents' fertility decisions respond to the gender of existing children.

Washington (2008) gives the following example. Suppose there are two types of parents in the economy. Type I parents hold strong son-preferring norms and more conservative gender attitudes. They ideally desire three children but continue childbearing until they have at least three children and at least two sons. Type II parents have no gender preference and also ideally desire three children, stopping once they reach that parity. In this setting, child gender composition is correlated with parental gender attitudes even after conditioning on parity. Among families with exactly three children, those with only one son are disproportionately drawn from parents without son preference, whereas families with two or three sons are a mixture of parents with and without such preferences.

If such preferences over child gender are also correlated with political attitudes, then differences in gender composition among families with the same parity would no longer be exogenous. Under such circumstances, estimates of the effect of daughters would conflate the causal effect of children's gender with selection driven by parental preferences rather than random variation in child gender. Washington (2008) argues that this concern can be diagnostically assessed by exploiting the quasi-random assignment of first-born gender. Under gender-based fertility stopping, the gender of the first child should predict completed fertility. Her finding that first-born gender does not predict total number of children provides evidence against systematic stopping behavior and supports the validity of the identification strategy

4.1.1 The Distribution of Fertility

Since birth-order information is not available, our data do not permit implementation of the stopping-rule test performed by Washington (2008), which can detect gender-based fertility stopping by regressing the total number of children on the gender of the first-born child. Nevertheless, the joint distribution of parities and the number of daughters in Table 2

provides informative descriptive evidence regarding the plausibility of strong gender-based fertility stopping behavior in our sample.

For each parity greater than one, we observe substantial dispersion in the number of daughters. In particular, even among families with three or more children, there are non-trivial shares of candidates whose children are predominantly or exclusively daughters. For example, among candidates with three children, families span the full range from zero to three daughters, and among those with four children, families with four daughters are observed. Such patterns are difficult to reconcile with extreme stopping rules in which parents cease childbearing upon the birth of a son.

4.1.2 Simulation-Based Tests for Gender-Based Fertility Stopping

To formalize the descriptive patterns above, we use bootstrap simulations to test the null hypothesis that specific gender-biased stopping rules are at play. In Table 3, we simulate distributions under eight plausible null hypotheses (R1–R8) that capture various forms of fertility preferences documented in the literature.

The first category includes deterministic son-preference rules. Rule R1 (Stop-at-First-Son) assumes parents cease childbearing immediately upon the birth of a son. Rule R2 (Min-2-then-Stop-at-Son) imposes a minimum family size of two, after which parents stop if they have a son. Rule R6 (Stop-at-Two-Sons) assumes parents continue until they have at least two sons. Rule R7 (Resignation Rule) models parents who persist in trying for a son but "resign" and stop if they reach a maximum capacity (set at 3 or 4 children) without having a son.

The second category allows for probabilistic stopping behavior. Rule R3 (Logistic Hazard) defines the probability of stopping as a function of having a son, specified by the logistic hazard model:

$$\Pr(S_i = 1) = \text{logit}^{-1}(\alpha_N + \lambda M_i), \tag{4}$$

where S_i denotes the event of stopping childbearing at parity N, α_N is a parity-specific

baseline derived from the observed data, and M_i is an indicator variable that takes the value of 1 if the family has at least one son. λ captures the strength of son preference. We test scenarios with $\lambda = 1.0$ and $\lambda = 2.0$, representing moderate and strong preferences, respectively.

The third category considers alternative preference structures. Rule R4 (Stop-at-First-Daughter) tests for daughter preference. Rule R5 (Balance Preference) assumes parents stop once they achieve a gender mix (at least one son and one daughter). Rule R8 (Homogeneity Preference) assumes parents stop if their children are all of the same sex (conditional on having at least two children).

We define statistical power as the probability of rejecting these alternative hypotheses at the 5% significance level, estimated via 200 Monte Carlo simulations for each rule. As reported in Panel B, the statistical power varies by rule strength; for the moderate son-preference rule (R3 with $\lambda = 1.0$), the power is 0.47, reflecting the difficulty of distinguishing this subtle bias from random variation. However, the power increases to 0.99 for the stronger preference ($\lambda = 2.0$) and is close to 1.00 for most deterministic rules. Crucially, regardless of the power, our goodness-of-fit tests consistently reject all gender-biased specifications ($p < 0.001$), including the moderate $\lambda = 1.0$ case. These results provide robust evidence that systematic son-biased stopping is absent in our sample.

While the lack of birth-order data prevents us from strictly ruling out highly nuanced forms of gender preference, the combination of descriptive evidence, formal goodness-of-fit tests, and simulation results strongly suggests that fertility decisions in our sample are not driven by child gender. We therefore proceed with the assumption that the number of daughters is plausibly exogenous conditional on the total number of children.

5 Results

5.1 Support for Separate Surnames

Table 4 examines whether candidates' support for allowing married couples to retain separate premarital surnames varies with their exposure to daughters. Across specifications, the association between daughters and support for separate surnames differs sharply by candidate gender.

Column (1) uses the number of daughters and its interaction with a male dummy. For female candidates, each additional daughter is associated with a 10.2 percentage-point lower probability of supporting separate surnames, and this effect is statistically distinguishable from zero. The interaction term is positive and statistically significant, implying that the marginal effect of daughters for male candidates, defined as the sum of the main and interaction effects, is estimated to be $\widehat{\beta}_1 + \widehat{\beta}_2 = 0.057$. A linear hypothesis test rejects the null that $\beta_1 + \beta_2 = 0$ at conventional levels, indicating that the marginal effect for male candidates is statistically distinguishable from zero. This suggests that an additional daughter increases the probability that male candidates take a position in favor of separate surnames.

Columns (2) and (3) indicate that the gender-heterogeneous relationship in column (1) is not an artifact of how daughter exposure is measured. In column (2), daughter exposure is coded as an indicator for having at least one daughter. The estimates remain negative for female candidates and the interaction with the male dummy is positive. However, the implied marginal effect for male candidates is close to zero and not statistically distinguishable from zero, with $\widehat{\beta}_1 + \widehat{\beta}_2 = 0.003$. Column (3) measures exposure as the share of daughters among children. The main effect is again negative for women and the interaction is positive. For this exposure coding, the marginal effect for male candidates is $\widehat{\beta}_1 + \widehat{\beta}_2 = 0.081$, and this estimate is statistically distinguishable from zero.

In summary, the three specifications point to a consistent gender asymmetry in how child gender composition relates to this policy stance. For female candidates, greater exposure to

daughters is associated with lower stated support for separate surnames across all treatment definitions, indicating that the estimated relationship for women is negative regardless of how daughter exposure is measured. For male candidates, the marginal effect of daughters is positive and statistically distinguishable from zero when exposure varies on the intensive margin, measured by the number of daughters (column (1)) or the share of daughters (column (3)), but small and statistically insignificant when exposure is defined only on the extensive margin as having at least one daughter (column (2)).

5.2 Support for Same-Sex Marriage

Now we turn to whether this daughter effect extends to a broader set of gender-equality-related issues that are not limited to women's rights, beginning with candidates' views on the legal recognition of same-sex marriage. Table 5 reports estimates for candidates' support for legally recognizing same-sex marriage. As in the previous table, the relationship between daughter exposure and policy support differs markedly by candidate gender.

Column (1) uses the number of daughters and its interaction with the male dummy. For female candidates, each additional daughter is associated with a 19.9 percentage-point lower probability of supporting same-sex marriage. The interaction term is positive and precisely estimated, implying a positive marginal effect of daughters for male candidates. T he estimated marginal effect for men is $\widehat{\beta}_1 + \widehat{\beta}_2 = 0.087$, and the null hypothesis that $\beta_1 + \beta_2 = 0$ is rejected at conventional levels.

Columns (2) and (3) show that this pattern is robust to alternative codings of daughter exposure. In column (2), exposure is defined on the extensive margin as having at least one daughter. The estimate for female candidates is strongly negative, while the interaction with the male dummy is strongly positive. The implied marginal effect for male candidates is $\widehat{\beta}_1 + \widehat{\beta}_2 = 0.079$, which is statistically distinguishable from zero. Column (3) measures exposure on the intensive margin as the share of daughters among children. The main effect remains negative for women and the interaction is positive, yielding a positive marginal effect

for male candidates of $\widehat{\beta}_1 + \widehat{\beta}_2 = 0.127$, which is statistically distinguishable from zero at conventional levels.

Overall, across all three specifications, daughter exposure is associated with lower support for same-sex marriage among female candidates, but higher support among male candidates, with the estimated marginal effects for men consistently positive and statistically significant. This provides evidence that the daughter effect is not confined to women's rights issues, but also spills over to a broader domain of gender-equality-related attitudes, including support for sexual-minority rights.

5.3 Support for Sexual Minorities

Table 6 examines candidates' support for promptly enacting a "promotion of understanding" bill regarding sexual minorities. The coefficient patterns again differ by candidate gender, but in contrast to same-sex marriage, the implied marginal effects for male candidates are small and not statistically distinguishable from zero.

Column (1) measures exposure by the number of daughters. For female candidates, each additional daughter is associated with a 13.5 percentage-point lower probability of supporting the bill. The interaction with the male dummy is positive, implying an estimated marginal effect for male candidates of $\widehat{\beta}_1 + \widehat{\beta}_2 = 0.004$, which is not statistically distinguishable from zero.

Columns (2) and (3) show a similar picture using alternative codings of daughter exposure. When exposure is defined on the extensive margin as having at least one daughter (column (2)), the main effect for women is negative, while the interaction is positive but imprecisely estimated (0.180, s.e. 0.116). The implied marginal effect for men remains small and statistically insignificant, with $\widehat{\beta}_1 + \widehat{\beta}_2 = -0.032$. Column (3) uses the share of daughters among children. The main effect is again negative for women (-0.293, s.e. 0.025) and the interaction is positive, yielding an implied marginal effect for male candidates of $\widehat{\beta}_1 + \widehat{\beta}_2 = 0.020$, which is also not statistically distinguishable from zero.

Overall, daughter exposure is consistently associated with lower support for this bill among female candidates, while for male candidates the estimated marginal effects are close to zero across all three exposure measures.

5.4 Discussions

Across the three outcomes, the estimates reveal two recurring features. First, the relationship between daughter exposure and candidates' positions is consistently heterogeneous by candidate gender. Second, the direction and strength of the implied "daughter effect" depends on the policy domain. For male candidates, daughter exposure is positively associated with support for optional separate surnames and for legal recognition of same-sex marriage, with the marginal effects generally emerging on the intensive margin. For the "promotion of understanding" bill, however, the implied marginal effects for men are close to zero across treatment definitions. For female candidates, the estimated association is negative across all three issues and across alternative measures of daughter exposure.

The contrast between same-sex marriage and the "promotion of understanding" bill is informative. Same-sex marriage is a salient rights claim with a clear policy status quo and a relatively direct implication for legal equality. By comparison, the "promotion of understanding" bill is framed as an informational or educational measure and may be perceived as less policy-consequential, more symbolic, or more easily reconciled with diverse ideological positions. The absence of a positive marginal effect for men on this item suggests that daughter exposure does not translate into uniformly greater support for all sexual-minority–related policies, but is instead concentrated on issues where candidates may view the policy choice as a clear stance on legal recognition and equal treatment.

The intensive-margin pattern for male candidates in the surnames and same-sex marriage analyses also helps characterize the form of the association. Effects are stronger when daughter exposure varies in degree (number or share of daughters) rather than when it is defined only as crossing the threshold of having at least one daughter. This is consistent

with an interpretation in which repeated or greater exposure to daughters is more strongly linked to positioning on these issues than merely having a daughter.

Finally, the negative associations among female candidates are noteworthy in light of the central claim that exposure to daughters might liberalize attitudes. In this setting, women's estimated relationships run in the opposite direction across all outcomes. One implication is that the "daughter effect," as captured by within-parity variation in child sex composition, is not a general mechanism operating similarly across candidate genders. Instead, the patterns are consistent with a mechanism that, if present, is concentrated among men, and even then appears issue-specific rather than universal across the broader set of gender-equality-related policies considered here.

6 Robustness Checks

This section evaluates the sensitivity of the main results to alternative econometric specifications, potential nonlinearities implied by the binary and ordinal outcomes, and sample and subgroup definitions. Across these exercises, the qualitative patterns documented in Section 5 remain intact: the estimated relationship between daughter exposure and policy positions is concentrated among male candidates for optional separate surnames and same-sex marriage, while estimates for the sexual-minorities bill are weaker and less stable.

6.1 Econometric Modeling

6.1.1 Binary Logit Model

Our baseline estimates rely on a linear probability model (LPM), which offers transparent interpretation and accommodates high-dimensional fixed effects. As a robustness check, we re-estimate the main specifications using a binary logit model. Because logit coefficients are not directly comparable to LPM coefficients, we report average marginal effects (AMEs) evaluated at the sample distribution of covariates. The AME estimates yield the same

substantive conclusions as the LPM. Daughter exposure is associated with higher predicted support among male candidates for optional separate surnames and same-sex marriage, while the implied marginal effects for female candidates are negative. For the sexual-minorities bill, the marginal effects for male candidates remain small and statistically indistinguishable from zero.

Panel (a) of Tables A.1 and A.2 present the estimated AMEs for optional separate surnames and same-sex marriage, respectively. The results generally corroborate our baseline findings. In Column (1), which uses the number of daughters, the patterns are consistent with the linear model. For male candidates, the estimated marginal effect of an additional daughter on support for separate surnames is 4.5 percentage points. While this estimate loses statistical significance compared to the baseline LPM estimate of 5.7 percentage points, the magnitude of the coefficient remains substantively similar. In contrast, for same-sex marriage, the estimated marginal effect is 6.8 percentage points and becomes statistically significant at the 1% level, showing stronger statistical evidence than the baseline result.

Columns (2) and (3) report results using alternative measures of daughter exposure: a binary indicator for having at least one daughter and the share of daughters among children. For optional separate surnames, the daughter dummy in Column (2) remains statistically insignificant, consistent with the linear model. The ratio of daughters in Column (3) yields a positive marginal effect of 5.9 percentage points; although this estimate is not statistically significant in the logit specification, it is comparable in magnitude to the significant linear estimate of 8.1 percentage points. Similarly, for same-sex marriage, both the dummy and ratio measures yield positive marginal effects of 6.1 and 7.6 percentage points, respectively. These estimates align with the direction of the baseline findings, even though they do not reach conventional levels of statistical significance. Overall, while the precision of the estimates varies, the qualitative evidence points to a positive association between daughters and progressive positions among male candidates.

Consistent with the null result in the baseline specification, the marginal effects for the

sexual minorities bill (Table A.3, Panel A) remain small and statistically indistinguishable from zero for male candidates across all specifications.

6.1.2 Ordered Logit Model

The survey items are measured on an ordered response scale, but the main analysis dichotomizes responses into "agree/strongly agree" versus other categories. To confirm that the findings are not an artifact of this dichotomization, we estimate ordered logit models using the full ordinal scale. We calculate the average marginal effects on the probability of selecting the most supportive response, "strongly agree."

Panel B in Tables A.1 through A.3 displays these results. The ordered logit estimates provide robust evidence for the daughter effect, particularly in capturing the intensity of support.

In column (1), which uses the number of daughters, the results reinforce and clarify the baseline findings. For optional separate surnames, the number of daughters is associated with a significantly higher probability of "strongly agreeing." The estimated marginal effect is 3.8 percentage points and is statistically significant at the 1% level. This result is particularly noteworthy because it identifies a significant effect that was less precise in the binary logit specification, suggesting that daughters drive fathers toward a stronger, more definite stance on this issue. Similarly, for same-sex marriage, the estimated marginal effect is 5.0 percentage points and is statistically significant at the 1% level, indicating that each additional daughter substantially increases the likelihood of strong support.

Columns (2) and (3) examine alternative measures. For optional separate surnames, the results for the daughter dummy and ratio measures in Columns (2) and (3) do not reach statistical significance, a pattern consistent with the binary logit results. However, for same-sex marriage, the association remains robust across specifications. The ratio of daughters in Column (3) yields a significant positive marginal effect of 6.4 percentage points at the 1% level, and even the binary indicator for having a daughter in Column (2) shows a marginally

significant positive effect of 3.4 percentage points at the 10% level. These patterns suggest that for same-sex marriage, the daughter effect is consistent across both the intensive and extensive margins of exposure.

Finally, regarding the sexual minorities bill, the marginal effects for male candidates (Table A.3, Panel B) remain statistically indistinguishable from zero across all specifications, consistent with the null findings in the other models.

6.2 Menopausal Sample

One potential threat to interpretation is that the observed number and sex composition of children may reflect incomplete fertility among younger candidates. To address this concern, we re-estimate the models on a restricted sample for whom completed fertility is more plausible. Specifically, we focus on candidates at ages beyond the typical reproductive ages where additional childbearing is unlikely. Specifically, we restrict the sample to elected female and male candidates of age of 50 or older.

Tables A.4, A.5, and A.5 show the results. Reassuringly, the main conclusions are largely unchanged in this "menopausal" subsample, though the estimates are noisier due to the small sample size. The positive marginal effects of daughter exposure for male candidates on optional separate surnames and same-sex marriage persist, while estimates for the sexual-minorities bill remain close to zero. This suggests that the baseline results are not driven by differences in fertility timing among younger candidates, at least for separate surnames and same-sex marriage.

7 Conclusion

This paper investigates whether politicians' family experiences shape their publicly stated policy positions by studying the association between exposure to daughters and elite attitudes toward gender- and family-related policies in Japan. Using original hand-collected data on

the number and gender of children for candidates elected to Japan's House of Representatives in 2021, linked to campaign-period responses in a widely used candidate survey, we analyze three salient policy debates: optional separate surnames for married couples, legal recognition of same-sex marriage, and early passage of legislation to promote understanding of sexual minorities. Our empirical strategy exploits plausibly exogenous variation in the number (and share) of daughters conditional on total number of children and a rich set of covariates and fixed effects.

Three findings emerge. First, the relationship between daughters and policy positions is sharply heterogeneous by candidate gender. For male candidates, greater exposure to daughters is associated with higher stated support for optional separate surnames and for same-sex marriage, with effects most evident when daughter exposure varies on the intensive margin (the number or share of daughters). Second, for female candidates, exposure to daughters is associated with less supportive positions across the outcomes we examine. Third, the daughter-associated shift among men is issue-specific rather than universal: it is pronounced for separate surnames and same-sex marriage but muted for the "promotion of understanding" bill regarding sexual minorities.

These patterns contribute to existing debates in several ways. Substantively, they extend the study of "daughter effects" to a non-Western, party-centered democracy in which organizational discipline and conservative social norms should, if anything, dampen the influence of idiosyncratic personal experiences. Methodologically, by linking independently compiled family composition data to campaign-period stated positions, the paper complements prior work that focuses on roll-call votes and other revealed behaviors. Conceptually, the gender-heterogeneous associations underscore that exposure to daughters does not map mechanically onto greater progressivism among elites; instead, any family-based socialization effect appears to be conditioned by politicians' baseline constraints and incentives, and may operate differently for men and women.

At the same time, several limitations point to directions for future research. First, al-

though conditioning on parity is a standard approach for leveraging quasi-random variation in child sex composition, we do not observe birth order. As a result, we cannot use first-born gender as a cleaner source of exogenous variation, nor can we implement first-born-based stopping-rule tests that would more directly assess the validity of using the full set of children. Our descriptive and simulation-based diagnostics provide no evidence of strong gender-based stopping behavior, but they cannot rule out more subtle forms of selective fertility. Second, the sample includes relatively few female legislators, so estimates for women should be interpreted with appropriate caution. Third, our outcomes are campaign-period survey responses that reflect both preferences and strategic positioning. Understanding when and how family experiences translate into legislative behavior remains an important next step.

Future work could strengthen inference and deepen mechanisms by extending the dataset to additional election cycles, incorporating richer measures of electoral vulnerability and district ideology, and connecting family composition not only to stated positions but also to subsequent legislative activities such as roll-call behavior, bill sponsorship, and committee participation. More broadly, studying how personal experiences interact with party discipline and electoral incentives across institutional settings may clarify when private life matters for public policy, and for whom.

Our findings suggest that even in a context where partisan coordination and conservative norms may constrain individual expression, politicians' family experiences can be meaningfully associated with their publicly articulated positions on consequential questions of gender and family policy. Even though the original positive findings in Washington (2008) do not replicate in later studies (Costa et al., 2019; Green et al., 2023), our evidence from Japan underscores the value of further out-of-sample tests.

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Tables

Table 1: Summary Statistics

	(a) All Elected Candidates					(b) Children Information Available					(c) Mean Diff.
	N	Mean	SD	Min	Max	N	Mean	SD	Min	Max	(b) − (a)
Male Dummy	465	0.9	0.3	0	1	298	0.91	0.29	0	1	0
Age	465	55.53	10.09	29	82	298	55.9	9.91	29	82	0.36
Endorsing Party Dummies:											
Liberal Democratic Party	465	0.56	0.5	0	1	298	0.6	0.49	0	1	0.04
Constitutional Democratic Party of Japan	465	0.21	0.41	0	1	298	0.21	0.41	0	1	0.01
Komeito	465	0.07	0.25	0	1	298	0.08	0.27	0	1	0.01
Japanese Communist Party	465	0.02	0.15	0	1	298	0.02	0.14	0	1	−0
Japan Innovation Party	465	0.09	0.28	0	1	298	0.05	0.21	0	1	-0.04^{**}
Democratic Party for the People	465	0.02	0.15	0	1	298	0.01	0.1	0	1	−0.01
Social Democratic Party	465	0	0.05	0	1	298	0	0.06	0	1	0
Reiwa Shinsengumi	465	0.01	0.08	0	1	298	0.01	0.1	0	1	0
Independent	465	0.02	0.15	0	1	298	0.02	0.14	0	1	−0
# of Children	298	2.01	1.06	0	6	298	2.01	1.06	0	6	–
# of Daughters	298	0.97	0.95	0	5	298	0.97	0.95	0	5	–
# of Terms Won	465	3.62	3.11	0	17	298	3.96	3.17	0	17	0.34
Incumbent Dummy	465	2.53	0.82	1	3	298	2.61	0.77	1	3	0.08
PR-Only Candidate Dummy	465	0.1	0.3	0	1	298	0.1	0.3	0	1	−0
Vote Share Won (SMD)	419	49.48	12.1	10.1	84.07	269	51.07	12.13	17.07	84.07	1.59^{*}
SMD Winner Dummy	465	0.62	0.49	0	1	298	0.64	0.48	0	1	0.02
Responded to Taniguchi-Asahi Dummy	465	0.96	0.19	0	1	298	0.95	0.21	0	1	−0.01
Children Info. Missing Dummy	465	0.36	0.48	0	1	298	0	0	0	0	-0.36^{***}
Support Separate Surnames Dummy	438	0.55	0.5	0	1	278	0.54	0.5	0	1	−0.01
Support Same-Sex Marriage Dummy	433	0.42	0.49	0	1	276	0.43	0.5	0	1	0.01
Support Sexual Minorities Dummy	434	0.66	0.47	0	1	276	0.65	0.48	0	1	−0.02

Notes: Panel (a) reports summary statistics for all candidates elected to the House of Representatives in 2021. Panel (b) reports summary statistics for the subsample of elected candidates for whom information on the number and gender of children is available (`missing_child_info` = 0). Panel (c) reports the difference in means between panels (b) and (a). Means of indicator variables can be interpreted as proportions. For panel (c), asterisks denote the results of a two-sided test of the null hypothesis that the mean difference between panels (b) and (a) is equal to zero. The asterisks ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 2: Distribution of Candidates by Number of Children and Number of Daughters

	# of Daughters								
# of Children	0	1	2	3	4	5	6	Total	Mean # of Daughters
0	25	–	–	–	–	–	–	25	–
1	34	25	–	–	–	–	–	59	0.42
2	36	64	27	–	–	–	–	127	0.93
3	12	20	25	11	–	–	–	68	1.51
4	1	4	6	2	3	–	–	16	2.12
5	0	1	0	0	0	0	–	1	1.00
6	0	0	0	0	1	1	0	2	4.50
Total	108	114	58	13	4	1	0	298	

Notes: The table reports the joint distribution of candidates by total number of children and number of daughters. Dashes indicate cells that are zero by construction.

Table 3: Systematic Tests for Identification Validity

Specification / Stopping Rule	**P-value**	**Power**
H_0: Gender-based Stopping		
R1: Stop-at-First-Son	< 0.001	1.00
R2: Min-2-then-Stop-at-Son	< 0.001	1.00
R3: Logistic Hazard ($\lambda = 1.0$)	< 0.001	0.47
R3: Logistic Hazard ($\lambda = 2.0$)	< 0.001	0.99
R4: Stop-at-First-Daughter	< 0.001	0.99
R5: Balance Preference	< 0.001	1.00
R6: Stop-at-Two-Sons	< 0.001	1.00
R7: Resignation Rule (Cap=3)	< 0.001	1.00
R7: Resignation Rule (Cap=4)	< 0.001	1.00
R8: Homogeneity Preference	< 0.001	1.00

Notes: Panel A reports p-values from within-family-size Binomial goodness-of-fit tests based on 5,000 bootstrap iterations. Panel B reports p-values from parametric bootstrap tests (2,000 iterations) for specific behavioral rules. Statistical power represents the probability of correctly rejecting each gender-based rule at the 5% significance level, given the observed sample size (N=273). Power is estimated by generating 200 synthetic datasets under each rule and calculating the rejection rate against the random fertility null. λ indicates the coefficient for having at least one son in the logistic stopping hazard, where higher values reflect stronger son preference. Cap denotes the maximum number of children a family reaches before stopping, regardless of gender composition.

Table 4: Daughters and Candidates' Support for Optional Separate Surnames

	Support Separate Surnames		
	(1)	(2)	(3)
# of Daughters	-0.102^{*}		
	(0.052)		
(# of Daughters) × Male Dummy	0.160^{**}		
	(0.065)		
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$		-0.163^{**}	
		(0.061)	
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\} \times$ Male Dummy		0.166^{*}	
		(0.080)	
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$			-0.136
			(0.094)
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}} \times$ Male Dummy			0.216^{*}
			(0.110)
Male Dummy	-0.080	-0.084	-0.080
	(0.154)	(0.161)	(0.160)
Party FE	Yes	Yes	Yes
Term FE	Yes	Yes	Yes
(# of Children) FE	Yes	Yes	Yes
(# of Children) × Male Dummy FE	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes
PR-Only Candidate Dummy	Yes	Yes	Yes
PR Block FE	Yes	Yes	Yes
Age and Age^2	Yes	Yes	Yes
$\widehat{\beta}_1 + \widehat{\beta}_2$	0.057^{***}	0.003	0.081^{***}
Observations	278	278	278

Notes: The dependent variable equals one if the respondent answered "strongly agree" or "agree" to the question item, "If a married couple so desires, the law should allow spouses to retain their respective premarital surnames after marriage", and zero otherwise. All models are estimated using a linear probability model. Standard errors are two-way clustered at the party and prefecture levels and reported in parentheses. The row labeled $\widehat{\beta}_1 + \widehat{\beta}_2$ reports the marginal effect of daughters for male candidates, computed as the sum of the main effect and the interaction with the male dummy; p-values are obtained from a linear hypothesis test against the null hypothesis that $\beta_1 + \beta_2 = 0$. The asterisks ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 5: Daughters and Candidates' Support for Same-Sex Marriage

	Support Same-Sex Marriage		
	(1)	(2)	(3)
# of Daughters	-0.199^{**}		
	(0.062)		
(# of Daughters) × Male Dummy	0.286^{***}		
	(0.072)		
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$		-0.459^{***}	
		(0.055)	
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$ × Male Dummy		0.538^{***}	
		(0.081)	
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$			-0.338^{***}
			(0.088)
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$ × Male Dummy			0.466^{***}
			(0.116)
Male Dummy	-0.257	-0.263	-0.257
	(0.155)	(0.162)	(0.160)
Party FE	Yes	Yes	Yes
Term FE	Yes	Yes	Yes
(# of Children) FE	Yes	Yes	Yes
(# of Children) × Male Dummy FE	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes
PR-Only Candidate Dummy	Yes	Yes	Yes
PR Block FE	Yes	Yes	Yes
Age and Age^2	Yes	Yes	Yes
$\widehat{\beta}_1 + \widehat{\beta}_2$	0.087^{**}	0.079^{*}	0.127^{*}
Observations	265	265	265

Notes: The dependent variable equals one if the respondent answered "strongly agree" or "agree" to the question item, "Marriage between two men or two women should be legally recognized", and zero otherwise. All models are estimated using a linear probability model. Standard errors are two-way clustered at the party and prefecture levels and reported in parentheses. The row labeled $\widehat{\beta}_1 + \widehat{\beta}_2$ reports the marginal effect of daughters for male candidates, computed as the sum of the main effect and the interaction with the male dummy; p-values are obtained from a linear hypothesis test against the null hypothesis that $\beta_1 + \beta_2 = 0$. The asterisks ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 6: Daughters and Candidates' Support for Sexual Minorities

	Support Sexual Minorities		
	(1)	(2)	(3)
# of Daughters	-0.135^{**}		
	(0.040)		
(# of Daughters) × Male Dummy	0.138^{***}		
	(0.039)		
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$		-0.212	
		(0.114)	
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\} \times$ Male Dummy		0.180	
		(0.119)	
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$			-0.293^{***}
			(0.025)
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}} \times$ Male Dummy			0.313^{***}
			(0.022)
Male Dummy	0.113^{**}	0.112	0.112^{*}
	(0.038)	(0.069)	(0.052)
Party FE	Yes	Yes	Yes
Term FE	Yes	Yes	Yes
(# of Children) FE	Yes	Yes	Yes
(# of Children) × Male Dummy FE	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes
PR-Only Candidate Dummy	Yes	Yes	Yes
PR Block FE	Yes	Yes	Yes
Age and Age^2	Yes	Yes	Yes
$\widehat{\beta}_1 + \widehat{\beta}_2$	0.004	-0.032	0.020
Observations	265	265	265

Notes: The dependent variable equals one if the respondent answered "strongly agree" or "agree" to the question item, "A 'promotion of understanding'bill regarding sexual minorities such as LGBT people should be enacted promptly", and zero otherwise. All models are estimated using a linear probability model. Standard errors are two-way clustered at the party and prefecture levels and reported in parentheses. The row labeled $\widehat{\beta}_1 + \widehat{\beta}_2$ reports the marginal effect of daughters for male candidates, computed as the sum of the main effect and the interaction with the male dummy; p-values are obtained from a linear hypothesis test against the null hypothesis that $\beta_1 + \beta_2 = 0$. The asterisks ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Appendix A. Robustness Check

Table A.1: Nonlinear Models for Optional Separate Surnames

	Support Separate Surnames		
	(1)	(2)	(3)
Panel A: Binary Logit (AME)			
# of Daughters	−0.041		
	(0.054)		
(# of Daughters) × Male Dummy	0.087		
	(0.063)		
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$		−0.164	
		(0.101)	
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\} \times$ Male Dummy		0.140	
		(0.100)	
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$			−0.093
			(0.076)
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}} \times$ Male Dummy			0.152*
			(0.092)
Male Marginal Effect	0.045	−0.024	0.059
	(0.031)	(0.065)	(0.062)
Panel B: Ordered Logit (AME on Strongly Agree)			
# of Daughters	0.010*		
	(0.006)		
(# of Daughters) × Male Dummy	0.027***		
	(0.009)		
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$		−0.038***	
		(0.011)	
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\} \times$ Male Dummy		0.031**	
		(0.015)	
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$			−0.048***
			(0.015)
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}} \times$ Male Dummy			0.099***
			(0.021)
Male Marginal Effect	0.038***	−0.007	0.051
	(0.014)	(0.026)	(0.035)
Controls	Yes	Yes	Yes
Party FE	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes
Observations	278	278	278

Notes: The dependent variable in Panel A equals one if the respondent answered "strongly agree" or "agree" to the question item, "If a married couple so desires, the law should allow spouses to retain their respective premarital surnames after marriage", and zero otherwise. Panel A reports Average Marginal Effects (AME) estimated using Binary Logit models. Panel B reports AMEs estimated using Ordered Logit models for the probability of selecting "Strongly Agree". Standard errors are clustered at the prefecture level and reported in parentheses. The row labeled 'Male Marginal Effect' reports the marginal effect of daughters for male candidates. The asterisks ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table A.2: Nonlinear Models for Same-Sex Marriage

	Support Same-Sex Marriage		
	(1)	(2)	(3)
Panel A: Binary Logit (AME)			
# of Daughters	-0.124^{***}		
	(0.043)		
(# of Daughters) × Male Dummy	0.192^{***}		
	(0.054)		
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$		-0.418^{***}	
		(0.023)	
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$ × Male Dummy		0.479^{***}	
		(0.058)	
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$			-0.217^{***}
			(0.061)
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$ × Male Dummy			0.293^{***}
			(0.088)
Male Marginal Effect	0.068^{***}	0.061	0.076
	(0.026)	(0.057)	(0.052)
Panel B: Ordered Logit (AME on Strongly Agree)			
# of Daughters	0.011		
	(0.007)		
(# of Daughters) × Male Dummy	0.039^{***}		
	(0.008)		
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$		-0.052^{***}	
		(0.012)	
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$ × Male Dummy		0.086^{***}	
		(0.008)	
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$			−0.006
			(0.015)
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$ × Male Dummy			0.070^{***}
			(0.011)
Male Marginal Effect	0.050^{***}	0.034^{*}	0.064^{***}
	(0.011)	(0.018)	(0.024)
Controls	Yes	Yes	Yes
Party FE	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes
Observations	276	276	276

Notes: The dependent variable in Panel A equals one if the respondent answered "strongly agree" or "agree" to the question item, "Marriage between two men or two women should be legally recognized", and zero otherwise. Panel A reports Average Marginal Effects (AME) estimated using Binary Logit models. Panel B reports AMEs estimated using Ordered Logit models for the probability of selecting "Strongly Agree". Standard errors are clustered at the prefecture level and reported in parentheses. The row labeled 'Male Marginal Effect' reports the marginal effect of daughters for male candidates. The asterisks ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table A.3: Nonlinear Models for Support for Sexual Minorities

	Support Sexual Minorities		
	(1)	(2)	(3)
Panel A: Binary Logit (AME)			
# of Daughters	−0.059		
	(0.086)		
(# of Daughters) × Male Dummy	0.057		
	(0.094)		
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$		0.008	
		(0.164)	
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\} \times$ Male Dummy		−0.081	
		(0.203)	
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$			0.014
			(0.147)
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}} \times$ Male Dummy			−0.018
			(0.193)
Male Marginal Effect	−0.002	−0.073	−0.004
	(0.038)	(0.082)	(0.082)
Panel B: Ordered Logit (AME on Strongly Agree)			
# of Daughters	0.005		
	(0.014)		
(# of Daughters) × Male Dummy	0.013		
	(0.010)		
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$		0.023	
		(0.023)	
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\} \times$ Male Dummy		−0.062***	
		(0.018)	
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$			0.041
			(0.027)
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}} \times$ Male Dummy			−0.023
			(0.018)
Male Marginal Effect	0.018	−0.039	0.017
	(0.018)	(0.038)	(0.044)
Controls	Yes	Yes	Yes
Party FE	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes
Observations	276	276	276

Notes: The dependent variable in Panel A equals one if the respondent answered "strongly agree" or "agree" to the question item, "A 'promotion of understanding'bill regarding sexual minorities such as LGBT people should be enacted promptly", and zero otherwise. Panel A reports Average Marginal Effects (AME) estimated using Binary Logit models. Panel B reports AMEs estimated using Ordered Logit models for the probability of selecting "Strongly Agree". Standard errors are clustered at the prefecture level and reported in parentheses. The row labeled 'Male Marginal Effect' reports the marginal effect of daughters for male candidates. The asterisks ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table A.4: Support for Optional Separate Surnames (Legislators Aged 50+)

	Support Separate Surnames		
	(1)	(2)	(3)
# of Daughters	−0.090		
	(0.069)		
(# of Daughters) × Male Dummy	0.132		
	(0.078)		
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$		−0.026	
		(0.037)	
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\} \times$ Male Dummy		0.029	
		(0.036)	
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$			−0.191
			(0.146)
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}} \times$ Male Dummy			0.253
			(0.178)
Male Dummy	−0.047	−0.057	−0.049
	(0.117)	(0.119)	(0.118)
Party FE	Yes	Yes	Yes
Term FE	Yes	Yes	Yes
(# of Children) FE	Yes	Yes	Yes
(# of Children) × Male Dummy FE	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes
PR-Only Candidate Dummy	Yes	Yes	Yes
PR Block FE	Yes	Yes	Yes
Age and Age^2	Yes	Yes	Yes
$\widehat{\beta}_1 + \widehat{\beta}_2$	0.042	0.004	0.062
Observations	183	183	183

Notes: The dependent variable equals one if the respondent answered "strongly agree" or "agree" to the question item, "If a married couple so desires, the law should allow spouses to retain their respective premarital surnames after marriage", and zero otherwise. All models are estimated using a linear probability model. Standard errors are two-way clustered at the party and prefecture levels and reported in parentheses. The row labeled $\widehat{\beta}_1 + \widehat{\beta}_2$ reports the marginal effect of daughters for male candidates, computed as the sum of the main effect and the interaction with the male dummy; p-values are obtained from a linear hypothesis test against the null hypothesis that $\beta_1 + \beta_2 = 0$. The asterisks ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table A.5: Support for Same-Sex Marriage (Legislators Aged 50+)

	Support Same-Sex Marriage		
	(1)	(2)	(3)
# of Daughters	−0.178		
	(0.253)		
(# of Daughters) × Male Dummy	0.225		
	(0.278)		
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$		−0.601	
		(0.320)	
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\} \times$ Male Dummy		0.619	
		(0.355)	
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$			−0.376
			(0.516)
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}} \times$ Male Dummy			0.422
			(0.534)
Male Dummy	−0.138	−0.132	−0.144
	(0.111)	(0.112)	(0.110)
Party FE	Yes	Yes	Yes
Term FE	Yes	Yes	Yes
(# of Children) FE	Yes	Yes	Yes
(# of Children) × Male Dummy FE	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes
PR-Only Candidate Dummy	Yes	Yes	Yes
PR Block FE	Yes	Yes	Yes
Age and Age^2	Yes	Yes	Yes
$\widehat{\beta}_1 + \widehat{\beta}_2$	0.048	0.018	0.046
Observations	182	182	182

Notes: The dependent variable equals one if the respondent answered "strongly agree" or "agree" to the question item, "Marriage between two men or two women should be legally recognized", and zero otherwise. All models are estimated using a linear probability model. Standard errors are two-way clustered at the party and prefecture levels and reported in parentheses. The row labeled $\widehat{\beta}_1 + \widehat{\beta}_2$ reports the marginal effect of daughters for male candidates, computed as the sum of the main effect and the interaction with the male dummy; p-values are obtained from a linear hypothesis test against the null hypothesis that $\beta_1 + \beta_2 = 0$. The asterisks ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table A.6: Support for Sexual Minorities (Legislators Aged 50+)

	Support Sexual Minorities		
	(1)	(2)	(3)
# of Daughters	0.044		
	(0.075)		
(# of Daughters) × Male Dummy	−0.022		
	(0.085)		
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\}$		0.124	
		(0.195)	
$\mathbb{1}\{\#\text{ of Daughters} \geq 1\} \times$ Male Dummy		−0.134	
		(0.191)	
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}}$			0.077
			(0.164)
$\frac{\#\text{ of Daughters}}{\#\text{ of Children}} \times$ Male Dummy			−0.060
			(0.163)
Male Dummy	0.329***	0.320***	0.326***
	(0.084)	(0.086)	(0.082)
Party FE	Yes	Yes	Yes
Term FE	Yes	Yes	Yes
(# of Children) FE	Yes	Yes	Yes
(# of Children) × Male Dummy FE	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes
PR-Only Candidate Dummy	Yes	Yes	Yes
PR Block FE	Yes	Yes	Yes
Age and Age^2	Yes	Yes	Yes
$\widehat{\beta}_1 + \widehat{\beta}_2$	0.022	−0.010	0.017
Observations	180	180	180

Notes: The dependent variable equals one if the respondent answered "strongly agree" or "agree" to the question item, "A 'promotion of understanding'bill regarding sexual minorities such as LGBT people should be enacted promptly", and zero otherwise. All models are estimated using a linear probability model. Standard errors are two-way clustered at the party and prefecture levels and reported in parentheses. The row labeled $\widehat{\beta}_1 + \widehat{\beta}_2$ reports the marginal effect of daughters for male candidates, computed as the sum of the main effect and the interaction with the male dummy; p-values are obtained from a linear hypothesis test against the null hypothesis that $\beta_1 + \beta_2 = 0$. The asterisks ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.