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ABSTRACT

How Does Gender Quota Shape Gender Attitudes?*

Starting in 2002, each electoral district in Taiwan is required to reserve one seat for women out of every four seats at the local councils, creating a 'zigzag' function of female councilors proportion across districts. Using this function, we estimate the effects of exposure to female political leadership on gender attitudes and behaviors. For son preference, we find that the gender quota reduced the propensity for parents with two daughters to give a third-parity birth, whereas there is no such effect on parents with two children of other sex compositions. This finding is confirmed by the evidence of self-reported son preference revealed in survey data. We further find that the gender quota encouraged female high school graduates to attempt college admission and pursue a college major in law or political science, which commonly prepare students to develop a career in politics. The gender quota also empowers married women to play a more active role in multiple household decisions. Our findings support the hypothesis that female political leaders perform as role models who elicit more egalitarian gender attitudes from women.

JEL Classification: J16, J13, I23, D10

Keywords: gender quota, son preference, gender role model, gender norms

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I. Introduction

Enhancement of women's status in a society does not necessarily synchronize with its economic growth or democratization. In the case of East Asia, females in Taiwan now enjoy a noticeably better political representation, economic opportunities, and social participation than their peers in Japan and South Korea: The proportion of seats held by women in Taiwan's national parliament is 41.6 percent in 2021, compared to 19 percent in South Korea and 9.9 percent in Japan. In Taiwan, 25.9 percent of high school principals are female in 2021, much higher than the corresponding figures for Japan (8.6 percent) and South Korea (13.0 percent). In 2020, the gender wage gap was 14.8 percent in Taiwan, compared to 30.7 percent in Japan and 30.4 percent in South Korea. More generally, regarding the Gender Gap Index created by the World Economic Forum to measure gender parity across 149 countries, Taiwan ranked 32nd in 2018, while Japan ranked 110th and South Korea 115th.[1] In addition to women's status, Taiwan appears to lead the way in enacting gender-friendly laws and policies. Abortion was virtually legalized in Taiwan as early as 1985, while it was banned in South Korea until 2020, and, in Japan, abortion is legal with strict limitations. Moreover, in 2019, Taiwan became the first country in Asia to recognize same-sex marriage nationwide.

The advancement of women's status and gender attitudes in Taiwan relative to Japan and South Korea is puzzling, given the shared Confucianism-rooted culture, persistent post-WWII economic growth, and development of democracy. It is also important to notice that, prior to WWII, the status of women in Taiwan was extremely stifling: Child brides and polygamy were prevalent, women were deprived of inheritance rights, and the gender ratio remained exceedingly high due to neglect of girls and even gendercide. This implies that the advancement in gender parity in Taiwan can only be achieved by institutional changes and other factors that occurred in the post-WWII era.

In this paper, we explore a unique institutional initiative in Taiwan – the re-

[1]The Gender Gap Index (GGI) is measured by evaluating gender parity over four key dimensions – economic participation and opportunity, educational attainment, health and survival, and political empowerment. Since 2006, the World Economic Forum (WEF) has reported GGI for 146 countries annually, providing a basis for time-consistent cross-country comparisons of gender equity. While WEF does not publish Taiwan's GGI, the Taiwan government compiles the formula and publishes the results annually. More details about GGI can be found at https://www.weforum.org/reports/.

served seats for women in the county/city councils – as a potential factor that might contribute to the improvement of women's status. The local council is the legislative government entity of the second-level administrative division in Taiwan. Councilors, with a term duration of four years, are elected through a single non-transferable vote (SNTV) system. Since 2002, a law amendment requires each electoral district to reserve one seat for women out of every four seats at every local council. Prior to the 2002 reform, the requirement was less stringent: No reservation for women if the number of seats to be elected in the district is less than five; one seat is reserved if the number ranges from 5 to 14; and two seats are reserved if the number exceeds 15. Since the number of seats assigned to each electoral district is exclusively determined by its population share within the county, the cross-district differences in the proportion of seats reserved for women provide an arguably exogenous source of variation in women's political representation. It is the variation that we exploit to estimate the impacts of exposure to female leadership on various gender-related outcomes.

Our interest is in evaluating the role of female leadership in alleviating gender norms against women, especially son preference, and gauging the relative importance of different mechanisms that might deliver these effects. Previous studies have revealed that exposure to strong female characters could have a profound impact on women, including altering their gender attitudes and decision-making authority within the household. In Taiwan, the local councils' reserved seats provide a unique opportunity to achieve these goals. Unlike national-level elections that heavily rely on public media exposure, electoral campaigns at the local level dominantly resort to face-to-face contacts and community networking. These in-person interactions and personal connections between female candidates and voters may intensify the image of female leadership perceived by voters, which might explain why prior studies found that the presence of lower-level female leaders, compared to higher-level ones, was more effective in empowering women against crime (Iyer et al., 2012) and increasing the survival of young girls in India (Kalsi, 2017).

Measuring the degree to which an individual's attitude of preferring a son to a daughter is intrinsically difficult. In the context of Taiwan, we utilize parents' propensity to have the third birth conditional on their first two children's gender composition as an indicator of son preference. Specifically, our data clearly show that parents with only two daughters exhibit a significantly higher propensity to have a third child than parents with only two sons and parents with just a son and a daughter. Therefore, the propensity gaps between parents with two daughters and other parents provide a unique behavioral measure of son preference. This echoes the rationale behind the strategy adopted by Angrist and Evans (1998) who used the gender composition of the first two children as an instrument for sibling size, assuming that parents of same-sex children are more likely to go on to have an additional child.

Our empirical work starts with estimating the extent to which the gender quota encouraged women to run for the council elections and the consequent seats filled by women. We carry out the estimation using data covering four elections from 1994 to 2005, when women's political engagement was relatively scarce. The results suggest that the gender quota effectively enticed more females to run for the councilor elections (with an elasticity of 0.75) and increased elected female councilors (with an elasticity of 1.01), and these effects remained strong through all four elections. Moreover, for each of the 1,803 candidates running for the 1998 election, we track her political career in the following 20 years. We find that the female councilors elected through reserved seats in 1998, despite they gained fewer votes than other winners did, performed as competitively as other winners in winning subsequent local or higher-level elections. These findings agree with prior studies showing that the gender quota not only produced sustainable female representation at the local councils (Huang, 2016; Batto, 2019) but also paved the way for local female politicians to move up the political ladder into higher offices (Batto, 2018; Huang, 2019).

Our identification strategy relies on comparing the son preference in the electoral districts being assigned a higher proportion of seats reserved for women to that in districts with a lower proportion of reserved seats. Using universal birth records, we show that the gender quota significantly reduced the propensity for parents with two daughters to give a third-parity birth; in sharp contrast, there were no such effects on parents with two sons or just a son and a daughter (regardless of the birth orders). For parents with only one child, we also find that the gender quota had a negative effect on the propensity to have a second-parity child for parents with one daughter, but the effect is only moderate and statistically insignificant. To shed more light on the impact of gender quota on son preference, we exploit survey data that report individuals' self-reported son preference. We find that women living in districts with a higher gender quota proportion exhibited a weaker preference for sons after the election. In contrast, the gender quota does not make such a difference in men's preference. Furthermore, the attenuation effect on women's son preference was concentrated on women over 45 years old, implying that the negative gender quota effect on two-daughter mothers may come from the weakening of son preference held by the grandmothers (namely, the two-daughter mothers' mothers and/or mothers in law), who commonly exert a strong influence over younger generation's birth decisions in Taiwan.

The gender-specific effects on son preference found above support a potential mechanism that could deliver the gender quota effect on fertility: Female political leaders can work as role models for women, making females' gender attitudes less traditional and more egalitarian, ultimately alleviating their desirability for sons (we name this gender role model hypothesis' henceforth). Related, since men's preference is found to be neutral to gender quota, it is likely that women's unilateral change in son preference translated into fertility decisions through intra-household bargaining (we name this intra-household bargaining hypothesis' henceforth). In Section V, we conduct multiple examinations to evaluate the two hypotheses. We first find that female high school graduates in districts with a higher gender quota proportion were more likely to attempt college admission by taking the national college entrance exam, but there is no such effect on male high school graduates. Moreover, when attempting college admission, female high school graduates with higher gender quota exposure were more inclined to select a college major in law or political science, which often prepares students for a career in politics; however, no such effect can be detected for male high school graduates. These gender-specific effects align with the predictions made by the gender role model hypothesis.

To examine the intra-household bargaining hypothesis, we exploit the family relationship data from the Survey of Social Development Trend (SSDT) in 1998 and 2002. The surveys asked both married couples to specify the primary decision maker on four types of household decisions: expenditure, parenting, saving and financing, and allocation of chores between household members. The results of our analysis suggest that married couples in higher exposure districts tend to report that the wives play a more important role relative to the husbands in making these household decisions. These findings lend credit to the hypothesis that exposure to female political leaders empowers women and translates women's preference into a halt of the son-targeting fertility rule through household bargaining.

The practice of the gender quota in Taiwan's local councilor elections provides a rare opportunity to observe its impacts on women's political representation and the subsequent effects on gender attitudes and behaviors. The findings of this study add valuable evidence to a series of previous studies that leverage the introduction of reserved village head positions for women in India and found its profound effects, including discouraging parents from practicing sex selection (Kalsi, 2017), increasing adolescent girls' school enrollment (O'Connell, 2018) and their aspirations for education (Beaman et al., 2012), reducing voter's resentment against female leaders (Beaman et al., 2009), promoting women's entrepreneurship (Ghani, Kerr and O'Connell, 2014), and empowering women to report crimes (Iyer et al., 2012). Our findings in this paper agree with these studies in that the gender quota in Taiwan empowers women to develop more egalitarian gender values and educational aspirations.

Our empirical analyses also contribute to a broader understanding of the environments where female role models may have an impact on women. Role models characterized by female teachers or mentors are often cited as an important factor affecting female students' educational and vocational outcomes, including academic performance (Dee, 2007; Carrell, Page and West, 2010; Muralidharan and Sheth, 2016; Lim and Meer, 2017), selection of STEM as college majors (Carrell, Page and West, 2010; Lim and Meer, 2020; Breda et al., 2023), and occupational preference and choice (Kofoed and mcGovney, 2019). Another strand of studies extends the gender role model effects to the political realm by showing that exposure to female political leadership attenuates traditional gender norms held by women (Kalsi, 2017), increases adolescent girls' school enrollment (O'Connell, 2018), and reduces gender gaps in educational attainment and aspiration (Beaman et al., 2012). Our findings in this paper add to this line of research by showing evidence supporting the gender role model hypothesis.

More broadly, our work is related to a large body of literature that examined the effectiveness of gender quota in boosting women's political representation. Our findings agree with abundant prior evidence suggesting that gender quota grants females an opportunity to gain low-level positions in the short term (Tripp and Kang, 2008; Jones, 2009; Schwindt-Bayer, 2009; Paxton, Hughes and Painter, 2010). In addition, we add valuable evidence to support the few existing studies examining whether gender quota helps females to ascend the political ladder to more powerful positions in the medium or long term (O'brien and Rickne, 2016; Baltrunaite et al., 2019; Bagues and Campa, 2021; Lippmann, 2021).

The remainder of this paper is organized as follows. Descriptions of the institutional background on the gender quota in Taiwan are provided in Section II. Section III offers explanations for the data and the key variables used in this study, followed by Section IV, which assesses the effectiveness of gender quota in enhancing women's political representation. Section V presents the estimation strategies and estimated effects of the gender quota on son preference, as well as effects on other outcomes which we use to evaluate the potential mechanisms. We discuss the results in Section VI, and conclude the paper in Section VII.

II. Institutional background

Gender quota was codified into Taiwan's Constitution in 1946, which requires a certain number of seats reserved for women at different levels of legislative entities.[2] Following the constitutional design, 5 to 10 percent of seats at the local councils should be reserved for women, starting the first council election in 1950. However, through the entire period from 1950 to around 1998, the actual seats taken by women at the local councils were kept at the minimum requirement level, mainly due to the authoritarian ruling of the Kuomintang (KMT) regime, which deliberately suppressed civil rights and manipulated elections (Huang, 2016). Only until the 1990s did the political environment witness a dramatic change, with the authoritarian regime being replaced by a democratic system and elections facing competition between different political parties. Starting in 1996, the Democratic Progressive Party (DPP), then the largest opposition

[2] Article 134 of the Constitution of the Republic of China (Taiwan) states, "In all elections, a certain number of seats should be reserved for women, and measures pertaining thereto shall that prescribed by law."

party, amended its party charter to require that one out of every four nominees must be a woman for all levels of election. The KMT followed suit by passing similar nomination rules in 2000.

In 1999, a major reform mandated that each electoral district reserves one out of every four seats for women at every county/city council. Prior to the reform, the quota was less demanding: No reserved seat for women is required if the number of winners in the electoral district is less than five; one seat is reserved if the number ranges from 5 to 14; and two seats are reserved if the number exceeds 15. The 1999 reform had a pivotal impact on female representation at the local councils. As shown in Figure 1 (the solid line), the overall proportion of female councilors experienced considerable growth from 15 percent in the late 1990s to 27 percent in the 2010s.

The assignment of gender quota was not limited to local councils; a certain number of seats at the Legislative Yuan, the national parliament in Taiwan, were also reserved for women when the legislators were elected through an SNTV system prior to 2004. Similar to the growing trend experienced by the female councilors after the 1990s, the proportion of female parliament members also exhibits a continuous growth from the 1990s to 2010s, as shown by the dashed line in Figure 1. The practice of gender quotas at the national and local levels has reshaped the landscape of female political representativeness in Taiwan after the mid-1990s. By elevating the winning odds for female candidates, the gender quota not only incentivized parties to nominate more females, but also more qualified ones, to run for elections. Huang (2019) pointed out that, compared to the male candidates who were replaced due to the reserved seats, the female winners elected through the reserved seats were as qualified in terms of education and political experience. In 2021, 41.6 percent of the parliament members were female, the highest share among Asian countries.

No similar design of Taiwan's gender quota has been seen in any other East Asian country. In Japan, for example, neither the National Diet (House of Representatives and House of Councilors) nor the local councils reserve seats for women. In South Korea, some seats are reserved for women at Korean National Assembly, its national legislative entity, and local councils but on a much smaller scale than that adopted in Taiwan. Both elections follow the mixed member system: Voters cast two separate ballots, one for party-listed candidates and another for district candidates, and the system has been in place since 1960 (Croissant, 2002). The Korean National Assembly has 300 members, 253 were elected through single-seat constituencies and 47 through party lists. The gender quota was introduced into the system in the early 2000s, which requires parties to nominate a woman in every other position on their party list, thus accounting for roughly 50 percent of the party-list seats, or 7.8 percent of the entire parliamentary seats. No district seat is reserved for women, but financial subsidies are provided to parties that nominate females accounting for at least 30 percent of district candidates. For local councils, the party-list seats account for a smaller portion, often around 10 percent, of the total seats, and 50 percent of the party-list nominees must be female. For district seats at the local councils, again, no gender quota is required, but parties are encouraged to nominate female candidates. Up to 2020, the proportion of female parliament members is 19 percent in South Korea and only 9.9 percent in Japan.

Unlike national election campaigns that heavily rely on public media and digital platforms, local council elections in Taiwan mainly resort to face-to-face connections built through field campaigning, such as community gathering, spot canvassing, and door-to-door visiting. This is because local politics in Taiwan typically revolves around grassroots issues, particularly interpersonal networks and relationships (Jacobs, 1979). Voters therefore have opportunities to closely contact with the candidates, often creating a more personal connection between them. Furthermore, the images of female candidates are extensively exposed to the general public through campaign signage, including building banners and street flags, which are widely displayed during the campaign period. It is not difficult to imagine that female candidates would attract considerable attention in campaigns conventionally dominated by male candidates, and the close voter-candidate contacts may intensify the image of female political leadership perceived by voters. After the campaign ends, the exposure of female leadership may continue through other channels. First, compared to their male colleagues, female councilors tend to offer more service to women by participating in local NGOs and feminist organizations that advocate women's rights or protect women from being harmed or abused. Second, as we will elaborate in Section VI, female councilors raise more concerns about gender-specific issues and design more female-friendly policies than male councilors, deepening their connections with local females.

III. Data and measurement

A. Election data

Our election data covered four council elections in 1994, 1998, 2002, and 2005.[3] The 1994 election data are collected from public media archives, and the remaining data are from the database of Taiwan's Central Election Committee. For each election, the data report every candidate's name and characteristics, including gender, age, and party membership, as well as electoral variables, including votes won, winning/losing outcome, and the electoral district each candidate participated in. We construct our instrumental variable by leveraging the assignment rules of the gender quota, which creates variation in the proportion of women's seats across districts. Specifically, let R_{de} denotes the proportion of seats reserved for women in electoral district d in the election year e, and the 'gender-quota function' derived from the assignment rules in 2002 and 2005 can be expressed as:

$$R_{de} = \text{GreatInt}\left(\frac{E_{de}}{4}\right)/E_{de} \tag{1}$$

where the function GreatInt(n) returns the greatest integer i such that $i \leq n$, and E_{de} refers to the number of seats to be elected in district d for election e. Equation 1 captures the fact that the gender quota rule demands no reserved seat for women in electoral districts with 1-3 winners, one reserved seat for districts with 4-7 winners, two reserved seats for districts with 8-11 winners, and so on, which amounts to 25 percent of the total seats at most. Following the same line, the gender-quota function for elections in 1994 and 1998 is:

$$R_{de} = \text{Round}\left(\frac{E_{de}}{10}\right)/E_{de} \tag{2}$$

[3]The 2005 council election was expedited because it was held together with the county mayor election. Prior to 2005, the mayor election was held one year before the council election.

where the function Round(n) returns the integer that is closest to n. Likewise, Equation 2 encapsulates the rules that assign no reserved seat for women for electoral districts with 1-4 winners, one reserved seat for districts with 5-14 winners, and two reserved seats for districts with 15-24 winners. The orange curve in Figure 2(A) graphically illustrates the gender-quota function for the 2002 election, where the orange curve describes the gender-quota function as in Equation 1, which exhibits a zigzag' shape with corners occurring at multipliers of 4. The blue curve in the same figure presents the average proportion of actual elected women, which closely imitates the same zigzag shape: the proportion of female winners remains low when the number of winners is below 4 thus no seat is reserved for women; thereafter, the proportion exhibits a peak at each of the corner points and declines between any two consecutive corners. Interestingly, the zigzag shape of the gender-quota function closely mimics the shape of the class-size function if the Maimonides' rule is used to put an upper limit of class size (Angrist and Lavy, 1999; Angrist et al., 2019; He and Bartalotti, 2020).

Figure 2(B) presents the corresponding curves using the 1998 election data, where the gender-quota function, as depicted by the orange curve, follows Equation 2 and exhibits only two corner points (at 5 and 15). Similar to the patterns of its counterpart in Figure 2(A), the blue curve in Figure 2(B) also heaps at the first corner point at 5, followed by a declining trend between 5 and 15. In both graphs, the positive correlation between the proportion of female winners and the gender quota function implies a contribution made by the gender quota to seats earned by women at the local councils.

As argued earlier in Section II, the gender quota increased the winning odds for female candidates and thus incentivized political parties to nominate more women to run for the local council elections. This can be seen in Figure 3(A), which presents the proportion of female candidates running for the 2002 election (the red curve) and the corresponding gender-quota function (the orange curve). Analogous to the blue curve in Figure 2(A), the red curve in Figure 3(A) shows that the proportion of female candidates exhibits a pattern that largely synchronizes with the gender-quota function. Finally, Figure 3(B) confirms the synchronization for the 1998 election.

B. Individual-level data

We leverage a wide range of administrative and survey data to explore the potential impacts of the gender quota. To construct our primary outcome variable, the decision on whether to have the third-parity child, we need to observe the entire birth records of every woman in our sample. To this end, we leverage the universal birth records documented in the Birth Registries, 1978 to 2006, which report every newborn child's ID, gender, birth date, birth parity, birth type (singleton or multiple births), birth weight (in grams), and her parents' characteristics at the time of the childbirth, including age, marital status, education level, and residential township. Siblings can be identified by their mother's citizen identification (ID) number, then the gender composition of a mother's first two children can be determined. For this study, we categorize mothers into four groups based on the sex composition of the mother's first two children – two daughters, two sons, an older son and a younger daughter, and an older daughter and a younger son.

We exploit the educational variables from the University Entrance Test records, 2000 to 2010, to estimate the effects of gender quota on educational aspiration and preference. These data report each student's ID, the high school graduated from, scores on the Advanced Subjects Test (AST, the primary university entrance test in Taiwan), and, if successfully admitted, the names of the university and program admitted to. We use the student's residence to identify the electoral district the student resided in at the time of the election, and use the identified district to merge with the election data. For this study, we construct two variables to shed light on individual educational aspiration. The first indicator is a dummy variable indicating whether a high school graduate attempts to apply to universities by taking up the AST. Second, the data also reveal the full list and ranking of the universities and programs each student applies to.[4] The ranking indicates the student's preference towards different academic disciplines, allowing us to investigate whether female political leaders serve as role models that inspire female students to select a college major in law or political science, which are both widely perceived as the primary disciplines that nurture future politicians. Due to data paucity, however, the indicator for AST taking up is only available from 2000 to

[4]Students could list up to 80 choices before 2004 and 100 choices after 2005 in their application list.

2003, and students' applications are only available from 2002 to 2010.

We also use individual self-reported son preference as an alternative outcome variable, which is reported in Taiwan's Social Change Surveys (TSCS), 2001 and 2006 (Chang, 2019; Fu, 2016). We link each respondent to the election data based on the year of the survey and the corresponding electoral district that her residence belongs to. Another survey data set we use is the Survey on Social Development Trends (SSDT), 1998 and 2002 (Directorate-General of Budget, Accounting and Statistics, Executive Yuan, 2004, 2006). All married couples in SSDT surveys are asked to specify who (the husband, wife, or both) is primarily responsible for the four household decisions – consumption, parenting, finance, and chore assignment. We use these answers to construct multiple variables indicating intra-household bargaining power. Using the same data-merging approach used for the TSCS, we link the election data and the SSDT data at the electoral district level. Our goal is to examine whether exposure to female political leaders reshapes the role of women as decision-makers in the household and, in turn, alters their birth decisions.

IV. The impacts of gender quota on female political representation

In Section III, Figure 2 and 3 have illustrated that both the proportion of seats taken by women and the proportion of female candidates running for the elections demonstrate a zigzag pattern that closely follows the shape of gender quota function. The pattern implies that the gender quota has effectively enhanced women's political representation at the city/county councils. Here, we move further to conduct a regression estimation of the correlations and reveal the statistical significance. The regression is specified as:

$$F_{dce} = \alpha_1 + \beta_1 R_{dce} + \pi_1 P_{dce} + \rho_c + \sigma_e + u_{dce} \tag{3}$$

where F_{dce} refers a treatment variable, either the proportion of female running candidates or the proportion of seats won by females in electoral district d of county c for election e; R_{dce} is the gender-quota function described as in Equation 1 for elections held in 2002 or 2005 ($e \geq 2002$), and as in Equation 2 for elections held in 1994 or 1998 ($e \leq 1998$); P_{dce} denotes the population size of electoral district d of county c in the year just before the year when election e was held; ρ_c and σ_e respectively represent fixed effects for counties and elections; u_{dce} is the error term.

It is noteworthy that controlling for P_{dce} is of particular importance because the gender quota is exclusively determined by the proportion of the district's population out of the entire county's population measured in the year prior to the election. After P_{dce} is controlled for, our estimation using Equation 3 returns the effect of gender quota on the outcome, indicated by β_1, net of any potential effects correlated with the population, such as the economic development level and employment opportunities of a district. Doing so reduces the potential threat of omitted variables bias that one may have a concern about. However, it may still be a concern that R_{dce} is correlated with some unobserved factors. To address this, we examine whether the vote share won by women in an election predicts the reserved seats for women in the next election. The examination is informative because if the gender quota function is indeed exclusively determined by population share, the gender quota of a district in any election should not respond to the voters' preference towards female candidates. It thus should be uncorrelated with the vote share won by women in the previous elections. Our examination is carried out by estimating the following regression:

$$R_{dce} = \alpha_1 + \beta_1 V_{dce-1} + \beta_2 C_{dce-1} + \beta_3 P_{dce} + u_{dce} \tag{4}$$

where V_{dce-1} denotes the vote share won by all female candidates in the previous election, C_{dce-1} refers to the proportion of female candidates in the prior election, and R_{dce} and P_{dce} are defined in the same way as in Equation 3. Table 1 presents the results from estimating Equation 4, where column 1 shows that the vote share won by female candidates in 1998 does not predict the assignment of reserved seats for women in the 2002 election; what matter are the proportion of female candidates running for the 1998 election, which reflects the 1998 gender quota (determined by 1997 population), and the population in 2001, which determined 2002 gender quota. Column 2 confirms these findings when the 2002 election data are used to predict the 2005 gender quota. The results presented in Table 1 support our claim that the assignment of gender quota does not reflect voters' preference towards female politicians.

Next, we use Equation 3 to estimate the effects of the gender quota on female winners and candidates, and the results are presented in Table 2. The estimate in column 1 shows a significantly positive effect of the gender quota on the proportion of seats won by women, with a one percentage-point (ppt) increase in the gender quota leading to a 1.01 ppt increase in the proportion of female seats. The elasticity is quite large given that, on average, 11.9 percent of seats were reserved for women, and only 18.6 percent of seats were finally won by women. Column 2 of Table 2 shows that a one ppt increase in the gender quota leads to a 0.753 ppt increase in the proportion of female candidates running for election, presumably because the gender quota motivates political parties to nominate more female candidates.

A natural question that arises is the extent and duration of the gender quota effects on female politicians, or, more specifically, whether the gender quota only helps female politicians gain a local position in the short term or helps them to climb the political ladder to higher offices in the long term. To shed light on this, we track all the male and female candidates, 1,803 in total, who ran for the 1998 election and compare the career development of the winners through reserved seats and that of other winners over a course of 25 years. The summary statistics in Table 3 show that 145 of the 292 female candidates won the election, with 20 of them winning through reserved seats; in comparison, 693 out of a total of 1,511 male candidates won the election.

Now, we categorize all winners into three groups: male winners, female winners through reserved seats (we name them reserved-seat winners hereafter), and other female winners who won a seat without relying on the gender quota privilege. For the three groups, we conduct a simple survival analysis and compare the continuation of their political positions from 1998 to 2023. Here, survival is defined as remaining in an elected office, either through winning a reelection at the local council or winning an election at a different level into another office, either city mayor, county premier, or national parliament legislator. Hazard is defined as a termination of servicing in any elected position, either because of losing an election or ceasing to run for any election. Figure 4 plots the Kaplan-Meier survival curves for the three groups separately, where the horizontal axis refers to years after 1998. The pink curve represents the survival curve for the reserved-seat winners, which is largely intertwined with the male winners' curve (in blue) and other female winners' curve (in light green) for the entire 25-year course. This suggests that, although the reserved-seat winners did not gain enough votes to win the election in the first place, they ended up with a long-term political career as successful as other winners. This suggests that the gender quota benefited the reserved-seat winners significantly and persistently.

To show statistical significance, we employ the Cox proportional hazards model to estimate the hazard ratios between the three groups while controlling for each candidate's party membership, age, and incumbency. As reported in Table 4, the estimated hazard ratio for the reserved-seat winners relative to the male winners (the benchmark) is 1.01, and it is statistically insignificant. This suggests that the average length of survival does not significantly differ between the two types of winners. For other female winners, the hazard ratio is 1.024 and also insignificant. These results confirm the patterns illustrated in Figure 4.

V. The effects of gender quota

A. The third-parity births

Son preference remains prevalent in Taiwan today, and more so during the period of our primary sample (1994 to 2006) when gender impartiality was more pronounced then. In our data, son preference can be detected by observing the practice of sex-selective abortion on higher parity births. Using the Birth Registries data from 1978 to 2006, we calibrate the annual male-to-female sex ratio by birth parity and present the results in Figure 5. The blue, red, and green curves respectively represent the sex ratio for first-, second-, and third-parity children. Before the mid-1980s, when ultrasound was yet available, the sex ratio did not differ much across birth parities. Starting in 1986, the sex ratio of third-parity births increased continuously, even exceeding 120 after 2002. During the same period, however, the sex ratios of the first- and second-parity births remain relatively stable. This finding is consistent with a large strand of prior studies showing that parental preference for having sons increases the male-to-female sex ratio in the context of East and South Asian countries (Sen, 1990; Das Gupta et al., 2003; Duflo, 2012; Lin, Liu and Qian, 2014; Almond, Li and Zhang, 2019).

The parents' preference for sons can also be manifested by Figure 6, which presents the Kaplan-Meier survival curves for mothers with two children, using the universal Birth Registries data from 1978 to 2006. The onset of risk is the timing (year and month) of the mother giving her second-parity birth, and the hazard event is defined as giving her third-parity birth. Mothers are split into four groups based on the sex composition of their first two children – two sons, two daughters, first son then daughter, and first daughter then son. Figure 6 shows that the propensity to give the third-parity birth is noticeably higher for mothers with two daughters (the blue curve) than the other three groups, highlighting the role of son preference in driving parents to practice son-targeting fertility rule.

Motivated by these findings, our strategy for estimating the effect of the gender quota on son preference focuses on exploring whether the gender quota changes the propensity of those parents with two daughters to have one more child relative to parents with two children of other sex compositions. It is noteworthy that we do not resort to the two-stage least squares (2SLS) method because the gender quota, though arguably exogenous, may result in changes in multiple factors that would have an impact on son preference. We have already shown that the gender quota not only granted women more seats in the local councils, but also encouraged more women to take the challenge to run for elections and, as a result, amplify the exposure of female leadership during the election campaigns. In light of this, we rely on the reduced-form estimation (if using the language of instrumental variable strategy) by leveraging the gender quota as an exogenous source of variation in the public exposure of female leadership. Our reduced-form regression is specified as follows:

$$\begin{aligned} Y_{idce} = & \alpha_1 + \beta_1 R_{dce} + \beta_2 R_{dce} \cdot BB_{idce} + \beta_3 R_{dce} \cdot BD_{idce} + \beta_4 R_{dce} \cdot DB_{idce} \\ & \alpha_2 BB_{idce} + \alpha_3 BD_{idce} + \alpha_4 DB_{idce} + \pi P_{dce} + \mathbf{X}_i \gamma + \rho_c + \sigma_e + \varepsilon_{idce} \end{aligned} \tag{5}$$

where Y_{idce} is a dichotomous variable indicating giving the third-parity birth ($= 1$) or not ($= 0$) for mother i who lived in electoral district d of county c during the period between election e and the next election; R_{dce} is the gender-quota function defined as in Equations 1 and 2. The three dummy variables, BB_{idce}, BD_{idce}, and DB_{idce}, respectively, indicate those mothers having two sons, a son then a daughter, and a daughter then a son (leaving mothers having two daughters, denoted as DD group henceforth, as the omitted group); $\mathbf{X}_i$ is a vector of idiosyncratic characteristics that are comprised of mother's and father's education, their ages at the year of election e, and the duration (measured by years) between the second-parity birth and election e; the population control, P_{dce}, and the fixed effects, ρ_c and σ_e, are defined as in Equation 3; finally, ε_{idce} represents the error term. Our coefficient of interest in Equation 5 is β_1, which captures the effect of gender quota on the propensity of the DD mothers to give the third-parity birth during the four years following election e. We are also interested in the sum of coefficients β_1 and β_2, which indicates the gender quota effect on the BB mothers; similarly, $\beta_1 + \beta_3$ and $\beta_1 + \beta_4$ capture the effects on the BD and DB mothers respectively.

The sample used for estimating Equation 5 contains all mothers with two children at the time of election e (1994, 1998, 2002, and 2005), and we observe each mother's decision to give the third-parity birth between elections e and $e + 1$ as our primary outcome variable. The summary statistics for the four different types of mothers are presented in Table 5. Consistent with the patterns shown in Figure 6, the DD mothers' propensity to give the third-parity birth (10.9 percent) is at least twice as high as those of the other three types of mothers. Importantly, all three key variables – the gender quota proportion, the proportion of female winners, and the proportion of female candidates – change minimally across mother types.

The results from estimating Equation 5 are presented in Table 6, where column 1 shows that the reduced-form estimate of $\hat{\beta}_1$ is -0.126. The magnitude of this estimate implies that a one standard deviation (0.072) increase in gender quota proportion leads to an 0.91 ppt (or 15 percent of mean) decrease in the propensity for the DD mothers to have the third-parity child. Meanwhile, the coefficient estimate $\hat{\beta}_2$ of the interaction term of gender quota proportion and the indicator for the BB mothers (BB_{idce}) is 0.174, implying a marginal effect on their propensity to have the third child is 0.048 ($\hat{\beta}_1 + \hat{\beta}_2 = -0.126 + 0.174 = 0.048$). Although a formal *F-test* on $\hat{\beta}_1 + \hat{\beta}_2 = 0$ rejects the null hypothesis that the marginal effect is zero, the magnitude of the effect is rather small. This is also the case for the BD and DB mothers, as $\hat{\beta}_1 + \hat{\beta}_3$ and $\hat{\beta}_1 + \hat{\beta}_4$ are 0.031 and 0.026, respectively, which are both statistically significant but small in size. A likely explanation of these fertility effects is that the exposure to female political leaders attenuated son preference and thus discouraged the DD mothers from giving a third-parity birth as an attempt to have a son. The exposure also encouraged the BB mothers, and the other two types of mothers to a lesser degree, to have a third child as an attempt to have a daughter, but these effects appear to be only moderate.

In columns 2 and 3 of Table 6, we split the entire sample into two groups – mothers with and without a higher school diploma – and repeat the estimation for the two subsamples separately. The estimate of β_1 for mothers with a high school degree is -0.141, which is larger in magnitude than that (-0.117) for mothers without a high school degree. We also carry out the same estimation separately for mothers who lived in urban areas and those who lived in rural areas, and the estimate of β_1 appears to be large in absolute value for urban mothers (-0.204) than for rural mothers (-0.115). Combined, these findings suggest that the gender quota effects on son preference are more of an urban phenomenon: Mothers with higher socio-economic status (SES) are more susceptible to exposure to female leadership, likely because they are less attached to traditional norms than lower SES mothers.

To advance our understanding of the gender quota effect on fertility decisions, we repeat the estimation of Equation 5 by replacing the dependent variable for a dummy variable indicating the gender of the third-parity child (boy=1; girl=0), using a sample of mothers who gave the third-parity birth during our sample period. Table 7 presents the results in the same format as Table 6, where the estimate of $\hat{\beta}_1$ in column 1 is 0.0984. This suggests that the gender quota significantly increases the likelihood of the DD mothers' third-parity children being boys. In light of the deterrence effect of gender quota on the DD mothers' fertility rate found above, the effect on the third-parity children's gender is likely a result of self-selection: The exposure to female leaders discouraged some DD mothers from having an additional child, but those who insisted on doing so were likely holding a stronger preference for sons, and thus were more likely to exercise sex-selective abortion, leading to a higher sex-ratio of their third-parity children. For other types of mothers, in contrast, the counterpart effects on children's gender are minimal and insignificant, as suggested by the corresponding *F-test* results displayed at the bottom of column 1. This finding is highly consistent with the results in column 1 of Table 6 in that the gender quota effects were concentrated on the DD mothers.

One may wonder whether the gender quota also affected parents' willingness to have a second-parity child and whether the effect differs between parents with a son and parents with a daughter. We estimate the effects by estimating Equation 5 with the sample being replaced for mothers with only one child during our sample period. To this end, Equation 5 is simplified for comparing the propensity to have a second-parity child for mothers who has a son and that for mothers who has a daughter (the benchmark group), using an interaction term of the gender quota proportion and a dummy variable indicating the mothers with a daughter. The estimation results are presented in Appendix Table A2, where the first coefficient estimate ($\hat{\beta}_1$) captures the effect of gender quota on the birth decisions of mothers with a son, and the second coefficient estimate ($\hat{\beta}_2$) captures the 'additional' effect for mothers with a daughter; thus, the overall effect on mothers with a daughter is represented by $\hat{\beta}_1 + \hat{\beta}_2$. Column 1 shows that the gender quota decreases the propensity to have a second-parity child for mothers with a son, but the estimate ($\hat{\beta}_1$) is minimal and statistically insignificant. For mothers with a daughter, $\hat{\beta}_1 + \hat{\beta}_2$ is negative, but the *F-test* result suggests an insignificant gender quota effect. Columns 2 to 5 present a similar scenario for the split samples: The gender quota appears to discourage mothers with a daughter from attempting another birth, but the effects are mostly insignificant; further, the gender quota effects are only little for mothers with a son.

B. Self-reported son preference

To further shed light on the effects of gender quota on son preference, we now turn to exploring the attitudinal variables reported in the Taiwan Social Change Surveys (TSCS), 2001 and 2006, which include a variable indicating self-reported son preference at the individual level. The variable contains the respondents' answers to a survey question asking about their evaluation of "the importance of having at least one son in order to continue the family bloodline". The same question was asked in the 2001 and 2006 waves of TSCS, but the answers were recorded on different scales. They are on a scale of 4 in 2001, ranging over 1 "very important", 2 "important", 3 "unimportant", and 4 "very unimportant"; and on a scale of 7 in 2006, ranging over 1 "extremely important", 2 "very important", 3 "slightly important", 4 "neutral", 5 "slightly unimportant", 6 "very unimportant", and 7 "extremely unimportant". To our end, we construct a dummy variable representing an answer of 1 or 2 in 2001, or any number from 1 to 3 in 2006, to indicate that the respondent attached at least some importance to having a son (= 1) or not (= 0). In timing, the 2001 and 2006 TSCS correspond to the 1998 and 2005 elections, and the exposure to gender quota is measured at the district level. The final sample used for estimation ends up with 3,697 individuals, which are comprised of 1,856 men and 1,841 women; their ages range from 19 to 92, and 46 percent of them stated that having at least a son is at least of some importance. The complete summary statistics of the TSCS sample are presented in Appendix Table A3.

To carry out the estimation, we employ a simplified version of Equation 5 with the dependent variable being replaced for the son preference dummy defined above, and, on the right-hand side, the three dummy variables indicating the first two children's sex composition being replaced by a dummy variable indicating female (F_i). The exact regression is specified as follows:

$$Y_{idce} = \alpha_1 + \beta_1 R_{dce} + \beta_2 R_{dce} \cdot F_i + \beta_3 F_i + \pi P_{dce} + \mathbf{Z}_i \gamma + \rho_c + \sigma_e + \varepsilon_{idce} \tag{6}$$

where R_{dce} is the gender-quota function defined as in Equations 1 and 2, and $\mathbf{Z}_i$ is a vector of individual characteristics, including age (at the time of the survey) and education. Since men are set as the omitted group, the coefficient β_1 captures the effect of gender quota on men's self-reported son preference, whereas $\beta_1 + \beta_2$ captures the effect for women.

Table 8 presents the estimation results, where column 1 suggests that the gender quota does not significantly affect men's evaluation of the importance of having a son, but it significantly diminished women's evaluation. Columns 2 and 3 present the results using a subsample comprised of individuals aged 19 to 45 (women's prime ages for childbearing) and the other subsample comprised of individuals aged 46 or above, respectively. A comparison between columns 2 and 3 shows that the attenuation effect of gender quota on women's son preference was more concentrated on the older women than on the younger ones. This implies that the negative effect of gender quota on the birth propensity of the DD mothers may, to a large extent, reflect the weakening of son preference held by their older generation, particularly their mothers-in-law, who usually exert a strong influence over the younger generation's birth decisions in Taiwan.

C. Educational aspiration and preference

The leadership image delivered by female politicians to the local voters contradicted the gender role stereotypes commonly perceived by the general public. The presence of local female councilors, therefore, might have provided alternative role models of what women can do in society and what careers they can choose to develop. In this section, we investigate this role model effect that can be a potential mechanism the gender quota operated through to attenuate son preference. Using the universal birth records merged with the University Entrance Test records, 2000 to 2003, we start with exploring whether the gender quota had any effect on female students' propensity to attempt college admission by taking up the Advanced Subjects Test (AST), the primary channel to university admission, at the age of 18 when they graduate from high school. We estimate the effects separately for male and female students, using a specification analogous to Equation 6, with the dependent variable being replaced by a dummy variable indicating the AST take-up. On the right-hand side, the election fixed effects σ_e are replaced for student cohort fixed effects, and we additionally control for the student's characteristics. The estimation results are presented in Table 9, where column 1 shows that female students who lived in the electoral districts with a higher gender quota proportion were significantly more likely to take up the AST. The coefficient of interest is 0.0492, which implies that a one-standard-deviation (0.073) increase in the gender quota proportion leads to a 0.36 ppt ($0.0492 \times 0.073 = 0.0036$) increase in the AST take-up rate for female students. This effect translates into a 1.23 percent increase from the mean, given that the average take-up rate for female students was 29.2 percent, as displayed at the bottom of column 1. In contrast, the estimated effect for male students is minimal and insignificant, as suggested by column 2 of the table.

An alternative educational outcome that we can explore refers to students' inclination to select some college majors that are relevant to developing a career in politics. In this regard, law and political science are perceived as the most pertinent in Taiwan, as many politicians, including all the presidents since 2000, hold a law or political science degree. To measure the degree to which a student favors these programs, we exploit the universal college entrance records from 2002 to 2010, which report the entire list of university programs to which each AST participant applied. For each participant, we calculate the proportion of law (or political science) programs out of all programs in the student's application list to indicate the degree of the student's interest in pursuing a career in politics; the higher the proportion is, the more enthusiastic the student is. We present the estimation results in Table 10, where the first estimate in column 1 suggests that the gender quota did not have a significant effect on the proportion of law programs applied by male students (the benchmark group). However, the second estimate shows a significant 'additional' effect of gender quota on female students' applications. The two coefficient estimates (-0.002 and 0.006) imply that a one-standard-deviation increase in the proportion of gender quota increased the proportion of law-related programs listed in female students' applications by 0.025 ppt ($(-0.002 + 0.006) \times 0.0613 = 0.00025$), and the effect is statistically significant as suggested by the *F-test* result shown in the bottom of column 1. Column 2 shows the estimation results for political science programs; similarly, the gender quota does not significantly change the male students' incentive to apply for political science programs, but the effect on female students is much greater and significantly positive.

Next, we explore another measure of a student's preference towards these two disciplines – the ranking of the first law (or political science) program listed in the student's application form, regardless of the number of law programs the student applied to. Naturally, the higher the ranking is, the more serious the student is about selecting law (or political science) as his/her college major. Column 3 of Table 10 shows that the gender quota does not significantly affect the ranking of the first law program listed by the male students, but the additional effect for female students (-3.113) is sizable and significant. Combined, the aggregate effect is around -2.4 (namely, the ranking is lifted by 2.4 positions), suggesting that a one-standard-deviation increase in the proportion of gender quota increased the ranking of the first law program by 0.1474 ($2.4 \times 0.0614 = 0.1474$). Finally, column 4 shows that the proportion of gender quota had no significant effect on the ranking of the first political science program listed by the male students. But, again, the estimated effect for female students ($-0.220 + (-3.336) = -3.556$) is negative, suggesting that the gender quota lifted the ranking of the first political science program listed in their applications.

The results presented in Table 10 provide direct evidence as to the role model effects on female students caused by exposure to local female politicians. Female students exhibited an inclination to follow female politicians' education path by choosing law or political science as their college major. To be sure, we are unable to distinguish whether these effects operated through a change in the student's own preference or a change in her parents' preferences, and it is also likely that both effects were simultaneously at play. In any case, the results shown in both Tables 9 and 10 suggest that the gender quota has gender-specific effects on educational aspiration and preferences towards college majors, lending support to our gender role model hypothesis.

D. Intrahousehold decision making

Next, we examine whether exposure to female political leaders is associated with women gaining higher intrahousehold bargaining power. To this end, we exploit data from the Survey on Social Development Trends (SSDT), 1998 and 2002. All married household members who were 20-year-old or older were asked to specify the primary decision maker(s) for each of the following four household decisions: (1) components of household expenditure; (2) saving and finance; (3) allocation of chores across household members; and (4) parenting (if not childless). For each question, the answer is either (i) myself, (ii) my spouse, (iii) both myself and my spouse, or (iv) other household members. For each question, we construct a dummy variable that is equal to 1 when the respondent considers that the wife is the primary decision maker or that both wife and husband are jointly the decision makers. This way, the dummy variable is designed to indicate the wife's decision power relative to the husband's, regardless of the respondent's gender. The sample is limited to married individuals aged between 16 and 45, which contains 16,384 to 17,358 individuals, depending on the missing answers to each question. The summary statistics of this sample are presented in Appendix Table A6.

Using the newly constructed sample, we repeat the estimation of Equation 6 with the dependent variable being replaced in turn by each of the four decision-making indicators, and, on the right-hand side, the election fixed effects σ_e being replaced by the survey year fixed effects. The results are presented in Table 11. Column 1 shows that the gender quota reduced the likelihood that male respondents reported the wife being the primary decision maker or both wife and husband being the joint decision makers on household expenditure, but the coefficient estimate (-0.085) is not statistically significant. In sharp contrast, the coefficient estimate of the interaction term of the female dummy and gender quota proportion (0.184) is positively significant, suggesting that the gender quota increased the likelihood that female respondents reported themselves (or jointly with their husbands) as the primary decision makers on household expenditure. The corresponding estimates of the interaction terms in columns 2 to 4 suggest that this is also the case for the other three decisions, although the coefficient for parenting is positive but insignificant. Lastly, column 5 presents the estimated effects of the gender quota proportion on the first principal component of the four decision dummy variables. Again, the results show that exposure to gender quota increased women's participation in household decision-making in a general sense. For male respondents, the estimated coefficients are negative across all the other three decisions, suggesting that men tended to report that their wives were less important in the decision-making. However, we are reluctant to draw a conclusion here since most of the estimates are insignificant. Combined, the results in Table 11 imply that the exposure to gender quota empowers women to take part in multiple dimensions of family decisions, a finding supporting the intrahousehold bargaining hypothesis.

VI. Discussion

An alternative channel that would deliver the gender quota effects on fertility decisions may come from the supply side – female politicians improved women's status by enacting public policies in favor of women and reallocating resources to them. After gaining more resources, women change their behaviors and attitudes. Indeed, in the case of reserved village council heads in India, Chattopadhyay and

Duflo (2004) found that village leaders elected through reserved seats invested more in water and roads that were mostly demanded by women. Using legislation data of the French Parliament from 2001 to 2017, Lippmann (2022) found that female legislators were more active than male ones on child and health issues that most concern women. In the case of reserved councilor seats in Taiwan, however, the supply-side argument is unlikely to explain the gender quota effects on fertility and educational outcomes that we found in Section V. First, the public policies and laws that would affect parents' decision to have a third-parity child should be those related to the costs of exercising sex-selective abortion. These policies and laws are all national-level affairs not determined by local councils. Second, no local public policies can explain the gender-specific effects of the gender quota on female students' propensity to select law or political science as their college major since tuition fees and all other educational costs apply equally to male and female students. Third, gender-related issues occupied only a tiny portion of the legislative agenda at the local councils. Using the entire interpellation documents submitted by the Taipei City councilors to the Council from January 2006 to June 2022, we calculate the proportion of these documents that contain at least one keyword relating to gender-specific issues, such as gender equality, sexual crimes, or maternal health. The overall proportion is only 1.5 percent, although the proportion is higher among the documents submitted by female councilors (2.3 percent) than those offered by males (1.1 percent). Details of the statistics are presented in Appendix Table A7. Combined, we believe that the attenuation effect on son preference of the gender quota found in Section V is unlikely to be driven by changes on the supply side.

To be sure, although the patterns of the gender quota effect on girls' educational outcomes shown in Section V are consistent with the role model hypothesis, we are unable to completely rule out other potential mechanisms that might deliver these effects. One such mechanism operates through parents' rational decisions on the investments in daughters' human capital: The image of successful female politicians revamps parents' perceptions of daughters' future achievements relative to those of sons; this incentivizes parents, both mothers and fathers, to increase daughters' human capital, leading to a higher proportion of female students attempting college admission. To examine this 'human-capital-investment' hypothesis, we shift the gear to investigate whether the gender quota has any sex-specific impact on another key aspect of human capital – health. The rationale behind this examination is that if the gender quota motivated parents to invest more in their daughters' human capital, an increase in gender quota would have improved daughters' health relative to sons.

To this end, we exploit Taiwan's Mortality Registries, 2000 to 2008, which report universal death records, including neonatal deaths. We implement the examination by estimating Equation 6, replacing the dependent variable with a dummy variable indicating whether the newborn child died within the first three years of life. The sample used for the estimation is comprised of all newborn children from 2000 to 2006. The estimation results are reported in Table 12, where column 1 shows that the coefficient estimate of gender quota proportion is insignificant at -1.33, suggesting that the exposure to gender quota does not much change the neonatal mortality rate for boys (the omitted group). The coefficient estimate of the interaction term of gender quota proportion and the female dummy is also insignificant at 1.59, implying a limited additional effect for girls. Both estimates change only slightly after we further control for birth parity fixed effects, as shown by the results in column 2. These findings do not support the human-capital-investment hypothesis.

VII. Conclusions

It is puzzling that, given the shared traditional culture and experience of post-WWII economic expansion, the gender gaps in political representation, economic opportunities, and social participation are significantly smaller in Taiwan than in Japan or South Korea. Using multiple administrative data, we evaluate the role of a unique design of the political system in Taiwan – the reserved seats for women at the county/city councils – in promoting women's political representation and altering gender attitudes and behaviors. Our identification strategy exploits the virtually exogenous cross-district variation in the proportion of female councilors created by the uneven distribution of reserved seats.

We have two sets of findings. The first set of findings provides evidence that the gender quota effectively encouraged women to run for the council elections and win more seats, and the winners gained a promising opportunity to move up to higher public offices. The second set of findings suggests that the gender quota impacted gender attitudes and decisions. Specifically, we find that the gender quota reduced fertility for parents with two daughters, but there was no similar effect on parents with two children of any other sex composition. We also find that the gender quota profoundly impacted females: It attenuated women's self-reported son preference, encouraged female students to attempt college admission and choose law or political science as their college major, and enhanced wives' decision power in households. All these findings support the hypothesis that female political leaders played as role models who inspired women to accept counter-stereotypical gender norms.

To what extent can the gender quota account for the differences in female status in Taiwan versus Japan and South Korea? Although the availability of our data is not sufficient enough to render a precise evaluation of the roles played by the gender quota, our description in Section II and analyses in Section IV and V suggest that its contributions are non-trivial to raising female political representation and shaping the gender attitudes held by women. The evidence provided in this paper connotes an important policy implication for countries, Asian or not, that are striving to build a more gender-egalitarian society.

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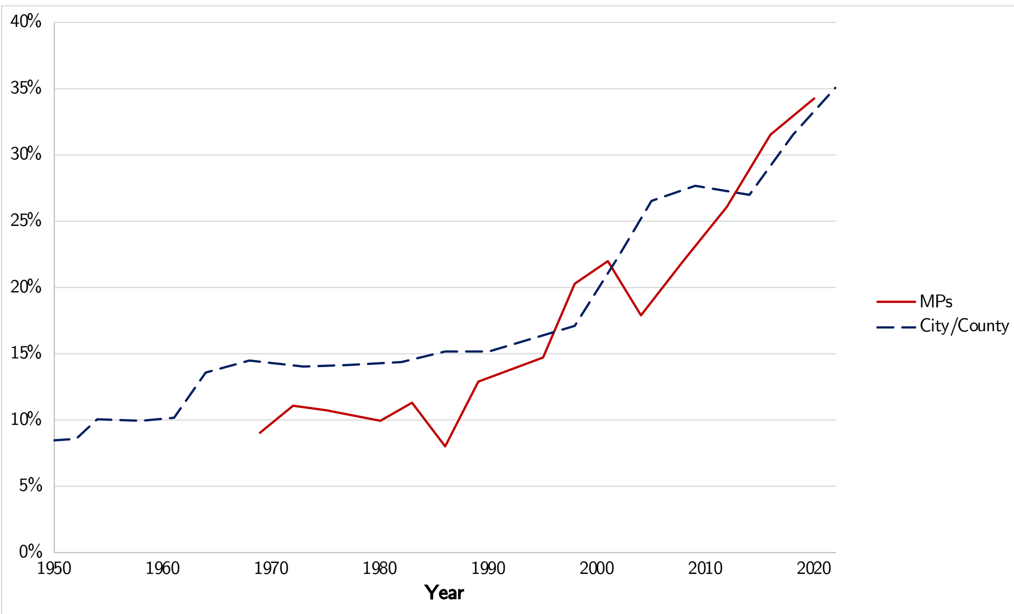
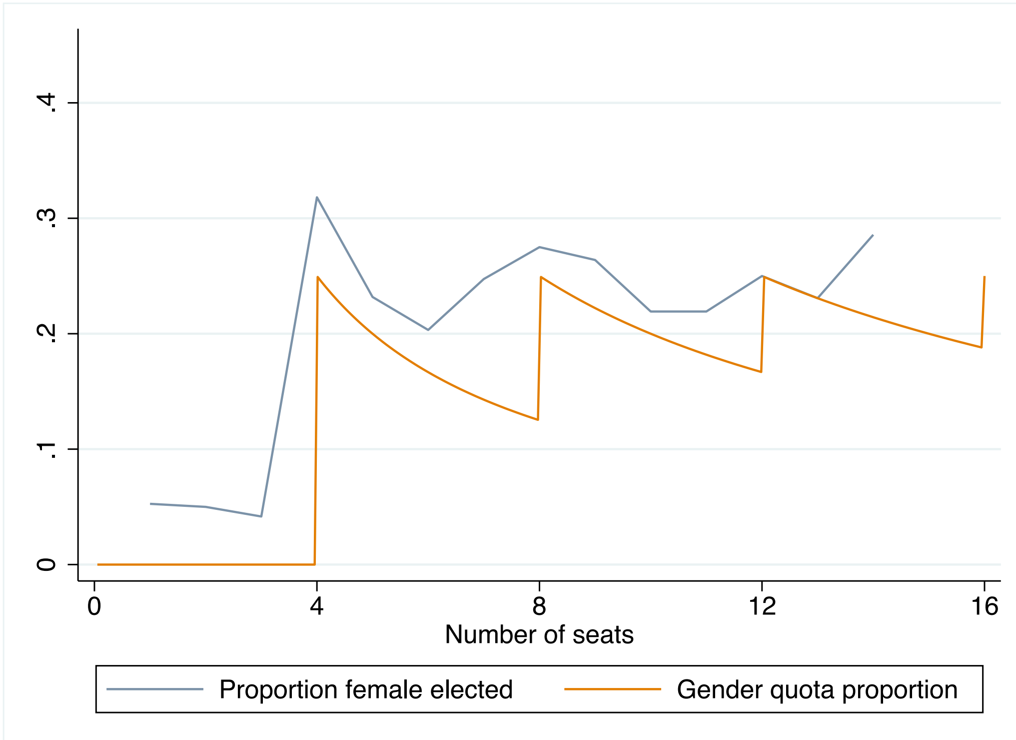
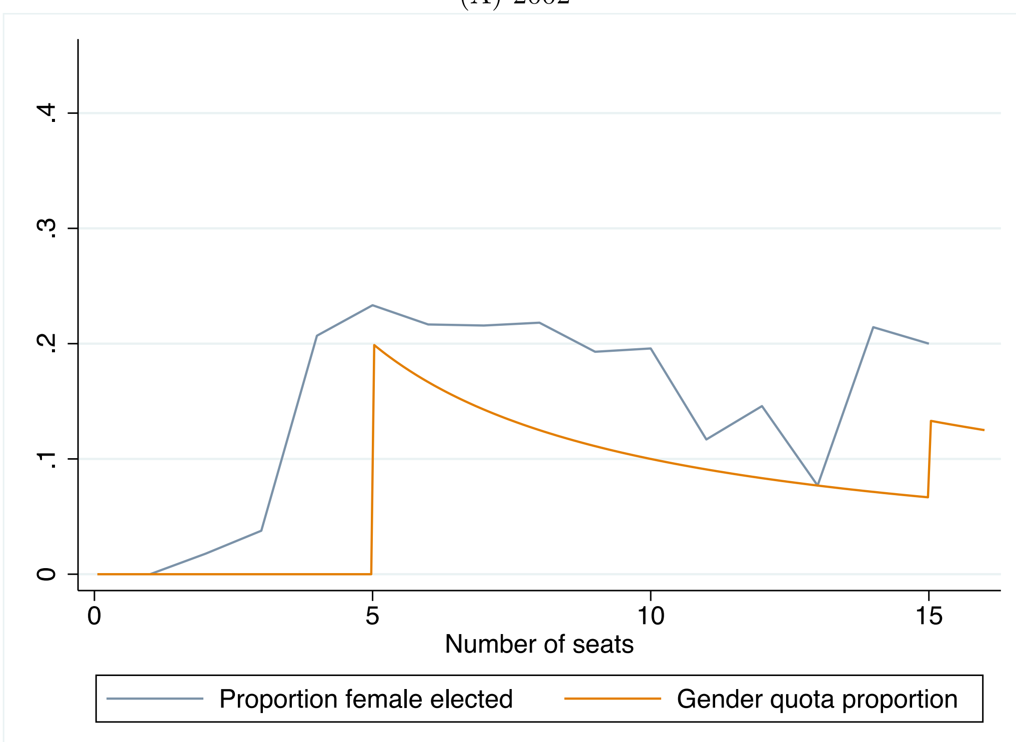


Figure 1. Proportion of Female Representatives in Taiwan

Notes: The dashed blue line indicates the proportion of female councilors among all local councilors in Taiwan. The solid red line indicates the proportion of female legislators (members of parliament, MPs) in the national parliament (Legislative Yuan). The first election of MPs in Taiwan was held in 1969, therefore no observations are presented before 1969. Data before 1989 are collected from Chou, Clark and Clark (1990), and data after 1990 are collected from the election database of the Central Election Committee.



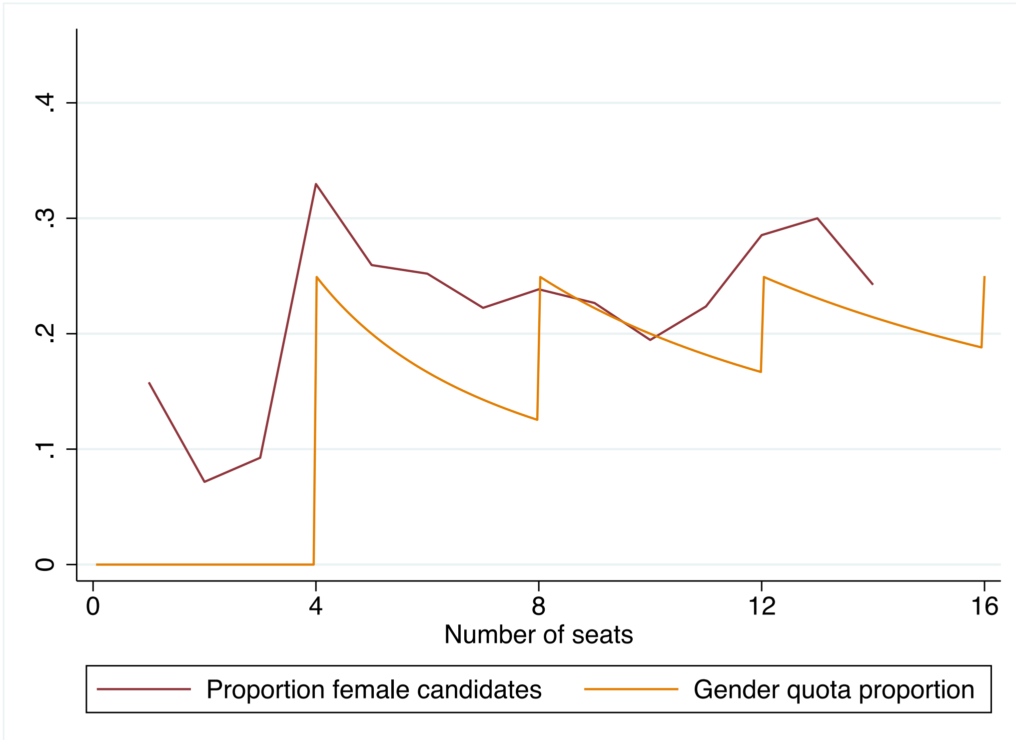
(A) 2002



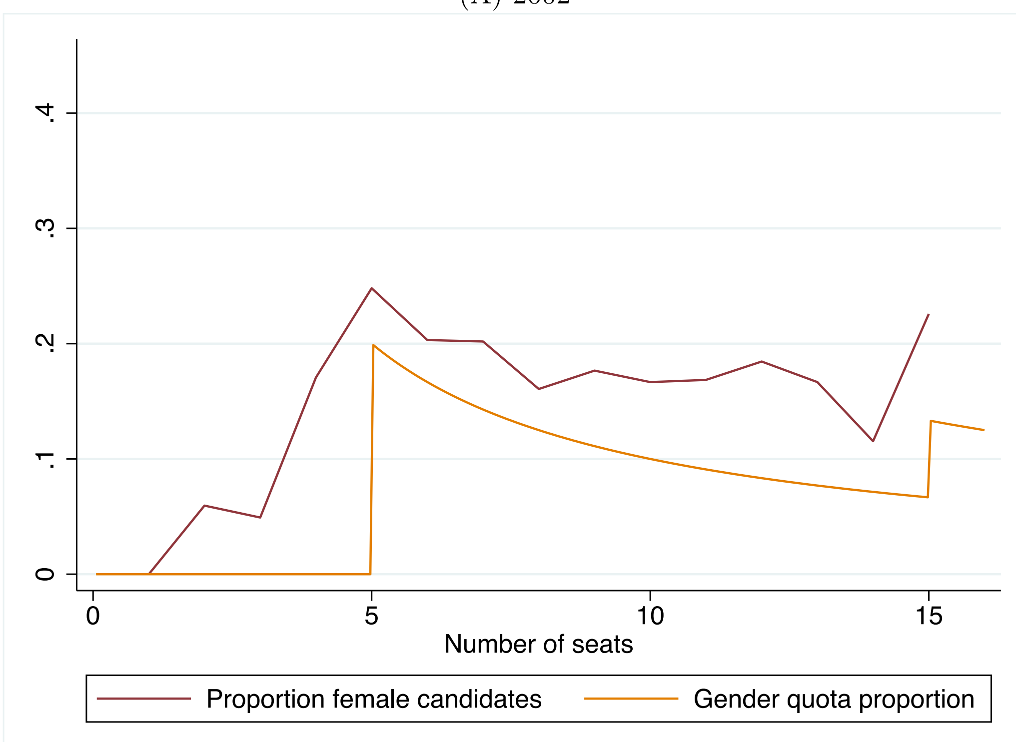
(B) 1998

Figure 2. Gender quota rule and proportion of actual seats won by women

Notes: The horizontal axis corresponds to the number of seats in an electoral district. In both figures, the blue line indicates the average proportion of seats won by females, and the orange line indicates the proportion of gender quota (the gender quota function).



(A) 2002



(B) 1998

Figure 3. Gender quota rule and proportion of female candidates running for elections

Notes: The horizontal axis corresponds to the number of seats in an electoral district. In both figures, the red line indicates the average proportion of candidates that are female, and the orange line indicates the proportion of gender quota (the gender quota function).

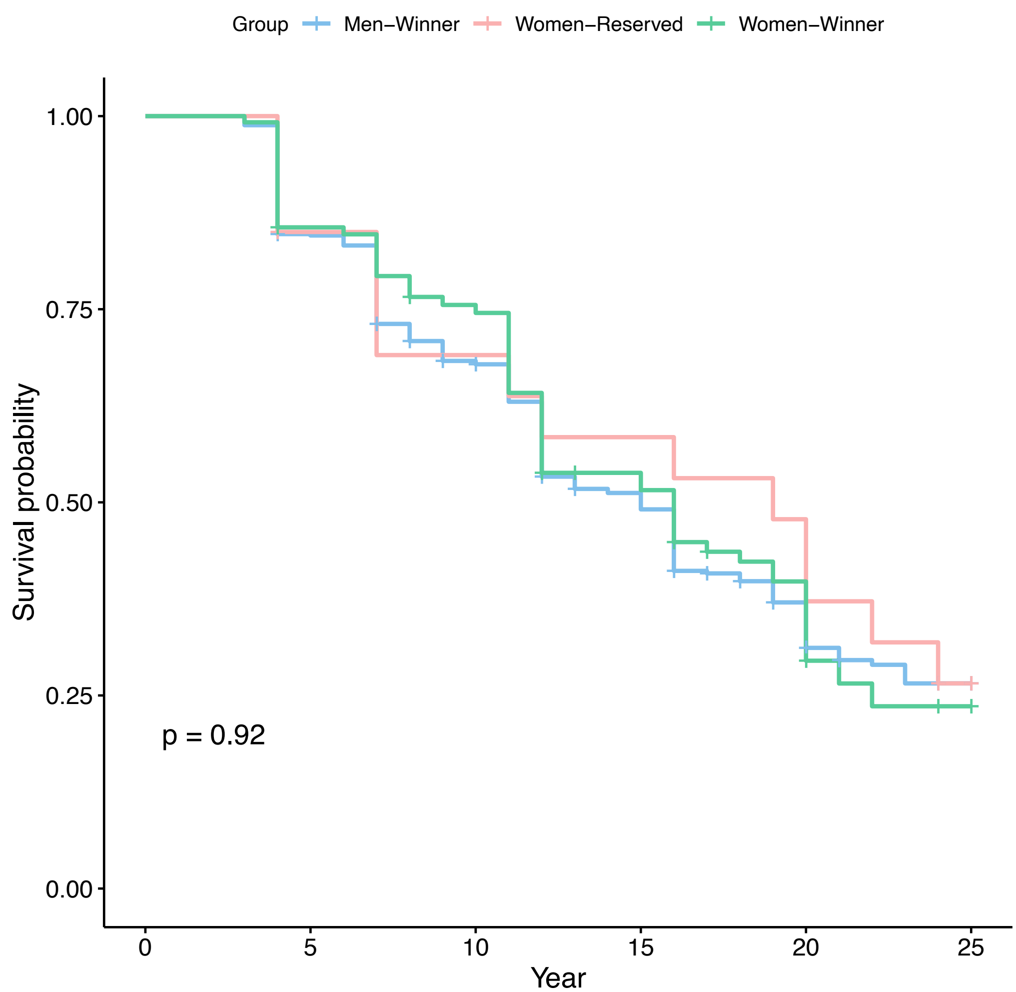


FIGURE 4. KAPLAN-MEIER SURVIVAL CURVES FOR THREE DIFFERENT TYPES OF WINNERS OF THE 1998 ELECTION

Notes: The sample is comprised of all the 826 winners (including reserved-seat winners) in the 1998 council election who did not run for the national legislator election in the same year. Each individual's political position is tracked from 1998 to 2023 using the election database of the Central Election Committee. The p-value of the log-rank test for the equality of survival probability between three types of winners is 0.92, implying that the differences in survival probability between the reserved-seat winners and the other two types of winners are statistically insignificant.

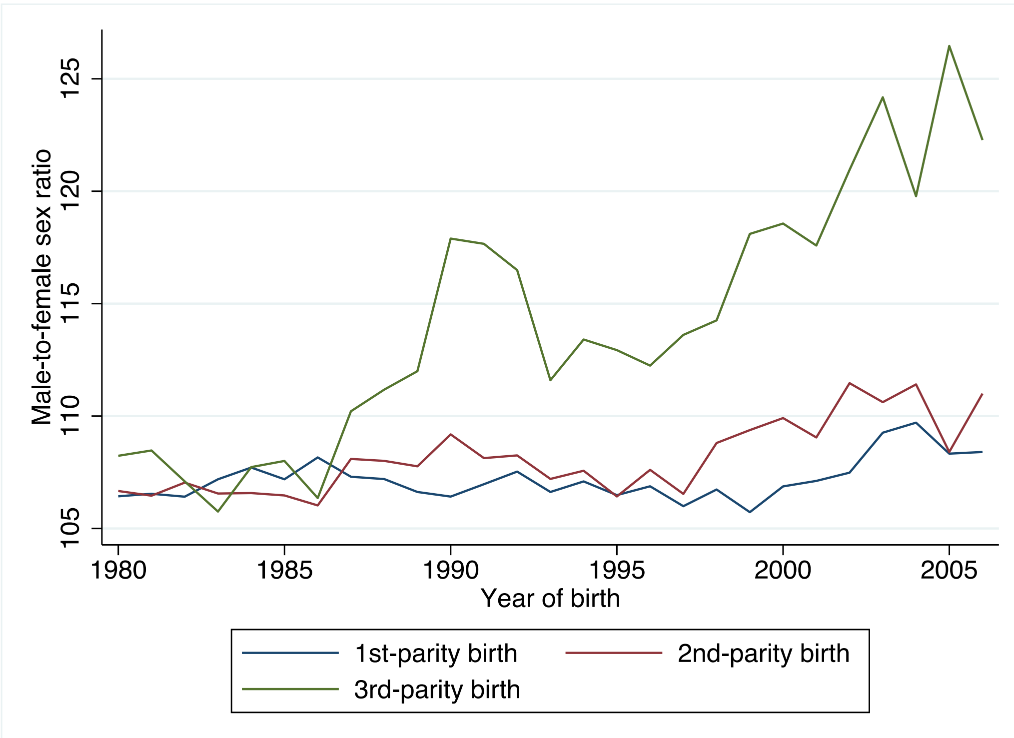


Figure 5. Male-to-female sex ratio at birth by birth order in Taiwan

Notes: The male-to-female sex ratio is defined as (male births / female births) $\times$ 100.

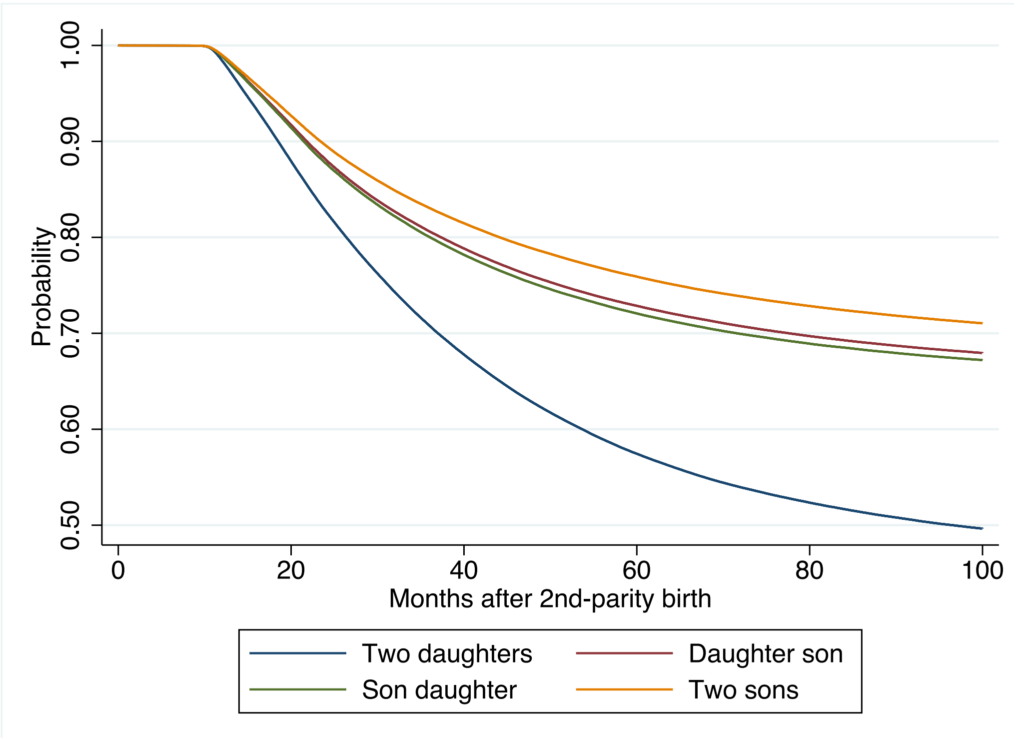


Figure 6. Kaplan-Meier survival curves for parents with two children to give the third-parity birth

Notes: The p-value of the log-rank test for the equality of survival probability between four types of mothers is 0.000.

Table 1—Examining the determinants of gender quota

	(1)	(2)
	Gender quota proportion in	
	2002	2005
Vote share won by female candidates in previous election	−0.007	0.011
	(0.109)	(0.146)
Proportion of female candidates in previous election	0.370***	0.239**
	(0.140)	(0.114)
Population size (10 million)	2.346***	3.070***
	(0.548)	(0.648)
Mean dep. var.	0.147	0.144
Obs.	159	160
Adj. R^2	0.310	0.320

Notes: This table reports the regression estimate of gender quota proportion in the 2002 and 2005 elections, predicted by the vote share gained by female candidates in the previous councilor election. The unit of observation is electoral district. The sample is comprised of all electoral districts of council election in Taiwan. Robust standard errors in parentheses. * $p<0.1$, ** $p<0.05$, *** $p<0.01$

Table 2—The effects of the gender quota on proportion of elected female councilors and proportion of female candidates running for council elections

	(1)	(2)
	Dependent variable:	
	Proportion female elected	Proportion female candidates
Gender quota proportion	1.01***	0.753***
	(0.080)	(0.077)
Population size (10 million)	−0.269	−0.605
	(0.69)	(0.60)
Mean dep. var.	0.186	0.184
Obs.	609	590
Adj. R^2	0.387	0.354

Notes: The sample is comprised of electoral districts in 1994, 1998, 2002, and 2005 council elections. Dependent variables in columns 1 and 2 are the proportion of female elected councilors and the proportion of female candidates, respectively. The observation number is smaller in column 2 due to missing values. Robust standard errors in parentheses. * p<0.1, ** p<0.05, *** p<0.01

Table 3—Winners and losers of the 1998 council election by gender

	Winner	Winner (reserved)	Loser	Loser (replaced)	
Women	125	20	147	–	292
Men	693	–	798	20	1,511
Total	818	20	945	20	1,803

Notes: This frequency table reports the number of winning and losing candidates in the 1998 council election by gender. "Winner (reserved)" indicates those female winning candidates who would otherwise lose the election if the quota policy had not existed. "Loser (replaced)" indicates those male candidates who would win the seats but eventually be replaced by the reserved winners.

TABLE 4—COMPARING THE POLITICAL CAREER DEVELOPMENT OF THE RESERVED WINNERS AND OTHER WINNERS OF THE 1998 COUNCIL ELECTION

	(1)	(2)
Woman (other winner)	0.977	1.024
	(0.125)	(0.133)
Woman (reserved winner)	0.903	1.010
	(0.247)	(0.279)
Age in 1998		1.020***
		(0.007)
Party membership: KMT		0.947
		(0.104)
Party membership: DPP		0.885
		(0.136)
Obs.	826	826
Nagelkerke pseudo R^2	0.00	0.01

Notes: This table reports Cox proportional hazards model estimates. Reported coefficients are hazard ratios. Standard errors of the coefficients are in parentheses. Hazard is defined as an individual exiting any elected positions (including local councilor, national legislator, and mayor). The sample is comprised of all the 826 winners (including reserved-seat winners) in the 1998 councilor election who did not run for the national legislator election in the same year. Each individual's political position is tracked from 1998 to 2023 using the election database of the Central Election Committee. The male winners are the omitted baseline group. "Woman (other winner)" is a dummy variable indicating those female candidates whose votes would qualify them as winners in the absence of the quota. "Woman (reserved winner)" is a dummy variable indicating those female winning candidates who would otherwise lose if the quota were absent. *** $p<0.01$.

Table 5—Summary statistics for the four different types of mothers

	Two daughters (N=685,362)		Daughter son (N=956,580)		Son daughter (N=912,783)		Two sons (N=1,033,753)	
Variables	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Giving 3rd-parity birth	0.109	0.312	0.051	0.220	0.051	0.220	0.047	0.211
Gender quota proportion	0.144	0.072	0.144	0.072	0.144	0.072	0.143	0.072
Proportion female elected	0.214	0.125	0.213	0.125	0.213	0.125	0.212	0.125
Proportion female candidates	0.202	0.099	0.202	0.099	0.201	0.099	0.201	0.099
Population size (thousand)	232.447	155.494	229.306	154.895	228.549	154.503	226.785	153.674
Age of mother in election year	35.568	5.319	36.247	5.179	36.247	5.174	36.379	5.155
Age of father in election year	38.546	5.486	39.236	5.356	39.220	5.350	39.352	5.331
Mother with high school diploma	0.510	0.500	0.484	0.500	0.482	0.500	0.472	0.499
Father with high school diploma	0.545	0.498	0.522	0.500	0.521	0.500	0.514	0.500
Urban	0.227	0.419	0.218	0.413	0.217	0.412	0.212	0.409

Notes: This table reports the summary statistics of the four types of mothers in our sample. The unit of observation is mother-by-election. The sample is comprised of mothers aged from 16 to 45 and their children's fathers aged from 16 to 50 in the year of election and have had a 2nd-parity birth from 1978 to 2006. "Urban" is a dummy variable indicating mothers having 2nd birth in Taipei City, Taoyuan City, Taichung City, Tainan City, and Kaohsiung City, which are the urbanized counties in Taiwan in our sample period. Due to missing data on the proportion of female candidates in the 1994 election, the observations for variable "proportion female candidates" are 672,713, 939,084, 895,485, and 1,013,674 for each type of mothers respectively.

Table 6—Estimating the effects of gender quota on son preference

	(1)	(2)	(3)	(4)	(5)
	Dependent variable: Giving 3rd-parity birth				
	Full Sample	Subgroups			
		High School	Non-HS	Urban	Non-urban
($\hat{\beta}_1$) Gender quota proportion	−.126***	−.141***	−.117***	−.204***	−.115***
	(0.015)	(0.018)	(0.016)	(0.037)	(0.017)
($\hat{\beta}_2$) Two sons × Gender quota proportion	.174***	.202***	.158***	.272***	.162***
	(0.017)	(0.020)	(0.017)	(0.042)	(0.018)
($\hat{\beta}_3$) Son daughter × Gender quota proportion	.157***	.183***	.146***	.246***	.147***
	(0.015)	(0.019)	(0.015)	(0.038)	(0.016)
($\hat{\beta}_4$) Daughter son × Gender quota proportion	.152***	.166***	.15***	.22***	.144***
	(0.015)	(0.019)	(0.015)	(0.037)	(0.016)
Two sons	−.079***	−.0923***	−.0674***	−.0817***	−.0806***
	(0.0028)	(0.0035)	(0.0029)	(0.0075)	(0.0030)
Son daughter	−.073***	−.0864***	−.0614***	−.0769***	−.074***
	(0.0024)	(0.0031)	(0.0025)	(0.0067)	(0.0026)
Daughter son	−.0723***	−.0827***	−.063***	−.0724***	−.0738***
	(0.0025)	(0.0031)	(0.0025)	(0.0065)	(0.0027)
Mean dep. var.	0.0609	0.0636	0.0583	0.0445	0.0654
Obs.	3,588,478	1,740,241	1,848,237	780,949	2,807,529
Adj. R^2	0.111	0.109	0.117	0.0839	0.116
F-test on marginal effect					
p-value of $H_0 : \hat{\beta}_1 + \hat{\beta}_2 = 0$	0.000	0.000	0.000	0.000	0.000
p-value of $H_0 : \hat{\beta}_1 + \hat{\beta}_3 = 0$	0.001	0.000	0.003	0.031	0.002
p-value of $H_0 : \hat{\beta}_1 + \hat{\beta}_4 = 0$	0.004	0.017	0.001	0.390	0.006

Notes: Mothers having two daughters (DD mothers) are omitted group. "Urban" group is defined as mothers having the 2nd-parity birth in Taipei City, Taoyuan City, Taichung City, Tainan City, and Kaohsiung City, which are the urbanized counties in Taiwan in our sample period. Clustered robust standard errors in parentheses. Standard errors are clustered at township-election level. * p<0.1, ** p<0.05, *** p<0.01.

Table 7—Estimating the effects of gender quota on the sex of third child

	(1)	(2)	(3)	(4)	(5)
	Dependent variable: The 3rd-party child being a male				
	Full Sample	Subgroups			
		High School	Non-HS	Urban	Non-urban
($\hat{\beta}_1$) Gender quota proportion	.0984***	.064	.0915**	.275**	.0876***
	(0.032)	(0.040)	(0.047)	(0.11)	(0.033)
($\hat{\beta}_2$) Two sons × Gender quota proportion	−.0778*	−.138**	.0347	−.07	−.0756
	(0.046)	(0.061)	(0.064)	(0.14)	(0.049)
($\hat{\beta}_3$) Son daughter × Gender quota proportion	−.0785*	.0244	−.118*	−.174	−.0654
	(0.045)	(0.061)	(0.065)	(0.16)	(0.046)
($\hat{\beta}_4$) Daughter son × Gender quota proportion	−.107**	−.0836	−.0768	−.124	−.102**
	(0.043)	(0.059)	(0.063)	(0.14)	(0.045)
Two sons	−.0646***	−.0757***	−.0567***	−.0726***	−.0634***
	(0.0066)	(0.0092)	(0.0087)	(0.022)	(0.0068)
Son daughter	−.0522***	−.0861***	−.0252***	−.057***	−.0505***
	(0.0063)	(0.0088)	(0.0089)	(0.021)	(0.0066)
Daughter son	−.0478***	−.0683***	−.0318***	−.0535***	−.0467***
	(0.0063)	(0.0088)	(0.0088)	(0.020)	(0.0066)
Mean dep. var.	0.539	0.547	0.530	0.544	0.537
Obs.	218,367	110,689	107,678	34,774	183,593
Adj. R^2	0.00473	0.00765	0.00200	0.00648	0.00448
F-test on marginal effect					
p-value of $H_0 : \hat{\beta}_1 + \hat{\beta}_2 = 0$	0.552	0.138	0.007	0.172	0.738
p-value of $H_0 : \hat{\beta}_1 + \hat{\beta}_3 = 0$	0.567	0.081	0.601	0.447	0.537
p-value of $H_0 : \hat{\beta}_1 + \hat{\beta}_4 = 0$	0.792	0.699	0.749	0.261	0.666

Notes: The sample is comprised of mothers having 3rd-parity birth. Mothers having two daughters as their first two children (DD mothers) are the omitted group. "Urban" group is defined as mothers having the 2nd-parity birth at Taipei City, Taoyuan City, Taichung City, Tainan City, and Kaohsiung City, which are the urbanized counties in Taiwan in our sample period. Clustered robust standard errors in parentheses. Standard errors are clustered at township-election level. * p<0.1, ** p<0.05, *** p<0.01

Table 8—Estimating the effects of gender quota on self-reported son preference

	(1) All Age	(2) Age 19-45	(3) Age > 45
Gender quota proportion	−0.0813	0.230	−0.526
	(.264)	(.313)	(.318)
Woman × Gender quota proportion	−0.494**	−0.384	−0.626*
	(.2)	(.288)	(.331)
Woman	−0.0671**	−0.0993**	−0.00758
	(.0305)	(.0463)	(.0505)
Mean dep. var.	0.46	0.356	0.594
Observations	3,697	2,077	1,620
Adj. R^2	0.131	0.0567	0.123

Notes: The dependent variable is a dummy variable indicating self-reported son preference constructed using data from the 2001 and 2006 waves of the Taiwan Social Change Survey. The variable refers to the respondents' answers to a survey question asking about their own evaluation of "the importance of having at least one son in order to continue the family bloodline." Male respondents are the omitted group. The sample in column 1 is comprised of all respondents. The samples in columns 2 and 3 are comprised of respondents aged from 19 to 45 and above 45, respectively. Standard errors in parentheses are robust and clustered at township-election level. * p<0.1, ** p<0.05, *** p<0.01

Table 9—Estimating the effects of gender quota on taking up college entrance test

	(1) Female	(2) Male
Gender quota proportion	0.0492**	0.0177
	(0.020)	(0.020)
Mean dep. var.	0.292	0.274
Observations	532,046	569,388
Adj. R^2	0.119	0.114

Notes: The dependent variable is a dummy variable indicating taking up the Advanced Subjects Test (=1) or not (=0) at the age of 18. The unit of observation is individual student. The samples in columns 1 and 2 are comprised of female and male students, respectively. Standard errors in parentheses are robust and clustered at township-election level. * p<0.1, ** p<0.05, *** p<0.01

Table 10—Estimating the effects of gender quota on preferences toward college majors

	(1)	(2)	(3)	(4)
	Proportion of applied programs in		Ranking of the 1st program in	
	Law	Political science	Law	Political science
($\hat{\beta}_1$) Gender quota proportion	−0.002	0.001	0.712	−0.220
	(0.0022)	(0.00081)	(1.42)	(2.21)
($\hat{\beta}_2$) Woman × Gender quota proportion	0.006**	0.002**	−3.113**	−3.336
	(0.0028)	(0.00079)	(1.47)	(2.26)
Woman	0.005***	0.001***	1.580***	1.139***
	(0.00052)	(0.00015)	(0.28)	(0.43)
Mean dep. var.	0.025	0.0062	22.3	24.7
Observations	735,312	735,312	219,699	124,407
Adj. R^2	0.0045	0.0068	0.022	0.013
F-test on marginal effect				
p-value of $H_0 : \hat{\beta}_1 + \hat{\beta}_2 = 0$	0.034	0.0020	0.099	0.092

Notes: The sample is constructed using AST participants' application records, which contains all AST participants from 2002 to 2010. The dependent variables in columns 1 and 2 are respectively the proportion of the applied law and political science programs among all programs that a student applied to. The dependent variables in columns 3 and 4 are respectively the ranking position of the first law and political science program in the application list. Clustered robust standard errors in parentheses. Standard errors are clustered at township-election level. * p<0.1, ** p<0.05, *** p<0.01

Table 11—Estimating the effects of gender quota on household decisions making

	(1)	(2)	(3)	(4)	(5)
	Dependent variable: The wife is the decision maker or both wife and husband are decision makers				
	Expenditures	Saving & finance	Allocation of chores	Parenting	1st principal component
Gender quota proportion	−0.085	−0.123	−0.135**	−0.042	−0.574*
	(0.112)	(0.086)	(0.058)	(0.083)	(0.345)
Female × Gender quota proportion	0.184***	0.234***	0.078*	0.030	0.669***
	(0.069)	(0.065)	(0.047)	(0.057)	(0.258)
Female	0.016	0.012	0.007	0.023**	0.079*
	(0.012)	(0.011)	(0.009)	(0.009)	(0.043)
Mean dep. var.	0.833	0.895	0.938	0.882	0.158
Observations	17,358	17,013	17,358	16,384	16,039
Adj. R^2	0.015	0.042	0.00401	0.0556	0.0172

Notes: The sample is comprised of married individuals aged from 16 to 45 in the Survey on Social Development Trends, 1998 and 2002. The dependent variable is a dummy variable indicating the respondent's answer being "it is the wife being the decision maker or both wife and husband being the decision makers.". "Female" is a dummy variable indicating the respondent is a female. Because we remove respondents who answered "not applicable" or "other", the number of observations varies across columns. Column 5 reports the estimated effects of the gender quota proportion on the first principal component of the four decision variables. The first principal component captures 48% of all variance. Observations are weighted to be nationally representative, with the weights being provided by the survey data. Clustered robust standard errors in parentheses. Standard errors are clustered at township-election level.
* $p<0.1$, ** $p<0.05$, *** $p<0.01$.

Table 12—Estimating the effects of gender quota on neonatal mortality

	(1)	(2)
	Death per 1,000 newborns	
Gender quota proportion	−1.33	−1.22
	(0.85)	(0.84)
Girl × Gender quota proportion	1.59	1.59
	(1.00)	(1.00)
Girl	−0.726***	−0.711***
	(0.19)	(0.19)
2nd-parity birth		1.08***
		(0.083)
3rd-parity birth		2.08***
		(0.17)
Mean dep. var.	3.07	3.07
Observations	2,590,558	2,590,558
Adj. R^2	0.000537	0.000673

Notes: The sample is comprised of all the newborn children from 2000 to 2006. The dependent variable is a dummy variable indicating death during the first three years of life ($= 1 \times 1,000$). Standard errors in parentheses are robust and clustered at township-election level. * p<0.1, ** p<0.05, *** p<0.01

Appendix

A1. Definition of law- and political-science-related college programs

To identify law-related (or political-science-related) college programs, we use the official categories of disciplines published by the Ministry of Education, which follows the International Standard Classification of Education – Fields of education and training (ISCED-F 2013). The field codes and corresponding programs are listed in Table A1. Notice that we exclude programs in the field of public administration from the political science category because most students in this field are interested in pursuing a career as a public servant rather than a career as a politician.

Table A1—List of law and political science programs

Law-related programs
ISCED-F 2013 field code: 0421 Law
Department of Financial and Economic Law
Department of Government and Law
Department of Judicial Studies
Department of Law
Department of Law (Division of Economic and Financial Law)
Department of Law (Division of Financial & Economic Law)
Department of Law (Division of Judicial Studies)
Department of Law (Division of Law)
Department of Law (Financial and Economic Law Program)
Department of Law (Legal Institutions Program)
Bachelor Programme of Extension Education (School of Law)
Political-science-related programs
ISCED-F 2013 field code: 0312 Political sciences and civics
Department of Global Politics and Economics
Department of Global Politics and Economics (English-taught Program)
Department of Government and Law
Department of International and Mainland China Affairs
Department of Political Economy
Department of Political Science
Department of Political Science, sub-division: international relations
Department of Political Science, sub-division: political theory
Department of Political Science, sub-division: public administration

A2. Additional Tables

Table A2—Estimating the effects of gender quota on son preference (2nd parity)

	(1)	(2)	(3)	(4)	(5)
	Dependent variable: Giving 2nd-parity birth				
	Full Sample	Subgroups			
		High School	Non-HS	Urban	Non-urban
($\hat{\beta}_1$) Gender quota proportion	−.00716	−.00353	.00337	−.0254	−.00504
	(0.015)	(0.018)	(0.016)	(0.038)	(0.016)
($\hat{\beta}_2$) 1st-parity is a daughter × Gender quota proportion	−.0118**	−.00178	−.0332***	−.0162	−.0113*
	(0.0058)	(0.0093)	(0.0058)	(0.018)	(0.0061)
1st-parity is a daughter	.0158***	.0192***	.0134***	.016***	.0159***
	(0.00093)	(0.0015)	(0.0010)	(0.0029)	(0.00098)
Mean Dep. Var.	0.175	0.234	0.118	0.156	0.180
Obs.	3,837,237	1,882,788	1,954,449	884,685	2,952,552
Adj. R^2	0.241	0.251	0.218	0.224	0.245
F-test on marginal effect					
p-value of $H_0 : \hat{\beta}_1 + \hat{\beta}_2 = 0$	0.202	0.773	0.070	0.304	0.319

Notes: Mothers having a son as their 1st-parity child are omitted group. "Urban" group is defined as mothers having the 2nd-parity birth in Taipei City, Taoyuan City, Taichung City, Tainan City, and Kaohsiung City, which are the urbanized counties in Taiwan in our sample period. Clustered robust standard errors in parentheses. Standard errors are clustered at township-election level. * p<0.1, ** p<0.05, *** p<0.01.

Table A3—Summary statistics of the Taiwan Social Change Survey sample

Variable	Obs.	Mean	Std. Dev.
Self-reported son preference	3,697	0.46	0.498
Gender quota proportion	3,697	0.145	0.0712
Woman	3,697	0.498	0.5
Population size (thousand)	3,697	234.826	151.543
Years of education	3,697	10.6	4.68
Age	3,697	45	16.3

Notes: "Self-reported son preference" is a dummy variable constructed using respondents' answers to a survey question asking about their own evaluation of "the importance of having at least one son in order to continue the family bloodline". The same question was asked in the 2001 and 2006 waves of TSCS, but the answers are recorded on different scales. They are on a scale of 4 in 2001, ranging over 1 "very important", 2 "important", 3 "unimportant", and 4 "very unimportant"; and on a scale of 7 in 2006, ranging over 1 "extremely important", 2 "very important", 3 "slightly important", 4 "neutral", 5 "slightly unimportant", 6 "very unimportant", and 7 "extremely unimportant". To our end, we construct a dummy variable representing an answer of 1 or 2 in 2001, or any number from 1 to 3 in 2006, to indicate that the respondent attached at least some importance to having a son (= 1) or not (= 0).

Table A4—Summary statistics for students at the school age of 18

	Female (N=532,046)		Male (N=569,388)	
Variable	Mean	Std. Dev.	Mean	Std. Dev.
Taking up AST	0.292	0.454	0.274	0.446
Gender quota proportion	0.141	0.073	0.141	0.073
Population size (thousand)	221.974	153.127	221.831	152.947
Age of student	18.353	0.478	18.351	0.477
Age of mother	43.267	3.621	43.280	3.631
Age of father	46.594	4.433	46.601	4.434
Mother with high school diploma	0.307	0.461	0.309	0.462
Father with high school diploma	0.396	0.489	0.397	0.489

Notes: The sample is comprised of students at the school age of 18 from 2000 to 2003. "Taking up AST" is an indicator variable that takes a value 1 if a student sign-up for AST or 0 otherwise.

Table A5—Summary statistics for students taking AST from 2002 to 2010

Variable	Obs.	Mean	Std. Dev.
Proportion of application: Law	735,312	0.0246	0.0683
Proportion of application: Political science	735,312	0.00621	0.0199
Ranking of 1st program in law	219,699	22.3	21.4
Ranking of 1st program in political science	124,407	24.7	21.4
Gender quota proportion	735,312	0.183	0.0613
Population size (thousand)	735,312	273	166

Table A6—Summary statistics of the Survey on Social Development Trends sample

Variable	Obs.	Mean	Std. Dev.
Wife or joint deciding household expenditures	17,358	.833	.373
Wife or joint deciding saving & finance	17,013	.895	.306
Wife or joint deciding allocation of chores	17,358	.938	.24
Wife or joint deciding parenting	16,384	.882	.323
Gender quota proportion	17,358	.139	.0725
Female	17,358	.538	.499
Population size (thousand)	17,358	245.130	159.803
Years of education	17,358	11.6	3.16
Age	17,358	36.2	5.81

Notes: Observations are weighted to be nationally representative, with the weights being provided by the survey data.

Table A7—Frequency of Gender-related Interpellations in Taipei City Council (2006 - 2022)

	(1) Women	(2) Men	(3) All councilors
Gender-related interpellations	154	165	319
Other interpellations	6,573	14,951	21,524
Total	6,727	15,116	21,843
Proportion of gender-related interpellations	0.023	0.010	0.015

Notes: This table reports the number of gender-related interpellations raised by Taipei City councilors from Jan. 2006 to Jun. 2022. The sample contains only written interpellations. An interpellation is defined as “gender-related” if the title contains any of following keywords: “gender”, “women”, “gender equality”, “men and women”, “sexual harassment”, “sexual violence”, “sexual predator”, “women’s restroom”, “female students”, “voyeurism”, “policewomen”, “expectant others”, “childbirth”, “infant care”, “childminder”, “pregnancy”, “uterus”, “breast cancer”, “breast milk”, or “breastfeeding”.