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ABSTRACT

Only-Child Matching Penalty in the Marriage Market*

This study explores the marriage matching of only-child individuals and the related outcomes. Specifically, we analyze two aspects: First, we investigate the marriage patterns of only children, examining whether people choose mates in a positive or negative assortative manner regarding only-child status. We find that, along with being more likely to remain single, only children are more likely to marry another only child. Second, we measure the matching premium or penalty using the difference in partners' socioeconomic status, measured by years of schooling, between only-child and non–only-child individuals. Our estimates show that among women who marry an only-child husband, only children are penalized, as their partners' educational attainment is 0.63 years lower. Finally, we discuss the potential sources of this penalty along with our empirical findings.

JEL Classification: J11, J12, J16

Keywords: marriage matching, only children, gender, machine learning

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1 Introduction

Globally, there has been an increase in only child families in many developed countries. For example, in the Americas and Asia, the percentage of one-child families among those with children has nearly doubled in recent decades.[1] According to Eurostat [2022], the percentage of single-child families in European countries in 2021 was almost half (49%), whereas The Office for National Statistics [2020] found this to be 43.7% in the UK in 2019. Such social trends are affected by many modern issues, such as economic concerns, becoming a parent at an older age, infertility status, marital lives, and careers with high pressure, growing childcare expenses, and the simple desire to have only one child. China's one-child policy has also contributed to only child families worldwide. There is, however, little understanding of only children's marriage outcomes.

Whether one can marry and to whom is hardly a matter of concern solely for only children; it also significantly impacts intergenerational relations. One of the major incentives for people to marry is the benefit derived from a larger family size, such as the provision of public goods, risk-sharing, and the advantages of economies of scale (Browning et al. [2014]). Many assume that they will either be single or the two of a couple in most cases. However, if we take a larger view of the family, whether one's marriage partner is an only child significantly impacts the size of each natal family. When an only child is young, parents can devote many resources to their dependent child. However, when parents become old and dependent, they cannot benefit from their family size. Unlike individuals with siblings, only children face the task of caring for their parents alone, typically after their prime marrying age has passed.[2] While the labor market may work and resolve these intergenerational burden gaps among regions or nations, the marriage market may act on such gaps among families. How does the marriage market affect this disparity between only child and non-only child households?

We expand the literature on marriage matching by highlighting sibling structure—particularly only-child status—as a key dimension of partner evaluation. Existing economic studies emphasize the positive assortativity in education and socio-economic status (SES), and its implications for inequality within households (e.g., Mare, 1991; Pencavel, 1998; Fernández and Rogerson, 2001; Greenwood et al., 2014; Eika et al., 2019). Recent advances in multidimensional matching theory (Chiappori et al., 2012, 2018) provide a framework to evaluate the match quality based on observable characteristics. However, these studies have largely overlooked the role of innate family structure.

[1] For instance, the percentage of such families in the US increased from 11% in 1976 to 21% in 2016; in Canada it increased from 12% in 1981 to 26% in 2019; and in Japan it increased from 10% in 2002 to 18.6% in 2015. In Singapore, 19.0% of married women had one child in 2010, while 24% did so in 2020.

[2] Indeed, the literature has shown that strong family ties or customs may negatively impact the younger generation's economic activities, especially in industrialized and urbanized economies [Alesina and Giuliano, 2010]. In families with only children, where externalities cannot occur, children are less likely to leave their parents [Konrad et al., 2002a, Rainer and Siedler, 2009], resulting in fewer labor market opportunities [Rainer and Siedler, 2009].

We build on Chiappori et al. [2018]'s framework to examine marriage matching and partner quality through the lens of only-child status. This is particularly relevant in aging societies, such as Japan, where caregiving responsibilities often fall disproportionately on individuals without siblings. In such contexts, sibling composition may function not only as a demographic trait but also as a signal of caregiving burdens and familial expectations. Social science research has begun to explore this dimension, showing that only children face greater difficulty finding partners due to presumed caregiving responsibilities [Yu and Hertog, 2018, Uchikoshi et al., 2023]. While such studies highlight sibling structure at the individual level, the market-level implications for partner matching remain unexplored.[3]

The purpose of this paper is twofold. First, we assess assortativity in marriage matching with respect to only-child status by conducting an analysis comprising three components. We first compare the observed marriage patterns with those predicted under a counterfactual scenario of random matching. Next, to examine the robustness of these patterns, we statistically evaluate the impact of only-child status on marital outcomes, particularly the likelihood of remaining single and that of marrying another only child. Finally, we analyze marital surplus based on the framework of Choo and Siow [2006], to provide one possible structural interpretation of the observed patterns. As a second component of the study, we estimate only child marriage matching outcomes following Chiappori et al. [2018], wherein the matching premium/penalty is measured by the difference in the partner's attractiveness. Here, educational attainment serves as a proxy for SES, and we examine how one's only-child status affects the marriage partner's SES.

In the first analysis, we find clear evidence of positive assortative matching with respect to only-child status. Although some mixed couples are observed, this reflects that the conditions for the symmetric case of Chiappori et al. [2018] are not fully satisfied. The analysis shows a greater tendency for only children to remain single and to form marriages with each other. Thus, the marriage market may contribute to reinforcing pre-existing disparities in family size. Moreover, our analysis of marital surplus indicates that surplus decreases along with increases in the number of only children within a couple. Taken together, these findings suggest that, although assortative matching by only-child status is present, it may not necessarily reflect the true preference among only children to marry one another.

Turning to the second analysis, we confirm the presence of a matching penalty associated with being an only child, and find that its magnitude appears to be gender asymmetric and contingent on the partner's only-child status. Specifically, while no significant penalty is observed in the pooled sample, only-child women face a pronounced matching penalty in the form of a reduction of approximately 0.63 years in their partner's years of education when

[3]Only a few studies in economics directly examine the effect of sibling composition on marriage outcomes. Angrist et al. [2010] found that individuals with many younger siblings are more likely to marry earlier and have more children, although only children were excluded. Vogl [2013] analyzed sibling-related marriage outcomes in Nepal, showing that younger sisters can accelerate marriage timing and reduce partner quality for older sisters. Appendix A provides a more detailed discussion of earlier studies in economics and the social sciences on sibling structure.

they marry a man who is also an only child. This magnitude substantially exceeds the gender gap in education within our sample, which is 0.34.

To deepen our analysis, we discuss the potential sources of the observed matching penalty based on our empirical findings and insights from the literature. In addition, we conduct two supplemental analyses. First, we explore heterogeneities based on respondents' birth year, age, and educational attainment, finding that although higher education mitigates assortativity, these characteristics do not alter the main findings regarding partner SES. Second, we examine alternative sibling structures by measuring the matching penalty for heirs defined under patrilineal and primogeniture systems. While heirs under the patrilineal definition consistently experience smaller penalties than only children, this pattern does not hold under the primogeniture definition. Together, these analyses offer further insight into the robustness and contextual variations in the documented matching patterns.

This study makes several contributions to the literature by investigating only children's marriage matching in Japan. On the one hand, this study explores the role of the marriage market by clarifying the nature of assortativity on sibling composition. If the burden of caring for parents differs between only and non-only children, then the degree of assortativity conveys a significant message in the context of an aging society. For example, suppose that the marriage market is characterized by negative assortative mating (i.e., only and non-only children are more likely to marry each other). In this case, the burden of family caregiving is moving more toward equalization. Conversely, in the case of positive assortative mating, the marriage market is accelerating inequality in this respect. Therefore, our findings may be meaningful for ascertaining whether the marriage market is driving equalization in family size. Recent social science studies have shown that sibling composition, such as being an only child or lacking male siblings, affects marriage patterns [Yu and Hertog, 2018, Uchikoshi et al., 2023]. This study contributes from an economic perspective by analyzing the market-level outcomes of one-child marriages, where the sibling structure may also reflect the underlying familial values and expectations.

Additionally, to our knowledge, this is the first study to measure marriage quality for only-child individuals. As noted earlier, few studies have examined marriage match quality with respect to only child status or sibling status composition, which are innate characteristics. Yu et al. [2012] and Angrist et al. [2010] examined the effects of sibling composition on marital status and age at first marriage as the marriage outcomes. Despite their insightful findings, these studies have failed to capture the perspective of the marriage market and matching with a partner. Thus, determining whether such an individual chooses their marital status or is forced to stay single remains challenging. Vogl [2013] examined women's arranged marriages in Nepal and the impact of certain sibling's presence on marriage outcomes and partner quality. Our study complements the literature by shedding light on only child marriages in developed countries with low fertility and aging populations from marriage market candidates' perspective. Furthermore, the only-child marriage patterns we uncovered reflect both sibling composition and education level, constituting a successful new application of a two-dimensional partner evaluation model [e.g., Chiappori et al., 2012, 2018].

Exploring the marriage matching of only children in Japan represents more than an exercise of academic curiosity. Many developed countries are experiencing population aging, with declining birth rates accompanied by an increasing number of only children. These phenomena have become a social problem owing to the burden of caring for older adults, with their longer life spans and extended caregiving periods, falls on their fewer children. Rainer and Siedler [2012] also suggested that the burden on the only child depends on the strength of social security and social expectations for informal family care. Japan is one of the countries with the most severely aging population and has an established social security system. At the same time, however, the duty of caring for older parents remains relatively strong in Japan and traditionally has been regarded as a family responsibility. Therefore, Japan is an interesting arena in which to explore how only children's relatively strong bonds with their parents may be linked to marriage matching outcomes and to shed light on the interdependence of parent–child relationships and marriage. Moreover, only children in the Japanese marriage market are also an appealing research population from the perspective of external validity as only children and children with siblings coexist in the same cohort, allowing cleaner observation of marriage market behavior.[4]

The remainder of this article is organized as follows. First, Section 2 explains the study background. Then, Section 3 describes the data. Section 4 elaborates on marriage patterns based on only child status, while Section 5 measures the marriage-matching premium/penalty. Section 6 discusses interpretations of the empirical results. In Section 7, we outline the supplementary analyses, while Section 8 presents the conclusions of the study.

2 Background

This section provides background on the distinctive character of the Japanese family and the only child.

2.1 Families in Japan

Confucianism has strongly affected the family system in Asian countries, where repaying parents is considered a virtue; the philosophical rationale for strong family ties is traced to Catholicism in Europe (Koyano [1996], Esping-Andersen [1997], Tsutsui et al. [2014]), resulting in a relatively large reliance on families rather than on society. In Japan, filial piety remains relatively strong, with caring for older parents traditionally being a family affair. According to Ministry of Health, Labour and Welfare [2020], in 2019, 28.2% of the primary caregivers for older adults requiring long-term care (LTC) were coresident couples of the younger generation

[4] While China's one-child policy presents an intriguing natural experiment, its uniform application complicates the interpretation of assortative matching patterns within standard theoretical frameworks. Specifically, incentives under the policy in China, such as permitting a second child only when both spouses are only children, may distort marriage choices [Lu, 2023]. See Wen [2023] as well.

(the older adults' children and their partners).[5] This figure is more than double the 12.1% of the care was provided by paid caregivers. Thus, the burden of care on the younger generation is still not being shouldered by the market.

Until the end of the WWII, Japan's inheritance system was patrilineal, with the eldest son inheriting the family head's entire estate. In 1947, the law was amended to allow family members other than the heir (typically, the eldest son) to inherit equally. However, for families that have existed for a long time or in farm families, a culture of inheritance by the family head (i.e., eldest son) still exists, along with heavier obligations attached to this role.

Moreover, the burden of care tends to be on a specific child and their spouse, perhaps with the couple living with a specific child's parents. If the child living with the parents is the eldest son and heir, his wife tends to make her marriage decision with the understanding that she may eventually care for the parents-in-law [Ogawa and Ermisch, 1996].[6]

However, it should be mentioned that the masculine norm is slowly fading and that the tendency to care for one's own parents has increased. Thus, while patrilineal co-residence remains more common (and although the practice of co-residence with parents is still stronger in Japan than in many advanced countries, co-residence itself has been declining over time), the gap between living with the husband's parents and with the wife's parents has been narrowing [National Institute of Population and Social Security Research, 2020].[7]

2.2 Only children in Japan

A survey that has been conducted since 1940 shows that the percentage of only child families has gradually increased since the 1990s [Cabinet Office, 2021], indicating an increase from 10% in 2002 to 18.6% in 2015. Meanwhile, the percentage of families with two children has remained stable at around 50% for almost 40 years. The rise in only child families can also be attributed to the decline in households with three or more children since the early 2000s.

In Japan's male-dominated society, relatively strong family norms persist, and the burden tends to be concentrated on the offspring, especially a specific child. Here, being a male or female only child is a unique characteristic in postmarital life (Yu and Hertog [2018]; Uchikoshi et al. [2023]). When an only child reaches adulthood, they automatically become the parents' heir. If the only child is a male, they automatically become the eldest son. If the only child is

[5] Comprehensive survey of living conditions, 2019.

[6] Ogawa and Ermisch [1996] show that in Japan's traditional family structure, the eldest son is strongly expected to support and care for his parents, which is a major motive for multigenerational co-residence. In practice, most multigenerational households with couples of childbearing age live with the husband's parents, and a majority of respondents cite "obligation as the eldest son" as the reason for co-residence. The study further provides evidence that caregiving responsibilities fall primarily on the eldest son's wife, thereby reducing women's wages and the probability of full-time employment.

[7] According to National Institute of Population and Social Security Research [2020], the share of households in which the wife (under age 70) lived with her parents increased from 6.5% in 1998 to 9.5% in 2018, whereas the share living with the husband's parents fell from 22.2% to 23.7%.

female, no eldest son exists to serve as the typical successor, while no other siblings exist to share the burden within the natal family. Consequently, the natal families of only child women have to give up not only intergenerational relations but also their surnames and family lines at the time of their marriage, all of which have been preserved over generations.[8] Encouragingly, the trend toward masculine domination is weakening (as discussed earlier). With this shift, only children, regardless of gender, are increasingly expected to bear the responsibility of caring for their own parents, particularly among younger generations.

Notably, an only child has the advantage that they can enjoy transfers from their parents throughout their lives. A prime example of postmarital income transfers are bequests from their parents.[9] Combined with the discussion on investment in education, which is examined later, Only children seem not only to be at a disadvantage but also to benefit from intergenerational relations.

Finally, despite the prevalent preference for sons in Asia, the male-to-female ratio among only children is nearly equal, as will be demonstrated with data in Section 3.[10] Thus, in Japan, the preference for sons may not be as pronounced as in other Asian countries where the sex ratio at birth is more distorted. Furthermore, families that exhibit a strong son preference often place significant value on upholding traditional family structures. These families are more inclined to pursue having multiple children to ensure the continuation of family leadership or adhering to the customary norm of having two children. Consequently, one-child families are more likely to reflect family planning based on the desired number of children rather than only on gender preference.

[8] In the past, when a woman had no brothers, it was common to adopt a son-in-law in order to maintain the family line.Wakabayashi and Horioka [2009] show that family patterns were generally consistent with the dynasty hypothesis, highlighting the strong linkage between parents and their successors. Specifically, self-employed parents were more likely to co-reside with their children, couples who adopted the wife's surname were less likely to live with the husband's parents, and even when the eldest child was female, families tended to co-reside with the eldest son. These findings underscore the enduring cultural centrality of the eldest son.

[9] According to The Yu-cho Foundation [2023], around 60% of respondents plan to divide their estates equally among their children, while about 10% would leave their entire estate to their only child. This suggests that only children may receive larger inheritances. The Dai-ichi Life Research Institute [2007], based on 715 samples, reported an average inheritance of 21.86 million yen, with 14.05 million yen for the eldest child and 13.03 million yen for younger siblings.

[10] Typically, the natural sex ratio at birth ranges from 103 to 107 boys for every 100 girls. However, this ratio tends to approach 50:50 as the children grow older. In Japan, for the generation under scrutiny, the sex ratio at birth has consistently hovered around 105, with the highest being 107.6 in 1966 [National Institute of Population and Social Security Research, 2021].

3 Data and summary statistics

3.1 The National Survey on Migration

The National Survey on Migration is a nationally representative survey from Japan, encompassing data on sibling configuration, birthplace prefecture, year of birth, and marital status of each family member. The data are collected by the National Institute of Population and Social Security Research through a random sampling method.[11] The surveyors distribute and collect the questionnaires for each household. This study uses the latest available waves from 1991, 1996, 2001, 2006, and 2011.[12] The collection rates for each wave are 89.4%, 95.8%, 85.5%, 74.0%, and 74.7%, respectively. In this study, we use the following information for the analyses.

Only child dummy: The survey asks for the number of surviving older brothers, older sisters, younger brothers, and younger sisters. When none of these are present, the only child dummy equals one for these individuals. All others are assigned a value of 0.

Age at the time of the survey: To account for changing trends over five-year intervals, we control for the age of respondents at the time of the survey, along with information on the year of birth.

Years of schooling: The number of years of schooling for the individual is calculated from the highest school graduated category.

Birth year: The respondent's year of birth is used.

Regional block: The survey asks for the prefecture where the respondent was born. In the analyses, the 47 prefectures are classified into ten blocks and used to account for regional characteristics.[13]

[11] The National Institute of Population and Social Security Research periodically assesses the accuracy of its valid individual data at both the regional block and population levels across five-year age groups, with those aged 85 years and older comprising one group (The National Institute of Population and Social Security Research 1993, 1998, 2005, 2009, 2013). To validate these data, comparisons are made with the corresponding population estimates from the Ministry of Internal Affairs and Communications' Statistics Bureau. Such comparisons were conducted during the 3rd survey in 1991, 5th survey in 2001, 6th survey in 2006, and 7th survey in 2011. However, the 4th survey in 1996 was compared with census data from 1995. It is important to note that the 7th survey in 2011 excluded the three Tohoku prefectures significantly affected by the Great East Japan Earthquake. Therefore, comparisons were not made with these affected prefectures. Across these five surveys, the largest discrepancy at the regional level was observed in the Tokyo area, which showed a difference of -3.3 points out of the 10 blocks surveyed in the seventh survey conducted in 2011. Additionally, the largest difference in age distribution was -1.0 points, which was found for the 20-24 age group in the fourth survey conducted in 1996.

[12] The sample size for the 1991 wave is relatively small. This is mainly due to the missing values for the place of origin. In addition, the 1991 wave uses slightly different questionnaires although it does contain our necessary variables, whereas the other waves have been changed to be more uniform.

[13] Specifically, Hokkaido for the "Hokkaido" block; Aomori, Iwate, Miyagi, Akita, Yamagata, and Fukushima for the "Tohoku" block; Saitama, Chiba, Tokyo, and Kanagawa for the "Minamikanto" block; Ibaraki, Tochigi, Gunma, Yamanashi, and Nagano for the "Kitakanto and Koshin" block; Niigata, Toyama, Ishikawa, and

3.2 Sample selection

Before examining the data, we highlight the key points regarding our samples. As explained later, this research explores marriage patterns of only children along with surplus analysis following the Choo and Siow [2006]'s framework. Then, we examine spousal SES matching based on the approach of Chiappori et al. [2018].

For the analyses of marriage patterns, we use information on the number of both single and married individuals. Ideally, the entire sample would be used for this analysis. However, due to data limitations discussed later, certain samples are excluded. Specifically, in the analyses of marriage patterns, while all single individuals from the survey are included, the sample of married individuals is restricted to household heads and their spouses. That is, married individuals who are not household heads or their spouses are excluded.

The issue with the dataset is that, while information on household members is available, detailed information about their spouses is not consistently provided. This information, such as the sibling composition, is only available for household heads and their spouses. Therefore, cases not involving household heads or their spouses must be excluded. [14]

Next, for the analysis of spousal matching characteristics, following Chiappori et al. [2018], we use data from household heads and their spouses for whom all relevant information was available. Consequently, singles and married individuals who are not household heads or their spouses are excluded from this analysis (the former exclusion is due to the definition of the analysis).

Formally, the sample is restricted to individuals aged between 23 and 65 in the survey year, for whom complete educational background information is available. The upper age limit is set due to the definition of the sibling structure variable. The number of siblings in the questionnaire is limited to those who are still alive. This restriction helps exclude older respondents who may have lost a sibling after entering the marriage market. Furthermore, the sample is limited to individuals for whom information on sibling composition, year of birth, and home prefecture is available.

Given these restrictions, unmarried individuals are limited to those who have never been married. The sample of married couples is restricted to those for whom information on both spouses is available; specifically, it includes couples where the respondent is either the household head or the spouse of the household head, and who are currently married. Individuals

Fukui for the "Hokuriku" block; Gifu, Shizuoka Aichi, and Mie for the "Tokai" block; Shiga, Kyoto, Osaka, Hyogo, Nara, and Wakayama for the "Kinki" block; Tottori, Shimane, Okayama, Hiroshima and Yamaguchi for the "Chugoku" block; Tokushima, Kagawa, Ehime, and Kochi for the "Shikoku" block; and Fukuoka, Saga, Nagasaki, Kumamoto, Oita, Miyazaki, Kagoshima, and Okinawa for the "Kyushu and Okinawa" block.

[14] The number of excluded cases is 4,144, which accounts for less than 10% of the total, and therefore, is considered a relatively modest proportion. Moreover, since we observe no clear difference in the probability of not being a household head or their spouse between only and non-only children, the main analysis is conducted using information from household heads and their spouses.

who are divorced, widowed, or separated are excluded. In addition, the sample is limited to couples in which both the respondent and their spouse reside in the same household. To ensure this, the sample is further restricted to households with at least two members. As a result of these sample restrictions, the number of observations decreased from 66,468 to 55,752.

3.3 Summary statistics

Table 1 is around here

Table 1 presents the descriptive statistics for each respondent's gender and only-child status. For continuous variables, the figures indicate the median within each subsample, while percentages or the 25th and 75th percentiles within the subsample are provided in parentheses. For categorical variables, the figures represent the number of samples and the percentage within each group. The following key points can be observed from these data.

First, the sample size of only children is necessarily small. Note that, as discussed in Section 2, no gender difference is observed for this variable, with the percentage of only children being about 7% each for men (=2,032/(26,945+2,032)) and women (=1,921/(24,854+1,921)). Second, only children's marital status is more likely to be single than that of non-only children at the descriptive level. Third, the average birth year of only children is later than that of non-only children. This may reflect the increasing prevalence of only-child families over time.

Finally, the key variable of interest, educational background, exhibits a noticeable difference based on only-child status. Specifically, only children tend to have slightly higher education levels. Regarding the years of education by gender, men generally have a higher level of education. Specifically, a large proportion of men in the 12-year high school graduate category as well as in the highest 16-year category. Conversely, women are concentrated in the high school diploma group and the second-highest educational level (14 years of schooling). This suggests that the polarization of educational attainment is more pronounced among men.

4 Marriage patterns for only-child status

This section explores marriage patterns based on three analyses. The first analysis descriptively examines whether the observed marriage patterns related to only-child status exhibit positive or negative assortative matching by comparing them with those generated through random matching. The second analysis formally tests these findings while also investigating the impact on the probability of remaining unmarried. Finally, we conduct supplementary analyses to gain a structural understanding of marriage patterns. Two key discussions emerge from these analyses.

The first key point is to understand the role of the marriage market by examining marriage patterns related to only-child status. Marriage patterns can exhibit positive or negative assortative mating [Becker, 1991], and different tendencies may be expected when viewed through the lens of only-child status.

The former occurs when individuals select partners who share characteristics. If caregiving responsibilities and economic incentives influence marriage decisions, non-only children, who generally face fewer caregiving burdens, may exit the marriage market first. Meanwhile, only children may be more likely to marry other only children.[15] Conversely, the latter arises when individuals avoid marrying those similar to them. A possible cause of this pattern could be the benefit of trade between providing domestic public goods. For example, an only child with high caregiving responsibilities may avoid marrying another only child. They may instead prefer a non-only child with lower caregiving burdens.

Additionally, financial resources and dynastic incentives may promote marriages between only and non-only children. Preferences in marriage may also be influenced not only by economic incentives but also by family norms and the intergenerational transmission of values. Recent studies have raised theoretical questions regarding individuals from families with strong norms of filial piety: Do they tend to choose partners who share similar values or do they instead prioritize their own family by selecting partners with weaker norms [Cigno et al., 2017, 2021].[16] Such intergenerational normative factors may explain the observed assortativity based on sibling structure, not because sibling composition is a familial value in itself, but because it can shape expectations about the family roles, particularly caregiving responsibilities.

If the analyses reveal that positive assortative mating holds, then, combined with the descriptive statistics showing that only children are more likely to remain unmarried, this suggests that the marriage market may function in a way that exacerbates disparities in the caregiving burden among the working-age population.

The second key point concerns examining marriage patterns in light of the framework of Chiappori et al. [2018]. Their model predicts perfect positive assortative matching, meaning that men and women are exactly matched on multiple characteristics such as education and smoking habits. This prediction holds only under the symmetric case, which requires that the

[15] If inheritance benefits outweigh caregiving burdens, only children may exit the market first, followed by non-only children. Shared family size preferences may also drive similar sibling structures in marriage.

[16] Recent studies have raised theoretical questions about the marriage partner of individuals from families with strong norms of filial obligation. Unlike other transmitted value factors such as ethnicity, religion, attitudes toward working women, or economic preferences (Bisin and Verdier [2000]; Bisin et al. [2004]; Fernández et al. [2004];Wu and Zhang [2021]), the transmission of strong family norms regarding filial piety can potentially create conflicts within families. According to Cigno et al. [2017] and Cigno et al. [2021], individuals from families with strong filial piety norms may choose to marry partners who also uphold these norms to preserve them. However, they may also prioritize their own family by marrying partners with weaker norms. If we consider only children as the group with the strongest parental care norms, assessing assortativity can help us provide some insights into this question. Although we do not directly uncover the underlying mechanism behind these results, our findings on partner choices can shed some light through the observable outcomes in the younger generations' adult pairings.

characteristics of men and women are similarly distributed and that these characteristics affect marital surplus in the same way for both genders. While our data indicate that the proportion of only children is the same for men and women, other conditions (e.g., the distribution of education) are not satisfied. The following analysis, therefore, evaluates how the observed marriage patterns deviate from this benchmark prediction.

4.1 Comparison with random matching

To capture the overall trend in marriage patterns among only children, we first compare the observed marriage patterns based on only-child status with those generated via random matching. For this purpose, we construct a counterfactual sample of randomly matched couples, following the permutation procedure employed by Chiappori et al. [2018]. We assign pseudo-random IDs drawn from a uniform distribution to men and women, rank them accordingly, and match them based on these ranks.

Table 2 presents the observed and randomly generated patterns. Only children are more likely to marry other only children. Although the percentage difference in marriages between only children may seem small at first glance, it becomes striking when viewed from the only child's perspective. For example, in the random matching scenario, the share of marriages between only children is 7.3% among men (96/1319) and 7.7% among women (96/1242). Meanwhile, in the observed data, it rises to 14.5% for men (191/1319) and 15.4% for women (191/1242). Formally, the chi-squared test strongly rejects the null hypothesis that only-child status and matching patterns are independent ($\chi^2(1) = 166.2$, $p < 0.001$).

Despite the tendency of positive assortative matching, mixed marriages between only and non-only children also occur. As discussed above, according to Chiappori et al. [2018], perfect assortative matching would arise under a symmetric case wherein male and female characteristics play the same role in the surplus function, and the distributions of these characteristics are the same across genders (Proposition 1, p.169). This requires: (i) the distribution of only-child status is the same across genders; (ii) the distribution of SES is the same across genders; (iii) the effect of only-child status on marital surplus is symmetric for husbands and wives; and (iv) the effect of education on marital surplus is also symmetric. At the observable level, Table 1 in Section 3.3 shows that the proportion of only children is nearly identical for men and women, suggesting symmetry in this dimension. Conversely, men tend to have higher educational attainment and greater variance in schooling. Thus, the distribution of education, which we use as an SES proxy, already violates one of the conditions for the symmetric case, which may hinder perfect assortative matching from fully holding.

4.2 Formal analysis of marriage patterns

This subsection statistically verifies the two trends previously observed in the marriage patterns of only children. Specifically, we examine the likelihood of each marriage pattern. Specifically,

the estimand is the average values of the following:

$$E[Type^{partner} \mid OnlyChild, X] - E[Type^{partner} \mid NotOnlyChild, X] \tag{1}$$

This equation shows the expected difference in the likelihood of marriage outcomes between only and non-only children, where $Type^{partner} \in \{Only\,Child, Not\,Only\,Child, Single\}$ is defined as follows: *Only Child* indicates that the partner is an only child, *Not Only Child* indicates that the partner has siblings, and *Single* indicates that the individual remains unmarried.

Then, we present the estimates of how marital status differs between only and non-only children, controlling for gender, year of birth, age, and birthplace, as shown in Eq.(1).[17]

In the estimation, we employ the double-debiased machine learning for augmented inverse probability weighting estimator [Robins and Rotnitzky, 1995, Chernozhukov et al., 2018]. While ordinary least squares (OLS) provides consistent estimators when the model is correctly specified, it can yield large biases under misspecification. Given the complexity of the confounding factors in the marriage market, strong reliance on linear specifications is risky, which motivates our use of double-debiased machine learning.

Double-debiased machine learning builds on augmented inverse probability weighting and incorporates cross-fitting, which mitigates the risk of overfitting and ensures asymptotic unbiasedness even when flexible machine learning methods are used. Nevertheless, double-debiased machine learning is not without limitations. It can suffer from large variance when the estimated propensity scores are close to zero or one, although this issue is not severe in our application (the minimum estimated propensity score is 0.038). Moreover, the protection against overfitting is only asymptotic; with small samples, this risk may not be fully alleviated. However, because our dataset is sufficiently large, these concerns are unlikely to affect our results. As shown in Table 1, our analysis is based on $N = 55{,}752$ individuals.[18]

Finally, the specific method can be summarized as follows: In the first stage, we estimate the nuisance functions as predictive models of $Type^{partner}$ and $OnlyChild$. To flexibly capture conditional means without relying on parametric assumptions, we implement a stacking algorithm that combines OLS, random forests, and Bayesian additive regression trees. In the second stage, these estimates are plugged into the augmented inverse probability weighting procedure to obtain the final estimator.

The estimator requires estimating the nuisance functions of $Type^{partner}$ and $OnlyChild$, $[E[Type^{partner} \mid OnlyChild, X]$ and $E[OnlyChild \mid X]$, where $OnlyChild = 1$ if an individual is an only child. X is the control variables including years of one's own schooling,

[17] Some may argue that using one's own education as a control variable can pose challenges to drawing causal inferences, as it is a post-treatment variable. However, our objective is not to measure causal effects but rather to shed light on how the presence or absence of siblings influences family formation among adults. Therefore, this study incorporates one's own education as a control variable, following the empirical framework of Chiappori et al. [2018]. Furthermore, even if education is not controlled for, it does not significantly alter the main findings based on our specification (see Appendix C.)

[18] In the couple-level analysis, the unit of observation is the household, yielding $N = 20{,}019$ married couples. This is a sufficiently large sample for our estimation strategy.

birth year, age, and region of birth. Note that we use robust standard error clustering at the household-level. We estimate this equation for each male and female sample. Technically, the estimator captures the percentage-point difference in the share of individuals in each marital status, defined by $Type^{partner}$, between only children and non-only children. Intuitively, it shows how being an only child affects the likelihood of each marriage state.

Figure 1 is around here

Figure 1 shows the estimated effects for *OnlyChild* conditional on the type of marriage partner (as well as single status), which are sorted by own gender. The left panels show the results for singles, while the right panels show the results for marriage with only children. Note that we only present the results for the two statuses as the effects of the remaining group of those marrying non-only children are automatically calculated using these two results.

Comparing the effects across the panels, we obtain four main insights: First, considering the panel as a whole, we observe similar trends in each panel regardless of gender. Second, the panel of singles reveals that only children are more likely to remain single than non-only children. The estimated effects are 0.07 for men and 0.05 for women. Although the point estimates are somewhat larger for men, the gender difference is not statistically significant.[19] Third, only children are more likely to marry only-child partners. The related values are 0.05 for men and 0.06 for women. Finally, inextricably associated with the above results, the only child status reduces the likelihood of marrying a non-only child partner by 0.12 for men and 0.11 for women.

These results highlight two important points. First, people choose mates in a positive assortative manner regarding only-child status.[20] Second, only children are less likely to get married. These results indicate that family size tends to increase with marriage for non-only children. Consequently, family size disparities are heightened by dynamics within the marriage market.

4.3 Systematic returns to marriage

Here, we estimate marriage returns, thus offering a structural perspective on the trends observed in the previous subsections. We first estimate the household-level marriage gain. According to Choo and Siow [2006], the total systematic gain to marriage for a type i male and type j female can be identified by estimating the logarithm of the following variable:

[19] The difference between men and women is not statistically significant ($p = 0.140$ for single status and $p = 0.571$ for only-child status).

[20] Although we here limit ourselves to likelihood comparisons to intuitively understand the state of the marriage market, we conduct a formal analysis to measure assortativity on only child status in Appendix B, following Chiappori et al. [2025]. Using multiple other indices, we further confirm positive assortative mating on only child status.

$$\Pi_{i*j} = \frac{\mu_{ij}}{\sqrt{\mu_{i0} \times \mu_{0j}}} \tag{2}$$

where μ_{ij} represents the number of marriages between type i males and type j females, whereas μ_{i0} and μ_{0j} denote the number of single males and females of each respective type. The larger this value, the greater the relative return on marriage for types i and j compared to remaining single. As noted in Choo and Siow [2006], the calculation normalizes by the number of singles, eliminating scale effects.

Additionally, individual-level systematic return can be computed. The return from marriage for a type i male to a type j female is given by:

$$n_{ij} = \ln\left(\frac{\mu_{ij}}{\mu_{i0}}\right) \tag{3}$$

Similarly, the systematic return from marriage for a type j female to a type i male is:

$$N_{ij} = \ln\left(\frac{\mu_{ij}}{\mu_{0j}}\right) \tag{4}$$

which can be identified through this estimation.

Table 3 shows the estimated household- and individual-level returns to marriage by type ij, specifically focusing on the combinations of the spouses' only-child statuses. Table 3 categorizes marriage patterns by only-child status, naming combinations in the order of male and female. Each row corresponds to a specific marriage pattern. For instance, a marriage between a non-only child male and an only child female is labeled as "Non-Only Child – Only Child Marriage." The three columns report the total return to marriage in Eq.(2), male's individual return in Eq.(3), and female's individual return in Eq.(4).

The highest total return is observed for marriages between non-only children, followed by mixed marriages where the female is an only child, then mixed marriages where the male is an only child. The lowest return is observed in marriages between two only children. This suggests that marriages involving only children tend to have lower returns. Furthermore, a comparison between mixed marriages where the husband is an only child and those where the wife is an only child reveals that the surplus size is generally similar. Thus, the presence of an only child has limited gender-based impact on the total household surplus. At the individual level, two main findings emerge: On the one hand, in homogeneous marriages concerning only-child status, females generally have higher returns. On the other hand, in mixed marriages, the only-child partner consistently attains higher individual returns than the non-only-child partner, regardless of gender. In addition, the gender difference in the surplus loss from marrying an only child appears small, with no clear pattern emerging.

5 Only-child matching premium/penalty in the marriage market

In this section, we examine whether an individual's only-child status is related to the quality of the marriage match. Ideally, we would like to measure the only child matching premium/penalty defined on a utility basis. However, this is not possible because it is unobservable. Thus, we alternatively use the partner's SES as its approximation. Human capital, as a key component of SES, is likely a monotonic form of attractiveness and linking it with the arguments on household inequality is practical [Mare, 1991, Pencavel, 1998, Fernández and Rogerson, 2001, Breen and Salazar, 2011, Greenwood et al., 2014, 2016, Eika et al., 2019]. Here, following Chiappori et al. [2018], we estimate the matching premium/penalty regarding partner SES.

5.1 Estimand

This subsection presents the estimands that are estimated in the following sections. Let $SES^{partner}$ denote the partner's SES. The estimand here is the average values of the following expression:

$$E[SES^{partner} \mid OnlyChild, X] - E[SES^{partner} \mid NotOnlyChild, X]. \tag{5}$$

Eq.(5) shows that only children's partners are likely to be more or less educated than non-only children's partners. Note that if we assume $E[SES^{partner}|OnlyChild, X] = \beta_0 + \beta_1 OnlyChild + \beta X$, we can use the same approach as in Chiappori et al. [2018]. However, the following section proposes a more flexible estimation strategy rather than relying on such a linear specification.

In the following analyses, we use SES as a proxy for a partner's attractiveness other than only-child status. As in Chiappori et al. [2018], we assume that, holding other conditions constant, individuals want a marriage partner with a higher SES. Thus, when Eq.(5) takes negative values, the married partner's SES is lower for the only child, suggesting the existence of an only child penalty and vice versa.

5.2 Estimation method

We estimate how the partner's SES differs between only and non-only children, controlling for X, as described in Eq.(5) and its estimand. The dependent variable is $SES^{partner}$, which captures attractiveness other than only-child status. Operationally, following Chiappori et al. [2018], we proxy $SES^{partner}$ by the partner's years of schooling. As in the previous analysis, $OnlyChild$, our variable of interest, is a dummy variable that equals one for an only child. X includes own years of schooling, birth year, age, and birth region. Again, we control for the individual's years of schooling to focus on the disparity in adulthood and to consider positive

assortativity on education for couples. Note that the same machine-learning procedure is followed as in the previous analysis in Section 4.2.

5.3 Results: Partner's years of schooling (Pooled sample analysis)

Figure 2 is around here

Figure 2 shows the estimated effects of *OnlyChild* using the total sample by gender. The panel presents the estimated effect of only-child status on the spouse's years of schooling. Figure 2 does not show a significant only-child-matching premium or a penalty in the pooled sample for either men or women.

5.4 Results: Partner's years of schooling (Subsample analysis by partner's status)

As we have seen, the difference is small and unclear in the results for the pooled data. However, considering that the benefits of larger family size may differ, the partner's only child status may matter in determining the premium/penalty. Therefore, we analyze the SES penalty using subsamples characterized by the partner's only child status.

Figure 3 is around here.

Figure 3 presents the results for this subsample analysis, restricted to the respondent's gender and the marriage partner's only-child status. Similar to the pooled-sample figure (Figure 2), the upper and lower parts of this figure correspond to the male and female samples, respectively. Furthermore, we demonstrate the results for the left-hand panels with individuals whose marriage partner is a non-only child, and the right panels show individuals whose marriage partner is an only child.

Figure 3 shows that the results vary considerably by the respondent's gender and the partner's only-child status. In the subsample in which the spouse is a non-only child, the effects for both men and women are close to zero, indicating that an individual's only-child status does not significantly affect their partner's years of schooling in such marriages. However, in the subsample where the partner is an only child, gender-specific patterns emerge: the effect for men remains near zero (-0.03, not significant at the 5% level), whereas the effect for women is notably negative at -0.63 and statistically significant. The penalty for female-only children married to male-only children (-0.63) is nearly twice the average gender gap observed in the full sample (0.34). Given that a gap of 0.34 years is already substantial, this comparison illustrates the considerable magnitude of the estimated effect. A more detailed interpretation of these findings is presented in Section 6.

6 Discussion

In the previous section, we showed that when only-child women marry only-child men, their husbands' educational attainment is 0.63 years lower than that of the husbands of non-only-child women. We now explore possible interpretations of why the outcomes differ depending on the partner's characteristics and whether gender-specific differences exist.

First, we consider why the only-child penalty is observed particularly when the marriage partner is also an only child. One possibility is that only children prefer partners who are also only children, such as due to shared values or experiences, even if this means compromising on SES. However, in Japan, only-child households are less common than in many Western countries. Further, the ideal family structure is often perceived as having two siblings [Raymo, 2022]. Thus, "marriages between only children" may not necessarily be considered aspirational.

Another explanation may be that only children, who may face disadvantages in the marriage market, end up marrying each other as a form of mutual compromise. This view is supported by prior research. For instance, Yu and Hertog [2018] found that in online matchmaking in Japan, only children tend to be avoided by others; moreover, even among only children, mutual avoidance may occur.

Turning to our own evidence, although the surplus analysis presented in Section 4.3 cannot be directly compared with the penalty estimations based on the framework of Chiappori et al. [2018], the findings do partly align with this interpretation. Specifically, being an only child reduces the couple's surplus, with couples composed of two only children showing the smallest total surplus, as illustrated in Table 3.

Next, we explain why the penalty appears only for women. Table 3 suggests that the effect of only-child status on surplus does not differ substantially by gender. Meanwhile, Figure 3 shows that the penalty in terms of spouses' educational attainment is observed only for women. Initially, these findings may appear inconsistent: Table 3 suggests limited gender difference in the impact of only-child status on surplus, while Figure 3 highlights a gender-specific penalty. However, several explanations may underlie this apparent inconsistency.

First, gender differences in the distribution of education amplify this asymmetry. Male educational attainment is more dispersed, making penalties for women more salient when measured through their partners' education. Second, education, as a component of SES, has different implications for marital surplus by gender. Research has consistently shown that men's SES is more highly valued in the marriage market than women's [Fisman et al., 2006, Hitsch et al., 2010, Low, 2014, Bertrand et al., 2015]. Because education, as a component of SES, carries different weights for men and women in the marriage market, the penalty becomes more salient for female-only children, who are more likely to experience it through the lower educational attainment of their spouses. In the Japanese context, Uchikoshi et al. [2024] further demonstrates that women are more sensitive to a partner's income than men. This reinforces the view that men's SES, including education, carries greater weight. Conversely, the distribution

of only-child status does not significantly differ between men and women. Thus, distributional differences in only-child status cannot account for the asymmetry.

In summary, the apparent inconsistency between Table 3 and Figure 3 can be reconciled by considering both the distributional properties of education and its role in shaping marital surplus. Therefore, the gender asymmetry observed in Figure 3 is less likely to arise from only-child status per se; rather, it reflects how education functions in the marriage market as a core element of SES.

Finally, Vogl [2013] offers an important comparison. The author studied arranged marriage markets in developing countries in an analysis focusing on specific sibling compositions. The finding revealed that women with sisters tend to marry earlier and their spouses tend to have lower educational attainment, which is a sign of lower search quality or compromise in partner selection.[21] Conversely, we find that although only children may be slightly more likely to remain single, when they do marry, their spouses tend to have lower educational attainment. This suggests that the mechanism behind our findings differs from those driven primarily by rushed or early marriage, yielding lower-quality matches.

7 Supplementary analyses

We have thus far observed that only-child individuals incur penalties in terms of lower partner SES. In addition to their heavier burden of caregiving, a widely discussed factor, only children also face a matching penalty, particularly when they marry another only child, potentially exacerbating the inequality. In this section, we deepen our understanding of the marital outcomes of only children from two perspectives: first, by conducting a heterogeneity analysis, and second, by examining alternative sibling configurations.

7.1 Heterogeneity of only child penalty

Here, we examine how the one-child matching penalty is affected by heterogeneity. In particular, we focus on two types of heterogeneity. The first is the demographic variables of birth year and age. The meaning of being an only child may change over time, and individuals may change their marriage behaviors depending on their own age. The second is educational background. A higher level of education may increase one's attractiveness and weaken the effect of being an only child. In addition, if caregiving duties are critical for their attractiveness, one may purchase these services from the market if they can sufficiently afford to do so.

To explore these issues, we examine the effects of heterogeneity separately for men and women. Specifically, we estimate a linear approximation model of the conditional mean difference in

[21] Vogl [2013]'s findings indicated that when a woman has a younger sibling of the same gender (i.e., a sister), the quality of the mate tends to be lower. According to the study, this outcome was attributed to the older sister being hurried into marriage to expedite the younger sister's marriage.

the heterogeneity analysis. Semenova and Chernozhukov [2021] extended the work of Chernozhukov et al. [2018] and proposed a method to estimate a linear approximation model of the mean difference using double-debiased machine learning. This method is expected to reduce misspecification errors and biases, as the estimation of the mean difference does. Asymptotic unbiasedness and normality hold for linear approximation models similar to the augmented inverse probability weighting estimation and approximate computation of confidence intervals is possible. For this reason, we employ this method for the heterogeneity analysis.

Figure 4 is around here.

Figure 4 shows the results of the heterogeneity analysis of marriage patterns. We first examine the effect of birth year. The results show that the more recent the respondent's birth cohort, the stronger the main effect: only children are somewhat more likely to remain single (though this effect is not statistically significant), less likely to marry non-only children, and more likely to marry other only children. Regarding age, being older at the time of the survey does not significantly affect the likelihood of remaining single, but it is associated with a stronger tendency for only children to marry other only children. For education, the strength of the main result weakens. We next interpret these results from the perspective of economic incentives to form larger families and parental monetary transfers.

Figure 4 shows counterintuitive results for birth year if we account for the circumstances in which Japan has experienced the gradual socialization of older adults' care through public policies (and slowly fading social norms regarding filial obligations). [22] However, this trend of socialization of care may also reflect the trend of women's advancement in society and the decrease in their labor in households. Women who were typically used to care for their parents and the parents-in-law are now more difficult to rely on as caregivers. Consequently, the difficulty of caregiving arrangements within households may lead to the enhanced trend that people choose mates in a positive assortative manner based on only child status (i.e., only children are less likely to be chosen as a marital partner). In terms of other factors, such as monetary transfer, only children's attractiveness may be weakened by the declining relative importance of parental monetary transfers compared with their own economic capabilities.

Regarding the heterogeneity in age, the main trend becomes more prominent as respondents get older at the time of the survey. Given that younger individuals remain single, this result reflects that the trend regarding the choice of marriage partner becomes even stronger when comparing those who are already married and young compared to those who are older. This suggests that individuals who marry at a later age may be selecting partners with a more realistic understanding of caregiving responsibilities. The results for education level may reflect that only children can overcome their disadvantages if one's educational background is higher, as discussed above.

Figure 5 is around here.

[22] Historically, children's burden of filial duty has been declining. First, in 1961, Japan enacted a national pension system. Additionally, with the enactment of the Long-Term Care Insurance Law in 2000, a system was put in place for society to support the care of older adults.

Finally, Figure 5 shows the estimates of the matching premium/penalty measured by the partner's SES. The main results are not significantly affected by either own birth year, age, or education. In summary, heterogeneity affects the choice of marital partner based only on sibling structure and not the partner's SES.

7.2 Excluding education from controls

This subsection examines whether our main result holds when education is not controlled for. As discussed in the main text, only children tend to be more educated, which may offset the penalty. Thus, we also explore how being an only child affects the partner's education overall and how this impact changes once positive educational assortativity is taken into account. In Appendix C, we follow the same procedure as in the main text but exclude the respondent's own education from the control variables.

According to Appendix C, only children are more likely to remain single and more likely to marry another only child, consistent with the main results obtained without controlling for education. However, the results for years of schooling show a modest deviation from the main findings. A clear difference emerges depending on the type of marriage partner: Being married to a non-only child significantly increases the partner's education level for only children, although the effect is relatively small. Conversely, the partner's education tends to decrease for both male and female only children who marry another only child. As the main results show, this is particularly evident for women, with effects that are significant and of similar magnitude.

In summary, although the impact is small, a notable difference from the main result is the presence of a premium for those married to non-only children. This pattern is likely due to the omission of education controls and reflects the well-documented positive correlation in educational attainment between spouses. However, this does not contradict our main finding of a penalty associated with marrying an only child.

7.3 The effects of alternative sibling positions

Thus far, we have examined the effects of being an only child on marriage patterns. However, the underlying factors driving these effects remain to be understood. To gain further insights, we conduct an additional analysis that considers alternative sibling positions and interprets them in terms of intergenerational relationships. If such relationships account for the disadvantages experienced by only children in the marriage market, policy interventions, such as promoting the socialization of informal care, could help mitigate these penalties.

In this context, we explore two alternative sibling positions: the effect of being the eldest son and that of the eldest child. In Japan, the eldest son (and their wife) traditionally bears certain obligations, including caring for their parents. Similarly, the eldest daughter with no male siblings is expected to assume this role. While the influence of the eldest son is widely

known, the concept of primogeniture, wherein inheritance goes to the eldest child, suggests that birth order may be more significant than simply being the eldest son. Recent trends indicate a growing preference for individuals to take care of their own parents and expect their own children to care for them. Additionally, considering the persistent gender gap in the provision of informal care, women are expected to care for their parents even if they have younger brothers. If two siblings have a significant age difference, the first-born effect may be even more pronounced than that of the eldest son. Considering these two cases, we test whether the effects of these sibling positions exist. To focus on sibling position, we analyze samples from families with only two siblings. This approach allows us to extend our main findings on the only child effect and make meaningful comparisons (for the detailed analysis, see Appendix D).

Two key findings emerge. First, patrilineal heiresses who marry only children exhibit effects that, while not statistically significant, are partially consistent with the main results regarding the marriage patterns of only-child women and their partners' educational background. Specifically, the observed penalty for patrilineal heiresses married to an only child aligns with our main findings, indicating a sizable impact; however, it is not statistically significant. The absence of a partner education penalty for heiresses married to individuals who are non-only children also coincides with our results for only children.

Second, among men expected to bear family responsibilities, especially patrilineal heirs, we observe clear tendencies in partner selection. Specifically, they tend to have a higher probability of remaining single and are more likely to avoid marrying only children. They may themselves be avoided due to their family responsibilities, which mirrors the patterns observed for only children. Moreover, the tendency of eldest sons with siblings to avoid only children does not contradict the main result of positive assortativity by only-child status. Instead, it may suggest that these eldest sons, who are likely to bear heavier familial duties, are more cautious in choosing a spouse who, as an only child, may prioritize caring for their own parents.

These findings do not dismiss the possibility that the strength of intergenerational relationships plays a role in the penalties faced by only children. Furthermore, the penalty for being an only child exceeds that of being an heir with one sibling, as suggested by the analyses based on the two definitions. This highlights the potential challenges faced by only children who lack the support of siblings; conversely, heirs can share their responsibilities among siblings.

8 Conclusion

We investigate marriage matching among only children, focusing on the strength of their intergenerational ties. Our analysis shows that assortative mating based on only-child status is significantly more pronounced in actual marriages than in random matches. Statistical tests confirm the presence of positive assortative mating, and we also find that only children are more likely to remain single.

Second, we measure the matching premium/penalty by applying the framework of Chiappori et al. [2018], finding an only-child matching penalty in terms of lower partner SES. Furthermore, this penalty is more pronounced for female only children who marry male only children, which we attempt to explain in the discussion section by combining our data analysis with existing literature.

Moreover, we conduct additional analyses to obtain a more profound understanding of the underlying cause of the penalty. Heterogeneity analyses reveal that one's own educational level alleviates the disparity in partner choice. Additionally, assortativity based on only child status becomes more pronounced among respondents born more recently and those who are older. Finally, other analyses exploring alternative sibling positions do not refute the possibility that heavier filial obligations influence only children's marriage patterns.

Several conclusions can be drawn from these findings. First, sibling composition, which is an inherent factor beyond an individual's control, tends to disadvantage certain individuals in terms of finding a suitable match in the marriage market. Specifically, being an only child leads individuals to compromise on the attractiveness of their marriage partners' SES. Second, the marriage market exacerbates disparities in family sizes. Given that only children tend to bear greater caregiving responsibility, our finding that they are more likely to remain unmarried or to marry other only children implies an increase in the inequality of caregiving burdens through the marriage market. Regarding the penalties faced by only children, socializing the burden of care may help address two negative aspects: the disadvantageous marriages of only children and the widening gap in caregiving burdens among the younger generations.

Before concluding, we will discuss the limitations of this study and future research directions. The only information we have regarding sibling composition is on *surviving* siblings. Since we may include those who have already lost siblings in our sample, we may underestimate the penalty for older generations. This issue also prevents us from considering the possibility that only child status may be driven by biological factors (e.g., infertility or the low probability of survival of all siblings), as Lu and Vogl [2022] noted. Since families with weak constitutions may be disadvantaged in the marriage market, such data may allow for alternative interpretations. Considering data limitations, parental information that could impact monetary transfers, such as bequests, should also be considered when examining sibling composition and marital outcomes. However, the dataset utilized here lacks information on parental SES. Including such additional parental information can offer further insights.

Relatedly, this study uses data from Japan, which is an East Asian country with strong traditions of filial piety, to focus on intergenerational relationships. Although the results appear reasonable, other sources of explanation for the only child penalty are possible. Future research can explore the effects of variations in policy changes on marriage patterns, if any, in other economies, à la Bau [2021], who demonstrated that pension reforms implemented in male-dominated and female-dominated societies affect cultural changes in marital practices. Scholars can also analyze the impact of alternative sibling structures on marital outcomes. While the current analysis primarily focused on only-child status, which is less influenced

by specific periods and cultures, a potential avenue for future research is to comprehensively investigate the effects of sibling structure in conjunction with those of masculine culture.

Furthermore, this study is restricted to analyzing marital status and spouse characteristics as dependent variables. However, it does not determine whether individuals are actually experiencing penalties or facing life challenges. Ideally, investigating whether sibling composition results in differences in utility levels could provide valuable insights. If subjective well-being indicators were available, this could open up another avenue for future research, broadening the understanding of our findings and deepening insights into how sibling configurations influence life outcomes.

Finally, our findings can provide valuable insights into the pace of population decline. Vogl [2020], focusing on the evolutionary process of intergenerational associations in fertility, raised the possibility of marriage assortativity as a mechanism that may contribute to the acceleration of population decline. Our discovery of assortativity represents a significant advancement in our understanding of demographics. Thus, it would be meaningful to examine the demographic impact of positive assortativity on sibship size in the marriage market, as it may contribute to a further decline in fertility rates.

Appendix A. Prior literature on only children and marriage

This appendix concisely reviews previous research on only children and marriage outcomes. Specifically, we focus on three dimensions: (1) the association between sibling composition and marriage outcomes, (2) educational attainment, and (3) relevant psychosocial mechanisms and other lifetime outcomes.

A.1 Studies on sibling composition and marital outcomes

While not always centered on only children, many social science studies have explored how family background—including sibship size and birth order—shapes marital behavior. For example, Yu et al. [2012] found that birth order and sibship gender composition affect age at first marriage in gender-asymmetric ways. In Japan, Kojima [1993] discussed marriage arrangements based on sibling structure.

More directly, Yu and Hertog [2018] analyzed online dating behavior in Japan, and found that only children receive fewer requests and are more likely to send requests, suggesting a disadvantage in the early stages of mate search. Using representative data from Japan, Uchikoshi et al. [2023] also analyzed the effect of the change in population structure of sibling composition (including only children) on demographic change using different indicators. The authors showed that children who are expected to care for their parents are less likely to marry. Specifically, they calculated the percentage of the male/female population with a particular sibling composition that is actually married to someone with that background.

In economics, Vogl [2013] examined arranged marriages in Nepal and showed that sibling composition, specifically the presence of younger sisters, can pressurize women into earlier and lower-quality marriages. Thus, sibling structure may influence not only whether and when individuals marry but also the quality of their matches.

A.2. Studies on sibling composition and educational outcomes

We use the education of one's partner as a key indicator of marriage quality. Studies have examined whether being an only child affects educational attainment, a potential confounder in assortative mating. Theoretical frameworks such as the quality–quantity trade-off model suggest that children in smaller families may benefit from more parental investment (Becker and Lewis, 1973, Becker and Tomes, 1976, Galor and Weil, 2000, Hazan and Berdugo, 2002, Moav, 2005).

Empirical findings, however, are mixed. Some studies report educational disadvantages for only children. For instance, Black et al. [2005] and Qian [2009] showed that only children tend to have lower educational outcomes than those with siblings.

Conversely, other studies highlight the advantages of being an only child or growing up in a smaller family. For example, Lee [2008] found that only children receive higher per-child expenditures. Meanwhile, Rosenzweig and Zhang [2009] showed that the presence of twin siblings, compared to being an only child, lowers educational outcomes. Further evidence using the one-child policy as a natural experiment suggests that additional siblings reduce educational attainment: Liu [2014] and Li and Zhang [2017] found negative effects of sibling size using one-child policy variation, while Qin et al. [2017] reach a similar conclusion using a regression discontinuity design.

Taken together, these findings suggest that while some studies emphasize the disadvantages of being an only child, others underscore the benefits of smaller family size. The impact of sibling composition on educational attainment is therefore complex and context-dependent.

A.3. Studies on sibling composition and other outcomes

Psychological studies have long debated whether only children differ from those with siblings in terms of personality and social development. Early stereotypes portrayed only children as spoiled or socially disadvantaged, a view famously summarized by G. Stanley Hall's remark that being an only child is "a disease in itself' " [as cited in Fenton, 1928]. While such views have been challenged, they continue to persist in public discourse [Mancillas, 2006, Griffiths et al., 2021].

From a developmental perspective, sibling relationships are seen as early training grounds for emotional regulation and social interaction. As noted by Feinberg et al. [2012], individuals with siblings may gain interpersonal skills relevant for later romantic relationships. However, recent research challenges the negative stereotype of only children, and highlights positive traits such as creativity, resilience, and adult success [Blake, 1989, Mellor, 1990, Polit et al., 1980, Polit and Falbo, 1987, 1988, Poston Jr and Falbo, 1990].

Several studies have examined how sibling composition influences labor market outcomes, often treating them as indicators of child quality.

Kessler [1991] found that only children had lower employment rates in adolescence but higher rates in their late twenties compared to middle children. Black et al. [2005] showed that having more siblings reduced full-time employment and earnings, especially for women; meanwhile, men's incomes declined, but their employment status remained unaffected. By contrast, Angrist et al. [2010] found no consistent effect of sibling number. Despite these mixed findings, socioeconomic success clearly affects marriage decisions, making it a relevant confounder in analyses of only child status.

In addition, many studies have addressed caregiving responsibilities for aging parents. Research across countries shows that only children bear a heavier caregiving burden than those with siblings [Coward and Dwyer, 1990, Dwyer and Coward, 1991, Spitze and Logan, 1991, Rainer and Siedler, 2012]. Economic theory suggests that siblings may free-ride by avoiding

proximity to parents [Konrad et al., 2002b, Rainer and Siedler, 2009]. Using German microdata, Rainer and Siedler [2009] found that only children are less able to move away from parents, and hence, face reduced labor market opportunities.

Appendix B. Assortativeness

This appendix formally checks the assortativity of only child status. In the main text, we focus on differences in marriage and partner-type probabilities between the two parties (i.e., only and non-only children). However, we use existing indices for robustness checks on the results of positive assortativity obtained from the difference in their marriage probabilities.

Here, we use four standard indices introduced in Chiappori et al. [2025]: odds ratio, likelihood ratio, minimum distance, and correlation. The odds ratio and likelihood ratio indices are interpreted as indicating positive assortativity when their values exceed one. In our context, the odds ratio index indicates the ratio of marriages between only children relative to mixed marriages with respect to sibling composition. Meanwhile, the likelihood ratio index represents the ratio of the probability of positive assortative matching in terms of sibling structure relative to what would occur randomly.

The remaining two indices, minimum distance and correlation, indicate positive assortativity when their values are positive. The minimum distance index represents the weight of the perfectly assortative component, while the correlation index measures the correlation between the wife's and husband's only child status.

All indices show positive assortativity in the pooled data. The values are 1.05 for the odds ratio, 1.01 for the likelihood ratio, 0.09 for the minimum distance, and 0.09 for the correlation. These results suggest a consistent, albeit weak pattern of positive assortative matching in only-child status.

Appendix C. Excluding education from controls

This appendix tests whether our main result holds without controlling for education. As discussed in the main text, only children tend to be more educated, and thus the penalty may no longer appear. Another interesting aspect is how only-child status affects the partner's education overall, particularly after accounting for the positive assortativity of education. Therefore, we follow the same procedure described in the main text and remove one's own education from the control variables. The sibling composition of the marriage partner is indicated by Figure 6, while the educational background of the marriage partner is indicated by Figure 7.

Figure 6 is around here.

Figure 7 is around here.

According to Figure 6, the tendency for only children to remain single and marry another only child persists, even without accounting for education. This provides further support for our main findings, demonstrating the robustness of the observed patterns. Conversely, the estimates related to years of schooling show some variation from the main result. As shown in Figure 7, the spouse type plays a key role here: Only children who marry non-only children tend to have partners with higher educational levels, although the effect size is modest. Conversely, when both spouses are only children, their partner's education tends to be lower. This pattern is statistically significant for women, with an estimate of -0.61, which is very close to the main result of -0.63.

Appendix D. The effects of alternative sibling positions

This subsection examines the effects of alternative sibling positions on their marriage matching outcomes. In our analysis with two-sibling respondents, one's position can be determined by the position of the other sibling. For instance, if one is a male and the first-born child, he is considered the heir if the other sibling is a younger brother, an elder sister, or a younger sister. If one is a female, the first-born daughter without a male sibling is considered the heiress. Thus, we define a dummy variable called "Patrilineal" that equals one for males when the other sibling is a younger brother, an older sister, or a younger sister, and zero if he has an older brother. For females, the dummy variable takes a value of one if the other sibling is a younger sister and zero if the other sibling is an elder brother, a younger brother, or an elder sister. However, if we focus on birth order, then both males and females become heirs if they are first-born children. Hence, we define a dummy variable called "Primogeniture' ' that takes a value of one when the other sibling is a younger brother or sister, and zero if they have an older brother or sister. Finally, we categorize individuals' marital status into three groups: married to an only child, married to a non-only child, and single. This allows us to compare the results with our main findings.

Figure 8 is around here.

Figure 9 is around here.

Figure 8 shows the results for the effects of heirs on their marital status according to the two definitions. The results indicate that under either definition, male heirs are more likely to remain single and less likely to marry only children. In contrast, the results differ for women depending on the definition used. Under the definition of a patrilineal heiress, the probability of remaining single increases. Moreover, while not statistically significant, the likelihood of marrying an only child also slightly increases. However, under the definition of primogeniture, no significant effect is observed on the likelihood of remaining single and the probability of marrying an only child is actually lower.

The difference between these two definitions lies in whether the younger sibling includes a brother. Patrilineal heiresses are also disadvantaged in marriage, similar to those faced by only children. This is likely because they are expected to assume family responsibilities. Conversely, primogeniture heiresses may have a younger brother, allowing for the possibility of sharing or delegating caregiving responsibilities.

In this sense, much like only children, patrilineal heiresses may be expected to assume broader family responsibilities, including caring for older parents within the household or as part of their marital role.

Meanwhile, no significant effects are observed for the education level of the marriage partner under either definition, as shown in Figure 9. However, in the case of heirs and heiresses of both definitions, the negative effect is greater when the marriage partner is an only child, which aligns with our main findings. Overall, penalties associated with each dummy variable are not as pronounced as the effect observed for individuals who are only children.

Tables

Table 1: Descriptive statistics of respondents' characteristics by gender and only-child status

Characteristic	Men		Women	
	Not only child N = 26,945[1]	**Only child** N = 2,032[1]	**Not only child** N = 24,854[1]	**Only child** N = 1,921[1]
MarriageStatus				
Married	18,777 (70%)	1,242 (61%)	18,700 (75%)	1,319 (69%)
Single	8,168 (30%)	790 (39%)	6,154 (25%)	602 (31%)
Age	43 (33, 54)	41 (32, 52)	42 (32, 52)	41 (32, 51)
EducationYear				
0	15 (<0.1%)	2 (<0.1%)	18 (<0.1%)	0 (0%)
6	41 (0.2%)	2 (<0.1%)	23 (<0.1%)	1 (<0.1%)
9	3,277 (12%)	214 (11%)	2,386 (9.6%)	132 (6.9%)
12	10,539 (39%)	705 (35%)	10,578 (43%)	744 (39%)
14	3,658 (14%)	332 (16%)	8,264 (33%)	667 (35%)
16	9,415 (35%)	777 (38%)	3,585 (14%)	377 (20%)
BirthYear	1,958 (1,948, 1,969)	1,961 (1,950, 1,970)	1,960 (1,949, 1,970)	1,963 (1,952, 1,971)
RegionalBlock				
Chugoku	1,682 (6.2%)	129 (6.3%)	1,507 (6.1%)	130 (6.8%)
Hokkaido	1,381 (5.1%)	119 (5.9%)	1,239 (5.0%)	109 (5.7%)
Hokuriku	1,444 (5.4%)	121 (6.0%)	1,332 (5.4%)	83 (4.3%)
Kinki	3,920 (15%)	346 (17%)	3,606 (15%)	318 (17%)
Kyusyu & Okinawa	3,985 (15%)	236 (12%)	3,708 (15%)	242 (13%)
Nouth Kanto & Koushin	2,456 (9.1%)	143 (7.0%)	2,219 (8.9%)	144 (7.5%)
Shikoku	1,080 (4.0%)	65 (3.2%)	1,000 (4.0%)	69 (3.6%)
South Kanto	5,643 (21%)	551 (27%)	5,347 (22%)	481 (25%)
Tohoku	2,342 (8.7%)	120 (5.9%)	2,151 (8.7%)	128 (6.7%)
Tokai	3,012 (11%)	202 (9.9%)	2,745 (11%)	217 (11%)

[1] n (%); Median (Q1, Q3)

Table 2: Observed versus randomly matched marriage patterns

Characteristic	Husband: Not Only Child N = 18,777[1]	Husband: Only Child N = 1,242[1]
Only Child Status with randomization		
Wife: Not Only Child	17,554 (88%)	1,146 (5.7%)
Wife: Only Child	1,223 (6.1%)	96 (0.5%)
Only Child Status		
Wife: Not Only Child	17,649 (88%)	1,051 (5.3%)
Wife: Only Child	1,128 (5.6%)	191 (1.0%)

[1] n (%)

Notes: This table compares the observed marriage patterns with those generated under random matching. The columns indicate whether the husband is an only child, and the rows indicate the same for the wife. Each cell reports the number of couples for each combination, with the proportion relative to the total number of couples shown in parentheses. Random couples are generated by applying a random permutation to the observed married individuals, following the procedure of Chiappori et al. [2018]. Men and women are each assigned IDs drawn from a uniform distribution, ranked accordingly, and then matched based on these ranks.

Table 3: Systematic returns to marriage

Men	Women	JointSurplus	Men's Surplus	Women's Surplus
OnlyChild	OnlyChild	-2.57	-1.42	-1.148
Not OnlyChld	Not OnlyChild	1.82	0.77	1.054
OnlyChild	Not OnlyChild	-1.48	0.285	-1.767
Not OnlyChld	OnlyChild	-1.35	-1.98	0.628

Notes: This table reports the estimated household-level and individual-level returns to marriage by spouse-type combination, based on each partner's only-child status. Marriage types are labeled according to the only-child status of the male and female spouse, respectively. For example, a union between a non-only child male and an only child female is labeled as "Non-Only Child – Only Child Marriage." Each row corresponds to a specific marriage pattern. The three columns report: the total return to marriage in Eq.(2); the male individual return in Eq.(3); and the female individual return in Eq.(4).

Figures

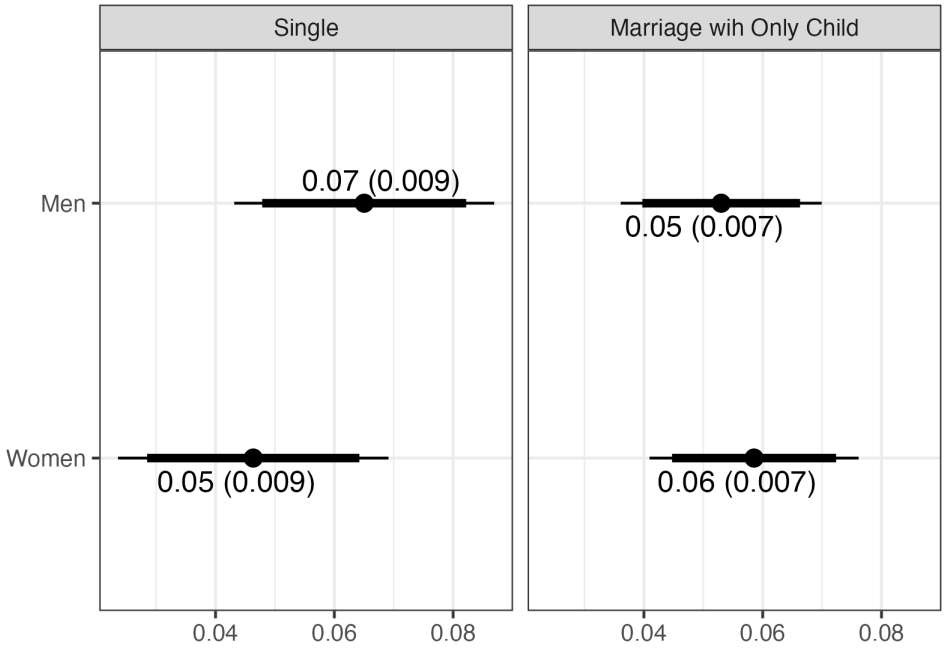

Figure 1: Effects of only child status on partner type

Notes: This figure shows the different marital statuses according to only child status, namely, single (left graph) and married to an only child (right graph), estimated by Eq.(1), along with the 95% confidence intervals; thin lines show Bonferroni-corrected confidence intervals. All nuisance functions are estimated using the stacking method [Wolpert, 1992, Breiman, 1996], which consists of OLS (including squared terms of age, birth year, and years of schooling), random forests [Breiman, 2001], and Bayesian additive regression trees [Chipman et al., 2006, 2010]. Standard errors clustered at the household level are in parentheses. We compare groups of only children and non-only children, and the effects represent the differences in the likelihood of being in each status. Specifically, they indicate the likelihood of being single for only children minus that for non-only children, and the likelihood of marrying another only child for only children minus that for non-only children, respectively.

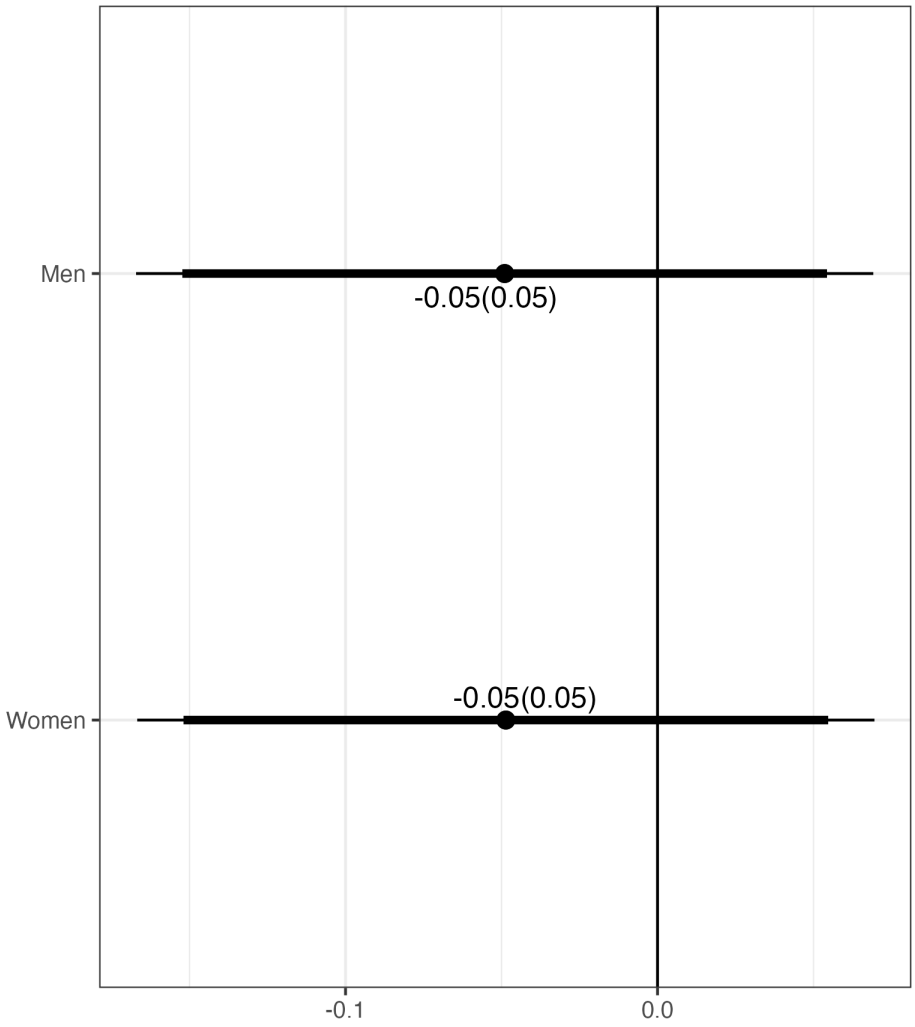

Figure 2: Effects of only-child status on partner's years of schooling (Pooled sample analysis)

Notes: This figure shows the difference in partners' years of schooling according to only child status, estimated by Eq.(5), along with the 95% confidence intervals; thin lines show Bonferroni-corrected confidence intervals. All nuisance functions are estimated using the stacking method [Wolpert, 1992, Breiman, 1996], which consists of OLS (including squared terms of age, birth year, and years of schooling), random forests [Breiman, 2001], and Bayesian additive regression trees [Chipman et al., 2006, 2010]. Standard errors clustered at the household level are in parentheses. We compare groups of only children and non-only children, and the effects represent the differences in their partners' years of schooling after controlling for other variables. Specifically, they indicate the years of schooling of only children's partners minus those of non-only children's partners.

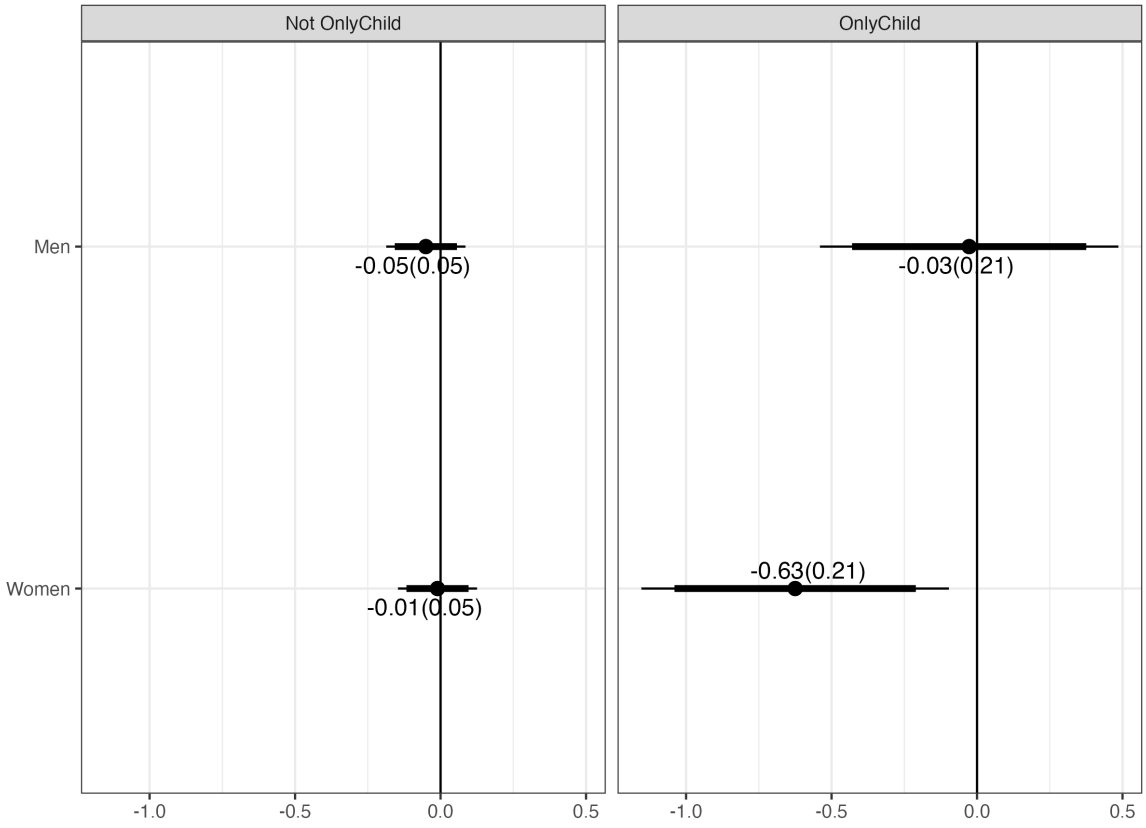

Figure 3: Effects of only-child status on partner's years of schooling (Subsample analysis)

Notes: This figure shows the difference in partners' years of schooling for subsamples defined by partner type, namely, married to a non-only child (left graph) and married to an only child (right graph), estimated by Eq.(5), along with the 95% confidence intervals; thin lines show Bonferroni-corrected confidence intervals. All nuisance functions are estimated using the stacking method [Wolpert, 1992, Breiman, 1996], which consists of OLS (including squared terms of age, birth year, and years of schooling), random forests [Breiman, 2001], and Bayesian additive regression trees [Chipman et al., 2006, 2010]. Standard errors clustered at the household level are in parentheses. We compare groups of only children and non-only children, and the effects represent the differences in their partners' years of schooling after controlling for other variables. Specifically, the effects are calculated as the years of schooling of only children's partners minus those of non-only children's partners.

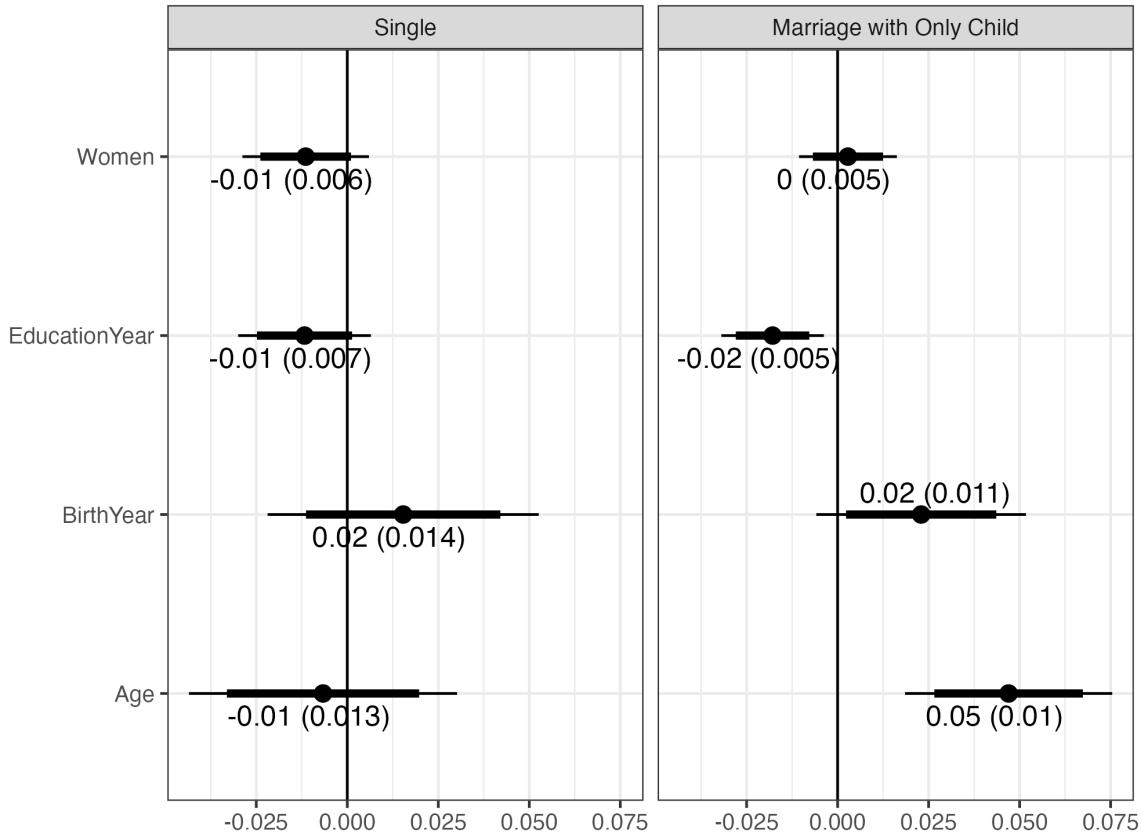

Figure 4: Effects of only child status on partner type (Heterogeneity)

Notes: This figure shows the best linear projection of the conditional difference in marital outcomes by only-child status, single (left graph) and married to an only child partner (right graph), along with the 95% confidence intervals; thin lines show Bonferroni-corrected confidence intervals. All nuisance functions are estimated using the stacking method [Wolpert, 1992, Breiman, 1996], which consists of OLS (including squared terms of age, birth year, and years of schooling), random forests [Breiman, 2001], and Bayesian additive regression trees [Chipman et al., 2006, 2010]. Standard errors clustered at the household level are in parentheses. We compare groups of only children and non-only children. Each variable is interpreted as follows. The coefficient for women is relative to men (the omitted category). Years of schooling, birth year, and age estimates represent the correlation between the dependent variable and a one-standard-deviation change from the mean of these variables, respectively.

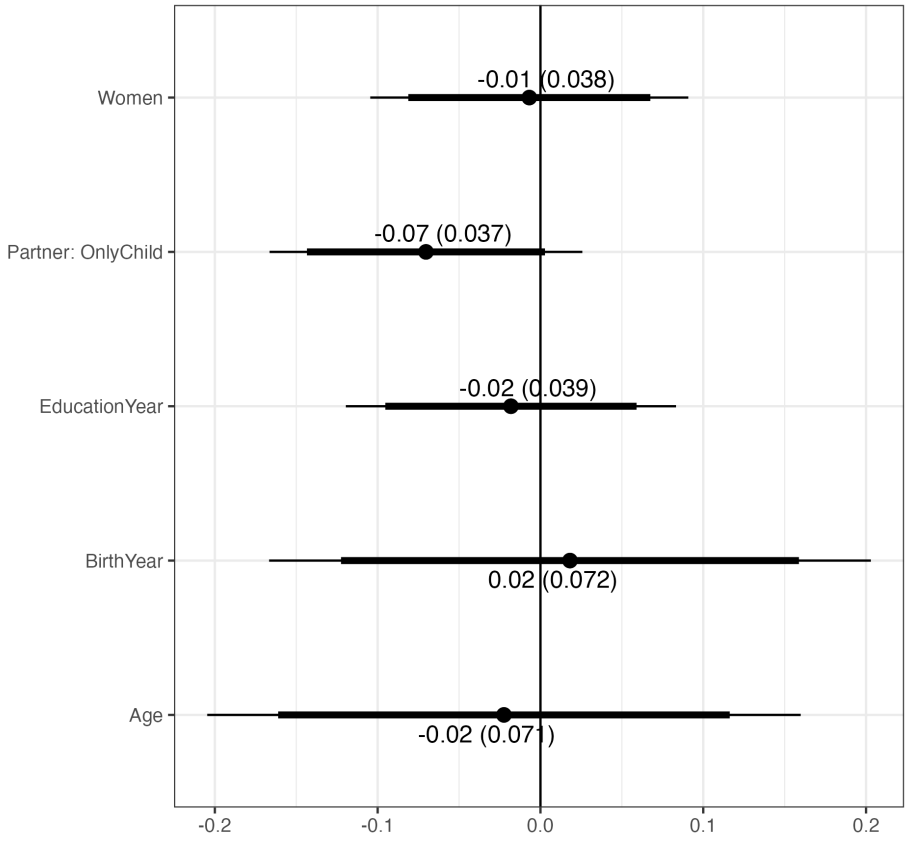

Figure 5: Effects of only child status on partners' years of schooling (Heterogeneity)

Notes: This figure shows the best linear projection of the conditional difference in partners' years of schooling, along with 95% confidence intervals; thin lines show Bonferroni-corrected confidence intervals. All nuisance functions are estimated using the stacking method [Wolpert, 1992, Breiman, 1996], which consists of OLS (including the squared terms for age, birth year, and years of schooling), random forests [Breiman, 2001], and Bayesian additive regression trees [Chipman et al., 2006, 2010]. Standard errors clustered at the household level are in parentheses. We compare groups of only children and non-only children. Each variable is interpreted as follows. The coefficient for women is relative to men (the omitted category). For the partner variable, the coefficient for only children is relative to marriages with partners who are non-only children. Years of schooling, birth year, and age estimates represent the correlation between the dependent variable and a one-standard-deviation change from the mean of these variables, respectively.

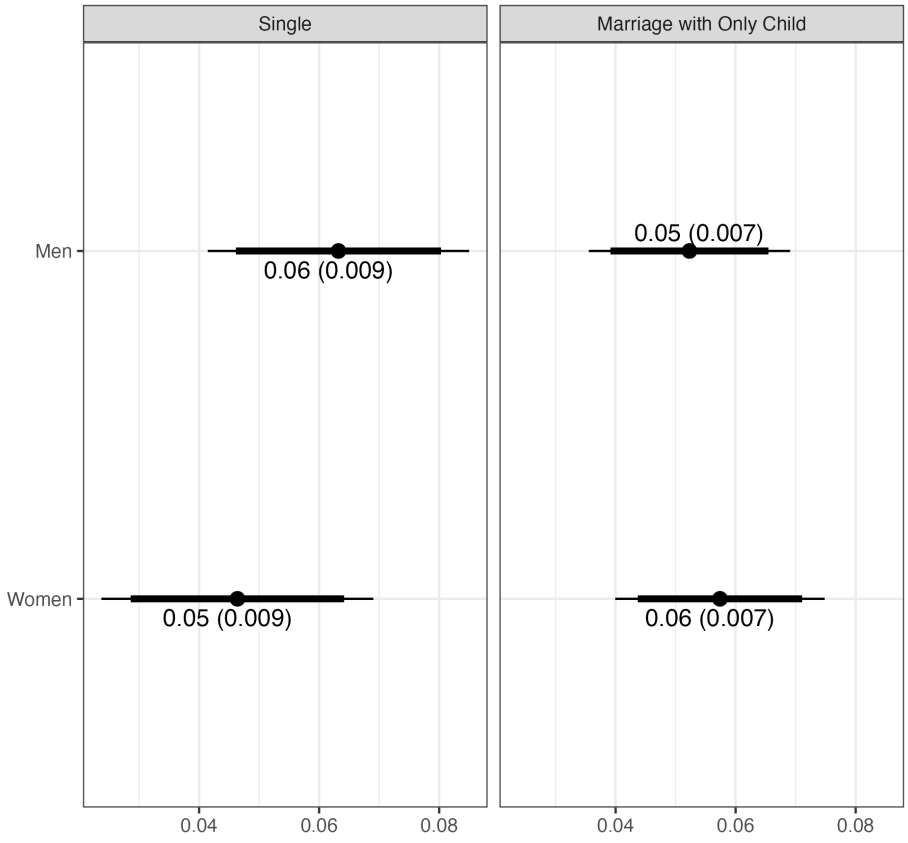

Figure 6: Effects of only child status on partner type (without controlling for education)

Notes: This figure shows the different marital statuses according to only child status, namely, single (left graph) and married to an only child (right graph), estimated by Eq.(1), along with the 95% confidence intervals; thin lines show Bonferroni-corrected confidence intervals. All nuisance functions are estimated using the stacking method [Wolpert, 1992, Breiman, 1996], which consists of OLS (including the squared terms of age and birth year but NOT years of schooling), random forests [Breiman, 2001], and Bayesian additive regression trees [Chipman et al., 2006, 2010]. Standard errors clustered at the household level are in parentheses. We compare groups of only children and non-only children, and the effects represent the differences in the likelihood of being in each status. Specifically, they indicate the only children's likelihood of being single minus the non-only children's likelihood of being single, and the only children's likelihood of being married to another only child minus that of non-only children, respectively.

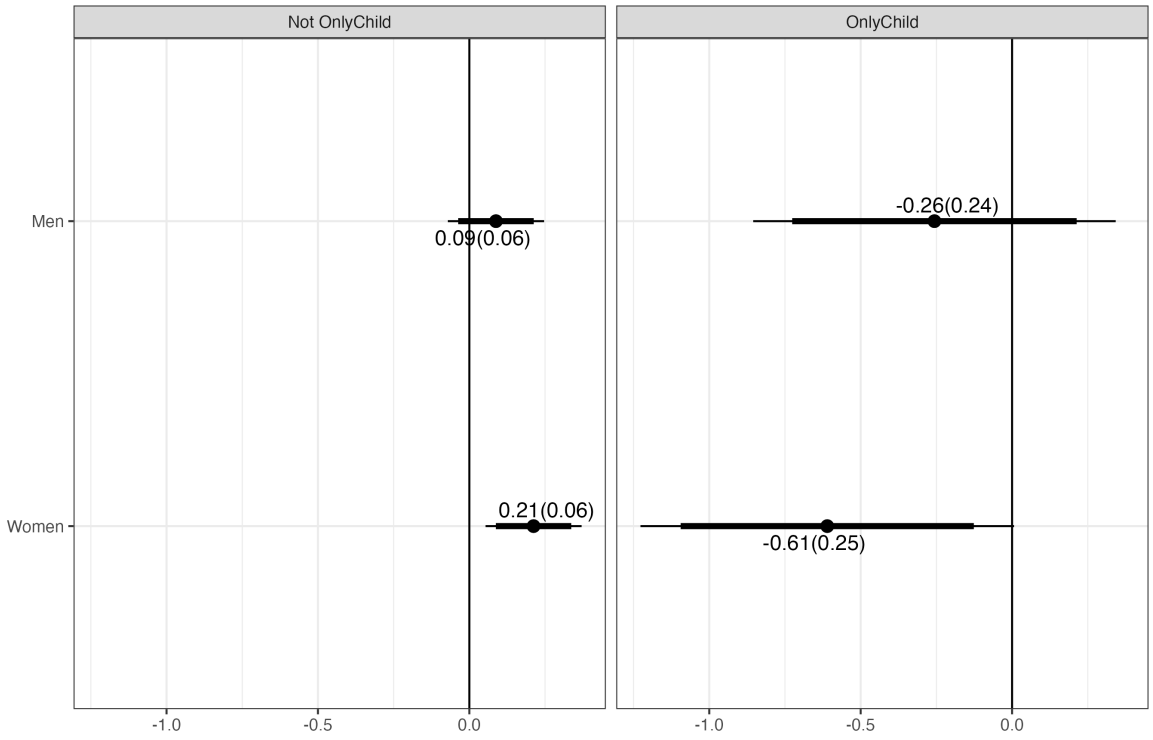

Figure 7: Effects of only-child status on partner's years of schooling (without controlling for education)

Notes: This figure shows the difference in partners' years of schooling according to only child status, for subsamples defined by partner type, namely, married to non-only children (left graph) and married to only children (right graph), estimated by Eq.(5), along with the 95% confidence intervals; thin lines show Bonferroni-corrected confidence intervals. All nuisance functions are estimated using the stacking method [Wolpert, 1992, Breiman, 1996], which consists of OLS (including the squared terms of age and birth year, but not years of schooling), random forests [Breiman, 2001], and Bayesian additive regression trees [Chipman et al., 2006, 2010]. Standard errors clustered at the household level are in parentheses. We compare groups of only children and non-only children. The effects represent the differences in their partners' years of schooling after controlling for other variables. Specifically, they indicate the years of schooling of only children's partners minus those of non-only children's partners.

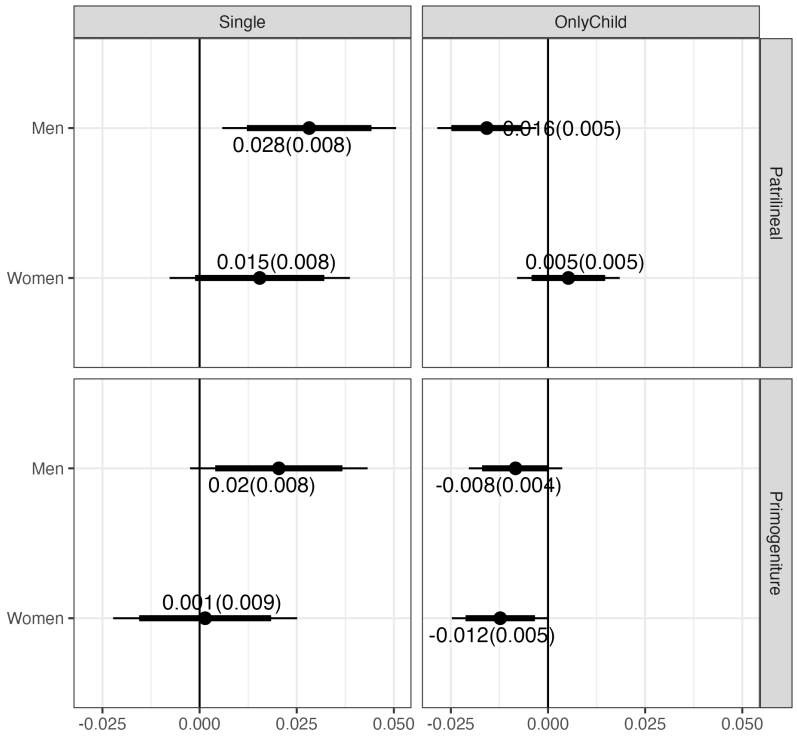

Figure 8: Effects of heir status on partner type

Notes: This figure shows the different marital statuses according to sibling positions, namely, single (left graph) and married to an only child (right graph), along with the 95% confidence intervals; thin lines show Bonferroni-corrected confidence intervals. The sibling positions are defined by dummy variables called "Patrilineal" (top graph) and "Primogeniture" (bottom graph) among respondents with one sibling. The patrilineal variable equals one for males when the other sibling is a younger brother, an older sister, or a younger sister, and zero if he has an older brother. For females, the dummy equals one if the other sibling is a younger sister and zero if the other sibling is an older brother, a younger brother, or an older sister. Primogeniture equals one when the other sibling is a younger brother or sister, and zero if they have an older brother or sister for both males and females. All nuisance functions are estimated using the stacking method [Wolpert, 1992, Breiman, 1996], which consists of OLS (including the squared terms for age, birth year, and years of schooling), random forests [Breiman, 2001], and Bayesian additive regression trees [Chipman et al., 2006, 2010]. Standard errors clustered at the household level are in parentheses. We compare groups of heirs and non-heirs for each definition. These effects represent the differences in the likelihood of being in each status. Specifically, they indicate the likelihood of being single for heirs minus that for non-heirs, and the likelihood of marrying an only child for heirs minus that for non-heirs, respectively.

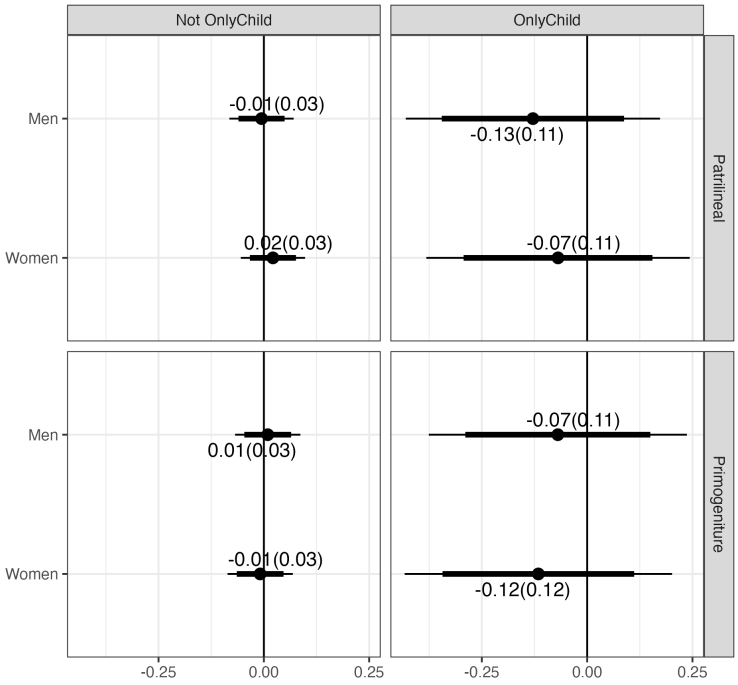

Figure 9: Effects of heir status on partner's years of schooling (Subsample analysis)

Notes: This figure shows the difference in the partners' years of schooling according to alternative sibling position for subsamples defined by partner type. Specifically, it shows the results among marriages with a non-only child partner (left graph) and marriages with an only child partner (right graph), along with the 95% confidence intervals; thin lines show Bonferroni-corrected confidence intervals. The sibling positions are defined by dummy variables called "Patrilineal" (top graph) and "Primogeniture" (bottom graph) among respondents with one sibling. The patrilineal variable equals one for males when the other sibling is a younger brother, an older sister, or a younger sister, and zero if he has an older brother. For females, the dummy variable equals one if the other sibling is a younger sister and zero if the other sibling is an older brother, a younger brother, or an older sister. Primogeniture equals one when the other sibling is a younger brother or sister, and zero if they have an older brother or sister for both males and females. All nuisance functions are estimated using the stacking method [Wolpert, 1992, Breiman, 1996], which consists of OLS (including squared terms of age, birth year, and years of schooling), random forests [Breiman, 2001], and Bayesian additive regression trees [Chipman et al., 2006, 2010]. Standard errors clustered at the household level are in parentheses. We compare groups of heirs and non-heirs in each definition, and the effects represent the differences in their partners' years of schooling after controlling for other variables. Specifically, they indicate the years of schooling of heirs' partners minus those of non-heirs' partners.

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