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Parental Loss in Early Years and Adult Family Formation: Evidence From U.S. Cohorts Born 1850–1910

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ABSTRACT

Objective: This paper examines the long-term impact of parental loss in early years on family formation as adults in U.S. cohorts born between 1850 and 1910, focusing on age assortative mating and the timing of first childbirth.

Background: Early parental loss can profoundly shape children's adult family formation trajectories. However, its impact remains unclear due to multiple mechanisms—such as social control and economic constraints—as well as selection into parental loss and stratified family formation patterns. Empirical evidence on the long-term effects of parental loss during young ages remains limited.

Method: The analyses leverage two separate data sources. Crowdsourced genealogical records are used to investigate the relationship between parental death and family formation. To address selection bias, the analysis employs family fixed effects, comparing siblings who experienced parental loss before or after reaching adulthood. Second, historical census data analyzes socio-demographic stratification of parental loss and family formation patterns.

Results: Initial OLS estimates link parental loss to increased age heterogamy and earlier childbirth, but fixed-effects models show these associations weaken or reverse. Census data confirm selection effects, revealing that disadvantaged individuals are more likely to experience parentless households during childhood, enter age-heterogamous unions, and start families earlier.

Conclusion: Parental loss in early years correlates with family formation in adulthood, but this appears driven by household selection. Despite data-related limitations, results underscore the role of selection in studying the consequences of parental loss and family structure.

1 | Introduction

The death of one or both parents in early years is a profoundly transformative event for children, influencing not only immediate outcomes but also shaping entire life trajectories. Researchers have explored the impact of early parental loss across a broad spectrum of short- and long-term outcomes, including health and mortality (Quanjer et al. 2023; Beekink et al. 1999), physical and cognitive development, overall well-being (Björkenstam et al. 2017; Hepworth et al. 1984), and social and economic mobility (Dribe et al. 2022; Adda et al. 2011). Notably, these effects have been documented in both historical (Quanjer et al. 2023; Dribe et al. 2022; Rosenbaum-Feldbrügge 2019) and contemporary contexts (Amato and Anthony 2014; Adda et al. 2011; Li et al. 2014).

Building on this foundation, the present article focuses specifically on how early parental loss influences family formation outcomes, of their children examining two key dimensions: partner selection—through age assortative mating—and the timing of family formation, specifically the age at first birth, looking at individuals born in the United States between 1850 and 1910.

The article first explores whether a relationship existed between the loss of a parent during younger ages and these two processes. It also investigates whether heterogeneity in this relationship exists across genders of both the lost parent and the child. Finally, it tries to disentangle the extent to which the associations are explained by the resources and disadvantaged conditions of families experiencing parental loss.

Unlike other consequences of parental loss, where negative outcomes are often predictable, the effects on family formation are more ambiguous due to the interplay of different mechanisms. On one hand, losing a parent may introduce economic constraints, limiting the resources available for pursuing personal preferences and romantic consideration in partner selection and family formation, forcing individuals to prioritize pragmatic material considerations such as financial stability and maintenance. On the other hand, the absence of parental authority can reduce parental control and free resources, potentially allowing individuals to make choices aligned with romantic and personal preferences.

As a result, the relationship between parental loss during early years and family formation patterns remains ambiguous and empirically unexplored. It is unclear whether early parental loss favors selecting a partner of a similar age with shared experiences and common life stages or encourages choosing an older, financially more stable partner. Similarly, it is unclear whether losing a parent in early ages accelerates family formation due to increased autonomy in setting up a household or the need to secure support from a partner, or it may delay it due to financial constraints and difficulties in setting up a new independent household. Moreover, these relationships are highly gendered, both in terms of the child's and the parent's gender. Men and women may respond differently to parental absence, economic pressure, and social expectations, while the loss of a father or mother may have distinct emotional and financial effects.

This ambiguity is particularly relevant in pre-second demographic transition contexts, such as Western societies between the 19th and 20th centuries, where strong familial bonds and significant family control heavily shaped children's life choices. Until the early twentieth century and the advent of industrialization, decisions about marriage and careers were influenced by family needs and objectives other than individual priorities. Partner selection and marriage timing were driven by practical concerns like property transfer, job security, and support for aging parents. Unlike today's emphasis on emotional bonds, family ties back then also prioritized economic stability and mutual obligations. All social classes shared this collective approach, valuing family members for their contributions to household stability and support during times of need (South 2001; Smith 1979).

Another crucial aspect when studying the relationship between parental loss and family formation processes is the likely high level of stratification in the two processes, possibly confounding the actual relationships. A social gradient in mortality in the U.S. was already evident in the late eighteenth and early nineteenth centuries. Lower-class and otherwise disadvantaged individuals often lived in unsanitary conditions, had less healthy lifestyles and poorer nutrition, and were employed in more hazardous occupations, all of which increased their mortality risk at all ages (Krieger and Fee 1996) and, consequently, their children's risk of parental loss. An association between parental loss and social background can therefore be reasonably inferred from patterns of socioeconomic, geographic, and racial mortality. Family formation patterns were also highly stratified by socioeconomic conditions, as individuals from lower socioeconomic backgrounds were more constrained by economic necessity, typically leading to lower age homogamy and faster family formation (South 2001).

Moreover, in the period we study, families across the socioeconomic spectrum faced distinct institutional and economic environments that shaped both their vulnerability to early mortality and the forms of oversight they exercised over children's behavior. Higher-status parents, with more stable employment, greater material resources, and stronger integration into community and religious institutions, were generally better positioned to monitor children's activities and enforce expectations around courtship, marriage timing, and household formation. In contrast, working-class and poor families, whose livelihoods were tied to unstable industrial or migratory labor, often experienced weaker forms of supervision due to long or irregular work hours, economic pressures that pushed adolescents into the labor market, and living arrangements that diluted parental oversight. These same conditions also heightened the risk of premature adult mortality.

We therefore compare the age assortative mating and timing of family formation between siblings, who experienced parental loss before or after reaching adulthood. The resulting estimate of parental loss removes the across-household selection component, which stems from varying propensities for parental death in households financially constrained or in adverse social conditions. Additionally, this within-household estimate accounts for unobserved family background characteristics shared between siblings. This method helps identify selection effects across households. Comparing the within-household estimate with the naïve OLS estimate highlights the direction of selection into parental loss at the household level (Abramitzky et al. 2012).

We leverage two separate sources of data. First, we use online FamiLinx crowdsourced genealogical data to establish the relationship between parental death and patterns of family formation. FamiLinx has the main advantage of including intergenerational information on parents and children, such as the death of the parent and the family formation timing of the child, and information on siblings. Second, we adopt US Census data to investigate social stratification in parentless households and selection into parental loss, that is, what social characteristics were associated with parentless households and family formation patterns.

Given the historical context and the complexity of the mechanisms linking parental loss to subsequent family outcomes, we do not begin the study with a specific directional hypothesis. In the theoretical and review sections of the article, we outline the potential mechanisms and logics that may suggest a specific directionality in the association. Theoretically, multiple and even opposite responses are plausible: parental loss

could hasten or delay union formation, widen or narrow age gaps between partners, and either accelerate or postpone the transition to first parenthood. Similarly, heterogeneous patterns by sex and age at bereavement may emerge in different directions. Because all these outcomes are theoretically defensible, the study adopts an exploratory approach aimed at describing empirical regularities rather than testing a predetermined expectation.

2 | Theoretical Framework: Social Control and Economic Incentives

Parents play a crucial role in shaping their children's life course decisions and transitions, either directly by influencing choices such as education and marriage or indirectly by instilling ideals, aspirations, and values. Thus, the loss of a parent during formative years, such as childhood or adolescence, might have a short-term impact on the children's cognitive, emotional, and social development, but also a long-term effect on their adult choices, including family formation trajectories (Kiernan 1992; Reneflot 2011; Dahlberg 2020; Beaujouan and Solaz 2023).

The impact of parental death on children's family trajectories can be understood through at least two key mechanisms. First, the loss of a parent during early childhood may free children from parental control in adult life decisions, allowing them to pursue their interests and preferences—a perspective we refer to as the *social control mechanism*. Second, the absence of parental support during young ages may alter children's economic incentives, likely imposing constraints on their life choices by reducing household income during formative years—a channel we term the *economic constraint mechanism*. We discuss how these two perspectives relate to age homogamy and the timing of family formation.

2.1 | Parental Loss and Partner Selection

In Western societies during the 19th and early 20th centuries, marriage was still a family matter driven by economic and instrumental considerations, aimed at ensuring the family's interests. Parents had a significant role in selecting suitable spouses for their children (Kalmijn 1998; van Leeuwen and Maas 2005), and the shock of the death of one (or both) could significantly affect such choices.

In partner selection, age homogamy has often been used as an indicator of the prevalence of romantic and emotional factors over practical considerations (Shorter 1975; van de Putte et al. 2009). Belonging to the same age cohort can enhance marital quality by facilitating shared values, attitudes toward marriage and family, preferences, and life experiences (Skopek et al. 2011). As a result, observed mating-age preferences typically reflect narrow interpartner age differentials, with a consistent directional bias favoring unions where the male is older than the female. Parents, however, might prioritize the economic security a partner could provide, even if it entails a larger age gap. In the past, this concern was likely stronger for daughters, who were less likely to work outside the household and more economically dependent on their family of origin. Moreover, women were often regarded as central to maintaining family status, which made them more vulnerable to parental pressure in marriage decisions compared to their brothers (Sassler 2005).

According to the *social control mechanism*, the loss of one or both parents may increase freedom in the partner selection process. If the preference for a partner of similar age holds, orphaned individuals should therefore have greater freedom to choose a partner according to their own age preferences. Thus, we can anticipate higher levels of age homogamy among children who experienced the loss of either parent, as both fathers and mothers played roles—albeit distinct ones—in shaping marital decision-making. The mechanisms likely differed, however: in patriarchal contexts, fathers' authority over household resources and marital arrangements positioned them as primary gatekeepers of age-appropriate matches, whereas mothers exerted influence mainly through their caregiving role and their interest in ensuring their children's well-being. These patterns are likely to differ by the child's gender, with stronger age-homogamous tendencies anticipated among daughters than sons. Historically, parents exercised closer supervision over daughters' marital choices, influenced by women's more limited economic opportunities, their greater social and familial accountability, and the expectation that marriage would be their primary route to security. Sons, by contrast, typically had greater autonomy in partner selection due to higher anticipated economic independence and broader access to formal employment.

The predictions derived from the economic constraints perspective are more complex. If the shock of parental death reduces a family's economic resources, children may be encouraged to seek a partner who can provide greater financial security, resulting in higher age heterogamy, as older individuals are often more economically established. This mechanism is particularly relevant in cases of father's loss, since men were typically the primary breadwinners of the household (van Poppel et al. 1998). Such a loss could be especially detrimental for daughters, who were less likely to work outside the parental home and had lower economic independence than sons. In this context, an increase in age differences between partners can be expected.

However, it should also be considered that the death of a parent might release economic resources through inheritance, potentially reducing financial constraints in partner selection and, in turn, diminishing the need to prioritize economic stability over other factors.

Finally, individuals who lost both parents at an early age represent a particularly disadvantaged group, as the complete loss of parental figures implies the disappearance of both emotional guidance and material support. For them, the economic constraints mechanism is expected to be especially salient, as the absence of parental resources may have substantially limited their opportunities and heightened their dependence on marriage for economic security. In this group, both men and women are likely to experience similar pressures, as the loss of both parents removes the differential influence typically exerted by mothers and fathers.

To sum up, parental loss can influence partner selection through several channels in more than one plausible way. According to a social control mechanism (a1), losing a parent increases autonomy and can lead to greater age homogamy. In contrast, an economic constraint mechanism (b1.1) suggests that parental loss creates financial and material strain, especially through the loss of the main provider, pushing individuals toward older, economically more secure partners, thereby reducing age homogamy. On the other hand, if parental loss increases available resources through inheritance (b1.2), it can increase freedom in partner choice and thus raise age homogamy. Finally, both mechanisms imply heterogeneity across genders (c1), suggesting that these associations are likely to be stronger for daughters than for sons. Historically, parents exercised more direct oversight of daughters' marital trajectories, and daughters generally had lower economic and social independence.

2.2 | Parental Loss and Timing of Family Formation

In pre-industrial societies, family obligations often took precedence over individual desires, leading to delays in marriage despite children's inclination toward earlier unions. Children were often expected to postpone or forgo marriage to support younger siblings, care for elderly parents, or maintain family stability (Hareven 1977). Parental control played a significant role in regulating marriage timing, especially to safeguard the transmission of property and wealth. Moreover, beyond pursuing the maximization of family interests, parents may want to postpone children's family formation to ensure children's best interests. Parents, especially among wealthier families, are likely to have high educational and career aspirations for their children that conflict with early marriage (Axinn and Thornton 1992).

From these premises and building on the social control hypothesis, the death of a parent, especially a father, could release children from such control and supervision, granting them the freedom to marry earlier and set up their household (Smith 1979).

Parental social control was also exerted through their management of the family economic resources, and the loss of a parent often granted children greater freedom, partly through the inheritance of those resources. In his seminal work, Hajnal (1965) argues that the death of a parent accelerates entry into marriage, and consequently also the age at first child, by allowing the children greater control over family estates and therefore greater freedom and material abilities to set up a new household earlier on. Indeed, the loss of a parent before or around marriable age is associated with a faster transition to first marriage for both men and women in various historical and contemporary European contexts (Voland and Willführ 2017; Kiernan 1992; Shenk and Scelza 2012; Hepworth et al. 1984; Rosenbaum-Feldbrügge and Debiasi 2019; Rosenbaum-Feldbrügge 2022).

Hajnal hypothesis, however, mostly refers to parental loss after childhood, when children have already reached independence, and the death of a parent likely resulted in immediate inheritances. When parental loss takes place before adulthood, it more likely results in a longer period lived in a deprived household and lower wealth accumulation (Oris and Ochiai 2002; Rosenbaum-Feldbrügge 2018). In this case, the lack of resources can make partnering and the establishment of a new household more problematic, possibly postponing family formation. Additionally, surviving children might have been expected to provide financial and caregiving support to a widowed parent, further hindering their ability to establish an independent family (Janssens 2002; Lundh and Kurosu 2014).

On the other hand, economic constraints could also result in anticipation if individuals were pushed to establish a new household and find a partner to seek material support. This was likely true for orphaned women who had less access to formal labor markets and independent sources of income.

Therefore, theoretically, multiple hypotheses are plausible. Existing empirical research, however, has most frequently found an anticipation effect of parental loss on the transition into marriage and parenthood across many contexts and periods (Pettay et al. 2020; Derosas et al. 2014; Lundh and Kurosu 2014; Hepworth et al. 1984; Kiernan 1992; Reneflot 2011; Shenk and Scelza 2012; Reneflot (2011), Kiernan 1992; Kiernan and Hobcraft 1997). Parental death was also associated with a propensity to leave parental house earlier, likely due to conflict with stepparents or due to a decline in household income and inability to support all family members (Bras and Neven 2007; Lundh and Öberg 2018; Steckel 1996; Kiernan 1992; Shenk and Scelza 2012).

A gendered pattern in the timing of family formation is also likely. In pre-industrial settings, daughters' lower economic and social independence made their transitions to marriage more closely tied to parental oversight, especially fathers' control over household resources. When a parent died, daughters often faced more constrained prospects within the parental home and fewer paths to self-support, making earlier marriage a more viable route to stability. Sons, with greater autonomy and broader access to work, were less dependent on parental structures when forming unions. Consequently, anticipatory shifts in marriage timing following parental loss are expected to be more pronounced for women.

To sum up, parental death during childhood can shape the timing of family formation through several distinct pathways, each pointing to different potential outcomes. Under a social control mechanism (a2), the loss of a parent reduces direct oversight allowing earlier transition into marriage. Economic hardship offers a more complex picture. When parental death results in scarce resources (b2.1), young people may find it difficult to accumulate the means needed to form an independent household or the need to support a surviving parent and family members, delaying marriage. At the same time, limited economic options may prompt some to marry earlier in search of material security (b2.2). Finally, these mechanisms are not neutral with respect to gender (c2), with daughters likely more affected due to both higher financial dependence and stronger familial control.

3 | Data and Sample

3.1 | Investigating the Relationship Between Parental Loss and Family Formation With Familinx

To establish the association between parental loss and family formation behaviors we use FamiLinx, a publicly available collection of crowdsourced genealogies downloaded from the website Geni.com (Kaplanis et al. 2018) (data are available at https://osf.io/fd25c/). Geni.com is a website where users can freely reconstruct family trees, including sociodemographic information for every individual included. In the case of similarities, the website offers the option to merge profiles and family trees. Data from Geni.com have been downloaded, cleaned for consistency, and organized into FamiLinx by Kaplanis et al. (2018). The final dataset contains approximately 86 million publicly available profiles that cover most of Western countries concentrated between the 15th and 20th centuries.

In the original study, FamiLinx data have been compared against several data sources by Kaplanis et al. (2018) showing good coverage of historical events and trends in Western countries and especially the U.S. Other studies have compared FamiLinx to various official statistics showing that, despite biases, they are representative of long-term trends in demographic outcomes and sociodemographic processes across Western countries (Blanc 2020; Cozzani et al. 2023; Colasurdo and Omenti 2024; Minardi et al. 2024; Stelter and Alburez-Gutierrez 2022).

Biases include the distortion typical of genealogical data, such as the overrepresentation of longer-lived individuals, vertical ties, and the underrepresentation of childless individuals and women. Moreover, previous studies suggested that records are positively selected based on social status and more representative of elites (Stelter and Alburez-Gutierrez 2022).

Despite limitations, genealogy data allows the reconstruction of intergenerational relationships and interconnected events across multiple generations. Therefore, they are among the very few data sources allowing to study multigenerational processes in historical periods and have been used to study topics such as mortality (Cozzani et al. 2023), fertility (Hsu et al. 2021; Gay et al. 2023; Blanc 2020), and assortative mating (Corti et al. 2024). In our case, we can link parents' years of death and the timing and characteristics of children's family formation events and know at which of their children's ages parents died. Genealogy data allows the reconstruction of intergenerational relationships and interconnected events across multiple generations. Unlike other historical sources, genealogies enable the identification of family members across generational lines and provide insights into the number and types of kin present during a specific period. In our case, we can link parents' years of death and the timing and characteristics of children's family formation events and know at which of their children's ages parents died.

With regards to our research question, the overrepresentation of higher status individuals should lead, if anything, to more conservative estimates, since wealthier families had more resources to adjust socially and economically to shocks induced by the loss of a parent. Evidence from the censuses helps to confirm this idea, highlighting that parental loss and family patterns associated with parental loss were less common in higher nonmanual classes.

Our baseline sample includes US-born individuals between 1850 and 1910 ($N = 2{,}291{,}538$). Cohorts born in those years experienced an unprecedented period of economic and social modernization, which reshaped family relationships. Economic changes such as the rise of a wage-based economy and migration flows toward urban areas weakened parent–child relations (Treiman 1970). Moreover, personal and romantic relationships significantly changed as a response to social modernization, with personal criteria and affection becoming more relevant than pragmatic considerations when choosing a partner (Shorter 1975). The precise cohort limits were also guided by practical considerations. First, these are birth cohorts whose family transitions and marriages can be observed in historical micro-censuses—beginning with the first available in 1850 (Ruggles et al. 2025)—which allows meaningful comparison with other reliable data sources. Second, we set the upper bound at 1910 to avoid including individuals who were likely still alive at the time of the 2011 data extraction (Kaplanis et al. 2018). Because demographic and genealogical information for living persons is typically sparse or imprecise, this cutoff maximizes usable sample size while minimizing missing or unreliable information.

Birth location in the Familinx data is provided as unstructured strings and as latitude and longitude. We obtain categorical information on country and state of birth through text similarity and automated reversed geo-parsing algorithms. We excluded individuals with missing or implausible information relevant for the analyses: we dropped individuals without their parents' death year ($N = 610{,}640$), and those for whom we could not identify at least one partner ($N = 192{,}962$). Moreover, we excluded individuals with parents of the same sex ($N = 188{,}071$), born after the mother's and/or father's death ($N = 123{,}544$). To study age assortative mating, we select individuals with a partner and their birth year, with no more than 50 years of difference. This first sample includes 101,631 individuals. We further select a sample including individuals with observed siblings, resulting in 53,749 cases.

To analyze the age at first birth, we further select individuals with at least one child and the first child's birth year, thus leading to a second sample of 55,669 individuals. Similarly to age assortative mating, the siblings' sample includes 28,238 individuals.

We calculate the variable of parental death (for both mothers and fathers) using the child's year of birth and the parent's year of death. By calculating the age of parental death as the difference between the parent's year of death and the child's birth year, we can identify the timing of the parental death relative to the child's age (we selected plausible values that cover a range from 0 to 100). We create a binary variable that takes the value of 1 if the child was younger than 16 years old at the time of parental death and 0 otherwise; we build this variable for mother's and father's death, as well as for the death of both parents. The choice of the threshold at age 16 aligns with previous research indicating that the effects of parental death differ significantly before and after the age of 16 (Rosenbaum-Feldbrügge 2019). Furthermore, considering the rarity of marriages before age 16

(Haines 1996), this operationalization offers a meaningful distinction for our analyses.

We operationalize age heterogamy as the absolute value of the age difference between the birth years of the male and female partners (Kolk 2015). We use additional operationalizations of age assortative mating; we build a dummy variable indicating whether the union is age homogamous or not (+3 years difference between the partners), and we also use a categorical variable indicating whether the union is age homogamous (0–3 years difference), hypergamous (man 3+ older) or hypogamous (woman 3+ years older). We chose this operationalization considering that the mean age difference in the period in our data is around 3 years. Moreover, we consider individuals within this age gap as sharing the same life-course phases, values, aspirations, and experiences. Results for this operationalization of age homogamy are shown in Table A2 of the Supporting Information Appendix. Given that the previous studies used various thresholds, we performed further checks changing the age cutoff at 5 and 2 years, and results do not change notably (see Tables A11 and A12 in the Supporting Information Appendix).

As for the age of the first child, we compute the difference between the year of birth of the first child and the individual's year of birth.

We consider a set of control variables related to family structure available in the data. Factors such as the number of siblings could influence the urgency to marry, while the presence of stepsiblings accounts for the eventuality of the surviving parent's remarriage. The sex ratio of siblings (brothers/sisters) indicates the gender composition of the family, while the birth order in the family of origin is included to account for differences between firstborns and following siblings, since they usually have different experiences and hereditary arrangements after the loss of a parent.

We employ linear regression models to examine the relationship between the absolute age difference between an individual and parental death. In the model, we include the dummy variable on the mother's death before age 16, the father's, and both parents. We run the same model with the age at the first child as a dependent variable.

We adopt a stepwise approach by first running empty models and then adding the variables of family structure. We run separate models by gender, as the consequences of parental loss and the control variables vary significantly among daughters and sons. This means that in the siblings' samples, we consider same-sex siblings' dyads. We further account for the sex composition of the family with the variable of sex ratio.

As a second step, we use sibling fixed-effect models to account for confounders in the relationship between parental loss and family formation outcomes. In these models, just siblings belonging to different age groups (before and after 16 in our case) at their parents' deaths contribute to the estimation of the effects. One of the main advantages of these models is that they allow for control of within-family characteristics that might act as confounders in these models. However, they significantly reduce the sample size in its numerosity and composition. In our case, we reduce the sample to families in which siblings have, as a mechanical consequence, a higher age difference than the complete sample and are thus selected according to certain characteristics.

3.2 | Investigating Selection Into Parentless Household and the Stratification of Family Formation Patterns Using the Censuses

The adoption of siblings' fixed effects is justified by the idea that both parental loss and patterns of family formation are socially stratified. However, the degree of stratification and the direction of the association between social status and other household characteristics and family formation are not straightforward. Using historical census data, we investigate what family-level factors are associated with both parentless households and family formation patterns.

As stratifying variables, we consider the household highest occupational class, ethnic/racial minority, macro-region of origin, migration background, nationality of origin, and urban or rural residence. While FamiLinx allows for reconstruction information on parental loss and family formation for the same individuals, it does not contain information on socioeconomic characteristics. Conversely, the census contains detailed information on socioeconomic characteristics but does not include information on parental loss; it must be proxied by parentless household and does not allow connecting childhood information with adult outcomes, except in some cases through record linkage.

We therefore separately analyze, on the one hand, the correlates of fatherless, motherless, and parentless households among children, and, on the other, the drivers of age assortative mating and age at first children among adults.

To investigate the social stratification of parentless households we select all individuals aged 16 or less and born between 1850 and 1910 from the 1880 to 1940 samples of the US historical censuses. We use three indicators of parental loss: widowed mother, widowed father, and parentless household. To ensure consistent identification of parental loss, the analysis focuses on the censuses from 1880 to 1920, the first to report marital status. Children are classified as having lost a father or mother when they reside with a surviving parent recorded as widowed or when the co-resident parental figure is identified as step or adoptive. The identification of step or adoptive parents is performed by IPUMS using direct and indirect information (Ruggles et al. 2025). Given that divorce was extremely rare during this period (0.84% among ever-married adults aged 18–45), the presence of a step- or adoptive parent is treated as indicative of the death of the corresponding biological parent. Potential misclassification arising from absent rather than deceased parents is likely limited since among children aged 16 or younger in the censuses from 1850 to 1920 only 2.16% of those living with a father are linked to a stepfather. Children missing both parents can only be identified through the absence of parents in the household and it is not possible to assume why parents are missing. Older children often lived away from their parents for work. To deal with this eventuality

models include a dichotomous variable indicating whether the child was employed, one indicating whether the child had no relationship to the head of the household, and one indicating if they were living in an institution. Moreover, as a robustness check we limit the sample to children aged 10 or less, therefore less likely to be away from the parental household for work or schooling. Descriptive patterns of children living with a widowed parent or without either parent, by age and gender, are presented in Supporting Information Appendix Figure A1.

Independent variables are based on the full household. The occupational class is defined as the highest class recorded in the household, ranking classes based on their average earnings in 1950. Racial/ethnic minorities were identified based on the census categories of the time in the variable race, reflecting historically situated constructions of race and ethnicity. The available categories included: white, Black/African American, Native American, Chinese, Japanese, Other Asian or Pacific Islander, Other race. We consider all categories different from white as racial/ethnic minority; however, the largest category by far is Black/African American. Macro regions—Northeast, Midwest, South, and West—are defined based on the household location. The migration origin of the household is disaggregated among the largest nationalities of origin during the years analyzed: US Natives, Canada, Central and South America, Northern Europe, Great Britain, Ireland, Germany, Western Europe, Eastern Europe, Southern Europe, Asia, Middle East, Africa, and Others. Since individuals in the household can have different migration backgrounds, the children's origin is defined through a hierarchical procedure. In the baseline case, the father's origin is assigned; if the father is missing, then the mother's origin is assigned; if both are missing, the modal household migration is assigned; if more than one mode is present in the household, the category other is assigned.

Linear probability models are used to examine the connection between parentless households and various variables, with dummy indicators representing the absence of a father, mother, or both parents serving as the dependent variables. Results from logistic regression do not alter the substantive findings.

The relationship between family formation—age assortative mating and age at first child—and stratifying variables is analyzed using two distinct samples. For age assortative mating, a sample of married individuals born between 1850 and 1910 aged 20 to 40 in the 1880 to 1920 censuses, using couples as the unit of analysis. For age at first child, a sample of individuals aged 15 to 30 born between 1850 and 1910 aged 20 to 40 in the 1880 to 1920 censuses.

Occupational class is defined as the highest occupation in the household of the couple. All other variables, racial/ethnic minority, macro region, origin, age, and year are defined based on the reference individual value.

Age at first child is based on the age of the eldest own child in the household. This may introduce biases since earlier children might have been dead or left the household. The age limit of 30 limits the likelihood that first children had already left the household. Since results can be biased by child mortality across groups, we also test the timing of marriage, limiting the analysis to the 1900 and 1910 censuses which included information on marriage duration. These robustness checks are reported in Table A.17 in the Supporting Information Appendix.

Age assortative mating is based on the absolute age difference between the partners. Linear regression is used to investigate the contribution of each independent variable on both outcomes limiting the sample to those married and those that already had a child. Given that age at first child and age at marriage are right censored in our samples, we also test all of these outcomes using Cox proportional hazards models including both parents and childless; results from these checks are presented in Tables A.16 and A.17 of the Supporting Information Appendix. Descriptive statistics for all variables and samples are reported in Tables A13 to A14, including the child sample, the sample for age at first child with non-childless and with all individuals, and the sample of couples.

We estimated separate models for men and women because the experience and consequences of parental loss are likely to differ by gender. In the historical context of our study, gendered family roles and unequal access to the formal labor market shaped men's and women's autonomy outside the household and their economic dependence on parents. These structural differences may also influence how control variables operate across genders. For these reasons, analyzing men and women separately provides a more accurate account of the mechanisms linking parental loss to later-life outcomes.

4 | Results

4.1 | Parental Death and Family Formation

Table 1 presents regression results for models examining the absolute age difference with the partner as the dependent variable, analyzed separately for men and women. The first column shows the results of a model with just the variables indicating parental death as independent factors. Among women (first column of the table), the death of any parent before age 16 is associated with a higher age difference with the partner. The direction and magnitude of coefficients remain similar also when introducing family-related factors such as the number of siblings, stepsiblings, the siblings' sex ratio, and the birth order (Column 2). When testing the siblings' sample (Column 3), associations maintain the same direction, magnitude, and significance for the mother's death and full orphanhood, while for the father's death, we lose significance and magnitude, but not the direction.

Results are similar with different specifications of age assortative mating. Table A.2 in the Supporting Information Appendix shows that among women, the loss of a parent negatively influences the likelihood of being in an age homogamous and hypogamous couple. Table A.4, which includes age assortative mating as a dummy variable distinguishing between age heterogamy and homogamy, shows the same pattern. The loss of a father decreases the likelihood of age homogamy by 2 percentage points among women.

However, when applying siblings' fixed effects (third and fourth column of Table 1), the coefficients of parental death change

TABLE 1 | OLS and siblings' fixed effects estimates on age difference with the partner. Men and women, FamiLinx data.

	Women					Men				
	OLS	OLS + controls	OLS siblings' sample	Siblings' sample FE	Siblings' sample FE + controls	OLS	OLS + controls	OLS siblings' sample	Siblings sample FE	Siblings sample FE + controls
Mother dead before age 16	0.22*	0.19**	0.40***	−0.04	0.16	0.11*	0.14**	0.26***	−0.24	0.29
	(0.08)	(0.08)	(0.12)	(0.33)	(0.35)	(0.07)	(0.07)	(0.10)	(0.23)	(0.24)
Father dead before age 16	0.25***	0.21***	0.05	−0.75**	−0.60*	0.16**	0.13*	0.20**	−0.47**	0.03
	(0.08)	(0.08)	(0.12)	(0.31)	(0.32)	(0.07)	(0.07)	(0.10)	(0.32)	(0.24)
Both parents dead before age 16	0.80***	0.72***	0.84***	−0.15	0.09	0.42***	0.35**	0.38	−0.07	0.67
	(0.18)	(0.18)	(0.29)	(0.62)	(0.63)	(0.19)	(0.15)	(0.24)	(0.47)	(0.48)
N siblings		0.02**	0.03**				0.04***	0.04***		
		(0.01)	(0.01)				(0.01)	(0.01)		
N stepsiblings		0.04***	0.03*				−0.02	−0.01		
		(0.02)	(0.02)				(0.02)	(0.02)		
Siblings' sex ratio		−0.13	0.078				0.11	−0.17		
		(0.11)	(0.17)				(0.09)	(0.17)		
Birth order		0.00	−0.01		−0.12*		−0.02**	−0.03**		0.15***
		(0.02)	(0.01)		(0.13)		(0.01)	(0.01)		(0.07)
Birth year dummy		YES	YES	YES	YES		YES	YES	YES	YES
Constant	5.02***	5.48***	4.97***			4.72***	4.96	5.312		
	(0.02)	(0.21)	(0.36)			(0.03)	(0.20)	(0.27)		
N	44,759	44,759	23,512	23,512	23,512	56,872	56,872	30,237	30,237	30,237
Adjusted R^2	0.001	0.009	0.009	0.14	0.09	0.0003	0.017	0.018	0.13	0.13

Note: ***$p < 0.001$; **$p < 0.01$; *$p < 0.05$. Robust standard errors at the family level in parentheses.

direction and, in most cases, lose conventional statistical significance. After including siblings fixed effects, father's death has a negative and significant effect on age difference for women, even controlling for birth order. This suggests that the positive effect observed in the pooled model was exclusively due to family-constant characteristics driving both selection into parental loss and age heterogamy. Once family characteristics are considered, the effect of parental loss reduces age distance from the partner, consistent with expectations from a social control hypothesis.

Among men results are similar. In OLS models, coefficients of parental death are negative and statistically significant, suggesting that the loss of a parent increases the age difference with the partner. In the siblings' sample (Column 3) coefficients maintain the same direction, magnitude, and significance. The only coefficient that notably changes is that of the father's death, which becomes negligible in magnitude and not significant. One plausible explanation is that restricting the sample further based on the availability of sibling information increases selectivity and bias. In Familinx, more precise information is associated with longer lifespans and larger family networks (Colasurdo and Omenti 2024; Minardi et al. 2024). As a result, father loss may have a weaker impact on these more selected individuals. When introducing sibling fixed-effects, the only effect that reaches statistical significance is that of the father's death, which becomes negative (Columns 4 and 5). Still, the effect disappears or becomes statistically non-significant after controlling for birth order and birth year.

Similar results are found with alternative specifications in Tables A.2 and A2 in the Supporting Information Appendix; the loss of a father decreases the likelihood of age homogamy and increases men's likelihood of being the youngest in the couple.

Table 2 shows models for age at first children. For women, losing a parent before age 16 is associated with earlier transition to childbearing. Being orphaned before age 16 decreases the age at first child by around 1 year (0.98 in the empty model, 1.28 with controls). Results are almost identical in the siblings' sample (Column 3). Such effects remain negative but no longer reach statistical significance in fixed-effects models. Among men, OLS models mirror the negative effect found for women; for instance, losing both parents before age 16 accelerates the transition to childbearing by 0.67 years (significant at the 10% level, column 2). However, the effects of the mother's and father's loss no longer reach statistical significance in the siblings' sample (Column 3). Similarly to women, such effects are no longer statistically significant when controlling for family characteristics in sibling fixed-effects models.

Differences between pooled OLS and siblings fixed effects models, especially with regards to age assortative mating, indicate the presence of family constant characteristics that influence both parental loss and mating patterns. In other words, the observed positive associations are not due to the loss of a parent per se but rather to family constant confounding factors that do not vary across siblings, such as socioeconomic background, urban residence, or geographical origins. For instance, individuals from more disadvantaged families are more likely to experience parental loss, to marry older individuals, or to have children at earlier ages. Once this is accounted for, the loss of a parent is little or even negative.

As an additional robustness check, we examined whether associations varied by the age at which parental loss occurred by splitting the measure into finer age groups. These models generally yielded patterns consistent with those reported in the main analysis, but the estimates were considerably less precise and did not reveal a clear gradient across age categories. This limitation was especially pronounced in the siblings sample, where the number of cases within each age-at-loss group was very small. As a result, it is difficult to draw meaningful conclusions about potential differences in the timing of parental loss. While our main analyses therefore rely on the more parsimonious cutoff at age 16, further research with larger samples could more fully investigate age-specific heterogeneity in these relationships.

4.2 | Correlates of Parentless Households and Family Formation Patterns

While the previous section has suggested the presence of family constant confounders influencing both parental loss and family formation patterns, due to the limited information contained in FamiLinx it is unclear what characteristics exactly drive this selection process. This section investigates how parentless households are distributed across socioeconomic and demographic characteristics, and how these same factors are related to age assortative mating and age at first child using historical censuses.

An association between parental loss and social background is explored in Table 3, which uses census data to estimate the likelihood of children aged 0–16 living in a household missing their father, mother, or both parents (a robustness check with children below 11 is reported in Table A.15). Fatherless households are substantially more frequent for households of persons not in employment or employed as service workers. Service workers include the lowest paid non-agricultural jobs and include all lower-class female segregated occupations. These associations also reflect the gender segregation of labor markets, as most female-accessible occupations were service jobs, and widows were most likely to be represented in this class. Few social differences are visible for children missing mothers; this is not surprising given that widowed fathers raising their children were extremely rare, and they were more often sent to their relatives, institutions, or work in other households (Gordon and McLanahan 1991). Parentless households emerge as a predominantly lower-class experience. They are also more likely among racial/ethnic minority families and rural contexts, showing some geographic stratification and differences among migration backgrounds, with some groups more exposed than others.

Table 4 shows the stratification of age assortative mating and age at first child for males and females. The sample is based on similar cohorts as the previous models. While in the model before characteristics referred to childhood household, in this case, they refer to the adult individual. A larger age difference between partners is most common in some urban contexts and among lower manual classes. Once again, racial/ethnic

TABLE 2 | OLS and siblings' fixed effects estimates on the age at first child. Men and women, FamiLinx data.

	Women					Men				
	OLS	OLS + controls	OLS siblings' sample	Siblings' sample FE	Siblings' sample FE + controls	OLS	OLS + controls	OLS siblings' sample	Siblings sample FE	Siblings sample FE + controls
Mother dead before age 16	−0.47*** (0.12)	−0.66*** (0.12)	−0.69*** (0.16)	−0.29 (0.50)	−0.41 (0.52)	−0.32*** (0.12)	−0.48*** (0.12)	−0.27 (0.18)	−0.18 (0.47)	0.24 (0.489)
Father dead before age 16	−0.40*** (0.12)	−0.57*** (0.12)	−0.67*** (0.16)	0.23 (0.44)	−0.03 (0.46)	−0.05 (0.12)	−0.22* (0.12)	−0.15 (0.17)	−0.37 (0.452)	0.04 (0.469)
Both parents dead before age 16	−0.98*** (0.25)	−1.29*** (0.25)	−1.58*** (0.40)	−0.37 (0.76)	−0.78 (0.75)	−0.35 (0.27)	−0.67** (0.27)	−1.21*** (0.43)	−0.34 (1.003)	0.22 (1.045)
N siblings		−0.14*** (0.01)	−0.13*** (0.02)				−0.13*** (0.01)	−0.06*** (0.02)		
N stepsiblings		−0.06*** (0.02)	−0.05** (0.02)				−0.08*** (0.03)	−0.07** (0.03)		
Siblings' sex ratio		−0.78*** (0.16)	−0.44* (0.24)				0.43** (0.17)	−0.26 (0.25)		
Birth order		0.06*** (0.02)	0.07*** (0.02)		0.30* (0.13)		−0.02 (0.02)	−0.03 (0.02)		0.22* (0.093)
Birth year dummy		YES	YES	YES	YES		YES	YES	YES	YES
Constant	24.02*** (0.04)	25.1*** (0.26)	28.84*** (0.43)			27.68*** (0.04)	27.8*** (0.27)	27.04*** (0.41)		
N	24,343	24,343	12,529	12,529	12,529	30,137	30,137	15,709	15,709	15,709
Adjusted R^2	0.001	0.016	0.016	0.29	0.30	0.0002	0.015	0.008	0.26	0.27

Note: ***$p < 0.001$; **$p < 0.01$; *$p < 0.05$. Robust standard errors at the family level in parentheses.

TABLE 3 | OLS estimates of parentless household. Census data.

	Missing father: widowed or remarried mother		Missing mother: widowed or remarried father		Orphan: Living in a household without parent	
	Coefficient	S.E.	Coefficient	S.E.	Coefficient	S.E.
Household class (Ref. Proprietors, Managers, Professionals)						
Craftsmen	−0.00***	(0.00)	0.00	(0.00)	−0.00**	(0.00)
Clerks and sales	0.03***	(0.00)	−0.00	(0.00)	−0.00	(0.00)
Operatives	0.03***	(0.00)	−0.00	(0.00)	0.00	(0.00)
Laborers	0.02***	(0.00)	0.00	(0.00)	0.00***	(0.00)
Farmers	−0.01***	(0.00)	−0.00	(0.00)	0.00***	(0.00)
Service workers	0.14***	(0.00)	−0.01***	(0.00)	0.02***	(0.00)
Farm laborers	0.05***	(0.00)	−0.00**	(0.00)	0.01***	(0.00)
Not in emp	0.28***	(0.00)	−0.01***	(0.00)	0.07***	(0.00)
Racial/ethnic minority	0.04***	(0.00)	0.01***	(0.00)	0.05***	(0.00)
Urban	0.02***	(0.00)	−0.00***	(0.00)	−0.00***	(0.00)
Region (Ref: Northeast)						
Midwest	0.01***	(0.00)	0.00	(0.00)	−0.00***	(0.00)
South	0.02***	(0.00)	0.01***	(0.00)	−0.00**	(0.00)
West	0.01***	(0.00)	0.00***	(0.00)	−0.00*	(0.00)
Origin (Ref: US born)						
Canada	−0.01***	(0.00)	0.00	(0.00)	−0.01***	(0.00)
Northern Europe	−0.00	(0.00)	0.00	(0.00)	−0.01***	(0.00)
Great Britain	−0.00*	(0.00)	0.00*	(0.00)	−0.01***	(0.00)
Ireland	0.02***	(0.00)	0.01***	(0.00)	−0.01***	(0.00)
Germany	−0.00**	(0.00)	0.00	(0.00)	−0.01***	(0.00)
Western Europe	−0.02***	(0.00)	−0.01***	(0.00)	−0.02***	(0.00)
Eastern Europe	−0.02***	(0.00)	−0.01***	(0.00)	−0.02***	(0.00)
Southern Europe	−0.03***	(0.00)	−0.00	(0.00)	−0.01***	(0.00)
Others	−0.00	(0.00)	0.01*	(0.00)	0.01	(0.00)
In employment	0.04***	(0.00)	0.01***	(0.00)	0.03***	(0.00)
No relation to Household head	−0.04***	(0.00)	−0.02***	(0.00)	0.67***	(0.00)
Year (Ref: 1880)						
1900	−0.00*	(0.00)	0.02***	(0.00)	0.01***	(0.00)
1910	−0.00***	(0.00)	0.02***	(0.00)	0.01***	(0.00)
1920	−0.01***	(0.00)	−0.01***	(0.00)	−0.01***	(0.00)
Age (Ref: 0)						
1–4	0.01***	(0.00)	0.01***	(0.00)	0.01***	(0.00)
5–9	0.04***	(0.00)	0.03***	(0.00)	0.02***	(0.00)
10–16	0.07***	(0.00)	0.05***	(0.00)	0.04***	(0.00)

(Continues)

TABLE 3 | (Continued)

	Missing father: widowed or remarried mother		Missing mother: widowed or remarried father		Orphan: Living in a household without parent	
	Coefficient	**S.E.**	**Coefficient**	**S.E.**	**Coefficient**	**S.E.**
Sex	0.00***	(0.00)	−0.00	(0.00)	0.01***	(0.00)
In institutions					0.86***	(0.01)
Observations	966,216		966,216		967,731	
R2	0.05		0.01		0.25	

Note: ***$p < 0.001$; **$p < 0.01$; *$p < 0.05$. Robust standard errors in parentheses.

minorities show larger heterogamy. Patterns by migration background also show the importance of family characteristics in terms of cultural background. Similar patterns of stratification are visible for age at first child, with the lower manual classes showing the earlier fertility. Again, racial/ethnic minorities showed a strong anticipation compared to whites, and migration background showed significant patterns.

Cox hazard models that account for right-censoring and for individuals who never marry or have children lead, with few exceptions, to similar conclusions. Groups in lower manual classes generally show earlier transitions. The discrepancies that do appear are largely attributable to differences in the proportion of individuals who experience the event. A notable example is Black individuals, who display hazard ratios below one—indicating later timing—yet show a negative association in conventional models. This apparent inconsistency reflects the higher prevalence of lifelong childlessness among Black individuals, while those who did have children tended to do so earlier. A comparable pattern emerges for marriage among women employed in lower service occupations: they were more likely to remain single, but those who did marry often entered marriage relatively early.

Altogether, these results help make sense of the change in association between parental loss and family formation after accounting for family constant characteristics, suggesting that parental loss was not a random occurrence but was highly clustered among specific more disadvantaged groups. These groups, in turn, also exhibited specific patterns of partner selection and age at first child.

5 | Discussion and Conclusions

Parental death during childhood represents a profound disruption to family life that can significantly affect children's long-term outcomes. In historical societies, parents were instrumental in guiding their children's choices, often pursuing economic and social interests, and their premature departure might change the course of their children's choices. On the one hand, the *social control* mechanism might predict less traditional family outcomes because of the released pressure coming from parents. On the other, the *economic constraints* mechanism posits that children might be limited by material constraints when planning the formation of their households.

This paper investigates how the loss of a parent at a young age impacts two outcomes—the age difference between partners and the timing of the first child—in the United States for cohorts born between 1850 and 1910 by using FamiLinx online crowdsourced genealogical data, which allows us to have information on both parents and children. To assess the degree of selection, we use US Census data to assess the level of selection into parental death and family formation outcomes, which indeed the data show.

Our results show an association between parental death and family formation. The loss of a parent increases age heterogamy in the children's couple, both for men and women. More specifically, as shown in additional analyses, women are less likely to marry someone of the same age or younger, whereas for men it becomes more likely to be the younger partner in the couple. Such results seem to support the *economic constraints* mechanism; when facing the loss of a parent, children seek an older partner, which most likely can guarantee economic security. The death of a parent also accelerates the transition to childbearing; for instance, women being orphaned at young ages have children around a year before their counterparts.

However, after the inclusion of family fixed-effects, accounting for family-specific characteristics, the main effects either reduced or even reversed. As for age assortative mating, sons are found to partner with partners of more similar age. Compared to the OLS estimates, the effect is reversed and supports the *social control* mechanism. For women, no significant effects are found for both age assortative mating and age at first child. Thus, results with FamiLinx point toward the role of selection in shaping the relationship between parental death and family formation. Such selection is confirmed with US Census data, which show that individuals from lower classes and racial/ethnic minorities were most likely to live in parentless households, display lower age homogamy with their partner, and have children earlier.

Thus, we can state that selection significantly shapes the relationship between parental death and family formation. Rather than parental death, it is the uneven distribution of family structures that influences life course trajectories and inequalities. We believe disentangling this question is crucial, as understanding these effects has significant implications for children's well-being and their choices throughout the life course.

TABLE 4 | OLS estimates of age distance between partners and age at first child. Census data.

	Age homogamy		Age at first child			
	Couples		Females		Males	
	Coefficient	S.E.	Coefficient	S.E.	Coefficient	S.E.
Household class (Ref. Proprietors, Managers, Professionals)						
Craftsmen	0.08***	(0.01)	−0.68***	(0.03)	−0.59***	(0.03)
Clerks and sales	−0.02	(0.02)	−0.10***	(0.03)	−0.25***	(0.04)
Operatives	0.15***	(0.02)	−0.92***	(0.03)	−0.76***	(0.03)
Laborers	0.24***	(0.02)	−1.06***	(0.03)	−0.82***	(0.04)
Farmers	0.13***	(0.02)	−0.77***	(0.03)	−0.63***	(0.03)
Service workers	0.15***	(0.03)	−0.82***	(0.05)	−0.57***	(0.06)
Farm laborers	0.25***	(0.04)	−0.86***	(0.04)	−0.76***	(0.04)
Not in emp	0.13	(0.12)	−0.97***	(0.09)	−0.90***	(0.15)
Urban	−0.10***	(0.01)	0.31***	(0.02)	0.09***	(0.02)
Racial/ethnic minority	0.10***	(0.02)	−1.20***	(0.02)	−1.05***	(0.03)
Region (Ref: Northeast)						
Midwest	0.10***	(0.01)	−0.13***	(0.02)	0.10***	(0.02)
South	0.26***	(0.01)	−0.83***	(0.02)	−0.42***	(0.02)
West	0.24***	(0.02)	−0.26***	(0.03)	0.21***	(0.04)
Origin (Ref: US born)						
Canada	0.04	(0.04)	0.23***	(0.06)	0.30***	(0.07)
Northern Europe	0.11**	(0.04)	1.04***	(0.07)	0.67***	(0.08)
Great Britain	−0.12**	(0.04)	0.43***	(0.07)	0.56***	(0.08)
Ireland	−0.17***	(0.04)	1.05***	(0.08)	0.77***	(0.11)
Germany	0.04	(0.03)	0.37***	(0.05)	0.44***	(0.06)
Western Europe	0.22***	(0.03)	0.14***	(0.04)	0.33***	(0.05)
Eastern Europe	0.12***	(0.04)	0.06	(0.05)	0.13*	(0.06)
Southern Europe	0.57***	(0.04)	−0.51***	(0.05)	0.27***	(0.06)
Others	0.35***	(0.06)	−0.18*	(0.08)	−0.01	(0.10)
Age (Ref: 25 for age difference; 15 for age at first child)						
+5	−1.05***	(0.04)	2.07***	(0.02)	3.03***	(0.06)
+10	0.17***	(0.04)	3.73***	(0.02)	5.35***	(0.06)
+15	1.17***	(0.05)				
Year (Ref: 1880)						
1900	1.98***	(0.02)	0.51***	(0.03)	0.39***	(0.03)
1910	1.86***	(0.02)	0.56***	(0.02)	0.28***	(0.03)
1920	1.83***	(0.02)	0.26***	(0.02)	0.00	(0.03)
1930	1.73***	(0.02)	0.17***	(0.02)	0.06	(0.03)
1940	1.77***	(0.02)				
Constant	0.80***	(0.04)	18.23***	(0.04)	18.75***	(0.07)
Observations	196,963		182,666		95,840	
R2	0.11		0.21		0.23	

Note: ***$p < 0.001$; **$p < 0.01$; *$p < 0.05$. Robust standard errors in parentheses.

Individuals who lost both parents at an early age appear to represent a particularly disadvantaged group, as the complete loss of parental figures likely entails a substantial reduction in emotional guidance and material support, heightening the risk of limited access to basic subsistence and necessitating major household changes. Previous research has shown that losing a parent can hinder development from childhood into adulthood across diverse contexts, reducing human capital and well-being (Beegle et al. 2010; Case and Ardington 2006). Although less frequently examined, the loss of both parents may place double orphans in an even more vulnerable position (Case et al. 2004; Ruiz-Casares et al. 2009).

In line with this literature, our results suggest that double orphanhood may exert a particularly strong influence on family formation patterns. For this group, both men and women face pronounced pressures, as the absence of both parents removes the differentiated influence typically exerted by mothers and fathers. However, considering the gendered nature of family formation and expectations, and women's historically limited access to formal labor markets, our findings hint that women may be especially affected by the loss of both parents, potentially relying more heavily on family formation as a means of securing stability. Because the experiences of double orphans are rarely documented—even in contemporary contexts, given the challenges of identifying such cases—our results underscore the need for further research on the gendered implications of double parental loss.

We are aware that our results might be challenged by data constraints and methodological issues. Genealogical data does not allow direct testing of the selection mechanism with other techniques rather than siblings' fixed effects, because it does not provide direct information on characteristics such as social class, geographical location, or race. However, genealogical data is still the only historical source that allows exploring intergenerational dependencies between parents and children, and we do fully exploit such features in our study. Further research and development of historical data sources should enhance the possibilities to study such phenomena that data include information on both intergenerational relationships and events and socio-economic characteristics. Also, with FamiLinx we cannot disentangle the effect of parental death from that of family structure, because we do not have information on living arrangements or family disruption for other reasons than parental death. However, due to the generative nature of genealogical data—which relies heavily on knowledge within families—it is likely that records with complete parental information predominantly correspond to two-parent households.

Secondly, besides the well-known advantages of siblings-fixed effects models (McLanahan et al. 2013), we must acknowledge some possible drawbacks that might influence the interpretation of our results. Siblings' fixed effects calculate their estimates just on the pool of siblings that experienced parental loss at different ages. We might thus estimate effects on a selected sample of families with particularly wide age distances between siblings. In our complete sample, the mean age difference between the youngest and oldest sibling within a family is around 9 years, while it rises to 13 years in the subsample used in the sibling fixed-effects estimation. And to the extent to which literature suggests that longer birth intervals were more common among lower social strata (van Bavel and Kok 2004; van Bavel 2004; Bengtsson and Dribe 2006), we might be calculating our effects on a selected sample of families (Engzell and Hällsten 2024).

The main analyses presented in this article have operationalized parental loss as losing a parent before turning 16 years old. As a robustness check, we also investigated the impact of parental loss at different ages (Tables A.5–A.10 in the Supporting Information Appendix). While results from these tests are broadly consistent with the main conclusions, we were unable to assess cross-age differences due to the limited sample size and statistical power. Nevertheless, previous literature has highlighted that the age at which a parent is missing is not inconsequential and can differently affect life courses. Therefore, future research should more directly investigate heterogeneities in this relationship depending on which age the event occurs.

The existing literature and our empirical analysis have largely focused on the nuclear family, thereby overlooking the fact that parent–child relationships are embedded within broader kinship and community networks. These networks can play a crucial role in shaping the consequences of parental loss for both single parents and children by providing material, emotional, and social support, as well as enforcing social control in the absence of parents. A substantial body of work demonstrates the importance of kin and community in many domains of family life in both historical and contemporary contexts (Furstenberg 2020; Furstenberg et al. 2020; Clark et al. 2017; Bratter and Whitehead 2018; Stack 1997). In the historical United States, for example, kin propinquity and availability were particularly significant (Nelson 2020). This suggests that how kin intervene to compensate for parental loss may be an important factor in understanding the link between parental loss and social inequality and represents a promising direction for future research.

The analysis of the interplay between parents' and children's lives provides valuable insights into the interconnectedness of family members and how it persists over the life course. In this paper, we examined the theoretical mechanisms through which the loss of a parent at a young age might influence subsequent steps in family formation, and we explored them empirically. We also addressed the issue of family selection into disruptive events and discussed methodological approaches to tackling this challenge.

Answering these types of questions is challenging methodologically and data-wise, yet it is essential for understanding how family lives and events are interconnected from an intergenerational perspective. Moreover, this link may not only last in the short run, but it can also have a long-lasting effect on younger family members. These dynamics are not confined to historical contexts but also resonate in contemporary societies marked by persistent family inequality and unequal life expectancy.

Evidence exists in contemporary society on the transformative and often negative effects of parental loss and family disruption on a wide range of children outcomes. Although much of the observed association is explained by the deprivation these events

can cause, parental loss is rarely considered deeply embedded in social and economic disadvantages. Results from this study highlight that, rather than being an exogenous factor, family disruption is one of the many mechanisms linking family disadvantages to children's life courses and exacerbating the consequences of inequalities. It suggests that the root of the negative association is not exclusively the absence of a parent per se but the underlying living conditions associated with it, and points to the importance of contextualizing family networks and structures within the family's social, economic, and cultural context for both contemporary studies and policy making. Moreover, although early parental loss is now relatively more uncommon in Western higher-income countries, it continues to be a crucial source of inequality in more deprived contexts, especially as a result of conflicts, epidemics, and extreme climate events. Future research should examine how these disruptive events can produce comparable experiences of early parental death (Alburez-Gutierrez et al. 2024), with potentially lasting consequences for bereaved children.

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Data Availability Statement

One of the datasets that supports the findings of this study is openly available at https://osf.io/fd25c/. The other data supporting the findings are available through IPUMS USA at https://usa.ipums.org/usa/index.shtml.

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Supporting Information

Additional supporting information can be found online in the Supporting Information section. **Table A1**. Descriptive statistics (mean, standard deviation, and percentage). FamiLinx data. **Table A2**. Multinomial logistic regression estimates, age assortative mating as categorical. FamiLinx data. **Table A3**. Distribution of age at parental death by 5-years categories. **Table A4**. OLS and siblings' fixed effects estimates of age assortative mating. Dummy variable, FamiLinx data. **Table A5**. OLS and siblings' fixed effects estimates of mother's death on age difference with the partner. Men and women, FamiLinx data, variable of parental age as categorical (5-years groups). **Table A6**. OLS and siblings' fixed effects estimates of father's death on age difference with the partner. Men and women, FamiLinx data, variable of parental age as categorical (5-years groups). **Table A7**. OLS and siblings' fixed effects estimates of both parents' death on age difference with the partner. Men and women, FamiLinx data, variable of parental age as categorical (5-years groups). **Table A8**. OLS and siblings' fixed effects estimates of mother's death on age at first child Men and women, FamiLinx data, variable of parental age as categorical (5-years groups). **Table A9**. OLS and siblings' fixed effects estimates of father's death on age at first child Men and women, FamiLinx data, variable of parental age as categorical (5-years groups). **Table A10**. OLS and siblings' fixed effects estimates of both parents' death on age at first child Men and women, FamiLinx data, variable of parental age as categorical (5-years groups). **Table A11**. Multinomial logistic regression estimates, age assortative mating as categorical (5-years difference). FamiLinx data. **Table A12**. Multinomial logistic regression estimates, age assortative mating as categorical (2-years difference). FamiLinx data. **Table A13**. Descriptive statistics for the children sample. **Table A14**. Descriptive statistics census samples with, parents, all, and couples. **Figure A1**. Share of children in household with widowed or without any parent by age and sex. **Table A15**. Linear probability model of widowed parent on children household characteristics for children younger than 11 years old. **Table A16**. Cox proportional hazards models on transition to first child for females and males. **Table A17**. OLS of age at marriage and Cox proportional hazards models on marriage by sex.