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Family Systems and Parents' Financial Support for Education in Early Adulthood

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Abstract

Young adults raised outside of two-parent families receive less financial support from their families for education compared with peers who always lived with both parents. We consider how parents' union status over time shapes contributions for young adult children's education. Our approach emphasizes the dynamic relationship between family structure and family economic resources. Marginal structural models with inverse probability weights estimate the association of parents' union status history with eventual financial transfers while not overcontrolling for the effects of union status operating indirectly through time-varying characteristics, such as coresident family composition and economic circumstances. The analytic sample includes parents of a recent cohort of young adults (Panel Study of Income Dynamics, 1983–2013, N = 2,754). Compared with parents who lived continuously with a child's other parent, unpartnered parents' transfers to children were 44 % to 90 % smaller, and repartnered parents' transfers were one- to two-thirds smaller, depending on how long the parent was unpartnered or repartnered. Through its influence on subsequent coresident family composition and family economic resources, parents' union status has indirect as well as direct associations with financial transfers to adult children for education.

Keywords Family structure · Family complexity · Education · Early adulthood · Financial transfers

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Introduction

Young adults raised outside of two-parent families receive less financial support from their families to attend college compared with otherwise similar peers who live with both biological or adoptive parents continuously through childhood (Henretta et al. 2012; Turley and Desmond 2011; Wojtkiewicz and Holtzman 2011). This disparity, evident only when children reach early adulthood, emerges from a lifetime of exposure to divergent family systems. Elemental aspects of these varied family systems include their *structure*, including the union status of a child's coresident parents and the biological relatedness of individuals in a shared family system; their *fluidity*, or the potential for change in family composition and coresidence over time; and their *dynamic* nature, or the extent to which change in one dimension of a family system precipitates change in another. When considered together, these three dimensions of a family system characterize the diversity of experience among the more than one-half of U.S. children living outside of their parents' ongoing shared first marriage (Livingston 2014).

We ask whether and how the structure, fluidity, and dynamic nature of family systems shape the financial support parents provide to their children for postsecondary education. We treat parents' union status as the linchpin of the family system, focusing particularly on whether a parent is partnered (either married or cohabiting) with a child's other biological parent, unpartnered, or repartnered. At the same time, we consider explicitly how parents' coresidence with their own children and the presence of other children in the family household both shape and follow from parents' union status over time. Our dynamic conceptual model thus treats family composition not as a fixed structure but instead as an evolving system with the potential to change the propensity for parents to invest in children's educational attainment over time. This framework brings together literatures developed in parallel during the last decade that have focused separately on how family structure, frequent family change, and patterns of exchange in complex families potentially compromise parents' investments in children's development.

Our analytic approach matches this conceptual approach. We use marginal structural models to adjust for time-varying confounding between parents' union status and other family circumstances, including parent-child coresidence, the presence of other children, and family socioeconomic status to preserve the direct and indirect effects of union status on subsequent financial transfers. We use long-running longitudinal data to consider the totality of parents' union status history in a contemporary cohort of young adults (born in 1983–1995). Data are from the U.S. Panel Study of Income Dynamics (PSID), a national household-based panel survey that has followed a sample of approximately 4,800 families and their descendants since 1968. We pair prospective annual or biennial longitudinal data on family organization and family income in parents' households with parents' reports of the total amount of money they have provided to support each child's education in early adulthood as of 2013.

Background

Children with single parents or in stepparent families receive less financial support for postsecondary education and are less likely to attend college compared with those with continuously partnered parents who have all their biological children in common (Wojtkiewicz and Holtzman 2011). Single parents' lower financial support to coresident children has largely been attributed to their lower family household income compared with two-parent family households (Thomson et al. 1994). Stepparent families, in contrast, have household earnings on par with two-parent families but contribute less in absolute dollars, as a proportion of family income, and as a proportion of student need (Turley and Desmond 2011). Within stepfamilies, however, the likelihood and magnitude of financial transfers to children varies by parents' relationships to children. Where unions include children from either parent's prior relationship and joint children, different-sex couples make transfers to children from the male partner's prior union 25 % to 50 % less often than they do to the children they had together (Henretta et al. 2014). The authors speculated that the lower probability of making transfers to stepchildren—and particularly to the father's prior children—may result from children's continued residence with their other biological parent outside of the stepfamily household. This explanation is consistent with research findings that nonresident fathers make fewer financial contributions to children compared with resident fathers (McLanahan et al. 2013) and that the likelihood of making such contributions continues to diminish as parents repartner and have additional children (Berger et al. 2012; Carlson and Berger 2013; Tach et al. 2010, 2014).

A critical point in this literature is that the same individual may be simultaneously a resident parent to one biological child and a nonresident parent to another. Further, the same individual may have coresident or nonresident stepchildren through a spouse or cohabiting partner. Thus, the probability that a parent provides financial support for a given child's postsecondary education is likely related to their union status (whether unpartnered or in a union with a child's other parent or a new partner), the parent's residential status with that child, and the parent's and child's degree of relatedness to other children present in the households they each occupy. Over time, each of these three conditions—parents' union status, residential status, and the presence of other biological children and stepchildren in a family system—potentially influences the other two. Moreover, these conditions together both arise from and contribute to financial strain and the family economic resources potentially available to a parent to support children's postsecondary education.

This perspective builds on a growing body of research that characterizes family composition and change as *family complexity,* which Carlson and Meyer (2014:7) defined as "occur[ring] when marriage and legal ties, living arrangements, fertility, and parenting in a child's family are not coterminous." In practical terms, this definition describes systems of family organization other than married, coresident two-parent families in which both parents have all their biological or adoptive children in common. Contemporary family complexity is characterized by nonmarital fertility, single parenthood and parental absence, parents' union dissolution, new family formation in residential and nonresidential parents' households, and parents' fertility with new partners or with former partners prior to a child's birth. Thus, the concept of family complexity powerfully connects a variety of family types and statuses. It recognizes that many of the family structure circumstances that family scholars have considered separately in fact overlap to create complex family systems.

Conceptualizing Family Complexity Over Time

Reflecting this definition, a substantial body of recent work has described complex family structure and its association with child well-being. This work has developed multidimensional measures of family structure from cross-sectional data or has developed such measures at one point in time to predict later well-being using longitudinal data. Using the 1996 and 2009 waves of the Survey of Income and Program Participation, Manning et al. (2014) considered the nexus of parents' union status and the biological relatedness of parents and siblings to children. They documented that more than 40 % of contemporary children reside outside of a simple nuclear family structure, including children living with single or cohabiting parents, stepparents, or half- and stepsiblings. Several studies in the last decade have also assessed these aspects of family structure collectively to demonstrate that complex family structure is associated with children's compromised academic and behavioral development across the early life course (Fomby et al. 2016; Gennetian 2005; Halpern-Meekin and Tach 2008; Tillman 2008).

A second line of research has focused on family complexity as fluid, treating change in the composition and coresidence of family members over time as elemental to complex family structure. That is, family complexity both results from and contributes to an evolving set of relationships between adults and children who are connected through union status, residential status, and fertility over time. Point-in-time measures of family complexity status like those described earlier implicitly reflect the prior change that led to current circumstances, but they do not articulate the process by which that change emerges.

A focus on fluidity highlights two aspects of complex family structure. First, family structure at one point conditions the probability of occupying any other family structure status in the future. For example, nonmarital fertility substantially reduces the likelihood that a woman will eventually marry either the biological father of her child or another partner (Carlson et al. 2004; Gibson-Davis 2011; Graefe and Lichter 2002; Raley 2001; Smock and Greenland 2010) and increases the likelihood of multipartner fertility, or having children with more than one partner over the reproductive life course (Cancian et al. 2011; Guzzo 2014). This point emphasizes how prior circumstances continue to influence family structure trajectories over the life course.

Second, this concept recognizes the relative durability or instability of particular family structure statuses. For example, research in this area has documented that marriage is a more enduring union status than cohabitation or being unpartnered. Importantly, research in this area has demonstrated the salience of both the structural and fluid aspects of family composition by documenting that cumulative time in a given family structure status and exposure to frequent changes between family statuses are independently associated with child and adolescent well-being (Fomby and Cherlin 2007; Wu 1996; Wu and Martinson 1993) as well as with educational attainment in early adulthood (Fomby 2013).

A third perspective has focused on the dynamic properties of family systems, or the extent to which change in family structure precipitates change in how family members engage with or invest in one another. In particular, this perspective has considered how nonresident parents' investments in children are shaped by their own and their former partner's repartnering and subsequent childbearing after union dissolution. This research has emphasized that parents and other actors in complex family systems confront competing obligations for time and financial support across households (Edin and Nelson 2013; Manning et al. 2003; Willis 2000) as well as ambiguity in social roles and relationship dynamics that arise from a lack of normative expectations about altruism and reciprocity (Cherlin 1978, 2004; Stewart 2005; Townsend 2002). Both mothers' and fathers' new relationship formation and subsequent childbearing are associated with declines in nonresident parents' financial contributions to children (Berger et al. 2012; Burton and Hardaway 2012; Carlson and Berger 2013; Carlson et al. 2008; Edin and Nelson 2013; Tach et al. 2010, 2014), a shortfall that is not necessarily recovered by financial contributions from coresident social parents or stepparents.

An overarching alternative to these perspectives is the expectation that *baseline selection* mechanisms explain why parents in complex families provide smaller or less frequent contributions to adult children's education compared with parents in stable unions characterized by single-partner fertility. Under this perspective, variation in the distribution of parents' economic resources to young adults is attributable not to family complexity during childhood but instead to circumstances arising from parents' characteristics that were present before a child's birth and that contribute both to parents' eventual complex family formation and to their limited long-term earning power (Fomby and Cherlin 2007; Furstenberg 2014; McLanahan and Percheski 2008; McLanahan et al. 2013). Thus, parents in complex families may have fewer resources to distribute to children because from the outset, they occupied a socioeconomic position that would be predictive of low earnings and low asset accumulation even in the absence of family complexity. Conventional regression approaches adjust for the confounding influence of these prior conditions on the association of interest by including relevant indicators, measured either retrospectively or prospectively, as covariates.

Time-Varying Confounding

Methods to account for baseline selection are a powerful corrective to associations that might otherwise be inferred to be causal. Yet family complexity and family resources are each time-varying, and even after accounting for baseline selection, they are likely to continue to influence one another over time. For example, a parent's union dissolution at Time 0 (t_0) is likely to influence both that parent's union status and household income at Time 1 (t_1). These values continue to influence their own and each other's values in later observations and subsequently to influence financial support for postsecondary education during the transition to adulthood. This dynamic interplay leads to a problem often referred to as time-varying confounding that complicates efforts to isolate the independent associations of family complexity and socioeconomic resources in childhood with parents' investments in children's postsecondary education.

The predicament arising from time-varying confounding is that family economic resources potentially confound the effect of family complexity on financial support for postsecondary education and thus should be controlled for. At the same time, however, family economic resources may be influenced by family complexity in the preceding period, suggesting that simply controlling for these factors would in fact remove some of the true long-term causal effects of family complexity that operate indirectly through family economic resources. As a result, when a time-varying covariate is included in a conventional regression model, it potentially "controls away" the indirect effect of the variable of interest on the outcome measure, yielding biased—usually underestimated—coefficients. Beyond that, any observed association between the two variables over time may be spuriously correlated if underlying factors associated with both family complexity and family economic resources are not taken into account.

A similar issue affects the measurement of family complexity in conventional regression models. An economical approach to summarizing family complexity is to construct a categorical indicator that captures its multidimensional nature simultaneously in terms of parents' union status, coresidence with the focal child, and relatedness of other household members to the focal child, and to compare the association of time in these complex family statuses with time in simpler forms of family organization with regard to the outcome. This approach, however, overlooks the dynamic process by which parents and children came to occupy a multidimensional state.

An alternative approach would be to focus on one dimension of family complexity while statistically controlling for the others in order to establish its independent association with the outcome. In the current case, a focus on parental union status would control for parent-child coresidence and the presence of other children in the parental household in order to describe how residing with a new partner or no partner shapes financial support for children's education in adulthood. Yet the focal child's presence in his or her parent's household at one point in time likely shapes that parent's union status at a later point in time (Carlson et al. 2004), and a parent's union status in turn shapes future coresidence with children (Tach et al. 2010). Thus, because the components of family complexity interact dynamically, their inclusion as covariates in a conventional regression model potentially overcontrols for the indirect effect of any single component of interest on the outcome.

Our analytic approach engages the dynamic interplay between multiple dimensions of family complexity and family economic resources in three ways: (1) it considers multiple key dimensions of complex family structure; (2) it rigorously accounts for parental selection into family systems; and (3) it adjusts for time-dependent confounding between parents' union status on the one hand and parents' family household composition and economic resources on the other. This approach yields refined estimates of the independent associations of parental background, family organization, and cumulative economic resources with parental investments in young adults' schooling to better inform policies targeting status attainment during early adulthood.

Family Complexity and Economic Resources for Education

We focus on the association between family complexity and parents' financial support for children's postsecondary education for four reasons. First, postsecondary education, particularly enrollment in a four-year college or university, is costly. In the 2014–2015 academic year, the average annual cost of undergraduate tuition, room, and board at a public four-year public institution was \$18,632; at private four-year nonprofit institutions, it was \$37,990 (Snyder et al. 2016). Second, the majority of financing for college tuition and related expenses is borne privately. Approximately 35 % of the cost of undergraduate tuition and expenses was covered by grants and scholarships to students in 2016–2017; the balance, or 65 %, was covered by a combination of parent and student income, savings, and borrowing, supplemented by contributions from kin (Sallie 2017). Rising costs have likely outpaced many families' capacity to save for college. Since 2005–2006, after adjusting for inflation, the cost of postsecondary education has increased by 34 % at public institutions and 26 % at private institutions (Isaacs et al. 2012; Snyder et al. 2016). Third, postsecondary education, particularly completion of a bachelor's degree, continues to yield substantially greater returns to income, wealth, and social mobility compared with earning a high school diploma or less (Isaacs et al. 2012; Julian and Kominski 2011). Further, evidence shows that these economic gains are strongest among those who are least likely to attend college (Brand and Xie 2010; but see Breen et al. 2015). Thus, strategies to mitigate financial barriers to enrollment and retention among students from economically disadvantaged families may be particularly remunerative in the long run.

Finally, we know relatively little about how contemporary family organization is related to children's experience of the transition to adulthood, including educational attainment. In earlier historical periods, family complexity emerged from remarriage and subsequent childbearing with a new spouse following widowhood, and more recently from remarriage following divorce. Today's young adults represent the first cohort also to experience family complexity emerging from the formation and dissolution of cohabiting unions and from nonmarital fertility (Carlson and Meyer 2014). Further, the various pathways through which contemporary family complexity occurs is largely patterned by social class (Guzzo 2014; McLanahan 2004), with children from more economically disadvantaged families more likely to experience family complexity outside of marriage compared with children from more advantaged backgrounds.

Data and Method

We describe variation in parents' investments in children's educational attainment during early adulthood using data from the U.S. PSID and its accompanying 2013 Rosters and Transfers module, which captured transfers of time and money between household heads and spouses/partners and their adult children. Measures of young adults' exposure to family complexity in childhood are derived from parents' birth and marriage histories, which have been collected from adult panel study members since 1985, and from children's residential histories constructed from household rosters.

PSID is a nationally representative, longitudinal and multigenerational study of U.S. families begun in 1968 to describe changes in household income and risk of poverty over time. The original sample of 4,802 U.S. families included an oversample of low-income households. (An immigrant refresher sample of about 500 families who moved to the United States since 1968 was added in 1997, but these families are excluded from our analysis.) When children who are born or adopted into PSID families grow up and establish their own households, they become PSID respondents themselves and pass on eligibility for the study to their own children. The study now includes as many as five generations of family members descended from the original 1968 household heads. Interviews were conducted annually through 1997 and biennially since then. As of 2013, 38 waves of data had been collected over 45 years, and the study achieved reinterview response rates of 96 % to 98 % in nearly every wave (McGonagle et al. 2012). In 2013, the study covered nearly 10,000 families and 25,000 individuals.

We focus on parents' financial transfers to their young adult children to support postsecondary education. The analytic sample includes PSID parents who completed the Rosters and Transfers module in the 2013 Core PSID interview and who had at least one child between 18 and 30 years old in that year (i.e., born between 1983 and 1995). Life course sociologists and developmental psychologists describe the life stage between ages 18 and 30 as the transition into adulthood, or a period marked by the intersection of biological age and the assumption of new roles and responsibilities as financially independent adults, workers, parents, and partners (Osgood et al. 2005). In the contemporary United States, parents often bear the financial investment in education, residential arrangements, and family formation required for young adults to achieve these social roles over a relatively short period (Furstenberg 2010).

The 2013 Rosters and Transfers module asked all household heads and their spouses or partners to roster all living parents and adult children (age 18 and older), regardless of whether the parent or child had ever lived in a PSID household. The respondent reported separately on the value of four types of income transfers to adult children: (1) past-year gifts or loans valued at $100 or more; and lifetime transfers since the child turned 18 (2) for school, including tuition, room and board, or books; (3) for help with a home purchase, including a down payment; or (4) for any other reason.

Dependent Variable

The dependent variable is lifetime transfers from a parental household in support of a young adult child's education since the child turned 18. This financial support might have been directed to any type of educational attainment, including completion of a high school diploma/GED or enrollment in a vocational school, certificate program, two-year or four-year nonprofit or for-profit college or university, or graduate school. Whether the money was actually spent for its intended purpose is not reported. Further, our modeling approach does not take into account whether young adults ever planned to enroll in postsecondary education. To be sure, in some cases, parents did not provide financial support for education in adulthood because their children never intended to enroll or remain in school after age 18. The available data do not allow us to disentangle the causal ordering between young adults' intentions and parents' financial support for education. We describe results from supplementary analyses limited to parents whose children ever attended college to assess the robustness of our conclusions.

Our focal analysis considers transfers where at least one parent was present in or descended from a 1968 PSID household and that parent was the household head or the spouse or partner of the head at the 2013 PSID interview. For most contemporary young adults, only one parent satisfies this condition. As a result, where children did not grow up living continuously with both parents, the transfer amount may be reported for the household headed by the person who was the childhood resident *or* nonresident parent, depending on which parent is descended from an original PSID family. Because status as a PSID descendant is exogenous to whether a parent lived continuously with his or her child, these data provide a relatively unbiased estimate of the average parental household transfer in the context of family complexity, drawing on information from both resident and nonresident parents. (Nonresident parents may be underrepresented to the extent that they were more likely to attrite from the study compared with adults who were stably partnered and coresident with children.) This approach highlights the dynamic interplay between family complexity and family economic resources: changes in family structure through childhood precipitate changes in parents' financial investments in children in both residential and nonresidential parental households, potentially culminating in lower total inputs to education.

Key Independent Variables

The main independent time-varying variables pertain to children's exposure to family complexity and family household income. The measures of *family complexity* were constructed from household rosters with move-in and move-out dates collected at each wave, as well as birth and union histories collected from all adult household members since 1985. From these rosters and histories, we considered for each parent-child pair the following aspects of the parent's family composition in each year when the child was between ages 1 and 18: (1) whether the parent was married to or cohabiting with the child's other biological parent, married to or cohabiting with another spouse/partner, or unpartnered (parents' union status); (2) whether the parent and child lived together (parent-child coresidence); and (3) the total number of other biological or adoptive children, stepchildren, and children of a parent's cohabiting partner in the parent's household (children in the family household).

In the analytic models described later, we permitted item (1) to influence items (2) and (3), and vice versa. That is, we expected the parent's union status at time t to be predictive of the likelihood of residing with the focal child and of the presence of other children in the family household at $t + 1$. Conversely, coresidence with the child and other children at time t were expected to influence parents' union status at time $t + 1$. Thus, reflecting the conceptual model of family complexity as a multidimensional, fluid, and dynamic process, the analytic model evaluates the association of a parent's duration in a given union status on his or her predicted transfer to an adult child after adjusting for parent-child coresidence and the presence of other children as potential time-varying confounders. We elaborate on this methodological approach later.

Total *family household income* refers to past-year family income in the household where a child resided in a given year. This PSID-constructed measure was calculated from respondent reports of household members' income from all sources, including annual taxable earned income and past-year public transfer income in the preceding year from all household family members aged 16 and older; past-year child support; past-year farm income; and past-year asset income (since 1993). All family income measures refer to the complete calendar year prior to the year of interview; for example, family income computed from the 2013 interview refers to the 2012 calendar year. Thus, we treated family income as a $t - 1$ measure such that family complexity in year t was a function of the past-year ($t - 1$) income, reported in year t. Since shifting to biennial interviewing after 1997, PSID has collected less detailed family income information for the year prior to the most recent calendar year (e.g., the year 2011 in the 2013 interview). In order to use a common metric for calculating family income in all years, we averaged reports of income from consecutive waves to estimate family income in the intervening year.

We accounted for a variety of baseline and time-varying confounders of parents' union status. Time-invariant variables include the child's year of birth (grouped into two-year intervals) and parent's race/ethnicity (non-Hispanic white, non-Hispanic black, or Latino), gender, and age at the child's birth. The following time-varying characteristics were measured in the child's household at baseline (at age 1 or 2) and in each year thereafter: (1) number of children under age 18 in the household, including full, half-, and stepsiblings of the focal child, children of the parent's cohabiting partner, and other minors; (2) whether the focal child was present in the parent's household; (3) family income ($1,000s, standardized to the year 2000); and (4) educational attainment, employment status, and homeownership status of the household head.

Sample Restrictions

The Rosters and Transfers module includes information on 3,983 parent-adult child pairs for children 18 to 30 years old in 2013. We excluded from our analytic sample the following pairs: 43 duplicate parent-child pairs; 652 parent-child pairs who were added to PSID in the 1997/1999 immigrant refresher sample (because these families were not observed until children were age 2 or older); and 15 pairs missing information on parent's race. We further restricted the sample to exclude 86 parent-child pairs for whom information on financial transfers for education was missing and 433 pairs in which the parent missed two or more consecutive core interview waves when the young adult child was 1 to 18 years old. The final sample size includes 2,754 parent-child pairs and, when weighted, is representative of parents of contemporary young adults in 2013 whose ancestral families were in the United States in 1968.

Method

To address the problem of time-varying confounding and to represent the dynamic nature of family complexity, we modeled financial support for postsecondary education during the transition to adulthood as a function of parental union status, parent-child coresidence, the presence of other children in the parent's household, and family economic resources using marginal structural models in which the parameters are estimated using inverse probability of treatment (IPT) weights. This class of statistical methods reduces bias in estimators compared with conventional methods by adjusting for time-dependent confounding between family complexity and family economic resources. The models provide average treatment effects of occupying a particular union status category for a given duration after accounting for the influence of time-varying confounding on the other factors. This approach was originally developed in a longitudinal framework in epidemiology to isolate the effect of a treatment on health outcomes when the level of treatment was determined by prior health status and was expected to influence both subsequent health status and subsequent treatment (Robins et al. 2000). The technique has been applied in social science to improve causal estimates of duration of exposure to low-income neighborhoods on educational attainment (Wodtke et al. 2011), early childbearing (Wodtke 2013), and smoking (Kravitz-Wirtz 2016a) and obesity (Kravitz-Wirtz 2016b) in early adulthood, and to marriage on criminal behavior (Sampson et al. 2006). We extend this approach to refine estimates of the relationship between parents' unpartnered or repartnered union status at child ages 1 through 18 and eventual financial support for children's postsecondary education during the transition to adulthood.

We implemented the analysis in two steps. First, we constructed IPT weights to generate a weighted pseudo-population in which treatment (in this case, parental union status) was no longer confounded by measured baseline or time-varying covariates related to family economic resources or other dimensions of family complexity. The denominator of the IPT weight for each parent-child pair was computed from a multinomial logit regression model estimating the probability that the parent occupied his or her actual union status at a given wave conditional on past union status, coresidence with the focal child, presence of other children in the household, family economic resources, and baseline confounders. In each year, the IPT weight gives more or less weight to parents who are underrepresented or overrepresented in their current union status category given their prior-year circumstances and baseline characteristics. The cumulative probability of a parent's position in a given union status over the child's early life course is the product of year-specific probabilities from the time the child was age 1 to age 18. The IPT weight ensures that the values of each of the covariates that contribute to the construction of the IPT weight are balanced in expectation across the three union status categories in each year. Thus, in the resulting weighted pseudo-population, exposure to a given union status operates as if it were randomized with respect to earlier covariates and mitigates the problems associated with time-varying confounding on observed covariates (Robins et al. 2000; Wodtke et al. 2011).

Without further adjustment, the numerator of the IPT weight would be 1, and the weight's value would represent the inverse of the probability of occupying one's observed union status category at each wave cumulated over 18 observation periods. To reduce variability in the IPT weight and improve efficiency in our estimation model, we generated a stabilized IPT weight, where the numerator represents the probability that a parent is in a given union status category in a given year as a function of their values on time-invariant and baseline covariates, but excluding time-varying covariates (Hernan et al. 2002). The resulting weights are roughly centered on 1, with relatively low variability. We further adjusted the stabilized IPT weight to trim outlier values at the 5th and 95th percentiles. Finally, we multipled the trimmed, stabilized weight by the 2013 PSID longitudinal individual weight for the parent to account for attrition and to make results generalizable to parents of young adults in the United States in that year.

In the second step, we estimated the predicted value of a parent's contribution to a child's postsecondary education as a function of the parent's union status history over the course of childhood, where union status history was measured as the proportion of time a parent occupied each status from the time the child was age 1 to age 18. Because the dependent variable is highly skewed, with a substantial amount of clumping at $0, we estimated the outcome using a zero-inflated negative binomial regression model. The zero-inflated model estimates separately the likelihood of making no transfer to children and the size of the transfer, if one was made. We used the same covariates to estimate both components of the model and applied Huber-White robust standard errors to account for the clustering of young adults within the same PSID parental households.

We summarize three specifications of the estimation model. The baseline model predicted parental financial support for young adults' educational attainment as a function of long-term exposure to a parent's union status only, weighted using the PSID longitudinal individual weight from 2013. Second, a conventional covariate-adjusted model introduced time-invariant covariates (i.e., parent and child characteristics at age 1) and the value of covariates including parent-child coresidence,

coresidence with other children, and family economic resources, measured when the focal child was age 14. To the extent that these covariates are confounded with parental union status, the coefficients associated with union status are expected to be reduced in magnitude compared with the baseline model but to remain biased. The third model applied the IPT weights constructed in the first step (as presented earlier) to predict financial transfers in a weighted pseudo-population in which exposure to family complexity in each wave was independent of baseline covariates and time-varying covariates in prior waves. Because the time-invariant and baseline characteristics informed the numerator and denominator of the IPT weight, they were included as covariates in the estimation model, but time-varying covariates were excluded. The influence of these variables as potential confounders on both selection into parents' union status and eventual financial support for postsecondary education was accounted for through the weighting process. Their influence as potential mediators of the association between parental union status and transfer amounts is not observed directly but operates through the overall union status effect.

To be unbiased and consistent, the parameters estimated using IPT weights (described later) require several assumptions: no unmeasured confounding (exchangeability), positivity, and correct specification of the model used to estimate the weights (Cole and Hernán 2008; Robins et al. 2000).

Although exchangeability assumptions are not testable in observed data, our analysis included a wide array of the most common joint predictors of treatment (parental union status) and outcome (parental investment in adult children's postsecondary education), as identified using theory and prior research and described previously.

The positivity assumption states that every unit (parent-child pair) must have at least a nonzero probability of exposure to every treatment group (parent partnered with child's other biological parent, repartnered, or unpartnered) across all levels and combinations of measured covariates. Theoretically, there is no reason that a parent-child pair could not *possibly* be exposed to every treatment group (a structural 0), given that union status is not formally conditioned on the basis of demographic or socioeconomic characteristics.

For the assumption of no model misspecification to hold, the mean of the IPT weights should be reasonably close to 1, and the range of values should be small. The best-behaved weights (mean ~ 1, small range) were achieved by bottom- and top-coding all weights outside the 5th and 95th percentiles, respectively. Alternate specifications based on different truncations of the weights did not substantively change our main results (Cole and Hernán 2008).

Results

Table 1 summarizes parents' year-to-year and lifetime experience of family complexity on three dimensions (weighted): union status, coresidence with the focal child, and total number of children in the parental household from the focal child's age 1 to age 18. When the focal child was age 1, about 66 % of parents resided with the child's other biological parent, and one-third were unpartnered; 1 % resided with a partner who was not the child's other biological parent. By child's age 18, the share of parents living with the child's other parent declined to 44 %, and the prevalence of families headed by unpartnered parents or stepfamilies increased commensurately (41.7 % and 13.9 % at age 17, respectively). The share of parents living with the focal child also declined, from 92 % at age 1 to 76.6 % at age 18. Family size increased from 2.07 children at child age 1 to a peak of 2.43 children at age 9, and then declined to 1.38 children at age 18. On average, parents spent 56.3 % of a child's years residing with the child's other biological or adoptive parent, 36.6 % unpartnered, and 7.1 % with a new partner; parents lived with the focal child about 88 % of the time (bottom row, Table 1). A substantial amount of change underlies these states. In each year, 6 % to 7 % of parents transitioned from one union status to another, and 2 % to 3 % of parents changed their coresidential status with children.

Table 2 summarizes time-invariant and baseline characteristics of the sample. Table A1 in the online appendix summarizes time-varying parental characteristics at each year of child age.[1]

Table 3 summarizes results of zero-inflated negative binomial regressions predicting the size of financial transfers from parents to children for educational purposes. Models 1 and 2 use the standard PSID longitudinal individual weight to account for oversampling, clustering, and attrition in order to produce population-representative estimates. Model 3 uses the IPT weight, which incorporates the PSID longitudinal individual weight. Time-varying covariates are measured at age 14 in Model 2 and at baseline in Model 3. Coefficients in panel 1 represent the log odds that a parent provided no financial transfer to a given child. Exponentiated values of the coefficients in panel 2 represent expected percentage change in the predicted size of a financial transfer (given that one was made) for a one-unit change in the associated independent variable. In Models 1 and 2, the reference category is parents who lived with the child's other biological parent at age 14; in Model 3, it is parents who lived with the child's other biological parent continuously.

In the unadjusted model (Model 1), unpartnered parents and parents living with a new partner at child age 14 were significantly more likely later to make no transfer to the focal child compared with parents living with the child's other biological parent (panel 1) and the predicted amount of any such transfer was significantly lower (panel 2). In Model 2, the magnitude of these associations was reduced by about one-half for both union status categories in panel 1 and by approximately one-third for unpartnered parents in panel 2 after we accounted for covariates that were expected to confound the association between parents' union status and subsequent financial transfers. However, coefficients remained statistically significant at $p < .05$ or less. Indicators of family economic resources at age 14—including average family income and the head's homeownership, employment status, and educational attainment—significantly predicted transfers in the expected direction.

Model 3 incorporates the IPT weight. When selection processes and time-varying confounding were accounted for, the proportion of a parent's time spent outside of a union with the focal child's other parent did not significantly predict whether the parent made any transfer to the child (panel 1). However, time as an unpartnered parent was a strong negative predictor of the conditional transfer amount (panel 2). The

[1] Because some children were first observed at age 2, the prevalence statistics for baseline family structure reported in Table 2 are not identical to those reported for family structure at age 1 in Table A1 (online appendix).

Table 1 Distribution of and change in family complexity components by child age for parents of young adults born 1983–1995: Panel Study of Income Dynamics (PSID)

	Union Status				Coresidence		Fertility
Child Age	Partnered With Other Biological Parent (%)	Single (%)	Repartnered (%)	Any Change (%)	Parent-Child Coresident (%)	Any Change (%)	Total Number of Children in Household
1	65.76	33.14	1.11	–	92.08	–	2.07
2	64.55	33.92	1.53	6.14	92.02	3.33	2.12
3	63.59	34.36	2.05	6.30	92.00	3.26	2.22
4	61.64	35.01	3.35	6.67	90.58	3.03	2.29
5	60.51	35.27	4.22	6.02	90.56	2.38	2.35
6	59.05	36.12	4.83	5.18	89.55	2.74	2.39
7	57.77	35.71	6.52	6.56	89.20	2.84	2.41
8	56.23	36.51	7.25	6.04	88.27	2.53	2.40
9	54.53	36.83	8.64	6.12	88.07	2.83	2.43
10	53.72	36.61	9.67	6.78	86.56	3.44	2.38
11	52.16	38.17	9.67	6.53	86.77	2.91	2.39
12	51.36	37.87	10.77	6.25	85.58	2.85	2.33
13	49.87	37.92	12.22	6.84	85.79	3.05	2.26
14	48.79	39.80	11.41	7.49	84.59	3.92	2.18
15	46.90	39.62	13.49	7.20	82.99	3.37	2.04
16	46.49	40.31	13.20	6.96	82.82	3.19	1.96
17	45.25	40.18	14.57	6.90	81.83	4.91	1.77
18	44.42	41.72	13.86	6.33	76.55	8.39	1.38
Lifetime	56.29	36.59	7.12		87.92		2.21
							(1.30)

Note: N = 2,754 (34,893 person-years).

exponentiated value of the associated coefficient (exp(–1.77) = 0.17) indicates that conditional transfers to children were predicted to be 83 % smaller for parents who were continuously unpartnered compared with parents who were continuously together in the weighted pseudo-population represented by the IPT-weighted analysis. Time in a union with a new partner was statistically nonsignificant but still fairly large in practical terms (exp(–0.66) = 0.52). Thus, the IPT-weighted model suggests that time outside of a two-parent union, and particularly time as an unpartnered parent, negatively predicts the magnitude of financial transfers to children for education. Further, this association operates in part through the tendency for union status, family economic position as measured by educational attainment and employment status, and other dimensions of family complexity to mutually reinforce one another over time.

Figure 1 facilitates interpretation of results from Model 3. The dashed horizontal bar represents the average estimated parent-to-child transfer for education when a parent was continuously married or cohabiting with their child's other biological or adoptive parent—a value of $9,007. Bar graphs represent the magnitude of difference when a

Table 2 Time-invariant/baseline sample characteristics: Panel Study of Income Dynamics (PSID)

Characteristic	Mean/%
Transfer Amount for School ($) (mean)	8,624
	(27,626)
Parent Union Status at Baseline (%)	
Partnered with child's other biological parent	65.58
Single	33.30
Repartnered	1.13
Coresidence at Baseline (%)	
Parent and child coresident	91.98
Parent and child not coresident	8.02
Total Number of Children in Household at Baseline (mean)	2.07
	(1.13)
Parent Type (%)	
Father (male)	38.20
Mother (female)	61.80
Parent's Race/Ethnicity (%)	
Non-Hispanic white	51.38
Non-Hispanic black	46.26
Non-Hispanic Asian	0.07
Non-Hispanic other	0.40
Latino	1.89
Parent Age at Child's Birth (mean)	27.35
	(6.15)
Year Child Born (mean)	1989
	(3.61)
Total Household Income at Baseline ($1,000s of year 2000) (mean)	44.95
	(42.02)
Parent Is Household Head or Spouse at Baseline (%)	
Yes	86.06
No	13.94
Parent Employment Status at Baseline (%)	
Employed or temporarily laid off	61.91
Unemployed	9.59
Not in the labor force	28.50
Parent Years of Completed Schooling at Baseline (mean)	12.77
	(2.13)
Homeownership Status at Baseline (%)	
Owner-occupied	47.49
Rent or other	52.51

Notes: Numbers in parentheses are standard deviations. $N = 2,754$.

Table 3 Zero-inflated negative binomial regressions predicting parent transfers to young adult children born 1983–1995 for educational purposes: Panel Study of Income Dynamics (PSID), 1983–2013

	Model 1			Model 2			Model 3		
	B	SE	*p*	B	SE	*p*	B	SE	*p*
1. Log Odds of Receiving No Transfer									
Proportion of time in union status (vs. partnered with child's other biological parent), child ages 1–18									
Single							0.83	0.52	
Repartnered							0.53	0.51	
Child is not coresident[a]				–0.69	0.32	*	–0.50	0.54	
Total number of children in household[a]				0.03	0.08		0.09	0.10	
Parent is child's mother (vs. father)				–0.41	0.16	*	–0.52	0.20	*
Parent's race (vs. non-Hispanic white)									
Non-Hispanic black				1.15	0.15	***	1.05	0.18	***
Non-Hispanic Asian				–22.07	0.69	***	–20.61	0.82	***
Non-Hispanic other				0.90	0.97		2.29	1.21	
Latino				0.00	0.49		–0.55	0.50	
Parent's age at child's birth				–0.03	0.01	*	–0.03	0.02	
Year child born				–0.04	0.01	**	–0.03	0.02	
Parent is not household head or spouse[a]				0.58	0.67		–0.05	0.46	
Parent union status (vs. partnered with child's other biological parent)[a]									
Single	1.67	0.16	***	0.72	0.17	***	0.36	0.39	
Repartnered	1.06	0.25	***	0.56	0.26	*	0.01	0.90	
Family economic circumstances[a]									
Total household income ($1,000s of year 2000)				0.00	0.00		0.00	0.00	
Total years of schooling completed				–0.28	0.03	***	–0.37	0.05	***
Employment status (vs. employed or temporarily laid off)									
Unemployed				0.95	0.34	**	0.37	0.37	
Not in the labor force				0.60	0.17	**	–0.14	0.15	
Home is not owner-occupied				–0.26	0.18		–0.14	0.25	
Intercept	–0.13	0.10		8.82	1.67	***	9.35	1.89	***
2. Transfer Amount									
Proportion of time in union status (vs. partnered with child's other biological parent), child ages 1–18									
Single							–1.77	0.64	**
Repartnered							–0.66	0.56	
Child is not coresident[a]				0.06	0.24		–0.42	0.42	
Total number of children in household[a]				–0.11	0.06		–0.21	0.09	
Parent is child's mother (vs. father)				0.01	0.10		–0.26	0.16	
Parent's race (vs. non-Hispanic white)									
Non-Hispanic black				–0.60	0.21	**	–0.17	0.33	
Non-Hispanic Asian				–0.51	0.74		–0.72	0.79	
Non-Hispanic other				–2.20	0.19	***	–1.71	0.57	**

Table 3 (continued)

	Model 1			Model 2			Model 3		
	B	SE	*p*	B	SE	*p*	B	SE	*p*
Latino				0.53	0.27		1.66	0.54	**
Parent's age at child's birth				0.04	0.01	**	0.04	0.02	*
Year child born				–0.09	0.01	***	–0.10	0.02	***
Parent is not household head or spouse[a]				0.78	0.88		–0.81	0.40	
Parent union status (vs. partnered with child's other biological parent)[a]									
Single	–1.64	0.24	***	–1.12	0.22	***	–0.24	0.41	
Repartnered	–0.83	0.22	**	–0.79	0.18	***	0.46	0.57	
Parent economic circumstances[a]									
Total household income ($1,000s of year 2000)				0.00	0.00	*	0.00	0.00	
Total years of schooling completed				0.19	0.03	***	0.20	0.04	***
Employment status (vs. employed or only temporarily laid off)									
Unemployed				0.04	0.27		–0.33	0.33	
Not in the labor force				–0.19	0.17		–0.03	0.18	
Home is not owner-occupied				0.21	0.22		0.03	0.16	
Intercept	10.6	0.09	***	13.04	1.85	***	16.52	1.92	***

Notes: Data are weighted using the PSID sampling weight in Models 1 and 2, and using both the PSID sampling and inverse probability weights in Model 3. $N = 2{,}754$.

[a] Measured at child age 14 in Models 1 and 2; measured at baseline in Model 3.

*$p < .05$; **$p < .01$; ***$p < .001$

parent was unpartnered (dark gray bars) or repartnered (light gray bars) during 25 %, 50 %, 75 %, or 100 % of childhood. Values beneath each bar show the estimated value of the parent-child transfer under each condition. Dotted gray vertical bars show the 95 % confidence interval for each estimated value.

In a weighted pseudo-population that is balanced on the distribution of covariates across parental union status categories, parents who spent at least 25 % of the child's years unpartnered or at least 50 % of that time repartnered made significantly lower transfers to adult children for education compared with continuously partnered parents. The magnitude of estimated differences is substantial in practical terms: when a parent spent 50 % of a child's years unpartnered, the estimated transfer ($2,869) was nearly 70 % less compared with when a parent was continuously partnered; a comparable repartnered parent's transfer ($5,503) was about 40 % less.

Supplementary Analysis

The preceding analysis includes parents of young adults who never attended college. Smaller estimated financial transfers from unpartnered or repartnered parents to children for educational purposes may result from lack of need for such support if young adults raised outside of stable two-parent families were more likely not to extend their educational attainment after age 18. To account for selection into educational

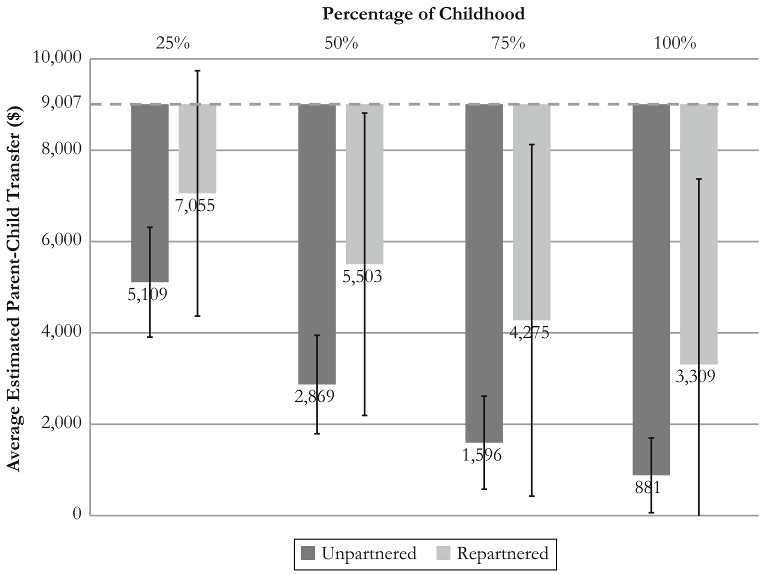

Fig. 1 Estimated parent-child transfer for education in early adulthood by parent's time in unpartnered or repartnered union status when child was age 1 to 18 compared with continuous coresidence with child's other parent: Panel Study of Income Dynamics (PSID), 1983–2013. N = 2,754.

attainment during adulthood, we restricted our analysis to the subset of parents who reported that their young adult child gained any educational attainment after high school or GED completion (N = 1,240). Results were consistent with those for the full sample of parents. Given any known postsecondary educational attainment, continuously partnered parents contributed an estimated average of $25,623 to young adult children, compared with estimated averages of $8,562 and $18,591 among parents who were unpartnered or repartnered, respectively, during one-half of their child's years before age 18. Thus, conditional on young adults' differential selection into postsecondary education, family composition history continued to contribute to parents' uneven financial transfers for related expenses.

We also considered parents' financial transfers exceeding $100 for purposes other than education to a young adult in the past year and since reaching age 18. Such transfers include gifts, loans, or contributions to expenses, such as a vehicle or rent. Compared with financial support for education, these transfers were relatively infrequent and smaller in value. To the extent that family composition history contributes to parents' general willingness or capacity to make transfers to children, we would expect the pattern of results pertaining to education financing to extend to transfers for other purposes.

Overall, differences in parents' estimated past-year and cumulative financial transfers to young adults for purposes other than education by family structure history did not achieve statistical significance in the marginal structural model framework. In general, though, time as an unpartnered parent was associated with *higher* transfers, and time as a repartnered parent was associated with *lower* transfers compared with remaining in a stable union with the child's other parent. We cautiously propose that parents' repartnered status is consistently associated with lower transfers to young adults, whereas time as an unpartnered parent may lead parents to direct resources to young adults for purposes other than education if those young adults are less likely to enroll or, given enrollment, are more likely to receive financial aid compared with youth who have stably partnered parents. Future research giving further consideration to these divergent patterns can provide guidance on how family structure conditions parent-child negotiations around financial transfers more broadly during the transition to adulthood.

Finally, the average parental transfer estimated from the model summarized in Table 3 is informed by both resident and nonresident parents. For comparison, Table A2 in the online appendix summarizes results from an IPT-weighted model restricted to the subset of parents who resided continuously with a child. Although not directly comparable with the results presented in Table 3, these results suggest that the experience of repartnering may have a weaker association with eventual transfer amounts when parents continue to coreside with children. This supports the expectation that part of the impact of stepfamily formation on parent-child transfers results from diminished contact with nonresident children after new union formation (Henretta et al. 2014).

Discussion

Contemporary family systems are varied in their characteristics and composition. Children in single-parent and stepparent families more often experience economic disadvantage across the early life course, with potential consequences for the intergenerational transmission of compromised status attainment. A substantial literature has sought to establish whether these observed associations are attributable to parents' characteristics that were present even before a child's birth (i.e., selection mechanisms) or to distinctive, ongoing processes in single-parent and stepparent families that constrain the investment of parents' economic resources in children.

This line of inquiry has been hampered by the very nature of family organization itself. In particular, complex family organization implicitly involves change in family structure over time, and such change alters the dynamics among resident and nonresident family members in ways that potentially transform the distribution of resources within families in the short and long term. Traditional regression-based techniques inadequately model the interplay between family composition and the family economic resources that both contribute to and arise from it. Further, much extant research lacks prospective measures of parental union status and family economic circumstances across the complete early life course, typically considers only those financial contributions provided by coresident parents, and rarely considers transfers to support the transition to adulthood.

We addressed these limitations by applying marginal structural zero-inflated, negative binomial regression models to long-running longitudinal data from the PSID to estimate parents' financial transfers to children for educational purposes during early adulthood as a function of parents' union status when children are between 1 and 18 years old. Our approach adjusts for time-varying confounding between parental union status on the one hand and coresident family composition and family economic circumstances on the other to remove selection effects and preserve the estimation of both indirect and direct effects in establishing the association of prior union status with subsequent financial transfers.

We report three main findings. First, as expected, in unadjusted estimates, parents who were unpartnered or repartnered during their children's adolescence (age 14) made less frequent and smaller transfers to support a child's education in early adulthood compared with when they were partnered continuously with a child's other biological parent. Given that today's young adults spend nearly half of childhood outside of a two-parent household, this pattern highlights a potentially significant driver of inequality in contemporary postsecondary educational attainment.

Second, estimates from the IPT-weighted models demonstrate that the disparity in transfer amounts increased with time as an unpartnered or a repartnered parent. Unpartnered parents were predicted to make transfers to children that were between 44 % and 90 % smaller compared with parents who lived continuously with a child's other parent, depending on how long the parent was unpartnered. Repartnered parents were predicted to make transfers that were between one-third and two-thirds smaller compared with continuously partnered parents. These estimates provide a relatively unbiased estimate of the average parental household transfer in the context of family complexity informed by both resident and nonresident parents.

Third, the IPT-weighted models show that parental union status has an enduring association with parents' financial support for education after baseline selection characteristics and time-varying confounding are adjusted for. We interpret the model results as evidence in support of the argument that both selection mechanisms and causal mechanisms operate to shape the association between parental union status and parents' financial transfers to adult children. That is, prior family socioeconomic disadvantage likely contributed to parents' entry into or persistence in a given union status, and union status in turn shaped the economic resources available in families to support young adults' education over time. This finding is consistent with prior literature showing that parents' repartnering, nonresidence with children, and subsequent fertility after union dissolution impact how resources are distributed within and transferred between parental households (Berger et al. 2012; Carlson and Berger 2013; Tach et al. 2014), but no prior work has demonstrated the durability of this pattern over the early life course to shape how parents invest in the transition to adulthood in a contemporary cohort of young adults.

In sum, this work demonstrates that the relationship between complex family organization and parents' investments in children is the product of both selection processes and causal mechanisms that cumulate over children's early life course in ways that are consequential for the transition to adulthood and potentially for long-term status attainment. We advance research in this field by highlighting the fluid and dynamic nature of family complexity and through the use of rich, long-running longitudinal data and rigorous analytic techniques.

Nevertheless, we note some limitations. First, our approach assumes that the impact of parental union status on eventual transfers is constant across all years of a child's life. A more developmental approach might consider whether parental union status early in the life course (when parents might initially invest in children's postsecondary education) or in adolescence (when recovering from shocks to a family economy might be difficult) are most consequential. An extension to the current work could capture developmental effects by considering parental union status during early childhood, middle childhood, or adolescence separately.

Second, our measure of other coresidential children in a parent's household as a dimension of family complexity is not fine-grained. Research on multipartner fertility and sibling complexity has highlighted relatedness among children as a driver of inequality in the distribution of resources within a family system (Henretta et al. 2014). Hence, we might wish to focus on the number of children in a parent's household who are stepsiblings or half-siblings to a focal child. However, such specificity is challenging to exploit in the current modeling framework because some sibship arrangements—particularly stepsibship—occur rarely in two-parent and single-parent PSID households and thus are unlikely to be time-varying confounders of those union status categories. Further, until 2013, the PSID birth history did not collect information about the second parent of children born to unpartnered respondents. As a result, it is not always possible to discern whether children share the same second parent or to develop a definitive measure of half-sibship. Here we regard the number of other children in a parent's household as an indicator of potential competition for resources and recognize that variation in sibship composition by parental union status may be concealed.

Finally, as with any analysis based on observational data, the estimated effect of family complexity on parental investments in adult children's postsecondary education may be biased by unmeasured confounders, such as birth intendedness; child and parental health status; and ecological factors, such as neighborhood and school quality. We sought to account for the factors that have been theoretically and empirically associated with both family complexity and parents' investments in young adults' education in prior research, but we recognize that this list may not be exhaustive.

Despite these limitations, this work informs the sociology of family and status attainment literatures by clarifying how selection mechanisms and family process influence each other over time to shape the distribution of resources to children raised in complex families as they enter adulthood themselves. This perspective is critical for identifying young adults in need of external financial support for postsecondary education and the transition to adulthood more broadly, as well as for understanding the circumstances that give rise to this need.

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