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Abstract

Men's and women's economic resources are important determinants of marriage timing. Prior demographic and sociological literature has often measured resources in narrow terms, considering employment and earnings and not more fine-grained measures of job quality. Yet, scholarship on work and inequality focuses squarely on declining job quality and rising precarity in employment and suggests that this transformation may matter for the life course. Addressing the disconnect between these two important areas of research, this paper analyzes data on the 1980-1984 U.S. birth cohort from the National Longitudinal Survey of Youth 1997 to examine the relationships between men's and women's economic circumstances and their entry into marital or cohabiting unions. We advance existing literature by moving beyond basic measures of employment and earnings and investigating how detailed measures of job quality matter for union formation. We find that men and women in less precarious jobs – jobs with standard work schedules and jobs that provide fringe benefits – are more likely to marry. Further, differences in job quality explain a significant portion of the educational gradient in entry into first marriage. However, these dimensions of job quality are not predictive of cohabitation.

Introduction

Since the 1970s, the economic opportunities for those with less than a college degree have deteriorated: earning power, job security, and jobs with good benefits have diminished, while precarious employment has become more prevalent (Kalleberg, 2009; Fligstein and Shin, 2004). Over the same period, entry into first marriage has declined precipitously and marriage has become increasingly stratified by class, with more educated men and women now more likely to marry than their less-educated counterparts (Wang and Parker, 2014; McLanahan, 2004; Ellwood and Jencks, 2004; Goldstein and Kenney, 2001; Schneider, Harknett, and Stimpson, 2018).

While a great deal of social scientific research has examined the influence of employment and earnings on family formation and in explaining the stark differences across education groups in family formation, this literature has taken a narrow approach to measuring employment. Unemployment and low earnings clearly matter for family formation (i.e. Burstein, 2007) and contribute to educational gaps in marriage (Harknett and Kuperberg, 2011), yet these measures do not capture important dimensions of job quality that have changed over time and are increasingly stratified by education.

Yet, it is precisely these aspects of employment contracts that are the focus of much scholarly and public discussion of the changing American economy (i.e. Kalleberg, 2011; Steverman, 2014). For instance, Hacker (2006) describes the transfer of risk from large institutional actors such as employers to households and workers – seen in the erosion of employment benefit packages. Along similar lines, scholars have also recently called attention to another transfer of risk from employers to employees seen in the rise in on-call and variable schedules for hourly workers by which employers effectively transfer payroll risk from firm to worker (Lambert, 2008; Boushey, 2016). This erosion of job quality has been most pronounced for those with less than a college degree (Kalleberg, 2009).

We use data from the National Longitudinal Survey of Youth 1997 to examine the role of job quality in union formation and in explaining stark educational differences in marriage. First, we assess the extent to which job quality matters for entry into marriage and cohabitation. We also extend the rich existing literature concerned with gender differences in the relationship between employment and marriage by investigating whether job quality operates differently on marriage for men and women. Finally, we examine whether accounting for measures of job quality can explain educational gradients in entry into first marriage.

Theory and Prior Research

We first discuss how employment affects entry into marriage and cohabitation, for men and for women, and discuss educational gradients in both marriage and cohabitation and prior approaches to explaining these gradients. We then turn to the literature that focuses more squarely on job quality and union formation before integrating this previous theory and empirical research to derive a set of hypotheses about how job quality is likely to shape union formation, matter differently by gender for union formation, and account for educational gradients in union entry.

Men's Employment and Entry into Marriage

Sociologists and demographers theorize that men's economic resources are a mark of marriageability (Davis and Blake, 1956; Wilson, 1987; Edin and Kefalas, 2005) and so make men with such resources more attractive as potential partners and may make men themselves feel more ready to marry. Research has shown that men's employment status and earnings are strongly related to transitions to first marriage (see comprehensive reviews by Burstein, 2007 and Ellwood and Jencks, 2004). Studies using a variety of data sets have consistently found that men are more likely to marry when they are employed (Sweeney, 2002; Sassler and Goldscheider, 2004; Harknett and Kuperberg, 2011; Harknett and McLanahan, 2004) and especially when they are employed full time

(Oppenheimer et al., 1997; Oppenheimer, 2003; Schneider, 2011; Shafer and James, 2013; Kuo and Raley, 2014). Similarly, men who earn more also have a higher risk of marriage (Sweeney, 2002; Clarkberg, 1999; Oppenheimer et al., 1997; Oppenheimer, 2003; Schneider, 2011; Shafer and James, 2013; Schneider and Reich, 2014; McClendon et al., 2014; Kuo and Raley, 2014).

Women's Employment and Entry into Marriage

There has been some debate regarding whether women's work and earnings would operate similarly to men's and increase entry into marriage (Oppenheimer, 1988), or would instead grant women a degree of autonomy that would discourage marriage (Becker, 1981). The weight of the research evidence suggests that in older cohorts, women's employment and earnings might have acted as a deterrent or substitute for marriage, either because economically independent women were empowered to forego marriage if that was their preference, because high-achieving women were violating strong norms encouraging male breadwinners and female homemakers and were therefore less attractive on the marriage market, or because non-marriage created an imperative for women to work more and earn more (Burgess et al., 2003; Xie et al., 2003).

For more recent cohorts, however, women's employment and earnings have come to be positively associated with entry into marriage. Research finds a trend toward gender convergence in the influence of employment and earnings on marriage (Clarkberg, 1999; Lichter et al., 1992; Shafer and James, 2013; Schneider, 2011; Schneider and Reich, 2014; Sweeney, 2002). More recent work using the 1980-1984 birth cohort captured in the NLSY97 finds that the positive relationship between women's earnings and marriage persists (McClendon et al., 2014; Kuo and Raley, 2014) as does the relationship between marriage and full-time employment (Addo, 2014). However, while the weight of evidence suggests that women's employment is positively associated with marriage, there is still evidence that the strength of this association is generally weaker than for men (Burstein, 2007; Shafer and James, 2013).

Influence of Employment on Entry into Cohabitation

Comparing marriage and cohabitation can bring the pathways through which employment influences union formation into sharper relief. If the main pathway through which employment affects union formation is that economic resources are needed to establish a joint household, then employment should be as influential for cohabitation as it is for marriage.

However, prior theory and research have considered marriage and cohabitation to be distinct institutions. Marriage has been seen as more fully "institutionalized," meaning that marriage has a clearer and more rigid set of norms and expectations compared with cohabitation (Nock, 1995). Because of the stronger normative expectations associated with marriage, including that marriage is more likely to come with the expectation of a long-term commitment, economic determinants are expected to play a larger role for marriage than for cohabitation.

Further, family scholars have theorized that contemporary marriage has a strong symbolic significance, as a "capstone" after economic milestones are achieved (Cherlin, 2004; Edin and Kefalas, 2005). Although in theory couples could marry before they achieve economic stability, family scholars have argued that marrying with scant financial resources is stigmatized and perceived as risky because of the potential for economic insecurity to destabilize the relationship (Gibson-Davis, Edin and McLanahan, 2005). In contrast, financial strain appears to be at least one factor that precipitates cohabitation (Sassler, 2004; Sassler and Miller, 2011; Sassler and Miller, 2017). Considering the connection between economic resources and cohabitation more broadly, Perelli-Harris and colleagues (2010) suggest that in many contexts, cohabitation really is an institution in response to economic insecurity.

Prior demographic research largely accords with this qualitative work. While earnings and employment are important positive determinants of entry into marriage, the effects on cohabitation are much less pronounced. Several studies find no association between employment and entry into cohabitation (Raley, 1996; Sassler and Goldscheider, 2004; Carlson, et al., 2004; Manning et al., 2014). But, in this line of research, an important exception is the work of Clarkberg (1999) which finds positive associations between earnings and cohabitation.

Prior research on gender differences in the association between employment and earnings and cohabitation is much more limited. In general, there is evidence of null effects of employment and earnings on cohabitation (Raley, 1996; Sassler and Goldscheider, 2004; Carlson, et al., 2004; Manning et al., 2014), though again an important exception is Clarkberg (1999) who finds that both men's and women's earnings positively predict cohabitation (vs. staying single) and that the effect is actually stronger for women than for men.

Educational Gradients in Union Formation

Marriage is strongly graded by educational attainment, with more highly educated individuals more likely to marry (Isen and Stevenson, 2010; Thornton et al., 2007). While this has been true for men since at least the middle of the 20th century, women's educational attainment, once negatively associated with marriage, has now become a positive predictor (Schneider, Harknett, and Stimpson, 2018; Goldstein and Kenney, 2001). In contrast, education is negatively related to cohabitation (Bumpass and Sweet, 1989; Sassler and Goldscheider, 2004; Thornton et al., 1995; Perelli-Harris, et al., 2016). Education may be positively related to marriage because it is a signal of long term economic prospects and stability (Sassler and Goldschneider, 2004) or because education fosters a cultural approach to parenting that prioritizes investment in children through marriage (Lundberg and Pollak, 2015). But, more simply, more highly educated men and women may also be more likely to marry because they possess the economic resources that have long been the normative pre-requisites of marriageability (Ishizuka, 2018).

In that case, the unequal distribution of economic resources by education in the United States has the potential to "explain" stratification in marriage entry along this axis of disadvantage.

That is, one reason why men and women with a college degree may be more likely to marry than their less educated counterparts is because education affords these individuals access to the economic resources that make them marriageable.

Job Quality and Union Formation

This existing body of empirical research has carefully shown that employment and earnings influence marriage and cohabitation and that those basic measures of work status may operate differently for men and for women. Both men's and women's employment and earnings have changed a great deal over the past several decades (i.e. Wilson, 1987; Oppenheimer, 1988). But, these decades have also witnessed profound transformation in the quality of jobs beyond simply levels of employment or pay. Along several dimensions, job quality appears to have deteriorated and, importantly, stratified along the axis of educational attainment over the past several decades (Kalleberg, 2009; Fligstein and Shin, 2004).

Prior literature identifies several important dimensions of job quality. First, at the most basic level, the structure of compensation plays a large part in defining what a "good job" is, with salaried positions offering greater flexibility and often greater compensation than hourly positions. Being salaried offers the additional benefit of a fixed and predictable amount of earned income each month, whereas those paid hourly are subject to greater income volatility. Education is strongly related to compensation type: almost two-thirds of salaried workers have a college degree and almost 80% of hourly workers having less than a college degree (Brenan 2017).

Multiple job holding is another indicator of job quality. Multiple job holding has been shown to be motivated by the need for more income and to be an adaptive response to economic insecurity (Hipple 2010, Zangelidis 2014). Although those with lower levels of education can be expected to be more in need of supplemental income and more subject to economic insecurity, in fact, multiple job holding appears more common for those with a college education (Lalé, 2015). Multiple job holding, then, is one notable exception to the typical pattern of higher education being associated with better job quality.

Beyond the structure of pay, Kalleberg (2011) notes that a defining feature of a "good job" is the provision of fringe benefits such as health insurance, retirement accounts or pension, and paid time off. The availability of these benefits has declined for all workers over time, but especially for less educated workers (Kalleberg, 2011). For instance, employer provision of health insurance has declined, especially for those with less education (Farber and Levy, 2000; EPI, 2012). Retirement benefits have also changed substantially, shifting from defined benefit to defined contribution plans and eroding all together as well (Hacker, 2006). This change too has been stratified by education (Kalleberg, 2009). There is also substantial stratification in paid time off, with more educated workers much more likely to have paid vacation, paid sick days, and paid parental leave than their less educated counter-parts (Glynn, Boushey, and Berg, 2016).

In addition to non-monetary compensation, job quality is also defined by work schedules. Scholars have long been concerned with non-standard work hours (i.e. Presser, 1999), but more recent research identifies variable work hours as a key dimension of precarious work. Many hourly employees with low educational attainment now work varying numbers of hours each week, often with different starting and stopping times and on different days each week (Henly, Shaefer and Waxman, 2006). This experience of unstable and unpredictable schedules is sharply stratified by education as many higher-SES workers have more stable schedules, or have more employee-control over their work schedules – a desirable flexibility in contrast to instability (Schieman and Plickert, 2008).

Finally, labor union membership may confer important economic benefits on workers in the near term and may also signal greater economic stability in the future (Schneider and Reich, 2014). While labor union membership has declined precipitously over the past several decades, approximately 11% of workers remain union members (Dunn and Walker, 2016). Those with some college and a college degree are more likely to be represented by a union than those with a high school degree only or less education, in keeping with the general pattern that higher educational attainment is associated with better job quality (Schmitt and Warner, 2009).

In sum, by many measures, employment has become more precarious and the educational divide in job quality appears pronounced and, in many instances, to have increased over time.

Consequences of Job Quality for Marriage and Cohabitation

Scholars of work have conjectured that the decline in job quality is likely to affect family formation and stability (Kalleberg, 2011, 2009), and demographers often note that economic precarity may be a determinant of family formation (Lichter, Qian, and Mellott, 2006; Cherlin, 2015). This perspective derives from the expectation that job quality may increase marriage through several pathways. First, to the extent that individuals with more economic resources – such as fringe benefits – are more marriageable, better quality jobs may simply be more economically valuable and so may increase marriageability. Second, in an era of increasingly precarious employment, having a job that offers fringe benefits or a stable schedule may serve as a marker of status and of the achievement normatively necessary for marriage (Cherlin, 2004; Schneider, 2011). Third, while income may convey information about current economic status, other aspects of job quality – such as fringe benefits or a union contract – may convey information about future prospects and economic stability that may be additionally valuable for marriage (Schneider and Reich, 2014).

In contrast to marriage, prior literature suggests that employment, and so, we would expect, job quality, is not strongly related to entry into a cohabitating union and that more economically-precarious individuals may actually enter into cohabitations rather than into marriage either because they lack the normative pre-requisites for marriage (Cherlin, 2004) or because the economies of scale of co-residence are a partial solution to financial fragility (Sassler, 2004; Sassler and Miller, 2011).

These theoretical perspectives on why job quality might matter for marriage and might matter more for marriage than cohabitation are largely divorced from the particulars of given indicators of job quality. Rather, they suggest that holding a better job – as captured by a set of indicators – conveys information about one's own and one's potential partner's long-term economic security and prospects.

But, there are good reasons to expect that certain dimensions of job quality may matter more for marriage than for cohabitation. One reason is that the legal distinction between marriage and cohabitation makes some benefits more valuable to spouses than to unmarried partners. Specifically, employer provided health and dental insurance is often structured in such a way as to directly benefit not just the employee, but also his or her spouse and dependents. While there are important exceptions for domestic partners (Polikoff, 2012), in general this benefit will only have a mechanical transferability to married partners and children that might render it particularly valuable for marriage versus cohabitation. This idea finds credence in the literature on health insurance and divorce where scholars have found evidence of "marriage lock" in which divorce is reduced because it would disrupt spousal health insurance (Sohn, 2015).

Another reason is that some aspects of job quality are only valuable over a long time-horizon. Specifically, the economic value of employer-provided retirement savings and of life-insurance is generally only realized later in life, which is decades away from the time of marriages that we observe in the NLSY97. While cohabitation could be a long-term relationship, demographic research makes clear that marriages are of significantly longer duration than cohabiting unions (Cherlin, 2010) and that married partners expect to remain together for far longer than cohabitating partners (Nock, 2005).

However, the demographic literature directly linking job quality with union formation is sparse. Using data from the NSLY79, Schneider and Reich (2014) find that union membership is a significant predictor of first marriage for men – a relationship they attribute to the health insurance benefits and longer job tenure that union membership can help workers secure. Drawing on data from the NLSY97, Kuo and Raley (2014) find that occupational autonomy is positively associated with marriage entry for women in their late 20s, but not for younger women and not for men, after controlling for earnings and employment. McClendon et al. (2014) find that paid parental leave is a significant predictor of marriage for women in the NLSY97 through round 13 (age 29). Piotrowski, Kalleberg, and Rindfuss (2015) take up the Japanese case and show that employment in non-regular jobs significantly reduces the risk of marriage entry relative to those working regular jobs, with the largest effects for men.

Even less work has examined how job quality affects entry into cohabitation. Of notable exception, using data from the NLSY79, Oppenheimer (2003) finds that economic instability is negatively related to marriage but positively related to cohabitation and, similarly, Clarkberg (1999) finds that having had more jobs and shorter job tenure are positively associated with cohabitation (though not with marriage) compared with remaining single. In a study measuring future earning potential, Xie et al. (2003) find that men's expected future earnings predict entry into marriage but not into cohabitation. Further evidence that economic insecurity is a deterrent to marriage but not to cohabitation comes from research showing that credit card and student loan debt are associated with entry into cohabitation but with delays in marriage (Addo, 2014).

In sum, these important exceptions aside, there has been very little quantitative, demographic effort to evaluate how job quality shapes marriage and cohabitation entry. But, given this forgoing work, we would expect that:

Hypothesis 1: Individuals who hold higher quality jobs will have a higher risk of marriage.

Hypothesis 2: Job quality will either negatively predict entry into cohabitation (compared with remaining single) or have a null relationship.

Prior research shows that income and employment are positively related to marriage entry and, for at least the past several decades, have had a positive association with marriage for both men and women. Theory and a limited amount of prior empirical research also suggest that job quality matters for marriage for both men and women. We hypothesize that job quality should be positively related to marriage for both men and women.

However, particular aspects of job quality may have stronger associations with marriage for women than for men. One of these is having access to a standard work-schedule. Women retain primary responsibility for domestic production and for childcare (Lyonette and Crompton 2014) and prior research has shown that non-standard and unstable and unpredictable work schedules make it particularly difficult to fulfill those gendered responsibilities (Presser 2005; Henly and Lambert 2014; Carillo et al., 2017). We might expect then that having a standard day schedule would have a stronger positive association with marriage for women than for men. In addition, we predict that "family-friendly benefits" such as parental leave, child care, and schedule flexibility may be particularly valuable for women, given that women typically shoulder a disproportionate share of caregiving responsibilities.

> *Hypothesis 3: Both men's and women's job quality will be positively associated with entry into marriage. However, standard work schedules and "family friendly benefits" will more strongly predict entry into marriage for women than for men.*

Given the paucity of evidence that economic resources are significantly linked to cohabitation, we do not offer a hypothesis on how gender may interact with economic influences on cohabitation. Rather, we consider our comparison of the relationship between job quality and cohabitation for men and for women to be an exploratory analysis, for which any significant findings would need to be replicated in future research.

Finally, there are well documented educational gradients in marriage entry and cohabitation and, as discussed above, there are also steep gradients in job quality by education. Higher educational attainment, and especially a college degree, is associated with better job quality on all indicators with the exception of multiple job holding. Higher educational attainment is also positively associated with entry into marriage and with relatively lower levels of nonmarital cohabitation. Job quality may confer both real and perceived economic security, and features of job quality such as fringe benefits can be considered economic resources. Given the robust literature showing that employment and earnings encourage marriage, we could reasonably expect differences in job quality by education to contribute to the higher rates of marriage for those with college education who enjoy the best quality jobs. Therefore, we expect that accounting for a richer set of job quality measures will diminish educational differences in marriage rates.

> *Hypothesis 4: The greater likelihood of marriage for more highly educated individuals will be partially accounted for by differences in job quality*

Data and Methods

Data

We investigate the relationship between job quality and union formation using rounds 1-16 of the National Longitudinal Survey of Youth 1997 (NLS97). The NLSY97 has followed a nationally representative sample of 8,984 youth born 1980-1984 (with an oversample of Hispanic or Latino and Black youth) with annual interviews from 1997 to 2011 and biennial interviews thereafter.

The NLSY97 provides a unique opportunity to study the impact of job quality on marriage and cohabitation entry. Rich employment data has been collected at fine-grained time intervals, allowing for an examination of the effects of labor market position that extends beyond employment and earnings to a range of job quality measures – fringe benefits, compensation structure, union membership, multiple job holding, and schedule regularity. Although respondents in the sample have only reached their early thirties, most have already entered first cohabitation or first marriage (approximately 74%).

Our sample includes all never-married-or-cohabited respondents over the age of 18 with non-missing data interviewed through round 16 of data collection (retention rate at round 16 of the study is 79.5%). Respondents enter the risk set at age 18 because labor market participation of youth is still quite low prior to exiting high school (only 49% of 17 year-old NLSY97 respondents were employed, and only 58% were employed or looking for work). The NLSY provides weekly employment data but monthly data on union formation and school enrollment, so our unit of analysis is the person-month. We draw on data from 4,162 men and 3,735 women. We observe a total of 336,535 person-months for men and 257,732 person-months for women before entry into either marriage or cohabitation. We also draw on a sample of 469,178 person-months for men and 402,448 person-months for women observed before entry into marriage (but that could be observed following a cohabitation).

Measures

Union Formation. We define two dependent variables that we employ in a competing risk analysis of union formation: entry into first marriage directly from non-coresidential status and entry into first cohabiting union from non-coresidential status. The NLSY provides data on the calendar year and month in which each applicable respondent entered their first marriage or their first cohabiting union. For the competing risks analysis, we code a variable to be 0 for all person-months prior to union entry, 1 for the month in which a first entry to marriage from single-status took place, and 2 for the month in which a first entry to cohabitation from single-status took place, with respondents censored once the event has occurred. We also construct a dichotomous measure of entry into first marriage where respondents are not censored at cohabitation and are coded as entering marriage regardless of whether they transition from non-coresidential status or from a cohabitation. Individuals who neither marry nor cohabit are censored at last interview/observation.

Table 1 summarizes these transitions for the observed sample, separately for men and women. We see that 14% of men and 16% of women in the sample transitioned directly to marriage from a non-coresidential state. Substantially larger shares – 27% of men and 33% of women – transitioned to marriage following cohabitation and about 29% of both men and women transitioned to cohabitation, but were not observed to enter a marriage. The remaining 30% of men and 22% of women never married or cohabited in the observation period.

Employment and Job Quality. To capture variation in job quality, we construct a set of time-varying individual-level measures of economic characteristics. In each survey round, the NLSY collects a battery of information on all jobs held by the respondent between the current and the previous survey round. We used the start, stop, and gap dates provided by the NLSY97 for each job to create weekly measures of employment data. For weeks in which respondents worked at more than one job, job characteristics from the respondent's "main job," identified by the NLSY, are used.[1] By averaging employment measures across weeks within each month, we collapse this weekly data to person-months in preparation to be merged with the monthly data on union formation. However, it is important to note that these weekly (and monthly arrays) by design do not capture any short-term within-job variation in our measures of job quality. Within-person variation in job quality is only observed when respondents change jobs or when new information on a job is collected in a subsequent round of the NLSY.

Our basic measure of employment captures respondent's earnings, where respondents who are not working have zero earnings. This measure of earnings is driven in part then by whether the young adults in our sample are employed or not. Because earnings and employment are therefore highly correlated (ρ=0.95), we do not include a separate measure of employment, and we interpret the earnings variable as a measure of both employment and earnings. The value of earnings is

[1] For employed person-weeks with no "main job" indicated, we designate the job with the largest number of hours as the main job.

defined based on the NLSY-calculated effective hourly rate of compensation, which includes earnings from overtime and performance pay. This hourly rate is multiplied by the number of hours worked to obtain weekly earnings. Earnings are inflation-adjusted to be expressed in 2013 dollars and logged.

We are able to leverage the rich data on jobs in the NLSY97 to examine job quality along five dimensions: (1) multiple job holding, (2) work schedule type, (3) compensation structure, (4) coverage by union contract, and (5) access to employer-provided benefits including health insurance, parental leave, and flexible scheduling.

First, we use the detailed weekly job calendar data to construct a measure of whether the respondent holds more than one job at a time. Second, we code three types of work schedules that deviate from a standard day shift: (1) non-standard shifts involving work in the evenings, at night, or on the weekend, (2) a split or rotating shift, or (3) an irregular schedule. Third, the NLSY97 collects hourly pay-rate data for those working at a job that pays by the hour – we use this variable to create an indicator of whether a job paid an hourly wage rather than a salary. Fourth, the NLSY97 also provides a dichotomous measure of whether a job is "covered by a contract that was negotiated by a union or employee association" that we use to define union jobs.

Finally, fifth, for each job listed by the respondent, the NLSY asks whether the employer provides a battery of fringe benefits: health, life, and dental insurance; paid and unpaid maternity leave; childcare; a "flexible work schedule"; a "retirement plan"; tuition reimbursement; and an employee stock ownership plan. We combine this series of dichotomous indicators for job benefits into a simple additive scale that has high reliability (Cronbach's alpha = 0.90). This scale is referred to as the "fringe benefits scale" in our tables and discussion of results. We also estimate models that include each of the benefit items separately, as well as models that examine an abridged additive scale that sums what we term "family-friendly" benefits: paid or unpaid parental leave, child care, and schedule flexibility. As with earnings, we capture benefits reported through the "main" job.

Education. We utilize the NLSY97's monthly educational history data to calculate a school enrollment indicator and a five-level variable that measures the respondent's highest attained degree (having completed less than 12 years of education, having received a high school diploma, having attended some college but no degree, having an Associate degree, or having received a Bachelor's degree or more education). The results are substantively similar when we combine the categories of some college and Associate degree.

Control Variables. Our analyses account flexibly for life-course effects by including a linear and quadratic term for the age of respondents in months, which is calculated as the difference between a respondent's birth month and the month of analysis. We also use dummy variables for the two recessions that occurred during our time period – the recession from March to November of 2001 and the Great Recession from December 2007 to June 2009 – in order to address any effect these economic shocks may have had on union formation. We also include measures of the number of children in the household and whether the respondent has recently experienced the birth of a child. Table 1 shows that almost half of men and 43% of women report a birth that precedes marriage or cohabitation. These time-varying characteristics are summarized for our analytic sample in Table 2.

We measure race-ethnicity using a four-category variable indicating whether a respondent identifies as Black, non-Hispanic; White, non-Hispanic; another race non-Hispanic; or Hispanic. Finally, we control for family background using a measure of respondents' mothers' educational attainment (collected at baseline).

Method

We estimate two main statistical models. First, we estimate a discrete time competing risks event history model of entry into first marriage or first cohabitation to assess the role employment precarity plays in union formation. The model is estimated with multinomial logistic regression. Respondents are censored when they either enter first marriage directly from a non-coresidential status or when they enter a first cohabiting union directly from a non-coresidential status. The estimated coefficients are the risk of entry into each of the two union types relative to remaining in a non-coresidential state. We also present tests of the significance of differences in the estimated coefficients for entry into marriage vs. cohabitation. Second, we estimate a discrete time event history model of entry into first marriage, irrespective of whether the respondent transitions from a non-coresidential state or from a cohabiting union. In this model, respondents are only censored at first marriage, not at first cohabitation, and the model is estimated with logistic regression.

In both models, standard errors are clustered at the level of the respondent. All time-varying covariates, with the exception of age, are lagged in order to ensure they are measured before the outcome events occur. In accord with our theoretical model, which assumes that sustained exposure to employment and job quality influence union transitions, we use 12-month averages of employment and job quality covariates. To ensure appropriate time-ordering with respect to union formation, we lag these predictors by 6 months, using average monthly employment and job quality data over the period 17 months to 6 months prior to the month the outcome is measured.[2] Following NLSY97 official guidance, we do not weight the analyses (NLSY, 2017).

[2] We also checked the robustness of the results to using a lag that measured the covariates at a point in time (rather than the 12 month average of the period 6-17 months prior). We tested using lags that were 12 months prior and 6 months prior to the event month. These models show very similar results. However, for men, the fringe benefits scale is a weaker predictor of marriage when using 12 month lags but a stronger predictor of marriage when using 6 month lags compared with the preferred models. Additionally, for both the 12 and 6 month lag, the benefits scale coefficient predicts cohabitation entry more strongly than in the main models.

Our analysis proceeds in three steps. In the first step, we focus on the association between our measures of job quality and union formation. Here, we begin by estimating the following model:

$$\log\left(\frac{p_{it}^{(u)}}{p_{it}^{(0)}}\right) = \alpha + \mu REC + \gamma X_i + \delta X_{it} + \beta E_{it} + \delta JQ_{it} \qquad (1)$$

We predict the probability of union formation, either through first marriage or first cohabitation, with a discrete time competing risks model. The model includes the time-varying measure of earnings/employment and measures of job quality, in addition to controls for time-invariant and time-varying individual socioeconomic and demographic characteristics and indicators of the two recessions. Because the measures of job quality are highly correlated, we estimate the model five times, entering each of the job quality measures separately (along with the other covariates). We also estimate the models separately for men and women. We then repeat this same approach but with a model that does not treat cohabitation as a competing risk and simply focus on predicting marriage, with a time-varying indicator of cohabitation status.[3]

In the second step of the analysis, we test for gender differences. We cannot easily test for differences in coefficients across separately estimated models or for interactions within a pooled model using the logistic regression models described above (Mood, 2010). While the competing risk models rely on multinomial logistic regression, the models of entry into marriage from either non-coresidential status or cohabitation can be estimated with a linear probability model (LPM), which does allow for interpretation of interaction coefficients (Angrist and Pischke, 2009). We then test for differences by gender in the association between job quality and entry into marriage using interactions between gender and job quality in the LPM.

[3] We also re-estimated the models restricted to respondents who were cohabiting in the month prior to measurement of the dependent variable. The results are unchanged despite a large reduction in sample size: working a split/rotating shift remains negatively associated with marriage for women, hourly work remains negatively associated with marriage for men, and fringe benefits remain positively associated with marriage for both men and women. These predictors of first marriage entry are similar whether respondents enter marriage from cohabitation or not.

Third, the same limitations of the logistic regression model that make it difficult to compare coefficients on separately estimated models for men and women also prevent us from easily comparing coefficients on the same measures across nested models. However, to test hypothesis 4, we need to compare if the coefficients on our set of indicators for educational attainment are significantly attenuated when we account for our measures of job quality. To make these comparisons, we make use of the method developed by Karlson, Holm, and Breen (2012) which allows us to make unbiased estimates of the extent to which our job quality measures mediate the association between union formation and education. Here, we assess how much the controlled total effect of education (adjusted for our controls as well as earnings/employment) is mediated by the inclusion of the full set of job quality measures. We present coefficients on education that summarize the total effect, direct effect, and indirect effect. We also show the percentage of the total effect that is attributable to the indirect pathway of job quality and tests of the statistical significance of that mediation.

Results

Job Quality and Entry into First Marriage

We begin by examining the relationship between our measures of job quality and union formation. The key results are presented in Table 3, with each of the panels, numbered 0 to 5, containing a separate regression model specification, varying the key independent variable that is included. The first two columns show the estimates for men from the competing risks models. In instances where the measure of job quality significantly affects the risk of transitioning to marriage versus remaining in non-coresidential status (with cohabitation as a competing risk) or affects the risk of transitioning to cohabitation versus remaining in non-coresidential status (with marriage as a competing risk), the coefficients are marked with asterisks. We underline the coefficients when the association between entry into marriage versus cohabitation is significantly different. The third column shows how job quality is associated with entry into marriage from either non-coresidential status or from cohabitation. We present identical models for women in the fourth through sixth columns.

As predicted by Hypothesis 1, men's job characteristics predict entry into marriage. While neither multiple job holding nor union coverage alters the risk of first marriage, we see that men who work a non-standard evening, night, weekend shift, a split, or a rotating shift are significantly less likely to marry compared with men who work a standard daytime shift, when cohabitation is a competing risk.

We also see that, controlling for a set of background characteristics and for earnings/ employment, the fringe benefits scale is positively and significantly associated with entry into first marriage ($b = 0.6561$, $p < 0.01$). Figure 1 (Panel A) plots the predicted probability of marriage entry in a given person-month by the extent of men's fringe benefits. Men who have the mean number of benefits have about a 0.0015 probability of transitioning to first marriage in a given month. Men with a standard-deviation greater number of benefits are 15% more likely to marry in a given month.

We also find support for Hypothesis 2. In Table 3, for men, none of the measures of job quality are significantly associated with entry into cohabitation (with marriage as a competing risk) and, further, the differences between these non-significant coefficients for cohabitation and the significant coefficients for marriage are in fact significant in the case of non-standard schedules and fringe benefits (denoted by the underlined coefficients).

Further in Table 4, across the various benefit types, we find some support for our expectation that health/dental insurance and retirement benefits would be more strongly related to marriage than cohabitation. For men, we see that health insurance and dental insurance are significantly related to marriage, but not cohabitation, and that the difference between the two is statistically significant. We see a similar relationship for life insurance. The results for retirement do not fully accord with our hypothesis. While retirement benefits are significantly related to marriage, they are also significantly related to cohabitation. While the coefficient is substantially larger for marriage, it is not significantly different. Of all of the specific benefits, only retirement and unpaid parental leave are positively related to cohabitation (versus remaining in a non-coresidential state), while schedule flexibility is negatively related to entering cohabitation.

The next three columns of Table 3 present similar results for women. As predicted by Hypothesis 1, women who work non-standard schedules are significantly less likely to marry (whether or not cohabitation is a competing risk) as are women who work split or rotating shifts (when cohabitation is not a competing risk). Similar to men, being paid hourly significantly reduces the risk of marriage for women.

We also find strong associations between fringe benefits and entry into marriage for women. Women whose jobs provide more fringe benefits, captured by higher values on the fringe benefits scale, have a significantly higher likelihood of transitioning into first marriage ($b = 0.7728$, $p < 0.001$), controlling for respondent characteristics and for earnings/employment. Similar to men, moving up by a standard deviation on the fringe benefits scale from the mean level of fringe benefits is associated with an approximately 19% increase in the probability of first marriage in a given person month (Figure 1, Panel B).

The evidence related to Hypothesis 2, on cohabitation, is perhaps more interesting for women than for men. First, we see in Table 3 that while schedule type and benefits are significantly related to marriage, there are no significant associations with cohabitation. Further, the differences between the significant association with marriage and the null association with cohabitation are themselves statistically significant for non-standard schedule and benefits. It is also notable that while a split shift is negatively, but not significantly, associated with marriage for women (versus staying in a non-coresidential state) and positively, but not significantly, associated with cohabitation (versus staying in a non-coresidential state), having a split shift is a significant predictor of transitioning to cohabitation versus transitioning to marriage. The same is true of union coverage.

In Table 4, as for men, we also find some support for our hypothesis that health and dental insurance and life insurance and retirement benefits would be more strongly associated with marriage than with cohabitation. We see that health and dental insurance benefits and retirement benefits are all positively and significantly related to marriage and that this positive association is significantly stronger than for cohabitation.

The full set of coefficients on earnings/employment from all of the models presented in Tables 3 and 4 are shown in Tables 6 and 7 in the appendix, which show that earnings/employment is a consistent significant predictor of union formation net of job quality measures across model specifications. As a robustness check, we estimated all of the models in Tables 3 and 4 only on those who were employed in the prior year, and find consistent results.

Gender Differences in Associations between Job Quality and Union Formation

As reported in Tables 3 and 4, we find strong evidence of associations between job quality and marriage entry for both men and women and little evidence of associations between job quality and cohabitation for either men or women. We have already reported that most of these differences in the power of job quality for predicting marriage versus cohabitation are statistically significant.

We also tested Hypothesis 3 by assessing whether the positive associations between job quality and marriage were significantly different between men and women. We find no evidence of any interactions between gender and job quality. There are similar null effects for multiple job holding for both men and women. While we hypothesized that non-standard or irregular shifts might have more negative consequences for women's marriage than for men's, we find no statistically significant differences in the estimated coefficients by gender as well as no significant differences by gender in the coefficients on both fringe benefits and being paid a salary. While union membership is not a significant predictor of marriage in the models that are estimated separately by gender (Table 3), we see that when pooling the data, union coverage is positively and significantly associated with marriage – for both men and women (results not shown).

Educational Attainment and Marriage Entry

We next turn to the relationship between educational attainment and first marriage and to the role of job quality in explaining any educational advantage in marriage entry. Table 5 presents the result of a decomposition analysis of the effect of education on marriage entry based on estimates from the competing risks models (Panels A and C) and the event history models that examine all transitions to marriage (Panels B and D). Here, we use the Karlson-Holm-Breen (2012) method to decompose the total effects of education (after conditioning on our controls and our measure of employment/earnings) into an indirect effect of job quality and the remaining direct effect. Panels A and B report the decomposition for men and C and D for women. While we use the language of "effects" in this section, we caution that these remain associational analyses.

Comparing across the top rows of Panels A and B, the total effect, we see that there is a strong and significant educational gradient in first marriage for men. The third row of Panels A and B shows how much of the educational effect is accounted for by our five measures of job quality. We see that these measures play a minor role in accounting for the differences in marriage between those with less than a high school degree and those who completed high school – just 1.5% in the competing risks models and 2.7% in the marriage only models, both of which are not significant. Job quality plays a somewhat larger role in accounting for differences between men with some college and less than high school – 5.7% and 6.5% – a statistically significant share. However, job quality plays a much more important role in accounting for differences between men with an Associate's degree (10.4% and 9.8%) and those with less than High School. This attenuation is also statistically significant at the $p < 0.001$ level. Job quality plays the largest role, however, in explaining the significant marriage advantage of men with at least a BA. Here, accounting for job quality explains about a fifth of the advantage ($p < 0.001$). In supplemental models, we find that fringe benefits and job schedules are the primary drivers of this mediation.

Panels C and D of Table 5 present parallel results for women. As for men, we see strong educational gradients in entry into marriage as reflected by the increasing and significant coefficients across the top rows of both panels – the total effects. While job quality does not significantly mediate the marriage advantage of women with a high school degree versus less than high school, there is statistically significant mediation of the marriage advantage of women with some college, an Associate's degree, or with a Bachelor's degree or more. As for men, this attenuation is most pronounced for the most highly educated women – about 25% of the advantage in the competing risks model (Panel C) and about 15% in the models that allow for entry into marriage from cohabitation (Panel D).

In all, Hypothesis 4 is strongly supported. Job quality substantially and significantly mediates the association between education and marriage, explaining on the order of 20% to 25% of the marriage gap between college-educated and less than high school educated men and women.

Discussion

Economic resources have long been appreciated as an important determinant of marriage, but prior research has mainly focused on narrow measures of economic circumstances – employment and earnings – and far less research has considered the influence of job quality. Yet, we know from research on the changing labor market that aspects of job quality – such as fringe benefits, work schedules, and whether workers are unionized or are paid hourly or salaried – are also important job features that affect financial security and work/life balance and have the potential to influence union transitions. However, very little prior research has examined this relationship. This paper takes advantage of rich data from the National Longitudinal Survey of Youth 1997, which follows the

1980-1984 birth cohort from their teenage years through their thirties, to address the question of how job quality matters for men and women's transitions to marriage and cohabitation.

We find support for our hypothesis that better job quality encourages marriage. After taking into account employment status and earnings, we find that men and women employed in jobs offering more fringe benefits are significantly more likely to marry. Scholars and advocates have recently turned their attention to another dimension of precarious employment – variable work schedules (i.e. Lambert, 2008; Henly et al, 2006). We show that split shifts, rotating schedules, and non-standard schedules are negatively associated with marriage, providing rare quantitative support for the intuition that such practices have negative consequences for families. In all, our findings are consistent with the conclusion that poor job quality is an impediment to marriage, and that job quality influences marriage entry above and beyond one's level of earnings.

Prior demographic research has debated whether women's economic resources would be positively or negatively related to marriage (Becker, 1981; Oppenheimer, 1988). Empirical work shows fairly conclusively that the association is positive for women in recent cohorts (Sweeney, 2002), but the evidence has continued to show that men's resources are more strongly related to marriage than women's (Shafer and James, 2013). We do not find evidence of that inequality in association with respect to job quality. Instead, it appears that women's job quality is just as strongly associated with marriage entry as men's.

Recent decades have witnessed a striking divergence in marriage behavior across education groups with declines in marriage concentrated among those without a college degree. The declines in labor market opportunities for those with less than a college education is a leading explanation for this divergence in marriage behavior across education groups. Here again, we find that job quality matters. The educational differences in entry to marriage are partially explained by differences across education groups in job quality as measured by fringe benefits, stable and regular schedules, and getting paid a salary. Those with higher levels of educational attainment are more likely to enjoy better quality jobs, e.g. jobs with better fringe benefits and more stable schedules and earnings, and these job features help to explain why those with higher levels of educational attainment are more likely to marry. Specifically, we find that differences in job quality accounted for between 20% and 25% of the marriage advantage enjoyed by men and women with at least a Bachelor's degree over those with less than a high school diploma.

These differences in job quality matter much less for cohabitation. In competing risks models, for both men and women, our measures of job quality – work schedule type, fringe benefits, and unionization were simply not predictive of cohabitation (one exception was that women who are paid hourly are less likely to transition to either marriage or cohabitation). We do not find that low job quality is a barrier to cohabitation as it is for marriage, but nor do we find much evidence that low job quality encourages cohabitation. In all, while precarious employment may slow marriage, it does not appear to significantly shape the formation of cohabiting unions.

The NLSY97 provides an opportunity to examine rich and nuanced job characteristics alongside standard measures of education and earnings and to estimate relationships between economic predictors that temporally precede marital and cohabitation outcomes. Nevertheless, men and women may alter their work effort and employment choices in anticipation of marital and cohabitation transitions. Therefore, a limit of our analysis is that we cannot take into account these anticipatory effects on economic circumstances, which may contribute to the relationship between economic predictors and marriage and cohabitation outcomes. If men and women increase their work effort and select into higher quality jobs in anticipation of a planned marriage, for instance, then our models will overestimate the influence of job quality on marriage. Additionally, while we are able to extend the age range for analysis beyond that of previous research using the NLSY97, we still do not yet observe unions that form after age 34. Finally, while we have drawn on a rich set of measures of job quality, we would ideally be able to also capture contingent employment relationships, control over work tasks, or perceived job insecurity (Kalleberg, 2009). We would also ideally be able to capture the "regular unpredictability" that appears to characterize many hourly jobs in terms of week-to-week variation in total hours worked and in work schedules (Schneider and Harknett, 2017). The NLSY97 has added new measures of these emergent precarious labor practices in recent waves, but these measures are not available prior to 2011. Future research that could identify data sources that capture that kind of within-job instability could very usefully examine if that form of precarity is associated with union formation for men and women.

In sum, we demonstrate that job quality matters for marriage for both men and women. We also find clear evidence that job quality plays a significant role in explaining educational divides in marriage entry and so lends credence to the predictions that the bifurcation of the labor market into "good jobs" and "bad jobs" helps to explain class differences in marriage behavior. Thus, understanding the mechanisms that sort individuals into bad jobs or provide ladders to better job opportunities is essential for understanding inequality in both economic and union formation domains.

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Table 1: Characteristics of Respondents in Sample

	Men	Women
Marriage and Cohabitation (%)		
Never married or cohabited	29.55	22.12
Married before cohabited	13.77	16.17
Married after cohabited	26.98	32.66
Cohabited but never married	29.70	29.05
Birth of Child before Marriage or Cohabitation (%)	48.63	42.60
Eventual Education (%)		
Less than HS	10.48	6.21
HS	31.40	21.04
Some College, No Degree	29.26	30.36
AA	6.94	9.08
BA+	21.91	33.31
Race-Ethnicity (%)		
White (Non-Hispanic)	50.77	49.13
Black (Non-Hispanic)	24.94	27.74
Other Race (Non-Hispanic)	3.58	3.75
Hispanic	20.71	19.38
Mother's Education (%)		
Less than HS	20.90	20.29
HS	34.31	34.40
Some College or AA	21.58	23.29
BA+	17.95	17.24
(missing data)	5.26	4.77
Number of Respondents	4,162	3,735

N=7,897 respondents eligible for inclusion in our competing risk models, who are present in the sample at least one month when they (1) are 18 or over, (2) haven't married or cohabited, and (3) have non-missing data

Table 2. Time-varying Characteristics for Respondents Used in Competing Risks Analysis

	Men				Women			
	Mean	SD	Min	Max	Mean	SD	Min	Max
First Marriage	0.001	0.039	0.000	1.000	0.002	0.047	0.000	1.000
First Cohabitation	0.007	0.081	0.000	1.000	0.009	0.093	0.000	1.000
Current Education								
Less than HS	0.32	0.45	0.00	1.00	0.28	0.43	0.00	1.00
HS	0.26	0.41	0.00	1.00	0.19	0.35	0.00	1.00
Some College, No Degree	0.31	0.45	0.00	1.00	0.38	0.46	0.00	1.00
AA	0.03	0.16	0.00	1.00	0.03	0.16	0.00	1.00
BA+	0.08	0.26	0.00	1.00	0.12	0.31	0.00	1.00
Logged Average Weekly Earnings	4.58	2.36	0.00	14.14	4.51	2.20	0.00	10.73
Job Characteristics:								
More Than One Job	0.08	0.22	0.00	1.00	0.10	0.24	0.00	1.00
Schedule - Evening, Night, Weekend	0.13	0.29	0.00	1.00	0.14	0.28	0.00	1.00
Schedule - Split, Rotating	0.09	0.24	0.00	1.00	0.10	0.25	0.00	1.00
Schedule - "Irregular"	0.07	0.21	0.00	1.00	0.08	0.21	0.00	1.00
Paid Hourly	0.46	0.41	0.00	1.00	0.50	0.41	0.00	1.00
Union	0.06	0.22	0.00	1.00	0.06	0.20	0.00	1.00
Benefits Scale	0.14	0.22	0.00	1.00	0.15	0.23	0.00	1.00
Variables used in benefits scale:								
health	0.28	0.41	0.00	1.00	0.26	0.40	0.00	1.00
life	0.17	0.34	0.00	1.00	0.17	0.34	0.00	1.00
dental	0.24	0.38	0.00	1.00	0.23	0.38	0.00	1.00
retirement	0.18	0.35	0.00	1.00	0.17	0.34	0.00	1.00
tuition reimbursement	0.10	0.27	0.00	1.00	0.11	0.28	0.00	1.00
stock	0.08	0.24	0.00	1.00	0.07	0.22	0.00	1.00
Family Friendly Benefits Scale	0.09	0.17	0.00	1.00	0.12	0.18	0.00	1.00
maternity leave - paid	0.09	0.26	0.00	1.00	0.13	0.30	0.00	1.00
maternity leave - unpaid	0.06	0.21	0.00	1.00	0.09	0.25	0.00	1.00
flexible schedule	0.19	0.34	0.00	1.00	0.22	0.35	0.00	1.00
child care	0.02	0.13	0.00	1.00	0.03	0.14	0.00	1.00
Controls:								
Student	0.40	0.43	0.00	1.00	0.51	0.43	0.00	1.00
Birth of Child	0.01	0.02	0.00	0.17	0.01	0.02	0.00	0.17
Number of Children in HH	0.42	0.27	0.00	2.09	0.42	0.27	0.00	2.17
Dot-com Recession	0.08	0.20	0.00	0.75	0.08	0.20	0.00	0.75
Great Recession	0.08	0.23	0.00	1.00	0.07	0.22	0.00	1.00
Age	22.77	3.58	18.00	34.00	22.57	3.59	18.00	34.00
Number of person-months	336,535				257,732			

Notes: N=594,267 person-months, in which eligible respondents (N=7,897) had not yet entered first marriage or cohabitation. For all characteristics besides the outcomes (marriage and cohabitation), variables are 12-month averages of lagged measures from 17 months prior to 6 months prior to the measurement month of the outcomes (because age increases linearly with time, we do not lag age as this would produce essentially identical results)

Table 3. Job Quality and Entry into First Marriage or Cohabitation for Men and Women

	Men			Women		
	Competing Risks		Marriage	Competing Risks		Marriage
	Marriage	Cohabitation		Marriage	Cohabitation	
Baseline - No Job Quality Predictors:						
(0) Earnings/Employment	0.0665*	0.1083***	0.1051***	0.0594*	0.0852***	0.0660***
	(0.0267)	(0.0115)	(0.0178)	(0.0247)	(0.0129)	(0.0147)
Regressions with Job Quality Predictors:						
(1) More Than One Job	0.1426	0.1116	0.003	-0.337	0.0017	-0.1368
	(0.1796)	(0.0913)	(0.1039)	(0.1821)	(0.0866)	(0.0972)
(2) Schedule						
Evening, Night, Weekend	-0.5308**	-0.0155	-0.2854**	-0.4771**	0.1328	-0.2612**
	(0.1805)	(0.0754)	(0.0977)	(0.1751)	(0.0767)	(0.0935)
Split, Rotating	-0.4772*	-0.1398	-0.1477	-0.2968	0.0954	-0.3018**
	(0.2012)	(0.0938)	(0.1042)	(0.1809)	(0.0823)	(0.0995)
"Irregular"	-0.0258	-0.082	-0.0814	-0.0322	-0.0772	-0.0026
	(0.2072)	(0.1093)	(0.1288)	(0.1889)	(0.103)	(0.1149)
(3) Paid Hourly	-0.2317	-0.0651	-0.1809**	-0.2761*	-0.1436*	-0.1410*
	(0.1256)	(0.0589)	(0.0675)	(0.1244)	(0.0634)	(0.0671)
(4) Union Contract	-0.0047	0.1287	0.1301	0.3015	-0.1299	0.0605
	(0.1948)	(0.0897)	(0.0943)	(0.1688)	(0.1107)	(0.1039)
(5) Fringe Benefits Scale	0.6561**	0.1563	0.6662***	0.7728***	0.2153	0.5742***
	(0.2003)	(0.1093)	(0.1112)	(0.21)	(0.115)	(0.1096)
N	336,535		469,718	257,732		402,448

Notes: Because measures of job quality are highly correlated with one another, the five sets of job quality measures (more than one job, schedule, etc.) are each included in separate models. The first row shows coefficients on earnings/employment from models that do not include job quality measures.
The competing risks results show coefficients for marriage vs. no union formation with cohabitation as a competing risk and for cohabitation entry vs. no union formation with marriage as a competing risk. Asterisks denote if the estimate is statistically significant compared to no union formation (*** p < 0.001; ** p < 0.01; * p < 0.05). Underlined values denote statistically significant differences between cohabitation and marriage at p < .05.
The third column (for men) and the sixth column (for women) of coefficients are from the logistic regression model that allows for entry into marriage irrespective of cohabitation status prior to marriage.
All regressions control for age, race/ethnicity, education, school enrollment, mother's education, recession context, fertility, and earnings/employment. The logistic regressions reported in the third and sixth columns also control for whether the respondent is currently cohabitating.

Table 4. Fringe Benefits and Entry into First Marriage or Cohabitation for Men and Women

	Men			Women		
	Competing Risks		Marriage	Competing Risks		Marriage
	Marriage	Cohabitation		Marriage	Cohabitation	
Health Insurance	0.3988**	0.1039	0.3951***	0.5906***	0.1606*	0.3818***
	(0.1268)	(0.0604)	(0.0721)	(0.1276)	(0.0666)	(0.0692)
Life Insurance	0.5124***	0.0823	0.3548***	0.3360*	0.1586*	0.2952***
	(0.1266)	(0.0639)	(0.069)	(0.1356)	(0.0732)	(0.0681)
Dental Insurance	0.4214***	0.0681	0.3685***	0.4172**	0.1077	0.3178***
	(0.1235)	(0.0616)	(0.0685)	(0.1315)	(0.0687)	(0.0677)
Retirement	0.4168***	0.1444*	0.3883***	0.4384***	0.0765	0.3480***
	(0.1261)	(0.0639)	(0.068)	(0.1318)	(0.0723)	(0.0681)
Tuition Reimbursement	0.3690**	0.0514	0.3462***	0.3757**	0.0812	0.2021**
	(0.1432)	(0.0775)	(0.0758)	(0.1379)	(0.0789)	(0.0732)
Stock Ownership	-0.0649	-0.0016	0.1364	0.3043	-0.0164	0.1649
	(0.1682)	(0.0809)	(0.0864)	(0.1667)	(0.0965)	(0.088)
Family Friendly Benefits Scale	0.322	0.0879	0.4505***	0.5075*	0.1908	0.4547***
	(0.2355)	(0.1288)	(0.1224)	(0.2388)	(0.1303)	(0.1194)
Maternity Leave - Paid	0.3573*	0.1484	0.3672***	0.3033*	0.0946	0.2546***
	(0.1474)	(0.0775)	(0.0752)	(0.1422)	(0.0736)	(0.0699)
Maternity Leave - Unpaid	0.1078	0.3065***	0.1773*	0.3217*	0.3198***	0.3594***
	(0.1821)	(0.0884)	(0.0861)	(0.1602)	(0.0839)	(0.0756)
Flexible Schedule	0.0927	-0.1758**	0.0638	-0.0141	-0.0619	-0.0368
	(0.1281)	(0.0659)	(0.069)	(0.1208)	(0.0631)	(0.0653)
Child Care	-0.4125	0.037	0.2152	0.4832*	0.0154	0.2616*
	(0.3608)	(0.1497)	(0.1383)	(0.2384)	(0.1534)	(0.1233)
N	336,535		469,718	257,732		402,448

Notes: Because measures of particular fringe benefits are highly correlated with one another, the individual fringe benefit measures (health insurance, life insurance, etc.) are each included in separate models.
The competing risks results show coefficients for marriage vs. no union formation with cohabitation as a competing risk and for cohabitation entry vs. no union formation with marriage as a competing risk. Asterisks denote if the estimate is statistically significant compared to no union formation (*** $p < 0.001$; ** $p < 0.01$; * $p < 0.05$). Underlined values denote statistically significant differences between cohabitation and marriage at $p < .05$.
The third column (for men) and the sixth column (for women) of coefficients are from the logistic regression model that allows for entry into marriage irrespective of cohabitation status prior to marriage.
All regressions control for age, race/ethnicity, education, school enrollment, mother's education, recession context, fertility, and earnings/employment. The logistic regressions reported in the third and sixth columns also control for whether the respondent is currently cohabitating.

Table 5. Decomposition of Effects of Education on Marriage into Direct and Indirect Effects via Job Quality, with Controls by Gender

Panel A. Decomposition for Men - Marriage with Cohabitation as Competeing Risk (N=336,535)

	High School			Some College			Associates			Bachelor's		
	Coeff.	Z	%	Coeff.	Z	%	Coeff.	Z	%	Coeff.	Z	%
Total Effect	0.3610	2.033	100%	0.5750	3.249	100%	0.8741	3.220	100%	1.0330	4.628	100%
Direct Effect	0.3558	1.993	98.5%	0.5422	3.052	94.3%	0.7829	2.882	89.6%	0.8046	3.521	77.9%
Indirect Effect	0.0053	0.476	1.5%	0.0329	2.211	5.7%	0.0912	3.438	10.4%	0.2284	4.207	22.1%

Panel B. Decomposition for Men - Marriage, Irrespective of Cohabitation (N=469,718)

	High School			Some College			Associates			Bachelor's		
	Coeff.	Z	%	Coeff.	Z	%	Coeff.	Z	%	Coeff.	Z	%
Total Effect	0.4384	4.434	100%	0.6665	6.683	100%	1.0268	6.826	100%	1.1748	9.607	100%
Direct Effect	0.4264	4.302	97.3%	0.6233	6.218	93.5%	0.9258	6.105	90.2%	0.9638	7.652	82.0%
Indirect Effect	0.0120	2.012	2.7%	0.0432	4.332	6.5%	0.1010	6.008	9.8%	0.2111	7.010	18.0%

Panel C. Decomposition for Women - Marriage with Cohabitation as Competeing Risk (N=257,732)

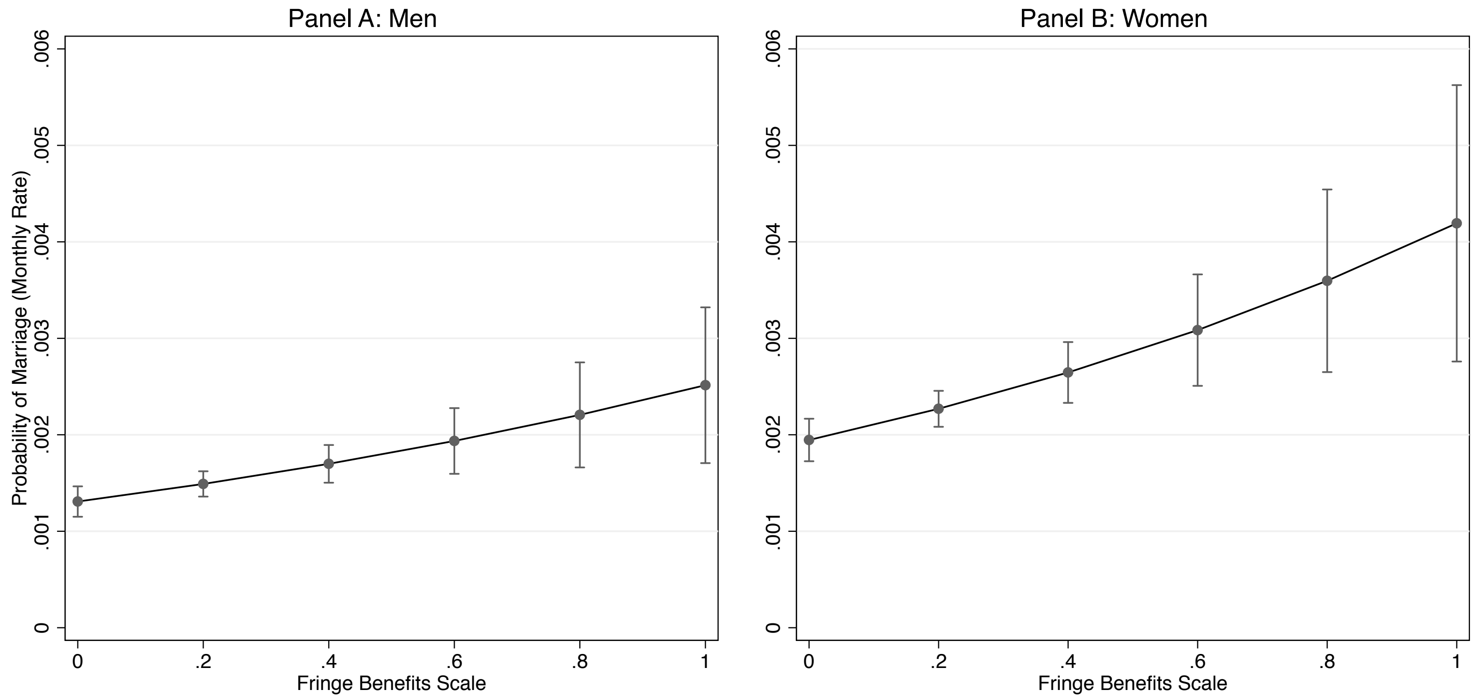
	High School			Some College			Associates			Bachelor's		
	Coeff.	Z	%	Coeff.	Z	%	Coeff.	Z	%	Coeff.	Z	%
Total Effect	0.0954	0.498	100%	0.4517	2.680	100%	0.4719	1.634	100%	0.9100	4.117	100%
Direct Effect	0.0983	0.513	103.0%	0.4254	2.518	94.2%	0.4031	1.395	85.4%	0.6664	2.971	73.2%
Indirect Effect	-0.0028	0.317	-3.0%	0.0263	2.004	5.8%	0.0688	2.785	14.6%	0.2436	4.202	26.8%

Panel D. Decomposition for Women - Marriage, Irrespective of Cohabitation (N=402,448)

	High School			Some College			Associates			Bachelor's		
	Coeff.	Z	%	Coeff.	Z	%	Coeff.	Z	%	Coeff.	Z	%
Total Effect	0.2748	2.640	100%	0.5446	5.470	100%	0.7504	4.916	100%	1.1211	9.252	100%
Direct Effect	0.2718	2.609	98.9%	0.5194	5.204	95.4%	0.6929	4.532	92.3%	0.9559	7.700	85.3%
Indirect Effect	0.0030	0.673	1.1%	0.0252	3.536	4.6%	0.0575	4.727	7.7%	0.1652	5.471	14.7%

Notes: All regressions control for age, race/ethnicity, education, school enrollment, mother's education, recession context, fertility, and earnings/employment. Regressions in Panels B and D also control for whether the respondent is currently cohabiting.

Figure 1. Predicted Probability of Marriage by Fringe Benefits Scale for Men and Women



Notes: Estimates from model with cohabitation as competing risk.