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Marriage and Work Among Prime-Age Men

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Abstract

Married men work more hours than men who have never been married. Fixed effect regressions reveal that part of this gap is attributable to an increase in work around the time of marriage. Two potential explanations for the increase are: (i) men hit by positive labor market shocks are more likely to marry; and (ii) marriage leads men to work more hours. Using a structural life-cycle model, we find that marriage substantially increases male hours of work. Counterfactual simulations suggest that declining marriage rates account for roughly half of the fall in prime age male hours observed over recent decades.

Keywords: Labor supply, family structure, marriage, marital wage premium.

JEL Classifications: D15, J1, J22, J31

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But if anyone does not provide for his relatives, and especially for members of his household, he has denied the faith and is worse than an unbeliever.

— 1 Timothy 5:8

How can you frighten a man whose hunger is not only in his own cramped stomach but in the wretched bellies of his children? You can't scare him – he has known a fear beyond every other.

— John Steinbeck, *Grapes of Wrath*

What does a man do, Walter? A man provides for his family.

— *Breaking Bad*

1 Introduction

This paper is motivated by two observations. First, married men work considerably more than men who have never been married. Among men ages 19 – 54 in the Current Population Survey (CPS), married men work 30% more annual hours than never-married men, a gap which has remained roughly constant since the 1970's.[1] Second, over the past half century, marriage rates and male hours of work have both fallen markedly. Between 1970 and 2018, marriage rates among men ages 19 - 54 fell by 30 percentage points, while hours of work fell by 7.7%. These patterns raise a natural question: did the decline in marriage contribute to the decline in male labor supply? The answer to this question depends on why married men work more than men who have never been married.

In this paper, we combine reduced form empirical analyses and a structural model to quantitatively assess the importance of two potential channels linking marriage and male labor supply. One potential channel is selection: men with better labor market outcomes may be more willing to marry and may also be more attractive marriage prospects. A second possibility is that marriage, or the prospect of marriage, increases the labor supply of men. Our results suggest that selection is quantitatively important, but does not explain the full gap in hours worked by marital status. In particular, we find that marriage increases male labor supply and, as suggested by Binder and

[1] Since 1975, the magnitude of the marital gap in annual hours worked for men has been larger, albeit with the opposite sign, than for women. An immense literature has sought to understand the effect of marriage and motherhood on the labor supply of women: important contributions include Becker (1985), Becker (1988), Becker (1991), Goldin (1992), Weil and Galor (1996), Goldin (2014), Doepke and Tertilt (2016).

Bound (2019) and Binder (2021), the observed decline in marriage explains a substantial portion of the observed decline in prime age male hours.

We begin by documenting a robust reduced form relationship between male marital status and hours worked. The gap in hours worked between married and never-married men remains large after controlling for observables. We also find that state-level variation in marriage rates strongly predict male hours worked, and that changes in state-level marriage rates over time predict changes in male hours. A simple accounting exercise, implicitly assuming marriage increases male labor supply, reveals that the decline in marriage rates between 1970-2018 predicts 72% of the decline in male hours observed over that time period.

Next, we document that an important share of the additional hours worked by married men can be attributed to an increase in work around the time of marriage. In particular, we regress hours worked on a set of dummy variables for distance-from-marriage, as well as individual fixed effects, on panel data from the National Longitudinal Survey of Youth 1979 (NLSY79). We find that men increase their hours by roughly 10% in the six years preceding marriage, and this increase persists for at least a decade after.

We then quantitatively assess potential explanations for *why* married men work more than men who have never been married and *why* male hours of work increase around the time of marriage. We develop a life-cycle model of male labor supply and saving, where men face uncertainty over wages, marital status, and fertility. The model distinguishes selection from marriage-related incentives through three forces. First, to align with the large fixed effects found in our reduced form analyses, the model includes permanent worker types, who differ in wages, the probability of getting married, and preferences for work. This channel captures selection into marriage based on time-invariant traits, which explain much but not all of the marital hours gap. Second, men face persistent wage shocks that affect both labor supply and the probability of marriage. This channel captures dynamic selection into marriage: men who receive a positive labor market shock, which is expected to persist for many years, are more likely to marry. Third, marriage and children raise labor supply through a "mouths-to-feed" motive. In the model, marriage and children can impact male labor supply via several channels. The most important of these is that men internalize the utility of their family members, which raises the marginal utility of consumption for men with families and induces them to work more (Blundell, Pistaferri and Saporta-Eksten, 2016; Fan, Seshadri and Taber, 2024). To capture the empirical finding that the increase in hours occurs in the years just before marriage, the model includes a gradual marriage process with an engagement stage, which signals to men that marriage is imminent and induces them to work more in anticipation of family-related consumption needs.

We estimate the model using data from the NLSY79. The estimated model closely matches the life-cycle evolution of marital status and children, as well as the correlation of earnings between spouses. Including permanent selection through the two worker types allows the model to match differences in the wages and average hours of never-married and ever-married men. To discipline dynamic selection via the wage shocks, we require the model to replicate the marginal effect of wages on marriage probabilities estimated in the CPS, where we instrument for individual wage changes with state-level variation.

The estimated model generates hours dynamics around marriage resembling those in our fixed-effect regressions, mainly through the mouths-to-feed channel. Although our CPS estimates imply positive dynamic selection into marriage from wages, this force is quantitatively weak in the model for two reasons. First, empirically, exogenous wage changes predict only modest increases in marriage. Second, in the model, the wage shocks that induce dynamic selection into marriage are persistent, implying a large income effect that offsets the substitution effect on labor supply. The absence of a strong dynamic selection effect leaves the mouths-to-feed effect as the main driver of the hours run-up. Additional evidence in favor of the mouths-to-feed effect comes from marriages preceded by pregnancies, which NLSY79 responses suggest are more likely to be unplanned and less likely to reflect favorable labor market shocks. Yet male hours rise more before these marriages than before marriages where children arrive later.[2]

Our model indicates that marriage has a substantial positive effect on male work hours. A counterfactual exercise shows that eliminating marriage altogether reduces average hours worked by prime-age men by 6.8%. Next, we quantify Binder and Bound's (2019) hypothesis that declining marriage rates contributed to the decline in male hours. To do so, we replace the NLSY79 marriage process with one estimated from the NLSY97, whose cohort married at much lower rates. Holding all other parameters fixed, this change accounts for 48% of the decline in male annual hours worked in CPS for these cohorts. If we add cross-cohort changes in male wages and spousal earnings, the model changes account for 94% of the corresponding CPS decline.

Our project lies at the intersection of three literatures that have remained largely isolated. The first studies the "male marriage premium," focusing mainly on hourly wages.[3] These papers typically find that wages are about 10% higher for married men after controlling for observables. The leading causal explanation is that marriage increases husbands' productivity by encouraging

[2]The hours response to pre-marital children is also evidence against marital selection along other dimensions, such as employment shocks (Kaplan, 2012).

[3]See, e.g., Korenman and Neumark (1991), Cornwell and Rupert (1997), Ginther and Zavodny (2001), Antonovics and Town (2004), Rodgers III and Stratton (2010), Budig and Lim (2016), Glauber (2018), Killewald and Lundberg (2017), and the meta-analysis by de Linde Leonard and Stanley (2015).

specialization in market work (Becker, 1991).[4] The leading non-causal explanation is selection: men with higher wages are more likely to marry. Our view is that this literature has not reached a firm conclusion about the relative importance of these explanations.[5] Our model includes both mechanisms. We allow wages to increase with hours of work, capturing specialization, and we allow marriage probabilities to depend on the man's permanent type and the realization of his persistent wage shock, capturing selection. In the model, specialization is the primary driver of wage increases.

We are aware of two papers in this reduced form literature that emphasize hours rather than wages. Akerlof (1998) studies men in the NLSY79 and shows that after marriage they receive higher wages, work more, and are less likely to abuse drugs and alcohol. Lundberg and Rose (2002) use the Panel Study of Income Dynamics (PSID) and show that after marriage and childbirth men receive higher wages and work more. However, neither study analyzes the time path of these variables or quantifies the channels behind the hours increase, which are central objectives of our paper.

The second literature studies interactions between gender, marriage, children, and the labor market in structural models. Becker (1985) and Becker (1991) are foundational theoretical contributions, while Greenwood, Guner and Knowles (2003) and Attanasio, Low and Sánchez-Marcos (2005) are early dynamic quantitative exercises. Some papers study only couples' labor supply and therefore cannot speak to differences between married and single individuals.[6] Others model male labor supply or earnings as exogenous.[7] Papers such as Guner, Kaygusuz and Ventura (2012*a*,*b*) and Borella, De Nardi and Yang (2023) include both marital dynamics and endogenous male labor supply. To our knowledge, however, none of these papers explains male labor market behavior around the time of marriage.

Within this literature, our work is most closely related to Siassi (2019) and Mazzocco, Ruiz and Yamaguchi (2014). Siassi (2019) studies income and wealth differences by marital status. In contrast, we emphasize that married men work more hours even as married women work less, and we study individual transitions around marriage rather than only cross-sectional differences. Mazzocco, Ruiz and Yamaguchi (2014) model labor supply, home production, savings, and marriage. Like us, they use panel data to show that hours rise around marriage, though they do not

[4]In a recent structural analysis, Pilossoph and Wee (2021) argue that the wage premium for married workers is due in part to different job search dynamics.

[5]For example, Borra, Browning and Sevilla (2021) find that upon marriage women increase their housework time much more than men. However, married men also increase their housework time, albeit by less and for different tasks.

[6]These include Knowles (2013), Blundell, Pistaferri and Saporta-Eksten (2016, 2018), Alon, Coskun and Doepke (2018), Chiappori, Dias and Meghir (2018), Bick and Fuchs-Schündeln (2018), and Ellieroth (2019).

[7]See, e.g., Greenwood et al. (2016), Caucutt, Guner and Rauh (2021), and Low et al. (2025).

document the similar patterns in wages and earnings.

Relative to both papers, our analysis differs in three ways. First, they model marriage as an instantaneous shock, while we allow marriage to be preceded by engagement. Second, we discipline the strength of dynamic selection using plausibly exogenous variation in state economic conditions, which is central to understanding why hours rise before marriage. Third, we use counterfactual exercises to quantify the importance of marriage for aggregate labor supply among prime-age men.

Finally, we contribute to the emerging literature on the joint secular declines in marriage and employment among prime-age men (Binder and Bound, 2019; Binder, 2021). In related event studies, Autor, Dorn and Hanson (2019) find that negative labor demand shocks reduce both marriage and fertility, while Kearney and Wilson (2018) find that fracking booms increase fertility but not marriage. Our approach differs by using a quantitative model to study how marriage and labor supply interact within a cohort, taking labor and marriage markets as given. We view our finding that marriage substantially increases male labor supply as a complementary input into understanding these broader social transformations.

The rest of the paper proceeds as follows. Section 2 documents that married men work more than single men and that their hours of work increase significantly before their first marriage. Section 3 develops a structural model that can explain these facts, and section 4 describes how the parameters of the model are set. Section 5 documents the properties of the baseline parametrized model. Section 6 uses the model to quantitatively analyze potential drivers of the marriage-related hours increase. We conclude in section 7.

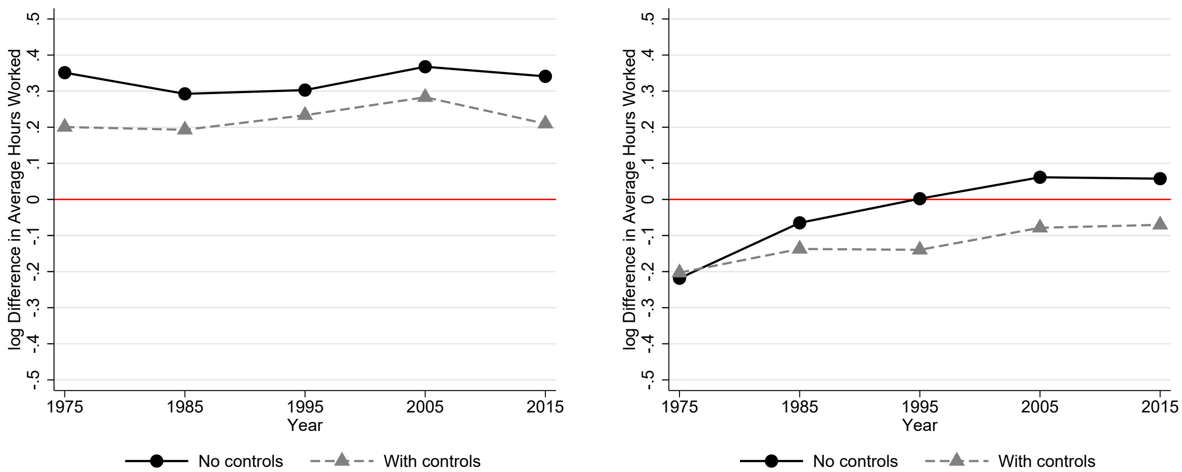
2 Evidence on Marriage and Male Labor Market Outcomes

We begin with empirical evidence on the relationship between marriage and male labor market outcomes. First, we use repeated cross-sectional data from the CPS to document a large and stable gap in hours worked between married and never-married men over the last four decades, our first motivating fact. We then show that the downward trend in male hours of work observed in the CPS has coincided with a downward trend in marriage rates, our second motivating fact. Finally, we use panel data from the NLSY79 to establish a direct relationship between marriage and labor market outcomes, tying the facts together.

Figure 1: Hours Worked by Marital Status: 1975–2018

(a) Men: log(married / never married)

(b) Women: log(married / never married)



Source: Men and women ages 19 - 54 in the 1975–2018 waves of the CPS ASEC. Data points are logs of ten-year centered averages, except for 1975, which averages across the years 1975-80. Annual hours worked are the product of usual weekly hours and weeks worked. The sample includes those with zero hours. The solid line plots the log difference in average annual hours worked by currently married individuals versus individuals who have never been married. The dashed line plots the difference in the estimated coefficient for married versus never-married individuals in a regression of log annual hours worked on marital status and controls for education, age, race and state of residence.

2.1 Marriage and Work in Cross-Sectional Data

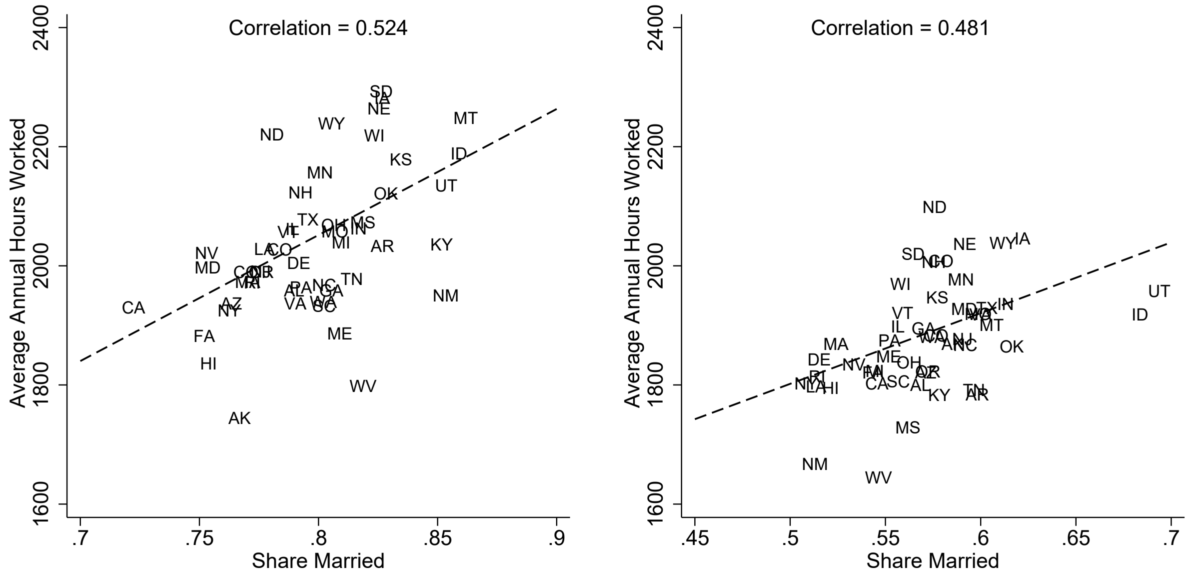
We use cross-sections for the years 1975 to 2018 taken from the Annual Social and Economic Supplement (ASEC) of the CPS (Flood et al., 2020). The ASEC includes information on both weekly hours and weeks worked in the previous calendar year, which allows us to construct a measure of annual hours worked. We restrict attention to the core working ages 19 - 54, and we include people with zero annual hours worked.

Figure 1 documents how annual hours of work differ by marital status. Our findings are consistent with a large number of earlier studies (see, e.g., Doepke and Tertilt 2016). Figure 2a shows results for men. The solid black line with circles shows the log ratio of average annual hours worked for currently married men relative to men who have never been married. Between 1975 and 2018, average annual hours worked by married men exceeded average hours of never-married men by 31 to 39 log points. Most of the gap remains after controlling for the men's education, age, race and state of residence (the gray dashed line with triangles) .

Figure 2b shows results for women. In 1975, married women worked nearly 30 log points less than never-married women. By the 2000's, however, the gap in the raw data had completely disappeared. After controlling for women's observables, the gap is always negative, but nonethe-

Figure 2: Cross-State Variation in Marriage and Hours Worked for Men

(a) 1975-1979 (b) 2015-2018



Source: Men ages 19 - 54 in the CPS ASEC. Annual hours worked are the product of usual weekly hours in the previous year and weeks worked in the previous year. The sample includes those with zero annual hours. The dashed line is the line of best OLS fit. Share married denotes the fraction who report being married at the time of the survey.

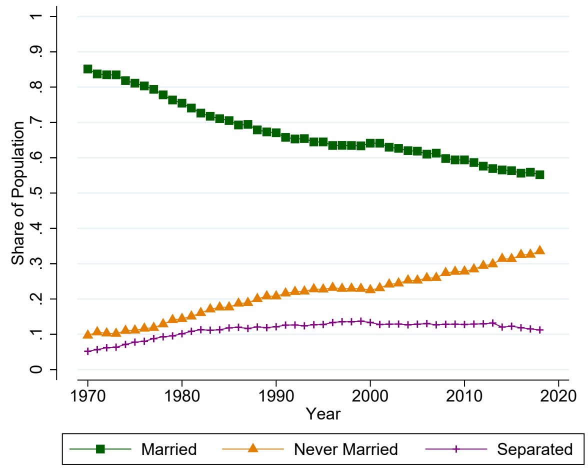
less, by 2018 it was only 12 log points. Since 1985, the magnitude of the marital hours gap among men has been larger than that of women, even after controlling for observables.

A clear cross-sectional relationship between marriage rates and hours worked is also apparent at the state level. Figure 2 plots the share of men who are currently married in each state against statewide average annual hours of work. Figure 2a plots data for the years 1975 to 1979. The correlation between hours worked and share married is 0.524, and the slope is significantly positive. Figure 2b shows the same scatter plot for the years 2015 to 2018, the most recent five-year window that excludes the large disruption from the pandemic. Even though both marriage rates and work have decreased in virtually every state, a similar positive relationship remains, with a correlation of 0.481.[8] Table A2 in Appendix B shows that this cross-state pattern continues to hold when we pool multiple years and control for age, education, and state.

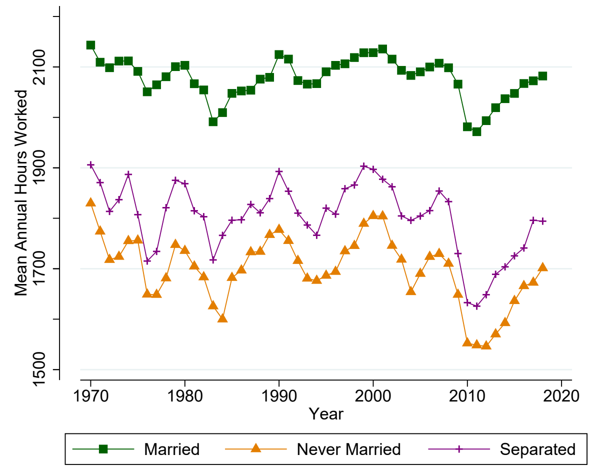
[8]In 2015, two noteworthy outliers are Utah and Idaho, with marriage rates of 69% and 68%, respectively. These two states have by far the largest fraction that is Mormon, a religion that emphasizes the importance of marriage.

Figure 3: Annual Hours of Work and Marital Status of Prime Age Men in the CPS

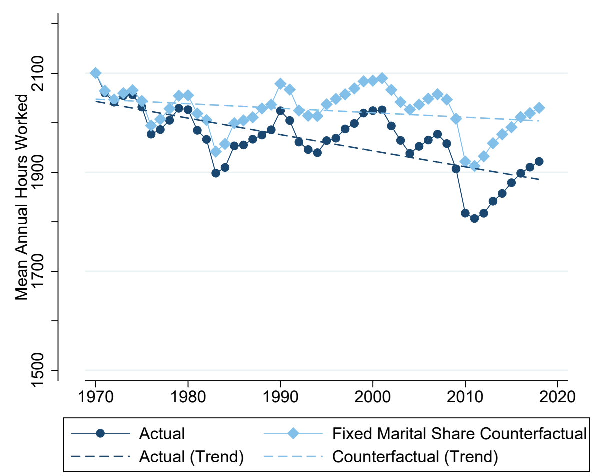
(a) Marital Status



(b) Annual Hours Worked by Marital Status



(c) Mean Hours vs. Shift-Share Counterfactual



Source: Men ages 19 - 54 in the 1970-2018 waves of the CPS ASEC. Fitted lines are log-linear trends. Counterfactual hours are constructed using mean hours by marital status, holding marital status shares fixed at their 1970 values. Figure 3b displays mean hours by marital status used to construct the counterfactual.

2.2 Marriage and Work over Time

Both marriage and male work hours have declined substantially over the past half century. Figure 3 documents these patterns in the CPS for the years 1970-2018. Figure 3a shows that the share of prime age men who are married has declined by about 30 percentage points. Figure 3b shows that over the same time period, married men consistently worked more than never-married men.

Figure 3c displays mean annual hours worked by all prime age men. Because the total change in hours is sensitive to starting and ending dates, we include a fitted trendline, which shows a decrease of 7.7%. A natural question is whether the decline in the marriage rate can account for a meaningful share of this decline in mean hours. To that end, Figure 3c also displays hours generated by a shift-share exercise in which we fix marital status shares at their 1970 values but allow hours conditional on marital status to vary over time as in the data. Between 1970 and 2018, the trend line for the shift-share counterfactual fell by 2.1%. The difference between this amount and the fall in the data trendline, 5.6 percentage points, implies that, in a reduced form sense, the decline in marriage rates can account for 5.6/7.7 = 72% of the decline in annual hours worked by prime age males.[9]

2.3 The Dynamics of Marriage and Work in Panel Data

Because the accounting exercise in the previous section implicitly assumes that hours differ by marital status for causal reasons – marriage causes men to work more – it provides an upper bound on the contribution of marriage to the decline in male labor supply. This motivates us to ask: what is the source of the hours difference? Does the typical man work more after he gets married? Or, alternatively, do men who eventually marry always work more, even before they are married? To answer these questions, we need to move beyond the cross-sectional comparisons in the preceding subsections and make use of panel data.

We turn to the NLSY79, a longitudinal study of 12,686 individuals born between 1957 and 1964 (Bureau of Labor Statistics, 2019*a*). Respondents were recruited and initially interviewed in 1979, when they were between 14 and 22 years old. They were then re-interviewed annually until 1994, then biennially afterward. The dataset contains a rich collection of information on family background, including detailed information on marriage and children, and labor market outcomes. Importantly, as of their initial interview 95% of male respondents in the NLSY79 had never been married, which allows us to observe changes in labor market outcomes around the date of marriage or the arrival of a child.[10]

We construct a "nearly-balanced" panel of men from the NLSY79 as follows. First, we drop the military oversample portion of the survey. This leaves us with 5,579 individuals who were

[9]Table A2 in Appendix B reveals that this relationship is also present at the state level: states which experienced larger declines in marriage rates also experienced larger declines in hours worked. Appendix Figure A1 shows that the decline in total hours of work took place along both the intensive and extensive margins of labor supply.

[10]Another candidate dataset with a long panel dimension is the PSID. Unfortunately, the PSID consistently collects detailed information only for household "heads" and "spouses." To the extent that younger individuals live with their parents, especially prior to marriage, this interviewing scheme limits our ability to study how labor market outcomes change in the years around marriage.

originally interviewed in 1979. Second, we restrict attention to men who we observe at age 50 or later, indicating that they remained in the survey for a substantial period of time. Third, among the remaining men, we restrict attention to those who were interviewed at least 20 times between 1979 and 2014 (out of a possible maximum of 26 interviews). These criteria balance our desire for a fairly complete life history against our need for a sufficiently large sample. This results in a final sample size of 2,731 men. High school graduates enter our sample once they turn 19. We also exclude observations from older ages if the man is currently enrolled in formal school; in particular, observations for men with a college degree do not enter into the sample until age 23 or older.

Table 1: Predictors of Male Annual Hours Worked in the NLSY79

	(1)	(2)	(3)
Constant	1751.0 ***	1808.4 ***	1866.0 ***
	(21.5)	(21.9)	(19.2)
Married	328.1 ***	283.0 ***	99.2 ***
	(10.5)	(10.5)	(13.1)
Separated / Widowed	71.7 ***	86.7 ***	7.7
	(15.1)	(15.1)	(17.7)
Less than High School	–	−228.6 ***	–
		(14.9)	
Some College	–	25.0 **	–
		(11.3)	
Bachelor's +	–	121.7 ***	–
		(11.1)	
Black	–	−296.5 ***	–
		(13.3)	
Hispanic	–	−150.3 ***	–
		(17.4)	
Age Cubic	*Y*	*Y*	*Y*
Year FEs	*Y*	*Y*	*Y*
Individual FEs			*Y*
R^2-adj	0.09	0.11	
N	42,930	42,930	42,930

Source: Males ages 19 - 54 in the NLSY79; see text for details. The sample includes those with zero annual hours.

To begin our panel analysis, we first regress annual hours worked on a dummy for current marital status, with and without controlling for individual fixed effects. The results are displayed in Table 1. Column (1) shows that the reference group of unmarried men work on average 1,751 hours per year. Married men of the same age and in the same calendar year work 328 hours more, a difference of 19%. Column (2) shows that adding controls for education reduces the increment to 283 hours. Column (3) utilizes the panel aspect of the data to include individual fixed effects in the controls. This further reduces the coefficient on marital status, but it remains statistically significant and economically meaningful at 99 hours. Based on these results, we conclude that a sizable share of the difference in hours worked by marital status is due to individual changes in hours that coincide with changes in marital status.

Next, we develop a fuller picture of how hours evolve around marriage by regressing annual hours on a sequence of dummies corresponding to the distance in years from the man's first marriage. Specifically, we run the following regression:

$$h_{i,t,d} = \beta_d^{distance} + \beta_t^{year} + \beta_i^{individual} + \varepsilon_{i,t}. \quad (1)$$

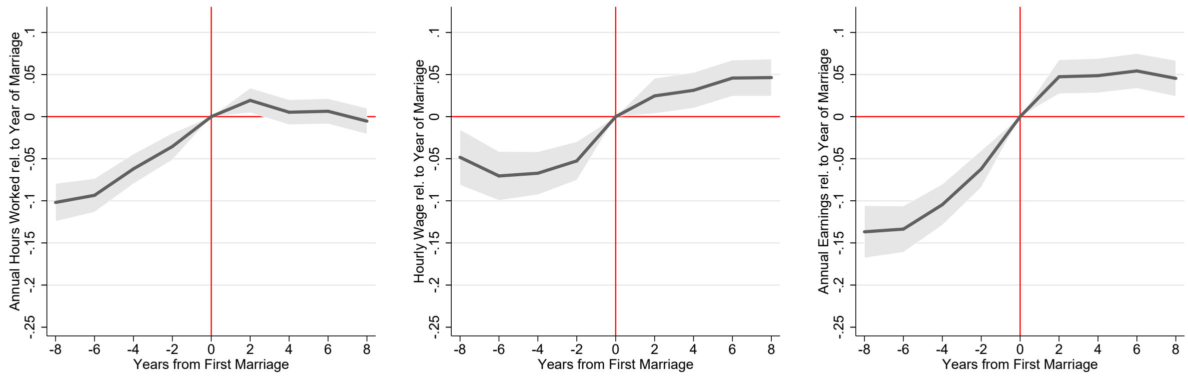
The terms $\beta_i^{individual}$ and β_t^{year} are individual and year effects. The term $\beta_d^{distance}$, $10 \leq d \leq 10$, is a "distance-from-marriage" effect, with $d = -10$ indicating ten years prior to the man's first marriage, and $d = 10$ indicating ten years after the man's first marriage. When running the regression, we exclude the coefficient at the time of marriage, $\beta_0^{distance}$, so that the reference group is men in the year they were first married. The regression excludes observations that are more than ten years away from the man's year of first marriage in either direction. To control for age effects that are independent from marriage, the hours measure $h_{i,t,d}$ equals annual hours worked divided by the average hours of married men of the same age. For example, a value of 1.1 at age 30 indicates that individual i's age-30 hours are 10% larger than the average for 30-year-old married men.[11]

Figure 4a plots the estimated values of $\beta_d^{distance}$. The figure shows that, relative to married men of the same age, annual hours increase 10% from six years before marriage to the year of marriage. Hours start increasing even before then, but the size of the increase is considerably smaller, and measured imprecisely. In any event, the shorter interval is arguably a more plausible time window for assessing the effects of marriage. Moving forward, six years after marriage,

[11] Although we control for age by normalizing hours relative to the average among men of the same age, we have encountered concerns that our estimated distance-from-marriage coefficients may nevertheless reflect age effects. To address these concerns, Figure A2 in Appendix B.3 displays the results of a placebo test of the regressions in Figure 4, in which we randomly scramble the age at first marriage of men in the regression sample, and use this to compute a placebo distance-from-marriage measure. If the estimates in Figure 4 reflected the effect of age rather than distance from marriage, we would expect the placebo regression to yield similar estimates for the distance-from-marriage coefficients, since the labor market outcomes and age of men are identical in the two regressions. However, the placebo estimates are virtually all insignificant. This provides confidence that our original estimates are in fact reflecting the effect of distance from marriage.

Figure 4: Labor Market Dynamics in the Years around Marriage

(a) Annual hours worked (b) Hourly wages (c) Annual earnings



Source: Males ages 19 - 54 in the NLSY79; see text for details. The solid line plots distance-from-marriage coefficients from the individual fixed effects regression equation (1). The shaded region corresponds to 95% confidence intervals.

mens' relative hours are essentially unchanged from the year they were married.[12]

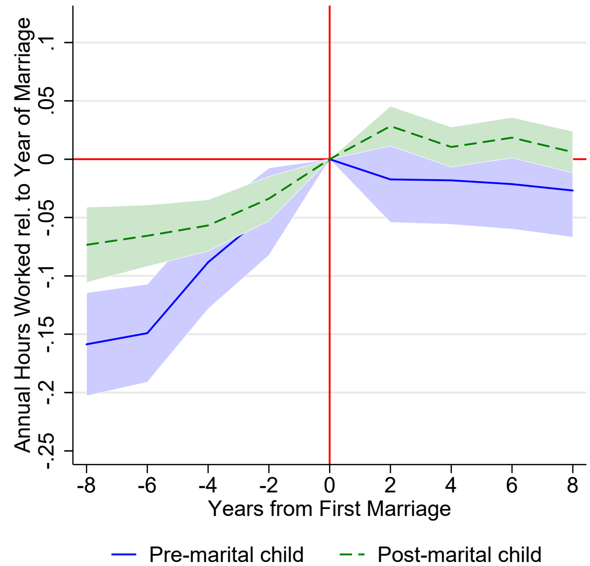
Figure 4b shows the results from a parallel regression of hourly wages on distance from marriage. Qualitatively, we observe a 'S-shape" to the wage coefficients, with a sharp increase in the years around marriage, and then a leveling-off several years after marriage. The wage and hours coefficients differ in at least one notable way. Much of the increase in hourly wages occurs *after* the increase in hours. In particular, nearly half of the increase in wages occurs after the year, while the increase in hours occurs almost entirely at or before marriage.

Figure 4c shows the results for annual earnings. The picture is very roughly the sum of the coefficients for annual hours and hourly wages in Figures 4a and 4b: earnings increase 18% from six years before marriage to two years after marriage, and are essentially flat afterward. We emphasize that, in a pure accounting sense, half of the increase in annual earnings is attributable to an increase in hours worked, rather than wages. From this perspective, understanding the "hours premium" appears to be at least as important as understanding the "wage premium" emphasized in the existing literature (see Section 1).

Marital status is highly correlated with the presence of children. A natural question is whether the changes in labor market outcomes that we have documented are more closely related to the onset of marriage or to the arrival of children. To investigate this, we run separate regressions for men whose first child appears before their first marriage, and men whose first child appears after their first marriage. Figure 5 displays the results, with the solid blue line corresponding to

[12]Figure A3 decomposes the change in annual hours worked around marriage into changes in hours per workweek and changes in annual weeks worked. Each margin generates roughly 50% of the increase in annual hours.

Figure 5: Hours Worked, Marriage and Children



Source: Males ages 19 - 54 in the NLSY79; see text for details. The solid line refers to men whose first child arrives before his first marriage ("pre-marital"). The dashed line refers to men whose first child arrives after his first marriage ("post-marital"). The lines plot distance-from-marriage coefficients from the individual fixed effects regression equation (1). The shaded regions correspond to 95% confidence intervals.

men with pre-marital children and the dashed green line corresponding to men with post-marital children. The results indicate that both groups of men experience significant increases in hours around the time of their first marriage. An important difference is that the increase in hours is more abrupt for men with pre-marital children. For example, in the six years before marriage, relative hours for men with pre-marital children increase by 16%, compared with 6% for men with post-marital children. Because the NLSY79 data also show that pre-marital pregnancies are more likely to be unplanned,[13] these results suggest that marriage and children have causal effects that encourage work. We also show in Appendix B.5 that labor market outcomes evolve similarly as children arrive.

[13]The NLSY79 asked whether pregnancies were unplanned in 1982 and every other year thereafter (even when the survey was annual). The responses to this question reveal that pre-marital pregnancies are three times as likely to be unplanned, and half as likely to be planned. In particular, among married men with one child or no children but one on the way, 72% reported that their first child was planned, 16% reported the child was unplanned, and 12% reported the child was neither planned nor unplanned. Among never-married men, 36% reported that their first child was planned, 45% reported the child was unplanned, and 19% reported the child was neither planned nor unplanned.

2.4 Summary

The data show a strong positive relationship between marriage and market work for prime-age men. In the cross-section, married men work substantially more than never-married men. This cross-sectional relationship has been fairly stable in the US since at least the mid 1970s.

Panel data regressions reveal that much of the cross-sectional difference in work by marital status is due to individual fixed effects. The data also reveal, however, that a substantial part of the difference is due to increases in hours around the time when men first marry. One possible explanation for the increase in hours is that marriage leads men both to work more once they marry and to work more in anticipation of marriage. An alternative explanation is dynamic selection, where persistent wage shocks that increase hours of work create or coincide with an increased likelihood of marriage. Because the two explanations have very different implications for how marriage affects male labor market outcomes, we would like to measure the importance of each. The results presented here provide some clues. For example, the increase in average hours is just as large as, and begins before, the increase in hourly wages that occurs around marriage. This suggests that shocks to wages are not the sole reason hours increase in the lead-up to marriage. We will quantify these two channels in the structural analyses below.

3 Model

3.1 The Life Cycle

We study a life-cycle model of male labor supply and saving. The man's age, j, is discrete. He is endowed with education level $e \in \{nc, c\}$ (non-college, or college). Non-college men enter the model at age $J_{nc} = 19$. College men are absent from the model during ages 19-22, and enter at age $J_c = 23$. Regardless of education, men retire exogenously at age J_R, and die at age J.

In addition to differences attributable to age, education, and realizations of multiple shocks, each man belongs to one of two permanent "types", indexed by $\ell \in \{1, 2\}$. Types differ by mean wage, the probability of becoming married, and taste for work. We include type differences because the hours regressions in Table 1 show that fixed effects substantially reduce the estimated coefficient on marriage. This suggests that permanent unobserved heterogeneity is important to understanding marriage-related differences in labor market outcomes.

3.2 The Man's Wage Process

In each period before retirement, $j < J_R$, men supply labor hours h_j. Male earnings are given by

$$me_j = w_j h_j^{1+\zeta}, \tag{2}$$

where w_j denotes a base hourly wage, and $h_j^{1+\zeta}$ is a "part-time penalty / overtime bonus" if $\zeta > 0$. The base wage w_j follows an AR(1) process with an age-, education-, and type-specific mean:

$$\log w_j = \alpha^w_{e,\ell,j} + \tilde{w}_j, \tag{3}$$

$$\tilde{w}_j = \rho^w_e \tilde{w}_{j-1} + \varepsilon^w_j, \tag{4}$$

$$\varepsilon^w_j \sim N(0, \sigma^\varepsilon_e), \text{ i.i.d.}, \tag{5}$$

$$\tilde{w}_0 \sim N(0, \sigma^{w_0}_e), \tag{6}$$

with $\tilde{w}$ independent of ℓ. In the notation above, and throughout the rest of this paper, superscripts are used to differentiate parameters, while subscripts indicate dependencies. For example, $\alpha^w_{e,\ell,j}$ is the mean-shifter for wages, w, for a man of education level e, type ℓ and age j.

3.3 Family Structure and Family Dynamics

The structure of the man's family is described by the triple $f = (r, a, n)$. The first component, r, denotes relationship status. Men can be single ($r = sn$), engaged ($r = en$), married ($r = mr$), or divorced ($r = dv$). To allow for selection, we allow the relationship transition probabilities to depend on a man's type (ℓ) and his wage shock ($\tilde{w}$); but they do *not* depend on the hours component h^ζ, which is endogenous. The second component denotes the age of any children, $a \in \{0, yc, oc, gc\}$, corresponding to no children, young children (ages 0-5), older children (ages 6-18), and grown children, respectively. We distinguish between young and older children because younger children are more expensive, imposing higher formal child care costs and discouraging spousal employment, as detailed in Sections 3.4.1 and 3.4.2. To simplify the model, we assume that all children in a household belong to the same age group. The third component denotes the number of children, $n = 0, 1, ..., 3$. As in Cubeddu and Ríos-Rull (2003), we treat family structure as an exogenous stochastic process.

3.3.1 Fertility Dynamics

A childless man ($n = 0$) of age j, education e, and relationship status r will have one (young) child next period with probability $\varphi^n_{0,r,e,j}$. As long as his children are young, additional offspring are possible. A man with $0 < n < 3$ young children will have $n+1$ children next period with probability $\varphi^n_{n,r,e,j}$ and n children with probability $1-\varphi^n_{n,r,e,j}$. Once a man's young children have aged, he does not have additional children. Divorced men have no additional children: $\varphi^n_{n,dv,e,j} = 0$. It bears noting that type affects fertility only indirectly, through the effects on relationship status r that we describe in the next subsection.

Children age stochastically and all at the same time.[14] Young children, $a = yc$, evolve to older children, $a = oc$, with probability $\varphi^a_{yc,e,j}$, and older children evolve to grown children, $a = gc$, with probability $\varphi^a_{oc,e,j}$. In families with young children, the aging shock occurs after the fertility shock. Having a grown child is an absorbing state. Finally, we assume that $\varphi^n_{n,r,e,j} = 0$, $\forall j \geq J_R - 2$, $\varphi^a_{yc,e,J_R-2} = 1$, and $\varphi^a_{oc,e,J_R-1} = 1$, which ensures that all children are grown by retirement. For clarity, we list the full set of child transition probabilities in Appendix C.1.

3.3.2 Relationship Status

Although some men enter the model married, most marry later. Childless men advance up the "relationship ladder" as follows: single men become engaged with probability $\phi^{en}_{e,\ell,j}(\tilde{w})$, and engaged men marry with probability $\phi^{mr}(\tilde{w})$. In addition, a never-married (single or engaged) man can have an out-of-wedlock birth, which makes marriage more likely; this "shotgun marriage" effect is motivated by the higher marriage rates observed in the first few years following an out-of-wedlock birth. We assume that a man with an out-of-wedlock birth faces "double jeopardy": because of the birth, his relationship status advances one stage with probability ϕ^{owb}_e; should this "shotgun advancement" not occur, he still faces the "regular" probability of advancing faced by childless men. For simplicity, we assume that only the first out-of-wedlock birth has this effect. Once married, men divorce with probability $\phi^{dv}_{n,e,j}$. Divorce is an absorbing state, i.e., we rule out re-marriage. Finally, relationships are fixed once an individual reaches retirement.

The transition probabilities allow for two types of selection. The first is permanent type-driven selection indexed by ℓ. The second is dynamic wage selection driven by shocks to the AR(1) process $\tilde{w}$. Both types of selection operate through the transition probabilities $\phi^{en}_{e,\ell,j}(\tilde{w})$ and $\phi^{mr}(\tilde{w})$. For clarity, we list the full set of relationship transition probabilities in Appendix C.2.

[14] Stochastic aging indirectly captures the uncertainty inherent in the costs of children, since there is variation ex-post in the time it takes for children to mature. The assumption that children age at the same time also simplifies the computation of the model by reducing the dimension of the children's' age space.

3.3.3 Cohabitation

Many couples begin pooling resources prior to marriage. For example, in the NLSY79, 49% of men reported cohabiting with their partner before their first marriage, with an average length of two years among those cohabiting. To capture this pattern, we allow engagement to take one of two forms, non-cohabitation, *en-n*, and cohabitation, *en-c*. For non-cohabiting couples, *en-n*, the man's budget constraint and preferences are identical to those of a single man. For cohabiting couples, *en-c*, the man's budget constraint and preferences are identical to those of a married man, except that the man and his partner file taxes separately rather than jointly. Throughout the paper, we will use the term "spouse" to denote partners of both cohabiting and married men.

When a couple first becomes engaged, they are assigned permanently to one of the two engagement types, cohabiting with probability $\phi^{en\text{-}c}$. Future fertility and relationship dynamics are independent of whether the engagement involves cohabitation.

3.4 Family Structure and Financial Resources

Relationships and children affect a man's financial resources in four ways. First, in larger households consumption must be spread across more individuals. We capture this effect through the use of equivalence scales that convert total consumption to per capita amounts. In doing this, we assume that children live in their father's household only if he is married or cohabiting. Second, spouses can generate earnings or, if they stay home with children, substitute for costly formal child care. Third, non-grown children are costly. Men with working spouses pay for formal child care, and men who are neither married or cohabiting pay child support. Finally, couples who divorce split their wealth in half. The possibility of such a split tends to reduce the husband's expected consumption, since, in the quantitative model, he usually has the higher earnings.

3.4.1 Spousal Earnings

At the time that a man becomes engaged, J_{en}, his future spouse draws the permanent earnings shock

$$\tilde{s} \sim N(\rho_e^s \tilde{w}_{J_{en}}, \sigma_e^s). \tag{7}$$

The spousal earnings shock $\tilde{s}$ is potentially correlated with the man's AR(1) shock at the time of engagement, $\tilde{w}_{J_{en}}$. We will estimate ρ^s from the data.

Men who are cohabiting or married combine their earnings with any earnings that their spouses receive. Given $\tilde{s}$, spousal earnings at age $j \geq J_{en}$, se_j, follow a two-stage process. The first stage uses a logit model to determine whether the spouse works:

$$\mathbb{1}_{se_j>0} = \begin{cases} 1 & \text{if } q_j < \frac{\kappa}{1+\kappa}, \\ 0, & \text{otherwise}, \end{cases} \tag{8}$$

$$\kappa = \exp\left(\tilde{s} + \alpha^{s,1}_{a,e,j}\right), \tag{9}$$

$$q_j \sim U[0,1], \text{ i.i.d.}, \tag{10}$$

where $\mathbb{1}_{\mathcal{A}}$ is the indicator function for event $\mathcal{A}$. The second stage determines the earnings of spouses who work:

$$se_j = \mathbb{1}_{se_j>0} \cdot \exp\left(\log(\underline{s}) + \max\{0, \beta^s_e \tilde{s} + \alpha^{s,2}_{e,j}\}\right). \tag{11}$$

In the above equations, β^s_e is a wage-scaling parameter, and $\alpha^{s,1}_{a,e,j}$ and $\alpha^{s,2}_{e,j}$ are mean-shifters that depend on the man's age and education. The mean-shifter for spousal participation, $\alpha^{s,1}_{a,e,j}$, also depends on the age of the children. Spouses with young children are least likely to work, spouses with no (or grown children) are most likely, and spouses with older children fall in between. The parameter $\underline{s}$, which places a lower bound on the earnings of working spouses, equals the earnings cutoff we use to define working spouses in the data.

In our model of spousal earnings, spouses with higher potential earnings are more likely to work. While a standard Tobit model would generate a similar relationship, we found that to fit the spousal earnings data well, we needed a more flexible specification. Even though the shock $\tilde{s}$ is permanent, spouses will move in and out of employment as $\alpha^{s,1}_{a,e,j}$ and q_j vary over the life cycle.

3.4.2 Child Costs

Married and cohabitating couples must provide care to their non-adult children. If the spouse does not work, $se_j = 0$, she provides the child care herself at no additional cost to the family. If the spouse works, then the couple must purchase child care in the market, at a total cost of $n_j \chi_{a_j} se_j$. We assume that child care costs are proportional to the number of children and to the spouse's earnings. The cost factor χ_a depends on the children's' age; older children ($a = oc$) are less expensive than younger children, and grown children ($a = gc$) impose no costs at all.

We assume that men who are neither married nor cohabiting do not live with their children and instead pay child support.[15] Child support equals the fraction $n\delta_a$ of the father's labor income. The parameter δ_a equals δ for young and older children, $a \in \{yc, oc\}$, and zero for grown children.

[15] We assume that when a single father marries, his children join his new household.

3.4.3 Divorce Costs

At the time of divorce, men lose half their assets. Divorced men also pay a fraction of their earnings, $n\delta_a$, in child support.

3.5 Preferences

Men have time-separable preferences over consumption and hours of work that vary with family structure, education and age:

$$u_{f,e,\ell,j}(c,h) = N_f \frac{(c/\eta_f)^{1-\gamma}}{1-\gamma} - \psi_{e,\ell,j} \frac{h^{1+1/\xi}}{1+1/\xi}, \tag{12}$$

with $\gamma, \xi > 0$. The parameter η_f is a household equivalence scale converting total household consumption, c, into the per capita amount consumed by the man. The shift term N_f captures the possibility that married and cohabiting men derive additional utility from the consumption of other household members. Alternatively, N_f may reflect social norms requiring men to "provide" for their families – we are agnostic as to the exact interpretation. Children living outside the man's household do *not* affect η_f or N_f. The labor disutility shifter $\psi_{e,\ell,j}$ varies with education, type and age. Future utility is discounted at the rate $\beta \in (0,1)$.

3.6 Total Income and Taxes

A household's total income, y, equals the sum of male earnings me, spousal earnings se (for married and cohabiting men) and capital income:

$$y_j = me_j + se_j + (R-1)k_j, \tag{13}$$

where k denotes the household's assets, and R is the constant gross rate of return. It faces payroll and income taxes:

$$T(me+se,k;f) = \tau^{ss}(me+se) + T_f^{inc}(y), \tag{14}$$

$$T_f^{inc}(y) = \left[\tau_f^0 + \tau_f^1 y^{\tau_f^2}\right] y, \tag{15}$$

where τ^{ss} is the payroll tax rate, and $T^{inc}(\cdot)$ is the income tax function. We allow the parameters of $T^{inc}(\cdot)$ (τ_f^0, τ_f^1, and τ_f^2) to differ by marital status and the number of dependent children.

3.7 Recursive Formulation

The state vector for an age-j man consists of the man's education level (e), type (ℓ), assets (k_j), wage deviation ($\tilde{w}_j$), relationship status (r_j), age and number of children (a_j, n_j) and, for men who are engaged or married, spousal earnings shocks ($\tilde{s}$ and q_j). We will continue to use $f = (r, a, n)$ as a compact index of family structure.

Appendix C.3 presents the full set of Bellman equations. In the interest of brevity, the description provided here is more condensed.

3.7.1 Single Men

The Bellman equation for a working-age ($j < J_R$) single man, $r = sn$, is

$$V_j^{sn}(e,\ell,k,\tilde{w},a,n) = \max_{c,k',h} \Big\{ u_{f,e,\ell,j}(c,h) \tag{16}$$

$$+ \beta \mathbb{E}_{\tilde{w}',a',n',r',\tilde{s}',q'} \left[V_{j+1}^{r'}(e,\ell,k',\tilde{w}',a',n',\tilde{s},q') \,\middle|\, \tilde{w},n,a,e,\ell \right] \Big\}$$

$$s.t. \quad c + k' \leq Rk + wh^{1+\zeta}(1 - n\delta_a) - T(wh^{1+\zeta}, k; f), \tag{17}$$

$$\log w = \alpha^w_{e,\ell,j} + \tilde{w}, \tag{18}$$

$$k' \geq k_{min}. \tag{19}$$

A man's expected continuation value depends on the potential evolution of his wage and family structure. A single man may father children or have his children age. The child support payments for these children act as an earnings tax. A single man may stay single or become engaged. His own wage $\tilde{w}$ will evolve and, should he become engaged, the spousal earnings shocks $\tilde{s}$ and q_j will be realized.

3.7.2 Engaged Men

A fraction $\phi^{en\text{-}c}$ of engaged men cohabit with their partner, while the remainder do not. Engaged men who do not cohabit with their partner have a Bellman equation identical to that of single men, except that their one step-ahead relationship possibilities do not include remaining single but include becoming married instead.

Engaged men who cohabit with their partner have a somewhat different Bellman Equation: (i) their flow utility function accounts for co-residing partners and children (through η_f and N_f); (ii) their budget constraint includes spousal earnings and the taxes paid on these earnings; (iii) because children are assumed to now reside with the father, child support costs, which were proportional to the man's earnings, are replaced with child care costs, proportional to the earnings of his partner. This yields the following budget constraint:

$$c+k' \leq Rk+wh^{1+\zeta}+se(1-n\chi_a)-T(wh^{1+\zeta},k;f^{en\text{-}c})-T(se,0;f^{en\text{-}c,sp}), \tag{20}$$

Cohabiting couples pool their labor income together, but each partner files taxes as an individual. We capture this by replacing f with $f^{en\text{-}c}$ and $f^{en\text{-}c,sp}$ in the tax functions.[16] Consistent with our assumption that spouses bring no wealth into the relationship, we assign all of the cohabiting couple's asset income to the man.

3.7.3 Married Men

The Bellman equation for a working-age married man, $r = m$, is identical to that of engaged-and-cohabiting men, except that his future relationship possibilities are to remain married or get divorced, and married couples file taxes jointly.

3.7.4 Divorced Men

The Bellman equation for a working-age divorced man, $r = dv$, is identical to that of singles, except that there is no uncertainty over future relationships (divorce is an absorbing state) or fertility (divorced men have no additional children).

3.7.5 Retired Men

Retired men ($j \geq J_R$) do not work, and if they are married or cohabiting, their spouses do not work. Their only income comes from their assets and from Social Security benefits received by the man ($b_{1,e}$) and his spouse ($b_{2,e}$). All children are grown in retirement, implying that there are no child care costs or child support payments due. Finally, relationships do not change in retirement, eliminating any uncertainty due to them. The resulting Bellman equation is completely

[16] We assume that the man, who is most likely to have the higher income, will be the partner who claims the children as tax dependents. This means that $f^{en\text{-}c} = (sn, n, a)$ and $f^{en\text{-}c,sp} = (sn, 0, 0)$.

deterministic:

$$V_j^{ret}(e,k,r) = \max_{c,k'} \quad u_{f,e,j}(c,0) + \beta V_{j+1}^{ret}(e,k',r), \tag{21}$$

$$s.t. \quad c + k' \leq Rk + b_{1,e} + b_{2,e}\mathbb{1}_{r\in\{mr,en\text{-}c\}} - T(b_{1,e} + b_{2,e}\mathbb{1}_{r\in\{mr,en\text{-}c\}}, k; f), \tag{22}$$

equation (19),

$$V_J \equiv 0. \tag{23}$$

Given that hours are identically zero and relationships are static, there are no type-related differences among retirees.

4 Model Parameters

We set the parameters of our model in three steps. First, we set a number of parameters to values consistent with the broader literature. In the second step, we estimate the stochastic processes for fertility, relationships, wages and spousal earnings, which can be identified outside our behavioral model, from the data. While the principal dataset used in this step is the NLSY79, we also utilize state variation in the CPS. The final step of our estimation process is to use the model to estimate the discount factor and the age-varying component of the disutility from work.

We also use the final step to estimate type-specific differences in wages, relationship transitions, and preferences for work, along with the type probabilities. These are set so that the model replicates observed differences in hours and wages between never-married and ever-married men. Although the stage-two and stage-three parameters are conceptually distinct in terms of identification, some of the type-specific parameters affect the parameter estimates in the second stage. We thus estimate the parameters for the two stages "semi-jointly", as described in the appendices.

4.1 Parameters Taken from Other Studies

Appendix Table A3 displays the values for parameters taken from other studies. The first panel of the table shows that non-college and college men enter the model one year after their modal graduation ages, 19 and 23, respectively. They retire at age $J_R = 65$ and die at age $J = 80$.

We set the consumption utility curvature parameter, γ, to 0.738, following Imai and Keane (2004). Estimates of this parameter vary widely (see the discussion in De Nardi, French and Jones (2010)). With separable utility, however, a value of γ greater than 1 would in a static model imply that the income effects of a wage change are large relative to the substitution effects. Given that many of the younger individuals in our model live nearly hand-to-mouth, consistent with large income effects, using $\gamma \geq 1$ would imply that young men would sometimes respond to wage increases by working fewer hours. This would rule out by construction the hypothesis that the higher hours of married men are due to their higher wages. Because we want to explore this hypothesis as an alternative to the mouths-to-feed mechanism, setting γ to a value less than 1 is appropriate. Sensitivity analyses in Appendix I show that the effects of marriage on hours are robust to this parameter.

Our choice of the Frisch elasticity, $\xi = 0.75$, lies in the middle of a wide range (Keane and Rogerson, 2012). In a recent paper, Bick, Blandin and Rogerson (2022), applying the approach for two-earner households developed by Bredemeier, Gravert and Juessen (2019), find elasticities ranging from 0.51 to 1.07.

We set N_f, which scales the utility from per capita consumption, to 2 for married and cohabiting men and 1 for the rest; we are effectively assuming that married and cohabiting men receive utility from their spouses' consumption. The literature provides little guidance for setting N_f. As we show in Section 6.3 below, the model is able to match the run-up in hours around the date of marriage only if N_f is larger for married men.

Appendix D describes how we set the remaining parameters.

4.2 Type-related differences

Each individual belongs to one of two unobserved types, indexed by $\ell \in \{1,2\}$. We will use $\ell = 1$ to index the "non-marrying" type and use p_1 to denote the probability that a man belongs to this type. The two types differ by mean wage, the probability of getting married, and taste for work. These differences are expressed as zero-mean deviations: this will at times greatly simplify our estimation procedure.

To account for type-specific differences in wages, we adjust the log shifter α^w as follows:

$$\alpha^w_{e,\ell,j} = \begin{cases} \alpha^w_{e,j} - \Delta^w, & \ell = 1, \\ \alpha^w_{e,j} + \frac{p_1}{1-p_1}\Delta^w, & \ell = 2. \end{cases} \tag{24}$$

Our normalization implies that Δ^w is positive: type-1 men have lower average wages.

The types also differ in the probability of remaining single. Recall that in the absence of a new child, the probability of engagement is ϕ^{en}, so that the probability of remaining single is

$\phi^{sn} = 1 - \phi^{en}$. Type effects enter as

$$\phi^{sn}_{e,\ell,j}(\tilde{w}) = \begin{cases} \phi^{sn}_{e,j}(\tilde{w})\big(1+\Delta^{sn}\big), & \ell = 1, \\ \phi^{sn}_{e,j}(\tilde{w})\big(1-\frac{p_1}{1-p_1}\Delta^{sn}\big), & \ell = 2. \end{cases} \tag{25}$$

Our prior belief is that Δ^{sn} will be positive: type-1 men are more likely to remain single. To simulate our model, we also need to know the distribution of relationships when individuals enter the model, $r_{J_{nc}}$ and r_{J_c}. To do this, we adjust the initial probability of being single in a manner parallel to that shown in equation (25), then adjust the remaining probabilities proportionally. We will use Δ^{s0} to denote the initial share deviation.

The third way the two types differ is in their disutility from work, $\psi_{e,j}$, namely:

$$\psi_{e,\ell,j} = \begin{cases} \psi_{e,j}\big(1+\Delta^{\psi}\big), & \ell = 1, \\ \psi_{e,j}\big(1-\frac{p_1}{1-p_1}\Delta^{\psi}\big), & \ell = 2. \end{cases} \tag{26}$$

Our prior belief is that Δ^{ψ} will be positive: type-1 men are less willing to work.

We estimate the 5 type-related parameters (p_1, Δ^{w}, Δ^{sn}, Δ^{s0}, Δ^{ψ}) using a grid search, targeting the differences in hours and wages between the never-married and the ever-married, by age. The data show that never-married men have lower wages and hours: this suggests that type-1 men, with lower wages and a greater distaste for work, should be less likely to form relationships and eventually marry.

Although the type-related parameters belong to the third and final stage, where we compare the endogenous behavior predicted by our model to the data, some of these parameters affect the parameter estimates for the exogenous stochastic processes found in the second stage. We therefore estimate the second- and third-stage parameters jointly. This is computationally challenging. Fortunately, there are a number of ways to streamline this process: we discuss the shortcuts in Appendix E.[17] Because the second and third stage parameters are distinct in terms of identification, in the remainder of this section we will discuss them sequentially, leaving a description of their mechanical interactions to the appendices.

4.3 Parameters Estimated Outside the Behavioral Model

We estimate three sets of parameters outside the behavioral model: (i) the parameters governing the stochastic process for male wages; (ii) the parameters determining the probability that a spouse

[17] Even with the shortcuts, we are only able to estimate the parameters using the Big Tex computer cluster.

works and her earnings when working; and (iii) the parameters determining the stochastic process for family structure.

4.3.1 The Male Wage Process

Using equation (2), we compute the hourly wage term w_j for an individual with annual earnings me_j and annual hours worked h_j as

$$w_j = me_j / h_j^{1+\zeta}. \tag{27}$$

This gives us a measure of hourly wages consistent with our model. We estimate the wage process for men from equations (3)-(5) in two steps, following French (2005). First, we run an individual fixed effects regression of log wages on a quadratic in age and a control for the national unemployment rate during January of that calendar year. The estimated coefficients for the quadratic in age, along with the average fixed effect, provide us with values for $\alpha_{e,j}^{w}$. In the second step, we calculate a residual wage for each individual, which is the difference between the log of his actual wage and the predicted wage $\alpha_{e,j}^{w}$ (plus the estimated unemployment effect). We then generate the covariance matrix for the first four lags of this residual, which we use to estimate the autocorrelation term ρ_e^w, the innovation standard deviation σ_e^{ε}, and the initial standard deviation $\sigma_e^{w_0}$ using the Simulated Method of Moments. The details of this procedure, including how it incorporates type-specific differences in α^w, are described in Appendix E.1. The first panel of Table 2 presents the estimates.

4.3.2 Spousal Earnings

The spousal earnings process in equations (7)-(11) requires estimates of: the dependence of the spouse's permanent earnings shock, $\tilde{s}$, on the husband's wage (ρ_e^s), along with the standard deviation of this shock's innovation (σ_e^s); the parameters of the mean-shifter for the probability that a spouse works ($\alpha_{a,e,j}^{s,1}$); the parameters of the mean-shifter for the earnings of working spouses ($\alpha_{e,j}^{s,2}$); and the relative importance of the permanent shock for spousal earnings (β_e^s).[18] We assume that both mean-shifters contain a quadratic polynomial in the husband's age, and that the mean-shifter for spousal employment contains coefficients for the presence of young or old children (the base case is no children or grown children). We also set the floor for earnings to $\underline{s} = \$3,630$ and

[18] In doing so, we utilize the estimated fertility and relationship probabilities described in the section immediately below.

Table 2: Parameters for Male Wages and Spousal Earnings

Description	Parameter	Value, Non-College	Value, College Graduates
Male Wages			
Mean-shifter, age quadratic	α^w_j	$(-1.555, 0.067, -0.072)$	$(-2.445, 0.129, -0.145)$
Autocorrelation	ρ^w	0.887	0.916
Standard deviation, innovation	σ^ε	0.213	0.190
Standard deviation, initial value	σ^{w_0}	0.086	0.239
Spousal Earnings			
Effect of man's wage on spousal employment	ρ^s	0.842	0.720
Standard deviation, innovation	σ^s	0.020	0.614
Effect of children on employment	$\alpha^{s,1}_{a,j}$: yc, oc	$(-1.115, -0.580)$	$(-0.065, 0.200)$
Effect of man's age on spousal employment, quadratic	$\alpha^{s,1}_{a,j}$	$(0.805, 0.088, -0.236)$	$(-0.5, 0.100, -0.100)$
Effect of man's age on spousal earnings, quadratic	$\alpha^{s,2}_j$	$(1.302, 0.056, -0.100)$	$(1.601, 0.055, -0.100)$
Effect of man's wage on earnings	β^s	19.97	1.01

Note: Superscripts are used to distinguish parameters, while subscripts are used to distinguish dependencies. We have omitted education (e) subscripts, as every parameter varies by education level. All parameters with the age subscript j utilize a quadratic in age. See sections 4.3.1-4.3.2 and Appendix E.1 for details.

censor all spousal earnings in the data below this level.[19]

The bottom panel of Table 2 presents our parameter estimates. We estimate these parameters using the Simulated Method of Moments, targeting age profiles for the share of spouses who work, the mean of log earnings among working spouses, the correlation of spousal earnings and male wages, and the standard deviation of log earnings among working spouses. We construct separate age-employment profiles for spouses with young and older children (determined by the age of the oldest child). Appendix F shows the model's fit of spousal employment and earnings. Consistent with the data, the model predicts that the employment rate for women with young children is about 20 percentage points less than the rate for women with no children. The employment rate for women with older children lies between these cases, but is closer to the childless rate. In contrast, among spouses who work, earnings are close to invariant, at least on average, over the number of children.

Appendix F shows that observed spousal earnings are quite variable, leading to large esti-

[19]This is the annual earnings from working ten hours per week for 50 weeks at \$7.26/hour, which is the median federal minimum wage over this time period.

mated values of the scaling parameter β^s. In contrast, our estimated values of ρ^s, which links the spousal earnings shock $\tilde{s}$ to male wages, are relatively small, so that much of the variation in spousal earnings is specific to the spouse. As a result, in our model the correlation between male wages and spousal earnings (among workers, conditional on age, education and number of children) never exceeds 0.3 and is often much lower. This is consistent with the observed correlations that we target. Our spousal earnings process thus generates positive, though quantitatively modest, assortative matching, at least along the intensive margin.

4.3.3 Family Structure Dynamics

The transitions governing the dynamics of relationships and children (Appendix equations A1-A3) are modeled as a set of logistic probabilities. We estimate these probabilities using the Simulated Method of Moments, matching family demographics over the life cycle in the NSLY79, along with a moment capturing the effect of wages on the probability of marriage in the CPS. We include the latter moment because in the model, the probability of "moving ahead" in a relationship depends in part on wages: the transition probabilities $\phi^{en}_{e,j}(\tilde{w}_j)$ and $\phi^{mr}(\tilde{w}_j)$ vary with the AR(1) wage shock $\tilde{w}_j$. Estimating these effects from the data requires variation in wages that is exogenous to other determinants of marriage. Although such variation is hard to find in the NLSY79, in the CPS we can instrument for a man's wage with the average wage in his state of residence. Appendix G.2 presents detailed results. The estimated coefficient in the CPS regression (Appendix Table A6) is 0.0144, which implies that a 10% increase in wages increases the probability of getting married by 0.144 percentage points. For context, the baseline probability of marriage in the estimation is 2.5%, implying that a 10% increase in wages increases the probability of getting married by $\frac{0.144}{2.5} = 5.76\%$. We assume that in our logistic transition probabilities, the coefficient on the wage shock $\tilde{w}$, which we denote by θ, is the same at all stages of a relationship, and we set θ so that the wage coefficient on a simulated version of the CPS regression matches the observed coefficient. Appendix G describes our specification and estimation procedure in more detail, and Appendix Table A5 provides the parameter estimates. We also consider much larger values of θ (more dynamic selection) as robustness checks in Section 6.2 and in Appendix J.

4.4 Parameters Set within the Behavioral Model

In the final stage of our estimation process we use the behavioral model to estimate the discount factor and the disutility of work. We set the discount factor β to 0.98 to match mean asset holdings at age 50. We allow the disutility of work to depend on education and age through the parameter

$\psi_{e,j}$, which we set to match the lifecycle profiles of hours for each education group. To limit the number of parameters estimated, we assume that $\psi_{e,j}$ is a quadratic function of age: $\psi_{e,j} = \psi_{e,0}(1+\psi_{e,1}\cdot j+\psi_{e,2}\cdot j^2)$. Appendix H shows that the estimated disutility from work is lowest in the middle of a man's career and highest at its beginning and end.

We also use the behavioral model to estimate the type-related parameters (p_1, Δ^w, Δ^{sn}, Δ^{s0}, Δ^{ψ}). We target differences in hours and wages between never-married and ever-married men, by age. Appendix E provides additional, technical details.

Table A7 in Appendix H presents the third-stage parameter estimates. The estimates imply that probably of belonging to type 1 is $p_1 = 0.62$, while $\Delta^w = 0.12$, $\Delta^{sn} = 0.12$ and $\Delta^{s0} = 0$. This in turn implies that the wages of type-1 men are 27% lower than those of type-2 men, and their annual likelihood of remaining single is 39% higher.[20] We also find that type-1 men have a much stronger distaste for work: see Figure A8. Given that type-1 men, who are less likely to form relationships and marry, also face lower wages and a greater distaste for work, we reproduce the observed tendency of never-married men to receive lower wages and work fewer hours.

5 Properties of the Baseline Model

5.1 Life-Cycle Patterns

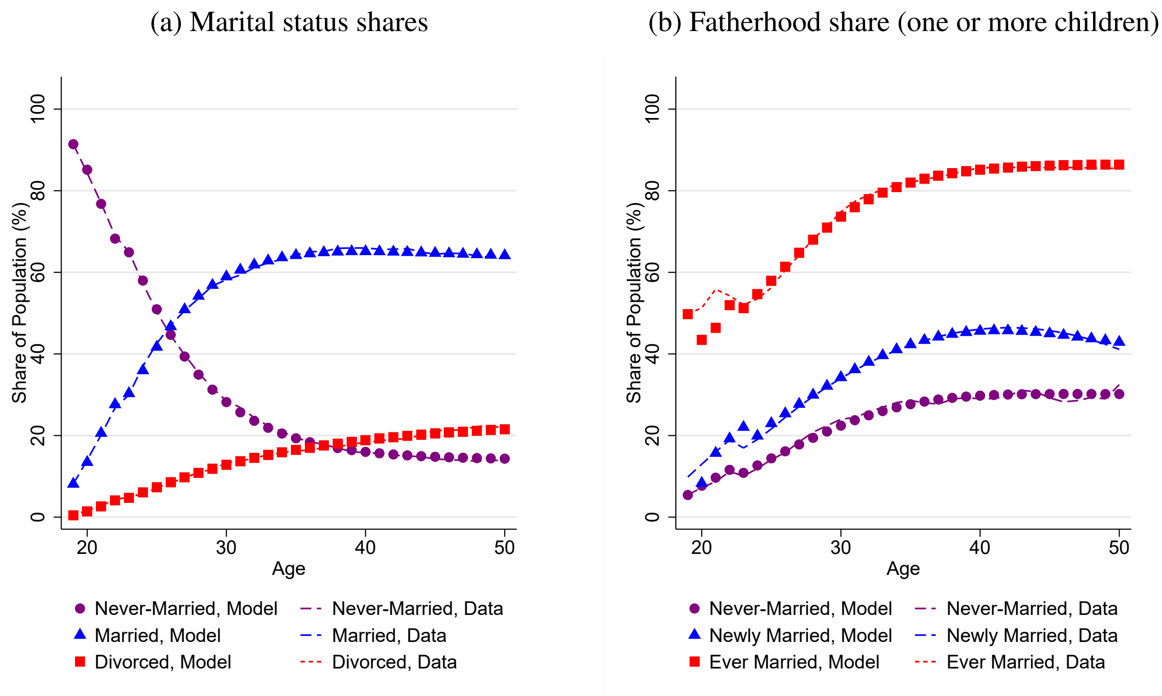
We begin with the life-cycle profiles targeted in the second and third stages of our parameter estimation process. Figure 6a compares the distribution of marital status in the model (markers) and data (lines). Figure 6b compares the rates of fatherhood (at least one child) by marital status.[21] In both figures, the model fits the data closely. Especially notable is the model's fit of the incidence of fatherhood among newly-married (men in their first year of marriage), which indicates that the model does a good job of capturing the effects of pre-marital children on relationship transitions. Figure A7 in the appendices shows that the model also does a good job of matching life-cycle patterns of family size.

Figure 7 presents two additional model-data comparisons. Figure 7a displays the life-cycle profiles for average hours worked by men who have never been married (but may marry in the future) and men who are currently married or divorced. In the data, mean hours increase rapidly from age 19 until ages 35-40, at which point they begin to gradually decline. In the model,

[20] Applying equation (24) yields the ratio $\exp(-0.12)/\exp\left(\frac{0.62}{0.38}0.12\right) = 0.729$. Applying equation (25) yields $(1+0.12)/(1-\frac{0.62}{0.38}0.12) = 1.393$.

[21] The profiles shift downward slightly at age 23 because that is the age at which college-educated men enter our sample.

Figure 6: Marital Status and Fatherhood by Age: Model and Data



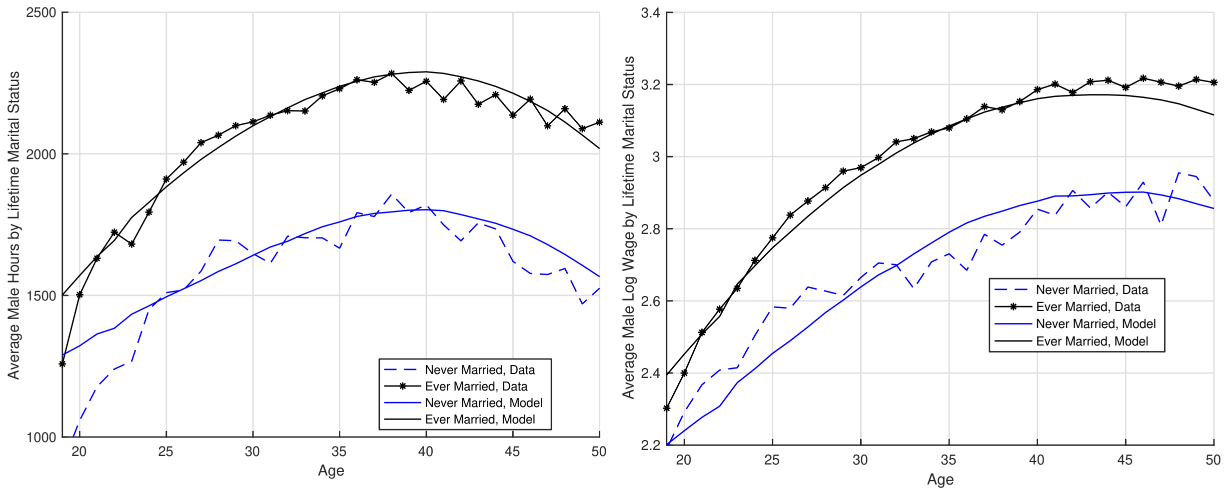
Source: Sample consists of men ages 19-50. Ages 19-22 include only men with less than a four -ear college degree; ages 23-50 include all education groups. Never-married men are those who at the age in question have yet to marry. Data correspond to the NLSY79. Model results are author calculations; see section 4.3.3 and Appendix G for details.

disutility from work is a quadratic in age, and the corresponding parameters are chosen to provide a tight fit to the hours profiles. (See Figure A8 in Appendix H.) Figure 7a shows that the model also matches the lower work hours of the never-married. Figure 7b likewise shows that, consistent with the data, the model predicts substantially lower wages for the never-married. The model replicates these differences by giving type-1 men, who are much less likely to ever marry, lower wages and a stronger distaste for work. In Appendix Figures A9 and A10, we disaggregate the hours and wage profiles even further, by education, and provide confidence intervals for the data profiles. For the most part, the model-predicted profiles lie within the confidence intervals for the data.

Figure 8 displays the life-cycle profile for mean (net) assets. While mean assets at age 50 is one of our targeted moments, assets at other ages are untargeted. Although the model profile tends to fall below its data counterpart, particularly before age 50, the two series track each other reasonably well.

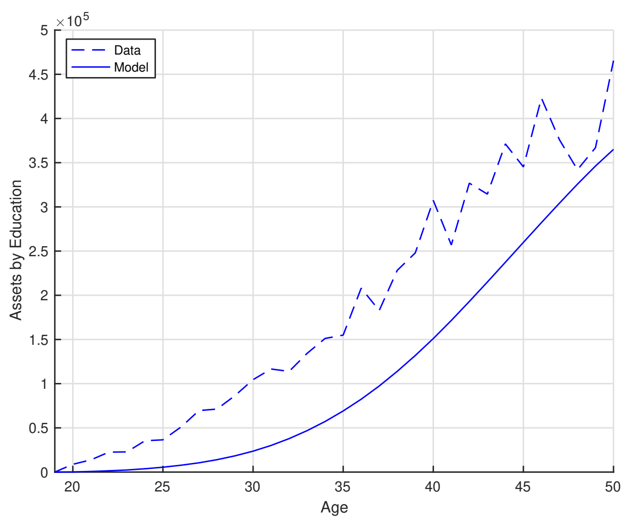
Figure 7: Mean Hours Worked and Wages by Age and Marital History: Model and Data

(a) Mean hours worked (b) Mean log wage



Source: Sample consists of men ages 19-50. Ages 19-22 include only men without a four-year college degree; ages 23-50 include all education groups. Data correspond to the NLSY79. Model results are author calculations; see section 4.4 and Appendices E and H for details.

Figure 8: Mean Assets by Age and Education: Model and Data



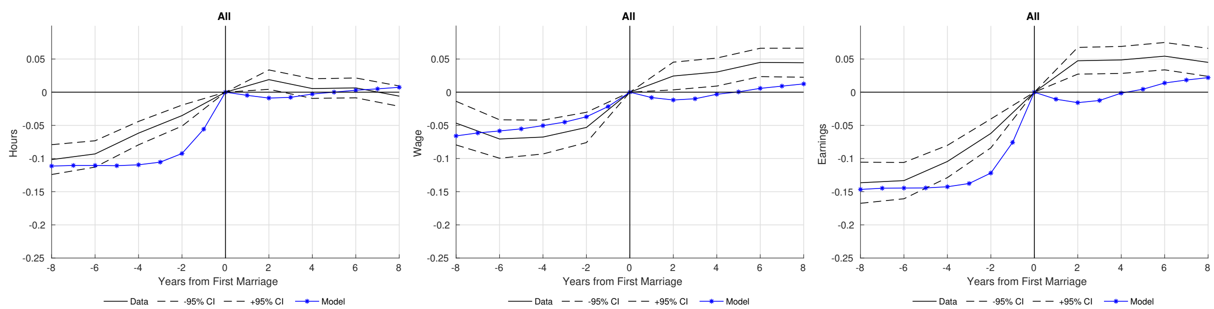
Source: Sample consists of men ages 19-50. Ages 19-22 include only men without a four-year college degree; ages 23-50 include all education groups. Data correspond to the NLSY79. Model results are author calculations; see text for details.

5.2 Marriage, Children and Male Labor Market Outcomes

The next test for the model is the extent to which it can generate the labor market outcomes observed around the time of marriage or (first) childbirth. Repeating exactly the analysis in Section 2.3, we first normalize each man's hours, wages and earnings by their age- and education-conditional averages. We then perform the fixed effect regressions described in equation (1).[22] Because we do not include these regressions in our estimation targets, comparing the data and model coefficients provides us with a validation exercise.

Figure 9: Labor Market Dynamics in the Years around Marriage: Model and Data

(a) Hours (b) Wages (c) Earnings



Source: Data results are from regressions using NLSY79 data; see section 2.3 for details. Model results are authors' calculations; see text for details.

We begin by assessing the model's ability to replicate observed hours dynamics. Figure 9a shows that the model generates the marriage-related hours growth found in the data: from six years before marriage to six years after, hours increase 12% in the model versus 10% in the data. Both model and data show a continuous increase in hours up to the year of marriage. The hours increase begins somewhat earlier in the data. The model generates pre-marital hours increases through cohabitation, which resembles marriage, and anticipatory effects. The onset of engagement, even if it does not involve cohabitation, signals that marriage has become more likely. Forward-looking men respond by increasing their hours in preparation. Our suspicion is that a relationship process that generated larger anticipatory effects would produce a more extended run-up. Consistent with the data, in the model hours change very little after marriage occurs.

Additional insight into these dynamics can be found in Appendix I, where we re-calculate the hours trajectories using different values of the coefficient of relative risk aversion (γ) and the Frisch elasticity of labor supply (ξ). Appendix I also acts as a robustness check. Over the range

[22] The regressions on the simulated data exclude year effects, as they have no counterpart in the model.

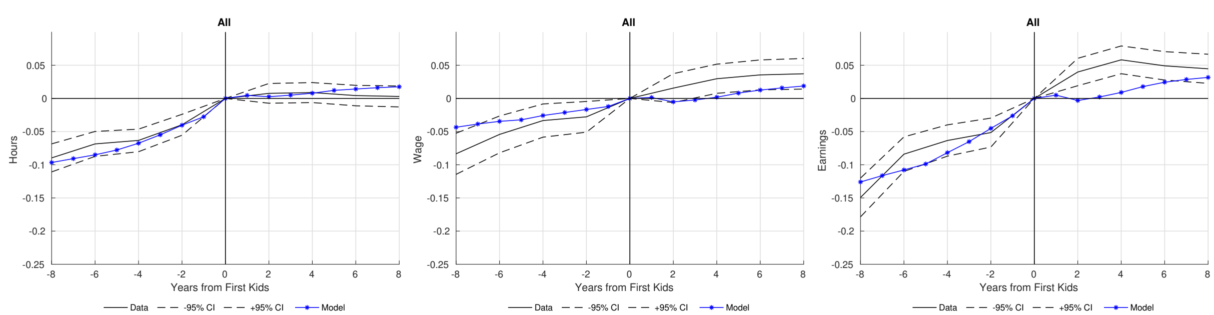
we consider (half or double baseline values), our qualitative results do not depend on specific parameter values, and the majority of the quantitative results are robust as well.

Figure 9b shows the model's implications for wages around the time of marriage. The model does a good job of matching the wage growth observed prior to marriage. In the model, wages rise in the run-up to marriage because higher wages increase the probability of marriage and because the increase in hours prior to marriage generates an increase in wages through the part-time wage penalty / overtime bonus. As we show in Section 6.2, most of the wage increase is due to the latter mechanism, rather than dynamic selection. A shortcoming of the model is that it generates very modest wage increases after the time of marriage. This could reflect the absence of human capital dynamics, such as learning-by-doing, that allow higher hours in one year to raise wages in subsequent years.

Figure 9c presents the earnings trajectory. Since log earnings are the sum of log hours and log wages and the correlation of hours and wages is fairly weak, the dynamics of mean log earnings are roughly the sum of the profiles for log hours and log wages. Overall, the model generates a 16% increase in earnings from six years before marriage to six years after marriage, compared with 18% in the data. In particular, the model matches the 13% increase in earnings prior to marriage, although the increase begins a bit earlier in the data. On the other hand, the model underpredicts the rise in earnings after marriage, because model wages do not increase greatly after marriage.

Figure 10: Labor Market Dynamics in the Years around First Child: Model and Data

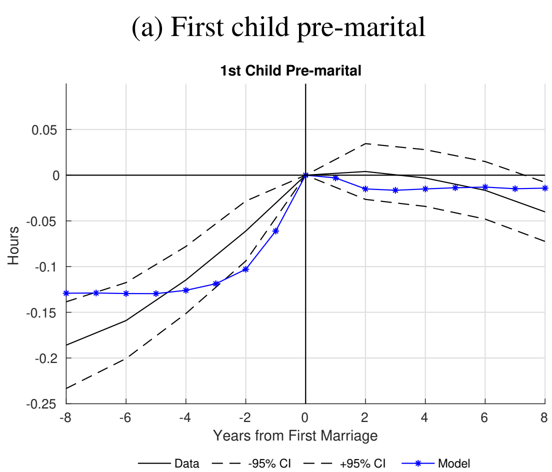
(a) Hours (b) Wages (c) Earnings

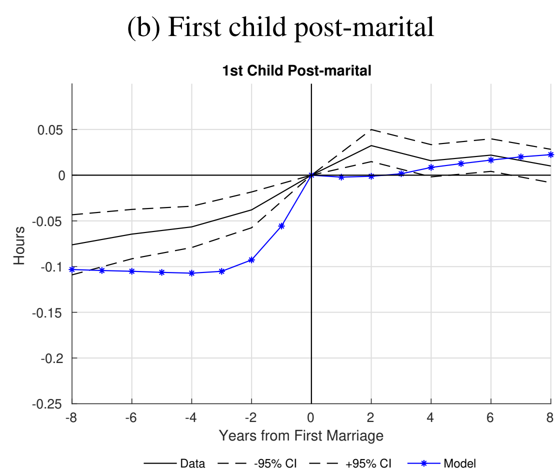


Source: Data results are from regressions using NLSY79 data; see section 2.3 for details. Model results are authors' calculations; see text for details.

It is natural to ask whether the effect of marriage on hours is primarily a response to the birth (or expected birth) of children. Figure 10a shows that hours do in fact rise steadily as the date of the first child approaches, and that the model generates a similar pattern. Figures 10b and Figure 10c show that the model does a good job of matching the wage and earnings growth observed around the time of the first child.

Figure 11: Hours of Work by Distance from Marriage and Timing of First Child: Model and Data





Source: Data results are from regressions using NLSY79 data; see Section 2.3 for details. Model results are authors' calculations; see text for details. Pre-marital children are those born in the year of marriage or before. Post-marital children are those born after the year of marriage.

Another way to assess the role of children is to compare men whose first child arrives on or before the year of marriage, and is thus "pre-marital," with men whose first child arrives after the year of marriage, and is thus "post-marital." Figure 11 shows that in both the data and the model, the run-up in hours prior to marriage is larger for men with pre-marital children, and the increase in hours after marriage is larger for men whose children are all post-marital. Figures 9, 10 and 11 thus suggest that both marriage and children are positively associated with male labor supply, and that our model replicates much of this relationship.

6 Quantitative Results: Why Do Married Men Work More?

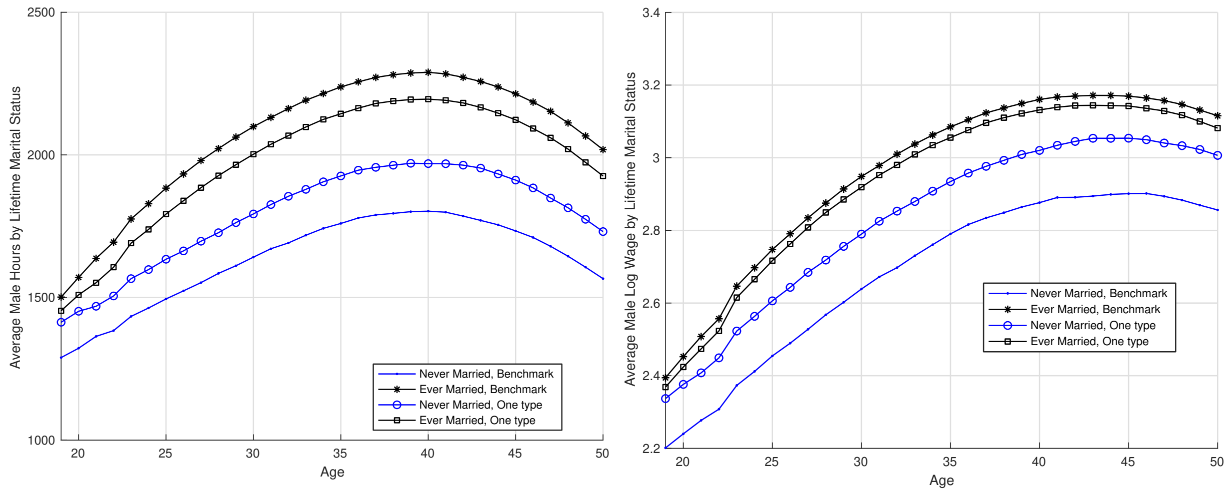
6.1 The Role of Permanent Selection

Our model contains two forms of marriage-related selection. The first type is permanent, type-related heterogeneity, where men more likely to marry face higher wages and a lower disutility from work. The second type is dynamic selection, where men with higher values of the AR(1) wage shock $\tilde{w}$ are more likely to marry.

Figure 12 isolates the effects of permanent selection. Figure 12a compares the hours profiles for ever- and never-married men generated by the baseline, two-type, model to the profiles gener-

Figure 12: Mean Hours and Wages by Age and Marital History: Baseline and One-type Model

(a) Hours worked (b) Log wages



Source: Author calculations; see text for details.

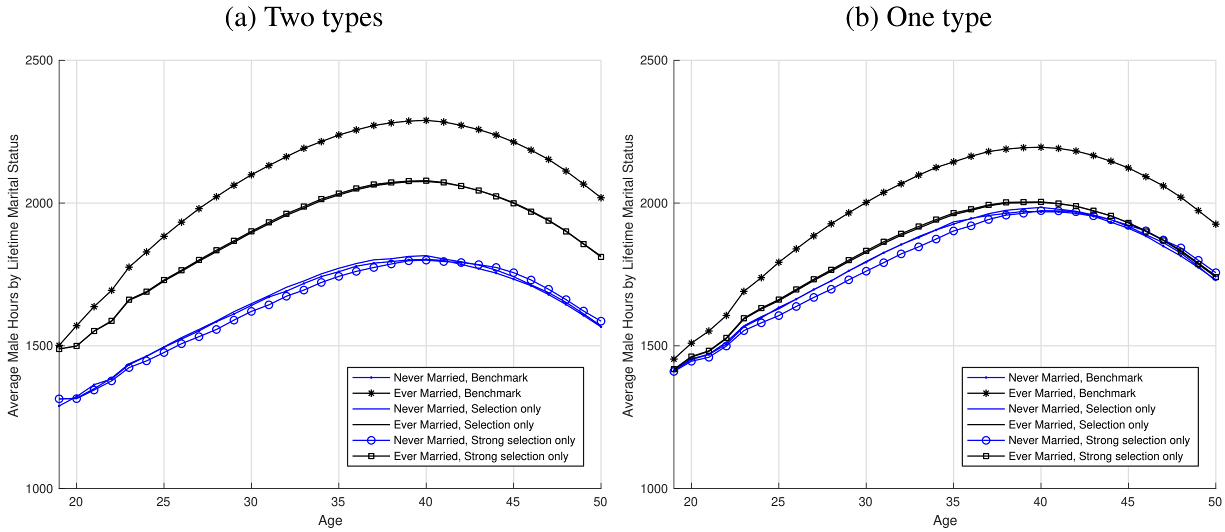
ated by a version of the model with no type differences. Comparing the two sets of model profiles gives us the effects of type differences on hours. At all ages, the hours gap more than doubles when type differences are included. The relative contribution of the permanent type differences is especially large at ages 25 and before, suggesting that other mechanisms – dynamic selection and the mouths-to-feed effect – become more important at older ages. This life-cycle pattern is consistent with the life-cycle incidence of marriage: because older men are more likely to be married, the mouths-to-feed effect will be stronger later in life. Even at earlier ages, however, the model without type differences still accounts for a non-trivial part of the hours gap.

Figure 12b presents the corresponding profiles for wages, which also reveal a large role for type differences.

6.2 The Role of Dynamic Selection

How much of the remaining gap is attributable to dynamic selection? In the model, men with higher AR(1) wage shocks move up the relationship ladder more quickly. This implies that married men have higher wages in part because of selection on $\tilde{w}$ and, because they respond to positive wage shocks by working more, they also have higher hours. To assess the quantitative impact of dynamic selection on labor supply, we run a specification where the causal effects of marriage on labor supply have been shut down, so that the only dynamic link between family structure and the labor market is dynamic selection. Specifically, in this alternative specification, spouses and

Figure 13: Mean Hours by Age and Marital History: Baseline and Selection-only Models



Source: In the selection-only model, spouses and children have no effect on husbands' consumption or the utility they receive from it. In the stronger selection-only model, the effect of the AR(1) shock $\tilde{w}$ on relationship transitions is over three times its baseline value. Panel (a) shows results for specifications with two permanent types. Panel (b) shows results for specifications with one permanent type. Author calculations; see text for details.

children have no effect on preferences or resources: $\eta_f = N_f = 1, \forall f$, and $se_j = \chi_a = \delta_a = 0$.[23] This leaves selection as the only dynamic link between family structure and the labor market. Married and engaged men work more because they receive better draws of $\tilde{w}$. To account for the possibility that our estimates understate the degree of wage selection in effect, we also consider a "stronger" selection-only alternative, where we increase substantially the parameter that governs dynamic selection (θ). As we detail below, this strengthens dynamic selection by a factor of more than three.

Figure 13a presents hours profiles for the full model and the two selection-only specifications. Removing the non-selection-related mechanisms shrinks the hours gap considerably, even in the stronger selection case, indicating that dynamic selection is not the most important dynamic mechanism. Figure 13b shows further that when permanent selection is also removed, very little of the hours gap remains. In short, permanent, type-related selection and non-selection dynamics account for most of the marriage-related hours gap; dynamic selection accounts for very little.

To gain further intuition, we examine the predictions of the same three specifications for

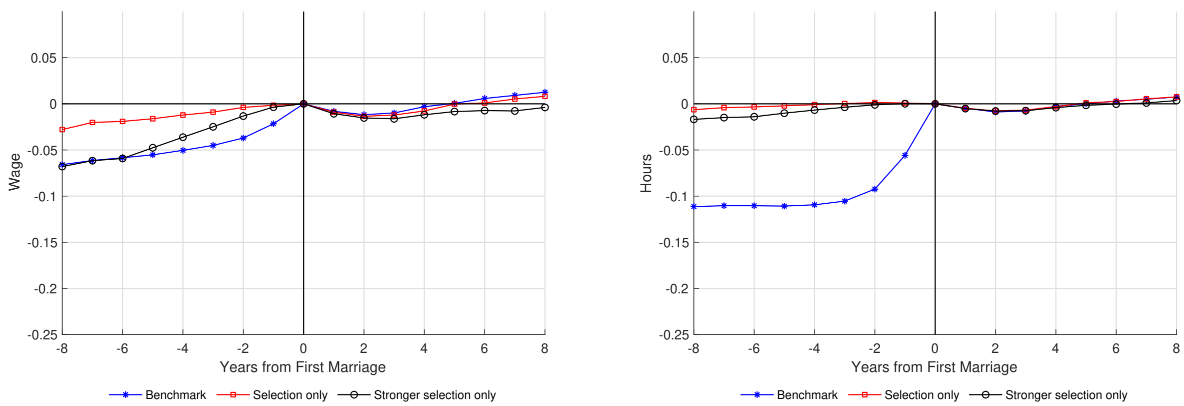
[23]In constructing this specification, we assume further that all men face the same tax schedule, using parameter estimates from Guner, Kaygusuz and Ventura (2014) that average across all unmarried men. We likewise assume that married men now keep all of their wealth upon divorce.

changes in wages and hours around the time of marriage.[24] Figure 14a shows that the pre-marital increase in wages in the selection-only specification (red line with squares) is much smaller than the one found in the baseline model (blue line with stars). This implies that the increase in wages in the benchmark model was not driven primarily by selection, but instead by the part-time wage penalty, where non-wage drivers pushed up wages by pushing up hours. These non-wage drivers are deactivated in the selection-only specification, where hours do not increase very much in the run-up to marriage (see Figure 14b).

Figure 14: Hours and Wages around Marriage: Benchmark and Selection-Only Models

(a) Hourly wages

(b) Annual hours worked



Source: Model results are authors' calculations. In the selection-only models, spouses and children have no effect on husbands' consumption or the utility they receive from it. In the stronger selection-only model, the effect of the AR(1) shock $\tilde{w}$ on relationship transitions is over three times its baseline value. See text for details.

Figure 14a also shows how we construct the "stronger" selection-only alternative. In particular, we manually increase the coefficient θ until the dynamic selection channel generates the full pre-marital increase in wages observed in the data. At this value of θ, a 10% increase in wages raises the probability of getting married by 16.7%, which is over three times the size of the baseline elasticity we estimate from the CPS.

Figure 14b compares the change in hours worked around marriage for the three specifications. Neither selection-only specification generates a substantial increase in hours prior to marriage. In the baseline selection-only model, hours increase less than 1% prior to marriage. In the stronger selection-only model, hours increase less than 2% prior to marriage, far below the increase seen in either the benchmark model or the data.[25]

[24] The simulated data used to generate Figure 14a includes type-related fixed effects but, as shown in Appendix Figure A12, shutting down the type differences has little effect on the artificial regression results. The dynamic effects of marriage operate more or less independently of the type differences. Given that $\tilde{w}$ is defined as a zero-mean deviation independent of type, this is not surprising.

[25] In Appendix Figure A12c, we show the results from a version of the model that is identical to the baseline except

It bears noting that in the data and baseline model, the increase in hours *prior* to marriage is larger than the increase in wages. If the increase in hours were solely a response to higher wages, the underlying uncompensated labor supply elasticity would have to be well in excess of 1. The uncompensated elasticity in our benchmark specification, however, is below 1. The benchmark Frisch elasticity is $\xi = 0.75$ (in the neighborhood of many empirical estimates), and the uncompensated elasticity is even smaller, especially when income effects are sizable, as is often the case for young men with little wealth.[26] One reason our results are consistent with large income effects is that the AR(1) shock $\tilde{w}$ is quite persistent, with an autocorrelation coefficient of around 0.9 (Table 2). This suggests that any wage shock with long-lived effects, such as a shock to wage growth, will generate modest hours responses. Intuitively, these persistent wage shocks are also the shocks most relevant for selection into marriage: because they have long-lived effects, they can have a significant effect on the perceived gains from a long-term commitment, whereas purely transitory shocks are unlikely to meaningfully alter marriage decisions.

Our finding of a modest wage elasticity is also consistent with the data shown in Figure 4, where hours remain more or less constant after marriage, even as wages continue to rise. The very fact that the relative magnitudes of the wage and hours changes are so different before and after the date of marriage argues against a simple selection story.

To summarize, the results shown in Figure 14 imply that wage selection plays a secondary role in generating the hours dynamics observed around the time of marriage. The degree of dynamic wage selection that we estimate is modest, and even if it were significantly larger, matching the observed hours dynamics would require large uncompensated wage elasticities. The larger hours increase associated with pre-marital children (see the discussion of Figure 10) also suggests that wage selection cannot be the sole explanation. If selection were the only mechanism, the increase in hours associated with pre-marital children, who are more likely to be unplanned and thus uncorrelated with labor market shocks, would be *smaller* than the increase associated with post-marital children, the opposite of what the data show.

6.3 The Labor Supply Effects of Marriage and Family Structure

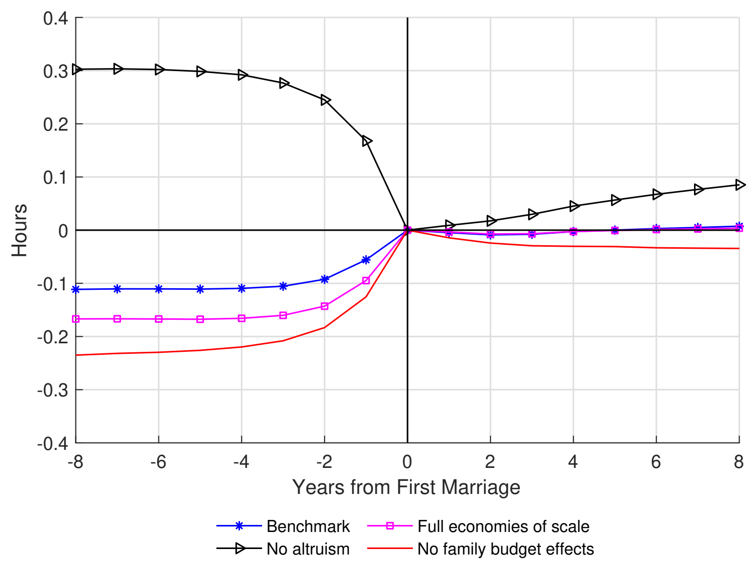
Marriage and family structure affect male labor supply through multiple channels. Having additional household members implies that a husband/father consumes only a portion of his house-

that θ is doubled. Relative to the baseline model, the "doubled selection" model generates an hours run up that is only modestly larger, again implying that dynamic selection plays a secondary role.

[26]To drill down on this point, it can be shown that with our specification, the Marshallian elasticity is given by $\frac{1-\gamma S}{1/\sigma+\gamma S}$, where S is the ratio of earnings to consumption (Keane 2011). Using our values of $\gamma = 0.738$ and $\sigma = 0.75$, and setting S to 0.75, we have an uncompensated elasticity of 0.24. (Setting S to 0.5 implies an elasticity of 0.37.)

hold's consumption expenditures (as $\eta_f > 1$). This effectively taxes the man's earnings, yielding the usual combination of income and substitution effects. Conversely, the utility shifter N_f implies that the husband receives utility from the consumption of other family members, which, all else equal, raises the shadow price of his earnings. Finally, other family members generate standard wealth effects: working spouses contribute earnings, while children impose child care costs if their mothers work. Together these family-related effects make up the mouths-to-feed effect.

Figure 15: Hours of Work by Distance to Marriage: Benchmark and Alternative Specifications



Source: Authors' calculations. "No altruism" shows the hours path for the specification where $N_f = 1, \forall f$. "Full economies of scale" shows the hours path for specification with $\eta_f = 1, \forall f$."No family budget effects" shows the hours path for the specification where spouses and children have no direct budgetary effects ($se(1 - n\chi_a) = 0$). See text for details.

To assess the quantitative importance of each of these mechanisms, Figure 15 presents the hours trajectories generated by three alternative specifications. In the first alternative (black line with triangles), we set $N_f = 1, \forall f$. This implies no altruism within a family: husbands/fathers do not receive utility from the consumption of other family members. Under this specification, marriage leads hours of work to *fall*. Recall that the flow utility from consumption equals $N_f \frac{1}{1-\gamma}(c/\eta_f)^{1-\gamma}$. If η_f is increasing in family size while N_f is not, spouses and children tax the man's earnings, and with $\gamma < 1$, hours will fall. This point is reinforced by the second alternative (pink line with squares), where we set $\eta_f = 1, \forall f$, resulting in "full economies of scale." With larger families imposing no consumption costs, but still generating altruism, the hours increase associated with marriage is even larger than in the baseline.

In the third alternative specification (solid red line), we assume that spouses have no earnings of their own and that children impose no direct child care costs, $se_j = \chi_a = 0$. This implies that families have no direct impact on the household budget, except through changes in the income tax function $T^{inc}(y)$. Under this specification, hours rise more than under the benchmark (starred blue line). For intuition, note that spouses always weakly expand the households' budget sets. Among working spouses ($se_j > 0$), the quantity $se(1-n\chi_a)$ is always positive: the cost of each child is at most $\chi_y = 28\%$ of their mother's earnings, and the largest possible number of children is $n = 3$. Meanwhile, non-working spouses have no income by construction and care for any children present. It follows that turning off the direct budget effects of families will on average tighten household budget constraints, which raises the marginal utility of consumption and induces husbands/fathers to work more.

The results from these alternative specifications imply that, through the lens of the model, men increase their labor supply around the time of marriage primarily due to altruism toward spouses and children (or its observational equivalent), and not primarily as a response to changes in family resources. The logic behind our finding is straightforward. The assumption that additional household members impose an earnings tax – η_f increases in household size – is an inherent feature of equivalence scales. With $\gamma < 1$, this tax discourages work. Moreover, as long as spouses cover the entire cost of child care, out of their earnings or through home production, married or cohabiting men enjoy higher household incomes. This too should discourage work. This leaves the size-dependent shifter N_f as the only feasible mechanism to generate a marriage-related increase in hours. Our study is not the first to employ this mechanism (see, e.g., Blundell et al. (2016) and Fan, Seshadri and Taber (2024)), but our results provide novel evidence in its support.

Taking Stock. The results presented in this section, along with those in Sections 5.2 and 6.2, lead us to three conclusions. First, the model generates a reasonable fit of male hours dynamics around the time of marriage. Second, selection into marriage on the basis of the AR(1) wage shocks cannot by itself explain the dynamic patterns observed in the data. Third, the mouths-to-feed mechanism appears to be an important driver of male labor supply. In particular, we find that the increase in male labor supply in the run-up to marriage is largely due to married men internalizing the consumption utility of their family.

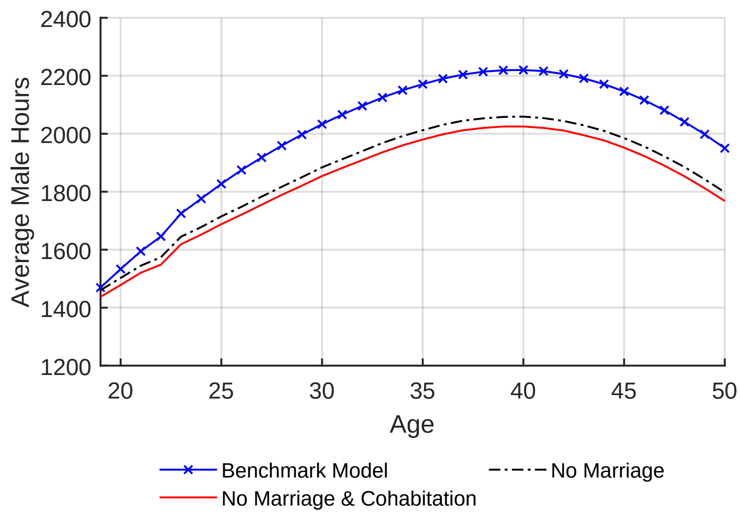
6.4 The Aggregate Impact of Marriage on Male Labor Supply

Our next set of numerical exercises uses the model to quantify the impact of marriage on aggregate labor supply.

6.4.1 Eliminating Marriage

We begin with an extreme experiment in which we eliminate marriage altogether. Specifically, we set the probability of engagement and marriage to zero, and we convert men who entered the baseline simulations as married or cohabiting into singles. Figure 16 compares the hours profile from this experiment ("No marriage & cohabitation") to that of baseline specification. Averaging across ages 19 - 50, eliminating marriage reduces male work by 168 hours per year, a 8.3% reduction. The aggregate effects are largest after age 35, when most men have married. For example, among 45-year-old men, eliminating marriage reduces hours worked by 194 hours, or 9.0%.

Figure 16: Hours of Work with and without Marriage



Source: Authors' calculations. In the "No Marriage" specification, we set the probability of marriage to zero. In the "No Marriage & Cohabitation" specification, we set the probability of engagement and marriage to zero. See text for details.

By way of comparison, in Siassi's (2019) framework, the feature most akin to our mouths-to-feed mechanism is a stronger bequest motive for individuals with children.[27] Removing this feature causes his estimate of the proportional earnings (not hours) gap to shrink by 2.3 - 2.8 percentage points (Siassi, 2019, Table 5), less than half the magnitude of our results.

In the model, cohabitation increases male labor supply in a manner similar to marriage. To understand the relative importance of cohabitation and marriage, we run a second experiment in which we eliminate marriage but not cohabitation. Specifically, we set the probability of marriage

[27] Because Siassi (2019) employs the flow utility function used in Greenwood, Hercowitz and Huffman (1988), which has no wealth effects, his framework is not directly comparable to ours.

to zero and convert men who began life married to singles, but we leave unchanged the probability of pre-marital cohabitation. Figure 16 also displays the hours profile from this experiment ("No marriage"). Averaging across ages 19 - 50, eliminating marriage but retaining cohabitation reduces male work by 138 hours, or 6.8%. This indicates that cohabitation has a sizable effect on aggregate hours worked ($8.3 - 6.8 = 1.5\%$), but that the effect of marriage is over four times as large.

6.4.2 Cross-cohort Comparisons

Next, we quantitatively assess the hypothesis by Binder and Bound (2019) and Binder (2021) that declining rates of marriage could help explain declines in male work rates in recent decades. To do this, we replace the baseline relationship processes with processes estimated from the NLSY97 (Bureau of Labor Statistics, 2019*b*), a longitudinal dataset comparable to the NSLY79 that follows a cohort born between 1980-1984, roughly two decades after the NLSY79 cohort, which was born between 1957-1964. Men in the NLSY97 cohort marry at lower rates than men in the NLSY79: for example, by age 25, only 27% of men in the NLSY97 have married, compared with 47% of men in the NLSY79.

Because the NLSY97 is based on a younger cohort than the NLSY79, the life-cycle moments we use to estimate the relationship process (see Section 4.3.3) are observed in the NLSY97 only through age 32. We therefore add to these moments imputed values for the share never-married between ages 33 - 50, assuming that the gap in marriage rates between the NLSY97 and NLSY79 remains constant from age 32 onward:

$$share_j^{97} = share_j^{79} + (share_{32}^{97} - share_{32}^{79}), \quad j \in \{33, ..., 50\}. \tag{28}$$

The imputation implies that, averaging across ages 19 - 50, 50.5% of men in the NLSY97 are never married, compared with 33.7% of men in the NLSY79. With the new moments thus assembled, we re-estimate the relationship process for the NLSY97 cohort.

The second row of Table 3 presents the hours associated with this experiment. Introducing the lower marriage rates from the NLSY97 reduces average hours of work for men ages 19 - 50 by 32 hours, or 1.59%. For context, recall from Figure 3 that between 1970 and 2018, hours of work by prime age men trended down by 7.7%. Background calculations show that the average year of birth in the NLSY79 is 1961.2, while the average year of birth in the NLSY97 is 1982, a difference of 20.8 years. This yields a 20.8-year trend decrease of $7.7\% \times \frac{20.8}{48} = 3.34\%$. Our results therefore suggest that recent declines in marriage rates, if exogenous, can account for a sizable share, 48%,

Table 3: Hours of Work with NLSY79 and NLSY97 Processes

Description	Hours	Change from Benchmark	
		Hours	Percentage
Benchmark model (NLSY79)	2,016.3		
NLSY97 marriage rates	1,984.2	-32.1	-1.59%
NLSY97 marriage rates + spousal earnings	1972.0	-44.3	-2.20%
NLSY97 marriage rates + spousal earnings + male wages	1952.8	-63.5	-3.15%

Note: Authors' calculations. Hours are annual averages for men ages 19 - 50. Spousal earnings in the NLSY97 are 13% larger than those in the NLSY79; male wages are 9% lower. See text for details.

of the decline in prime-age male hours over the past five decades. Applying this fraction to the total trend decline of 7.7% yields an hours decrease of 3.70%.[28] Notably, the model-predicted decline, while substantial, is smaller than the decline of 5.6% found in our motivating shift-share exercise (Section 2.2). While the shift-share exercise treats the correlation between marriage and hours as completely causal, the model allows for multiple forms of selection, the size of which we discipline with data.

The model experiment thus indicates that the decline in marriage rates can account for much of the decline in hours observed in the data, but not all. One potentially important mechanism is spousal earnings, which in our model are one component of the mouths-to-feed effect. Table 3 shows that imposing NLSY97 spousal earnings (13% higher) causes hours to fall further, to 2.20%. If we include the cross-cohort decline in male wages (9%), we have an overall decrease of 63 hours or 3.15%, equal to 94% of our CPS comparison value of 3.34%. This suggests that the decline in marriage rates and the changes in male and female earning power can together account for almost the entire decline in male hours since 1970.

7 Conclusion

Married men work substantially more hours than men who have never been married. Panel data reveal that although much of the gap can be attributed to fixed effects, a substantial portion reflects an increase in work around the time of marriage. Two potential explanations for this increase are: (i) men hit by positive labor market shocks are more likely to marry; and (ii) the prospect of

[28] By way of comparison, Binder (2021), using difference-in-differences estimators, finds that the combination of no-fault divorce and increased employment opportunities for women can account for 25% of the decline in labor force participation of young men without a college degree.

marriage increases mens' labor supply. We quantify the relative importance of these two channels using a structural life-cycle model of marriage and labor supply. The model allows for both permanent and dynamic selection into marriage and for marriage to directly affect the returns to work. A version of the model estimated to life-cycle marriage and fertility moments in the NLSY79 replicates the marriage-related hours dynamics observed in the same dataset.

The model provides a framework to evaluate the link between marriage and labor market outcomes. Through the lens of the model, dynamic selection into marriage driven by persistent wage shocks explains only a small part of the increase in hours around marriage. Instead, the primary driver of the increase in hours is that having a family increases the marginal utility men receive from household expenditures. Counterfactual experiments within the model show that declining marriage rates can account for half of the decline in prime age male hours since 1970. Moreover, once we expand the experiment to include cross-cohort changes in male wages and spousal earnings, our model generates most of the decline.

Our findings highlight several promising areas for future research. First, our analysis focuses on the impact of marriage on male labor market outcomes in the US. This raises the prospect that differences in marriage rates could explain cross-country differences in male labor market outcomes. While Bick and Fuchs-Schündeln (2018) study the labor market implications of cross-country differences in the taxation of married couples (as opposed to singles), to our knowledge, no one has considered marriage itself as a source of variation. Another important open question involves the joint determination of family structure (relationships and fertility) and labor supply (Gayle and Shephard, 2019; Caucutt, Guner and Rauh, 2021). Our analysis takes the life-cycle process for family structure as given; relaxing this assumption would add computational challenges to an already complex model.

References

Aaronson, Daniel, and Eric French. 2004. "The Effect of Part-time Work on Wages: Evidence from the Social Security Rules." *Journal of Labor Economics*, 22(2): 329–252.

Akerlof, George A. 1998. "Men without Children." *Economic Journal*, 108(447): 287–309.

Alon, Titan, Sena Ekin Coskun, and Matthias Doepke. 2018. "Trends in Work and Leisure: It's a Family Affair." Society for Economic Dynamics.

Antonovics, Kate, and Robert Town. 2004. "Are All the Good Men Married? Uncovering the Sources of the Marital Wage Premium." *American Economic Review*, 94(2): 317–321.

Attanasio, Orazio, Hamish Low, and Virginia Sánchez-Marcos. 2005. "Female Labor Supply as Insurance against Idiosyncratic Risk." *Journal of the European Economic Association*, 3(2-3): 755–764.

Autor, David, David Dorn, and Gordon Hanson. 2019. "When Work Disappears: Manufacturing Decline and the Falling Marriage-market Value of Young Men." *American Economic Review: Insights*, 1(2): 161–78.

Becker, Gary S. 1985. "Human Capital, Effort, and the Sexual Division of Labor." *Journal of Labor Economics*, 3(1, Part 2): S33–S58.

Becker, Gary S. 1988. "Family Economics and Macro Behavior." *American Economic Review*, 78(1): 1–13.

Becker, Gary S. 1991. *A Treatise on the Family.* Harvard University Press.

Bick, Alexander, Adam Blandin, and Richard Rogerson. 2022. "Hours and Wages." *Quarterly Journal of Economics*, 137(3): 1901–1962.

Bick, Alexander, and Nicola Fuchs-Schündeln. 2018. "Taxation and Labour Supply of Married Couples across Countries: A Macroeconomic Analysis." *Review of Economic Studies*, 85(3): 1543–1576.

Binder, Ariel. J. 2021. "The Effect of Declining Marriage Prospects on Young Men's Labor-Force Participation Rate." SSRN Working Paper 3795585.

Binder, Ariel J, and John Bound. 2019. "The Declining Labor Market Prospects of Less-Educated Men." *Journal of Economic Perspectives*, 33(2): 163–90.

Blundell, Richard, Luigi Pistaferri, and Itay Saporta-Eksten. 2016. "Consumption Inequality and Family Labor Supply." *American Economic Review*, 106(2): 387–435.

Blundell, Richard, Luigi Pistaferri, and Itay Saporta-Eksten. 2018. "Children, Time Allocation, and Consumption Insurance." *Journal of Political Economy*, 126(S1): S73–S115.

Blundell, Richard, Monica Costa Dias, Costas Meghir, and Jonathan Shaw. 2016. "Female Labor Supply, Human Capital, and Welfare Reform." *Econometrica*, 84(5): 1705–1753.

Borella, Margherita, Mariacristina De Nardi, and Fang Yang. 2023. "Are Marriage-Related Taxes and Social Security Benefits Holding Back Female Labour Supply?" *The Review of Economic Studies*, 90(1): 102–131.

Borra, Cristina, Martin Browning, and Almudena Sevilla. 2021. "Marriage and Housework." *Oxford Economic Papers*, 73(2): 479–508.

Bredemeier, Christian, Jan Gravert, and Falko Juessen. 2019. "Estimating Labor Supply Elasticities with Joint Borrowing Constraints of Couples." *Journal of Labor Economics*, 37(4): 1215–1265.

Budig, Michelle J, and Misun Lim. 2016. "Cohort Differences and the Marriage Premium: Emergence of Gender-Neutral Household Specialization Effects." *Journal of Marriage and Family*, 78(5): 1352–1370.

Bureau of Labor Statistics, U.S. Department of Labor. 2019*a*. "National Longitudinal Survey of Youth 1979 Cohort, 1979-2016 (rounds 1-27)." Produced and distributed by the Center for Human Resource Research, The Ohio State University, Columbus, OH.

Bureau of Labor Statistics, U.S. Department of Labor. 2019*b*. "National Longitudinal Survey of Youth 1997 Cohort, 1997-2017 (rounds 1-18)." Produced and distributed by the Center for Human Resource Research, The Ohio State University, Columbus, OH.

Caucutt, Elizabeth M, Nezih Guner, and Christopher Rauh. 2021. "Incarceration, Unemployment, and the Racial Marriage Divide."

Chiappori, Pierre-André, Monica Costa Dias, and Costas Meghir. 2018. "The Marriage Market, Labor Supply, and Education Choice." *Journal of Political Economy*, 126(S1): S26–S72.

Citro, Constance, and Robert Michael. 1995. "Adjusting Poverty Thresholds." *Measuring Poverty and New Approach, National Academic Press, Washington*.

Cornwell, Christopher, and Peter Rupert. 1997. "Unobservable Individual Effects, Marriage and the Earnings of Young Men." *Economic Inquiry*, 35(2): 285–294.

Cubeddu, Luis, and José-Víctor Ríos-Rull. 2003. "Families as Shocks." *Journal of the European Economic Association*, 1(2-3): 671–682.

de Linde Leonard, Megan, and Tom D Stanley. 2015. "Married with Children: What Remains when Observable Biases are Removed from the Reported Male Marriage Wage Premium." *Labour Economics*, 33: 72–80.

De Nardi, Mariacristina, Eric French, and John B Jones. 2010. "Why Do the Elderly Save? The Role of Medical Expenses." *Journal of Political Economy*, 118(1): 39–75.

Doepke, Matthias, and Michèle Tertilt. 2016. "Families in Macroeconomics." In *Handbook of Macroeconomics*. Vol. 2, 1789–1891. Elsevier.

Ellieroth, Kathrin. 2019. "Spousal Insurance, Precautionary Labor Supply, and the Business Cycle." *Unpublished Manuscript, Indiana University*.

Fan, Xiaodong, Ananth Seshadri, and Christopher Taber. 2024. "Estimation of a Life-Cycle Model with Human Capital, Labor Supply, and Retirement." *Journal of Political Economy*, 132(1): 48–95.

Flood, Sarah, Miriam King, Renae Rodgers, Steven Ruggles, and Robert J. Warren. 2020. "Current Population Survey." IPUMS, Minneapolis, MN. https://cps.ipums.org/cps/ (accessed May 2020).

French, Eric. 2005. "The Effects of Health, Wealth, and Wages on Labour Supply and Retirement Behaviour." *Review of Economic Studies*, 72(2): 395–427.

Gayle, George-Levi, and Andrew Shephard. 2019. "Optimal Taxation, Marriage, Home Production, and Family Labor Supply." *Econometrica*, 87(1): 291–326.

Ginther, Donna K, and Madeline Zavodny. 2001. “Is the Male Marriage Premium Due to Selection? The Effect of Shotgun Weddings on the Return to Marriage.” *Journal of Population Economics*, 14(2): 313–328.

Glauber, Rebecca. 2018. “Trends in the Motherhood Wage Penalty and Fatherhood Wage Premium for Low, Middle, and High Earners.” *Demography*, 55(5): 1663–1680.

Goldin, Claudia. 1992. *Understanding the Gender Gap: An Economic History of American Women.* Oxford University Press.

Goldin, Claudia. 2014. “A Grand Gender Convergence: Its Last Chapter.” *American Economic Review*, 104(4): 1091–1119.

Greenwood, Jeremy, Nezih Guner, and John A Knowles. 2003. “More on Marriage, Fertility, and the Distribution of Income.” *International Economic Review*, 44(3): 827–862.

Greenwood, Jeremy, Nezih Guner, Georgi Kocharkov, and Cezar Santos. 2016. “Technology and the Changing Family: A Unified Model of Marriage, Divorce, Educational Attainment, and Married Female Labor-force Participation.” *American Economic Journal: Macroeconomics*, 8(1): 1–41.

Greenwood, Jeremy, Zvi Hercowitz, and Gregory W Huffman. 1988. “Investment, Capacity Utilization, and the Real Business Cycle.” *American Economic Review*, 402–417.

Guner, Nezih, Remzi Kaygusuz, and Gustavo Ventura. 2012*a*. “Taxation and Household Labour Supply.” *Review of Economic Studies*, 79(3): 1113–1149.

Guner, Nezih, Remzi Kaygusuz, and Gustavo Ventura. 2012*b*. “Taxing Women: A Macroeconomic Analysis.” *Journal of Monetary Economics*, 59(1): 111–128.

Guner, Nezih, Remzi Kaygusuz, and Gustavo Ventura. 2014. “Income Taxation of U.S. Households: Facts and Parametric Estimates.” *Review of Economic Dynamics*, 17(4): 559 – 581.

Imai, Susumu, and Michael P Keane. 2004. “Intertemporal Labor Supply and Human Capital Accumulation.” *International Economic Review*, 45(2): 601–641.

Kaplan, Greg. 2012. “Inequality and the Life Cycle.” *Quantitative Economics*, 3(3): 471–525.

Keane, Michael, and Richard Rogerson. 2012. “Micro and Macro Labor Supply Elasticities: A Reassessment of Conventional Wisdom.” *Journal of Economic Literature*, 50(2): 464–76.

Keane, Michael P. 2011. “Labor Supply and Taxes: A Survey.” *Journal of Economic Literature*, 49(4): 961–1075.

Kearney, Melissa S, and Riley Wilson. 2018. “Male Earnings, Marriageable Men, and Nonmarital Fertility: Evidence from the Fracking Boom.” *Review of Economics and Statistics*, 100(4): 678–690.

Killewald, Alexandra, and Ian Lundberg. 2017. “New Evidence against a Causal Marriage Wage Premium.” *Demography*, 54(3): 1007–1028.

Knowles, John A. 2013. “Why are Married Men Working so Much? An Aggregate Analysis of Intra-household Bargaining and Labour Supply.” *Review of Economic Studies*, 80(3): 1055–1085.

Korenman, Sanders, and David Neumark. 1991. “Does Marriage Really Make Men More Productive?” *Journal of Human Resources*, 26(2): 282–307.

Low, Hamish W, Costas Meghir, Luigi Pistaferri, and Alessandra Voena. 2025. "Marriage, Labor Supply, and the Dynamics of the Social Safety Net." Federal Reserve Bank of Chicago. Working Paper 2025-27.

Lundberg, Shelly, and Elaina Rose. 2002. "The Effects of Sons and Daughters on Men's Labor Supply and Wages." *Review of Economics and Statistics*, 84(2): 251–268.

Mazzocco, Maurizio, Claudia Ruiz, and Shintaro Yamaguchi. 2014. "Labor supply, wealth dynamics and marriage decisions." Society for Economic Dynamics.

Pilossoph, Laura, and Shu Lin Wee. 2021. "Household Search and the Marital Wage Premium." *American Economic Journal: Macroeconomics*, 13(4): 55–109.

Purcell, Patrick J, Howard M Iams, and Dave Shoffner. 2015. "Education, Earning Inequality, and Future Social Security Benefits: A Microsimulation Analysis." *Soc. Sec. Bull.*, 75: 15.

Rodgers III, William M, and Leslie S Stratton. 2010. "Male Marital Wage Differentials: Training, Personal Characteristics, and Fixed Effects." *Economic Inquiry*, 48(3): 722–742.

Siassi, Nawid. 2019. "Inequality and the Marriage Gap." *Review of Economic Dynamics*, 31: 160–181.

Weil, David N, and Oded Galor. 1996. "The Gender Gap, Fertility, and Growth." *American Economic Review*, 86(3).

For Online Publication: Appendices

A Data Appendix

This appendix summarizes the data sources, sample construction, and key variable definitions used in the paper. Table A1 lists the four datasets and the role each plays.

Table A1: Data Sources

Dataset	Waves used	Role in paper	Where used
CPS ASEC	1970–2019	Cross-sectional hours gaps; aggregate trends; shift-share; state analyses	Figs. 1–3, Table A2
Linked monthly CPS	1982–2019	Effect of wages on marriage probability (IV)	Table A6
NLSY79	1979–2016 (rounds 1–27)	Event studies around marriage and first child; model calibration	Table 1, Secs. 2.3, 4–6
NLSY97	1997–2017 (rounds 1–18)	Cross-cohort marriage, wage, and spousal earnings processes	Table 3

Unless noted otherwise, all dollar amounts are deflated to constant 2016 dollars using the CPI-U.

A.1 CPS ASEC

Source. We use the cleaned version of the Annual Social and Economic Supplement (ASEC) of the Current Population Survey, from IPUMS-CPS (Flood et al., 2020).

Sample. Individuals ages 19–54 with valid marital status. We include individuals with zero annual hours of work.

Key variables.

- *Annual hours worked*: usual hours worked per week last year (`UHRSWORKLY`) multiplied by weeks worked last year (`WKSWORK1`), set to zero for individuals who did not work. Because `UHRSWORKLY` is unavailable before the 1976 survey, series that begin in 1970 use intervalled weeks (`WKSWORK2`, assigned interval midpoints) and full-time/part-time status (`FULLPART`, assigned 45 and 25 hours, respectively). After 1976, the mean of annual hours worked implied by these intervalled series closely approximate the mean of annual hours worked using `UHRSWORKLY` and `WKSWORK1`.

- *Marital status*: "married" combines married, spouse present and married, spouse absent; "never married" is never married/single; "separated" combines separated, divorced, and widowed.

- *Education*: less than a four-year college degree versus college degree or more, based on `EDUC`. The regressions underlying Figure 1 and Table A2 use a four-category classification (less than high school, high school, some college, bachelor's or more).

- *State analyses*: state-year cells are weighted averages over the individual-level sample; "share married" is the fraction currently married at the survey date.

A.2 Linked Monthly CPS

Source and linking. The marriage-transition regressions in Appendix G use the monthly CPS, also from IPUMS-CPS. The CPS rotation design interviews each household in months-in-sample 1–4, and again one year later in months-in-sample 5–8. We keep observations from months-in-sample 4 and 8, for which earnings are collected as part of the outgoing rotation group. We link individuals across the two interviews using the IPUMS person identifier `CPSIDP`.

Sample. Men who appear in both interviews, are unmarried in the first interview, reside in the same state in both, and report valid usual hours and weekly earnings.

Key variables.

- *Marriage transition*: an indicator equal to one if the man is never married in the first interview and married (or separated/widowed) in the second, and zero if he is never married in both.

- *Hourly wage*: weekly earnings divided by usual weekly hours. We drop wage observations below half the federal minimum wage or above $100. Wages and earnings enter the regressions through the inverse hyperbolic sine transformation, demeaned.

- *Instrument*: the state-year mean of the (transformed) individual wage.

A.3 NLSY79

Source. The National Longitudinal Survey of Youth 1979 (Bureau of Labor Statistics, 2019*a*), a panel of 12,686 individuals born 1957–1964, interviewed annually from 1979 to 1994 and biennially thereafter. We use rounds 1–27 (1979–2016).

Sample construction. Our "nearly-balanced" panel of men is built as follows:

1. Start with the 6,403 male respondents; drop the military oversample, leaving 5,579 men.

2. Drop men with missing race or education information, and a small number of men with internally inconsistent childbearing histories.

3. Keep men observed with valid labor market information at age 50 or older, and who were interviewed at least 20 times (of a possible 26) between 1979 and 2016.

This yields 2,731 men. Within this panel, we use person-year observations at ages 19 and above, and we exclude person-years in which the man is currently enrolled in school; as a result, men with a college degree effectively enter the sample at age 23 or older. In the event-study regressions (Equation 1) we additionally require that the man first married between ages 18 and 46 and keep observations within ten years of the first marriage.

Key variables.

- *Annual hours and weeks*: hours and weeks worked in the past calendar year, from the NLSY work-history variables. Annual hours are capped at 5,200; hours per week (annual hours divided by weeks worked) are capped at 100.

- *Annual earnings*: wage and salary income in the past calendar year (using the truncation-revised series where available), deflated by the CPI-U and capped at $200,000.

- *Hourly wage*: annual earnings divided by annual hours worked, set to missing for men with zero hours, wages above $100, or wages below one-half the (real) federal minimum wage.

- *Marriage*: age at first marriage is the year the first marriage began (from the marriage-history variables) minus birth year; distance from marriage is age minus age at first marriage. Current marital status comes from the survey-year marital status variable; in the small number of cases (805 of roughly 450,000 person-year observations) where a respondent reports "never married" after previously reporting a marriage, we recode the later observations as separated.

- *Cohabitation*: for men married at the time of interview, the 1990–2000 rounds ask whether the respondent cohabited with his spouse before marriage and, if so, when cohabitation began. We use these questions to compute the incidence (49%) and average duration (two years) of pre-marital cohabitation reported in Section 3.3.3.

- *Children*: the number of children ever born and each child's birth year come from the fertility history. We treat a first child as "pre-marital" if born in or before the year of first marriage. Reported child counts are made non-decreasing over time. Whether a pregnancy was planned is taken from the question asked in 1982 and biennially thereafter.

- *Spousal earnings, weeks, and hours*: from the household questions on the spouse's labor income and work time, used in the years the couple is married. Spousal earnings are deflated by the CPI-U; in the estimation of the spousal earnings process, working spouses are defined as those earning at least $\underline{s} = \$3{,}630$ per year (half the current federal minimum wage, 20 hours per week, 50 weeks per year), and earnings below this floor are censored at it.

- *Event-study normalization*: as described in Section 2.3, each labor market variable is normalized by the average among ever-married men of the same age, education group, and sex.

- *Child support*: incidence and amounts of child support paid by fathers with non-resident children, used to set the child support parameter δ_a (Appendix D).

A.4 NLSY97

The National Longitudinal Survey of Youth 1997 (Bureau of Labor Statistics, 2019*b*) follows a cohort born 1980–1984; we use rounds 1–18 (1997–2017). We process the NLSY97 with the same code and definitions as the NLSY79, adjusting only for differences in variable names and question design. Because the cohort is younger, the sample covers ages 19–32, and the nearly-balanced panel requires at least 15 interviews rather than 20. The NLSY97 is used only to re-estimate the relationship and fertility processes and to measure cross-cohort changes in male wages and spousal earnings for the cross-cohort counterfactuals in Section 6; marriage shares at ages 33–50 are imputed by assuming the NLSY97–NLSY79 gap in the never-married share remains constant from age 32 onward.

B Supplemental Estimates of the Relationship between Marriage and Male Labor Market Outcomes

B.1 State Average Hours Worked

Table A2: Predictors of State Average Male Annual Hours Worked in the CPS

	(1)	(2)	(3)
Constant	1450.0***	1443.3***	1537.9***
	(36.6)	(56.0)	(74.5)
Share Married	706.2***	768.7***	609.3***
	(62.1)	(84.3)	(123.1)
Average Age	–	0.4	0.0
		(0.5)	(0.3)
Share Less than High School	–	0.3	−5.4
		(14.1)	(8.1)
Some College Share	–	11.1	4.6
		(11.7)	(6.8)
Bahelor's + Share	–	6.0	3.7
		(11.7)	(6.8)
Share Black	–	−23.1	4.2
		(15.1)	(9.1)
Year FEs		*Y*	*Y*
State FEs			*Y*
R^2-adj	0.22	0.42	0.83
N	459	459	459

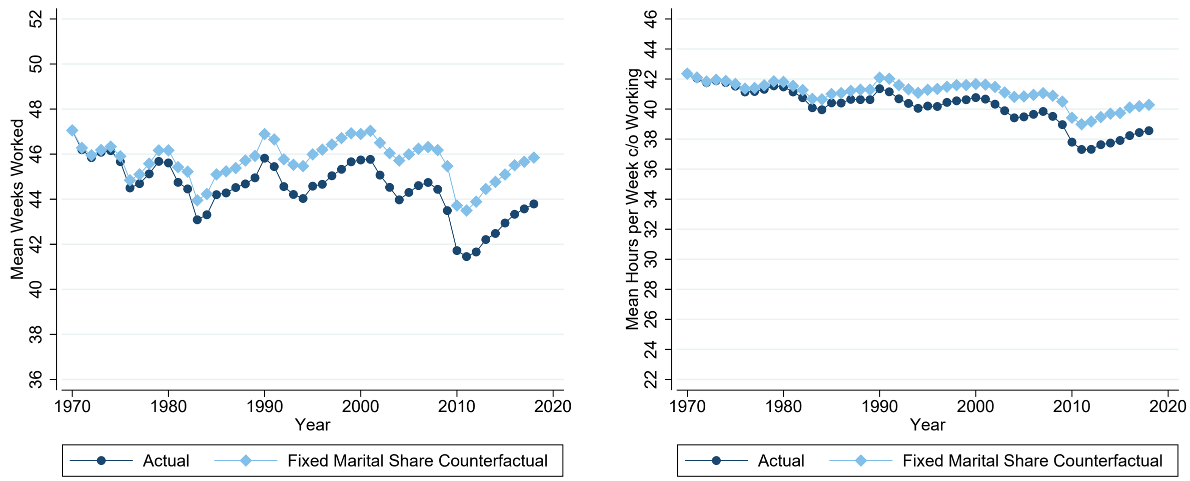
Source: Men ages 19 - 54 in the 1975-2019 waves of the CPS ASEC born between 1957-1964 (to correspond with the NLSY79 birth cohorts). The sample includes individuals with zero annual hours. The unit of observation is a 5-year average of average hours for a given state. There are nine five-year time periods in our sample: 1975-1979, 1980-1984, ..., 2015-2019; the data include all 50 states and the District of Columbia; together this implies 459 observations.

B.2 Workweeks and Hours per Workweek by Marital Status in the CPS

Figure A1 shows that the secular decline in annual hours reflects declines along both the extensive (number of workweeks) and intensive (weekly hours when working) margins. The total effect of these declines is partially offset by a secular increase in the covariance between weeks and hours. (Recall that $E(\text{weeks} \times \text{hours}) = E(\text{weeks})E(\text{hours}) + cov(\text{weeks}, \text{hours})$.)

Figure A1: Weeks of Work and Weekly Hours (for workers) by Marital Status: 1970-2018

(a) Weeks of work (b) Hours per workweek

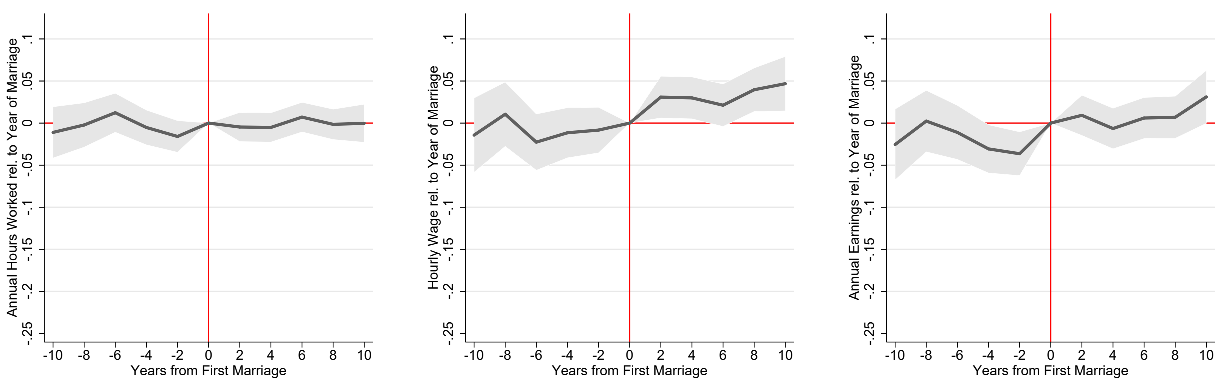


Source: Men ages 19 - 54 in the 1970-2019 waves of the CPS ASEC. The sample includes individuals with zero annual hours.

B.3 Placebo Test for Distance-from-Marriage Regressions

Figure A2: Placebo Test: Labor Market Dynamics in the Years around Marriage

(a) Annual hours worked (b) Hourly wages (c) Annual earnings

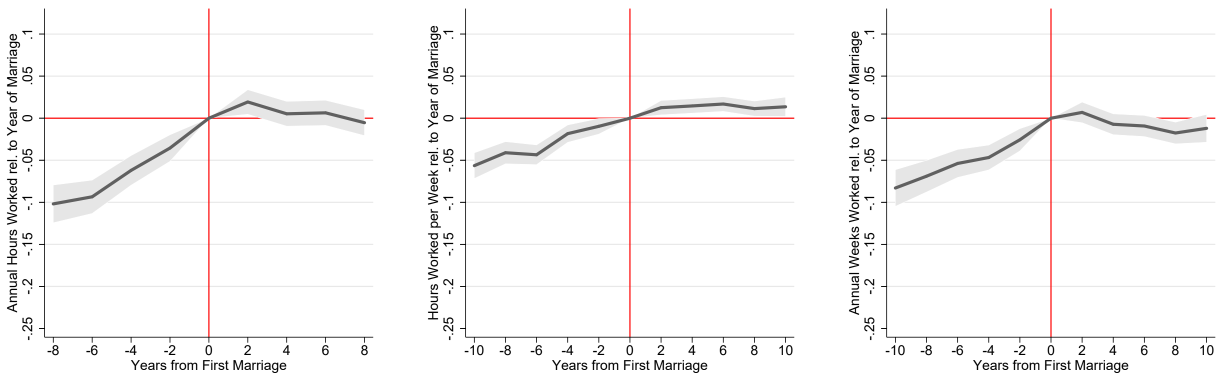


Source: Males ages 19 - 50 in the NLSY79, see text for details. The solid line plots distance-from-marriage coefficients from the individual fixed effects regression equation (1). The shaded region corresponds to 95% confidence intervals. The results in this table correspond to a placebo test of the results in Figure 4, in which age of first marriage was randomly reassigned among individuals in the regression sample.

B.4 Decomposing the Distance-from-Marriage Effects on Hours

Figure A3: Labor Market Dynamics in the Years around Marriage

(a) Annual hours worked (b) Hours per workweek (c) Annual weeks worked

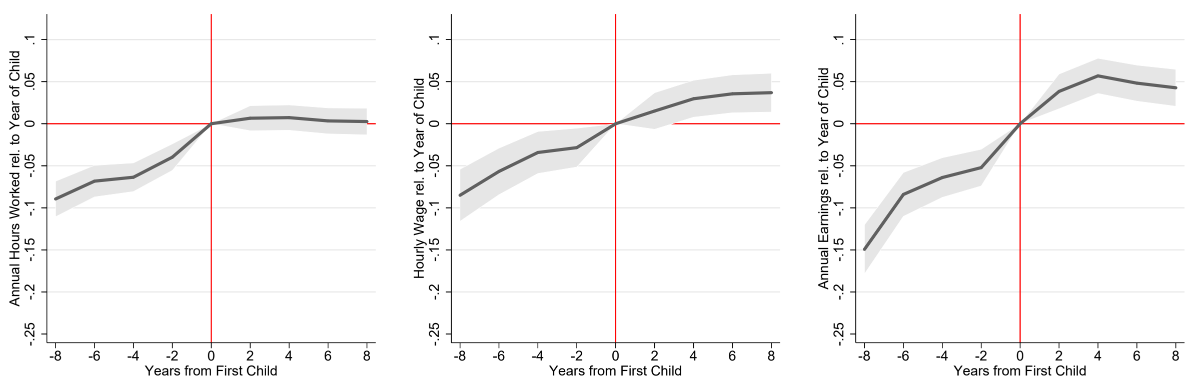


Source: Males ages 19 - 54 in the NLSY79, see text for details. The solid line plots distance-from-marriage coefficients from the individual fixed effects regression equation (1). The shaded region corresponds to 95% confidence intervals.

B.5 Decomposing the Distance-from-Child Effects on Hours

Figure A4: Labor Market Dynamics in the Years around First Child

(a) Annual hours worked (b) Hourly wages (c) Annual earnings



Source: Males ages 19 - 54 in the NLSY79; see text for details. The solid line plots distance-from-first-child coefficients from the individual fixed effects regression, by regressing on a sequence of dummies corresponding to the distance in years from the man's first child. The shaded region corresponds to 95% confidence intervals.

C Additional Model Details

C.1 Fertility Dynamics

Let $\phi^{an}_{a,n,r,e,j}(a',n')$ denote the probability that a man of age j with education e, relationship status r, and current child status (a,n) will have child status (a',n') next period. Collectively, our assumptions imply that:

$$\phi^{an}_{0,0,r,e,j}(yc,1) = \varphi^{n}_{0,r,e,j}, \tag{A1a}$$

$$\phi^{an}_{0,0,r,e,j}(0,0) = 1 - \varphi^{n}_{0,r,e,j}, \tag{A1b}$$

$$\phi^{an}_{yc,n,r,e,j}(oc,n+1) = \varphi^{a}_{yc,e,j} \cdot \varphi^{n}_{n,r,e,j}, \tag{A1c}$$

$$\phi^{an}_{yc,n,r,e,j}(yc,n+1) = (1 - \varphi^{a}_{yc,e,j}) \cdot \varphi^{n}_{n,r,e,j}, \tag{A1d}$$

$$\phi^{an}_{yc,n,r,e,j}(oc,n) = \varphi^{a}_{yc,e,j} \cdot (1 - \varphi^{n}_{n,r,e,j}), \tag{A1e}$$

$$\phi^{an}_{yc,n,r,e,j}(yc,n) = (1 - \varphi^{a}_{yc,e,j}) \cdot (1 - \varphi^{n}_{n,r,e,j}), \tag{A1f}$$

$$\phi^{an}_{oc,n,r,e,j}(gc,n) = \varphi^{a}_{oc,e,j}, \tag{A1g}$$

$$\phi^{an}_{oc,n,r,e,j}(oc,n) = 1 - \varphi^{a}_{oc,e,j}, \tag{A1h}$$

$$\phi^{an}_{gc,n,r,e,j}(gc,n) = 1. \tag{A1i}$$

C.2 Relationship Dynamics

Our structure gives rise to the following transition probabilities for single men:

$$\Pr_j(r_j = en \mid r_{j-1} = sn,\ n_j = 0,\ \tilde{w},\ e,\ \ell) = \phi^{en}_{e,\ell,j}(\tilde{w}), \tag{A2a}$$

$$\Pr_j(r_j = sn \mid r_{j-1} = sn,\ n_j = 0,\ \tilde{w},\ e,\ \ell) = 1 - \phi^{en}_{e,\ell,j}(\tilde{w}), \tag{A2b}$$

$$\Pr_j(r_j = en \mid r_{j-1} = sn,\ n_j = 1,\ n_{j-1} = 0,\ \tilde{w},\ e,\ \ell) = \phi^{owb}_{e} + \left(1 - \phi^{owb}_{e}\right)\phi^{en}_{e,\ell,j}(\tilde{w}), \tag{A2c}$$

$$\Pr_j(r_j = sn \mid r_{j-1} = sn,\ n_j = 1,\ n_{j-1} = 0,\ \tilde{w},\ e,\ \ell) = \left(1 - \phi^{owb}_{e}\right)\left(1 - \phi^{en}_{e,\ell,j}(\tilde{w})\right), \tag{A2d}$$

$$\Pr_j(r_j = en \mid r_{j-1} = sn,\ n_{j-1} > 0,\ \tilde{w},\ e,\ \ell) = \phi^{en}_{e,\ell,j}(\tilde{w}), \tag{A2e}$$

$$\Pr_j(r_j = sn \mid r_{j-1} = sn,\ n_j > 0\ \tilde{w},\ e,\ \ell) = 1 - \phi^{en}_{e,\ell,j}(\tilde{w}), \tag{A2f}$$

and the following probabilities for married and divorced men:

$$\Pr_j(r_j = dv \mid r_{j-1} = mr,\ n,\ e) = \phi^{dv}_{n,e,j}, \tag{A3a}$$

$$\Pr_j(r_j = mr \mid r_{j-1} = mr,\ n,\ e) = 1 - \phi^{dv}_{n,e,j}, \tag{A3b}$$

$$\Pr_j(r_j = dv \mid r_{j-1} = dv,\ n,\ e) = 1. \tag{A3c}$$

The transition probabilities of men who are currently engaged take the same form as those of single men, with the term $\phi^{en}_{e,j}(\tilde{w})$ replaced by $\phi^{mr}(\tilde{w})$. The probability of getting married conditional on being engaged, $\phi^{mr}(\tilde{w})$, does not vary with age or type: the age pattern of marriage depends solely on life-cycle changes in the rate of engagements and out-of-wedlock births. This is a concession to our data, which does not report engagement – we observe only whether an individual is married or has children.

C.3 Bellman Equations

The state vector for an age-j man consists of the man's education level (e), type (ℓ), assets (k_j), wage deviation ($\tilde{w}_j$), relationship status (r_j), age and number of children (a_j, n_j) and, for men who are engaged or married, spousal earnings shocks ($\tilde{s}$ and q_j). We will continue to use $f = (r,a,n)$ as a compact index of family structure. In contrast to the main text, in the Bellman equations that follow we will write out the integrals explicitly, to present the model's stochastic structure in detail.

C.3.1 Single

The Bellman equation for a working-age ($j < J_R$) single man, $r = sn$, is

$$\begin{aligned}
V^{sn}_j(e,\ell,k,\tilde{w},a,n) = \max_{c,k',h} \Bigg\{ & u_{f,e,\ell,j}(c,h) \; + \; \beta \int_{\tilde{w}'} \Bigg(\sum_{(a',n')} \phi^{an}_{a,n,sn,e,j}(a',n') \qquad \text{(A4)} \\
& \times \Big[\Pr_j\left(sn \,|\, sn,n',n,\tilde{w}',e,\ell\right) V^{sn}_{j+1}\left(e,\ell,k',\tilde{w}',a',n'\right) + \Pr_j\left(en \,|\, sn,n',n,\tilde{w}',e,\ell\right) \\
& \times \int_{\tilde{s}} \int_{q'} V^{en}_{j+1}\left(e,\ell,k',\tilde{w}',a',n',\tilde{s},q'\right) dF^q(q')\, dF^s_e(\tilde{s}\,|\,\tilde{w}') \Big] \Bigg) \, dF^w_{e,j}(\tilde{w}' \,|\, \tilde{w}) \Bigg\},
\end{aligned}$$

$$s.t. \quad c + k' \le Rk + wh^{1+\zeta}(1 - n\delta_a) - T(wh^{1+\zeta}, k; f), \tag{A5}$$

$$\log w = \alpha^w_{e,\ell,j} + \tilde{w}, \tag{A6}$$

$$k' \ge k_{min}. \tag{A7}$$

The cumulative distribution function $F^w_{e,j}(\tilde{w}' \,|\, \tilde{w})$, which describes the distribution of next period's idiosyncratic wage shock, $\tilde{w}_{j+1}$, given the current shock $\tilde{w}_j$, is based on Equations (4)–(6). The cumulative distribution functions related to spousal earnings, $F^s_e(\tilde{s}|\tilde{w})$ and $F^q(q)$, follow from equations (7) and (10), respectively.

C.3.2 Engaged

The continuation value conditional on becoming engaged is itself an expectation over the value of engagement with and without cohabitation: $V^{en} = \phi^{en\text{-}c} \times V^{en\text{-}c} + (1-\phi^{en\text{-}c}) \times V^{en\text{-}n}$. The Bellman equation for a working-age man who is engaged and not cohabiting is

$$
\begin{aligned}
V_j^{en\text{-}n}(e,\ell,k,\tilde{w},a,n,\tilde{s},q) = \max_{c,k',h} \Bigg\{ & u_{f,e,\ell,j}(c,h) \;+\; \beta \int_{\tilde{w}'} \int_{q'} \Bigg(\sum_{(a',n')} \phi^{an}_{a,n,en,e,j}(a',n') \\
& \times \Big[\Pr\big(en\text{-}n \,|\, en\text{-}n, n', n, \tilde{w}', e\big)\, V_{j+1}^{en\text{-}n}\big(e,\ell,k',\tilde{w}',a',n',\tilde{s},q'\big) \\
& + \Pr\big(mr \,|\, en\text{-}n, n', n, \tilde{w}', e\big)\, V_{j+1}^{mr}\big(e,\ell,k',\tilde{w}',a',n',\tilde{s},q'\big) \Big] \Bigg) \, dF^q(q')\, dF^w_{e,j}(\tilde{w}' \,|\, \tilde{w}) \Bigg\} \qquad \text{(A8)} \\
s.t. \quad & \text{equations (A5)-(A7).}
\end{aligned}
$$

The Bellman equation for a working-age man who is engaged and cohabiting, $r = en\text{-}c$, is somewhat different: (i) the flow utility function accounts for the co-residing partner and children (through η_f and N_f); (ii) the budget constraint includes spousal earnings and the taxes paid on these earnings; (iii) with children now residing with their father, child support costs proportional to the man's earnings are replaced with child care costs proportional to the earnings of the spouse. This yields

$$
\begin{aligned}
V_j^{en\text{-}c}(e,\ell,k,\tilde{w},a,n,\tilde{s},q) = \max_{c,k',h} \Bigg\{ & u_{f,e,\ell,j}(c,h) \;+\; \beta \int_{\tilde{w}'} \int_{q'} \Bigg(\sum_{(a',n')} \phi^{an}_{a,n,en,e,j}(a',n') \\
& \times \Big[\Pr\big(en\text{-}c \,|\, en\text{-}c, n', n, \tilde{w}', e\big)\, V_{j+1}^{en\text{-}c}\big(e,\ell,k',\tilde{w}',a',n',\tilde{s},q'\big) \\
& + \Pr\big(mr \,|\, en\text{-}c, n', n, \tilde{w}', e\big)\, V_{j+1}^{mr}\big(e,\ell,k',\tilde{w}',a',n',\tilde{s},q'\big) \Big] \Bigg) \, dF^q(q')\, dF^w_{e,j}(\tilde{w}' \,|\, \tilde{w}) \Bigg\}, \qquad \text{(A9)} \\
s.t. \quad & c + k' \le Rk + wh^{1+\zeta} + se(1-n\chi_a) - T(wh^{1+\zeta}, k; f^{en\text{-}c}) - T(se, 0; f^{en\text{-}c,sp}), \qquad \text{(A10)} \\
& se \text{ satisfies equations (8)-(11),} \qquad \text{(A11)} \\
& \text{equations (A6)-(A7).}
\end{aligned}
$$

Cohabiting couples pool their labor income together, but each partner files taxes as an individual. We capture this by replacing f with $f^{en\text{-}c}$ and $f^{en\text{-}c,sp}$ in the tax functions. (We assume that the man, who is most likely to have the higher income, will be the partner who claims the children as tax dependents. This means that $f^{en\text{-}c} = (sn,n,a)$ and $f^{en\text{-}c,sp} = (sn,0,0)$.) Consistent with our assumption that spouses bring no wealth into the relationship, we assign all of the cohabiting couple's asset income to the man.

C.3.3 Married

The Bellman equation for a working-age married man, $r = mr$, is

$$V_j^{mr}(e,\ell,k,\tilde{w},a,n,\tilde{s},q) = \max_{c,k',h} \Bigg\{ u_{f,e,\ell,j}(c,h) \; + \; \beta \int_{\tilde{w}'} \Bigg(\sum_{(a',n')} \phi^{an}_{a,n,mr,e,j}(a',n') \tag{A12}$$

$$\times \Bigg[\phi^{dv}_{n,e,j} V^{dv}_{j+1}\Big(e,\ell,\frac{1}{2}k',\tilde{w}',a',n'\Big)$$

$$+ \; (1-\phi^{dv}_{n,e,j}) \int_{q'} V^{mr}_{j+1}\big(e,\ell,k',\tilde{w}',a',n',\tilde{s},q'\big)\, dF^q(q') \Bigg] \Bigg) \, dF^w_{e,j}(\tilde{w}' \,|\, \tilde{w}) \Bigg\},$$

$$s.t. \quad c + k' \leq Rk + wh^{1+\zeta} + se(1-n\chi_a) - T(wh^{1+\zeta} + se, k; f), \tag{A13}$$

equations (A6)-(A7) and (A11).

With probability $\phi^{dv}_{n,e,j}$ the man divorces, keeping half of the household assets. With probability $(1-\phi^{dv}_{n,e,j})$ the man remains married. Married couples file taxes jointly.

C.3.4 Divorced

The Bellman equation for a working-age divorced man, $r = dv$, is

$$V_j^{dv}(e,\ell,k,\tilde{w},a,n) = \max_{c,k',h} \Bigg\{ u_{f,e,\ell,j}(c,h) \tag{A14}$$

$$+ \beta \int_{\tilde{w}'} \Bigg(\sum_{(a')} \varphi^{a}_{a,e,j}(a') V^{dv}_{j+1}\big(e,\ell,k',\tilde{w}',a',n\big) \Bigg) dF^w_{e,j}(\tilde{w}' \,|\, \tilde{w}) \Bigg\},$$

$$s.t. \quad \text{equations (A5)-(A7).}$$

Divorced men face the same budget constraints as singles. Their Bellman equation is much simpler, however, due to our assumptions that: (i) divorced men never remarry; and (ii) divorced men have no additional children.

C.3.5 Retired

Retired men ($j \geq J_R$) do not work, and if they are married or cohabiting, their spouses do not work. Their only income comes from their assets and from Social Security benefits. In retirement, all children are grown and relationships do not change in retirement, eliminating any uncertainty due to either factor. The resulting Bellman equation is completely deterministic:

$$V_j^{ret}(e,k,r) = \max_{c,k'} \quad u_{f,e,j}(c,0) + \beta V_{j+1}^{ret}(e,k',r), \tag{A15}$$

$$s.t. \quad c + k' \leq Rk + b_{1,e} + b_{2,e}\mathbb{1}_{r\in\{mr,en\text{-}c\}} - T(b_{1,e} + b_{2,e}\mathbb{1}_{r\in\{mr,en\text{-}c\}}, k; f), \tag{A16}$$

equation (A7),

$$V_J \equiv 0. \tag{A17}$$

Given that hours are identically zero and relationships are static, there are no type-related differences among retirees.

D Parameters Taken from Other Studies

Table A3 displays the values for parameters taken from other studies. The first panel of the table shows that non-college and college men enter the model one year after their modal graduation ages, 19 and 23, respectively. They retire at age $J_R = 65$ and die at age $J = 80$.

In the main text we discuss how set three key parameters: the consumption utility curvature parameter, γ, which we set to 0.738; the Frisch elasticity, ξ, which we set to 0.75; and the consumption utility scaler, N_f, which we set to 2 for married and cohabiting men and 1 for the rest. We turn now to describing how we set the remainder of Table A3.

Our formulation of the equivalence scale η_f comes from Citro and Michael (1995):

$$\eta_f = \begin{cases} 1, & \text{if } r \notin \{mr, en\text{-}c\}, \\ (2 + \mathbb{1}_{a<gc} \cdot 0.7n)^{0.7}, & \text{if } r \in \{mr, en\text{-}c\}. \end{cases} \tag{A18}$$

Grown children, who are assumed to live outside the household, do not enter the formula.

We set the gross interest rate R to 1.02, a standard value. Spouses who work surrender a fraction of their earnings to pay for formal child care. Using Borella, De Nardi and Yang's (2023) estimates, we set χ_a to 28% and 7% of her earnings, per child, for young and old children, respectively. Men who are neither married nor cohabiting pay child support costs equal to $\delta_a = 1.7\%$ of their earnings for each non-grown child. We take this number from the NLSY79.[29]

[29] The NLSY79 data show that 28.73% of men with non-resident children pay child support, which as a fraction of earnings has a median value of 9.4%. These data also reveal that men with non-resident children have an average of 1.6 such children. Dividing the product of the first two numbers by the third gives us $(0.2873 \cdot 0.094)/1.6 = 0.017$.

Table A3: Parameters Taken from Other Studies

Description	Parameter	Value	Target/Source
Demographics			
Starting age, non-college	J_{nc}	19	
Starting age, college	J_c	23	
Retirement age	J_R	65	
Terminal age	J	80	
Preferences			
Coefficient of RRA	γ	0.738	Imai and Keane (2004)
Frisch elasticity	ξ	0.75	Various (see Section 4.1)
Consumption utility shifter, married or cohabiting	$N_{f:r\in\{mr,en\text{-}c\}}$	2	See Section 4.1
Consumption utility shifter not married or cohabiting	$N_{f:r\notin\{mr,en\text{-}c\}}$	1	See Section 4.1
Equivalence scale	η_f	eqn. (A18)	Citro and Michael (1995)
Budget			
Interest rate	R	102%	
Child care cost per young child	χ_y	28%	Borella, De Nardi and Yang (2023)
Child care cost per older child	χ_o	7%	Borella, De Nardi and Yang (2023)
Child support cost per non-grown child, non-resident fathers	δ_a	1.7%	See appendix text
Part-time wage penalty	ζ	0.400	Aaronson and French (2004)
Borrowing limit	k_{min}	\$0	
Government			
Income Tax	$\tau_f^0, \tau_f^1, \tau_f^2$		Guner, Kaygusuz and Ventura (2014)
SS tax	τ^{ss}	5.2%	2013 value
SS benefit, man, non-college	$b_{1,nc}$	\$20,570	See appendix text
SS benefit, man, college	$b_{1,c}$	\$29,520	See appendix text
SS benefit, spouse, non-college	$b_{2,nc}$	\$15,620	See appendix text
SS benefit, spouse, college	$b_{2,c}$	\$21,810	See appendix text

Note: Child care and child support costs are expressed as fractions of earnings and are per child. Quantities are expressed in 2013 dollars.

We set ζ, the parameter governing the part-time wage penalty, to 0.4, following Aaronson and French (2004). At this value, a person working half-time suffers a 25% decrease in wages.

We set k_{min} to \$0, ruling out unsecured borrowing.

We set the income tax parameters to the values reported by Guner, Kaygusuz and Ventura (2014, tables 10 & 11), who estimate tax rates as a function of income, marital status and number of children, using administrative data. The Social Security tax rate τ^{ss} equals 5.2%, the value in effect in 2013. Our estimates of Social Security benefits, b_{1e} and b_{2e}, are based on microsimulation estimates from Purcell, Iams and Shoffner (2015, Table 3), adjusted for real wage growth.[30]

E Implementing Type-related Differences: Details

Each individual belongs to one of two unobserved types, indexed by $\ell \in \{1,2\}$. We will use $\ell = 1$ to index the "non-marrying" type and use p_1 to denote the probability that a man is of this type. It proves convenient to express the type-related differences as zero-mean deviations to (a subset of) the model's forcing processes and parameters. We estimate the type-related parameter vector using a grid search, targeting differences in wages and hours between never-married and ever-married men. As noted in the main text, however, type differences also affect the exogenous processes for wages, relationships, and spousal earnings. This requires us to estimate some of the parameters for the exogenous processes jointly with some of the type-related parameters and make ex-post adjustments to other type-related parameters.

E.1 Wages

We assume the type-specific mean $\widehat{\alpha}^w_{e,\ell,j}$ is given by

$$\alpha^w_{e,\ell,j} = \alpha^w_{e,j} + \tilde{\Delta}^w_i, \tag{A19}$$

$$\tilde{\Delta}^w_i = \begin{cases} -\Delta^w, & \ell_i = 1, \\ \frac{p_1}{1-p_1}\Delta^w, & \ell_i = 2. \end{cases} \tag{A20}$$

Our normalization implies that Δ^w is positive: type-1 men have lower average wages. For future reference, we note that the variance and autocovariances of the fixed effect $\tilde{\Delta}^w_i$ equal

$$\left(\sigma^{\Delta w}\right)^2 = \frac{p_1}{1-p_1}(\Delta^w)^2. \tag{A21}$$

[30] We use Purcell, Iams and Shoffner's (2015) alternative specification, which assumes that the real wages of non-college and college graduates grow at annual rates of 0.7% and 1.6%, respectively. Because these estimates are for people born between 1965 and 1979, on average 11.5 years younger than those in the NLSY, we deflate them by 11.5 years of real wage growth. (We also convert the numbers into 2013 dollars using the CPI.)

We now proceed in a way very similar to that used by French (2005). We begin by running an individual fixed effects regression of log wages (after adjusting for hours) on a quadratic in age and a control for the national unemployment rate during January of that calendar year. The estimated coefficients for the quadratic in age, along with the average fixed effect, give us predicted wages, $\widehat{\alpha}^w_{e,j}$. Subtracting $\widehat{\alpha}^w_{e,j}$ (along with an adjustment for aggregate unemployment) from observed wages produces a panel of estimated wage residuals, $\{\widehat{w}^*_{i,j}\}_{i,j}$. Because $\{\widehat{w}^*_{i,j}\}_{i,j}$ is constructed by subtracting $\widehat{\alpha}^w_{e,j}$, not $\widehat{\alpha}^w_{e,\ell,j}$, it includes $\tilde{\Delta}^w_i$.

Next, we assume that the stochastic process for the "true" wage residual, $w^*_{i,j}$, is the sum of the type effect $\tilde{\Delta}^w_i$, the AR(1) process $\tilde{w}_{i,j}$ and the white noise process $\bar{w}_{i,j}$:

$$w^*_{i,j} = \tilde{\Delta}^w_i + \tilde{w}_{i,j} + \bar{w}_{i,j}, \tag{A22a}$$

$$\tilde{w}_{i,j} = \rho^w_e \tilde{w}_{i,j-1} + \varepsilon^w_{i,j}, \tag{A22b}$$

$$\varepsilon^w_{i,j} \overset{iid}{\sim} N(0, \sigma^\varepsilon_e), \tag{A22c}$$

$$\tilde{w}_{i,0} \sim N(0, \sigma^{\tilde{w}_0}_e), \tag{A22d}$$

$$\tilde{w}_{i,j} \perp\!\!\!\perp \tilde{\Delta}^w_i, \ \forall i, j, \tag{A22e}$$

$$\bar{w}_{i,j} \overset{iid}{\sim} N(0, \sigma^{\bar{w}}_e), \tag{A22f}$$

$$\bar{w}_{i,j} \perp\!\!\!\perp \tilde{\Delta}^w_i, \tilde{w}_{i,j+s}, \ \forall i, j, s. \tag{A22g}$$

The variances and autocovariances of $w^*_{i,j}$ are thus

$$\mathrm{var}_e(w^*_{i,j}) = \left(\sigma^{\Delta w}\right)^2 + \mathrm{var}_e(\tilde{w}_{i,j}) + (\sigma^{\bar{w}}_e)^2, \tag{A23}$$

$$\mathrm{cov}_e(w^*_{i,j}, w^*_{i,j+k}) = \left(\sigma^{\Delta w}\right)^2 + (\rho^w_e)^k \mathrm{var}_e(\tilde{w}_{i,j}), \ k > 1. \tag{A24}$$

We can then use equation (A24) to back out the autocorrelation ρ^w_e:

$$\widehat{\rho}^w_e = \sqrt{\frac{\mathrm{cov}_e(\widehat{w}^*_{i,j}, \widehat{w}^*_{i,j+4}) - \left(\sigma^{\Delta w}\right)^2}{\mathrm{cov}_e(\widehat{w}^*_{i,j}, \widehat{w}^*_{i,j+2}) - \left(\sigma^{\Delta w}\right)^2}}, \tag{A25}$$

We use the second and fourth lags, and take the square root, because the NLYS79 is biennial in later years. Unless the process for $\tilde{w}_{i,j}$ is stationary, $\mathrm{cov}_e(w^*_{i,j}, w^*_{i,j+k})$ will differ by j as well as k. We thus calculate equation (A24) at each value of j and take averages. With $\widehat{\rho}$ in hand, we can estimate the variance of the AR(1) component at each age j as

$$\widehat{\mathrm{var}}_e(\tilde{w}_{i,j}) = \left[\mathrm{cov}_e(\widehat{w}^*_{i,j}, \widehat{w}^*_{i,j+2}) - \left(\sigma^{\Delta w}\right)^2\right] \frac{1}{(\widehat{\rho}^w_e)^2}, \tag{A26}$$

Under the error components model (A22), the variances of $\tilde{w}_{i,j}$, $\text{var}_e(\tilde{w}_{i,j}), j = 0,1,...,J$ are:

$$\text{var}_e(\tilde{w}_{i,0}) = (\sigma_e^{\tilde{w}_0})^2, \tag{A27a}$$

$$\text{var}_e(\tilde{w}_{i,1}) = (\rho_e^w)^2 \text{var}(\tilde{w}_{i,0}) + (\sigma_e^{\varepsilon})^2, \tag{A27b}$$

$$\text{var}_e(\tilde{w}_{i,j}) = (\rho_e^w)^2 \text{var}(\tilde{w}_{i,j-1}) + (\sigma_e^{\varepsilon})^2, \quad j > 1. \tag{A27c}$$

We can then estimate $\widehat{\sigma}_e^{\varepsilon}$ and $\widehat{\sigma}_e^{\tilde{w}_0}$ by fitting the sequence of variances implied by equation (A27) to the sequence implied by equation (A26), $\left\{\widehat{\text{var}}_e(\tilde{w}_{i,j})\right\}_{j=0}^{J}$.

In addition to affecting the wage process, Δ_w and p_1 will also affect the processes for relationships and spousal earnings, both through their dependence on the AR(1) wage shock $\tilde{w}_j$ and directly (for p_1). Fortunately, the other type-related effects can be modeled in a way that detaches them from the second-stage processes, leaving us with just p_1 and Δ^w. Therefore, for every (p_1, Δ^w) pair on our estimation grid, we re-estimate the second stage parameters for the exogenous processes. While doing this, we set the other type-related parameters equal to zero, giving us a two-dimensional grid over (p_1, Δ^w) that is orders of magnitude smaller than the full parameter grid.

E.2 Relationship transitions

The types also differ in the probability of remaining single. Recall that in the absence of a new child, the probability of engagement is $\phi_{e,j}^{en}$, so that the probability of remaining single is $\phi_{e,j}^{sn} = 1 - \phi_{e,j}^{en}$. To simplify the notation, we will ignore the dependence of $\phi_{e,j}^{sn}(\tilde{w})$ on p_1 and Δ^w (described immediately above); these parameters would otherwise appear as subscripts on ϕ^{sn}.

Type effects enter as

$$\phi_{e,\ell,j}^{sn}(\tilde{w}) = \begin{cases} \phi_{e,j}^{sn}(\tilde{w})\left(1 + \Delta^{sn}\right), & \ell = 1, \\ \phi_{e,j}^{sn}(\tilde{w})\left(1 - \frac{p_1}{1-p_1}\Delta^{sn}\right), & \ell = 2. \end{cases} \tag{A28}$$

Our prior belief is that Δ^{sn} will be positive: type-1 men are more likely to remain single.

The type effects in equation (A28) will, over the life cycle, cause the distribution of singles to shift toward type-1 individuals. This will in turn cause the aggregate transition probabilities to deviate from their one-type values. If left uncorrected, these discrepancies would require us to re-estimate the second stage parameters (male wages, fertility, relationships and spousal earnings) at each value of Δ^{sn}. Given that we must already re-estimate the second-stage parameters at each value of Δ^w and p_1, adding another dimension of re-estimation would raise major computational challenges. We instead introduce a bias correction that preserves the relationship patterns generated by the second stage relationship parameters, and thus continues to match the moments targeted in the second stage.

To see how the correction works, we must first introduce some additional notation. Let n^{-} denote the lagged number of children, and let $\pi_{e,n^{-},n,j}(\tilde{w})$, denote the probability that a single person remains single *after* accounting for out-of-wedlock births:

$$\pi_{e,n^{-},n,j}(\tilde{w}) = \begin{cases} (1-\phi_e^{owb})\phi_{e,j}^{sn}(\tilde{w}) = (1-\phi^{owb})(1-\phi_{e,j}^{en}(\tilde{w})), & \text{if } n^{-}=0 \text{ and } n>0, \\ \phi_{e,n^{-},n,j}^{sn}(\tilde{w}) = 1-\phi_{e,j}^{en}(\tilde{w}), & \text{otherwise,} \end{cases} \tag{A29}$$

Applying equation (A28) yields the type-specific probabilities

$$\pi^{*}_{e,n^{-},n,j,1}(\tilde{w}) = \pi_{e,n^{-},n,j}(\tilde{w})(1+\Delta^{sn}), \tag{A30}$$

$$\pi^{*}_{e,n^{-},n,j,2}(\tilde{w}) = \pi_{e,n^{-},n,j}(\tilde{w})\big(1-\tfrac{p_1}{1-p_1}\Delta^{sn}\big). \tag{A31}$$

For sufficiently large values of Δ^{sn}, $\pi^{*}_{e,n^{-},n,j,1}$ may exceed one or $\pi^{*}_{e,n^{-},n,j,2}$ may be less than zero. We therefore check that $\pi^{*}_{e,n^{-},n,j,\ell}(\tilde{w})$ lies in the $[0,1]$ interval and discard any values of Δ^{sn} that violate this constraint.

Turning to the adjustment itself, we will de-clutter the notation by suppressing the dependence of π on $\tilde{w}$, e, n^{-}, and n. Let $s_{j,\ell}$ denote the number of type-ℓ singles in period j, so that the total number of singles is $s_j = s_{j,1}+s_{j,2}$. Denote the share of type-1 singles in period j by $\lambda_j = s_{j,1}/s_j$. Using this notation, and equations (A30) and (A31), the number of singles in period j+1 implied by the *biased* transition probabilities $\pi^{*}_{j,1}(\tilde{w})$ and $\pi^{*}_{j,2}(\tilde{w})$ is

$$\begin{aligned} s^{*}_{j+1} &= s^{*}_{j+1,1}+s^{*}_{j+1,2} \\ &= s_{j,1}\pi_j(1+\Delta^{sn}_{j,})+s_{j,2}\pi_j\big(1-\tfrac{p_1}{1-p_1}\Delta^{sn}_{j}\big) \\ &= s_j\Big(\lambda_j\pi_j(1+\Delta^{sn}_{j})+(1-\lambda_j)\pi_j\big(1-\tfrac{p_1}{1-p_1}\Delta^{sn}_{j}\big)\Big) \\ &= s_j\pi_j\Big(1+\Delta^{sn}_{j,}\big[\lambda_j-(1-\lambda_j)\tfrac{p_1}{1-p_1}\big]\Big) \\ &:= s_j\pi_j\delta_j^{-1}. \end{aligned} \tag{A32}$$

The bias is $\delta_j^{-1}-1$: it is the extent to which the share of singles will be too large relative to the share implied by one-type transition probabilities. Note that $\delta_j = 1$ if $\lambda_j = p_1$ or $\Delta^{sn}_{j,ch} = 0$, as it should. Moreover, $\lambda_j > p_1 \Rightarrow \delta_j^{-1} > 1$: a preponderence of low types results in too many singles at period j+1.

To remove the bias, we need to replace π_j with $\pi_j \delta_j$. The share single (averaged across types) in every period will then equal that of the one-type case. We do this recursively.

1. Begin period j with $\lambda_j = \frac{s_{j,1}}{s_{j,1}+s_{j,2}}$, and the one-type transition probability π_j.

2. As described by equation (A28), find the type-specific transition probabilities $\pi^*_{j,1} = \pi_j\left(1+\Delta^{sn}_j\right)$ and $\pi^*_{j,2} = \pi_j\left(1-\frac{p_1}{1-p_1}\Delta^{sn}_j\right)$.

3. Find the adjustment
$$\delta_j = \left(1+\Delta^{sn}_j\left[\lambda_j-\left(1-\lambda_j\right)\tfrac{p_1}{1-p_1}\right]\right)^{-1}. \tag{A33}$$

4. Correct the type-specific probabilities found in step 2 by replacing $\pi^*_{j,\ell}$ with $\pi^{**}_{j,\ell} = \pi^*_{j,\ell}\delta_j$.

5. Move to period j+1 with
$$\lambda_{j+1} = \frac{\pi^{**}_{j,1}\lambda_j}{\pi^{**}_{j,1}\lambda_j+\pi^{**}_{j,2}\left(1-\lambda_j\right)}. \tag{A34}$$
This uses $s_{j+1,1} = s_{j,1}\pi^{**}_{j,1}$ and $s_{j+1,2} = s_{j,2}\pi^{**}_{j,2}$.

6. The sequence begins with $s_{0,1}$ and $s_{0,2}$ from the initial distribution of states, which we discuss immediately below.

It bears noting that even when $\pi^*_{e,n^-,n,j,\ell}(\tilde{w})$ lies between 0 and 1, $\pi^{**}_{e,n^-,n,j,\ell}(\tilde{w})$ may not. Values of Δ^{sn} that push $\pi^{**}_{e,n^-,n,j,\ell}(\tilde{w})$ out of the $[0,1]$ interval are discarded.

The initial distribution of states includes not only the relationship status r, but the wage deviation ($\tilde{w}$), spousal shock ($\tilde{s}$, set to zero when single), and the age and number of children (a,n). To simplify, we assume that the initial value of $\tilde{w}$ is independent of the initial values of r and (a,n). Let $sh_{r,a,n}$ denote the fraction of the initial distribution with relationship status r and child configuration (a,n). As before, we introduce type differences as changes to the probability of being single:

$$sh_{sn,a,n,1} = sh_{sn,a,n}(1+\Delta^{s0}), \tag{A35}$$
$$sh_{sn,a,n,2} = sh_{sn,a,n}\left(1-\tfrac{p_1}{1-p_1}\Delta^{s0}\right), \tag{A36}$$
$$\frac{sh_{m,a,n,2}}{sh_{m,a,n,1}} = \frac{sh_{en,a,n,2}}{sh_{en,a,n,1}} = \frac{sh_{dv,a,n,2}}{sh_{dv,a,n,1}}. \tag{A37}$$

To implement equation (A37), let $\varsigma_{a,n} = sh_{sn,a,n}\big/\sum_r sh_{r,a,n}$ denote the initial fraction of men with child configuration (a,n) who are single. It then follows from equation (A35) that

$$sh_{r,a,n,1} = sh_{r,a,n} \frac{1 - \varsigma_{a,n} * (1 + \Delta^{s0}_{a,n})}{1 - \varsigma_{a,n}} \tag{A38}$$
$$= sh_{r,a,n} \left(1 - \Delta^{s0}_{a,n} \frac{\varsigma_{a,n}}{1-\varsigma_{a,n}}\right), \ \ r \in \{en, m, dv\}.$$

It likewise follows from equation (A35) that

$$sh_{r,a,n,2} = sh_{r,a,n} \left(1 + \Delta^{s0}_{a,n} \left(\frac{p_1}{1-p_1}\right) \frac{\varsigma_{a,n}}{1-\varsigma_{a,n}}\right), \ \ r \in \{en, m, dv\}. \tag{A39}$$

E.3 Preferences

The third way the two types differ is in their disutility from work, $\psi_{e,j}$. Previously, the disutility of work was $\psi_{e,j} = \psi_{e,0}(1 + \psi_{e,1} \cdot j + \psi_{e,2} \cdot j^2)$. The type-specific adjustments are:

$$\psi_{e,\ell,0} = \begin{cases} \psi_{e,0}\left(1 + \Delta^{\psi}\right), & \ell = 1, \\ \psi_{e,0}\left(1 - \frac{p_1}{1-p_1}\Delta^{\psi}\right), & \ell = 2, \end{cases} \tag{A40}$$

$$\psi_{e,2,1} = \psi_{e,1,1} = \psi_{e,1}, \ \ \psi_{e,2,2} = \psi_{e,1,2} = \psi_{e,2}. \tag{A41}$$

Our prior belief is that Δ^{ψ} will be positive: type-1 men are less willing to work. This the easiest of the type-related parameters to estimate, as it has no effect on the second-stage processes.

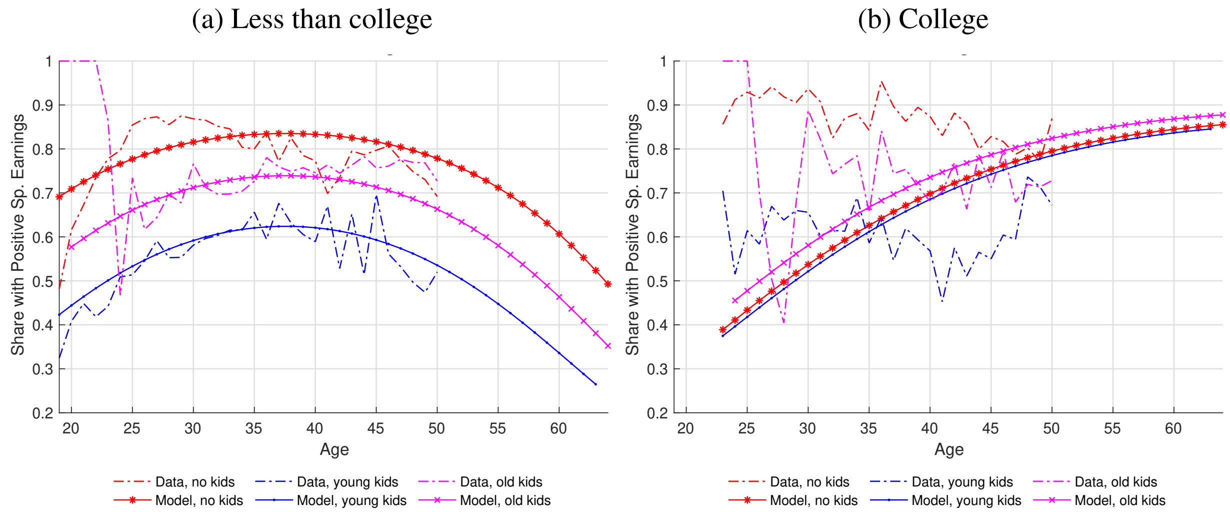
F Model Fit of Spousal Employment and Earnings

The parameters of the spousal earnings process shown in the lower panel of Table 2 are set to match the age profiles of spousal employment, the mean and standard deviation of log earnings among working spouses, and the correlation of spousal earnings and male wages. Because the NLSY79 does not provide earnings information for unmarried partners, the simulated profiles and the data targets both use data only for the years the couples are married.

Figure A5 compares the model's predictions of spousal employment to those found in the NLSY79. The model replicates the lower rates of employment among women with children, especially young children. Figure A6 provides the corresponding comparison for the logged earnings of employed spouses. In contrast to employment, spousal earnings vary relatively little by family composition – in the model the only differences are (miniscule) selection effects.

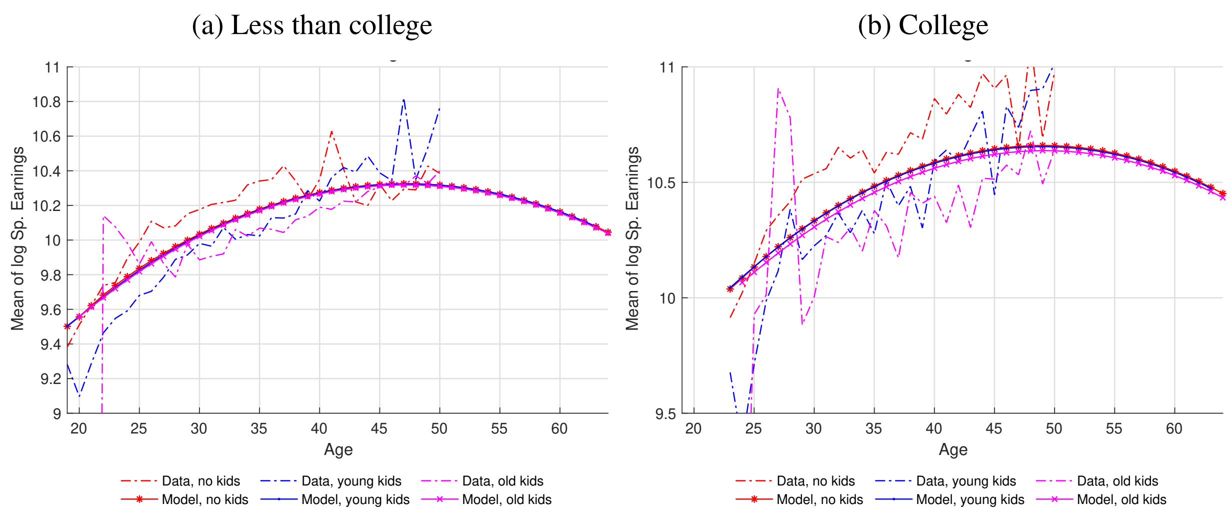
The first row of Table A4 shows the standard deviation of these earnings, again conditional

Figure A5: Spousal Employment by Age, Education and Age of Children, Model and Data



Source: Data are for spouses of married men ages 19-50 (less than college) or 23-50 (college graduates) in the NLSY79; see text for details. Model results are authors' calculations; see text for details.

Figure A6: Spousal Earnings (If Employed) by Age, Education and Age of Children, Model and Data



Source: Data are for spouses of married men ages 19-50 (less than college) or 23-50 (college graduates) in the NLSY79; see text for details. Model results are authors' calculations; see text for details.

on working. This statistic helps pin down β_e^s, the coefficient on the shock $\tilde{s}$ in the spousal earnings equation, and σ_e^s, the volatility of this shock. While we target standard deviations on an age-by-age basis, the data moments are noisy, and we therefore report just the unconditional average across ages. Spousal earnings are quite volatile: the standard deviation of logged earnings ranges between 0.65 and 0.68.

Table A4: Standard Deviation of Spousal Earnings and Correlation with Male Wages, Model and Data

	Less than College		College Graduates	
Statistic	Data	Model	Data	Model
Standard deviation, $\ln(se_j)$	0.640	0.612	0.667	0.646
Correlation $(\ln(se_j), \tilde{w}_j)$				
No children	0.283	0.295	0.181	0.182
Young children	0.177	0.171	0.091	0.118
Older children	0.111	0.144	0.011	0.092

Note: Data are for married men (and spouses) ages 19 - 50 (less than college) or 23 - 50 (college graduates) in the NLSY79 and are restricted to observations with positive values; see text for details. Model results are authors' calculations; see text for details. Standard deviations and correlations are calculated age-by-age; reported above are unconditional averages.

Our final target is the correlation between male wages and spousal earnings, conditional on both parties working, as a function of their childrens' age. This moment, along with the earnings' variance, helps identify ρ_e^s and σ_e^s, the coefficient on male wages and the idiosyncratic variation, respectively, in the distribution of the shock $\tilde{s}$. Table A4 shows that among couples with no children, the correlations are around 0.18 for those with a college degree and close to 0.30 for those without. The correlation coefficients for couples with children are even lower.

G Estimating the Dynamics of Family Structure

G.1 Main Estimates

Fertility dynamics are governed by $\phi^n_{n,r,j}$, the probability of a new child being born given existing children n, relationship status r, and man's age j. New children stop arriving once the existing children age. Let $\phi^a_{yc,j}$ denote the probability that all the young children of an age-j man become old children; and $\phi^a_{oc,j}$ the probability that all the old children become grown. All of these probabilities are modeled as logistic functions of a quadratic polynomial in the man's age.

The parameters governing relationship dynamics are $\phi^{en}_j(\tilde{w})$, the probability of a single man of age j becoming engaged; $\phi^{en\text{-}c}$, the probability that a engaged man cohabits with his partner; $\phi^{mr}(\tilde{w})$, the probability of an engaged man becoming married; ϕ^{owb}, the effect of an out-of-wedlock birth on the probability of a relationship advancing; and $\phi^d_{n,j}$, the probability of a married

Table A5: Parameters for Family Dynamics

Parameter Name	Parameter	Value, Non-College	Value, College
Children (logistic coefficients)			
First child, pre-marital	$\phi^{n}_{sn,0,j}$	$(-3.684, 0.135, -0.919)$	$(-4.482, 0.170, -1)$
First child, married	$\phi^{n}_{mr,0,j}$	$(-0.365, -0.110, -0.081)$	$(-2.132, 0.203, -1.408)$
Second child, pre-marital	$\phi^{n}_{sn,1,j}$	$(-2.348, 0.181, -1.319)$	$(-2.007, -0.156, 0.061)$
Second child, married	$\phi^{n}_{mr,1,j}$	$(-1.105, 0.026, -0.112)$	$(-1.291, 0.069, -0.331)$
Third child, pre-marital	$\phi^{n}_{sn,2,j}$	$(-1.619, 0.066, -0.792)$	$(-15, -1.901, -10)$
Third child, married	$\phi^{n}_{mr,2,j}$	$(-1.107, -0.110, 0.140)$	$(-2.238, 0.084, -0.753)$
Young children age	$\phi^{a}_{yc,j}$	$(-4.432, 0.211, -0.341)$	$(-5.250, 0.262, -0.349)$
Old children age	$\phi^{a}_{oc,j}$	$(-15, 0.604, -0.608)$	$(1, -1, -10)$
Relationship Dynamics (logistic coefficients)			
Engagement	ϕ^{en}_{j}	$(-2.125, 0.028, -0.700)$	$(-1.514, -0.138, 0.205)$
Marriage	ϕ^{mr}	0.0	0.0
Impact of wage shock	θ	0.228	0.201
Effect of pre-marital child	ϕ^{owb}	-0.196	3.172
Divorce, $n < 3$	$\phi^{d}_{n,j}$	$(-1.778, -0.282, 0.500)$	$(-4.321, 0.044, -1.077)$
Divorce, $n = 3$	$\phi^{d}_{n,j}$	$(-1.657, -0.150, 0.113)$	$(-15, 1.238, -3.530)$
Relationship Dynamics (probabilities)			
Initial relationship distribution	(sn, en)	$(0.814, 0.100)$	$(0.714, 0.100)$
Fraction cohabiting	$\phi^{en\text{-}c}$	0.5	0.5

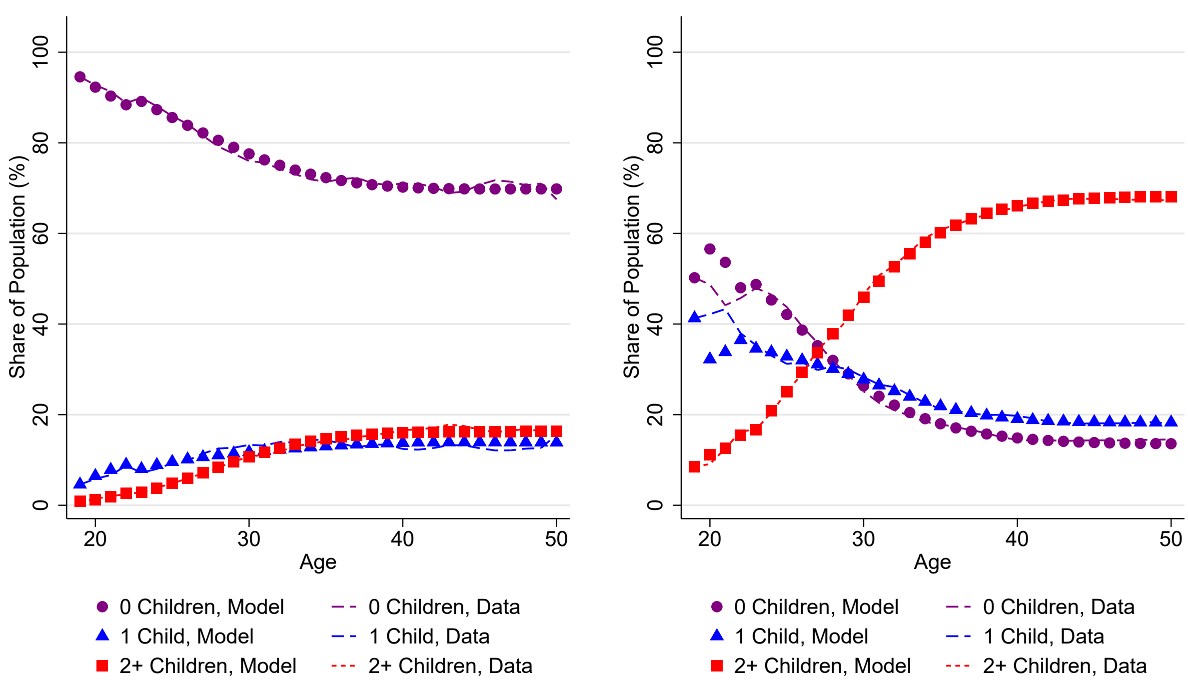
Note: Superscripts are used to distinguish parameters, while subscripts are used to distinguish dependencies. All parameters with the age subscript j utilize a quadratic in age. See section 4.3.3 and Appendix G for details.

man of age j with n children becoming divorced. All of these probabilities are modeled as logistic functions, but their arguments differ. The probabilities of becoming engaged and getting divorced depend on the man's age via a quadratic polynomial. The probability of transitioning from engagement to marriage is age-invariant; because we do not observe the engagement status of unmarried men, we simplify the model by assuming that all the age-related variation prior to marriage is captured in the rate at which engagements form. Data limitations also lead us to assume that the effect of a pre-marital birth (ϕ^{owb}_{e}) is the same for single and engaged men, and that cohabitation does not effect the rate at which engaged couples marry. (We set the fraction of engaged couples who cohabit, $\phi^{en\text{-}c}$, to 1/2 for all ages and education levels, consistent with the cohabitation fraction in the NLSY79.) Because the divorce probability is notably higher for men with three or more children in the data, we estimate two sets of divorce probabilities, one for men with less than three children, and a second for men with three children.

We estimate these parameters separately for each education group, using the simulated method of moments. Specifically, we target the following empirical age profiles: the share of men who are never married and who have n children, for $n \in \{0, 1, 2, 3+\}$; the share of men who have ever been married and who have n children, the share of divorced men who have n children, and the share of newly married men who have had at least one child. The latter age profile is informative about the effect of pre-marital children on marriage probability. We also target the coefficient on male wages in a regression of marriage transition probabilities; see section 4.3.3 and Appendix G.2 for details. Table A5 shows the resulting parameter estimates. Figures 6a and 6b in the main text and Figure A7 in this Appendix show model fits.

Figure A7: Number of Children by Marital Status, Model and Data

(a) Never married (b) Ever married



Source: Sample consists of men ages 19-50. Ages 19-22 include only men with less than a four -ear college degree; ages 23-50 include all education groups. Never-married men are those who at the age in question have yet to marry. Data correspond to the NLSY79. Model results are author calculations; see section 4.3.3 and Appendix G for details.

G.2 Background Estimates: State-Level Wages and Marital Transitions

Table A6 shows results from a linear probability regression of marriage on male wages. We transform wages using the inverse hyperbolic sine function. This allows for a (near-) logarithmic relationship when wages are positive, but also accommodates values of zero. The first column of the table shows the results from an OLS regression. The estimated coefficient on wages is 0.0092. The second column shows the results for an IV regression where we instrument for each individual's wages with the average wages in his state of residence. The F-statistic for state-level wages in the first-stage regression is 254.69, highly significant and indicative of instrument relevance. Instrumenting for wages causes the coefficient to increase to 0.0144. This is the coefficient value we target when estimating our model of relationship dynamics. In particular, we set the model parameter θ, which governs how wages affect relationship transitions, so that the wage coefficient in our model-simulated data also equals 0.0144.

Table A6: Predictors of State-Level Marital Transitions

	OLS	2SLS
Constant	0.0206**	0.0252**
	(0.0082)	(0.0110)
asinh(*wage*)	0.0092***	0.0144*
	(0.0008)	(0.0083)
Age	0.0065***	0.0059***
	(0.0014)	(0.0017)
$Age^2/100$	−0.0279***	−0.0250***
	(0.0096)	(0.0107)
College +	0.0124***	0.0089***
	(0.0030)	(0.0062)
New child	0.5300***	0.5288***
	(0.0102)	(0.0104)
Year FEs	*Y*	*Y*
State FEs	*Y*	*Y*
R^2-adj	0.10	0.10
N	26,290	26,290

Source: 1982-2019 waves of the CPS ORG. Sample is men ages 19-54.

H Third-stage Parameters

Table A7 presents the third-stage parameter estimates. As discussed in Section 4.4 and Appendix E, β is set to match mean assets at age 50, while the type-related parameters are set to match the gaps in hours and wages between the ever- and never-married.

Table A7: Third-stage Parameters

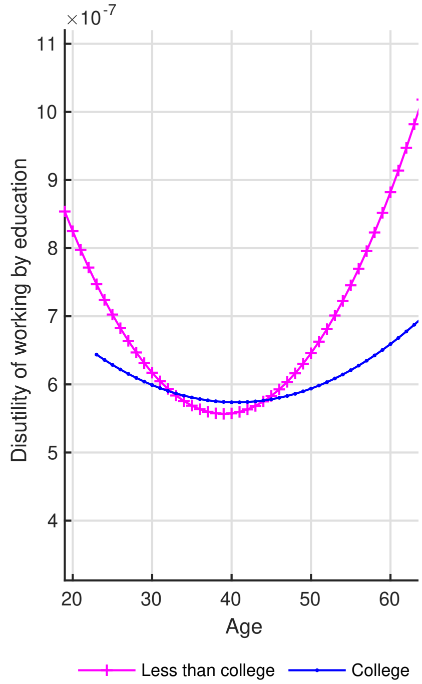
Parameter Name	Parameter	Value
Not Type-related)		
Discount factor	β	0.98
Work disutility age quadratic, non-college	$\psi_{nc,j}$	$(4.77\times10^{-7}, -8.96\times10^{-7}, 1.33\times10^{-3})$
Work disutility age quadratic, college	$\psi_{c,j}$	$(4.92\times10^{-7}, -1.25\times10^{-3}, 3.95\times10^{-4})$
Type-related		
Probability of being type 1	p_1	0.62
Wage shifter	Δ^w	0.12
Relationship transition probability shifter	Δ^{sn}	0.12
Initial relationship share shifter	Δ^{s0}	0.00
Work disutility shifter	Δ^{Ψ}	0.17

Note: Author's calculations. See section 4.4 and Appendix E for details. In the absence of the type adjustment, work disutility is $\psi_{e,j} = \psi_{e,0}(1+\psi_{e,1}\cdot j+\psi_{e,2}\cdot j^2)$, with j defined as calendar age less 39.

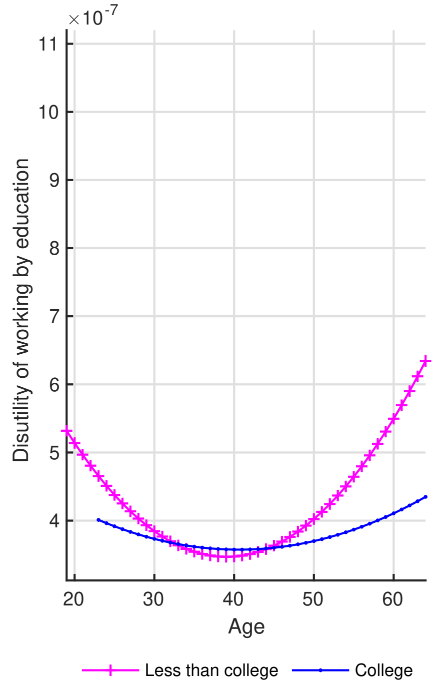
Figure A8 shows how the estimated disutility of working, $\psi_{e,\ell,j}$, varies across the life cycle. As discussed in Section 4.4, for each education group, our estimate of work disutility is the quadratic function that allows the model to best fit the life-cycle hours profiles found in the NLSY79. The figure distinguishes between type-1 (left panel) and type-2 individuals (right panel), showing that type-1 individuals dislike work much more intensely.

Figure A8: Work Disutility by Age, Education and Type

(a) Type 1



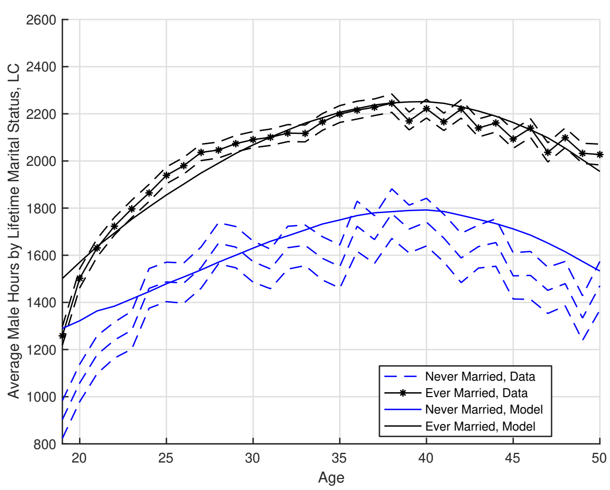
(b) Type 2



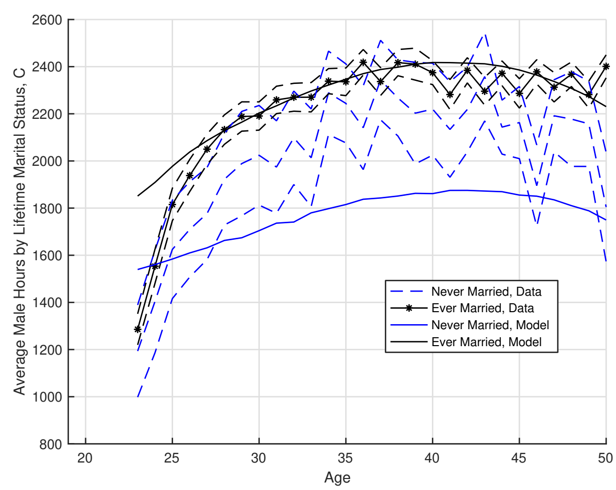
Source: Author calculations. See Section 4.4 for details.

Figure A9: Mean Hours Worked by Age, Education, and Marital History: Model and Data

(a) Less than college



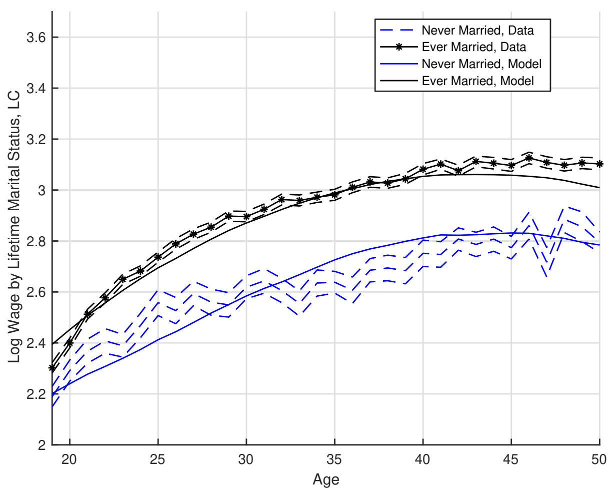
(b) College



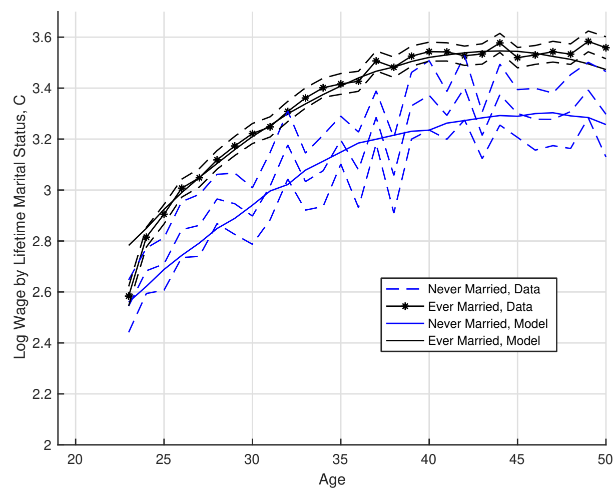
Source: Sample consists of men ages 19-50 for men without a four-year college degree and men ages 23-50 for men with a four-year college degree. Data correspond to the NLSY79. The dotted bands represent the 95% confidence intervals for the data. Model results are author calculations; see section 5.1 and Appendices E and H for details.

Figure A10: Mean Log Wages by Age and Marital History: Model and Data

(a) Less than college



(b) College



Source: Sample consists of men ages 19-50 for men without a four-year college degree and men ages 23-50 for men with a four-year college degree. Data correspond to the NLSY79. The dotted bands represent the 95% confidence intervals for the data. Model results are author calculations; see section 5.1 and Appendices E and H for details.

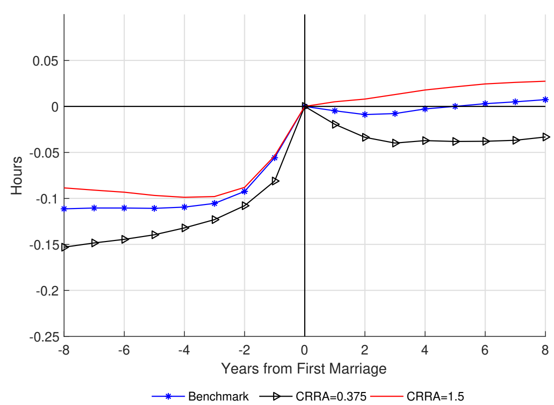
I Sensitivity Analyses

We now assess the sensitivity of our results to two key parameters, namely the coefficient of relative risk aversion (γ) and the Frisch elasticity of labor supply (ξ), halving or doubling each parameter from its benchmark value. Figure A11 shows that over this range our qualitative results do not depend on specific parameter values, and the majority of the quantitative results are robust as well.

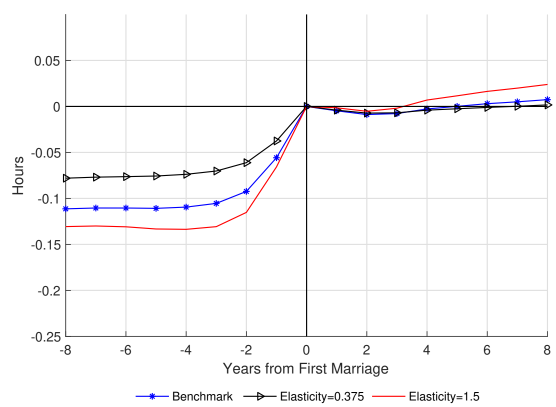
Figure A11a displays the change in hours around marriage for different values of γ. As γ increases from 0.375 to 1.5, the run-up in hours prior to marriage declines somewhat, from 15% to 9%, while the change in hours after marriage switches from a 3% decrease to a 3% increase. The net result is that the total change in hours around marriage varies little within this parameter range. To interpret these results, it is helpful to rewrite the marginal utility of consumption as $MU_c = N_f \eta_f^{\gamma-1} c^{-\gamma}$. Under our parameterization, when an unmarried man without children marries, N_f increases from 1 to 2, while η_f increases from 1 to roughly 1.6. For any $\gamma > 0$, $N_f \eta_f^{\gamma-1}$ will increase, raising the marginal utility of consumption and encouraging work. If the couple then has children, η_f grows even larger, but N_f remains at 2. Noting that most children arrive after marriage, it follows that after marriage the product $N_f \eta_f^{\gamma-1}$ falls when γ is less than one and rises when it is greater than one. This is indeed consistent with Figure A11a, which shows hours falling after marriage when $\gamma = 0.375$ or the benchmark value of 0.738 (at least initially), but rising when $\gamma = 1.5$.

Figure A11: Hours of Work by Distance from Marriage: Effects of Preference Parameters

(a) Coefficient of relative risk aversion, γ



(b) Frisch elasticity, ξ



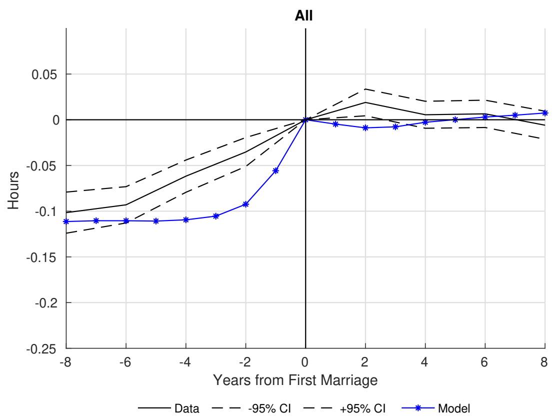
Source: Authors' calculations. See text for details.

The effects of changing ξ, shown in Figure A11b, are more straightforward to interpret. As the Frisch elasticity increases from 0.375 to 1.5, the hours responses grow in size, but their signs remain the same.

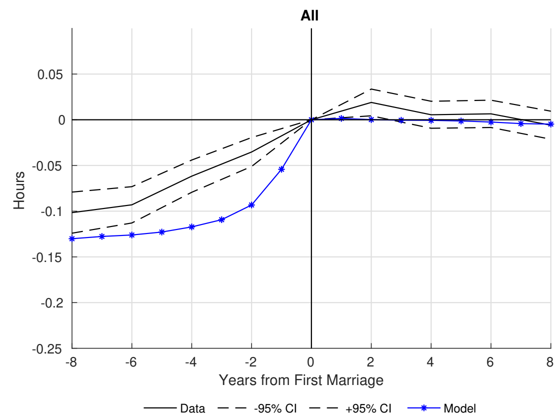
J Alternative Specifications

Figure A12: Hours of Work by Distance from Marriage: Alternative Specification

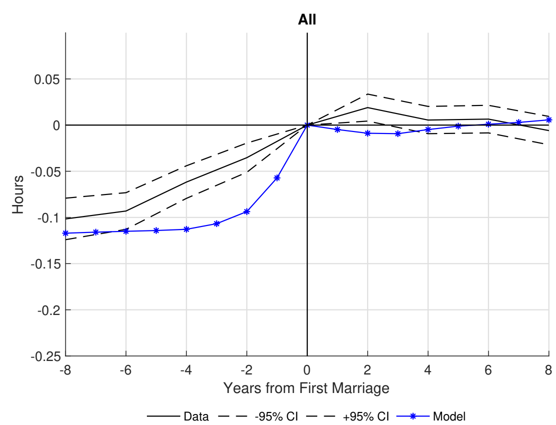
(a) Baseline model



(b) One-Type Alternative



(c) Dynamic selection doubled



Source: Authors' calculations; see text for details.