Federal Reserve Bank of Dallas

Marriage and Work Among Prime-Age Men

Adam Blandin, John Bailey Jones and Fang Yang

Working Paper 2313 August 2023 (Revised December 2025)

Research Department

https://doi.org/10.24149/wp2313r1

Working papers from the Federal Reserve Bank of Dallas are preliminary drafts circulated for professional comment. The views in this paper are those of the authors and do not necessarily reflect the views of the Federal Reserve Bank of Dallas or the Federal Reserve System. Any errors or omissions are the responsibility of the authors.

Marriage and Work Among Prime-Age Men[*]

Adam Blandin[†], John Bailey Jones[‡] and Fang Yang[§]

First version: August 23, 2023
This version: December 8, 2025

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

[*] We are grateful to Stefania Albanesi, Alex Bick, Bettina Brueggemann, Katrina Mullen, Luigi Pistaferri and seminar participants at the 2019 Arizona State University Reunion Conference, 2020 Southern Economic Association Annual Meeting, 2021 Midwest Economic Association Meeting, and 2022 Duke/FRB Richmond/UVa Jamboree, CAU Conference on Structural Approach in Public Economics, 2022 BSE Summer Forum Workshop, 2022 SED Annual Meeting, 2023 System Applied Micro Conference, the 2023 annual meeting of the Society of Economics of the Household, 2024 BSE Summer Forum Workshop, University of Oklahoma, University of Alberta, Bank of Mexico, USC, and 2025 SAET conference for useful comments. We thank Benjamin Hoham for outstanding research assistance. Yang gratefully acknowledges the support from E. J. Ourso College of Business Summer Research Funding at LSU. This research was supported through computational resources provided by the Big-Tex High Performance Computing Group at the Federal Reserve Bank of Dallas. The opinions and conclusions are solely those of the authors and do not reflect the views of the Federal Reserve Bank of Dallas, the Federal Reserve Bank of Richmond, or the Federal Reserve System.

[†] Adam Blandin, Vanderbilt University, adam.blandin@vanderbilt.edu.

[‡] John Bailey Jones, Federal Reserve Bank of Richmond, john.jones@rich.frb.org.

[§] Fang Yang, Federal Reserve Bank of Dallas, fang.yang@dal.frb.org.

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, *The 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. The first is that in the cross section, married men work substantially more than men who have never been married. For example, among men ages 19 - 54 in the Current Population Survey (CPS), currently married men work at least 30% more annual hours than men who have never been married, a gap which has remained roughly constant since 1975 (see Section 2). 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.[1] The second observation is that over the past 50 years, marriage rates and male hours of work have both fallen markedly. CPS data show that between 1970 and 2018, marriage rates among men ages 19 - 54 fell by 30 percentage points, while hours of work fell by 7.7%.

One potential driver of the correlation between male marital status and hours worked is selection: men with better labor market outcomes are more willing to marry and are more attractive marriage prospects. An alternative possibility is that marriage, or the prospect of marriage, increases the labor supply of men. In this paper, we combine reduced form empirical analyses and a structural model to quantitatively assess the importance of these two potential channels. 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,

[1] 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).

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

We begin by documenting a strong reduced form relationship between male marital status and hours worked. Married men work roughly 30% more hours than men who have never been married, and this relationship has remained remarkably stable for at least the last half century. A simple accounting exercise reveals that the decline in marriage rates between 1970-2018 can account for 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* 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, and use it to evaluate two categories of explanations. The first category we refer to as dynamic selection: we allow men receiving higher wage shocks, who respond by increasing their labor supply, to experience a higher likelihood of marriage. The second category we refer to as a "mouths-to-feed" effect: the prospect of marriage raises male labor supply. In the model, marriage and children can impact male labor supply via several channels, but the most important is that men internalize the utility of family members, raising the man's marginal utility of consumption, as in Blundell, Pistaferri and Saporta-Eksten (2016) and Fan, Seshadri and Taber (2024). This effect leads men to work more when married and to increase their hours of work in anticipation of starting a family.

We calibrate the model using data from the NLSY79. The calibrated model closely matches the life-cycle evolution of marital status and children, as well as the correlation of earnings between spouses. By including two permanent worker types, who differ in wages, the probability of getting married, and preferences for work, the model is also able to match differences in the wages and average hours of never-married and ever-married men. To discipline the strength of the dynamic selection effect, we require the model to replicate the marginal effect of wages on marriage probabilities found in the CPS, where we instrument for individual wage changes with state-level variation.

The calibrated model generates hours dynamics around the time of marriage resembling those in our fixed effect regressions, mainly through the mouths-to-feed channel. Our estimates from the

CPS imply that positive wage shocks are associated with positive selection into marriage. However, in the model this force is quantitatively weak. Moreover, because empirically hours increase more than wages prior to marriage, the model would require larger-than-typical labor supply elasticities for wage-based selection to be the sole explanation. Additional evidence against the selection explanation comes from the behavior of men whose marriages are *preceded* by pregnancies. Responses in the NLSY79 indicate that pre-marital pregnancies are more likely to be unplanned, and thus less likely to be driven by labor market shocks. Yet male hours increase by more for marriages preceded by pregnancies than marriages where children arrive later.[2]

Our model indicates that marriage has a substantial positive effect on male work hours. We conclude our analysis with two sets of numerical experiments quantifying the impact of marriage on aggregate male hours of work. First, a simple counterfactual exercise shows that eliminating marriage altogether reduces average hours worked by prime-age men by 6.8%. Next, we quantitatively assess Binder and Bound's (2019) hypothesis that the that the decline in marriage rates in recent decades has contributed to the decline in male hours over the same time period. We do so by re-estimating the marriage and family formation processes with data from the NLSY97, a longitudinal dataset similar to the NSLY79 that follows a cohort born 21 years later. 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. When we simulate our structural model using the NLSY97 marriage process, but leave all other parameters constant, we find that average hours worked by prime-age men fall 1.59%. For context, over the typical 21-year interval in our CPS sample, trend hours fell by 3.34%.[3] In other words, our model counterfactual with reduced marriage rates explains 48% of the observed decline in male hours. Moreover, once we augment the experiment to include cross-cohort changes in male wages and spousal earnings, our model generates a decline of 3.15%, 94% of its CPS counterpart.

Our project lies at the intersection of three literatures that are related but have nevertheless remained largely isolated from each other. The first is a series of reduced form analyses attempting to explain the "male marriage premium." Most of these focus on the difference in hourly wages between married and never-married men.[4] On average, these papers find that wages are about 10% higher for married men after controlling for observables. The leading causal explanation for the

[2]The hours response to pre-marital children is also evidence against marital selection along other dimensions, such as unobserved heterogeneity in wage *growth* (Guvenen, 2009) or employment shocks (Kaplan, 2012).

[3]Between 1970 and 2018, trend hours in the CPS fell by 7.7%; this implies that in a 21-year interval, hours fell by $7.7\% \times \frac{20.8}{48} = 3.34\%$. (20.8 is the precise cross-cohort gap in birth years.)

[4]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).

male marriage premium in wages is that marriage increases husbands' productivity by allowing them to specialize in market work rather than home production (Becker, 1991).[5] The leading non-causal explanation is that men with higher wages are more likely to marry. Our view is that this literature has not reached a firm conclusion about which explanation is more important. Our model includes both mechanisms: we account for specialization by allowing wages to increase with hours of work; and we account for selection by allowing the probability of marriage to depend on transitory wage shocks. We find that in our model, specialization is the primary driver of the wage increases.

We are aware of two papers within this reduced form literature that emphasize differences in hours, rather than wages, between married and never-married men. 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) study men in the Panel Study of Income Dynamics (PSID). They show that after marriage and the birth of their children, men receive higher wages and work more. However, neither study analyzes the time path of these variables, and they do not attempt to quantify the channels that might generate the hours increase, which is one of our main objectives.

The second literature to which our paper contributes is a collection of structural analyses that explore the interactions of gender, marriage, children and the labor market. 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 of these papers examine only the labor supply decisions of couples, and so cannot speak to differences between married and single individuals.[6] Other papers model male labor supply and earnings as exogenous.[7] There is also work featuring both marital dynamics and endogenous male labor supply, such as Guner, Kaygusuz and Ventura (2012*a*,*b*), and Borella, De Nardi and Yang (2023). But to our knowledge, none of these papers have attempted to explain male labor market behavior around the time of marriage.

Within this literature, our work is most closely related to two existing papers. Siassi (2019) seeks to explain differences in income and wealth by marital status. An important difference between our paper and his is that Siassi (2019) estimates a single marital gap for men and women together, while we emphasize that married men work more hours, even as married women work

[5] 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.

[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), Low et al. (2017), and Caucutt, Guner and Rauh (2021).

less. In addition, Siassi (2019) focuses on cross-sectional differences by marital status, while we also assess individual transitions near the time of marriage. The second closely related paper is Mazzocco, Ruiz and Yamaguchi (2014), who model the link between labor supply, home production, savings and marriage. Like us, they use panel data to document that hours increase in the years around marriage, though unlike us they do not document the qualitatively similar patterns in hourly wages and annual earnings.

Relative to both these papers, our analysis differs in three key dimensions. First, the two existing papers model marriage as an instantaneous shock, while we allow marriage to be preceded by engagements. Second, our paper disciplines the quantitative strength of the selection channel using plausibly exogenous variation in state economic conditions, which is central to understanding *why* hours increase in the run-up to marriage. Third, we conduct a series of counterfactuals demonstrating that marriage is an important determinant of overall labor supply by prime-age men.

Finally, we contribute to the nascent literature studying linkages between the secular declines in marriage and employment among prime-age men (Binder and Bound, 2019; Binder, 2021). In a pair of related event studies, Autor, Dorn and Hanson (2019) find that negative labor demand shocks reduce both marriage and fertility rates, while Kearney and Wilson (2018) conclude that fracking booms increase fertility but not marriage. Our approach is somewhat different from the latter two papers: we use a quantitative model to understand how marriage and labor supply interact within a particular cohort of men (the NLSY79), taking their labor and marriage markets as given. We view our findings that marriage leads to a substantial increase in male labor supply as a complementary input into the larger project of understanding these 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

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

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.

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. We then use panel data from the NLSY79 to establish a direct relationship between marriage and labor market outcomes, tying the facts together.

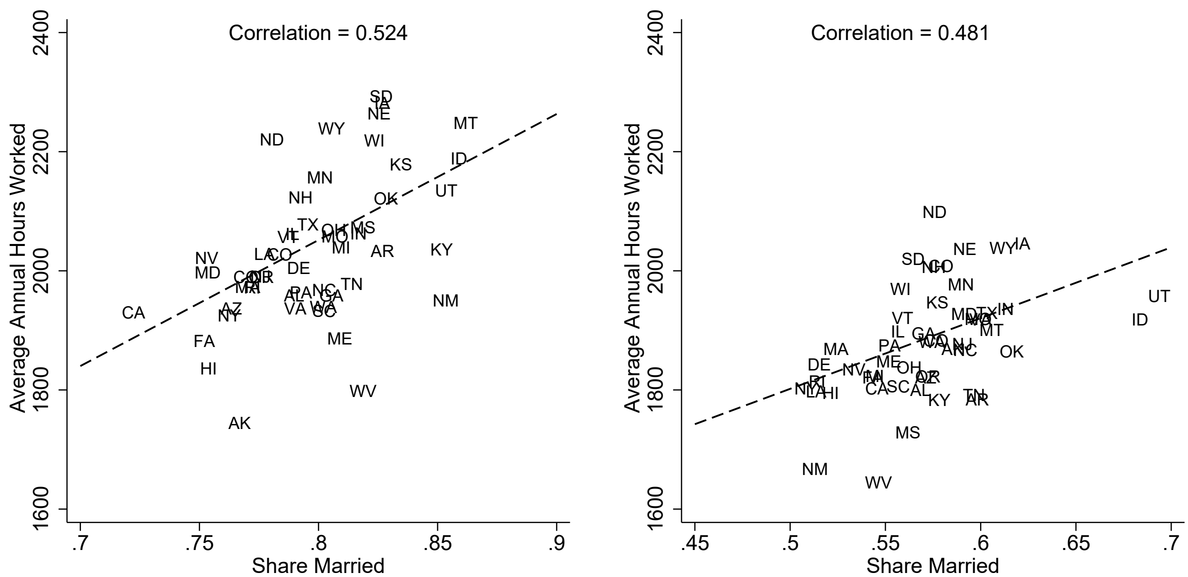
2.1 Marriage and Work in Cross-Sectional Data

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

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

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



Source: Males 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 fit using OLS. Share married refers to the share of men in the sample who were currently married at the time of the survey.

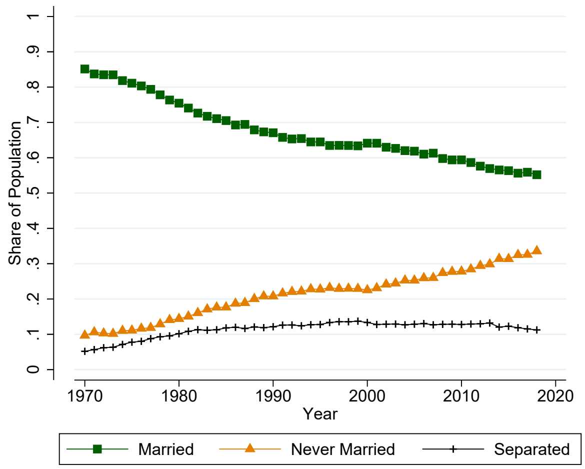
men by 31 to 39 log points. Most of the gap remains after controlling for the mens' education, age, race and state of residence (the grey dashed line with triangles) .

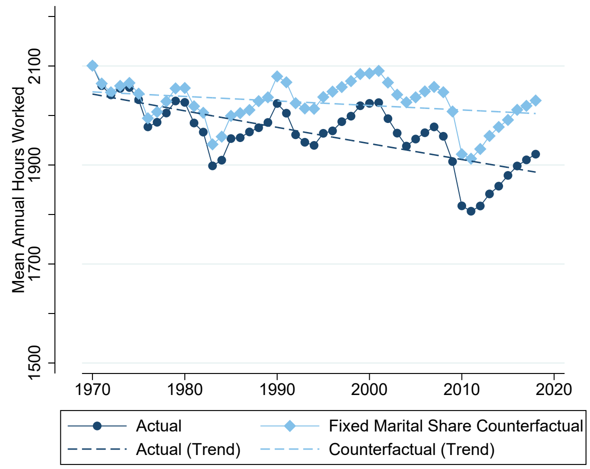
Figure 2b shows results for women. In 1975, married women worked nearly 30 log points less than never-married women. By the 2000's, however, the gap in the raw data had completely disappeared. After controlling for women's observables, the gap is always negative, but nonetheless, by 2018 it was only 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 marital status 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 4 in Appendix A shows that this cross-state pattern continues to hold

[8]In 2015, two noteworthy outliers are Utah and Idaho, with marriage rates of 69% and 68%, respectively. A likely

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





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 17 displays mean hours by marital status used to construct the counterfactual.

after controlling for age, education, and state.

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 displays mean annual hours worked by 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%.

Because married men consistently work more than never-married men over this time period (see Appendix Figure 17), a natural question is whether the decline in marriage can account for an meaningful share of the decline in mean hours. To this end, Figure 3b 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

contributing factor is that these two states have by far the largest population share that is Mormon, a religion which emphasizes the importance of marriage.

prime age males.[9]

2.3 The Dynamics of Marriage and Work in Panel Data

The accounting exercise in the previous section implicitly assumes that hours differ by marital status. What is the source of this 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 this question, we need to move beyond the cross-sectional comparisons in the preceding subsections and make use of panel data.

We therefore 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 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.

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

[9] Table 4 in Appendix A 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.

[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.

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) makes use of 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.

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

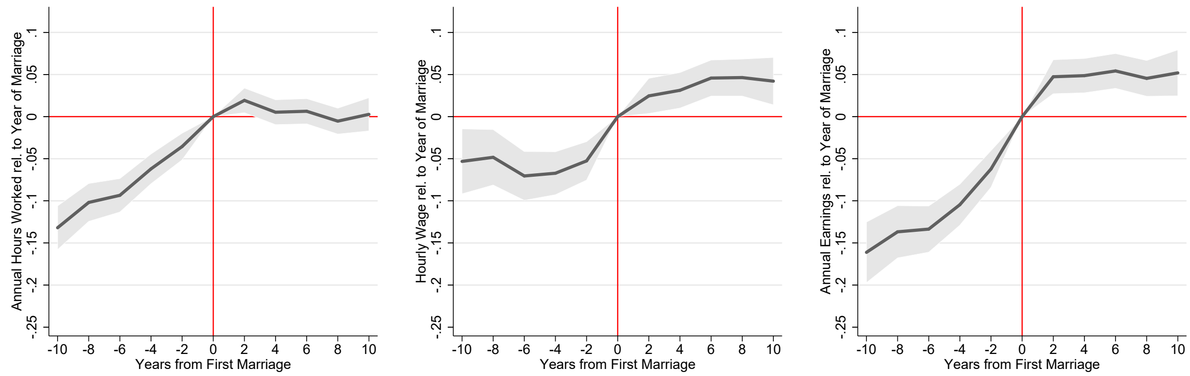
	(1)	(2)	(3)
Constant	1751.0 ***	1808.4 ***	1866.0 ***
	(21.5)	(21.9)	(19.2)
Married	328.1 ***	283.0 ***	99.2 ***
	(10.5)	(10.5)	(13.1)
Separated / Widowed	71.7 ***	86.7 ***	7.7
	(15.1)	(15.1)	(17.7)
Less than High School	–	−228.6 ***	–
		(14.9)	
Some College	–	25.0 **	–
		(11.3)	
Bachelor's +	–	121.7 ***	–
		(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.

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

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.

marriage. Specifically, we run the following regression:

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

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

[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 18 in Appendix A.4 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.

of marriage. Hours start increasing even before then, but the size of the increase (3% over the preceding four years) 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, ten years after marriage, 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 in the year of marriage or later, while the increase in hours occurs almost entirely in or before marriage.

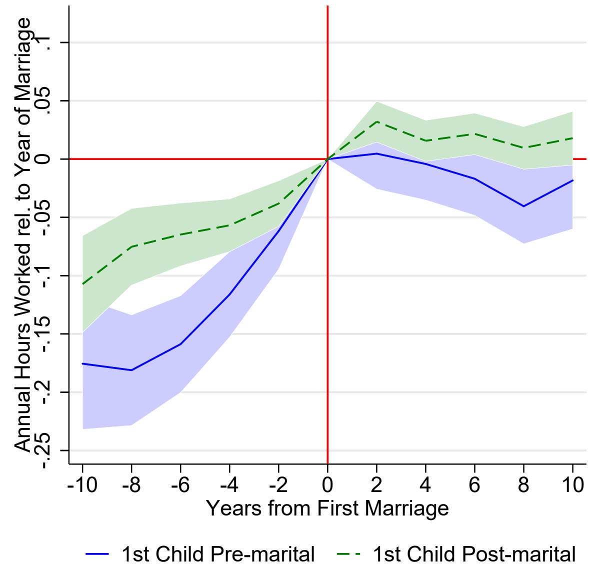
Figure 4c shows the results for annual earnings. The picture is very roughly the sum of the coefficients for annual hours and hourly wages in Figures 4a and 4b: earnings are essentially flat from ten years before marriage to six years before marriage, then increase 18% from six years before marriage 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 to ask 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 for men whose first child appears after their first marriage. Figure 5 displays the results, with the solid blue line corresponding to 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

[12]Figure 19 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.

[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 pregrnancies are three times as likely

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.

have causal effects that encourage work.

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, and remains after controlling for observables and individual fixed effects. Panel data reveal that a substantial part of the cross-sectional difference in work by marital status is driven by 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 events that increase hours of work create or coincide with an increased likelihood of marriage. Because the two explanations have very different implications for how

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.

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 including fixed effects substantially reduces the estimated coefficient on marriage. This suggests that unobserved heterogeneity is an important factor in 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 unobserved type (ℓ) and his wage shock ($\tilde{w}$). 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, $n = 0, 1, \ldots, \bar{n}$. 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 $0 < n < \bar{n}$ 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; this implies a newborn can age immediately into an older child. 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

[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 childrens' age space.

transition probabilities in Appendix B.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 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 transitory wage selection driven by shocks to $\tilde{w}$. Both types of selection operate through the engagement probability $\phi^{en}_{e,\ell,j}(\tilde{w})$. Because $\tilde{w}$ is defined as a zero-mean deviation independent of type, it has little, if any, effect on the type-specific probability of marriage. For clarity, we list the full set of relationship transition probabilities in Appendix B.2.

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. 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 spouse draws the permanent earnings shock

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

The 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:

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

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

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

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

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

In the above equations, β_e^s is a wage-scaling parameter, and $\alpha_{a,e,j}^{s,1}$ and $\alpha_{e,j}^{s,2}$ are mean-shifters that depend on the man's age and education. The mean-shifter for spousal participation 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 childer 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 childrens' 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.

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 worked each period 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

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

other household members. 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:

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

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

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^0_f, τ^1_f, and τ^2_f) 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 B.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

$$V^{sn}_j(e,\ell,k,\tilde{w},a,n) = \max_{c,k',h} \Big\{ u_{f,e,\ell,j}(c,h) + \beta \mathbb{E}_{\tilde{w}',a',n',r',\tilde{s}',q'} \left[V^{r'}_{j+1}(e,\ell,k',\tilde{w}',a',n',\tilde{s},q') \,\middle|\, \tilde{w},n,a,e,\ell \right] \Big\} \tag{16}$$

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

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

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

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. He 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 future relationship possibilities are to remain engaged or to become married.

Engaged men who do cohabit with their partner have the same Bellman Equation as engaged non-cohabiting men, except that: (i) the budget constraint includes spousal earnings and the taxes paid on these earnings; (ii) any children are assumed to now reside with the father, so child support costs, which were proportional to the man's earnings, are replaced with child care costs, proportional to the earnings of the spouse. 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.

[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 $f^{en\text{-}c,sp} = (sn, 0, 0)$.

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 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 set the discount factor so that the mean asset holdings at age 50 generated by the model match those in the NLSY79. We set the work disutility parameters to match life-cycle labor supply profiles.

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-specifc 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 5 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 dominate 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 Section 5.3 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 the gross interest rate R to 1.02, a standard value.

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 6.2 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 C 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. Differences between the types are expressed as as zero-mean deviations. Specifying the differences in this way 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 D.[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 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 works and her earnings when working; (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

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

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 $\alpha^w_{e,j}$. 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 $\alpha^w_{e,j}$ (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^w_e, the innovation standard deviation σ^{ε}_e, and the initial standard deviation $\sigma^{w_0}_e$. The details of this procedure are described in Appendix D.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^s_e), along with the standard deviation of this shock's innovation (σ^s_e); the parameters of the mean-shifter for the probability that

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

Table 2: Parameters for Male Wages and Spousal Earnings

Description	Parameter	Value, Non-College	Value, College Graduates
Male Wages			
Mean-shifter, age quadratic	α_j^w	$(-1.555, 0.067, -0.072)$	$(-2.445, 0.129, -0.145)$
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			
Dependence on husband's wages	ρ^s	0.842	0.720
Standard deviation, innovation	σ^s	0.020	0.614
Effect of children on employment	$\alpha_{a,j}^{s,1}$: yc, oc	$(-1.115, -0.580)$	$(-0.065, 0.200)$
Effect of man's age on spousal employment, quadratic	$\alpha_{a,j}^{s,1}$	$(0.805, 0.088, -0.236)$	$(-0.5, 0.100, -0.100)$
Effect of man's age on spousal earnings, quadratic	$\alpha_j^{s,2}$	$(1.302, 0.056, -0.100)$	$(1.601, 0.055, -0.100)$
Effect of spousal shock 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 D.1 for details.

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). 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 censor all spousal earnings in the data below this level.[18]

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 E 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

[18] 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.

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 E shows that observed spousal earnings are quite variable, leading to large estimated 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 in equations (29)-(31) are modeled as a set of logistic probabilities. We estimate these probabilities to match 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 idiosyncratic 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 F.2 presents detailed results. The estimated coefficient in the CPS regression (Appendix Table 8) 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 F describes our specification and estimation procedure in more detail, and Appendix Table 7 provides the parameter estimates.

4.4 Parameters Set within the Behavioral Model

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 $\psi_{e,j}$. We set $\psi_{e,j}$ so that the life-cycle profiles of hours for each education group generated by the model match those found in the data. 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 G 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 D provides additional, technical details.

Table 9 in Appendix G 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.[19] We also find that type-1 men have a much stronger distaste for work: see Figure 23. 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 over the life cycle in the model (markers) and data (lines). Figure 6b compares the rates of fatherhood (at least one child) by marital status over the life cycle in the model and data.[20] In both figures, the model fits the data closely. Especially notable is the model's fit of the incidence of fatherhood among the 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 22 in the appendices shows that the model also does a good job of matching life-cycle patterns of family size.

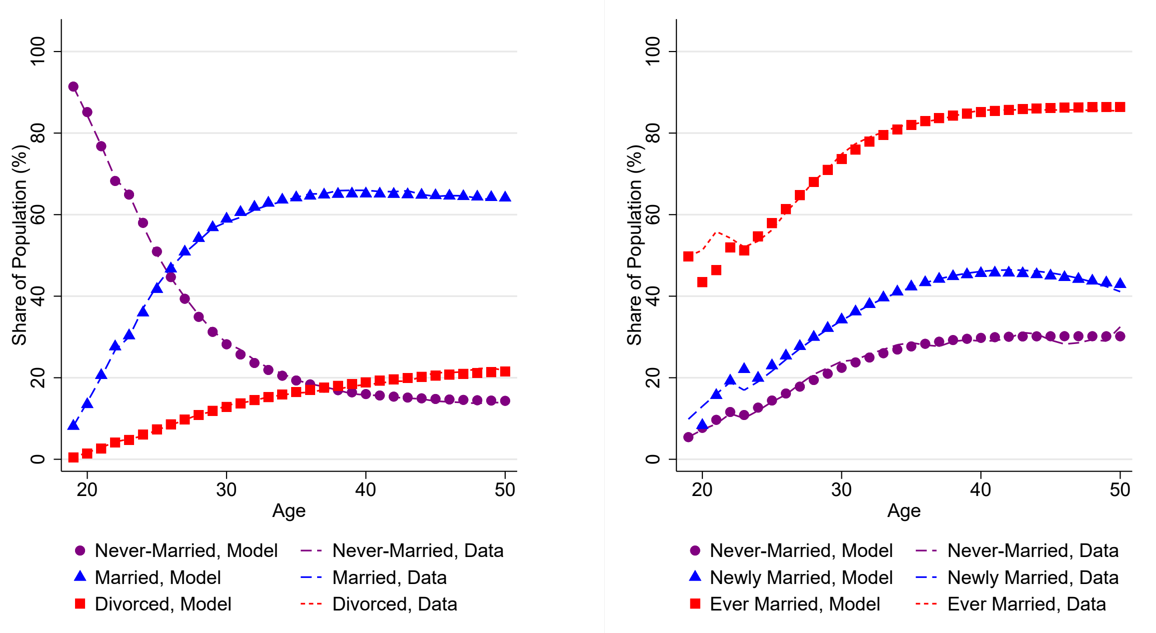
[19] Applying equation (24) yields the ratio $\exp(-0.12)/\exp\left(\frac{0.62}{0.38}0.12\right) = 0.729$. Applying equation (25) yields $(1+0.12)/(1-\frac{0.62}{0.38}0.12) = 1.393$.

[20] 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

(a) Marital status shares

(b) Fatherhood share (one or more children)



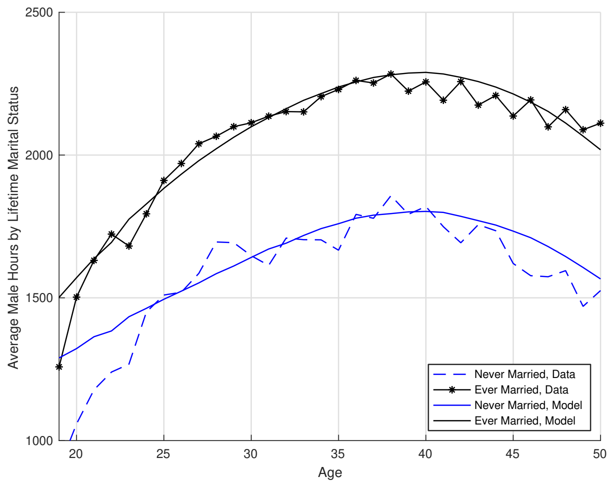
Source: Sample consists of men ages 19-50. Ages 19-22 include only men with less than a four -ear college degree; ages 23-50 include all education groups. Never-married men are those who at the age in question have yet to marry. Data correspond to the NLSY79. Model results are author calculations; see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

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, 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 23 in the appendices.) Figure 7a also shows that the model matches the lower work hours of the never-married. Figure 7b shows that the model does a good job of matching the profiles for wages. Consistent with the data, the model predicts substantially lower wages for the never-married.

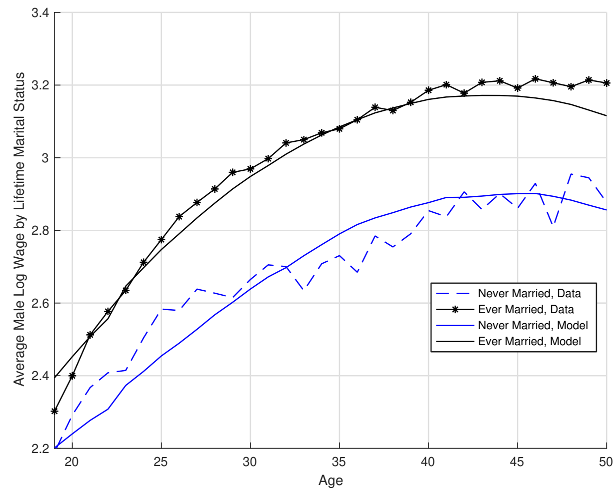
Figure 8 displays the life-cycle profiles for mean (net) assets. While the intertemporal discount factor β is chosen to match mean assets at age 50, assets at other ages are untargeted. Although the model profile tends to fall below its data counterpart, particularly before age 50, the two series track each other reasonably well.

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

(a) Mean log hours worked

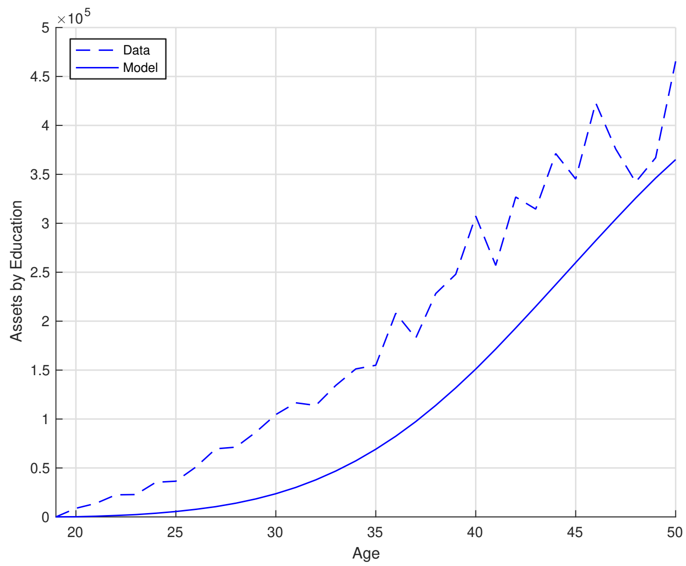


(b) Mean log wage



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 G and D 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 section 4.4 for details.

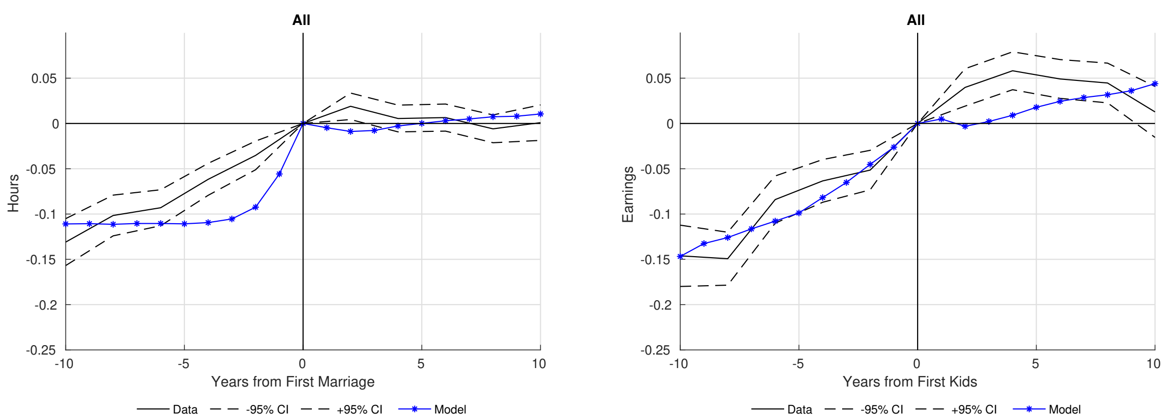
5.2 Marriage, Children and Male Labor Market Outcomes

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. To replicate the panel 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).[21] Because we do not include these regressions in our estimation targets, comparing the data and model coefficients provides us with a validation exercise.

We begin by assessing the model's ability to replicate observed hours dynamics. Figure 9a shows that the model generates most of the marriage-related hours growth found in the data: from six years before marriage to ten years after, hours increase 12% in the model versus 10% in the data. Both model and data show a continuous increase in hours in the years before marriage, although the increase begins somewhat earlier in the data. Consistent with the data, in the model hours change very little once marriage occurs.

Figure 9: Hours of Work by Distance from Marriage or First Child: Model and Data

(a) Distance from date of marriage (b) Distance from date of first child



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

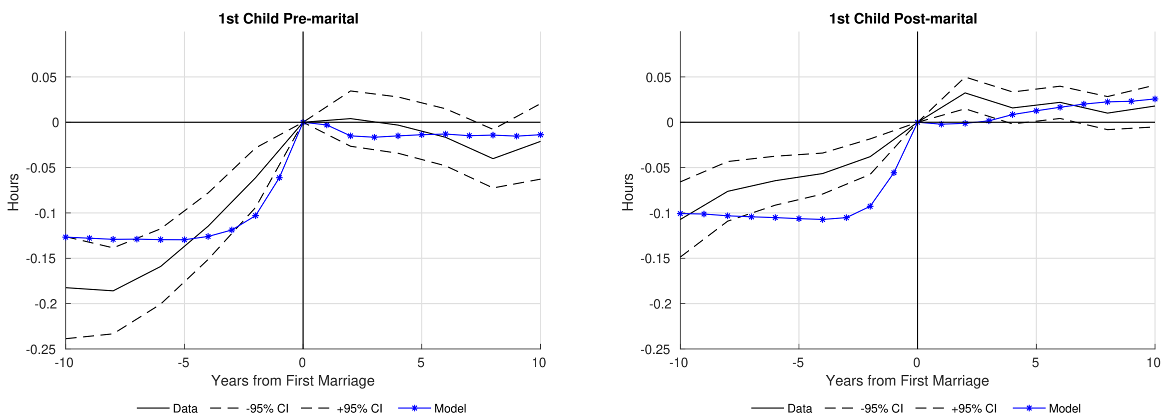
It is natural to ask whether the effect of marriage on hours is primarily a response to the birth (or expected birth) of children. Figure 9b shows that hours do in fact rise steadily as the date of the first child approaches, and that the model generates a similar pattern. Another way to assess the role 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,

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

and is thus "post-marital." Figure 10 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 and 10 thus suggest that both marriage and children are positively associated with male labor supply, and that our model replicates much of this relationship.

Figure 10: Hours of Work by Distance from Marriage and Timing of First Child: Model and Data

(a) First child pre-marital (b) First child post-marital



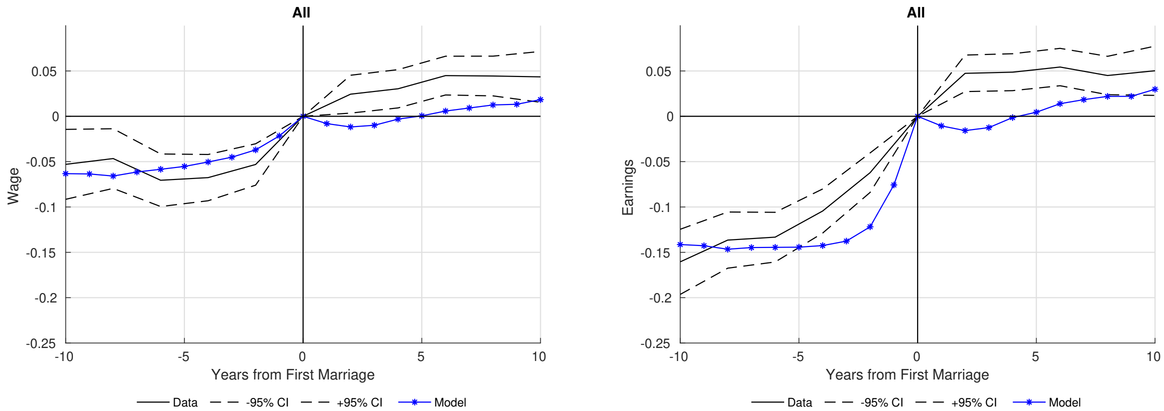
Source: Data results are from regressions using NLSY79 data; see Section 2.3 for details. Model results are authors' calculations; see text for details. Pre-marital children are those born in the year of marriage or before. Post-marital children are those born after the year of marriage.

Figure 11a 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. 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 11b 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 17% increase in earnings from six years before marriage to ten 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

Figure 11: Wages and Earnings by Distance to Marriage: Model and Data

(a) Wages (b) 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.

greatly after marriage.

5.3 Sensitivity Analyses

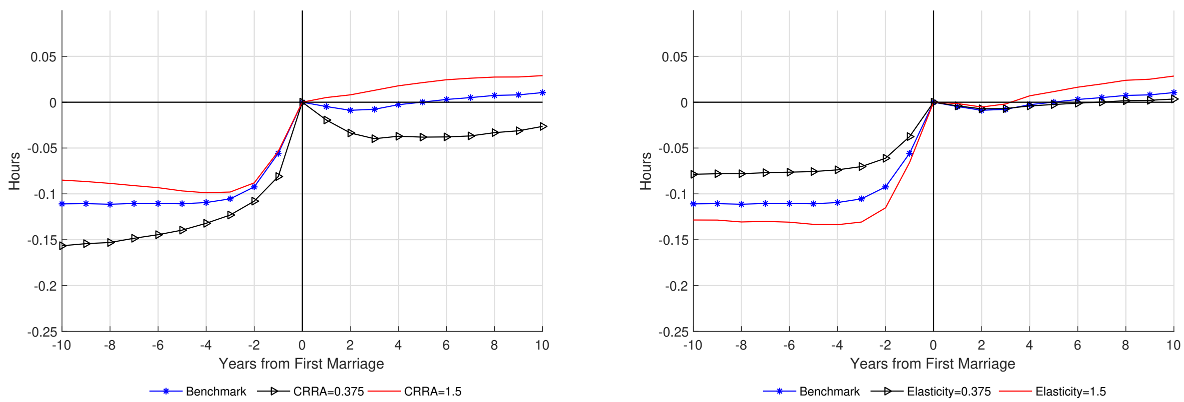
We now assess the sensitivity of our results to two key parameters, namely the coefficient of relative risk aversion (γ) and the Frisch elastiticity of labor supply (ξ), halving or doubling each parameter from its benchmark value. Figure 12 shows that over this range our qualitative results do not depend on specific parameter values, and the majority of the quantitative results are robust as well.

Figure 12a 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 calibration, when a unmarried man 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 12a, which shows hours falling after marriage when $\gamma = 0.375$ or the benchmark value of 0.738 (at least initially), but rising when $\gamma = 1.5$.

The effects of changing ξ, shown in Figure 12b, 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.

Figure 12: Hours of Work by Distance from Marriage: Effects of Preference Parameters

(a) Coefficient of relative risk aversion, γ (b) Frisch elasticity, ξ



Source: Model results are authors' calculations. See text for details.

6 Quantitative Results: Why Do Married Men Work More?

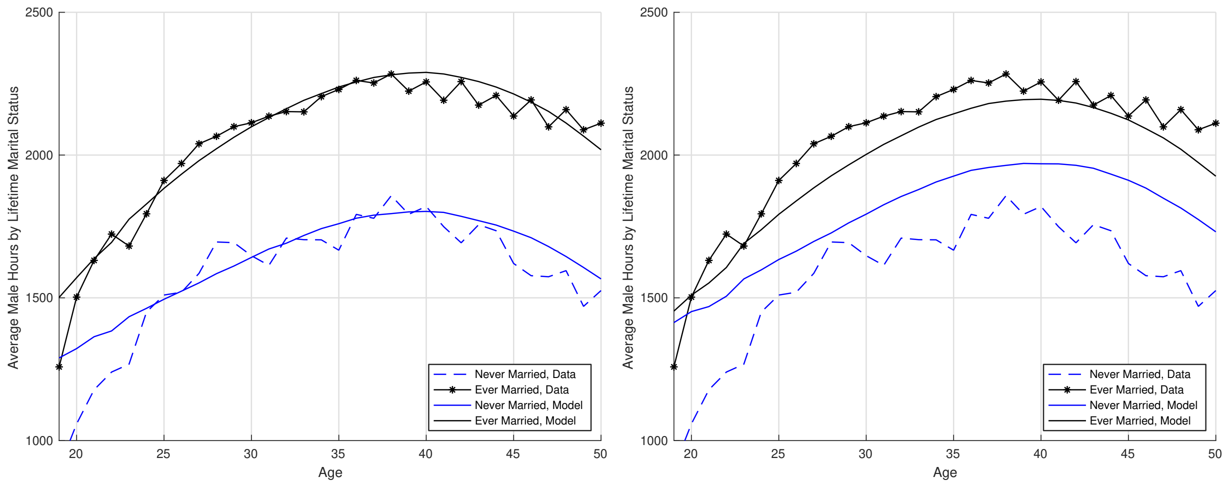
6.1 The Role of Selection into Marriage

Our model contains two forms of marriage-related selection. The first type is permanent, type-related heterogeneity, where men less likely to marry face lower wages and a higher disutility from work. The second type of selection is transitory, where men with higher values of the AR(1) wage shock $\tilde{w}$ are more likely to marry. Figure 13 isolates the effects of the first type of selection. Figure 13a of the figure presents the log hours profiles for ever- and never-married men generated by the data and the baseline, two-type, model, while Figure 13b presents the profiles generated by a version of the model without permanent type differences. Comparing the two sets of model profiles gives us the effects of type differences. At all ages, the hours gap more than doubles when type differences are included, indicating that they drive much of the gap. 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

Figure 13: Mean Hours Worked by Age and Marital History: Model and Data

(a) Baseline model: Two types

(b) One type



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 D and G for details.

the hours gap.

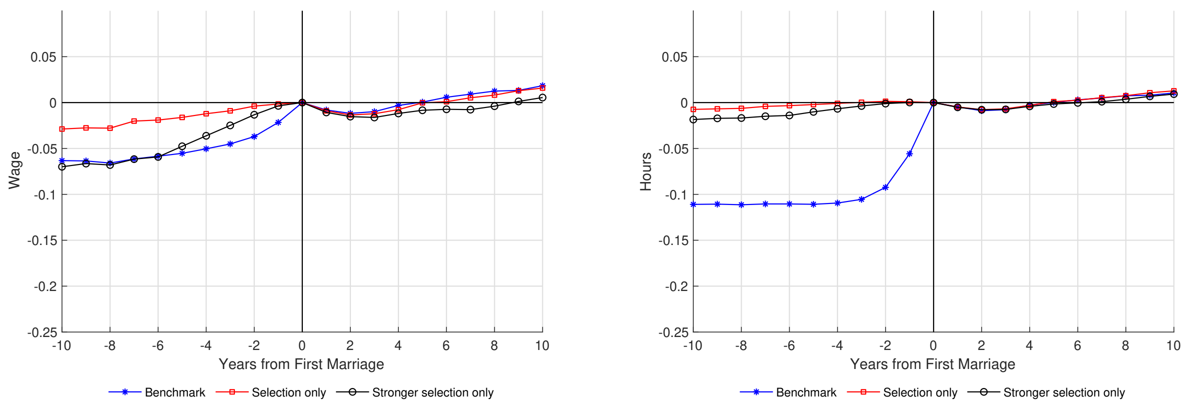
How much of the gap in the one-type model is attributable to dynamic selection? In the model, men with higher transitory wage shocks move up the relationship ladder more quickly. This implies that married men have higher wages in part because of selection on $\tilde{w}$ and, because they respond to positive wage shocks by working more, they also have higher hours. To assess the quantitative impact of dynamic selection on labor market dynamics around marriage, Figure 14 presents results from a specification where the only link between family structure and the labor market is selection. Specifically, in this alternative specification, spouses and children have no effect on preferences or resources: $\eta_f = N_f = 1, \forall f$, and $se_j = \chi_a = \delta_a = 0$.[22] Under this specification, the main link between family structure and the labor market is selection.

Figure 14a shows that the pre-marital increase in wages – from 6 years before to the date of marriage – 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; instead, it was driven mainly by the part-time wage penalty, where higher hours lead to higher wages. This channel is weaker in the

[22] 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.

Figure 14: Labor Market Dynamics 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 wages on relationship transition probabilities is over three times its baseline value. See text for details.

selection-only specification, where hours do not increase very much in the run-up to marriage (see Figure 14b).

To account for the possibility that our estimates understate the degree of wage selection in effect, we introduce a "stronger" selection-only alternative, where we manually increase the coefficient θ until the 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 elasticity we estimate from the data.

Figure 14b compares the change in hours worked around marriage in the benchmark model, the selection-only model, and the stronger selection-only model. 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. For intuition, note 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. The benchmark assumption 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 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 chidren (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.

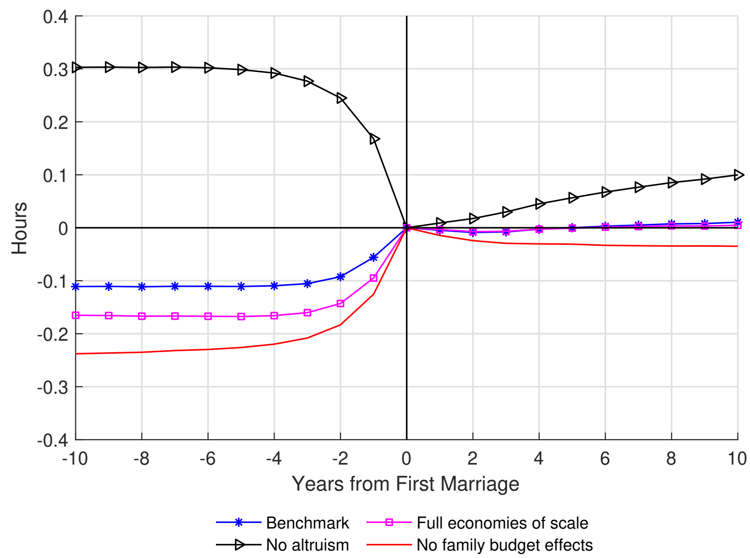
6.2 The Labor Supply Effects of Marriage and Family Structure

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 household'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.

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, marriage and children effectively 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.

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



Source: Model results are authors' calculations. "No altruism" shows hours path for specifciation where $N_f = 1, \forall f$. "No family wealth effects" shows hours path for specification where spouses and children have no direct budgetary effects ($se(1 - n\chi_a) = 0$). See text for details.

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.[23] 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,

[23] Introductions to the theory and construction of equivalence scales include Cowell and Mercader-Prats (1999) and Lewbel and Pendakur (2008).

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 thus far lead us to three conclusions. First, the model generates a reasonable fit of male hours dynamics around the time of marriage. Second, reverse causality, i.e., selection into marriage on the basis of transitory wage shocks, cannot by itself explain the 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.

6.3 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.

6.3.1 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%.

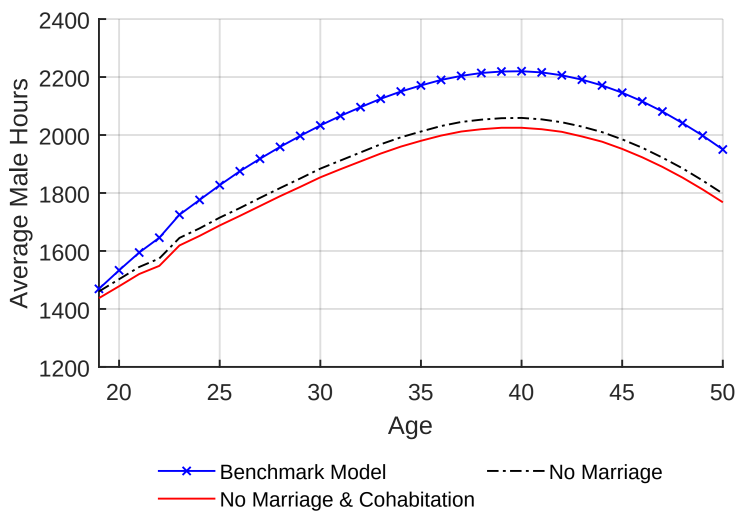
By way of comparison, in Siassi's (2019) framework, the feature most akin to our mouths-to-feed mechanism is a stronger bequest motive for individuals with children.[24] Removing this feature causes his estimate of the proportional earnings (not hours) gap to shrink by 2.3 - 2.8 percentage points (Siassi, 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

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

which we eliminate marriage but not cohabitation. Specifically, we set the probability of marriage 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.

Figure 16: Hours of Work with and without Marriage



Source: Model results are 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.

6.3.2 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.

Table 3: Hours of Work with NLSY79 and NLSY97 Processes

		Change from Benchmark	
Description	Hours	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.

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 1980, 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%, 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%.[25]

The model experiment indicates that the decline in marriage rates can account for much of the decline in hours observed in the data, but not all. In our model, spousal earnings are one

[25] 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.

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 much of the gap is accounted for by an increase in work at the individual level 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 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. A version of the model calibrated 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, selection into marriage based on labor market shocks explains only a small part of the increase in hours around marriage. Instead, the primary driver of the increase in hours is that men internalize the utility that their family members receive from consumption. Counterfactual experiments within the model show that marriage rates are an important determinant of aggregate hours worked by prime age men, and declining marriage rates can explain 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, as 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.

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.

Cowell, Frank A, and Magda Mercader-Prats. 1999. "Equivalence Scales and Inequality." In *Handbook of Income Inequality Measurement*. 405–435. Springer.

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.

Guvenen, Fatih. 2009. "An Empirical Investigation of Labor Income Processes." *Review of Economic Dynamics*, 12(1): 58–79.

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.

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.

Lewbel, Arthur, and Krishna Pendakur. 2008. "Equivalence Scales." *New Palgrave Dictionary of Economics*, 3: 26–29.

Low, Hamish, Costas Meghir, Luigi Pistaferri, and Alessandra Voena. 2017. "Marriage, Labor Supply and the Social Safety Net."

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

A.1 State Average Hours Worked

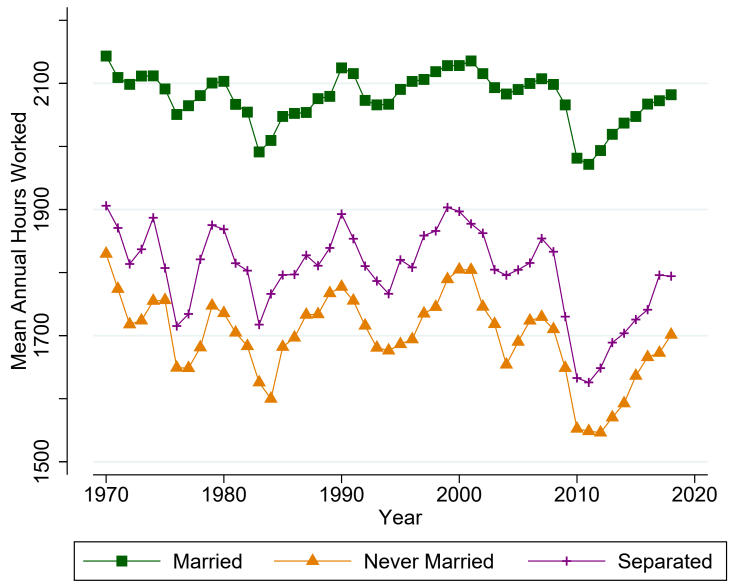
Table 4: Predictors of State Average Male Annual Hours Worked in the CPS

	(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. 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.

A.2 Annual Hours Worked by Marital Status in the CPS ASEC

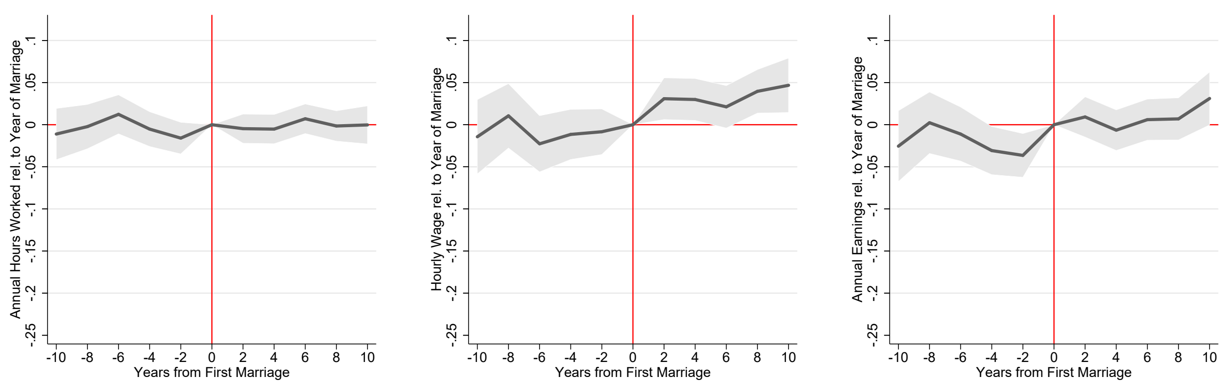
Figure 17: Annual Hours Worked by Marital Status: 1970-2018



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

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

Figure 18: Placebo Test: Labor Market Dynamics in the Years around Marriage

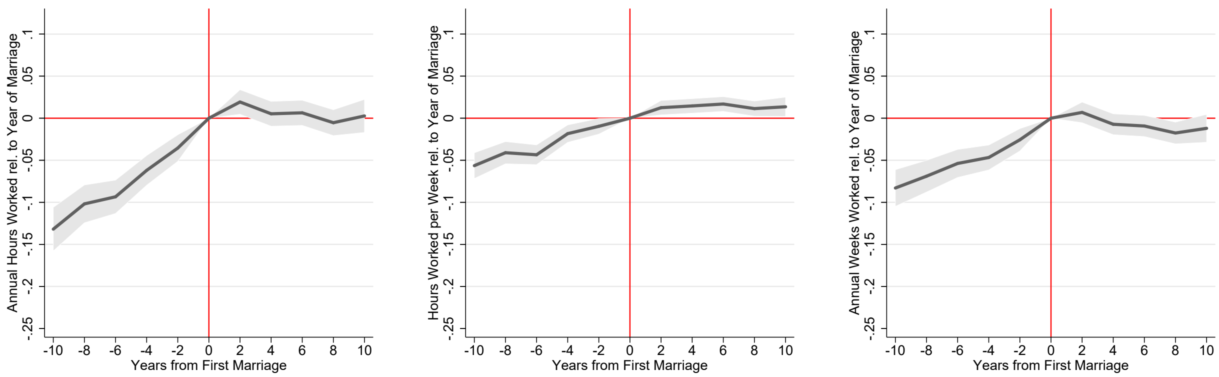


Source: Males ages 19 - 54 in the NLSY79, see text for details. The solid line plots distance-from-marriage coefficients from the individual fixed effects regression equation (1). The shaded region corresponds to 95% confidence intervals. The results in this table correspond to a placebo test of the results in Figure 4, in which age of first marriage was randomly reassigned among individuals in the regression sample.

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

Figure 19: Dynamics of Weekly Hours and Annual Weeks Worked 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 Additional Model Details

B.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:

$$\phi^{an}_{0,0,r,e,j}(yc,1) = \varphi^{n}_{0,r,e,j}, \tag{29a}$$

$$\phi^{an}_{0,0,r,e,j}(0,0) = 1 - \varphi^{n}_{0,r,e,j}, \tag{29b}$$

$$\phi^{an}_{yc,n,r,e,j}(oc,n+1) = \varphi^{a}_{yc,e,j} \cdot \varphi^{n}_{n,r,e,j}, \tag{29c}$$

$$\phi^{an}_{yc,n,r,e,j}(yc,n+1) = (1-\varphi^{a}_{yc,e,j}) \cdot \varphi^{n}_{n,r,e,j}, \tag{29d}$$

$$\phi^{an}_{yc,n,r,e,j}(oc,n) = \varphi^{a}_{yc,e,j} \cdot (1-\varphi^{n}_{n,r,e,j}), \tag{29e}$$

$$\phi^{an}_{yc,n,r,e,j}(yc,n) = (1-\varphi^{a}_{yc,e,j}) \cdot (1-\varphi^{n}_{n,r,e,j}), \tag{29f}$$

$$\phi^{an}_{oc,n,r,e,j}(gc,n) = \varphi^{a}_{oc,e,j}, \tag{29g}$$

$$\phi^{an}_{oc,n,r,e,j}(oc,n) = 1 - \varphi^{a}_{oc,e,j}, \tag{29h}$$

$$\phi^{an}_{gc,n,r,e,j}(gc,n) = 1. \tag{29i}$$

B.2 Relationship Dynamics

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

$$\Pr_j(r_j = en \mid r_{j-1} = sn,\ n_j = 0,\ \tilde{w},\ e,\ \ell) = \phi^{en}_{e,\ell,j}(\tilde{w}), \tag{30a}$$

$$\Pr_j(r_j = sn \mid r_{j-1} = sn,\ n_j = 0,\ \tilde{w},\ e,\ \ell) = 1 - \phi^{en}_{e,\ell,j}(\tilde{w}), \tag{30b}$$

$$\Pr_j(r_j = en \mid r_{j-1} = sn,\ n_j = 1,\ n_{j-1} = 0,\ \tilde{w},\ e,\ \ell) = \phi^{owb}_e + \left(1 - \phi^{owb}_e\right)\phi^{en}_{e,\ell,j}(\tilde{w}), \tag{30c}$$

$$\Pr_j(r_j = sn \mid r_{j-1} = sn,\ n_j = 1,\ n_{j-1} = 0,\ \tilde{w},\ e,\ \ell) = \left(1 - \phi^{owb}_e\right)\left(1 - \phi^{en}_{e,\ell,j}(\tilde{w})\right), \tag{30d}$$

$$\Pr_j(r_j = en \mid r_{j-1} = sn,\ n_{j-1} > 0,\ \tilde{w},\ e,\ \ell) = \phi^{en}_{e,\ell,j}(\tilde{w}), \tag{30e}$$

$$\Pr_j(r_j = sn \mid r_{j-1} = sn,\ n_j > 0\ \tilde{w},\ e,\ \ell) = 1 - \phi^{en}_{e,\ell,j}(\tilde{w}), \tag{30f}$$

and the following probabilities for married and divorced men:

$$\Pr_j(r_j = dv \mid r_{j-1} = mr,\ n,\ e) = \phi^{dv}_{n,e,j}, \tag{31a}$$

$$\Pr_j(r_j = mr \mid r_{j-1} = mr,\ n,\ e) = 1 - \phi^{dv}_{n,e,j}, \tag{31b}$$

$$\Pr_j(r_j = dv \mid r_{j-1} = dv,\ n,\ e) = 1. \tag{31c}$$

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.

B.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.

B.3.1 Single

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

$$V_j^{sn}(e,\ell,k,\tilde{w},a,n) = \max_{c,k',h} \Bigg\{ u_{f,e,\ell,j}(c,h) \,+\, \beta \int_{\tilde{w}'} \Bigg(\sum_{(a',n')} \phi^{an}_{a,n,sn,e,j}(a',n') \tag{32}$$

$$\times \Big[\Pr_j\left(sn\,|\,sn,n',n,\tilde{w}',e,\ell\right) V^{sn}_{j+1}\left(e,\ell,k',\tilde{w}',a',n'\right) + \Pr_j\left(en\,|\,sn,n',n,\tilde{w}',e,\ell\right)$$

$$\times \int_{\tilde{s}}\int_{q'} V^{en}_{j+1}\left(e,\ell,k',\tilde{w}',a',n',\tilde{s},q'\right) dF^q(q')\, dF^s_e(\tilde{s}\,|\,\tilde{w}') \Big] \Bigg)\, dF^w_{e,j}(\tilde{w}'\,|\,\tilde{w}) \Bigg\},$$

$$s.t. \quad c + k' \leq Rk + wh^{1+\zeta}(1 - n\delta_a) - T(wh^{1+\zeta}, k; f), \tag{33}$$

$$\log w = \alpha^w_{e,\ell,j} + \tilde{w}, \tag{34}$$

$$k' \geq k_{min}. \tag{35}$$

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.

B.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

$$V_j^{en\text{-}n}(e,\ell,k,\tilde{w},a,n,\tilde{s},q) = \max_{c,k',h} \Bigg\{ u_{f,e,\ell,j}(c,h) \,+\, \beta \int_{\tilde{w}'}\int_{q'} \Bigg(\sum_{(a',n')} \phi^{an}_{a,n,en,e,j}(a',n') \tag{36}$$

$$\times \Big[\Pr\left(en\text{-}n\,|\,en\text{-}n,n',n,\tilde{w}',e\right) V^{en\text{-}n}_{j+1}\left(e,\ell,k',\tilde{w}',a',n',\tilde{s},q'\right)$$

$$+ \Pr\left(mr\,|\,en\text{-}n,n',n,\tilde{w}',e\right) V^{mr}_{j+1}\left(e,\ell,k',\tilde{w}',a',n',\tilde{s},q'\right) \Big] \Bigg)\, dF^q(q')\, dF^w_{e,j}(\tilde{w}'\,|\,\tilde{w}) \Bigg\}$$

$s.t.$ equations (33)-(35).

The Bellman equation for a working-age man who is engaged and cohabiting, $r = en\text{-}c$, is identical to the previous Bellman equation except that: (i) the budget constraint includes spousal earnings and the taxes paid on these earnings; (ii) 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

$$V_j^{en\text{-}c}(e,\ell,k,\tilde{w},a,n,\tilde{s},q) = \max_{c,k',h} \Bigg\{ u_{f,e,\ell,j}(c,h) \,+\, \beta \int_{\tilde{w}'} \int_{q'} \Bigg(\sum_{(a',n')} \phi^{an}_{a,n,en,e,j}(a',n') \tag{37}$$

$$\times \Big[\Pr\big(en\text{-}c \,|\, en\text{-}c, n', n, \tilde{w}', e\big)\, V_{j+1}^{en\text{-}c}\big(e,\ell,k',\tilde{w}',a',n',\tilde{s},q'\big)$$

$$+ \Pr\big(mr \,|\, en\text{-}c, n', n, \tilde{w}', e\big)\, V_{j+1}^{mr}\big(e,\ell,k',\tilde{w}',a',n',\tilde{s},q'\big)\Big]\Bigg)\, dF^q(q')\, dF^w_{e,j}(\tilde{w}' \,|\, \tilde{w})\Bigg\},$$

$$s.t. \quad c + k' \leq Rk + wh^{1+\zeta} + se(1-n\chi_a) - T(wh^{1+\zeta}, k; f^{en\text{-}c}) - T(se, 0; f^{en\text{-}c,sp}), \tag{38}$$

$$se \text{ satisfies equations (8)-(11)}, \tag{39}$$

equations (34)-(35).

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.

B.3.3 Married

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

$$V_j^{mr}(e,\ell,k,\tilde{w},a,n,\tilde{s},q) = \max_{c,k',h} \Bigg\{ u_{f,e,\ell,j}(c,h) \,+\, \beta \int_{\tilde{w}'} \Bigg(\sum_{(a',n')} \phi^{an}_{a,n,mr,e,j}(a',n') \tag{40}$$

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

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

$$s.t. \quad c + k' \leq Rk + wh^{1+\zeta} + se(1-n\chi_a) - T(wh^{1+\zeta} + se, k; f), \tag{41}$$

equations (34)-(35) and (39).

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.

B.3.4 Divorced

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

$$V_j^{dv}(e,\ell,k,\tilde{w},a,n) = \max_{c,k',h} \left\{ u_{f,e,\ell,j}(c,h) \right. \tag{42}$$

$$\left. + \beta \int_{\tilde{w}'} \left(\sum_{(a')} \varphi_{a,e,j}^{a}(a') V_{j+1}^{dv}\left(e,\ell,k',\tilde{w}',a',n\right) \right) dF_{e,j}^{w}(\tilde{w}' \,|\, \tilde{w}) \right\},$$

$$s.t. \quad \text{equations (33)-(35).}$$

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.

B.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:

$$V_j^{ret}(e,k,r) = \max_{c,k'} \quad u_{f,e,j}(c,0) + \beta V_{j+1}^{ret}(e,k',r), \tag{43}$$

$$s.t. \quad c + k' \leq Rk + b_{1,e} + b_{2,e} \mathbb{1}_{r \in \{mr,en\text{-}c\}} - T\left(b_{1,e} + b_{2,e} \mathbb{1}_{r \in \{mr,en\text{-}c\}}, k; f\right), \tag{44}$$

$$\text{equation (35)},$$

$$V_J \equiv 0. \tag{45}$$

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

C Parameters Taken from Other Studies

Table 5 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 $J = 80$.

Table 5: 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) (see main text)
Frisch elasticity	ξ	0.75	Various (see main text)
Consumption utility shifter, married or cohabiting	$N_{f:r\in\{mr,en\text{-}c\}}$	2	See main text
Consumption utility shifter not married or cohabiting	$N_{f:r\notin\{mr,en\text{-}c\}}$	1	See main text
Equivalence scale	η_f	eqn. (46)	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.

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 5.

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

$$\eta_f = \begin{cases} 1, & \text{if } r \notin \{mr, en\text{-}c\}, \\ (2 + \mathbb{1}_{a<gc} \cdot 0.7n)^{0.7}, & \text{if } r \in \{mr, en\text{-}c\}. \end{cases} \tag{46}$$

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

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.[26]

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.[27]

D 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

[26] 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 $(0.2873 \cdot 0.094)/1.6 = 0.017$.

[27] 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.)

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.

D.1 Wages

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

$$\alpha^w_{e,\ell,j} = \alpha^w_{e,j} + \tilde{\Delta}^w_i, \tag{47}$$

$$\tilde{\Delta}^w_i = \begin{cases} -\Delta^w, & \ell_i = 1, \\ \frac{p_1}{1-p_1}\Delta^w, & \ell_i = 2. \end{cases} \tag{48}$$

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

$$\left(\sigma^{\Delta w}\right)^2 = \frac{p_1}{1-p_1}(\Delta^w)^2. \tag{49}$$

We now proceed in a way very similar to that used by French (2005). We begin by running an individual fixed effects regression on log wages (after adjusting for hours). The coefficients from this regression, 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

$$w^*_{i,j} = \tilde{\Delta}^w_i + \tilde{w}_{i,j} + \bar{w}_{i,j}, \tag{50a}$$

$$\tilde{w}_{i,j} = \rho^w_e \tilde{w}_{i,j-1} + \varepsilon^w_{i,j}, \tag{50b}$$

$$\varepsilon^w_{i,j} \overset{iid}{\sim} N(0, \sigma^\varepsilon_e), \tag{50c}$$

$$\tilde{w}_{i,0} \sim N(0, \sigma^{\tilde{w}_0}_e), \tag{50d}$$

$$\tilde{w}_{i,j} \perp\!\!\!\perp \tilde{\Delta}^w_i, \ \forall i, j, \tag{50e}$$

$$\bar{w}_{i,j} \overset{iid}{\sim} N(0, \sigma^{\bar{w}}_e), \tag{50f}$$

$$\bar{w}_{i,j} \perp\!\!\!\perp \tilde{\Delta}^w_i, \tilde{w}_{i,j+s}, \ \forall i, j, s. \tag{50g}$$

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

$$\mathrm{var}_e(w^*_{i,j}) = \left(\sigma^{\Delta w}\right)^2 + \mathrm{var}_e(\tilde{w}_{i,j}) + (\sigma^{\bar{w}}_e)^2, \tag{51}$$

$$\mathrm{cov}_e(w^*_{i,j}, w^*_{i,j+k}) = \left(\sigma^{\Delta w}\right)^2 + (\rho^w_e)^k \mathrm{var}_e(\tilde{w}_{i,j}),\ k > 1. \tag{52}$$

We can then use equation (52) to back out the autocorrelation ρ^w_e:

$$\widehat{\rho}^w_e = \sqrt{\frac{\mathrm{cov}_e(\widehat{w}^*_{i,j}, \widehat{w}^*_{i,j+4}) - \left(\sigma^{\Delta w}\right)^2}{\mathrm{cov}_e(\widehat{w}^*_{i,j}, \widehat{w}^*_{i,j+2}) - \left(\sigma^{\Delta w}\right)^2}}, \tag{53}$$

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 (52) 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

$$\widehat{\mathrm{var}}_e(\tilde{w}_{i,j}) = \left[\mathrm{cov}_e(\widehat{w}^*_{i,j}, \widehat{w}^*_{i,j+2}) - \left(\sigma^{\Delta w}\right)^2\right] \frac{1}{(\widehat{\rho}^w_e)^2}, \tag{54}$$

Under the error components model (50), the variances of $\tilde{w}_{i,j}$, $\mathrm{var}_e(\tilde{w}_{i,j}), j = 0, 1, ..., J$ are:

$$\mathrm{var}_e(\tilde{w}_{i,0}) = (\sigma^{\tilde{w}_0}_e)^2, \tag{55a}$$

$$\mathrm{var}_e(\tilde{w}_{i,1}) = (\rho^w_e)^2 \mathrm{var}(\tilde{w}_{i,0}) + (\sigma^{\varepsilon}_e)^2, \tag{55b}$$

$$\mathrm{var}_e(\tilde{w}_{i,j}) = (\rho^w_e)^2 \mathrm{var}(\tilde{w}_{i,j-1}) + (\sigma^{\varepsilon}_e)^2, \quad j > 1. \tag{55c}$$

We can then estimate $\widehat{\sigma}^{\varepsilon}_e$ and $\widehat{\sigma}^{\tilde{w}_0}_e$ by fitting the sequence of variances implied by equation (55) to the sequence implied by equation (54), $\left\{\widehat{\mathrm{var}}_e(\tilde{w}_{i,j})\right\}_{j=0}^J$.

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 modelled 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.

D.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 $\phi^{en}_{e,j}$, so that the probability of remaining single is $\phi^{sn}_{e,j} = 1 - \phi^{en}_{e,j}$. To simplify the notation, we will ignore the dependence of $\phi^{sn}_{e,j}(\tilde{w})$ on p_1 and Δ^w (described immediately above); these parameters would otherwise appear as subscripts on ϕ^{sn}.

Type effects enter as

$$\phi^{sn}_{e,\ell,j}(\tilde{w}) = \begin{cases} \phi^{sn}_{e,j}(\tilde{w})\big(1+\Delta^{sn}\big), & \ell = 1, \\ \phi^{sn}_{e,j}(\tilde{w})\big(1-\frac{p_1}{1-p_1}\Delta^{sn}\big), & \ell = 2. \end{cases} \tag{56}$$

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

The type effects in equation (56) 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:

$$\pi_{e,n^-,n,j}(\tilde{w}) = \begin{cases} (1-\phi^{owb}_e)\phi^{sn}_{e,j}(\tilde{w}) = (1-\phi^{owb})(1-\phi^{en}_{e,j}(\tilde{w})), & \text{if } n^- = 0 \text{ and } n > 0, \\ \phi^{sn}_{e,n^-,n,j}(\tilde{w}) = 1-\phi^{en}_{e,j}(\tilde{w}), & \text{otherwise,} \end{cases} \tag{57}$$

Applying equation (56) yields the type-specific probabilities

$$\pi^*_{e,n^-,n,j,1}(\tilde{w}) = \pi_{e,n^-,n,j}(\tilde{w})(1+\Delta^{sn}), \tag{58}$$

$$\pi^*_{e,n^-,n,j,2}(\tilde{w}) = \pi_{e,n^-,n,j}(\tilde{w})\big(1-\frac{p_1}{1-p_1}\Delta^{sn}\big). \tag{59}$$

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 (58) and (59), 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

$$\begin{aligned} s^*_{j+1} &= s^*_{j+1,1} + s^*_{j+1,2} \\ &= s_{j,1}\pi_j(1+\Delta^{sn}_{j,}) + s_{j,2}\pi_j\big(1 - \tfrac{p_1}{1-p_1}\Delta^{sn}_j\big) \\ &= s_j\Big(\lambda_j\pi_j(1+\Delta^{sn}_j) + (1-\lambda_j)\pi_j\big(1 - \tfrac{p_1}{1-p_1}\Delta^{sn}_j\big)\Big) \\ &= s_j\pi_j\Big(1+\Delta^{sn}_{j,}\big[\lambda_j - (1-\lambda_j)\tfrac{p_1}{1-p_1}\big]\Big) \\ &:= s_j\pi_j\delta_j^{-1}. \end{aligned} \tag{60}$$

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 (56), find the type-specific transition probabilities $\pi^*_{j,1} = \pi_j\big(1+\Delta^{sn}_j\big)$ and $\pi^*_{j,2} = \pi_j\Big(1 - \frac{p_1}{1-p_1}\Delta^{sn}_j\Big)$.

3. Find the adjustment
$$\delta_j = \Big(1+\Delta^{sn}_j\big[\lambda_j - \big(1-\lambda_j\big)\tfrac{p_1}{1-p_1}\big]\Big)^{-1}. \tag{61}$$

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
$$\lambda_{j+1} = \frac{\pi^{**}_{j,1}\lambda_j}{\pi^{**}_{j,1}\lambda_j + \pi^{**}_{j,2}\big(1-\lambda_j\big)}. \tag{62}$$
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:

$$sh_{sn,a,n,1} = sh_{sn,a,n}(1+\Delta^{s0}), \tag{63}$$

$$sh_{sn,a,n,2} = sh_{sn,a,n}\big(1-\tfrac{p_1}{1-p_1}\Delta^{s0}\big), \tag{64}$$

$$\frac{sh_{m,a,n,2}}{sh_{m,a,n,1}} = \frac{sh_{en,a,n,2}}{sh_{en,a,n,1}} = \frac{sh_{dv,a,n,2}}{sh_{dv,a,n,1}}. \tag{65}$$

To implement equation (65), 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 (63) that

$$\begin{aligned} sh_{r,a,n,1} &= sh_{r,a,n}\frac{1-\varsigma_{a,n}*(1+\Delta^{s0}_{a,n})}{1-\varsigma_{a,n}} \\ &= sh_{r,a,n}\Big(1-\Delta^{s0}_{a,n}\tfrac{\varsigma_{a,n}}{1-\varsigma_{a,n}}\Big),\ \ r\in\{en,m,dv\}. \end{aligned} \tag{66}$$

It likewise follows from equation (63) that

$$sh_{r,a,n,2} = sh_{r,a,n}\Big(1+\Delta^{s0}_{a,n}\Big(\tfrac{p_1}{1-p_1}\Big)\tfrac{\varsigma_{a,n}}{1-\varsigma_{a,n}}\Big),\ \ r\in\{en,m,dv\}. \tag{67}$$

D.3 Preferences

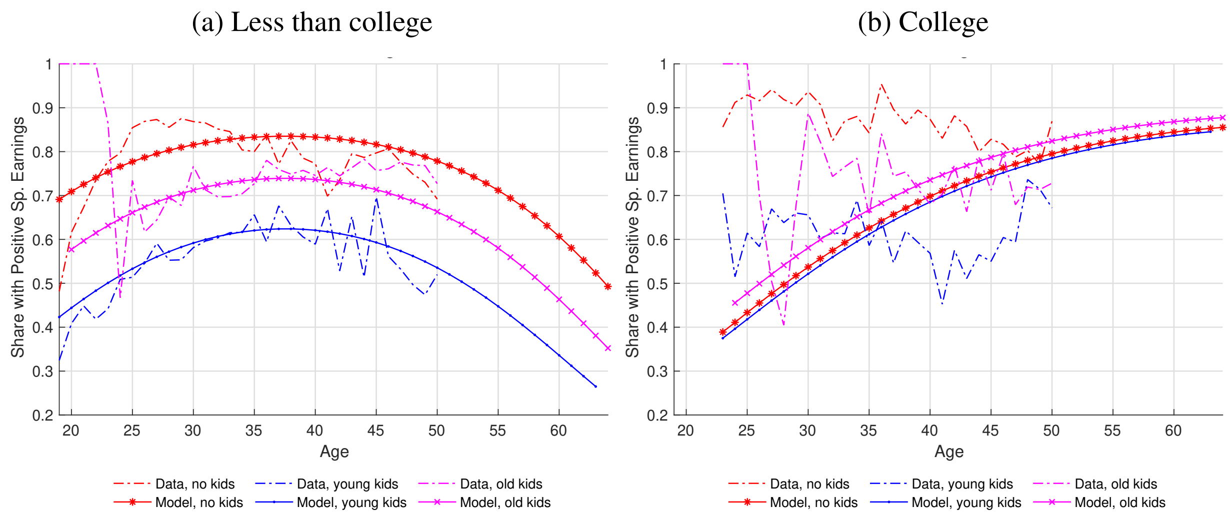
The third way the two types differ is in their disutility from work, $\psi_{e,j}$. Previously, the disutility of work was $\psi_{e,j} = \psi_{e,0}(1+\psi_{e,1}\cdot j+\psi_{e,2}\cdot j^2)$. The type-specific adjustments are:

$$\psi_{e,\ell,0} = \begin{cases} \psi_{e,0}\big(1+\Delta^{\psi}\big), & \ell=1, \\ \psi_{e,0}\big(1-\tfrac{p_1}{1-p_1}\Delta^{\psi}\big), & \ell=2, \end{cases} \tag{68}$$

$$\psi_{e,2,1} = \psi_{e,1,1} = \psi_{e,1},\ \ \psi_{e,2,2} = \psi_{e,1,2} = \psi_{e,2}. \tag{69}$$

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.

Figure 20: 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.

E Model Fit of Spousal Employment and Earnings

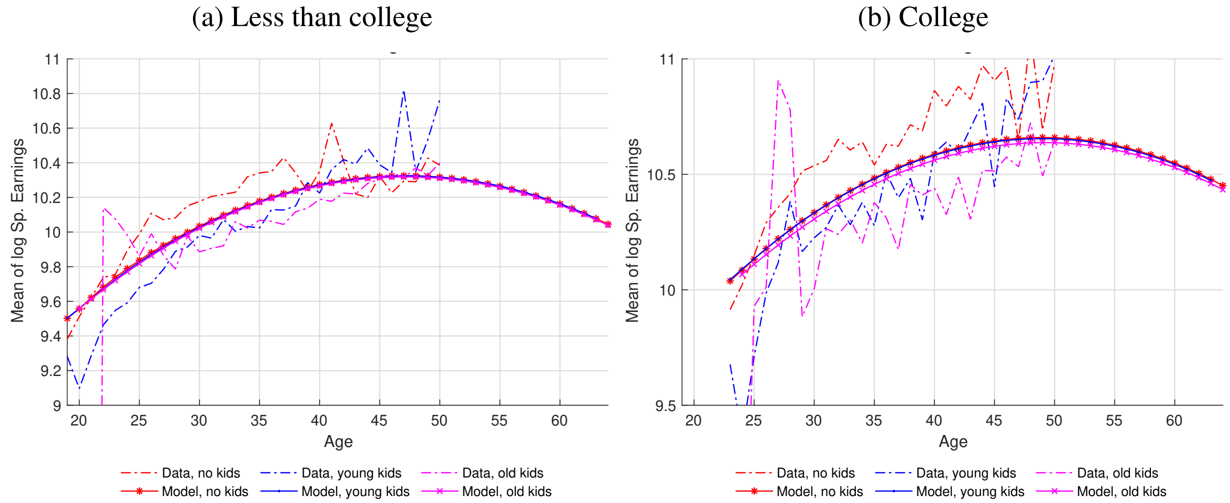
The parameters of the spousal earnings process are set to match the age profiles of spousal employment, the mean and standard deviation of log earnings among working spouses, and the correlation of spousal earnings and male wages. Because the NLSY79 does not provide earnings information for unmarried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

Figure 20 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 21 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 6 shows the standard deviation of these earnings, again conditional 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.

Our final target is the correlation between male wages and spousal earnings, conditional on

Figure 21: 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.

Table 6: 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.

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 6 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.

F Estimating the Dynamics of Family Structure

F.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.

Table 7: 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

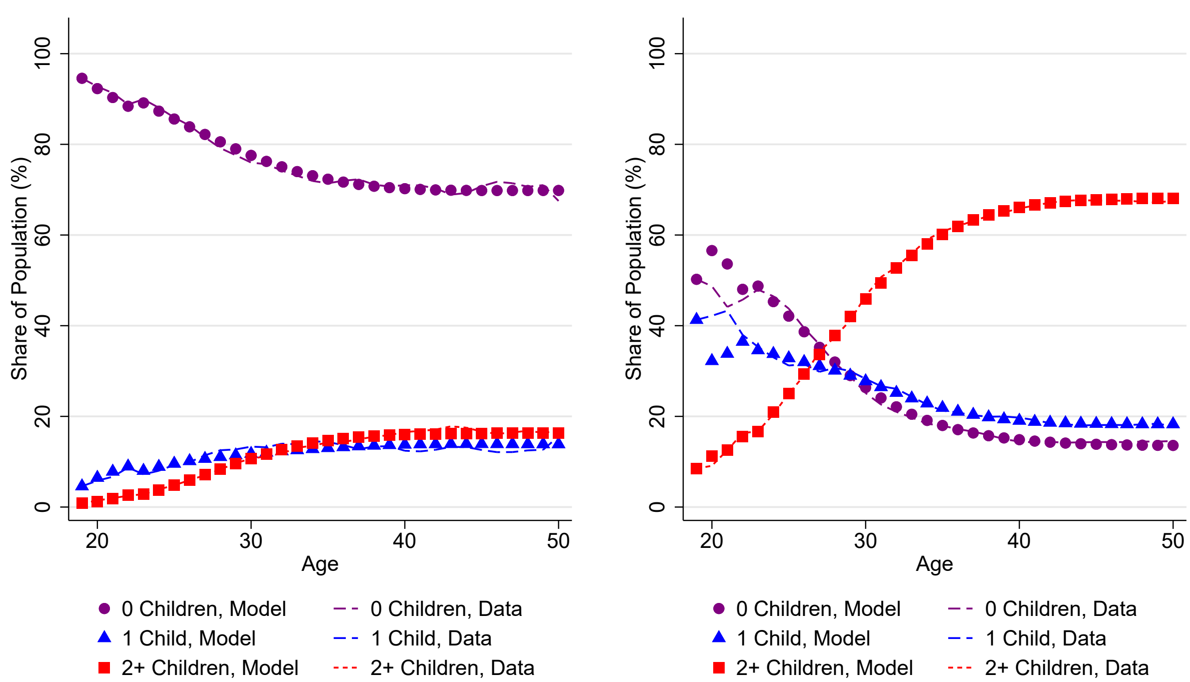
Note: Superscripts are used to distinguish parameters, while subscripts are used to distinguish dependencies. All parameters with the age subscript j utilize a quadratic in age. See section 4.3.3 and Appendix F for details.

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 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 transtioning 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.

Figure 22: 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 F for details.

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 F.2 for details. Table 7 shows the resulting parameter estimates. Figures 6a and 6b in the main text and Figure 22 in this Appendix show model fits.

Table 8: 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)
$\text{Age}^2/100$	−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.

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

Table 8 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 accomodates 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.

G Third-stage Parameters

Table 9 presents the third-stage parameter estimates. As discussed in Section 4.4 and Appendix D, 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.

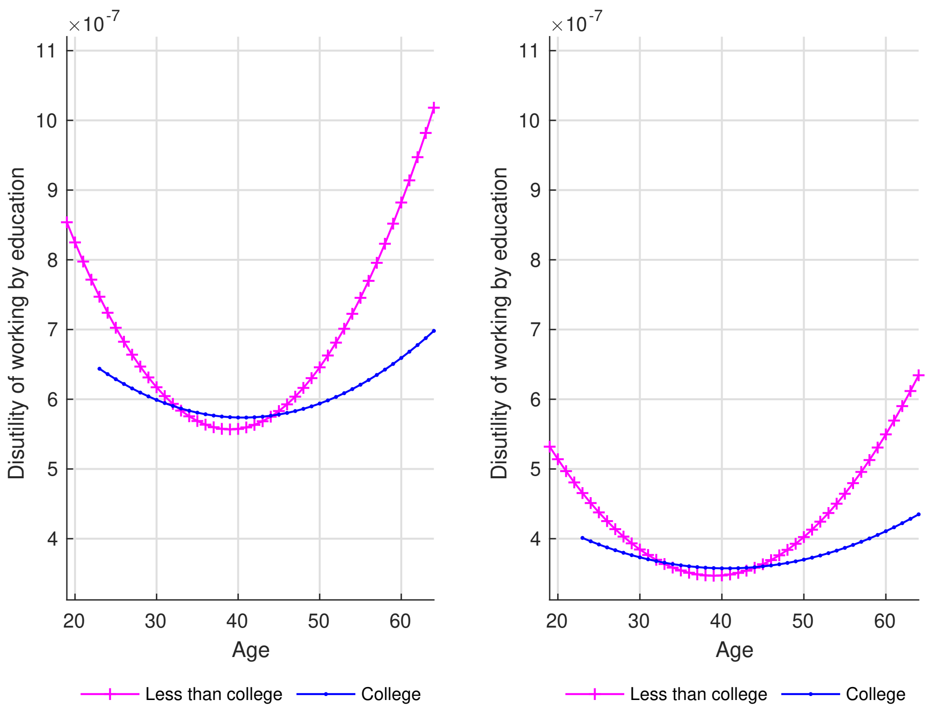
Table 9: Third-stage Parameters

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 D 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 23 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 23: Work Disutility by Age and Education



Source: Author calculations. See Section 4.4 for details.