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Abstract

Gender differences in preferences for parental leave contribute to the early-career growth in the gender wage gap among highly educated millennial Americans. Estimating a hedonic job-search model, I show that, while both men and women experience wage growth by entering firms offering better pay and benefits, the wage gap increases as women accept lower wages upon receiving job offers from employers providing parental leave. The wage-gap growth could decline by 60% if preferences for paid parental leave did not differ by gender. It could also decline if mandating and subsidizing the provision of paid leave muted workers' leave-wage trade-off.

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1 Introduction

Paid and unpaid parental leave are seldom available to workers in the United States. Although the Family and Medical Leave Act (FMLA) of 1993 guarantees up to 12 weeks of unpaid parental leave to eligible workers in large establishments, the Federal Employee Paid Leave Act (FEPLA) of 2019 provides paid leave to eligible federal workers, and several US states implemented paid family leave laws, parental leave coverage remains scattered, conditional on requirements that young workers may not satisfy, and unequally distributed. As of 2023, only 27% of civilian American workers had access to paid family leave (BLS, 2023), while the FMLA requirements exclude from unpaid leave coverage at least 40% of the US workforce (Brown, Herr, Roy, & Klerman, 2020).[1]

Anecdotal evidence suggests that only some large US employers are primarily filling gaps in parental leave coverage (Cain Miller, 2018b) by providing this benefit in an effort to attract and retain employees (Cain Miller, 2018a; Michelson, 2021). In fact, in a recent contribution, Goldin, Kerr, and Olivetti (2020) observe that firms are especially likely to provide paid parental leave when they employ a sufficiently large number of individuals who made substantial human capital investments when young and who are, therefore, more likely to return to work following a leave period.

In this paper I ask how the provision of employer-sponsored paid and unpaid parental leave, and lack thereof, affects the gender wage gap among college graduate, millennial American workers, and its increase during workers' early careers.[2]

I focus on college graduate workers for several reasons. First, the secular increase in women's college graduation rate was associated with longer-term expectations of labor market commitment among young women (Goldin, Katz, & Kuziemko, 2006) and, although women with bachelor's and postgraduate degrees have progressively postponed their fertility to avoid overlaps between education completion, early-career outcomes, and family formation (Goldin, 2004; Nitsche & Brückner, 2020), the share of women with at least one biological child by age 44 has been rising in recent years (Nitsche & Brückner, 2020). Second, low labor supply and career interruptions can be especially detrimental

[1]Details on paid and unpaid parental leave policies in the United States can be found in Section A in the Online Appendix.

[2]The term "millennial" refers to the cohort born between 1981 and 1996.

for women's career progression in relatively high-pay occupations that mostly employ college graduate workers (Bertrand, Goldin, & Katz, 2010; Gicheva, 2013; Goldin, 2014).

These facts suggest that, if college graduate women in recent cohorts are increasingly committed to both achieving fulfilling, uninterrupted careers and having children (Goldin, 2024), they may make choices that enable them to reconcile these ambitions since the very start of their careers, thus selecting in jobs and establishments providing benefits, such as paid parental leave, to facilitate the reconciliation of career and family. Yet, in a scenario where parental leave is relatively scarce, young women may disproportionately search for employers offering this benefit and be willing to accept lower wages in exchange for its provision. While young men may value the provision of parental leave as well when choosing jobs, if parental leave is more salient for young women, their stronger willingness to pay (preferences) for this benefit may contribute to the increase in the gender wage gap observed during workers' early careers, as a result of workers' search for (and entry in) firms that offer parental leave.

To the best of my knowledge, Goldin and Katz (2011) were the first who noted that any "aspects of workplace family friendliness" (Goldin and Katz, 2011, p. 47), including schedule or hours flexibility, employer-sponsored child care and paid leave could be conceptualized as workplace amenities that workers in the high end of the labor market may be willing to pay for and that could, therefore, contribute to gender pay differences if their provision is more salient for women than for men.

To answer my research question, I use restricted-access geocoded data from the National Longitudinal Survey of Youth 1997 (NLSY97) and study the first six years of labor market experience (early careers) of millennial American college graduates born between 1980 and 1984 and entering the US labor market between 2000 and 2011.

First, I provide reduced-form evidence suggesting that the search for parental leave may affect women's job-search outcomes and the early-career growth in the gender wage gap. Specifically, I document that the gender wage gap increases during early careers, primarily among workers who change employer at least once during this period. This fact is not driven by the potential wage losses of women who have at least one child during their early career. Furthermore, I show that the growth in the gender wage gap among workers who change at least one job during their early career can be entirely explained by the stronger wage gains that men obtain following their first job change. Finally, I show that the likelihood of being offered valuable benefits such as paid and unpaid leave increases upon changing employer, and that the availability of both paid and unpaid parental leave significantly reduces the chances of undergoing a job change for women, while only the availability of paid leave reduces men's job change probability.

Second, I use a hedonic job-search model to estimate men's and women's willingness to pay for paid and unpaid parental leave, quantify their impact on the early-career growth in the wage gap, and study the effects of a policy mandating and subsidizing the provision of paid parental leave.

The model builds on the seminal contribution of Bonhomme and Jolivet (2009). It is a random search model with on-the-job search where, in any given month, workers may receive job offers consisting of a wage and a set of benefits and work arrangements. Upon receiving a job offer, employed workers decide whether to accept it by comparing its implied utility with the utility they obtain at their current job. Workers' utility depends on the wage, work arrangements and benefits that employers offer. Employers' wage offers depend on whether benefits are provided, on workers' skill level and occupation, and can be heterogeneous depending on whether the employer is located in a state that implemented paid parental leave laws or not.

The model estimation identifies workers' willingness to pay for benefits using the wage-benefits outcomes of employed workers' job-to-job transitions, conditional on search frictions, identified by the frequency of different labor market transitions, and on the properties of the distribution of wage offers received by workers, identified by the wage-benefits outcomes of previously unemployed workers. Conditional on search frictions and job offers, this method infers stronger preferences for a benefit the lower is the average wage accepted by workers entering firms offering it, compared to their previous wage.

The Bonhomme and Jolivet (2009) revealed-preferences approach used to identify preferences through job-to-job transitions, conditional on estimated search frictions and wage offers, has several important features. First, it overcomes the biases affecting estimators of preferences based on the cross-sectional correlation between wages and amenities among employed workers.[3] Second, it reduces concerns that unobserved gender differences in

[3]Section B in the Online Appendix contains a discussion of econometric challenges in identifying individual preferences for non-wage benefits.

wage offers or in search frictions may bias gender differences in estimates of willingness to pay for benefits such as paid or unpaid parental leave. In fact, if women are offered lower wages or face stronger search frictions compared to men, and these factors are unaccounted for, estimates of women's willingness to pay for benefits may be misleading. Young women may be offered lower wages if employers expect them to accumulate human capital more slowly than men (Amano-Patiño, Baron, & Xiao, 2020; Xiao, 2021). Furthermore, gender differences in wage offers may increase in firms offering parental leave, if employers expect women to take up leave more often and for longer periods (Olivetti & Petrongolo, 2017). Also, women may limit their job search due to a stronger willingness to trade-off commuting time for wages (Le Barbanchon, Rathelot, & Roulet, 2021) or to stronger implicit costs of commuting (Caldwell & Danieli, 2024) which might trigger monopsonistic wage discrimination (Manning, 2003) and result in lower wage offers. In addition, if search behavior differs across genders (Bowlus, 1997; Cortés, Pan, Pilossoph, Reuben, & Zafar, 2023), or discrimination in hires and layoffs exists (Egan, Matvos, & Seru, 2022), women may also face stronger search frictions and receive fewer job offers than men. All these factors can explain why women experience lower wage growth, compared to men, as they change employer, thus potentially biasing upward the estimates of women's willingness to pay for benefits, if not accounted for.

The model estimation can also separately identify workers' preferences for different benefits and job characteristics. In my preferred specification, I estimate men's and women's willingness to pay for paid and unpaid parental leave, and let preferences for schedule flexibility and long work hours to also differ by gender. This approach mitigates potential bias in estimated gender differences in preferences for parental leave. Such bias could arise if women have stronger preferences for employers offering part-time work (Bowlus & Grogan, 2009; Liu, 2016) and flexible work arrangements (Mas & Pallais, 2017; Wiswall & Zafar, 2018; Xiao, 2021), who may also provide other family-friendly benefits. Similarly, the bias could arise if men disproportionately select into high-pay, long-hours jobs (Bertrand, Goldin, & Katz, 2010; Cortés & Pan, 2019; Gicheva, 2013; Goldin, 2014) where access to paid parental leave may be less common.

The main estimation results show that both young women and men highly value paid and unpaid parental leave, but women are willing to pay substantially more than men in exchange for their provision. Concerning paid parental leave, I find that a woman earning the same wage as a comparable man in an establishment where paid parental leave is not available, would accept 44% of the wage that the man would accept to enter a workplace where the benefit is available. Regarding unpaid leave, a woman earning as much as a comparable man in an establishment where the benefit is not available would accept 89% of the wage that the man would accept to enter a firm that offers the benefit. Even though workers, and most prominently women, are willing to pay for the provision of parental leave, I estimate that firms providing these benefits offer to both men and women wages at least as high as firms where paid and unpaid leave are not available.[4] Crucially, a substantial gender gap in workers' valuation of paid parental leave also exists between all men and women who do not have children during their early careers.

These results are consistent with the hypothesis that, upon entering the labor market, both men and women search for profitable employment relationships, and experience wage growth as they progressively enter better jobs, the latter offering higher wages (Topel & Ward, 1992) and more valuable benefits (Hwang, Mortensen, & Reed, 1998; Sockin, 2022). Due to their stronger willingness to pay for parental leave, however, young women accept lower wages compared to men, upon receiving job offers from employers who provide it.

Consistent with this interpretation, counterfactual analyses show that the early-career growth in the gender wage gap would drop by 64% if willingness to pay for paid parental leave did not differ across genders. A decline in the early-career growth in the gender wage gap may also occur if a policy mandating and subsidizing the provision of paid parental leave muted the effect of preferences for this benefit on accepted wages. The widespread availability of paid parental leave could lessen workers' need to trade off wages for leave provision, thus reducing the gap in accepted wages between men and women entering leave-providing firms, and fostering women's early-career wage growth.

This paper is rooted in the academic discussion regarding paid parental leave provision in the United States. The vast literature on the topic highlighted the varied effects of leave policies. Regarding the impact of leave provision on women's labor market outcomes, some contributions documented the positive effects on post-childbirth women's job continuity of policies granting relatively short paid leave periods (Baum & Ruhm,

[4] Women are offered lower wages in establishments where paid parental leave is available than in establishments where paid leave is not provided. The estimated parameter, however, is not statistically significant.

2016; Byker, 2016; Rossin-Slater, 2018; Rossin-Slater, Ruhm, & Waldfogel, 2013), and the non-adverse effect of a higher wage replacement rate (Bana, Bedard, & Rossin-Slater, 2020), whereas recent work on the long-term effects of the 2004 California paid family leave policy found that it did not increase mothers' employment or earnings while, in fact, exacerbating child penalties especially for women at the lower end of the wage distribution (Bailey, Byker, Patel, & Ramnath, in press).[5]

If the potential effects of parental leave policies are numerous, nuanced, and dependent on institutional factors and on the degree of competition in different labor markets (Olivetti & Petrongolo, 2017), the absence of policies mandating parental leave can also be consequential. Blau and Kahn (2013) highlighted that the lack of family-friendly policies might have contributed to the stagnation in female labor supply growth over last three decades in the United States. In this paper, I provide evidence suggesting that the decentralization of parental leave provision may contribute to the early-career growth in the gender wage gap. To the extent that the choice to provide parental leave is left to employers and costs are not subsidized, only some employers for whom the costs of leave provision are affordable will provide it (Goldin, Kerr, & Olivetti, 2020). Consequently, workers for whom the availability of parental leave is more salient will disproportionately pay for it by accepting lower wages in exchange for its provision. This fact can especially penalize young women who are strongly attached to the labor market, for whom the availability of parental leave may represent a form of employment insurance and career continuity in the event of a childbirth.[6]

[5] It is also still a matter of debate whether policies extending parental leave coverage and duration have small (Bartel, Rossin-Slater, Ruhm, Slopen, & Waldfogel, 2023a) or large (Ginja, Karimi, & Xiao, 2023) effects on employers' costs, and whether they have positive (Bartel, Rossin-Slater, Ruhm, Slopen, & Waldfogel, 2023b) effects, or null or negative effects (Dustmann & Schönberg, 2012) on parents and children's health and human-capital outcomes, while recent evidence suggest that paid parental leave may entail profitability gains for firms (Bennett, Erel, Stern, & Wang, 2023). A vast literature studied the impact of family leave policies on workers' outcomes in European and OECD countries. Among the most recent contributions, Kleven, Landais, Posch, Steinhauer, and Zweimüller (2024) show that expansions of parental leave coverage during the 20th century did not contribute to gender convergence in Austria; Stearns (2018) finds that expanded access to paid parental leave in Great Britain had positive short-run effects on mothers' labor supply but no discernible effects on earnings. Olivetti and Petrongolo (2017) and Bartel, Rossin-Slater, Ruhm, Slopen, and Waldfogel (2023b) provide comprehensive reviews of the literature on the impacts of family leave policies.

[6] This hypothesis is consistent with evidence provided by Goldin and Mitchell (2017) who, using data from the Survey of Income and Program Participation from the 1990s, show that women tho had access to paid parental leave around childbirth were substantially more likely to participate in the labor force prior to childbirth compared to women who quit their job around childbirth.

Studying the impact of workers' willingness to pay for parental leave on the gender wage gap, this paper belongs to the growing body of work that analyzes the impact of preferences for non-wage job attributes (Flabbi & Moro, 2012; Hotz, Johansson, & Karimi, 2018; Liu, 2016; Mas & Pallais, 2017; Morchio & Moser, 2024; Wiswall & Zafar, 2018; Xiao, 2021), of location and commuting (Caldwell & Danieli, 2024; Le Barbanchon, Rathelot, & Roulet, 2021), and of firm heterogeneity (Barth, Kerr, & Olivetti, 2021; Card, Cardoso, Heining, & Kline, 2018; Card, Cardoso, & Kline, 2016) on wages and on gender inequality in labor market outcomes. The recent contribution by Morchio and Moser (2024) is especially close to this work. Exploiting rich linked employer-employee data from Brazil to identify the parameters of a hedonic search model, the authors find that part of of the gender earnings gap among Brazilian workers can be attributed to women's stronger valuation of non-wage benefits. Consistent with the finding I obtain on young, college graduate, American workers, the authors also show that, while both men and women in Brazil appear to value the possibility of accessing paid parental leave, women tend to select in firms with more generous leave policies.

Finally, this paper provides a comprehensive analysis of the evolution of the gender pay gap during the early careers of *millennial* American workers. By focusing on this recent cohort, this paper complements the literature studying the impact of wage gains from job changes (Keith & McWilliams, 1999; Loprest, 1992), search frictions, job search and quit behavior (Bowlus, 1997; Light & Ureta, 1992; Royalty, 1998), returns to actual labor market experience (Light & Ureta, 1995), human capital accumulation and wage offers (Amano-Patiño, Baron, & Xiao, 2020) on the early-career growth in the gender wage gap among young US *baby-boom* workers during the 1990s.[7]

[7]While I focus on college graduate workers throughout the paper, in Section E in the Online Appendix I repeat all descriptive and reduced-form analyses on workers who enter the labor market without a college degree. The results I obtain provide suggestive evidence that the search for paid parental leave may not be a crucial determinant of the early-career gender wage gap growth in this group of workers. The early-career average wage profile of workers without a college degree is substantially flatter compared to the wage profile of college graduate workers. In line with recent evidence suggesting that workers without a college degree face a flatter job ladder (Gabe, Abel, & Florida, 2019) and that, while more likely to reallocate to better firms through job-to-job transitions, they are also more likely to fall off the job ladder (Haltiwanger, Hyatt, & McEntarfer, 2018), I also find that workers without a college degree do not experience wage increases following their first early-career job change. In addition, I find that the rising gender wage gap during the early-careers of workers without college degree appears to be mostly driven by the wage losses and labor supply declines experienced after childbirth by women who have children during this time span, underscoring the importance of child penalties in determining the labor market outcomes of women within this group (Kleven, Landais, & Søgaard, 2019).

The paper is structured as follows. Section 2 illustrates the data, the stylized facts describing the early careers of millennial college graduate American workers, and the reduced-form relation between benefits, job changes, and the early-career gender wage gap. Section 3 outlines the hedonic search model and its estimation, and shows the estimation results and the outcomes of several counterfactual exercises. Section 4 concludes.

2 Data, stylized facts and reduced-form evidence

In this section I describe the data used throughout this paper and the features of the early careers of millennial American college graduates. I show that workers' transitions across employers (job changes) are a major determinant of the early-career growth in the gender wage gap in this group, and that the availability of valuable benefits affects workers' job-change decisions.

2.1 Features of the NLSY97 and sample selection

I use restricted-access geocoded data from rounds 1 to 15 (2015) of the National Longitudinal Survey of Youth 1997 (NLSY97), a US-representative panel following 8984 individuals born between 1980 and 1984 annually from 1997 to 2011 and biennially from then on.

The survey records comprehensive information on individuals' demographic characteristics, family background, family-formation decisions, education and labor market history. Regarding the latter, the NLSY97 contains detailed annual information on workers and on their employers and jobs. Using individual-specific employer identifiers, the NLSY97 collects data on employees' employer-specific wages and work hours, and on the availability of benefits such as paid parental leave, unpaid parental leave, employer-sponsored child care, health insurance, life insurance, dental care coverage, retirement plans, tuition refund, sick leave, paid vacation, stock ownership, and of flexible work arrangements.

Since access to benefits in the NLSY97 is self-reported by employees, it is possible that the variables are measured with error, depending on the extent to which workers are aware that benefits are available to them. Regarding paid and unpaid parental leave availability in particular, it is possible that workers for whom these benefits are not valuable are less likely to be aware to have access to parental leave at their current workplace.

In Section F in the Online Appendix, I discuss measurement error in benefit availability in the NLSY97, show that a small share of male and female workers are likely to mistakenly report to not have access to unpaid parental leave, and outline the imputation method that I implement to address measurement error in parental leave reporting.[8] Specifically, I rely on details of the Family and Medical Leave Act of 1993 and of state-level paid leave policies implemented in California (2004), New Jersey (2009), and Rhode Island (2014) to identify workers who, under the assumption that all employers comply with family leave legislation, mistakenly report to not have access to leave. I then use a matching estimator to identify workers who are not covered by parental leave laws and are likely to mistakenly report to not have access to the benefit. For workers who report leave availability with error, I impute it based on establishment size (for unpaid leave), on workers' state of residence (for paid leave), and on the availability of paid and unpaid leave reported by observationally similar workers who correctly record the availability of these benefits.

The imputation of paid and unpaid parental leave availability should reduce concerns that any estimated gender differences in the share of workers' with access to these benefits, or any gender differences in estimated workers' willingness to pay for their provision, are strongly affected by underlying gender differences in workers' awareness regarding the availability of benefits that can be differently relevant for men and women. I use adjusted measures of paid and unpaid parental leave throughout this paper.[9]

I match employer-employee specific information to the weekly arrays of the NLSY97, available for all years between 1997 and 2015. The weekly arrays show each worker's week-specific employment status, employer, and work hours. The arrays allow to follow workers throughout their careers, and to study their labor market transitions, employment gaps, and outcomes within employers and across employers. Throughout this paper, I define transitions across employers as job changes. Finally, I match workers in the sample to their year-specific state of residence using restricted-access NLSY97 geocodes.

[8] As far as benefits such as health insurance, life insurance, dental care, childcare, paid vacation and sick leave are concerned, I show that the shares of NLSY97 employees who report to have access to these benefits are in line with evidence on benefit access arising from other surveys, such as the National Compensation Survey of the Bureau of Labor Statistics and the Job Search Supplement of the Survey of Consumer Expectations of the Federal Reserve Bank of New York. As far as paid parental leave is concerned, I show that NLSY employees' responses are in line with access to this benefit recorded in 2012 the Family and Medical Leave Worksite Survey (Klerman, Daley, & Pozniak, 2012).

[9] Robustness exercises using raw measures of paid and unpaid leave are included in the Online Appendix.

The sample I study consists 266 male and 379 female hourly paid and salaried workers who enter the labor market after college or graduate school completion, and whose labor market histories can be observed for the first six years of labor market experience. I refer to this time-span as workers' early careers.[10]

2.2 Sample characteristics and stylized facts

Table 1 reports the demographic characteristics of workers in the final sample. Young men and women are approximately 24 years old at labor market entry. While a higher proportion of women are married or cohabit (38%) or have a child (8%) at labor market entry, compared to men, marriage rates and the likelihood of having children become increasingly similar across genders throughout workers' early careers.

Importantly, by the sixth year in the labor market, 77% of women are married or cohabit while only 31% of women have at least one child, and the average age of mothers at first childbirth is slightly below 28 years of age. The age at first childbirth, and the fertility rate of women in the sample are considerably lower than the age at first childbirth of college graduate women in recent years. This implies that a large share of the 69% of women who do not have children during their early careers are likely to have children in subsequent years. This information is relevant, as it is plausible that some of these women may take incoming family-formation and fertility choices into account during their early-careers.[11]

[10]Female workers represent 58.8% of the final sample. Applying weights, the estimated share of female workers in the underlying population is 57.9% (standard error 1.9%). Considering the full number of degrees conferred by post-secondary institutions, between 2000 and 2011, the share of women among individuals receiving Bachelor's degrees fluctuated between 57.1% and 57.5%, while the share of women among individuals receiving Master's degrees fluctuated between 58.2% and 60.3% (National Center for Education Statistics, 2023). Thus, the gender composition of the sample I study reflects the well-documented over-representation of women in the population of post-secondary degree recipients among recent cohorts of Americans (Goldin, Katz, & Kuziemko, 2006). In the raw sample of NLSY97 individuals entering the labor market between 2000 and 2011 after completing college or graduate school, 41.6% are males and 58.4% (population estimate 57.8%) are females. The similar gender composition of the final sample studied in this paper relative to the overall sample of NLSY97 college graduates suggests that sample cleaning decisions did not disproportionately select either men or women. Section C in the Online Appendix contains details regarding sample selection. Tables A1, A2 and A3 in the Online Appendix show the impact of sample cleaning decisions on the gender composition of the sample I study.

[11]As documented by Martinez and Daniels (2023), between 2015 and 2019, 78% of US women with a college degree or graduate degree between 40 and 49 years old had ever had a biological child. Among 22-to-49 years old women in the same education group, 42.9% of women had their first child after their 30th year of age. This implies that a substantial share of college graduate women in my sample who do not have children during their early careers will eventually do. This suggests, in turns, that these women

Table 1: Time-invariant sample characteristics

	Men	Women	Difference p-Value
Age at labor market entry	24.41	24.47	0.343
Graduate degree by labor market entry	0.07	0.10	0.136
African American	0.06	0.08	0.092
Marries/cohabits by labor market entry	0.28	0.38	0.010
Marries/cohabits by 3^{rd} yr in labor market	0.57	0.67	0.020
Marries/cohabits by 6^{th} yr in labor market	0.71	0.77	0.224
Has child by labor market entry	0.05	0.08	0.077
Has child by 3^{rd} yr in labor market	0.15	0.18	0.238
Has child by 6^{th} yr in labor market	0.29	0.31	0.415
Age at first childbirth	28.43	27.68	0.053
N	266	379	

Notes: National Longitudinal Survey of Youth 1997 (NLSY97), Rounds 1-15. The sample includes workers who enter the labor market between 2000 and 2011 after college graduation or after graduate school completion. All individuals in the sample have non-missing observations for demographic characteristics and for wages, work hours, and employer and job characteristics throughout the first six years of labor market experience (early career). The number of observations in the table refers to the number of female (379) and male (266) unique individuals in the final sample, observed during the first week in employment at labor market entry. All workers are subsequently observed for six years (72 months, 512 weeks). Custom population weights applied.

Table 2 summarizes workers' early-career histories and the evolution in their labor market outcomes. As shown in panel (a), the early careers of both male and female Millennial college graduates are very dynamic. Slightly less than 50% of men and women in the final sample change at least one employer (job) by six years since labor market entry, the first job-change occurring around the third year of labor market experience for both male and female college graduates.

As shown in panels (b) and (c) of Table 2, workers enter larger firms as they change jobs, suggesting that workers' aim to climb the job ladder by moving toward higher-pay jobs could be a main determinant of job changes, in line with seminal work by Topel and Ward (1992).[12]

may, as long as their career progresses, account for prospective family formation decisions when making job-change choices. This hypothesis is in line with evidence that women's desired fertility can affect their labor force participation and occupational choices, and induce wage losses that represent career costs that women incur even before childbirth (Adda, Dustmann, & Stevens, 2017).

[12]While declining over the last several decades, a large-firm pay premium is still evident in the United States (Bloom, Guvenen, Smith, Song, & von Wachter, 2018).

Table 2: Time-varying sample characteristics

	Men	Women	Difference p-Value
	(a) Labor market history		
Total n. of years employed	5.49	5.36	0.116
Tot n. spells out of work	1.88	2.10	0.044
Tot n. weeks out of work	26.73	33.08	0.116
Total n. of jobs held	1.71	1.73	0.748
Changes employer by 6th year in labor market	0.48	0.47	0.824
Year of experience first job change	3.23	3.22	0.797
	(b) Outcomes – first early-career job		
Average weekly hours worked	42.17	41.80	0.645
Weekly hours > 40	0.25	0.24	0.917
Total n. of weeks employed in t	45.50	47.38	0.003
Hourly rate of pay (in 2005 US Dollars)	16.57	15.67	0.229
Hourly pay – Executive/Managerial	15.70	15.67	0.229
Hourly pay - Professional	18.55	16.76	0.096
Hourly pay - Sales/Office	15.37	12.41	0.047
Hourly pay – Ever changes job	15.01	13.51	0.104
Hourly pay – Never changes job	18.03	17.60	0.678
Employer with 1-49 employees	0.43	0.39	0.415
Employer with 100+ employees	0.44	0.45	0.916
Employer with 500+ employees	0.22	0.23	0.396
	(c) Outcomes – last early-career job		
Average weekly hours worked	44.02	42.57	0.025
Weekly hours > 40	0.45	0.35	0.010
Total n. of weeks employed in t	49.96	48.84	0.055
Hourly rate of pay (in 2005 US Dollars)	23.55	20.92	0.004
Hourly pay – Executive/Managerial	23.26	21.41	0.238
Hourly pay - Professional	25.90	22.37	0.017
Hourly pay - Sales/Office	20.16	16.03	0.014
Hourly pay – Ever changes job	22.98	19.48	0.004
Hourly pay – Never changes job	24.07	22.21	0.186
Employer with 1-49 employees	0.38	0.32	0.132
Employer with 100+ employees	0.54	0.51	0.570
Employer with 500+ employees	0.31	0.27	0.469
N	266	379	

Notes: NLSY97, sample selection as in Table 1. Panels (b) and (c) refer to workers observed, respectively, during the first week employed in their first early-career job, and during their last week employed in their last early-career job. Custom population weights applied.

Concerning labor market attachment, work hours and job continuity, panel (a) of Table 2 shows that women spend approximately 33 weeks out of work, overall, during the their early careers, while men's employment gaps duration sums up to 27 weeks. While men's and women's weekly work hours and annual weeks worked are remarkably similar at labor market entry, as shown in panel (b), by the end of their last early-career job, women work almost two hours less than men per week and one week less per year, as reported in panel (c). The rising gaps in weeks and hours worked are not driven by women's labor supply decline, they are rather determined by men's faster increase in weeks and hours worked.

In line with the literature, panels (b) and (c) of Table 2 also show that the gender pay gap expands over time in the labor market (Amano-Patiño, Baron, & Xiao, 2020; Barth, Kerr, & Olivetti, 2021; Loprest, 1992; Manning & Swaffield, 2009) due to the faster wage growth experienced by young men, across and within occupation classes.

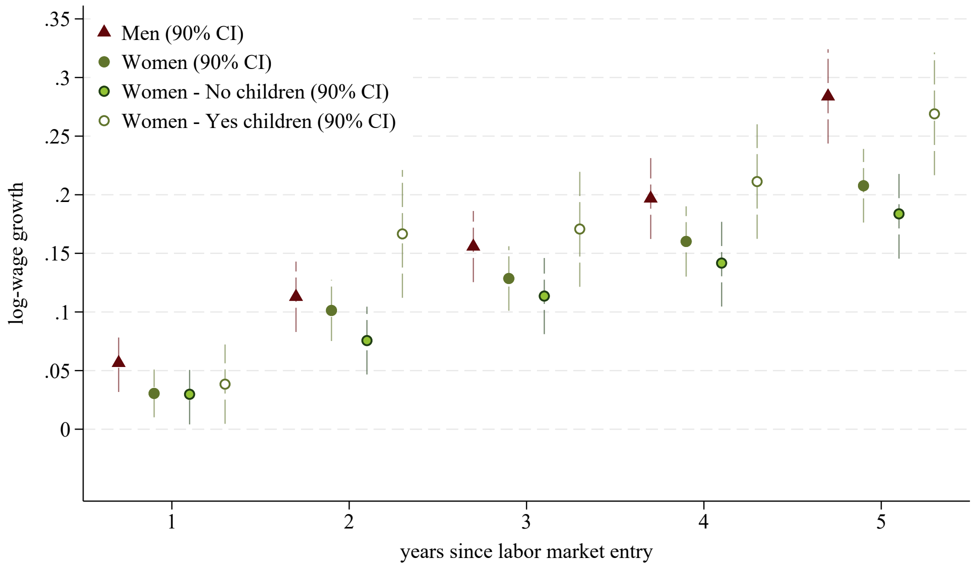
It is worth noting that the rise in the gender wage gap with experience is not driven by the post-childbirth wage losses of the 30% of women in the final sample who have children during their early careers. Figure 1 shows the coefficients of returns to experience estimated separately for men, all women, and women by fertility status in fixed-effect regressions that control for several worker- and employer-specific characteristics. The figure shows that the gender wage gap rises steadily with experience, and that the 70% of women who do not have children in their early careers obtain lower returns to experience compared to men. It implies that, while childbirth events generate penalties that contribute to expand labor market gaps between men and women (Angelov, Johansson, & Lindahl, 2016; Kleven, Landais, & Søgaard, 2019), among recent cohorts of college graduate workers, the roots of those gaps and of their growth over workers' careers may exist since labor market entry.[13]

This result is crucial, and radically distinguishes the early-career paths of college graduate workers from the early-career paths of workers without a college degree. As shown in Figure A9 in the Online Appendix, the early-career gender wage gap among workers without a college degree, and its growth, are entirely determined by the wage losses of

[13] As the early-career increase in the gender wage gap among college graduate workers does not appear to be the ex-post outcome of family-formation decisions, it also cannot be entirely explained by the rising gender gap in weeks and hours worked. Table A4 in the Online Appendix shows that, among workers who do not have children throughout their early careers, the gender wage gap increases over time within occupations and in spite of strong similarities in work hours and labor force attachment between men and women in this group. While this evidence suggests that wage premia for long work hours may not fully explain the observed increase in the gender wage gap during workers' early careers, such premia impact wages, predominantly among career-oriented workers in certain managerial and professional occupations (Bertrand, Goldin, & Katz, 2010; Cortés & Pan, 2019; Gicheva, 2013; Goldin, 2014) where college graduates represent the vast majority of the employed workforce. For this reason, I account for gender differences in work hours throughout this paper.

women who have children during the first few years in the labor market.

Figure 1: Estimated returns to experience

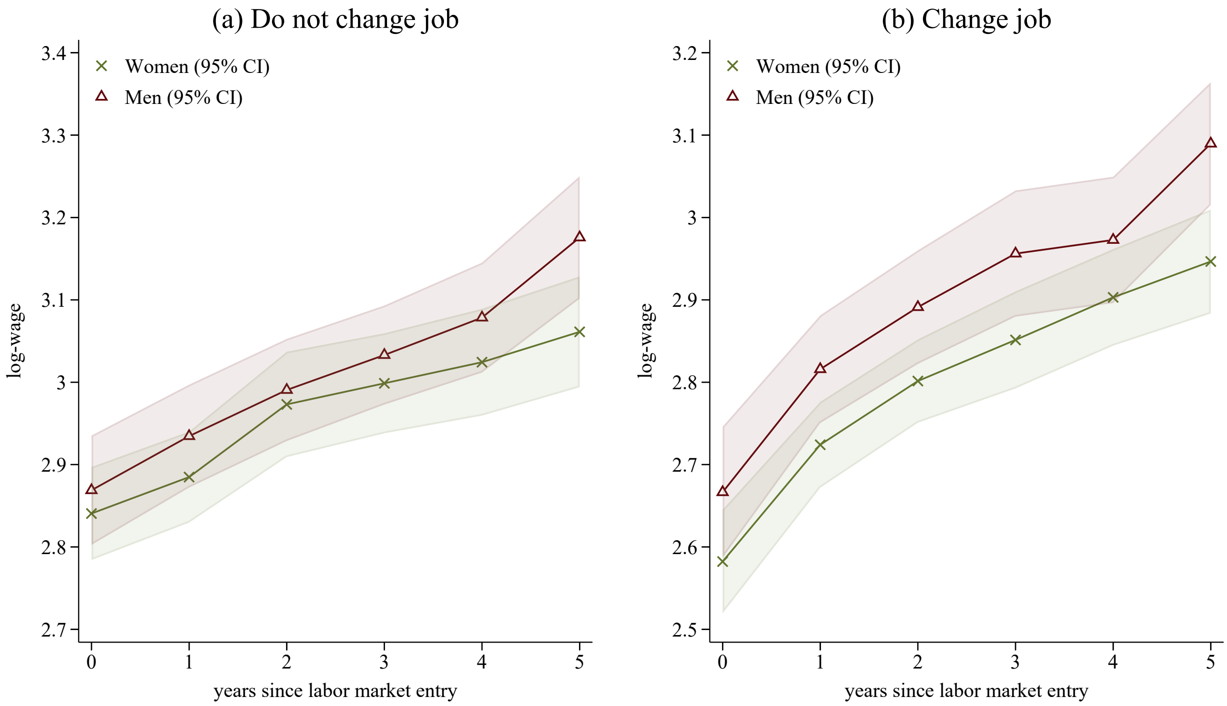


Notes: NLSY97, sample selection as in Table 1. The figure depicts the returns to experience estimated in log-wage regressions through fixed-effect estimator, and clustering standard errors at the individual level. The regressions take the following form $w_{i,t,k} = \alpha + \sum_{k=0}^{5} \beta_k \mathbb{I}\{\text{year in sample} = k\} + \delta_t + x'_{i,t,k}\gamma + \epsilon_i + u_{i,t,k}$. Control variables include occupation and industry dummies, workers' state of residence, establishment dimension, and for the set of benefits and work arrangements available at workers' current workplace: paid and unpaid parental leave (adjusted for measurement error), paid vacation, paid sick leave, child care, schedule flexibility, retirement plan, health insurance, life insurance, dental care coverage, stock ownership, tuition refund.

The early-career increase in the gender wage gap due to gender differences in returns to experience among college graduate workers appears to be, at least partly, driven by the rising difference in wages between college graduate men and women who change at least one employer during their early career.

As shown in Figure 2, male and female workers who move across employers experience faster wage growth compared to workers who remain with the same employer throughout their early careers, suggesting that, as in the seminal contribution of Topel and Ward (1992), returns to *search capital* (Topel, 1991) are an important determinant of early-career wage growth. The gender wage gap, however, also increases faster within this group. The gender wage gap among employees who work for the same employer throughout their entire early careers is small, not statistically significant, and roughly constant during the most part of workers' early careers. As shown in Figure A1 in the Online Appendix, this result is also not driven by the small number of college graduate women who have children during their early career. This evidence suggests that, even before childbirth (or regardless of it), women and men may approach job search differently, leading to women receiving lower returns when switching employers. As a result, the gender wage gap widens over time in the labor market.

Figure 2: Experience wage profiles by job-change status



Notes: NLSY97, sample selection as in Table 1. Panel (a) shows the average log-wage profiles of workers who do not change employer (job) during their early careers. Panel (b) depicts the log-wage profiles of workers who change at least one employer during the same time period. The areas depict 95% confidence intervals of the gender-specific mean (log) wages. Custom population weights applied.

2.3 Job changes, wage growth and the gender wage gap

To quantify the impact of returns to workers' transitions across employers on the early-career growth in the gender wage gap, in this section I estimate the wage gains that workers obtain after their first early-career job (employer) change. If workers move across employers to enter a more profitable employment relationship, their wage should increase compared to the hourly pay they would have received had they not changed employer, conditional on work experience and on other employee- and job-specific characteristics (Burdett & Mortensen, 1998; Topel & Ward, 1992). Even if workers' job changes are driven by workers' aim to move to a more desirable workplace (Sorkin, 2018), or to improve their overall working conditions and the set of benefits offered to them, job changes can still lead to wage growth if, as predicted by hedonic search theory, firms where valuable benefits and work arrangements are available are more productive than firms where benefits are not offered (Hwang, Mortensen, & Reed, 1998).

To estimate wage gains from job changes, I use the following regression model.

$$w_{i,j,k,t} = \sum_{t=2000}^{2016} \beta_t \mathbb{1}\{\text{year} = t\}_t + \sum_{\tau=2001}^{2016} \gamma_\tau \mathbb{1}\{t = \tau\}_i + \delta \mathbb{1}\{t \geq \tau\}_{i,t} + x'_{i,j,k,t}\psi + \varepsilon_i + u_{i,j,k,t} \tag{1}$$

Where $w_{i,j,t}$ is the log-wage that employee i receives at firm (employer) j in week k in year t, $\mathbb{1}\{\text{year} = t\}$ is a non-parametric time trend, and τ denotes the year in which the first job change occurs for workers who change at least one job during their early career, so that $\mathbb{1}\{t = \tau\}$ is an indicator variable taking value 1 if worker i changes their first job in year t. The indicator $\mathbb{1}\{t \geq \tau\}$ takes value 1 in all years following the year in which the first job change takes place for workers who change at least one employer during their early career. $x_{i,j,k,t}$ is a vector of individual- and job-specific control variables, which include a cubic function of workers' experience. Following Light and Ureta (1995), I calculate aggregate experience using the annualized sum of weeks that i spent in employment between labor market entry and week k in year t. Hence, the variable implicitly controls for periods spent out of work, reducing concerns that workers with apparently similar amounts of experience have different levels of job continuity or have accumulated different levels of human capital.[14] I estimate the regression separately for men, women, and women who do not have children during their early career, using a fixed-effect estimator, and clustering standard errors at the individual level.[15]

The parameter of interest is δ, which captures the post-τ change in the average difference in (log) wages between workers who change employer for the first time in year τ and workers who do not. Under the assumption that, absent the job change, the wages

[14] Actual aggregate experience is calculated as $\text{exp}_{i,t} = (\sum_{j=1}^{K} \text{weeks worked in year } j \text{ up to weeek } k)/52$ where $j = 1$ is the year of labor market entry, and K is t.

[15] Following instructions provided by the Bureau of Labor Statistics concerning the use of weights in regression analyses using NLY97 data (BLS), I do not use population weights in my preferred estimated regression. Figure A3 in the Online Appendix, shows that the using weights does not qualitatively change the main estimation results.

of workers who change employer for the first time in year τ and the wages of workers who do not would have grown at the same rate (parallel trends), δ identifies the wage gains obtained by job changers following the first move across employers.[16]

In Figure 3, I report the gender-specific estimated δ-parameters. In panel (a), I report the coefficients estimated on the entire sample of workers. Panel (b) shows the coefficients estimated on the subsample of workers who change at least one job during their early career. This sample restriction reduces concerns that unobserved heterogeneity in expected wage-growth prospects between workers who do not leave their first early-career workplace for at least six years and workers who decide to change employer between one and five years after labor market entry may bias the results.[17]

From panel (a), it is unclear whether, following their first early-career job change, young men experience stronger wage growth compared to women. In the full-sample of workers, there is a 4 log-point difference in returns to the first job change between men (10 log-points, s.e. 0.03) and women (6 log-points, s.e. 0.03), which is marginally not statistically greater than zero at conventional levels (p-value .171). This result may be due to a gender difference in the downward bias in the δ-estimate reported in panel (a).

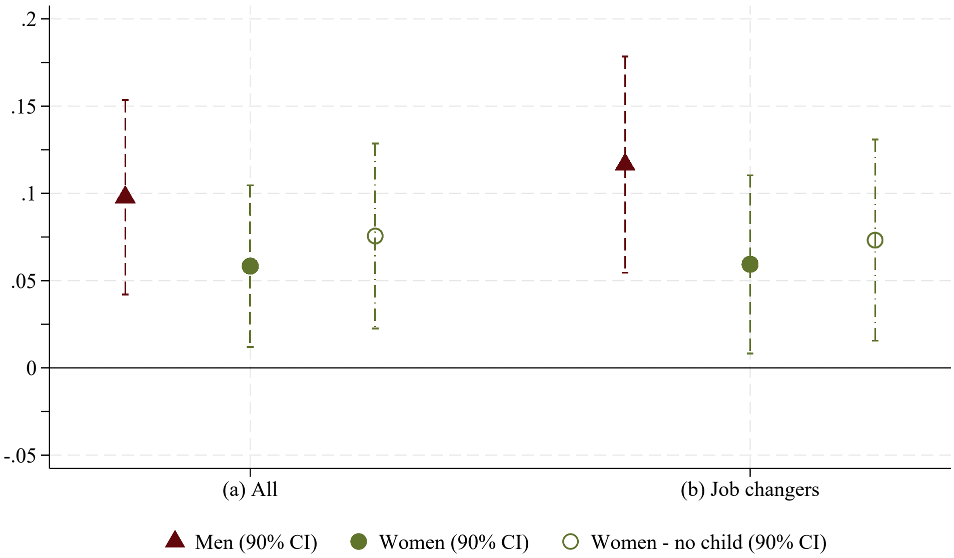
Consistent with stochastic models of mover-stayer heterogeneity (Singer & Spilerman, 1976), if workers who expect their wage to grow slowly in their first early-career workplace are more likely than others to change employer, the counterfactual post-τ time trend in wages among workers who change employer in τ should be flatter than the wage-growth observed among workers who decide to remain in their first workplace throughout their early career. If so, including workers who never change employer during their early career in the control group of workers who do not change job in $t = \tau$ may bias upward the estimated counterfactual time-trend in the wages of workers who change job in τ, thus biasing downward the δ-estimate of the wage gains that workers obtain following their first early-career job change. If male workers expect faster wage growth than female workers on average, then the downward bias in the pseudo diff-in-diff returns to job changes should be especially strong for men. A graphical illustration of this argument

[16] Job changes can lead to wage gains and to faster wage growth either because workers' enter more productive establishments, or because the match between workers and firms improve (Abowd, Kramarz, & Margolis, 1999; Card, Cardoso, & Kline, 2016; Jinkins & Morin, 2018).

[17] For an analysis of workers heterogeneity and the duration of employer-employee matches, see the work by Singer and Spilerman (1976) and its discussion in Topel and Ward (1992).

can be found in Figure A2 in the Online Appendix.[18]

Figure 3: Estimated wage gains following first early-career job change



Notes: NLSY97, sample selection as in Table 1. Panel (a) shows the δ coefficients estimated, separately by gender, in the full sample of workers. Panel (b) shows the δ coefficients estimated on the subsample of workers who change at least one employer throughout their early career. The x variables in the regression include dummy variables for all the benefits offered at time-t workplace, work hours (in log-terms), a dummy variable taking value 1 if time-t workplace employs more than 50 workers, dummy variables for workers' occupation and industry, year-t unemployment rate (in log terms) in worker i's US region of residence, dummies indicating whether, in year t, a worker resides in a state that implemented paid leave laws between 2000 and 2016, and a cubic function of workers' aggregate work experience up to year t. In figure A3 in section D.1 in the Online Appendix, I report the coefficient estimates when custom population weights are used. In figures A12 (unweighted) A13 (weighted) in section E in the Online Appendix I report the estimated δ coefficients for workers without a college degree.

In line with this argument, restricting the sample to workers who change at least one employer during their early career, men's estimated δ becomes larger, and the 6 log-point gender difference in the wage return to workers' first job change is statistically greater than zero with p-value equal to .115 assuming zero covariance between $\hat{\delta}^f$ and $\hat{\delta}^m$.[19] While it cannot be rejected that, following their first job change, women who do not have children during their early career obtain wage gains comparable to those earned by men,

[18]In a recent contribution, Koşar and van der Klaauw (2023) estimate that movements across employers are associated with a positive (though not statistically significant) change in workers' wage-growth expectations.

[19]The test statistic is $z = \frac{(\hat{\delta^m} - \hat{\delta^f})}{\sqrt{s.e(\hat{\delta^m})^2 + s.e.(\hat{\delta^m})^2}} = \frac{(.12 - .06)}{\sqrt{.04^2 + .03^2}} = 1.55$. Because job changes are associated with wage increases for both men and women, and the coefficient standard errors are similar across genders, the covariance between δ^f and δ^m is likely positive. With a covariance as small as .0005, the difference between the two estimated coefficients is statistically greater than 0 with p-value of 0.09.

this result is affected by lack of statistical power.[20]

Overall, gender differences in the wage growth attributable to workers' first job change can explain between 67% and 100% of the early-career growth in the gender wage gap among the 48% of workers in the final sample who change at least one job during the early-career.[21]

2.4 Benefits, work arrangements and job changes

College graduate men and women may experience different wage gains when changing jobs for several reasons, including differences in the likelihood of receiving valuable job offers (search frictions), differences in the wages offered to men and women by employers that provide different bundles of benefits and work arrangements (job offers), and differences in the price that male and female workers accept to pay in exchange for the provision of valuable benefits (preferences for amenities).

While the impact of these factors on men's and women's wage gains from job changes cannot be disentangled in reduced-form analyses relying on observational data, in this section I provide evidence suggesting that workers do value non-wage benefits, and that the provision of certain benefits and work arrangements affects their job-change decisions.

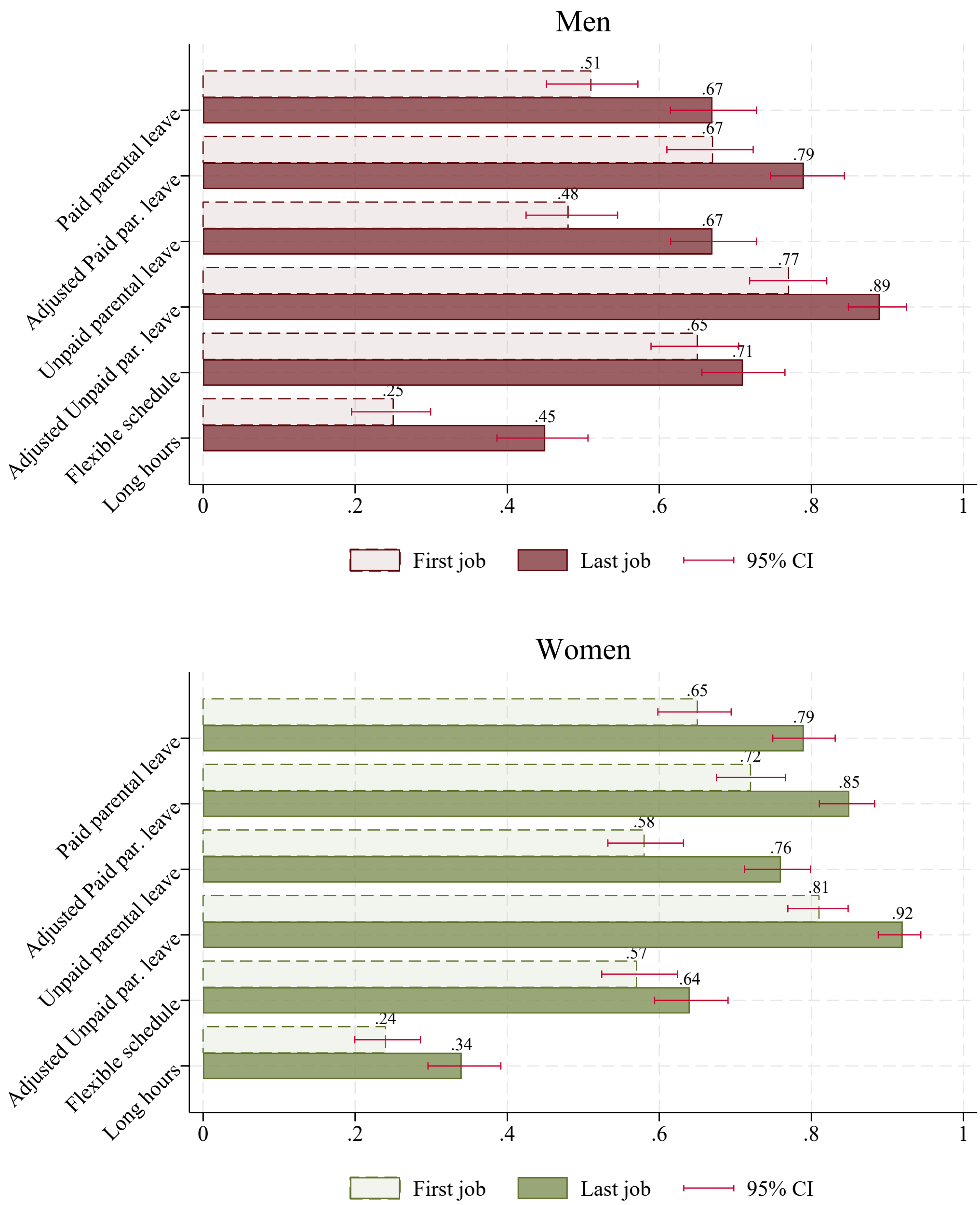
Figure 4 compares, separately among men and women, the shares of workers who are offered different benefits in their first early-career job and in their last early-career job. I assume that a benefit is always provided at an employee's workplace if they report that it is provided at least once during their tenure at the workplace. Thus, any changes in

[20] To further corroborate the result in Figure 3 panel (b), in figures A4 and A5 in the Online Appendix, I report the results of an event study regression that estimates pre-τ and post-τ (log) wage changes experienced by workers. Figure A4, reporting the coefficients estimated on the full sample of workers, shows a negative, though not statistically significant, trend in men's (log) wages in the years preceding their first movement across employers. Limiting the sample to workers who change at least one job during their early-careers, there is no detectable pre-trend in male workers' wages in the years preceding the first early-career job change (Figure A5), while the post-τ estimated wage gains associated with job changes increase.

[21] According to the estimates in Figure 3, among male (female) workers who change at least one employer during their early career, wages increase between 10 log-points (6 log-points), panel (a), and 12 log-points (6 log-points), panel (b). Thus, the first early-career job change is associated with an increase in the gender wage gap between 4 and 6 log-points. Table 2 panels (b) and (c) show that, among workers who ever change job during their early career, the gender wage gap grows from $(\log 15.01 - \log 13.51) = 10.5$ log-points at labor market entry to $(\log 22.98 - \log 19.48) = 16.5$ log-points, a 6 log-point change, by the end of workers' last early-career job. Hence, from 4/6, or 67%, to 6/6, or 100% of the growth in the early-career gender wage gap among workers who change at least one job during their early career, can be explained by the gender differences in the wage gains associated with the first job change.

the availability of benefits reported in the figure are due to workers who change at least one job throughout their early career.

Figure 4: Shares of employees who receive selected non-wage benefits



Notes: NLSY97, sample selection as in Table 1. The figures depict the share of men and women who report to work for an employer that ever offers a certain benefit or work arrangement during the worker's tenure, respectively, at the beginning of their first early-career job (lighter colors), and the end of their last early-career job (darker colors). Custom population weights applied. The shares of workers in establishments where other benefits are available are reported in figure A6 in the Online Appendix. The other benefits available in the NLSY97 are: child care, paid vacation, paid sick leave, health insurance, dental care, life insurance, retirement plan, stock ownership, tuition refund.

As far as paid and unpaid parental leave are concerned, the figures depict both the raw and adjusted shares of workers who report that the benefits are available to them. The adjusted shares are computed by imputing the workplace availability of paid and unpaid parental leave following the method that I outlined in Section 2.1 and explain in Section F in the Online Appendix.

As shown in the figure, the share of employees working for employers who provide paid and unpaid parental leave, and enable workers to work on a flexible work schedule, rises considerably during workers' early career. Figure A6 in the Online Appendix shows that access to other benefits, such as retirement plans or paid vacation, also rises as workers change jobs. This evidence suggests that job changes may be driven not only by workers' aim for higher wages, but also by their willingness to enter workplaces where benefits that match their needs and preferences are available.

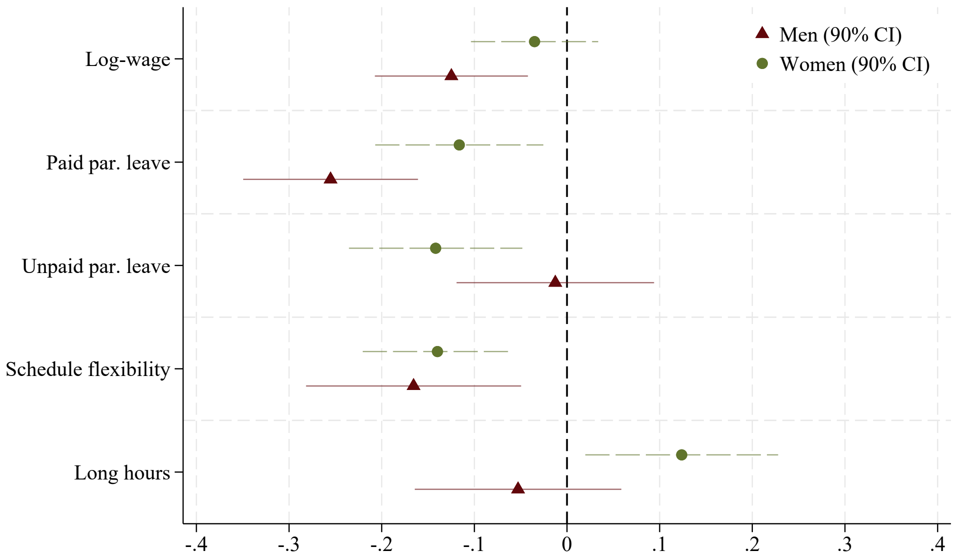
Consistent with the hypothesis that workers take the availability of certain benefits into account when making job-change decisions, Figure 5 reports selected coefficients of a linear probability model of job change, estimated separately for men and women. The coefficients capture the relationship between workers' probability of changing employer between two consecutive years and the characteristics of workers' previous jobs. The estimated regressions are

$$\mathbb{1}\{J_{i,t} \neq J_{i,t-1}\}_{i,t} = \alpha + \sum_{k=1}^{K} \delta_k \mathbb{1}\{\text{Benefit } k \text{ provided}\}_{i,j,t-1} + x'_{i,j,t-1}\psi + \varepsilon_i + u_{i,j,t} \quad (2)$$

Where $\mathbb{1}\{J_{i,t} \neq J_{i,t-1}\}_{i,j,t}$ is an indicator variable taking value 1 if worker i changes employer between years $(t-1)$ and t, and $\mathbb{1}\{\text{Benefit } k \text{ provided}\}_{i,j,t-1}$, is a dummy variable taking value 1 if i's $(t-1)$ employer provides benefit k. $x'_{i,j,t-1}$ is a vector of control variables capturing $(t-1)$ characteristics of i and of their employer, j. ε_i is a worker fixed effect and $u_{i,j,t}$ is an error term. I estimate the model using a fixed-effect estimator and cluster standard errors at the individual level.

Results in Figure 5 show that employees working in firms providing benefits are less likely to change employer by the following year. Among women in particular, the provision of paid and unpaid parental leave are associated with declines by, respectively, 12 percentage-points and 14 percentage-points in the probability of changing employer. For men, only the coefficient for paid parental leave is negative and significant. Schedule flexibility is associated with a significant reduction in the chances of changing employer for both men and women. While working more than 40 hours per week (long hours) does not affect men's job-change probability, working long hours is associated with a 12 percentage-points increase in women's job-change probability.[22]

Figure 5: Linear probability model of job changes - Selected coefficient estimates



Notes: NLSY97, sample selection as in Table 1. The figure reports selected coefficients of a fixed-effect linear probability model of job changes, and 90% confidence intervals. Each coefficient captures the difference in the average probability of changing employer in t between employees whose $(t-1)$-employer offered a benefit and employees whose $(t-1)$ employer did not provide it, controlling for the provision of other benefits, $(t-1)$ (log) wage, a quadratic in $(t-1)$ experience, occupation, industry, employer dimension, regional unemployment rate, and on the total number of weeks spent out of employment until $(t-1)$. The estimated coefficients attached to all the non-wage benefits included as regressors in the linear probability model are reported in Figure A7 in the Online Appendix.

The characteristics of college graduate workers in the final sample, the features of their early careers, the gender differences in the relation between job changes and wage growth, and in the relation between benefits, work arrangements and job changes suggest that hedonic job-search dynamics may shape the early-career path in the gender pay gap.

During their early careers, almost 50% of college graduate workers change at least one job and enter larger firms. As workers change jobs, they experience increases in work hours and in hourly wages that are larger among men, an increase in the likelihood of

[22]The coefficients associated with other benefits are either very close to zero (paid sick leave, stock ownership, tuition refund) or very noisily estimated (see Figure A7 in the Online Appendix). Figure A8 in the Online Appendix compares the estimated coefficients attached to paid and unpaid parental leave when measurement-error adjusted and raw values of the variables are used.

being offered valuable benefits such as paid or unpaid parental leave, and an increase in the likelihood of having flexible work arrangements.[23]

Consistent with hedonic job search theory (Hwang, Mortensen, & Reed, 1998), the steeper early-career wage path of workers who change jobs and the improvement in work arrangements and in benefits offered to workers, suggest that employees who change employer progressively climb the job ladder to enter employment relationships offering both higher pay and better benefits and working conditions. If certain work arrangements or benefits, such as the the provision of paid or unpaid parental leave, are more salient to young women, however, the latter may be willing to pay more for those amenities compared to men, thus experiencing slower wage growth when changing job. Such dynamics would then contribute to the increase in the gender wage gap in years of experience.

Yet, by relying on reduced-form evidence alone it is not possible to either quantify men's and women's preferences (willingness to pay) for benefits and work arrangements, or evaluate their impact on the gender wage gap and on its early-career growth. Even conditional on workers' initial conditions, the likelihood of changing employer and the wage-effect of a job change may be determined by factors unrelated to workers' preferences. The likelihood of receiving and accepting valuable job offers may differ by gender if men and women face different chances of receiving valuable job offers (search frictions), or if comparable men and women are offered different wages (job offers).

In the next section, I use an adaptation of the Bonhomme and Jolivet (2009) model to estimate gender-specific preferences for paid and unpaid parental leave, accounting for potential gender differences in preferences for schedule flexibility and long work hours, in search frictions, and in the job offers received by workers.

[23]Figures A29, A30, and A31 in Online Appendix Section F, show that workers hired by larger employers are more likely to be offered paid and unpaid parental leave, irrespective of whether the availability of these benefits is imputed (adjusted for measurement error) or not.

3 Hedonic search model

3.1 Model setup

There are two separate labor markets, one for male (m) and one for female (f) workers. I denote workers' gender by g. Within each labor market, there are continuous masses of workers and firms. Both employed and unemployed workers search for jobs. An employed worker obtains an outside offer at monthly rate λ_1^g, while the monthly arrival rate of job offers for unemployed workers is λ_0^g. If a worker loses their job, they either become unemployed (at rate q^g per month), or contemporaneously obtain an outside job offer (rate λ_2^g per month) that they accept.[24] The monthly rates of job-offer arrival and of job loss define search frictions.

A job is an employer-employee specific bundle $(w_{i,j}, \mathbf{a}_{i,j})$, where $w_{i,j}$ is the (log) hourly pay of worker i at employer j, and $\mathbf{a}_{i,j} = [a_{i,j}^1, ..., a_{i,j}^K]$ is a vector of indicator variables, each taking value 1 if j offers, respectively, schedule flexibility, long hours, paid parental leave, unpaid parental leave.

The unobserved cumulative distribution of job offers available to workers of gender g is $F^g(w_{i,j}, \mathbf{a}_{i,j}|b, c, pol)$. Workers take $F(.)$ as given, thus the model is in partial equilibrium. To control for within-gender observed heterogeneity in the job-offers that workers receive, I let $F(.)$ depend on workers' ability, denoted b, and on the 1-digit SOC occupation class in which a worker spends the most time during the first few years in the labor market (workers' career), c. Furthermore, I allow the distribution of job offers that workers face to vary depending on whether firm j is located in a state with implemented paid parental leave legislation (pol), or not.[25]

When employed, worker i obtains utility from their (log) wage and from the benefits

[24]The λ_2^g parameter that Bonhomme and Jolivet (2009) add to the Hwang et al. (1998) set-up is of particular interest here. On the one hand, it allows to quantify potential gender differences in the relative likelihood of *constrained* and *unconstrained* job moves. On the other hand, it can highlight gender differences in the ability of workers who received a job termination notice to elicit job offers that would avoid entering unemployment.

[25]In the time-frame that I study, the following states implemented paid family leave laws: California from 2004, New Jersey from 2009, and Rhode Island in 2014. Table A9 in the Online Appendix shows that 14% of men and approximately 17% of women are ever employed in California, Rhode Island or New Jersey in jobs ending after the implementation of paid leave legislation.

and work arrangements offered by their employer, j. The utility function is

$$u_i^g(w_{i,j}, \mathbf{a}_{i,j}) = w_{i,j} + \delta^{g'} \mathbf{a}_{i,j} \tag{3}$$

For each amenity $a_{i,j}^k$, the parameter δ_k^g measures workers' preferences for a^k. For each gender g, workers' marginal willingness to pay for $a_{i,j}^k$ is $e^{-\delta_k^g}$. It represents the lowest wage that a worker would accept to work for an employer that provides benefit a^k as a share of the hourly pay earned at an otherwise identical workplace where a^k is not provided. The larger δ_k^g, the lower the wage that worker i accepts in exchange for the provision of a^k.

The estimation of the model requires the characterization of the steady-state distribution of accepted wages and amenities among employed workers of gender g, $g^g(u|.)$, which can be shown to take the following form

$$g^g(w, \mathbf{a}|b, c, pol) = (1 + k)\frac{f^g(w, \mathbf{a}|b, c, pol)}{[1 + k\bar{F}_u^g(w + \delta'\mathbf{a}|b, c, pol)]^2} \tag{4}$$

Where $f(.)$ denotes the density of job offers received by workers, and $\bar{F} = 1 - F(.)$. Search frictions are measured by $k = \frac{\lambda_1}{q+\lambda_2}$. The higher k, the higher the arrival rate of utility-enhancing job offers relative to the sum of the constrained job-to-job transition rate, λ_2, plus the rate of employment loss, q.[26]

Equation (4) contains one of the key insights of the Bonhomme and Jolivet (2009) model. It highlights that the relation between wages and amenities observed in the data depends not only on workers' preferences (through δ), but also on search frictions (through k) and on the distribution of job offers that workers face (through f and $\bar{F}$). The result is especially relevant in the context of this paper. It implies that, to correctly estimate gender differences in workers' willingness to pay for non-wage benefits and work arrangements, potential gender differences in search frictions and in the wage-amenities bundles that employers offer to male and female workers must be properly accounted for.

[26]Section G.1 in the Online Appendix contains a complete proof for the derivation of equation 4 following Bonhomme and Jolivet (2009).

3.2 Model estimation

I estimate the model using a 72-month panel dataset following the 645 college graduate workers in the sample of interest from labor market entry to the end of the sixth year of labor market experience. I construct the dataset using the weekly arrays of the NLSY97. For each month, I define a worker to be either employed or out of work based on the most frequent employment status observed in the four-week period. If a worker i is employed, I use the most frequent worker-specific employer identifier appearing in the weekly arrays to determine the worker's employer, j. If a worker is not employed, I assume they are unemployed.[27]

For employed workers, I retain information on the wage, benefits and work arrangements available at their current employer, j. The benefits and work arrangements of interest are measured by a set of dummy variables, taking value one if the employer, respectively, provides paid parental leave (pl), or unpaid parental leave (ul), allows for schedule flexibility (fs), and requires the employee to usually work long hours, that is, more than 40 weekly hours (lh).[28]

In the estimated model, wages, benefits and work arrangements are not allowed to change within employer over time. For this reason, I let the job-specific worker's wage be the average (log) wage that an employee i receives while working for a given employer j, and I let each benefit dummy variable take value 1 if employer j ever offers it to worker i during their tenure at their current workplace.

As explained in the previous section, job offers are heterogeneous based on workers' ability (b) and on their career (c). I proxy workers' ability using the (log of the) CAT-ASVAB test score percentile, available in the NLSY97. A worker's career is modeled as a dummy variable indicating the occupation class recurring most frequently during the

[27] Bowlus (1997) shows that part of the gender pay gap between US college graduate workers in the baby boom generation depended of the low search intensity of women who temporarily exited the labor force rather than being unemployed. In the sample of millennial college graduates that I study, however, the number of employment gaps too small for me to be able to separately estimate heterogeneous search frictions depending on the nature of out-of-employment gaps. As shown in Table 2, women and men in the sample of interest spend on average around two spells (groups of consecutive weeks) out of employment during their early careers.

[28] The baseline estimation of the model relies on imputed measures of paid parental leave and unpaid parental leave. The imputation follows the method that I outlined in Section 2 and explain in Section F in the Online Appendix.

worker's first six years of labor market experience. I define the following careers: clerical, executive, professional, other. The clerical career (cl) includes sales and office occupations (2-digit SOC occupation groups 43 and 45). The executive career (ex) includes management, business and financial occupations (2-digit SOC occupation groups 11 and 13). The professional career (pr) includes professional specialty occupations (2-digit SOC occupation groups 15 to 29). I classify the remaining occupations as "other" (ot).

I define time-constant careers for identification purposes. Allowing workers to switch occupations over time would require to model them as job-specific characteristics, and to estimate workers' preferences for occupations alongside workers' preferences for benefits and work arrangements. While this could be done in principle, it is not feasible given the relatively small number of observations in my sample. The definition of careers that I use, instead, assumes that workers choose their career before entering the labor market, and that job markets are segregated by careers. This choice enables me to account for within-gender heterogeneity in job offers, while keeping a parsimonious number of parameters to be estimated.[29]

It is worth noting that the partial equilibrium feature of the Bonhomme and Jolivet (2009) model is crucial to estimate the parameters of interest given the characteristics of the data I use. While the NLSY97 data identify, for each worker, the employer-specific features of their job and movements across employers, the employers' identities are unknown. It implies that the data do not record whether different employees work for the same employer. Consequently, it is not possible to use the NLSY97 to model and estimate employers' decisions to offer certain wages and benefits to their employees. For this reason, the features of labor demand and of the wage-benefit offers received by male and female workers can only be modeled in reduced form. Specifically, wage and benefit

[29]Assuming that the distribution of job offers that workers receive is heterogeneous, within genders, across broad occupation classes, is equivalent to the assumption that a college graduate worker who decides to begin a career as an engineer (professional career) may be looking for employment positions as teaching assistant (professional career) but they may not be looking for employment positions as, say, postal service mail carrier (clerical career) or as tax collector (managerial career), during their first few years of labor market experience. Assuming heterogeneity in the job offer distribution across broadly defined occupation classes (careers) is crucial to correctly estimate preferences for non-wage benefits. Wage offers, returns to experience and returns to job changes differ across occupations, and across genders within occupations, not necessarily due to whether paid and unpaid parental leave are more or less likely to be offered to workers in a certain occupation group. Not allowing job offers to differ across occupations would bias estimates of preferences for non-wage benefits.

offers take the following form.

$$w^*_{i,j}(b,c,pol) = \varphi^w_0 + \mu^w_1 b_i + \mu^w_2 pol_i + \rho' \mathbf{a}^*_{i,j} + \sum_{c\in\{\text{ex, pr, ot}\}} \varphi^w_c \mathrm{c}_i + \sigma_w \varepsilon^w_{i,j} \tag{5}$$

$$a^{k*}_{i,j}(b,c,pol) = \mathbf{1}\{\mu^{a^k}_0 + \mu^{a^k}_1 b_i + \mu^{a^k}_2 pol_i \sum_{c\in\{\text{ex, pr, ot}\}} \varphi^{a^k}_c \mathrm{c}_i + \varepsilon^{a^k}_{i,j} > 0\}$$
$$\text{for} \quad a^k_{i,j} \in \{a^{fs}_{i,j}, a^{lh}_{i,j}, a^{pl}_{i,j}, a^{ul}_{i,j}\} \tag{6}$$

Where $\varepsilon^w_{i,j}$ and $\varepsilon^{a^k}_{i,j}$ for $a^k_{i,j} \in \{a^{fs}_{i,j}, a^{lh}_{i,j}, a^{pl}_{i,j}, a^{ul}_{i,j}\}$ are independent standard normal shocks. φ^w_0 and $\mu^{a^k}_0$ are, respectively, the mean offered wage, and a constant factor affecting the likelihood of amenity a^k provision, in the sales and clerical career (the base group). μ^w_1 denotes the relation between wage offers and workers' ability, while $\mu^{a^k}_1$ denotes the relationship between worker i's ability and the likelihood that worker i is offered benefit a^k. μ^w_2 denotes the difference in average wages offered by employers in states with implemented paid leave laws and other states, while the $\mu^{a^k}_2$ identifies the change in the likelihood that worker i is offered benefit a^k when a job is located in a state with implemented paid leave legislation. The parameters φ^w_c and $\varphi^{a^k}_c$ capture, respectively, the difference between the average wage offered in career c (executive, professional, other) and the average wage offered in the clerical career, and the career-specific changes in the inverse cumulative distribution function of amenity a^k, compared to the base group. Equation (5) shows that wage offers $w^*_{i,j}(b,c,pol)$ depend on the amenities that employers offer through the coefficient vector ρ. For each amenity a^k, ρ^k represents the average difference in the offered wages between employers who provide a^k and employers who do not.

It is worth noting that, whether the values of the ρ-parameters correspond to workers' preferences parameters δ's, is an empirical question. In a frictionless hedonic labor market, workers and employers match based on workers' preferences for the amenities that each employer provides, and on employers' marginal cost of amenity provision.[30] If so, the magnitudes of the ρ parameters closely correspond to the magnitude of the δ param-

[30]See Rosen (1974), and its discussion in Greenstone (2017).

eters, while their sign should be opposite. As shown by Hwang, Mortensen, and Reed (1998), however, if search frictions exist and finding job offers takes time, not all workers are able to immediately select into jobs providing the benefits they value the most. In this context, productive firms offer benefits to attract and retain a greater number of employees. In equilibrium, profits are equal across employers, and employees who work for the most productive (and larger) employers, who offer valuable benefits, earn higher wages than employees working for (smaller) employers who do not provide any benefit. Thus, if search frictions exist, the ρ-parameters corresponding to valuable benefits are non-negative, and workers obtain a *de facto* wage premium for working in firms providing valuable benefits, even if they may be willing to accept lower wages to work in those firms.

Consistent with the impact of search frictions on the relation between wages and amenities that Hwang, Mortensen, and Reed (1998) theorized, equation (4) shows that stronger search frictions (smaller k) cause the empirical density of job offers accepted by employed workers, $g(.)$, to resemble the unobserved distribution of job offers determined by labor demand, $f(.)$, while not necessarily reflecting workers' preferences.

I can now find the gender-specific likelihood function describing the distribution of wages and amenities among employed workers and workers' labor market transitions. I drop the superscript g to simplify notation. Since I observe all workers from labor market entry, I assume that all workers experience one initial period of unemployment. I denote this period as $t = 0$.[31] Following this initial period of unemployment, in any of the subsequent 71 months, workers can either remain unemployed ($e_{t+1} = 0$) or become employed ($e_{t+1} = 1$). The labor market transitions that workers experience between any two periods t and $(t+1)$ affect each worker's contribution to the $(t+1)$ likelihood function,

[31] This assumption, which differs from the Bonhomme and Jolivet (2009) framework, is instrumental for me to model the job-search period that all workers experience when first entering the labor market, and its impact on the job offers that men and women receive. This assumption is important to account for the fact that the distribution of job offers that men and women receive may differ since labor market entry for several reasons. As documented by Cortés, Pan, Pilossoph, Reuben, and Zafar (2023), stronger risk-aversion among female highly educated workers leads them to accept job offers earlier than their male counterparts at labor market entry, incurring wage losses. Young women may also receive different information when browsing for jobs compared to men (Wasserman & Gallen, 2021) which can determine gender differences in received and accepted job offers. Furthermore, women tend to ask lower wages compared to men (Roussille, 2024). More broadly, with this assumption I can account for the impact on wages of gender differences in behavioral traits Shurchkov and Eckel (2017) and in other unobserved factors including, for example, different preferences for commuting (Le Barbanchon, Rathelot, & Roulet, 2021), other than the sources of job-offer heterogeneity that I explicitly model.

which, as in Bonhomme and Jolivet (2009), is

$$l_{t+1} = q^{ju_t}[1-\lambda_0]^{uu_t} \times \lambda_0^{uj_t} f_{t+1}(w_{t+1}, \mathbf{a}_{t+1}|.)^{uj_t} \times [1-\lambda_1 \bar{F}(u_t|.) - \lambda_2 - q]^{s_t} \times$$
$$\times [\lambda_1 \mathbf{1}\{w_{t+1} + \delta'\mathbf{a}_{t+1} > w_t + \delta'\mathbf{a}_t\} + \lambda_2]^{jj_t} f_{t+1}(w_{t+1}, \mathbf{a}_{t+1}|.)^{jj_t} \tag{7}$$

Where $s_t, jj_t, ju_t, uj_t, uu_t$ are dummy variables indicating, respectively, workers who, between months t and $t+1$, remain in the same job, change job, enter unemployment, exit unemployment, remain unemployed. Finally, the likelihood function, capturing the labor market transitions and outcomes of all N^g workers in each gender-specific sample for all months $(t+1) \in \{1, 72\}$ of their early career, is

$$L(.) = \prod_{i=1}^{N^g} \prod_{t=0}^{71} l_{t+1}(e_{t+1}, w_{t+1}, \mathbf{a}_{t+1}, s_t, jj_t, ju_t, uj_t, uu_t | e_t, w_t, \mathbf{a}_t, b, c, pol) \tag{8}$$

Following Bonhomme and Jolivet (2009), the functional forms for $f(w^*, \mathbf{a}^*|.)$ and $\bar{F}_u(u|.)$ and, consequently, the functional form of the likelihood function (8) can be found by exploiting the assumptions of normality and independence of the random shocks in the wage and benefits offers.[32]

The gender-specific likelihood functions depend on the following parameter vector.

$$\begin{aligned} \underbrace{\Gamma}_{44\times1} = [& \underbrace{\theta}_{36\times1}, \underbrace{\lambda}_{4\times1}, \underbrace{\delta}_{4\times1}] \\ & [\varphi_0^w, \underbrace{\rho'}_{[\rho^{fs}\ \rho^{lh}\ \rho^{pl}\ \rho^{ul}]}, \sigma_w, \mu_1^w, \mu_2^w, \underbrace{\varphi^{w\prime}}_{[\varphi_{ex}^w\ \varphi_{pr}^w\ \varphi_{ot}^w]}, [\varphi_0^{fs}\ \mu_1^{fs}\ \mu_2^{fs}\ \underbrace{\varphi^{fs\prime}}_{[\varphi_{ex}^{fs}\ \varphi_{pr}^{fs}\ \varphi_{ot}^{fs}]}], ..., \\ & [\varphi_0^{ul}\ \mu_1^{ul}\ \mu_2^{ul}\ \underbrace{\varphi^{ul\prime}}_{[\varphi_{ex}^{ul}\ \varphi_{pr}^{ul}\ \varphi_{ot}^{ul}]}]], [\lambda_0,\ \lambda_1,\ \lambda_2,\ q], [\delta^{fs},\ \delta^{lh},\ \delta^{pl},\ \delta^{ul}] \end{aligned} \tag{9}$$

Where θ is the (36×1) vector of parameters characterizing the unobserved distribution of job offers that workers receive, λ is the (4×1) vector of search friction parameters, and δ is the (4×1) parameter-vector of preferences.

I estimate Γ separately for male and female workers using the sequential maximum likelihood algorithm proposed and explained by Bonhomme and Jolivet (2009). The al-

[32]Section G.2 in the Online Appendix contains the derivations of the the functional forms for $f(w^*, \mathbf{a}^*|.)$ and $\bar{F}_u(u|.)$.

gorithm allows to estimate the parameter vector of workers' preferences for benefits and work arrangements δ through a revealed-preferences approach that exploits the changes in the wages earned by workers who, in any given month, undergo job-to-job transitions between employers offering different sets of benefits. This method identifies δ conditional upon the previous estimation of θ using the wage-benefit outcomes of workers who are at labor market entry or undergo an unemployment-to-employment transition, and conditional on the estimation of search frictions λ through the maximization of the likelihood of observing different types of labor market transitions in the data.

In sections G.3 and G.4 in the Online Appendix, I provide an in-depth explanation of parameter identification, and I discuss why it is crucial to separately identify the parameters characterizing the distribution of job offers that workers receive (θ) in order to correctly identify gender differences in willingness to pay for non-wage benefits (δ).

3.3 Estimation results

The estimation results for parameters describing workers' willingness to pay for amenities are reported in Table 3, together with block-bootstrapped standard errors and likelihood-ratio test p-values. The estimated coefficients show that both men and women highly value the possibility to work for employers offering paid and unpaid parental leave. Willingness to pay for these benefits, however, is substantially higher for women.

In Table 4, I use the estimated coefficients to calculate, for every benefit or work arrangement, the lowest wage that a worker would be willing to accept in order to move from an employer that does not offer the benefit to one that does, as a share of the wage received in the workplace where the benefit is not available. Concerning paid parental leave, the coefficients imply that a woman earning the same wage as a comparable man in a firm that does not offer paid parental leave, would be willing to accept 44% of the wage that the man would accept to enter a workplace where the benefit is available. Regarding unpaid leave, a woman earning the same wage as a comparable man in an establishment where the benefit is not available would accept 89% of the wage that the man would accept to enter a firm that offers the benefit.[33]

[33]For paid parental leave, $\exp\{-1.590\}/\exp\{-.759\} = .4356 \sim 44\%$. In order to enter a workplace where paid parental leave is offered, women (men) would be willing to accept as little as 20.4% (46.8%) of the wage they earn at a workplace where paid leave is not offered. A woman earning around (real) \$16 per

Table 3: Estimated preference coefficients

	Schedule flexibility	Long hours	Paid parental leave	Unpaid parental leave
(a) Women	.733	-.013	1.590	1.119
	(.449)	(.667)	(.417)	(.422)
	[1.000]	[1.000]	[.003]	[.006]
(b) Men	.605	.687	.759	1.008
	(.493)	(.720)	(.560)	(.554)
	[1.000]	[.586]	[.029]	[.003]

Notes: NLSY97, sample as in Table 1. Sequential maximum likelihood estimates of preference parameters. The table reports the estimated parameters identifying workers' willingness to pay for, respectively, schedule flexibility, long hours, paid parental leave, and unpaid parental leave. Block-bootstrapped standard errors are in parentheses. P-values of the likelihood ratio tests ($H0 : \delta^{a_k} = 0$, $H1 : \delta^{a_k} \neq 0$) are in brackets. The number of observations used in the estimation are 27288 for women (379 women observed over 72 months since labor market entry) and 19152 men (266 men observed over 72 months since labor market entry).

Table 4: Estimated marginal willingness to pay for amenities

	Schedule flexibility	Long hours	Paid parental leave	Unpaid parental leave
Women	48.0%	101.3%	20.4%	32.7%
Men	54.6%	50.3%	46.8%	36.5%

Notes: NLSY97, sample as in Table 1. Calculated willingness to pay for benefits and work arrangements based on estimates in table 3. The table shows the lowest wage that a worker would accept to move from a firm that does not provide a benefit to a firm that does, as a share of their no-benefit wage. This is equal to $\exp(-\delta_a)$. To calculate it, notice that workers' utility is $u_a = w_a + \delta_a$ if an employers provides a benefit and $u_n = w_n$ if an employer does not provide the benefit. Thus $u_a \geq u_n$ if $w_a \geq \exp(-\delta_a)w_n$.

As far as long hours are concerned, the preference coefficient is negative for women, implying that women would require to be compensated with a higher hourly wage in order to work long hours. Conversely, men in the sample of interest attach a positive value to the possibility of working long hours. The coefficients, however, are not statistically different from zero for either men or women.

Both male and female workers in the sample of interest do appear to value schedule

hour, the average wage of women in the sample of interest at labor market entry, would accept any wage higher than \$3.26 per hour to work in an establishment where paid parental leave is offered. In fact, this implies that virtually every job where paid parental leave is offered ensures women a higher utility than jobs where the benefit is not available. A man earning the same initial wage would not accept less than \$7.49. Though less wide, gender differences in willingness to pay for unpaid parental are also non-negligible.

flexibility. The difference in magnitude between the estimated gender-specific parameters implies that the lowest wage that women are willing to accept to work for an employer that offers schedule flexibility is 88% of the wage that a man would accept. While, this result corroborates earlier evidence highlighting the relationship between gender differences for various types of flexible work arrangements, including work-from-home arrangements (Mas & Pallais, 2017), part-time work (Liu, 2016) and schedule flexibility (Wiswall & Zafar, 2018; Xiao, 2021), and the gender wage gap, the preference coefficient that I estimate is not statistically significant for men and marginally not statistically significant at 10% level for women.

Overall, the results show that, while gender differences do exist in preferences for work hours and flexible work arrangements, the provision of unpaid and, mostly, paid parental leave appears to be a key determinant of gender differences in workers' selection into different firms, and of the gender wage gap, during early careers.[34]

It is worth noting that estimated gender differences in workers' willingness to pay for paid and unpaid parental leave are qualitatively unaffected when the model is estimated using different sets of benefits (see Table A16 and Table A17 lines 4 and 5 in the Online Appendix). Furthermore, gender differences in willingness to pay for paid parental leave persist when the model is estimated on workers who enter the labor market with any level of education, even if additional benefits (such as paid vacation, or the contribution to a retirement plan) are added to the model (see Table A17 lines 1 to 3). In addition, gender differences in estimated preferences for paid parental leave are robust to the estimation of the model using raw measures of paid and unpaid leave that do not account for potential measurement error in these variables (see Table A17, line 4). Estimating gender differences in workers' willingness to pay for paid and unpaid parental leave using raw measures of access to these benefits, however, is likely to *underestimate* rather than *overestimate* gender differences in preferences for these benefits. This finding is consistent

[34]This result is not surprising. According to a recent survey by the Boston College Center for Work & Family among employees in four large firms that do offer parental leave, 75% of surveyed workers reported that parental leave provision increased the likelihood that they would remain at their current workplace, and almost all male and female workers who took leave used it to its largest extent. Yet, flexible work arrangements appeared to be valuable for a subset of the interviewed workers: upon returning to work after a parental leave period, 50% of women and 27% of men reported an increased reliance on flexible work arrangements, while the remaining 50% of women and 73% of men did not change their work schedule or location (Boston College Center for Work and Family, 2019).

with the hypothesis that especially male workers who are unaware of the availability of paid and unpaid parental leave are also those who value these benefits the least. Finally, and perhaps most importantly, gender differences in workers' willingness to pay for paid parental leave persist when women's preferences are estimated on the subset of college graduate female workers who do not have children throughout their early careers (see Table A17, panel (a), line 6). This evidence suggests that, in a context where access to parental leave is not guaranteed, workers may progressively select into jobs that do offer this benefit. Young women, however, are more likely than their male counterparts to consider the availability of parental leave when accepting job offers, thus receiving lower wages compared to men in exchange for the provision of this valuable benefit, even substantially before the birth of their first child.

Tables A10 and A11 in the Online Appendix report, respectively, the estimated search friction coefficients, and the estimated parameters of the job offer distribution that men and women face. Estimated search frictions differ by gender during the early careers of college graduate workers: employed male workers are 4 percentage-points more likely than their female counterparts to receive at least one utility-enhancing job offer per year, and they are 4 percentage-point less likely to exit employment at least once per year. The distributions of job offers differ by gender: women are offered lower wages compared to men across and within careers (occupations), obtain offers implying lower wage-increases, compared to men, from employer that require long work hours or offer paid parental leave, and obtain offers implying higher wages-increases, compared men, from employers offering unpaid parental leave and schedule flexibility. Overall, consistent with Sockin (2022), employers providing valuable benefits offer wages at least as high as employers that do not provide benefits. While employers providing paid parental leave in states without implemented paid leave policies offer lower wages to women compared to employers where the benefit is not available, the estimated parameter is small and not statistically different from zero. Thus, in line with hedonic job-search models, throughout their careers workers progressively enter employment relationships that offer both higher wages and more valuable bundles of benefits and work arrangements. Yet, gender differences in willingness to pay for certain benefits imply that women accept lower wage gains, compared to men, when entering employment relationships in which certain benefits, and most prominently paid parental leave, are provided.

3.4 Counterfactual analyses

In this section I use the estimated parameters to predict the early-career evolution of men's and women's wages, and to quantify the impact of gender differences in workers' willingness to pay for parental leave on the the early-career gender wage gap growth.

To compute workers' predicted and counterfactual wages, I simulate cross-sections of 1000 labor market entrants of each gender. I then use the estimated parameters to predict workers' yearly transitions across employment statuses and across jobs, and their wage-benefits outcomes. I perform the simulations separately by careers, as defined in the previous section. For simplicity, I simulate samples ignoring within-gender heterogeneity in ability. For each year in the labor market, the simulation generates a distribution of employed workers across jobs defined by their pay level, benefits, and work arrangements. I use the simulated distribution to compute the year-of-experience- and career-specific average (log) wage. The year-t average (log) wage of workers of a given gender is the weighted average of the career-specific simulated average wages, with weights equal to the share of workers of a given gender in each career. To study changes in the labor market outcomes of college graduate male and female workers in different counterfactual scenarios, I repeat the simulation of workers' early careers using different parameter values.

Figure 6 panel (a) verifies that the model predicts the early-career growth in the gender wage gap among college graduate workers observed in the data well, even not accounting for within-gender heterogeneity in ability. The model over-predicts the early-career growth in the gender wage gap among college graduate workers by 0.85 log-points.[35]

The figure shows that the gender wage gap is predicted to expand by 5 log-points over workers' early careers, and that women's stronger willingness to pay for paid parental leave can explain the bulk of the lower wage growth experienced by female workers compared to their male counterparts. The early-career gender wage gap would increase by 1.8 log-

[35] Figures A38 and A39 in the Online Appendix show that the ability of the model to predict the early-career growth in the gender wage gap worsens when the model is estimated substituting alternative benefits (paid vacation) for, respectively, long hours and schedule flexibility. This evidence supports the modeling choice that I made in accounting for flexible work arrangements and long work hours when studying workers' job search, in line with the extensive literature showing that part of the growth in the gender wage gap during workers' early careers can be attributed to gender differences in the likelihood to work long hours (Bertrand, Goldin, & Katz, 2010; Gicheva, 2013; Goldin, 2014) and in workers' valuations of different types of flexible work arrangements (Wiswall & Zafar, 2018; Xiao, 2021)

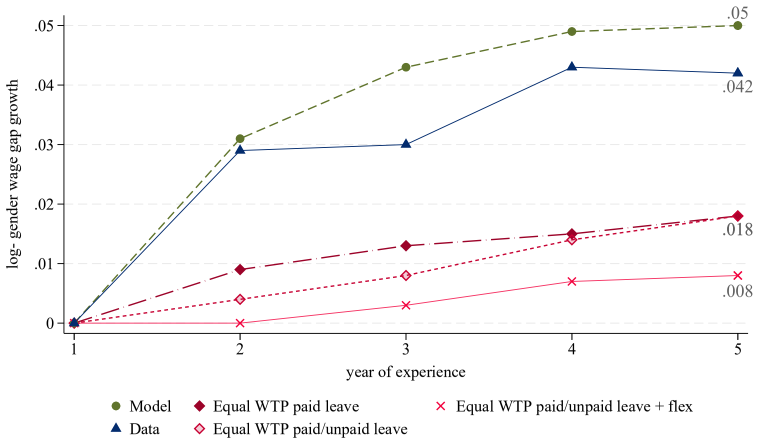
points over workers' early careers if men's and women's preferences for paid parental leave were identical. Thus, women's stronger willingness to pay for paid parental leave explains around 64% of the early-career growth in the gender wage gap. The growth in the gender gap would decline by one additional log-point if men's and women's willingness to pay for both paid and unpaid parental leave as well as for schedule flexibility were the same. In fact, gender differences in preferences for schedule flexibility can explain, approximately, 16% of the early-career growth in the gender wage gap.[36]

Workers' preferences for parental leave affect earned wages by impacting the lowest wages that employees accept in order to enter a firm that provides the benefit. To the extent that paid parental leave availability is more salient for women, they accept lower wages, and experience lower wage growth, as they are hired by employers providing it. Since the likelihood of being employed in leave-providing firms rises over time as workers search for better jobs, the wage gap due to gender differences in willingness to pay for paid parental leave increases in years of experience. In fact, as shown in Figure 6 panel (b), the gender wage gap rises largely due to women's strong willingness to pay for paid leave even though both men's and women's wages increase over time in the labor market.

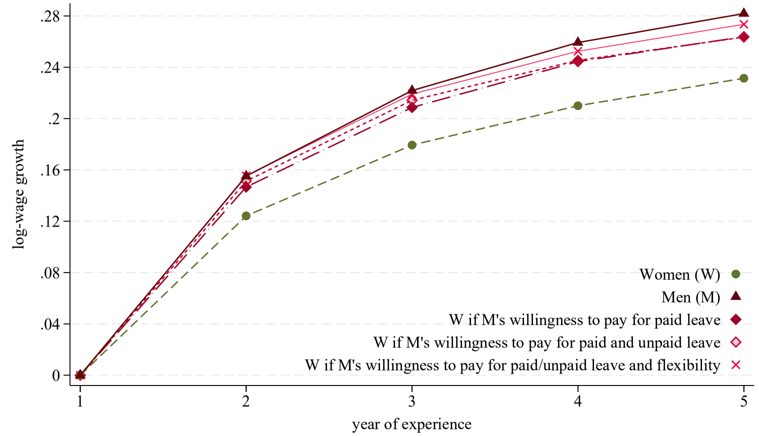
Last, I use the estimated parameters of the model to predict changes in the early-career growth in the gender wage gap in the counterfactual scenario in which, as paid parental leave becomes available in all workplaces, neither men nor women trade off wages for the provision of the benefit. The results of this exercise, reported in Figure 6 panel (c), show that the early-career growth in the gender wage gap could decline in this scenario. As shown in Figure 6 panel (d), while both men and women experience faster wage growth when paid parental leave becomes universally available, the model predicts that such change would disproportionately impact women's early-career wage growth. Since women trade off the provision of paid parental leave for lower wages, compared to men, when accepting jobs at workplaces that offer the benefit, making the trade-off unnecessary could enable them to experience a greater increase in the wage gains obtained upon entering firms that offer the benefit compared to their male counterparts.

[36]The result on the effect of schedule flexibility on the early-career growth in the gender wage gap is in line with one of the key findings in the work by Wiswall and Zafar (2018). In their study of the college-major choices, job-search decisions and labor market outcomes of American college students, they find that women's stronger willingess to pay for flexible work arrangements can explain around 25% of the gender earnings gap of workers at age 30.

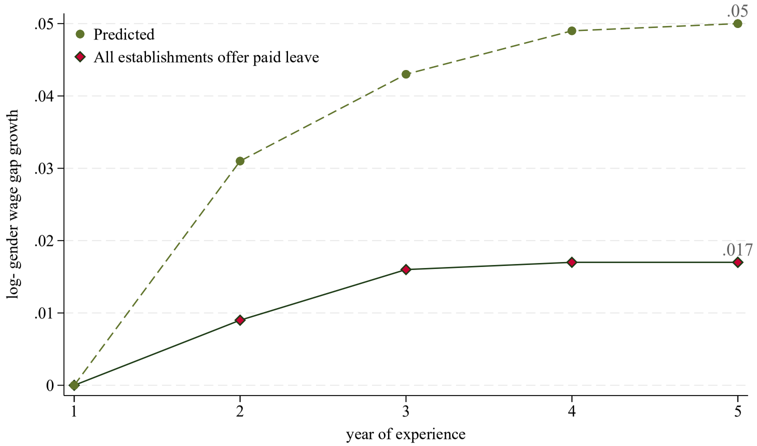
Figure 6: Model fit and counterfactual exercises



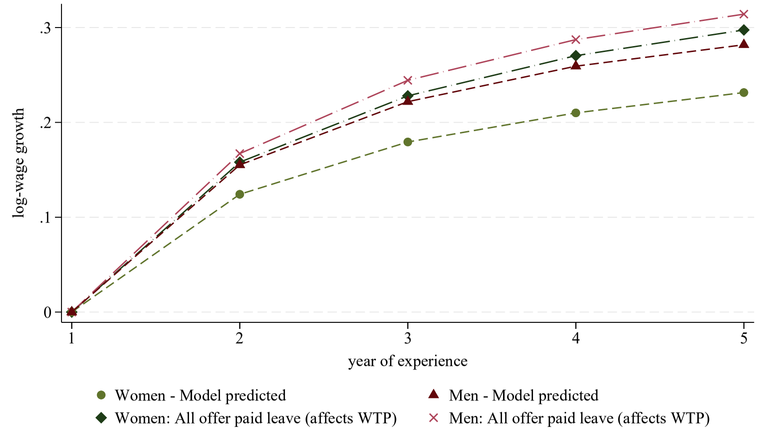
(a) Preferences and gender gap growth



(b) Preferences and log-wage growth



(c) All firms offer paid leave - gender gap growth



(d) All firms offer paid leave - men's and women's wage growth

Notes: NLSY97. The figures show counterfactual exercises using simulated samples of workers constructed using the estimated gender-specific parameters of the structural job-search model.

It is worth noting that, in this counterfactual exercise, I assume that the wages that firms offer to male and female workers do not change when paid leave is widely available. This is equivalent to modeling the scenario as a policy change that mandates the provision of paid leave while entirely subsidizing the employers' costs of providing the benefit, so as to minimize the chances that employers change their labor demand decisions.

One major limitation of this analysis is that it cannot account for the possibility that firms change their wage offers following the implementation of a policy mandating parental leave. General equilibrium considerations are not possible in this context, given the partial equilibrium nature of the model that I estimate, and given that employers' identity is unknown in the NLSY97. While the vast literature on parental leave policies that I reviewed in the introduction answered several questions concerning the labor-supply and labor-demand effects of parental leave policies, the evidence that I provide signals that the scarcity of parental leave may itself be a significant institutional factor influencing workers' labor market outcomes and, potentially, gender wage differences.

4 Conclusions

In this paper I studied the first six years of labor market experience of millennial college graduate Americans to understand whether and how workers' search for employers offering parental leave affects the early career growth in the gender wage gap. As parental leave is not guaranteed to most employees in the United States, workers may search for firms offering such benefit and be willing to accept lower wages in exchange for its provision. If parental leave is more valuable to young women than to young men, the gender wage gap may grow as workers search for, and are hired by, employers who provide it.

Using a hedonic search model, I showed that women's stronger willingness to pay for paid and unpaid parental leave is a key determinant of the early-career growth in the gender wage gap. I estimated that women are willing to pay 66% more than men to obtain access to paid parental leave, and 11% more than men to have access to unpaid parental leave. The lower wages that women accept to work in firms offering paid parental leave explain around 60% of the early-career growth in the gender wage gap. I also found that a policy mandating and subsidizing the provision of paid parental leave could reduce the growth in the early-career gender wage gap. The widespread availability of parental leave could mute workers' trade-off between leave access and wages, and reduce the gap in accepted wages between men and women entering leave-providing firms.

Keeping in mind their partial equilibrium nature, the results in this paper suggest that the scarcity of parental leave availability may be consequential. If the decision to offer this benefit is decentralized to employers, and the costs of providing it are not subsidized, only employers that can afford the cost will provide it (Goldin, Kerr, & Olivetti, 2020), and workers for whom parental leave is more salient will pay a higher price for its provision, thus accepting lower wages compared to potentially equally productive workers. This may be detrimental to the labor market outcomes of recent generations of young college graduate women, whose substantial labor market commitment pairs with family-formation ambitions (Goldin, 2024), and for whom the availability of parental leave may represent a first-order form of employment insurance and career continuity in the event of a childbirth.

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Online Appendix

to

The Search for Parental Leave and the Early-Career Gender Wage Gap

Ilaria D'Angelis

Not for Publication

Contents

Online Appendix

A Parental leave policies in the United States

As of 2024, 23 US states have passed paid parental leave laws. The year of effectiveness of each state law is in parentheses: California (2004), Colorado (2023/24), Connecticut (2021/22), Delaware (2025/26), District of Columbia (2020), Maine (2025/26), Massachusetts(2019/21), Maryland (2024/26), New Hampshire (2023), New Jersey (2009), New York (2018), Oregon (2023), Rhode Island (2014), Vermont (2023/25), Washington (2019/20), Virginia (2022), Arkansas (2023), Tennessee (2024), Alabama (2024), Minnesota (2026), Texas (2023), Florida (2023), Kentucky (2024) (Bipartisan Policy Center, 2024).

Employees' eligibility to paid parental leave varies by states. Most state laws require employees to have earned a minimum income threshold in wage in the twelve months prior to the paid leave period; some laws have either minimum-hours requirements for eligibility or employment length requirements. Some laws do not provide employment protection (California), or limit employment protection to certain employees (Oregon). Two states exempt employers with less than 25 (Oregon) or 50 (Rhode Island) employees from paid leave provision.

Crucially, paid leave is mandatory and funded through social insurance only in some states: California, New Jersey, Rhode Island, New York, Washington, Massachusetts, Connecticut, Oregon, Colorado, Maryland, Delaware, Minnesota, Maine, and in the District of Columbia. Paid leave is voluntary, that is, can be offered by private insurance in: New Hampshire, Virginia, Vermont, Arkansas, Tennesee, Alabama, Texas, Florida. It is mandatory and offered by private insurance companies in Kentucky.

Regarding unpaid leave, the Family and Medical Leave Act (FMLA) of 1993 exempts firms not employing at least 50 employees for 20 weeks per year from the requirement of providing the unpaid leave to workers. Eligibility for FMLA coverage is limited to employees having worked for an employer for at least one year and for 1250 hours minimum in the previous year. Furthermore, if employers offer paid leave, they can require employees to use padi leave allowances during their FMLA-covered leave (Wage and Hour Division: United States Department of Labor, 2023).

For comparison, all European Union countries mandate paid maternity leave and parental leave, and the average compensation for parents on leave is 50% of their previous earnings (Janta & Stewart, 2018; van Belle, 2016).

B Econometric challenges in the identification of preferences for non-wage benefits

In Rosen (1974) theory of compensating wage differentials, in a competitive labor market equilibrium homogeneous workers and firms, workers with strong preferences for a valuable amenity accept wage cuts in exchange for its provision. The consequent equilibrium cross-sectional correlation between valuable benefits and wages is negative. The literature provided evidence that this implication is counterfactual. Hwang, Reed, and Hubbard (1992) noted that estimating workers' preferences for job attributes through the cross-sectional relation between wages and amenities leads to substantial biases due to workers' unobserved skill heterogeneity. Brown (1980) further noted that employee-level panel data fixed-effect regressions also provide biased (towards zero) compensating differential estimates as they cannot control for employer heterogeneity and, over time, workers' may search and progressively enter more productive jobs offering both higher wages and better amenities. Hwang, Mortensen, and Reed (1998) showed that the lack of evidence on compensating wage differentials through reduced-form wage regressions suggests that labor market is frictional and not perfectly competitive. Several authors provided evidence that properly accounting for job search dynamics changes the empirical estimates of workers' preferences for amenities (Bonhomme & Jolivet, 2009; Gronberg & Reed, 1994; Hwang, Mortensen, & Reed, 1998; Sullivan & To, 2014). The empirical implications of hedonic search models have been used to estimate workers' willingness to pay (preferences) for job attributes by Bonhomme and Jolivet (2009), Flabbi and Moro (2012), Hotz, Johansson, and Karimi (2018), Liu (2016), Morchio and Moser (2024), Sullivan and To (2014), Sorkin (2018), Xiao (2021). Khandker (1988) was the first to introduce non-wage attributes in a search model.

C Sample cleaning

The sample I study consists 266 male and 379 female hourly paid and salaried workers who enter the labor market between 2000 and 2011 after college or graduate school completion and whose labor market histories can be observed for the first six years of labor market experience. I refer to this time-span as workers' early careers. To reconstruct workers' careers, I define the year of labor market entry as the first year such that, for two consecutive years, a worker is employed for more than 26 weeks per year (Loprest, 1992) and for at least 30 hours per week on average (Blau & Kahn, 2017). Applying similar sample-cleaning restrictions as Gicheva (2013), I exclude individuals who ever report hourly wages below 1/2 the year-specific federal minimum wage or above $240 (in 2005 US dollars). I also drop workers who, when employed, ever report working less than 1 hour per week or more than 112 hours per week at their job, and who are ever self-employed, employed in agriculture or in the military. I also exclude individuals for whom, in employment relationships lasting more than 14 weeks, information about relevant job-specific information (hourly wage, benefits, establishment dimension) is ever missing and not-imputable using adjacent-years information for the same job. Because information regarding non-wage benefits, alternative work arrangements, and access to paid or unpaid leave is not available in the NLSY97 for jobs that last less than 14 weeks, I ignore these employment relationships in my analysis. Finally, I drop individuals with missing CAT-ASVAB test-score percentile, and with missing information regarding their current US state of residence.

The tables below show that the sample-selection choices that I implement affect a small number of male and female workers, and do not cause any gender imbalances in the composition of the final sample of interest compared to the gender composition of recent cohorts of American college graduates.[1]

[1]See footnote 10 in the paper.

Table A1: Sample cleaning steps - initial sample: workers entering the labor market with a Bachelor's, Master's, or Ph.D - share of observations dropped relative to number of observations remained after the previous cleaning step

	Men	Women	N. Obs.
Hourly pay ever below 1/2 min wage	0.043	0.041	1353
	(0.202)	(0.199)	
Hourly pay ever above $240 or missing	0.062	0.056	1274
	(0.242)	(0.229)	
Job-specific weekly hours ever below 1	0.000	0.000	1274
	(0.000)	(0.000)	
Job-specific weekly hours ever above 112 or missing	0.015	0.017	1253
	(0.122)	(0.131)	
Ever in military during early career	0.012	0.005	1243
	(0.107)	(0.074)	
Ever self-employed during early career	0.113	0.084	1124
	(0.317)	(0.277)	
Missing ASVAB test-score percentile	0.147	0.132	969
	(0.355)	(0.338)	

Notes: NLSY97. Standard deviations in parentheses.

Table A2: Sample cleaning steps – initial sample: workers entering the labor market with a Bachelor's, Master's or Ph.D -share of dropped individuals for whom the values of the variables of interest are missing

	Share of dropped individuals with ever missing value in variable of interest	N. dropped individuals
Hourly pay ever above $240 or missing	0.772	79
Job-specific weekly hours ever above 112 or missing	0.952	21

Notes: NLSY97. The table reports the number of individuals dropped from the sample at specific data-cleaning steps, and the share of those individuals for whom the value of the variable of interest is ever missing.

Table A3: Sample cleaning steps – initial sample: workers entering the labor market with a Bachelor's, Master's or Ph.D - share of women at each sample cleaning step

	Female share Pop. estimate using weights	Female share Sample	N. Obs. men and women
College graduate sample at labor market entry	0.578	0.584	1412
	(0.494)		
Hourly pay never below 1/2 min wage	0.578	0.585	1353
	(0.494)		
Hourly pay never above $240 or missing	0.581	0.586	1274
	(0.494)		
Job-specific weekly hours never below 1	0.581	0.586	1274
	(0.494)		
Job-specific weekly hours never above 112 or missing	0.580	0.586	1253
	(0.494)		
Never in military during early career	0.581	0.587	1243
	(0.494)		
Never self-employed during early career	0.588	0.595	1124
	(0.492)		
No missing ASVAB test-score percentile	0.594	0.600	969
	(0.491)		
Final sample	0.579	0.588	645
	(0.494)		

Notes: NLSY97. The final sample includes 379 women and 266 men, each observed for 312 weeks (72 months) since labor market entry. The final sample only includes individuals who satisfy all the previous cleaning steps and for whom the relevant employment and job characteristics information (including benefits provided) is never missing during workers' early career. Standard deviation for population estimates in parentheses.

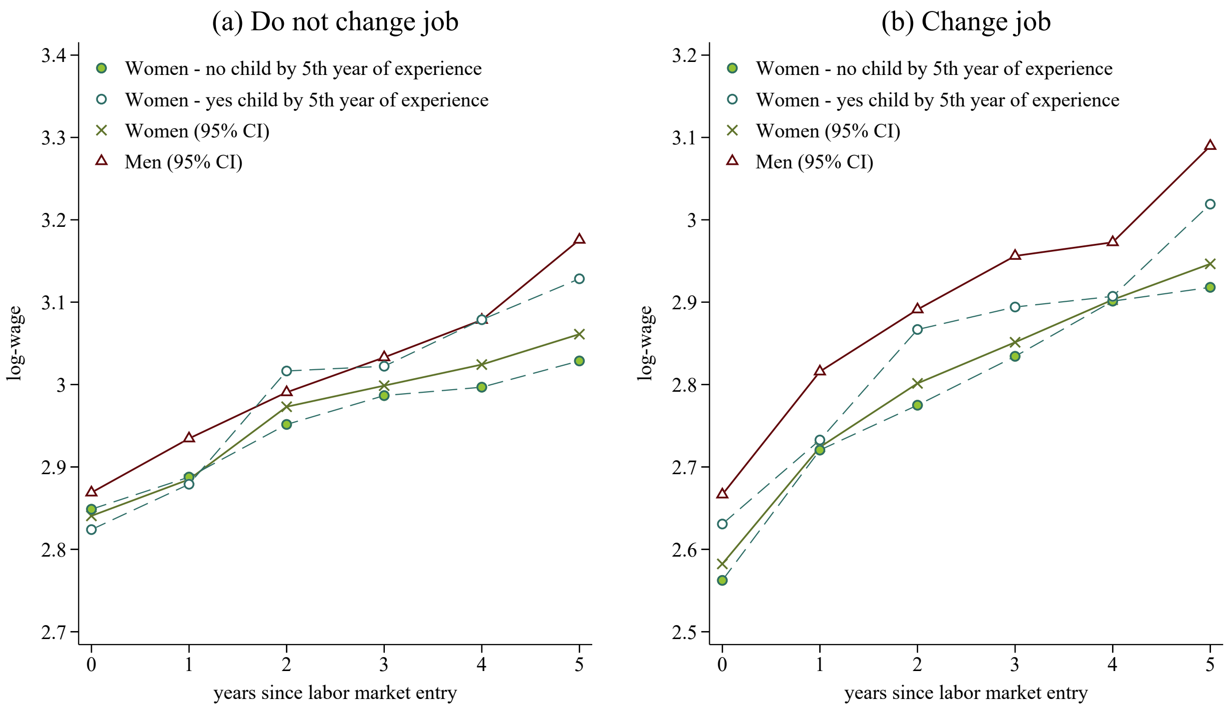
D Additional descriptive statistics and reduced-form analyses

Table A4: Time-varying sample characteristics – workers without kids during early career

	Men	Women	Difference p-Value
	(a) Labor market history		
Total n. of years employed	5.46	5.37	0.380
Tot n. spells out of work	1.93	1.97	0.771
Tot n. weeks out of work	28.24	32.51	0.380
Total n. of jobs held	1.75	1.77	0.918
Changes employer by 6th year in labor market	0.48	0.48	0.953
Year of experience first job change	3.41	3.23	0.554
	(b) Outcomes – first early-career job		
Average weekly hours worked	42.36	41.78	0.496
Weekly hours > 40	0.27	0.25	0.608
Total n. of weeks employed in t	46.10	47.14	0.137
Hourly rate of pay (in 2005 US Dollars)	16.58	15.50	0.265
Hourly pay – Executive/Managerial	15.06	15.67	0.483
Hourly pay - Professional	18.46	16.74	0.259
Hourly pay - Sales/Office	16.44	12.12	0.015
Hourly pay – Ever changes job	15.07	13.11	0.148
Hourly pay – Never changes job	17.97	17.75	0.805
Employer with 1-49 employees	0.44	0.40	0.291
Employer with 100+ employees	0.44	0.45	0.841
Employer with 500+ employees	0.20	0.23	0.129
	(c) Outcomes – last early-career job		
Average weekly hours worked	44.16	43.67	0.379
Weekly hours > 40	0.44	0.38	0.221
Total n. of weeks employed in t	49.92	49.27	0.357
Hourly rate of pay (in 2005 US Dollars)	23.70	20.55	0.008
Hourly pay – Executive/Managerial	23.09	21.15	0.284
Hourly pay - Professional	25.70	21.94	0.063
Hourly pay - Sales/Office	21.03	16.37	0.026
Hourly pay – Ever changes job	23.32	19.29	0.009
Hourly pay – Never changes job	24.05	21.73	0.210
Employer with 1-49 employees	0.38	0.32	0.152
Employer with 100+ employees	0.52	0.54	0.705
Employer with 500+ employees	0.30	0.29	0.974
N	187	255	

Notes: NLSY97, sample selection as in Table 1. The table refers to the subsample of workers who do not have children throughout their early careers.

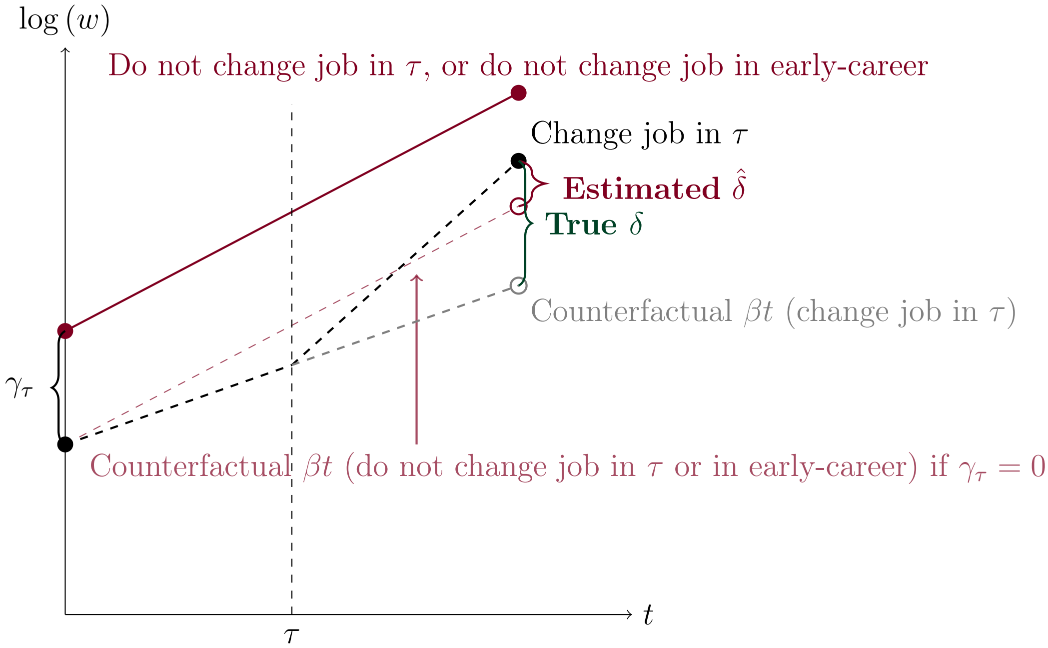
Figure A1: Experience wage profiles by job change and family composition status



Notes: NLSY97, sample selection as in Table 1. The figures depict the average log-wage profiles of men and women over years of experience. Panel (a) refers to workers who remain with the same employer throughout their early careers, panel (b) refers to workers who change at least one employer during their early careers. Both figures compare men to, respectively, all women, women who do not have children during their early career, and women who have their first child during their early career.

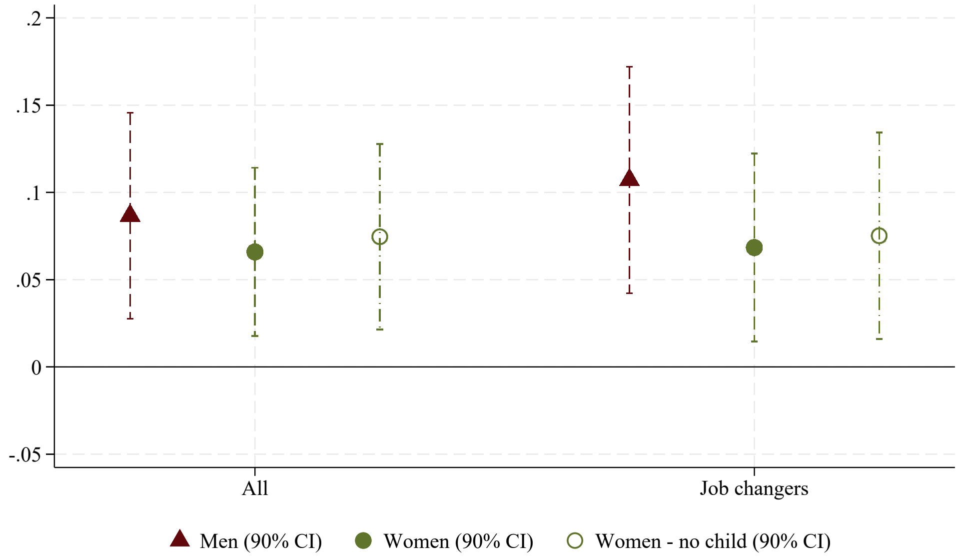
D.1 Diff-in-diff and event study: gains from first job change

Figure A2: Bias in $\hat{\delta}$ if workers with low wage-growth prospects are more likely to move across employers



Notes. The figure shows that, if workers who undergo their first job change in any year $t = \tau$ during their early career are selected among workers who face worse wage-growth prospects at their initial employer, then the estimated counterfactual time-trend in wages is biased upwards if it is estimated on a sample that includes workers who do not change any job (employer) during their early career. If so, then the pseudo diff-in-diff estimated wage gains from job changes are biased downwards. As workers who change at least one job during their early career, but do not change job in τ, are likely to face similar wage-growth prospects at their initial employer compared to workers who change their first job in τ, then restricting the sample to workers who change at least one employer during their early careers should be instrumental to estimate a more credible time-trend in wages, and a δ coefficient that more closely identifies the true wage gains experienced by workers following their first early-career job change. In the figure, solid dots and lines indicate observed wages and trends, while dashed lines and empty dots indicate unobserved wages and trends.

Figure A3: Estimated wage gains following first early-career job change - weighted regression results



Notes: NLSY97, sample selection as in Table 1. Panel (a) shows the δ coefficients estimated, separately by gender, in the full sample of workers. Panel (b) shows the δ coefficients estimated on the subsample of workers who change at least one employer throughout their early career. The coefficients are estimated through fixed-effect estimator, standard errors are clustered at the individual level and custom population weights are applied. The estimated regression controls for the same variables listed in the notes to figure 3.

In what follows I estimate workers' wage gains following their first early-career job change using an event study regression.

The main scope of this analysis is to explore whether the small wage gains from job changes estimated on the full sample of men may be due to an underlying unobserved difference in time-trends in wages between workers who change at least one employer during their early career and workers who do not. If so, then it is reasonable to consider the gender difference in wage gains from job changes estimated on the full sample of workers as a lower bound of the true gender difference in wage gains from job changes.

In the regression that I estimate, I denote τ as the year in which the first early-career job change occurs, for workers who change at least one job in this period of time. $d \in \{(\tau - 3), ...(\tau + 1)\}$ denotes whether year t is a certain number of years before or after the year the first job change of an individual. I denote ι the year or labor market entry. The estimated regression takes the following form.

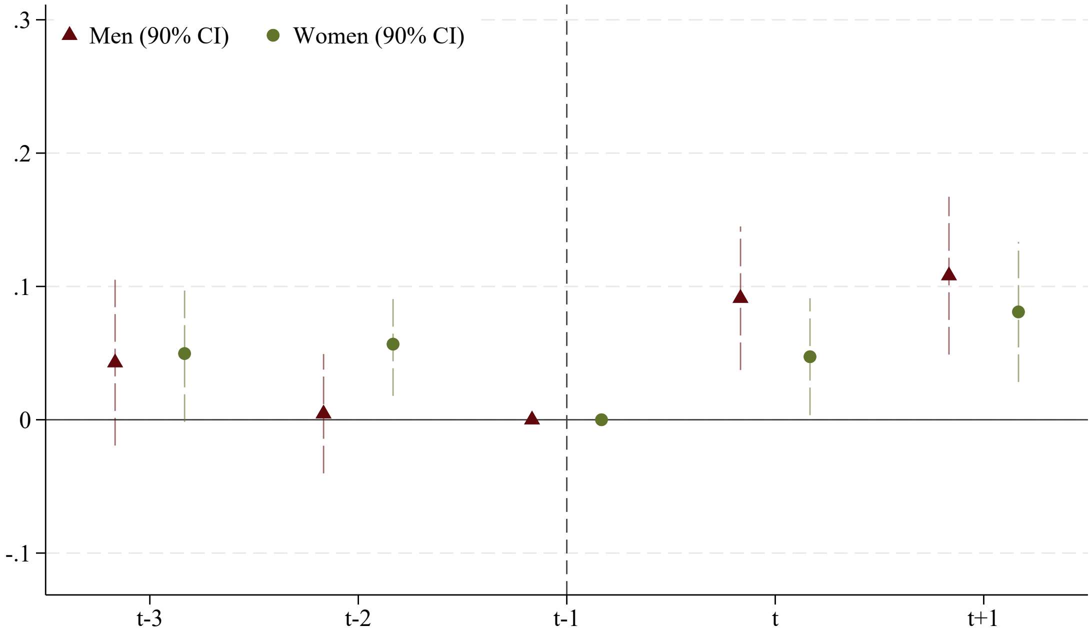
$$
\begin{aligned}
w_{i,j,k,t} = \sum_{t=2000}^{2016} \beta_t \mathbb{1}\{year = t\}_t + \sum_{d=\tau-3}^{\tau+1} \delta_d \mathbb{1}\{t = d\}_{i,t} + \sum_{y=\iota}^{\iota+5} \gamma_y \mathbb{1}\{t = y\}_{i,t} + \\
+ \xi \mathbb{1}\{\forall t \in \{\iota, ..., \iota+5\}, t > \tau+1\}_i + x'_{i,j,k,t}\psi + \varepsilon_i + u_{i,j,k,t} \qquad (1)
\end{aligned}
$$

Where β_t captures the time-trend in wages among workers who do not change jobs throughout their early career, and γ_y captures returns to experience (years in the labor market since labor market entry). δ_d, the coefficient of interest, captures the difference in (log) wages between $d \in \{(\tau-3), (\tau-2), \tau, (\tau+1)\}$ and the omitted category, $\tau - 1$, for workers who change at least one job during their early careers. I can keep in the regression sample workers who never change employer throughout their early career by including a dummy variable that takes value 1 for those workers such as, for any year of labor market experience ι between labor market entry and five years later, t is not in the range between $\tau - 3$ and $\tau + 1$. Thus, the dummy takes value one for all workers who, by the end of their early-career, have not (yet) changed at least one employer.

The regression controls for the same time-varying firm-specific and worker-specific variables included in regression model 1 in the paper, with the exception of workers' experience that is included here as a set of dummy variables. I estimate the model through fixed-effect estimator, and cluster standard errors at the individual level.

The δ_d coefficients, separately estimated for men and women, are reported in Figure A4. Concerning women, it appears that they experience a wage drop in the year preceding their first early-career job change, and gain afterwards. As far as men are concerned, it appears that workers' wages progressively decline as they approach their first early-career job transition, with respect to the time-trend in wages estimated among workers who do not change any job throughout their early career. While the trend is not statistically significant, this evidence suggests that, among men who change at least one job during their early careers, wages grow slowlier over years in the labor market with respect to workers who do not change any employer in the same period of time.

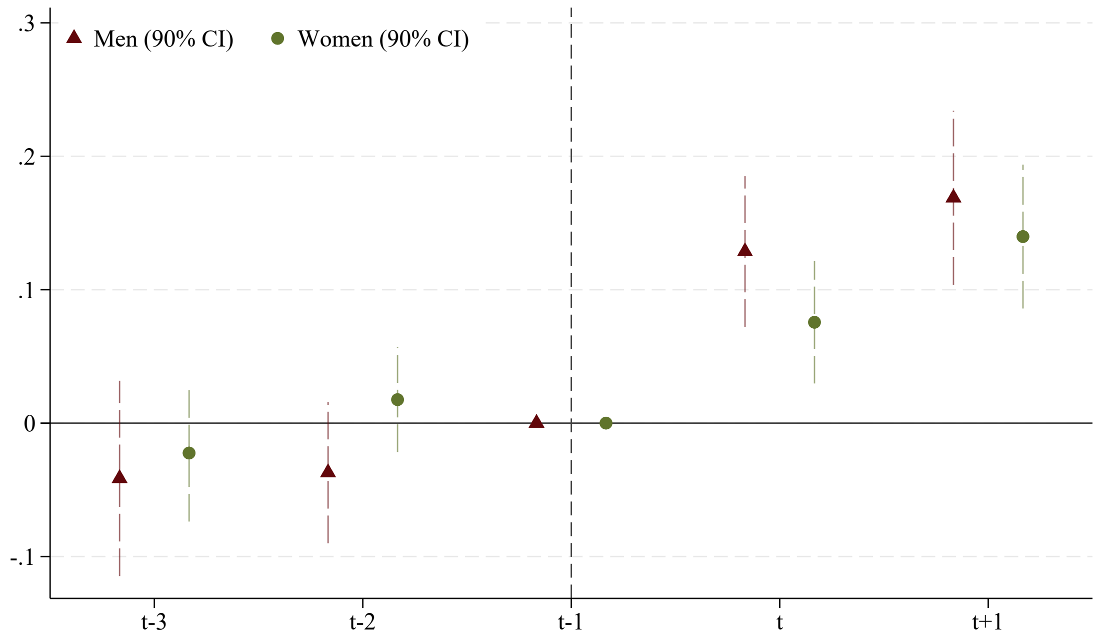
Figure A4: Event study estimates of wage gains from first job change - full sample



Notes: NLSY97, sample selection as in Table 1. The figure depicts the $\delta_{\tau\pm d}$ coefficients in regression 1. A control dummy variable coded as $\mathbb{1}\{t = \tau + 100\}$ is included to maintain workers who do not change jobs during their early careers as a control group in the regression. Thus, the coefficients β_t capture the time-trend in wages among workers who do not change job throughout their early career.

The negative pre-trend in men's wages fades out as the event study regression is estimated on the subsample of workers who change at least one job during their early career. It suggests that, once workers' selection into job-change decisions is taken into account, male workers do appear to gain substantially than women by switching jobs during their early careers.

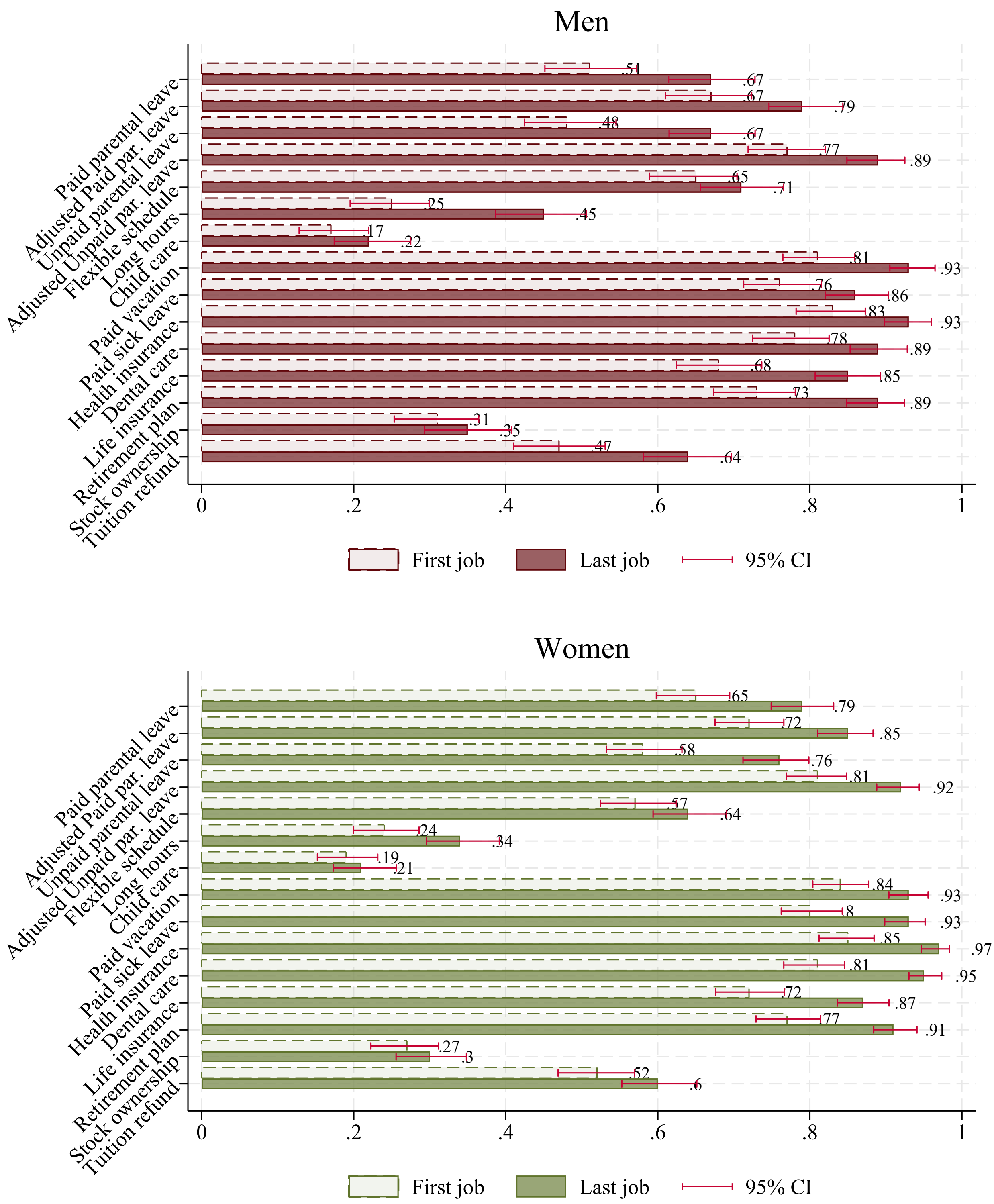
Figure A5: Event study estimates of wage gains from first job change - workers who change at least one job during their early career



Notes: NLSY97, sample selection as in Table 1. The figure depicts the $\delta_{\tau \pm d}$ coefficients in regression 1.Workers who do not change at least one job during their early-careers are excluded. Thus, the coefficients β_t capture the time-trend in wages among workers change at least one job during their early career but for whom $t \neq \tau \pm d$.

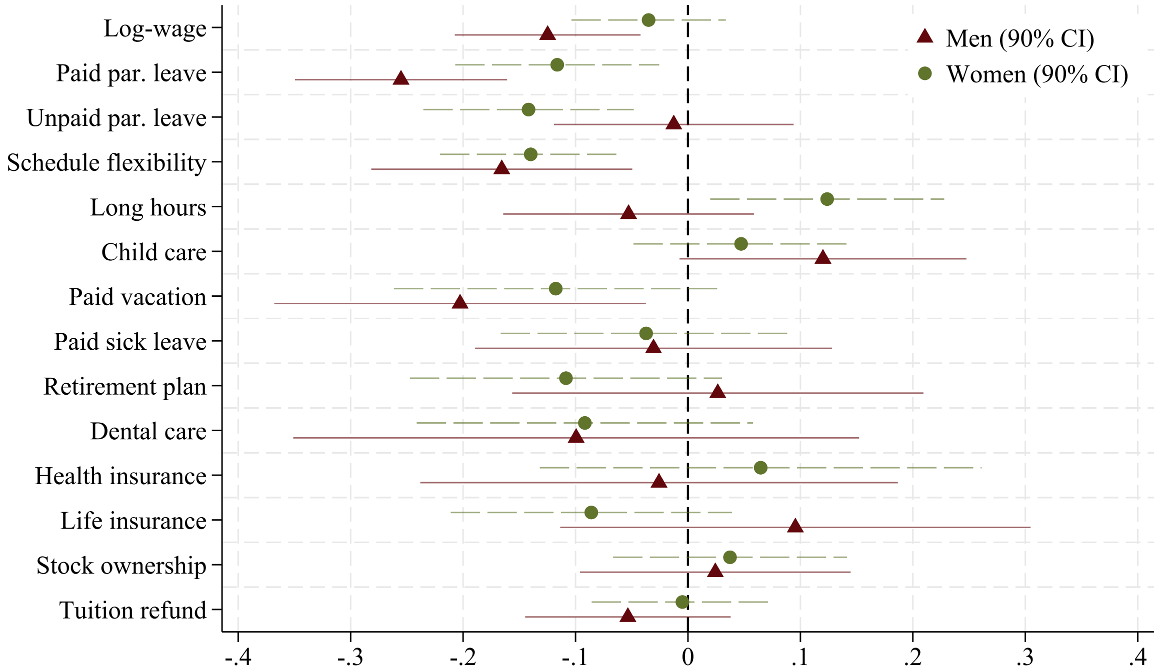
D.2 Benefits, work arrangements and job changes

Figure A6: Shares of employees who receive non-wage benefits



Notes: NLSY97, sample selection as in Table 1. The figures depict the share of men and women who report to work for an employer that ever offers a certain benefit or work arrangement during the worker's tenure, respectively, at the beginning of their first early-career job (lighter colors), and the end of their last early-career job (darker colors). Custom population weights applied.

Figure A7: Linear probability model of job changes - Selected coefficient estimates



Notes: NLSY97, sample selection as in Table 1. The figure reports selected coefficients of a fixed-effect linear probability model of job changes, and 90% confidence intervals. Each coefficient captures the difference in the average probability of changing employer in t between employees whose $(t-1)$-employer offered a benefit and employees whose $(t-1)$ employer did not provide it, controlling for the provision of other benefits, $(t-1)$ (log) wage, a quadratic in $(t-1)$ experience, occupation, industry, employer dimension, regional unemployment rate, and on the total number of weeks spent out of employment until $(t-1)$.

Figure A8: Linear probability model of job changes - Paid and unpaid leave coefficients with and without measurement error correction

Notes: NLSY97, sample selection as in Table 1. The figure reports selected coefficients of fixed-effect two linear probability models of job changes, and 90% confidence intervals. The estimated models include all the control variables included in regression 2. The two estimated models differ depending on whether raw or imputed (measurement-error corrected) variables capturing the provision of paid and unpaid parental leave at workers' current employer are used.

E Descriptive statistics and reduced-form analyses: workers without college degree

Table A5: Time-invariant sample characteristics - workers without college degree

	Men	Women	Difference p-Value
Age at labor market entry	21.40	21.96	0.000
Graduate degree by labor market entry	0.00	0.00	.
African American	0.14	0.16	0.011
Marries/cohabits by labor market entry	0.23	0.37	0.000
Marries/cohabits by 3^{rd} yr in labor market	0.52	0.67	0.000
Marries/cohabits by 6^{th} yr in labor market	0.64	0.76	0.000
Has child by labor market entry	0.13	0.30	0.000
Has child by 3^{rd} yr in labor market	0.25	0.45	0.000
Has child by 6^{th} yr in labor market	0.34	0.56	0.000
Age at first childbirth	24.46	22.77	0.000
N	907	912	

Notes: National Longitudinal Survey of Youth 1997 (NLSY97), Rounds 1 to 15. The sample includes workers who enter the labor market between 2000 and 2011 and who do not have a Bachelor's degree (or higher degree) by labor market entry. All individuals in the sample have non-missing observations for demographic characteristics and for wages, work hours, and employer and job characteristics throughout the first six years of labor market experience (early career). The number of observations in the table refers to the number of female (912) and male (907) unique individuals in the final samples, observed during the first week in employment at labor market entry. All workers are subsequently observed for six years (72 months, 512 weeks).

Table A6: Time-varying sample characteristics - workers without college degree

	Men	Women	Difference p-Value
	(a) Labor market history		
Total n. of years employed	5.10	5.00	0.151
Tot n. spells out of work	2.65	2.56	0.687
Tot n. weeks out of work	46.92	51.96	0.151
Total n. of jobs held	2.11	2.04	0.133
Changes employer by 6th year in labor market	0.62	0.60	0.439
Year of experience first job change	3.21	3.37	0.014
	(b) Outcomes – first early-career job		
Average weekly hours worked	40.29	39.59	0.001
Weekly hours > 40	0.11	0.06	0.000
Total n. of weeks employed in t	44.85	46.74	0.000
Hourly rate of pay (in 2005 US Dollars)	11.11	10.16	0.000
Hourly pay – Executive/Managerial	10.89	9.59	0.035
Hourly pay - Professional	13.11	13.40	0.880
Hourly pay - Sales/Office	10.98	9.52	0.000
Hourly pay – Ever changes job	10.13	9.26	0.001
Hourly pay – Never changes job	12.69	11.51	0.018
Employer with 1-49 employees	0.56	0.58	0.199
Employer with 100+ employees	0.31	0.32	0.817
Employer with 500+ employees	0.11	0.13	0.080
	(c) Outcomes – last early-career job		
Average weekly hours worked	39.86	37.64	0.000
Weekly hours > 40	0.17	0.11	0.000
Total n. of weeks employed in t	47.78	45.44	0.000
Hourly rate of pay (in 2005 US Dollars)	14.42	13.35	0.003
Hourly pay – Executive/Managerial	12.95	11.65	0.204
Hourly pay - Professional	19.35	17.19	0.088
Hourly pay - Sales/Office	14.32	12.91	0.001
Hourly pay – Ever changes job	13.46	12.48	0.031
Hourly pay – Never changes job	15.99	14.65	0.024
Employer with 1-49 employees	0.53	0.52	0.941
Employer with 100+ employees	0.33	0.37	0.299
Employer with 500+ employees	0.11	0.14	0.166
N	907	912	

Notes: NLSY97. Sample selection as in Table A5. Panels (b) and (c) refer to workers observed, respectively, during the first week employed in their first early-career job, and during their last week employed in their last early-career job.

Table A7: Time-varying sample characteristics - workers without college degree and no kids during early career

	Men	Women	Difference p-Value
	(a) Labor market history		
Total n. of years employed	5.10	5.25	0.019
Tot n. spells out of work	2.62	2.39	0.103
Tot n. weeks out of work	47.00	38.95	0.019
Total n. of jobs held	2.09	2.02	0.349
Changes employer by 6th year in labor market	0.60	0.58	0.387
Year of experience first job change	3.23	3.45	0.106
	(b) Outcomes – first early-career job		
Average weekly hours worked	40.02	39.71	0.224
Weekly hours > 40	0.10	0.08	0.230
Total n. of weeks employed in t	44.91	47.66	0.000
Hourly rate of pay (in 2005 US Dollars)	11.04	10.82	0.623
Hourly pay – Executive/Managerial	11.16	10.31	0.512
Hourly pay - Professional	12.34	14.03	0.212
Hourly pay - Sales/Office	10.89	9.85	0.063
Hourly pay – Ever changes job	10.12	9.49	0.183
Hourly pay – Never changes job	12.45	12.69	0.979
Employer with 1-49 employees	0.57	0.58	0.388
Employer with 100+ employees	0.30	0.29	0.349
Employer with 500+ employees	0.11	0.12	0.420
	(c) Outcomes – last early-career job		
Average weekly hours worked	39.60	38.67	0.148
Weekly hours > 40	0.17	0.14	0.269
Total n. of weeks employed in t	47.53	47.17	0.655
Hourly rate of pay (in 2005 US Dollars)	14.28	14.68	0.858
Hourly pay – Executive/Managerial	13.34	14.07	0.766
Hourly pay - Professional	18.42	18.19	0.770
Hourly pay Service/Sales/Admin Support	14.08	13.63	0.319
Hourly pay – Ever changes job	13.37	13.55	0.812
Hourly pay – Never changes job	15.68	16.26	0.896
Employer with 1-49 employees	0.54	0.51	0.662
Employer with 100+ employees	0.33	0.36	0.839
Employer with 500+ employees	0.11	0.13	0.598
N	572	365	

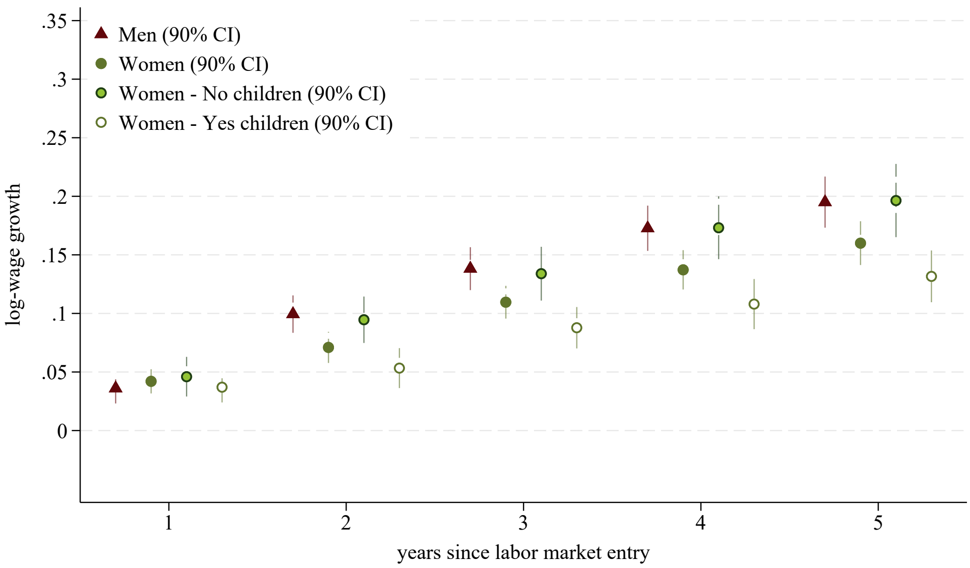
Notes: NLSY97. Sample selection as in Table A5. The table refers to the subsample of workers who do not have children throughout their early careers.

Table A8: Gender gaps in hours and weeks worked - workers without college degree

	Men v. Women no kids	Men v. Women not married	Do not change job Men v. Women	Change job Men v. Women
(a) Outcomes – first early-career job				
Average weekly hours worked	0.66***	0.82***	0.88***	0.72***
Weekly hours > 40	0.04***	0.05***	0.05***	0.05***
Total n. of weeks employed in t	2.38***	-1.94***	-1.09***	-1.54***
(b) Outcomes – last early-career job				
Average weekly hours worked	1.14***	1.70***	0.63***	2.83***
Weekly hours > 40	0.03***	0.07***	0.04***	0.07***
Total n. of weeks employed in t	0.35***	0.85***	2.24***	1.84***
N	1272	1161	702	1117

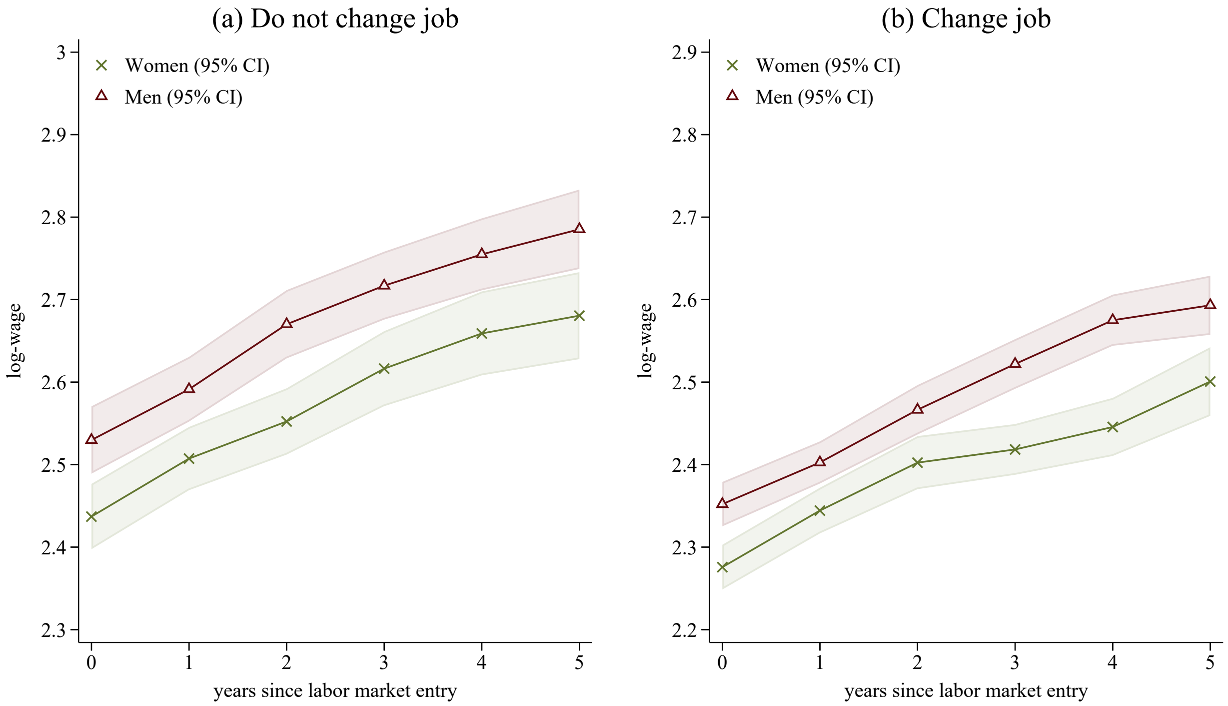
Notes: NLSY97. Sample selection as in Table A5. The first two columns compare all men to, respectively, women who do not have children by the sixth year of labor market experience, and women who are neither married nor cohabit by the same year. The last two columns restrict the sample to, respectively, men and women who do not change employer during their early careers, and men and women who change at least one employer during the same time period.

Figure A9: Estimated returns to experience - workers without college degree



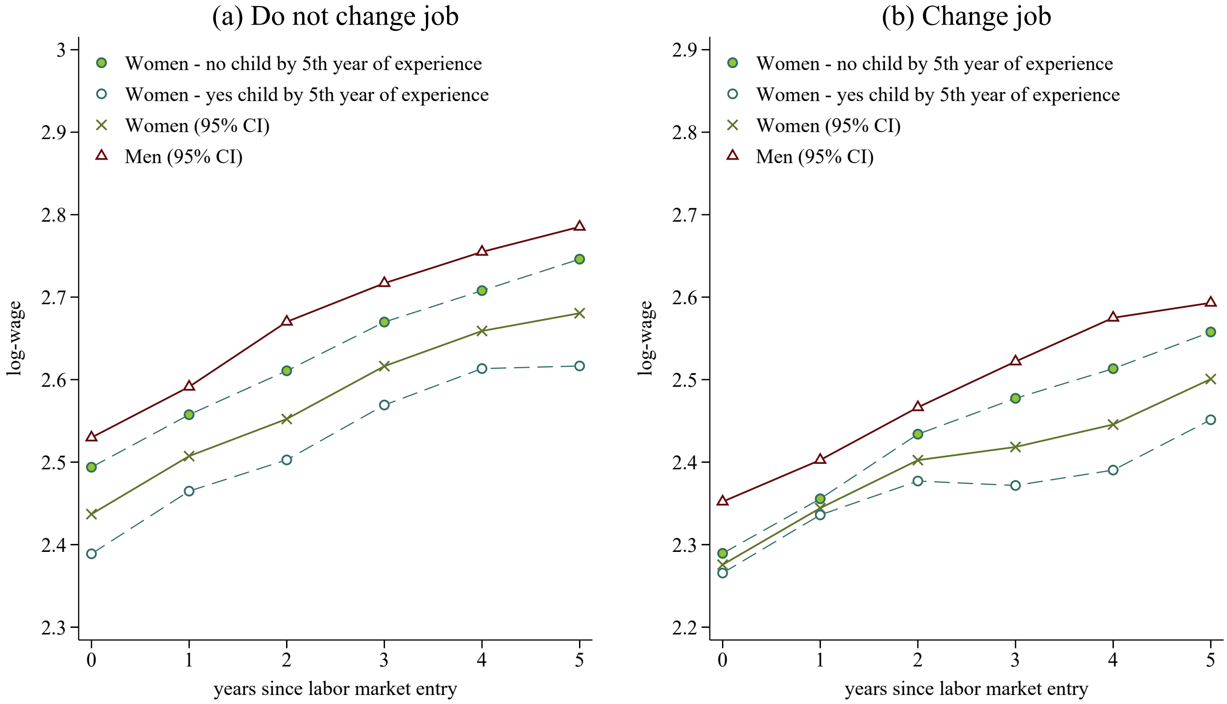
Notes: NLSY97, sample selection as in Table A5. The figure depicts the returns to experience estimated in log-wage regressions through fixed-effect estimator, and clustering standard errors at the individual level. The regressions take the following form $w_{i,t,k} = \alpha + \sum_{k=0}^{5} \beta_k \mathbb{I}\{\text{year in sample} = k\} + \delta_t + x'_{i,t,k}\gamma + \epsilon_i + u_{i,t,k}$.

Figure A10: Experience wage profiles - workers without college degree



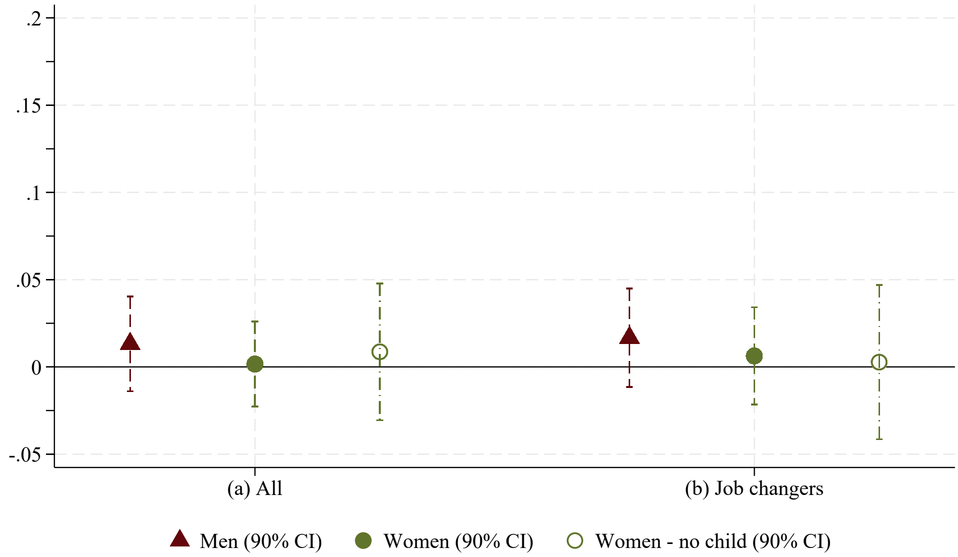
Notes: NLSY97. Sample selection as in Table A5. The figure depicts the average log-wage profiles of men and women. Panel (a) focuses on workers who do not change employer throughout their early career. Panel (b) focuses on workers who change at least one employer throughout their early career.

Figure A11: Experience wage profiles - workers without college degree, women with and without children



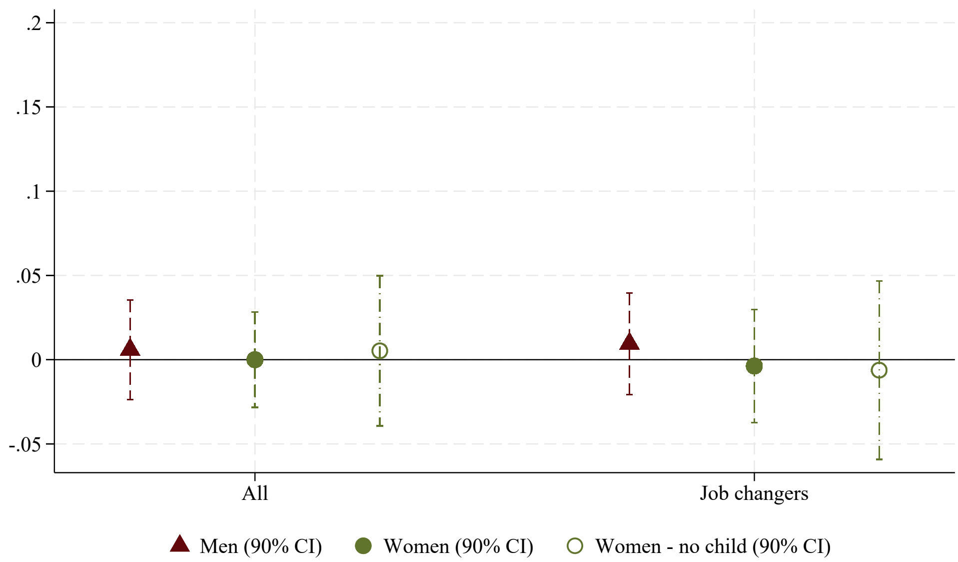
Notes: NLSY97. Sample selection as in Table A5. The figure depicts the average log-wage profiles of men and of, respectively, all women, women who do not have children by five years since labor market entry (44% of women), women who do have at least one child by the end of their early career (56% of women). Weights applied.

Figure A12: Estimated wage gains following first early-career job change - workers without a college degree



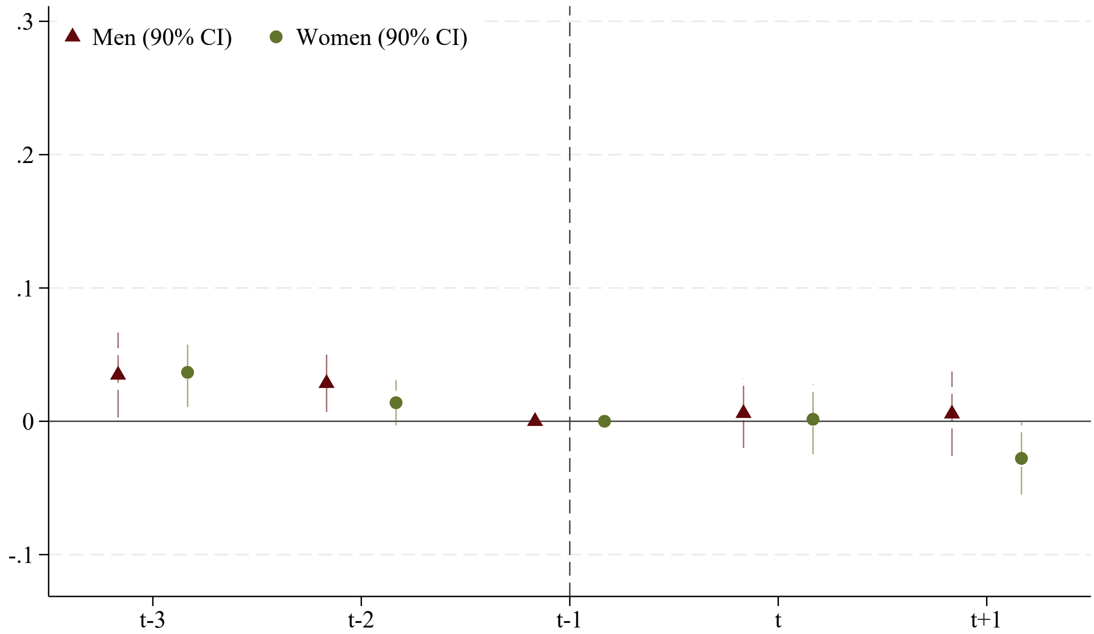
Notes: NLSY97, sample selection as in Table A5. Panel (a) shows the δ coefficients estimated, separately by gender, in the full sample of workers. Panel (b) shows the δ coefficients estimated on the subsample of workers who change at least one employer throughout their early career. The coefficients are estimated through fixed-effect estimator, standard errors are clustered at the individual level. The estimated regression controls for the same variables listed in the notes to figure 3.

Figure A13: Estimated wage gains following first early-career job change - workers without college degree - weighted regression results



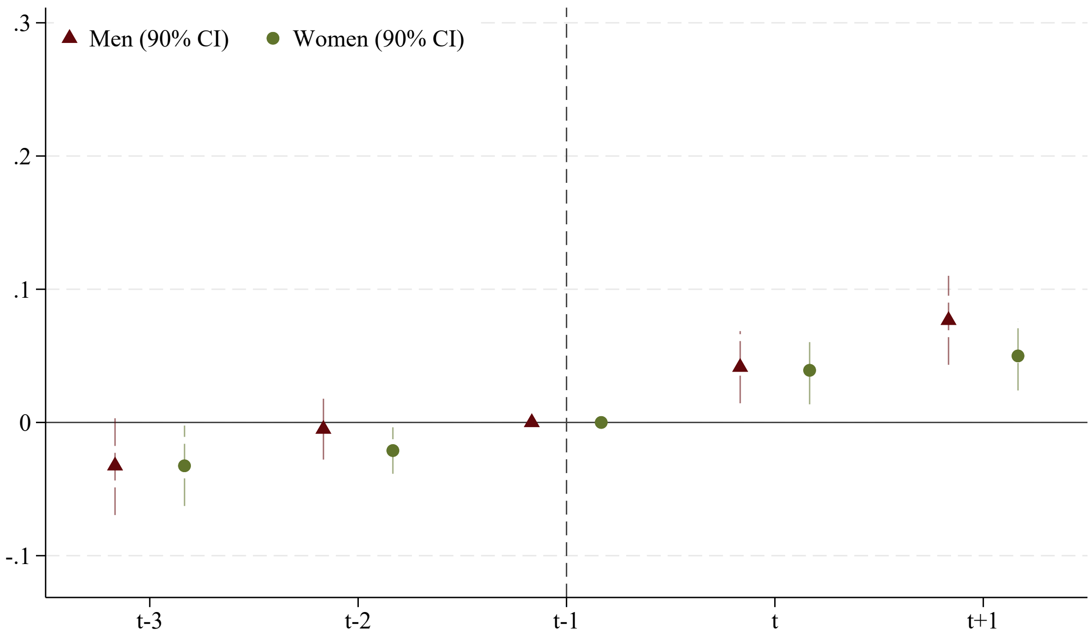
Notes: NLSY97, sample selection as in Table A5. Panel (a) shows the δ coefficients estimated, separately by gender, in the full sample of workers. Panel (b) shows the δ coefficients estimated on the subsample of workers who change at least one employer throughout their early career. The coefficients are estimated through fixed-effect estimator, standard errors are clustered at the individual level and custom population weights are applied. The estimated regression controls for the same variables listed in the notes to figure 3.

Figure A14: Event study estimates of wage gains from first job change - full sample - workers without college degree



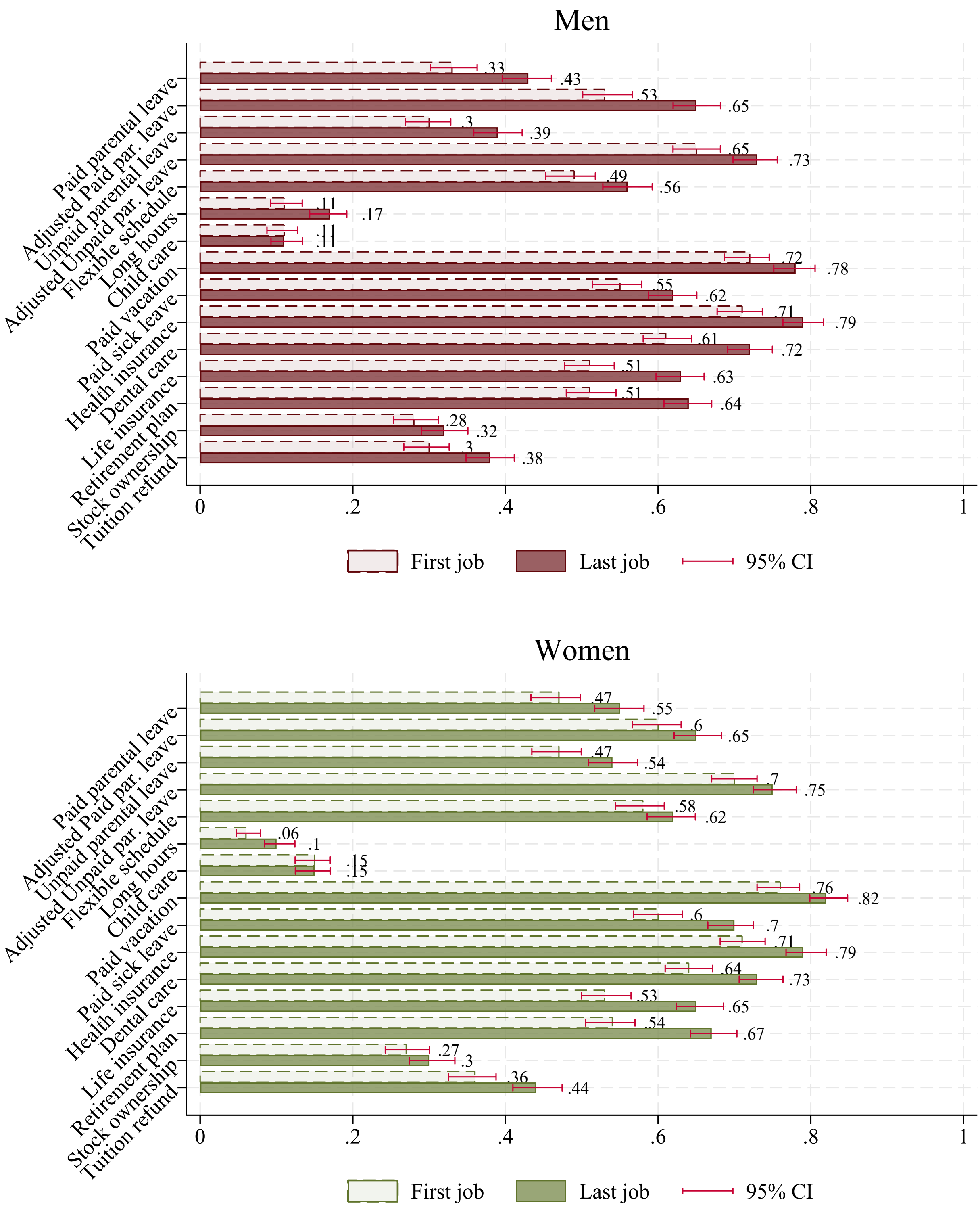
Notes: NLSY97, sample selection as in Table A5. The figure depicts the $\delta_{\tau \pm d}$ coefficients in regression 1. A control dummy variable coded as $\mathbb{1}\{t = \tau + 100\}$ is included to maintain workers who do not change jobs during their early careers as a control group in the regression. Thus, the coefficients β_t capture the time-trend in wages among workers who do not change job throughout their early career.

Figure A15: Event study estimates of wage gains from first job change - workers who change at least one job during their early career - workers without college degree



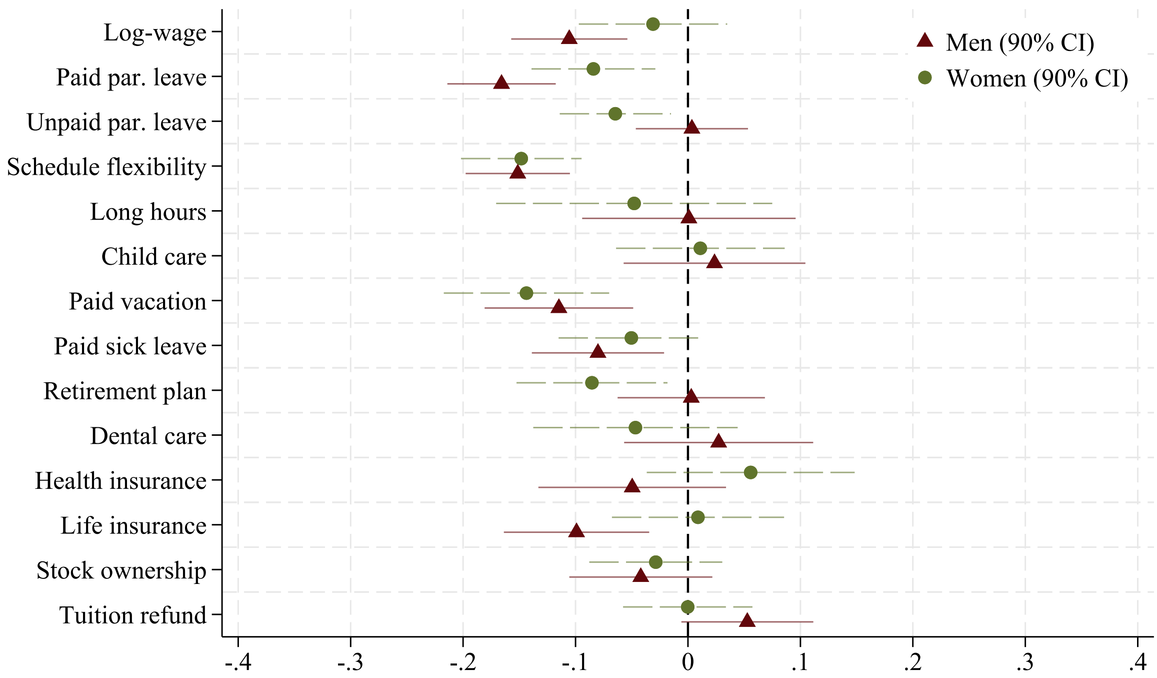
Notes: NLSY97, sample selection as in Table A5. The figure depicts the $\delta_{\tau \pm d}$ coefficients in regression 1.Workers who do not change at least one job during their early-careers are excluded. Thus, the coefficients β_t capture the time-trend in wages among workers change at least one job during their early career but for whom $t \neq \tau \pm d$.

Figure A16: Shares of employees who receive non-wage benefits - workers without college degree



Notes: NLSY97, sample selection as in Table A5. The figures depict the share of men and women who report to work for an employer that ever offers a certain benefit or work arrangement during the worker's tenure, respectively, at the beginning of their first early-career job (lighter colors), and the end of their last early-career job (darker colors). Custom population weights applied.

Figure A17: Linear probability model of job changes - selected coefficient estimates - workers without college degree



Notes: NLSY97, sample selection as in Table A5. The figure reports selected coefficients of a fixed-effect linear probability model of job changes, and 90% confidence intervals. Each coefficient captures the difference in the average probability of changing employer in t between employees whose $(t-1)$-employer offered a benefit and employees whose $(t-1)$ employer did not provide it, controlling for the provision of other benefits, $(t-1)$ (log) wage, a quadratic in $(t-1)$ experience, occupation, industry, employer dimension, regional unemployment rate, and on the total number of weeks spent out of employment until $(t-1)$.

Figure A18: Linear probability model of job changes - paid and unpaid leave coefficients with and without measurement error correction - workers without college degree

Notes: NLSY97, sample selection as in Table A5. The figure reports selected coefficients of fixed-effect two linear probability models of job changes, and 90% confidence intervals. The estimated models include all the control variables included in regression 2. The two estimated models differ depending on whether raw or imputed (measurement-error corrected) variables capturing the provision of paid and unpaid parental leave at workers' current employer are used.

F Measurement error in paid and unpaid parental leave variables

In this section, I explain how I tackled possible measurement error (and gender differences in measurement error) in measures of paid and unpaid parental leave in the NLSY97. I also explain the imputation that I use to correct potential measurement error in paid and unpaid parental leave variables. Throughout my paper, I rely on imputed measures of paid and unpaid parental leave. In this Online Appendix, I show that the most important results in my paper are qualitatively unaffected when using raw measures of paid and unpaid parental leave.

F.1 Benefit availability across surveys

In the NLSY97, the availability of non-wage benefits at workers' current workplace is self-reported by employees. In this section, I verify whether NLSY97 employees' answers are consistent with evidence about benefits availability collected from other surveys: the

National Compensation Survey (NCS) of the Bureau of Labor Statistics, and the Job Search Supplement of the Survey of Consumers Expectations (SCE) of the Federal Reserve Bank of New York.In this section, I use publicly available NLSY97 data rather than NLSY97 data merged with restricted-access geocodes. The NLSY97 data used here only differ from the final sample used in my paper in that the data in this section do not exclude individuals with missing information on state of residence (which is not observed in publicly available NLSY97 data).

The National Compensation Survey is an establishment-based survey conducted annually by the Bureau of Labor Statistics. NCS respondents do not answer a unified questionnaire. Rather, BLS field economists collect information through conversation with firm officials and by accessing firm documents (BLS, 2017). The BLS annually releases summary tables that estimate the availability of benefits among different categories of workers and establishments using NCS data, while NCS microdata are not publicly available.[2] Information is collected on the following types of benefits: healthcare benefits, retirement plans, life insurance coverage, short-term and long-term disability insurance coverage, paid leave (including paid family leave), unpaid family leave, health promotion benefits, financial benefits, pretax benefits and quality of life benefits.

The BLS does not disclose information regarding non-response rates and imputation rates, but provides detailed documentation that explains the methods used to calculate benefit availability and to impute missing values from non response in NCS microdata. NCS data are collected at the establishment level and, for every establishment, BLS economists collect information on workers in a limited number of randomly selected occupations: from up to 4 in small establishments (less than 40 employees), to 8 in establishments with at least 250 employees. Having identified occupations, BLS economists define a quote as a group of workers who, within the same establishment, are in the same occupation, and have the same bargaining status, full-time or part-time status, work level and type of pay. Information on benefits is then collected in such a way that, if a worker in a quote within an establishment has access to a benefit, all workers in the same quote

[2]Using NCS microdata, the BLS calculates the shares of individuals for whom non-wage benefits are available among: the entire civilian workforce, all private sector workers, for workers by category of establishment size, by full-time and part-time status, by average-wage, by Census region, by aggregate occupation and industry groups.

are considered to have access to the same benefit. This feature, also noted by Goldin, Kerr, and Olivetti (2020), implies that NCS data do not allow to estimate whether the availability of benefits differs across demographic groups of workers. Regarding missing values, when a company official participates in the survey but does not provide answers to certain inquiries, the values for the variables of interest are imputed. For example, if information on the availability of a benefit in a certain quote is missing, the availability of the benefit of interest is imputed using responses about benefit availability recorded among similar occupations in establishments with similar characteristics.[3]

As far as paid and unpaid parental leave are concerned, the Bureau of Labor Statistics defines them in their glossary of terms (BLS, 2024). Importantly, unpaid family/parental leave is defined in accordance to the provisions of the Family and Medical Leave Act of 1993.

The Job Search Supplement of the Survey of Consumers Expectations takes place annually since October 2013 on a random cross section of SCE respondents.[4] The Job Search Supplement was designed and initially administered by Jason Faberman, Andreas Mueller, Ayşegül Şahin and Giorgio Topa, and a complete description of the survey design can be found in Faberman, Mueller, Şahin, and Topa (2022). As the Survey of Consumers Expectations, the Job Search Supplement is nationally representative of 18 year-old and older heads of households. The Job Search Supplement asks detailed informaton on labor market transitions, and on the job search determinants and outcomes of both employed and unemployed workers. Importantly, for all currently employed workers, the survey asks information regarding the non-wage benefits available at their current workplace. The benefits whose availability is recorded in the survey are the following: traditional pension plan, employer contribution to a retirement account, health insurance, dental or vision insurance, health care or dependent care flexible spending account, housing or housing subsidy, life or disability insurance, commuter benefits, childcare assistance, stocks, options, or other company equity (available since 2014), quality of life benefits (gym memberships, tuition reimbursement, etc.).

[3] Additional information regarding benefit-incidence calculations, reweighting and imputation can be found in the Handbook of Methods for National Compensation Measures (BLS, 2017).

[4] Detailed information regarding the Survey of Consumers Expectations can be found in Armantier, Topa, van der Klaauw, and Zafar (2017).

Because in the paper I study the first six years of labor market experience of NLSY97 workers who enter the labor market between 2000 and 2011, to compare access to benefits across surveys, I use NCS tabulations for all workers in the civilian workforce in year 2012. In the Job Search Supplement of the SCE, I retain currently employed workers between 16 and 64 years old who are not self-employed, who usually work between 1 and 112 hours per week, and with non-missing information regarding their occupation. I use cross sections from all survey years between 2013 and 2016. I then construct two samples: the first sample includes heads of household born between 1980 and 1984; the second sample includes all heads of household between 18 and 64 years old. The first sample, which I label "SCE - 1980-1984", includes individuals born in the same years as individuals in the NLSY97 survey. Regarding the publicly available NLSY97 data that I use in this section, I clean data following the same steps used to isolate the main sample of interest in my paper. Because information on states of residence is not observable in public available NLSY97 data, the sample used in this section cannot exclude individuals with missing data regarding the state of current residence. Furthermore, in this section I include workers of all levels of education unless otherwise noted. The years covered in the NLSY97 data I use go from 2000 to 2016.[5]

In terms of comparability across surveys, some aspects are worth noting. First, the survey samples are representative of different populations. The NLSY97 is nationally representative of the cohort born between 1980 and 1984. The samples in the Job Search Supplement of the Survey of Consumers Expectations are representative of 18 years-old and older heads of household. The tabulations using National Compensation Survey microdata produced by the Bureau of Labor Statistics are meant to be representative of the entire workforce. Given differences in the underlying populations, at least to some extent, differences in estimated access to benefits across samples may capture differences in the availability to benefits across different types of workers. Second, the National Longitudinal Survey of Youth 1997 only collects information on the availability of benefits among workers whose employment relationship lasts more than 14 weeks. Not measuring benefit availability for temporary workers, the NLSY97 is likely to overestimate the availability of certain benefits. Benefits such as health or life insurance, in fact, are less likely

[5]Using NCS tabulations from different years between 2000 and 2016 would not affect any results.

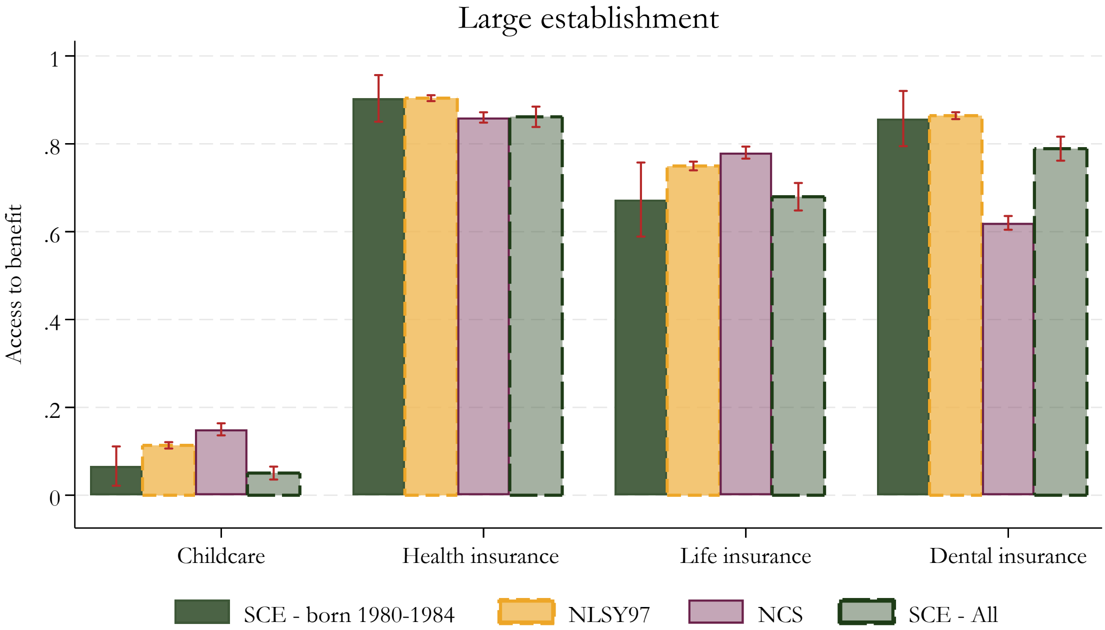
to be available to temporary workers and to workers in non-standard labor contracts (Berdahl & Moriya, 2021). Third, survey questions regarding benefit availability are not always comparable across surveys. For example, questions regarding health insurance, questions are similar between the NLSY97 and the Job Search Supplement of the SCE. NCS-based tabulations made available by the Bureau of Labor Statistics, however, record the availability "healthcare benefits" or "medical health benefits" depending on the survey year. Regarding life insurance, the benefit measured appears to be comparable between the NLSY97 and the NCS, while the SCE survey asks whether current employees' workplaces offer life or disability insurance. Similarly, questions regarding dental care benefits appear similar between the NLSY97 and the NCS, while the SCE asks whether dental or vision insurance is available. Finally, NCS respondents do not respond to a unified questionnaire, which creates some additional level of difficulty in comparing evidence arising from the NCS to evidence arising from different surveys.

The figures below show the cross-survey comparison in the availability of benefits recorded in all surveys, and whose definition is comparable across surveys. Figure A19 contains estimates for all workers, while Figure A20 splits workers in two groups based on workplace dimension. The figures show that, with the exception of dental care coverage,the availability of benefits appears to be remarkably similar across surveys. It is especially reassuring that the availability of benefits recorded from employees' responses in NLSY97 and Job Search Supplement data does not strongly differ from the availability of benefits estimated by the BLS using NCS establishment-based data. It suggests that the estimated incidence of benefits in employee-based surveys should not be strongly affected by measurement error. It is also worth noting that responses provided by NLSY97 individuals are typically comparable to the responses provided by workers of similar age in the Job Search Supplement of the SCE.

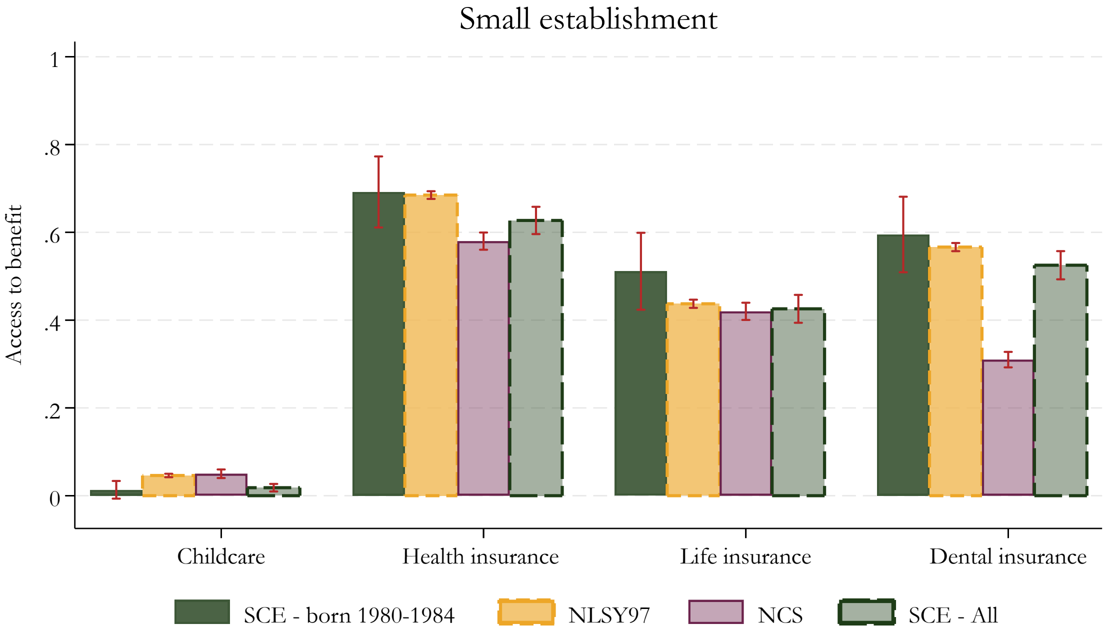
Figure A19: Access to benefits across surveys

Notes: SCE: Job Search Supplement of the Survey of Consumers Expectations, 2013-2016; NCS: National Compensation Survey, publicly available tabulations provided by the BLS for year 2012; NLSY97: National Longitudinal Survey of Youth 1997, sample selected as in Table 1 in the manuscript, with the exclusion of requirements regarding workers' education and non-missing values in state of residence. The latter piece of information is not available in public-use NLSY97 data. Sample weights are applied to estimates using the SCE and the NLSY97. For NLSY97, specifically, custom weights that account for the panel dimension of data are used.

Figure A20: Access to benefits by establishment size



Panel (a)



Panel (b)

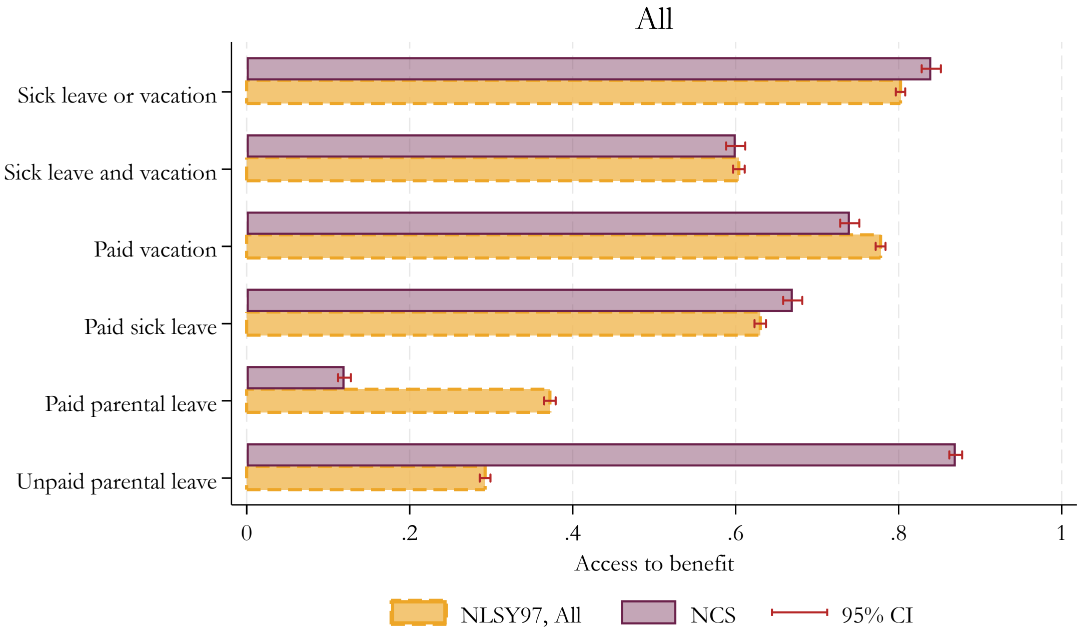
Notes: Samples as in Figure A19. Panel (a) contains estimates of benefits availability in large establishments (100+ employees). Panel (b) contains estimates of benefits availability in small establishments (less than 100 employees). The threshold for establishment sizes is so defined to ensure comparability across surveys. The NCS tabulations are available for establishments with less than 50 employees, less than 100 employees, and more than 100 employees (but not for all establishments with at least 50 employees). Establishment size in the Job Search Supplement of the SCE is recorded as a categorical variable whose first category includes establishments with 1-to-99 employees.

F.2 Availability of paid parental leave and unpaid parental leave in NLSY97 and NCS data

The Job Search Supplement of the Survey of Consumers Expectations does not collect information regarding workers' access to paid sick leave, paid vacation days, paid parental leave and unpaid parental leave. Information on access to these benefits, instead, is collected in the National Compensation Survey. The figures below compare access to the benefits listed above recorded in the NCS and NLSY97 samples described in the previous section. For NCS data, access refers to the share of employees for whom the benefit is available at their current workplace. For NLSY97 data, access refers to the share of workers who report that their employer currently offers them a certain benefit.

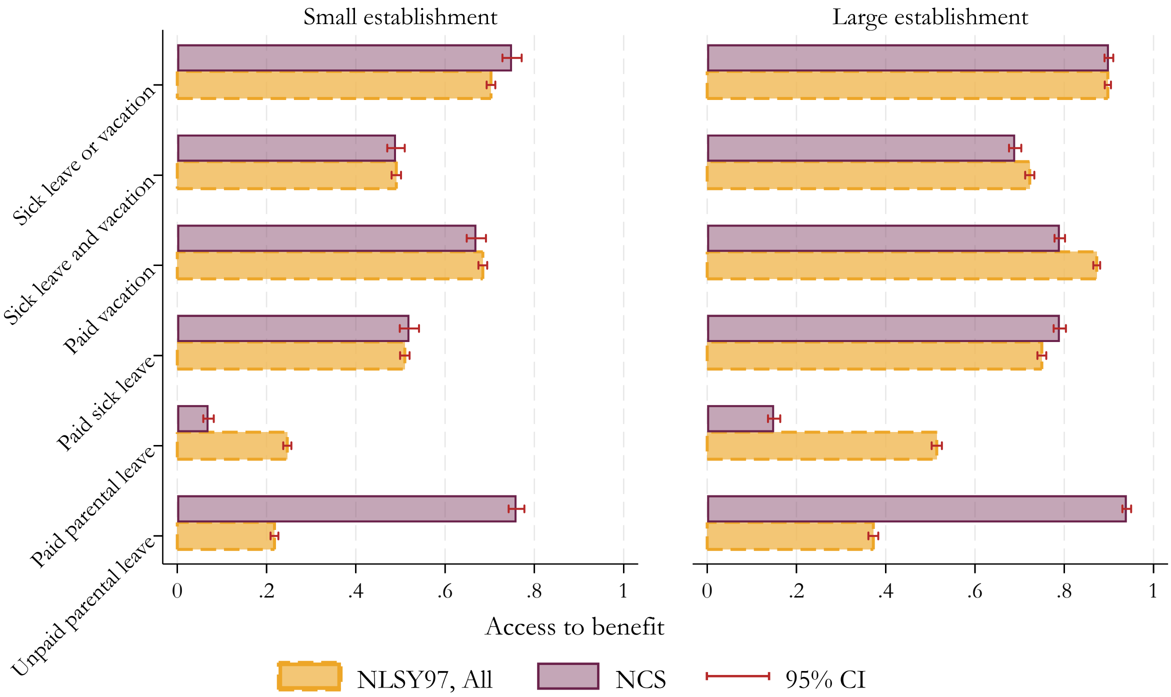
Figure A21 refers to all workers, Figure A22 separates employees working in establishments with more than 100 employees (right) from employees working in establishments with less than 50 employees (left). In this section only, I let the small-establishment threshold be 50 employees to isolate establishments that are less likely to be covered by the provisions of the Family and Medical Leave Act of 1993.

Figure A21: Access to benefits in NLSY97 and NCS



Notes: Samples as in Figure A19.

Figure A22: Access to benefits in NLSY97 and NCS by establishment size



Notes: Samples as in Figure A19. Small establishment: less than 50 employees. Large establishment: 100 or more employees. NCS tabulations for establishments with 50 or more employees are not available.

As figures A21 and A22 show, access to paid sick leave and vacation appear to be similar between the NLSY97 and the NCS. That is, the share of young employees who report having access to some days of paid sick leave and/or paid vacation in the NLSY97 sample I constructed is comparable to the national share of employees in the civilian workforce who are estimated to have access to the same benefits based on the National Compensation Survey tabulations.

As far as paid and unpaid parental leave are concerned, however, differences arise. As shown in Figure A21, the NCS estimates that 87% of employees in the civilian workforce had access to unpaid parental leave in 2012. Instead, slightly less than 30% of employed workers in the NLSY97 report that unpaid parental leave is currently available to them at their workplace. Regarding paid parental leave, it appears that the availability of this benefit is more prevalent in the NLSY97 sample than in the US workforce in similar years. As Figure A22 shows, differences between the NLSY97 and the NCS in estimated access to unpaid and paid parental leave remain when workers are split between small-establishment and large-establishment employees.

F.3 Addressing inconsistencies in unpaid parental leave availability between NLSY97 data and NCS data

In this section, I aim to uncover the extent to which the vast differences emerging in parental leave coverage between the NLSY97 and the NCS are due to measurement error in the variables collected in the NLSY97. Because the most concerning aspect in NLSY97 data is the apparent undercount of workers having access to unpaid parental leave, I will focus this benefit first.

First, it is worth noting that, in the NLSY97, differences in the availability of leave arise depending on how the availability of a benefit is defined. In figures A21 and A22, I report the shares of NLSY97 workers who report that leave is currently available to them at their current workplace. Considering unpaid parental leave in particular, there are several reasons why a worker may reply that the benefit is not available to them: the benefit is not provided by their employer (correct answer), the benefit is provided but the worker is not aware of its provision (incorrect answer), the benefit is provided but the worker is not currently covered because they are not eligible (correct answer), the employee works for an employer that offers paid leave but not unpaid leave (correct answer), the employee works for an employer that offers both FMLA-compliant unpaid leave and paid leave but requires workers who take FMLA leave to use their paid leave at that time (correct answer).

According to the Family and Medical Leave Act of 1993 (FMLA), 12 weeks of job-protected unpaid leave should be available each year to employees in most workplaces with at least 50 workers for the following qualifying reasons: (1) medical reasons related to pregnancy, (2) care of a newborn child or of a child placed for adoption or foster care, (3) serious medical condition of the employee or of their child, partner or spouse. Some details of the FMLA are worth noting.[6]

1. Not all employers with at least 50 employees are covered by the FMLA. Among private-sector employers, only employers who employ at least 50 employees in 20 or more workweeks in either the current calendar year or the previous calendar year

[6]For exhaustive information about the FMLA and its administration, please see the Family and Medical Leave Act website of the Department of Labor (Wage and Hour Division: United States Department of Labor, 2023).

are required to comply with the FMLA provisions.

2. Not all employees in FMLA-covered establishments are eligible for receiving FMLA-protected unpaid leave. To be eligible for taking unpaid leave under the FMLA, a worker

   2.a Must work at a location where the employer has at least 50 employees within 75 miles.

   2.b Must have at least one (possibly non consecutive) year of tenure at their current employer before their leave starts.

   2.c In the year prior to the beginning of leave, the employee must have worked at least 1250 hours for their current employer.

3. If an employer offers some form of paid leave (including paid parental leave, paid sick leave, paid vacation), three things can happen

   3.a Employers can offer paid time off on top of FMLA-covered unpaid leave

   3.b Employees can choose to take paid leave during FMLA leave.

   3.c Employers can require that workers use their paid leave during FMLA leave.[7]

These features of the FMLA are interesting for several reasons. First, based on employees' responses, the NLSY97 records whether an individual currently works in an establishment employing more than 50 employees. Yet, according to point 1. mentioned above, employing at least 50 employees is not a sufficient condition for an establishment to be covered by the FMLA. Thus, one should not expect that all NLSY97 employees currently working in sufficiently large establishments are offered unpaid parental leave, even assuming that all establishments fully comply with the act. Second, it is possible that some workers in the NLSY97 who report that unpaid parental leave is currently not available to them do not satisfy the eligibility criteria for receiving FMLA-protected unpaid leave mentioned in points 2.a, 2.b, and 2.c above. If any of these two circumstances occur, workers in the NLSY97 survey who state that unpaid parental leave is not

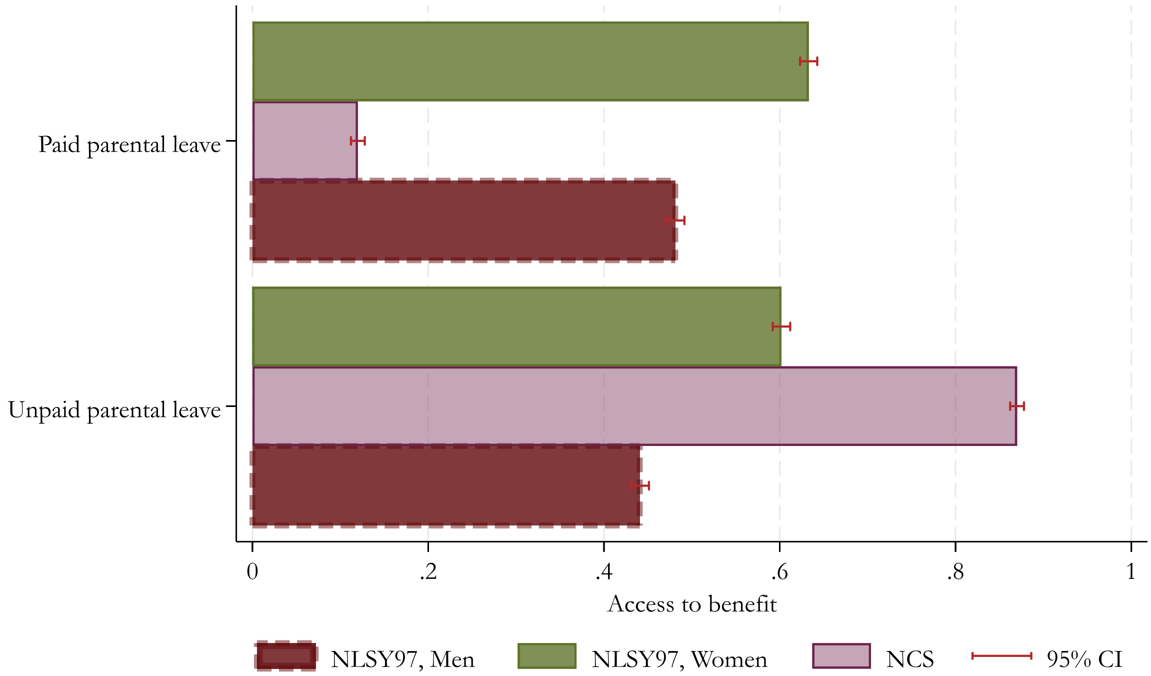
[7]See the Department of Labor Fact Sheet #28: the Family and Medical Leave Act (Wage and Hour Division: United States Department of Labor, 2023).

currently available to them would be correctly reporting that the benefit is not available to them, even though unpaid parental leave is offered at their current workplace.

To verify whether NLSY97 workers respond to survey questions regarding paid and unpaid leave based on their own eligibility status, in figures A23 and A24 I compare the availability of paid and unpaid leave in the NCS to the shares of NLSY97 employees who report to be offered, separately, paid parental leave and unpaid parental leave, at least once during their tenure at their current workplace. Because workers with longer tenure are more likely to satisfy FMLA eligibility requirements, this definition of leave availability should lead to an increase in NLSY97 estimated access to paid and unpaid parental leave.

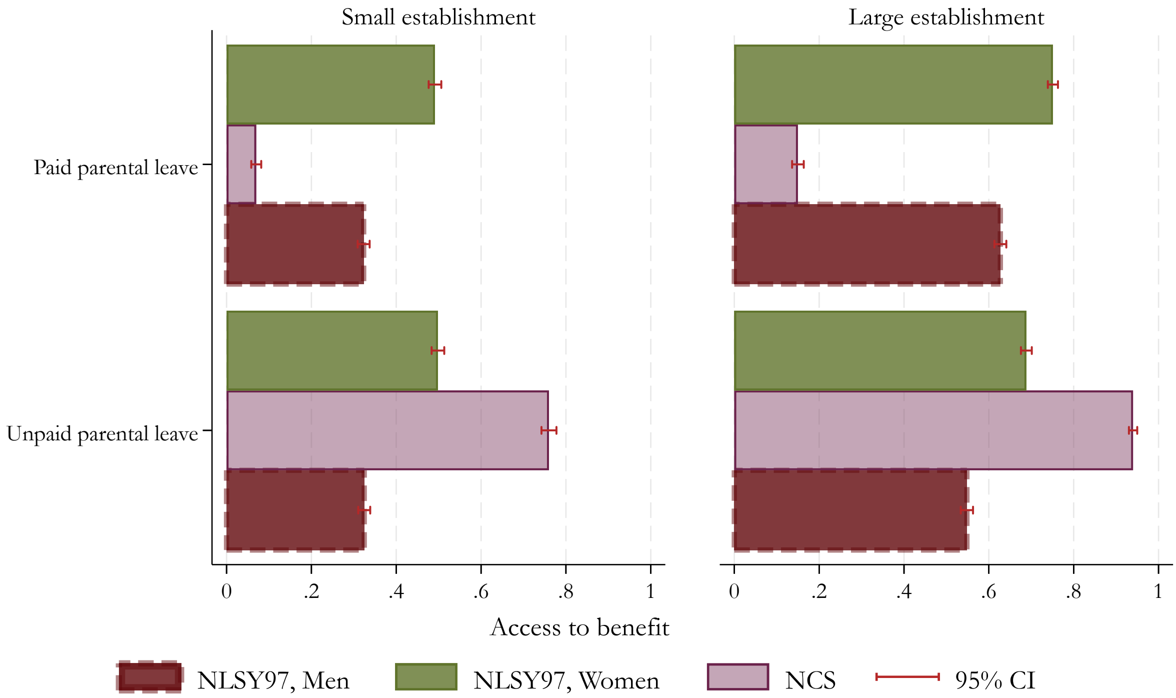
The figures show that, while substantial differences still exist in the estimated access to unpaid parental leave based on the NCS and on NLSY97 data, the difference is considerably smaller when changing the definition of benefit availability in NLSY97 data. Among workers in establishments of all size, less than 30% of NLSY97 individuals report to currently be offered unpaid parental leave, while around 43% of men and 60% of women report to be offered unpaid parental leave at least once during their tenure at their current employer. This evidence suggests that workers in the NLSY97 may be reporting availability of parental leave based on both the actual availability of the benefit and their own eligibility. As workers with longer tenure are more likely than newly hired workers to satisfy the FMLA eligibiliy criteria, they are also more likely to report that the benefit is available to them. Thus, defining both paid and unpaid parental leave to be available to NLSY97 workers if they ever report that the benefit is available to them during their tenure at their current workplaces considerably reduces the extent of underreporting of (especially unpaid) leave availability in NLSY97 data.

Figure A23: Access to benefits in NLSY97 and NCS - Men and Women



Notes: Samples as in Figure A19

Figure A24: Access to benefits in NLSY97 and NCS by establishment size - Men and Women



Notes: Samples as in Figure A19. Small establishment: less than 50 employees. Large establishment: 100 or more employees (NCS), 50 or more employees (NLSY97).

I further investigate the remaining discrepancy in unpaid leave availability between the NCS and NLSY97.

According to the NCS, as of 2012, more than 85% of workers in the civilian workforce were estimated to have access to the benefit. In the National Compensation Survey, information regarding access to this benefit is collected by BLS officiers inquiring regarding firms' compliance with the Family and Medical Leave Act of 1993. As previously noticed, however, an employer may comply with the FMLA (and, in fact, offer unpaid parental leave) even though "employees may use employer provided paid leave at the same time that they take FMLA leave if the reason they are using FMLA leave is covered by the employer's paid leave policy. An employer may also require an employee to use their paid leave during FMLA leave." (Wage and Hour Division: United States Department of Labor, 2023)

Employers providing paid leave can, in fact, require that workers use their entitled paid time off during FMLA-covered leave. Thus, if an employer complies with the FMLA, offers paid parental leave to their employees, and requires employees to use paid parental leave while on FMLA leave, employees may be correctly reporting that unpaid parental leave is, de facto, not offered to them at their current workplace. This implies that an employee in the NLSY97 who declares that paid parental leave is offered at the FMLA-covered establishment where they work, while unpaid parental leave is not, may be giving a correct answer.

To explore whether this feature of the FMLA may matter in explaining the difference in unpaid leave provision betweeen the NCS and the NLSY97, using the NLSY97 I define a variable capturing FMLA compliance. It takes value 1 if an employee works in an establishment where, at any point during their tenure, they are offered either paid parental leave, or unpaid parental leave, or both. I compute the share of employees in FMLA-compliant establishments on three groups of workers: newly hired workers (employees who have been working in an establishment for less than one year), incumbent workers (employees who have been working at their current workplace for at least one year), and FMLA-eligible workers (employees who have worked for at least one year at their current workplace and, in the previous calendar year, have worked for at least 1250 hours).

In NCS data, I use the share of employees who are estimated to be offered unpaid parental leave as a measure of employees in the US workforce who work in FMLA-compliant establishments. For NLSY97 data, I produce results separately for men (burgundy bars) and women (green bars).

Panels (a) and (b) in Figure A25 report the result of this exercise. Both panels include two graphs: the left-hand side graph shows results for small establishments, and the right-hand side graph shows results for large establishments. Several aspects are worth noting. First, the share of workers who report to work in establishments that offer either paid or unpaid parental leave at any point during workers' tenure is higher among employees who are more likely to satisfy the FMLA eligibility criteria. The difference in the incidence of parental leave between new hires and incumbents is due to the answers of workers whose maximum tenure at one firm does not exceed one year.

Second, comparing workers of the same gender and within the same tenure categories, the share of employees who report to be offered either paid or unpaid leave is substantially higher when workers work in large establishments, that is, in establishments that are more likely to be covered by the FMLA provisions. This evidence is reassuring that answers provided by NLSY97 individuals regarding parental leave availability are in line with the fact that only large establishments are legally required to comply with the Family and Medical Leave Act.

Third, the difference in the reported availability of paid or unpaid leave between NLSY97 men and women in the same tenure categories is larger in small establishments (not covered by the FMLA) than in large establishments. Among FMLA eligible employees, in particular, 75% of women and 53% of men report to be offered either paid or unpaid leave in establishments with less than 100 employees, while 92% of women and 80% of men report to be offered either paid or unpaid leave in FMLA-covered establishments with at least 100 employees.[8]

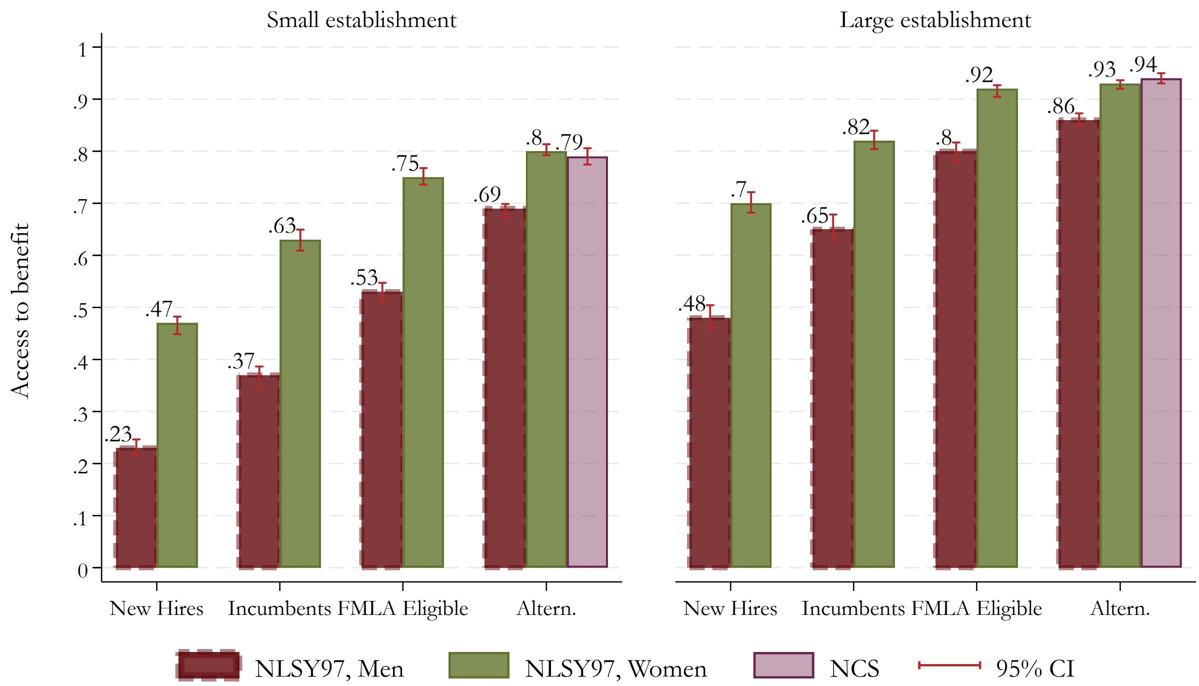
Finally, the share of both men and women who report to ever be offered either paid or unpaid leave during their tenure at their current workplace is highest among FMLA-eligible employees in both small and large establishments. Furhermore, gender differences in the incidence of parental leave is the smallest within this group.

It is worth noting that, because I define parental leave to be available in the NLSY97

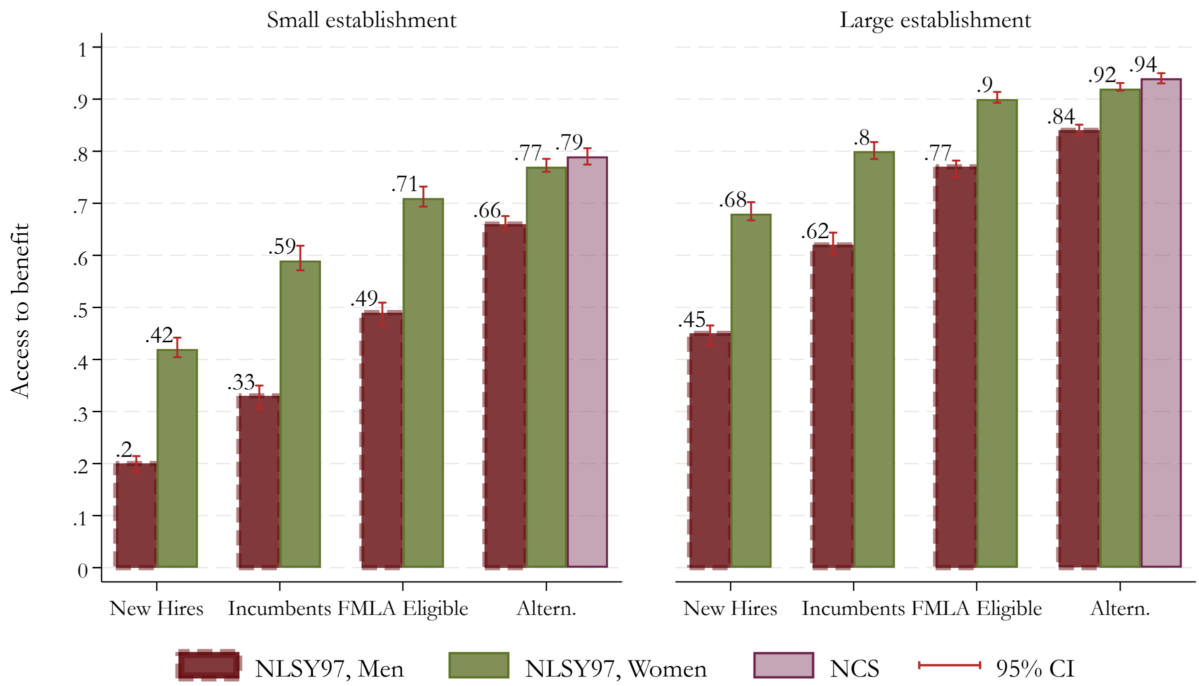
[8]As panel (b) shows, defining small establishments as establishments with less than 50 employees, the share of FMLA-eligible women who declare to be offered either paid or unpaid leave is 71%, and the share of men who report to be offered the same benefits is 49%.

if a worker ever reports paid or unpaid leave to be offered to them at any point during their tenure at their current workplace, the difference in the incidence of parental leave between new hires, incumbents, and FMLA-eligible workers is due to the answers given by workers who leave their current job before reaching the next tenure step. This fact can signal two possibilities: first, workers are less likely to be aware that parental leave is offered at their employer when they are not eligible for FMLA coverage. It implies that, for some of the new hires and incumbent employees who report that parental leave is not offered at their current establishment, the variable is measured with error. Second, workers who are employed in establishments that do not offer either paid or unpaid leave are more likely to leave their employer early during their tenure. It implies that workers are correctly reporting parental leave availability, but some establishments may not be complying with FMLA policies.

Figure A25: Share of individuals in FMLA-compliant establishments - Men and women



(a) Large-establishment threshold in NLSY97: 100 employees



(b) Large-establishment threshold in NLSY97: 50 employees

Notes: Samples as in Figure A19.

To distinguish between the two cases above, I perform one additional exercise. In the last three bars of the left-hand and right-hand panels in Figure A25, I show the share of workers in the civilian workforce who, according to NCS tabulations, worked in establishments offering unpaid leave in 2012. I compare this figure to an imputed measure of FMLA compliance which assumes that every FMLA-eligible NLSY97 worker in either large or small establishments is offered either paid or unpaid parental leave. The imputation grounds on the assumption that parental leave availability is misreported for the 20% of women and 31% of men among FMLA-eligible workers in small establishments, and the 8% of women and 14% of men among FMLA-eligible workers in large establishments, who do not report that parental leave is available to them.

According to this alternative measure of FMLA compliance (named "Altern." in the figure), the share of all women who are offered parental leave according to NLSY97 data is virtually identical to the share of all workers in the civilian workforce for whom parental leave is available according to NCS tabulations. This result suggests that, to the extent that it is plausible to assume that all FMLA-covered (large) establishments comply with all FMLA requirements, 8% of FMLA-eligible NLSY97 women and 14% of FMLA-eligible NLSY97 men employed in FMLA-covered establishments are likely mistakenly reporting that they do not have access (either currently, or throughout their tenure) at their current workplace. The results also imply that parental leave access in the NLSY97 is measured with error for around 51% FMLA-eligible men and 29% of FMLA-eligible women in establishments not covered by the FMLA.[9]

The imputation assumption used in this exercise, however, may be somewhat extreme. By assuming that all workers who are eligible for FMLA leave are ever offered either paid or unpaid leave during their tenure at their current workplace, I am in fact assuming full compliance with FMLA provisions among both establishments that are covered by the FMLA and establishments that are not covered by the FMLA (thus are not supposed to abide by it). This assumption, is most likely incorrect.

In their report on employees' FMLA awareness and employers' FMLA compliance, prepared for the Department of Labor, Klerman, Daley, and Pozniak (2012b) find the following evidence regarding FMLA implementation in 2012.

[9]It is worth noting that the residual differences in leave availability reported by men and women are driven by a higher share of men among individuals with less than one year of tenure at their current employer.

- Among worksites that were not covered by the FMLA in 2012, 75.6% allowed leave for "the care of a newborn" under all circumstances. Among employees in worksites that were not covered by the FMLA, 75.9% were estimated to work at worksites that allowed leave for the care of a newborn.

- Among worksites covered by the FMLA, 87.5% allowed leave for the care of a newborn. Among employees in FMLA-covered worksites, 97.5% were estimated to work at worksites that offered leave for the care of a newborn.

This evidence suggests that, while it may be plausible to assume full compliance with at least the FMLA provisions regarding parental leave for the care of a newborn child among establishments that are covered by the FMLA, it is not correct to assume, as I previously did, that all worksites that are not covered by the FMLA offer parental leave to their employees as soon as they become eligible for FMLA benefits. Consequently, while it can be plausible to assume that, for all FMLA-eligible workers in FMLA-covered workplaces who report to not have access to either paid or unpaid leave at their current workplace, leave availability is measured with error, the set of workers in establishments uncovered by the FMLA who mistakenly report that parental leave is not available at their current workplace is fundamentally unknown.

The imputation that I use to assign the availability of unpaid leave to workers directly addresses this problem. Before describing it, in the next section I discuss discrepancies in paid parental leave availability between the NLSY97 and the NCS.

F.4 Addressing inconsistencies in paid parental leave availability between NLSY97 data and NCS data

Regarding paid parental leave, the share of workers in the NLSY97 who report having access to it at least once during their tenure at their current workplace is substantially larger than the share of civilian workers who are offered paid family leave according to NCS tabulations (Figure A23). One reason for the discrepancy may be due to the difference in the definition of the benefit between the NLSY97 and the NCS. In the NLSY97, workers are asked whether they have access to paid parental leave, while NCS data report whether firms offer "paid family leave". It is possible that, while some firms do offer paid leave to parents of newborn children, they do not offer other types of paid family leave, thus leaving workers beyond childbearing age without access to paid leave. This may imply that younger workers, such as individuals in the NLSY97, are more likely to be offered paid family leave than the average worker in the US civilian workforce. In fact, in their recent study of firms' decision to provide paid family leave, Goldin, Kerr, and Olivetti (2020) find that employees working in firms that offer the benefit are younger than employees working in firms where the benefit is not available. Nevertheless, the discrepancy in access to paid parental leave between the NLSY97 and NCS data is large and worth of further investigation.

In order to verify the credibility of NLSY97 measures of paid parental leave availability, I compare them to evidence collected by Klerman, Daley, and Pozniak (2012b) in the 2012 Family Medical Leave report that the authors prepared for the Department of Labor. The report provides evidence on the implementation of the FMLA and on the availability of both unpaid and paid parental leave to US workers in 2012. The evidence provided is based on data from a worksite survey of 1,812 establishments, and on data from an employee survey of 2,852 workers, both designed and administered by Abt Associates.[10]

For year 2012, using their worksite survey, Klerman, Daley, and Pozniak (2012b) find that 53.5% of FMLA-covered worksites offered paid maternity leave to at least some workers (32.7% offer paid maternity leave to all workers), while 44.6% of FMLA-covered establishments offered paid paternity leave to at least some employees (23.2% offer paid paternity leave to all employees). Using weights to adjust for worksite dimension, the authors estimate that, in 2012, 69.3% of workers in FMLA-covered establishments worked in establishments offering paid maternity leave to at least some employees. Regarding paid paternity leave, 44.8% of employees in FMLA-covered establishments were estimated to work at worksites offering the benefit to at least some employees. Interestingly, to the extent that female workers may disproportionately select in establishments that offer maternity leave, while male workers may select in establishments that offer paternity leave, these numbers are in line with the fact that, as shown in Figure A24, around 75% of women and 60% of men in the NLSY97 report to have access to paid parental leave in large establishments between 2000 and 2016.

Considering worksites not covered by the FMLA, Klerman, Daley, and Pozniak (2012b)

[10]FMLA surveys were also administered in 1995, 2000 and 2018.

find that, in 2012, 21.7% of establishments in this category offered paid maternity leave to at least some employees, while 14.4% of these establishments offered paid paternity leave to at least some employees. In addition, the authors estimate that 37.6% of employees in worksites not covered by the FMLA were employed in worksites offering paid maternity leave to at least some workers, while 25.3% of employees worked in establishments providing paid paternity leave to at least some workers. As shown in Figure A24, around slightly less than 50% of women and around 36% of men in the NLSY97 report to have access to paid parental leave in large establishments between 2000 and 2016.

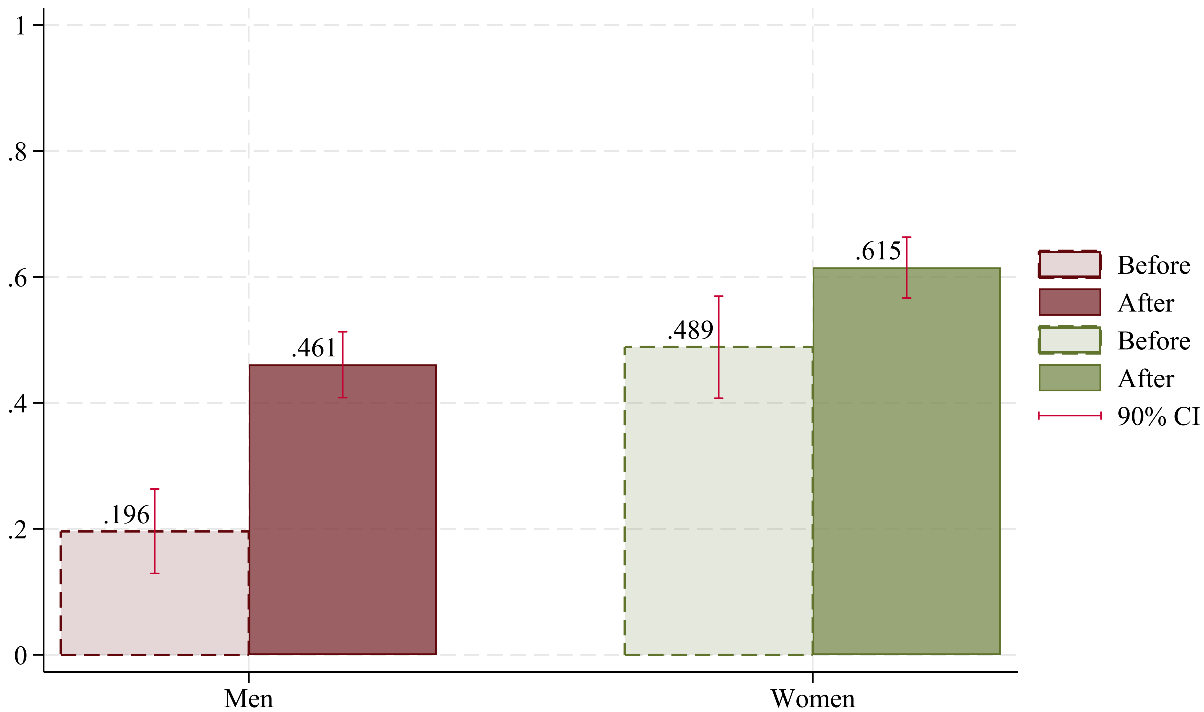
The incidence of paid parental leave observed by Klerman, Daley, and Pozniak (2012b) using the 2012 FMLA worksite survey is similar in magnitude to the incidence of paid parental leave recorded in NLSY97 data (see Figure A24), and it substantially differs from the availability of paid family leave recorded in NCS tabulations for the same year. Furthermore, the shares of workers employed in either FMLA-covered or FMLA-uncovered worksites that, according to the 2012 FMLA worksite survey, offered, respectively, paid maternity leave and paid paternity leave, are in line with the fact that more women than men in the NLSY97 report that paid parental leave is available to them.

This evidence suggests that estimates of gender differences in paid parental leave access arising in NLSY97 data are unlikely to be strongly affected by measurement error due to gender differences in workers' awareness regarding firm-level paid leave policies. To further support this point, it is worth noting that the estimates of paid parental leave availability in the Klerman, Daley, and Pozniak (2012b) report are not based on employees' responses, but on responses recorded in their worksite survey. Worksite respondents in the Klerman, Daley, and Pozniak (2012b) surveys are individuals in each worksite who are knowledgeable regarding FMLA administration Klerman, Daley, and Pozniak (2012a). Thus, estimates of paid maternity leave availability and paid paternity leave availability in the FMLA surveys are not affected by measurement error due to gender differences in employees' awareness regarding firm-level paid leave policies between employed men and women.

Finally, it is worth noting that workers in the sample I study appear to be aware of policy changes regarding paid parental leave, further supporting the hypothesis that a large number of workers in the NLSY97 know whether paid parental leave is available to them. Figure A26 reports that the shares of employees residing in California, Rhode

Island, or New Jersey who report that paid parental leave is available to them, are higher when employees work in jobs ending after the state-specific year of paid leave law implementation, than when they are employed in jobs ending before the implementation of paid parental leave laws.

Figure A26: Access to paid parental leave - California, Rhode Island, and New Jersey



Notes: NLSY97, sample selection as in Table 1. The figure compares the share of workers who report that paid parental leave is ever available to them in jobs that ended before the state-specific year of implementation of paid parental leave legislation in California, New Jersey, and Rhode Island, to the share of workers who ever report that paid parental leave is available to them, in the same states, in jobs that end after the implementation of paid parental leave legislation. Custom population weights applied to calculate the population estimates. Paid parental leave legislation was implemented in 2004 in California, in 2009 in New Jersey, and in 2014 in Rhode Island.

F.5 Imputing paid and unpaid parental leave availability

To summarize, in the previous sections I provided suggestive evidence that, while measures of paid parental leave in the NLSY97 are not necessarily affected by (gender differences in) measurement error, it is more plausible that some gender differences exist in workers' awareness about the availability of unpaid parental leave. In fact, 8% of FMLA-eligible women in FMLA-covered establishments and 14% of FMLA-eligible men in FMLA-covered establishments are likely to misreport that parental leave is not available to them. In addition, among the 51% of FMLA-eligible men and 29% of women in small establishments who report that parental leave is not available to them, some unknown share of workers may be mistakenly underreporting actual parental leave access.

Grounding on this evidence, I developed an imputation method to identify NLSY97 workers who plausibly report parental leave availability at their current workplace with error. The imputation works as follows. First, I assume that all FMLA covered establishments (50 or more employees) fully comply with FMLA provisions, thus offering either paid or unpaid parental leave. This assumption implies that all FMLA eligible workers in FMLA covered establishments who report to never have access to either paid or unpaid leave at their current workplace throughout their tenure are assumed to misreport parental leave availability. In the final sample of interest, 5% of women and 13% of men report leave availability with error (see Figure A28).

Second, I use the observable characteristics of workers in FMLA covered establishments who mistakenly report that parental leave is not available to them, to identify the unknown group of workers who misreport parental leave availability in establishments that are not covered by the FMLA (less than 50 employees) using a nearest neighbor matching estimator. Workers are matched based on their sex (exact match), and on the following characteristics: occupation and industry classes recurring most frequently during workers' early careers, US region of residence, education category, year of labor market entry, year, CAT-ASVAB test percentile, (log of) hours usually worked in the current year, whether the worker changes at least one job during their early career, whether the worker has at least one child by three years and six years since labor market entry, whether the worker marries or cohabits by three years or six years since labor market entry. Once workers are matched, I let workers in establishments not covered by the FMLA misreport whether they have access to parental leave if they never report to be offered either paid or unpaid leave during their tenure at their current establishment, and their nearest neighbors in the group of workers in FMLA-covered establishments mistakenly report that parental leave is never available to them throughout their tenure at their current workplace.

Third, having identified the potential set of workers who mistakenly report whether either paid or unpaid parental leave are available to them, based on observable characteristics, I match these workers to individuals of the same gender, with the same FMLA eligibility status, in establishments of the same dimension, who are imputed to correctly report whether parental leave is available to them. Finally, I assign whether, respectively, paid parental leave and unpaid parental are currently available to workers who misreport leave-availability based on the answers concerning the two separate benefits among the closest-neighbor workers who correctly report FMLA compliance.

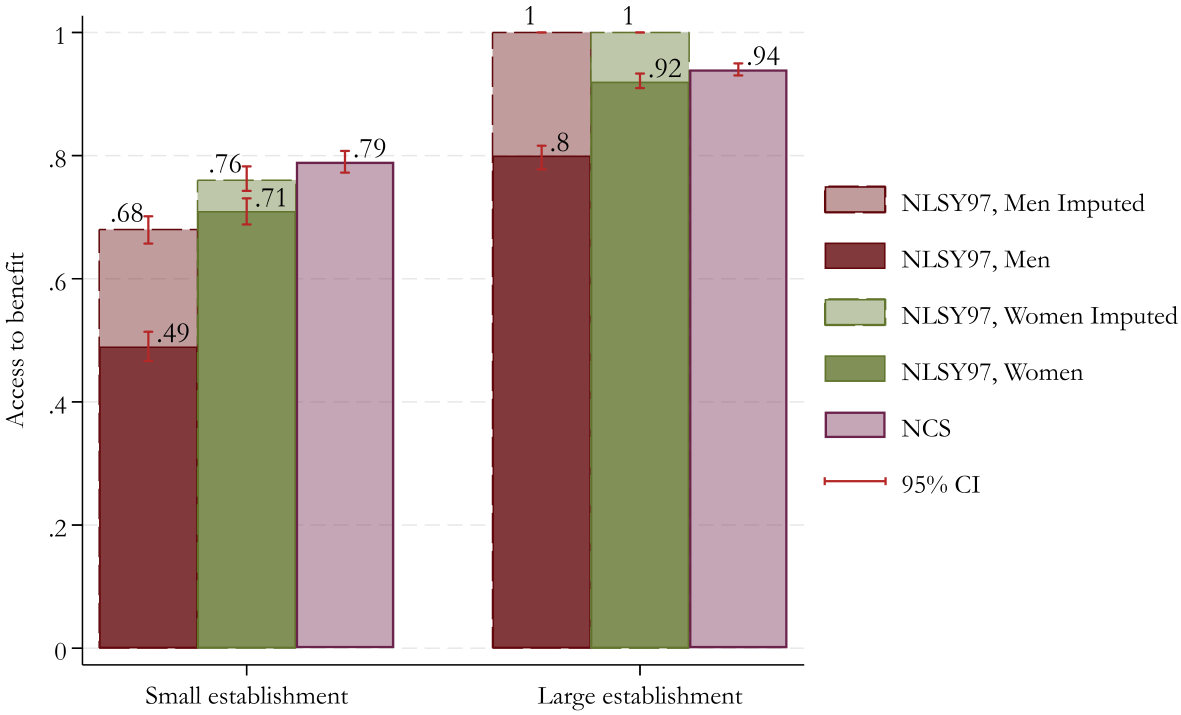
After imputing parental leave availability using the matching estimator explained above, I further impute that all workers in FMLA covered establishments have access to unpaid parental leave, and all workers in employed in jobs beginning from 2004 in California, from 2009 in New Jersey, and from 2014 in Rhode Island, have access to paid parental leave.

Finally, I define, respectively, paid and unpaid parental leave to be available at a worker's current workplace if, at least once during their tenure, a worker either reports that the benefit is available to them, or the worker is imputed to have access to the benefit.

In practice, the imputation imposes full compliance with FMLA provisions in establishments with at least 50 employees, and assumes that FMLA compliance (the availability of either paid or unpaid leave) is reported with error by all FMLA-eligible employees in FMLA-covered establishments who report to have never been offered either paid or unpaid leave at their current workplace. Then, using a nearest neighbor estimator I impute errors among workers who report that parental leave is not available to them, but for whom mistakes cannot be observed (i.e. workers in establishments that are not covered by the FMLA).

In Figure A27 I show the impact of the matching-estimator imputation only (i.e. before imposing that all establishments comply with the FMLA and with California, New Jersey, and Rhode Island paid leave policies) for the subsample of FMLA-eligible workers of all levels of education in the publicly available NLSY97 data. The imputation increases the share of NLSY97 male and female workers employed in workplaces that comply with FMLA provisions, irrespective of whether workplaces are covered or not by the FMLA. Because fewer women are likely to mistakenly misreport leave availability with respect to men, imputations have a larger impact on men than on women and reduce the gap in parental leave availability between women and men in small establishments from 22 percentage points to 8 percentage points. Nevertheless, the imputations applied do not make the incidence of parental leave equal across genders in establishments that are not covered by FMLA provisions.

Figure A27: Share of FMLA-eligible workers ever offered paid or unpaid leave at current workplace



Large-establishment threshold in NLSY97: 100 employees

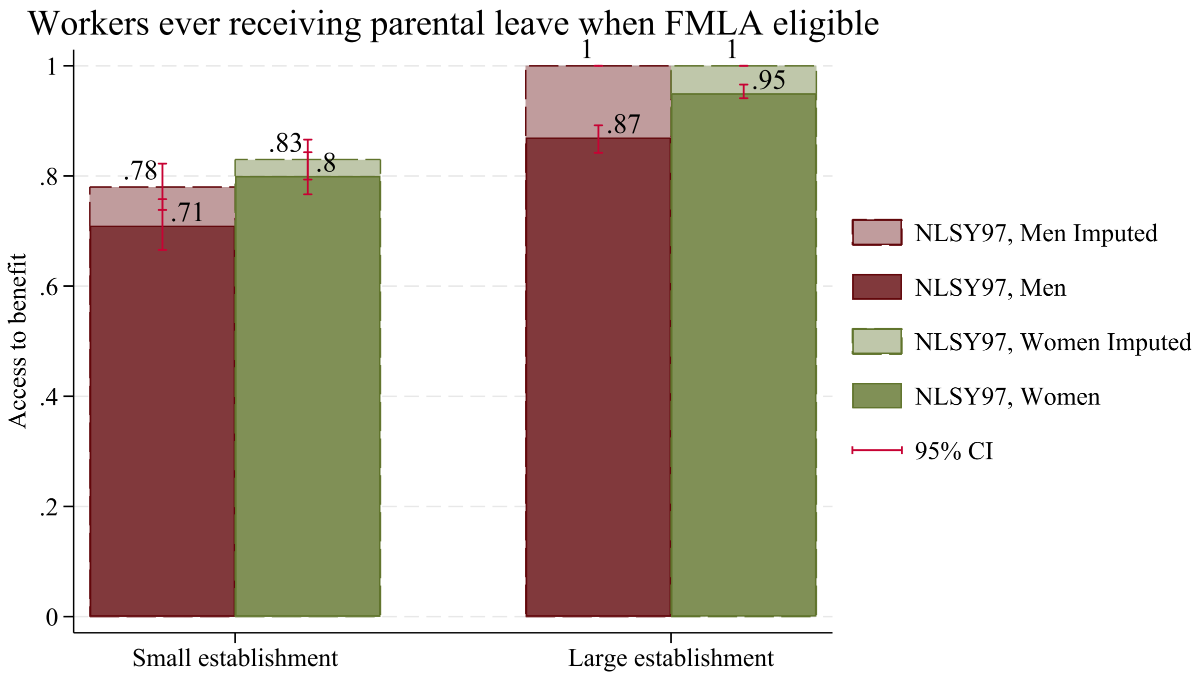
Notes: Samples as in Figure A19. Small establishment: less than 50 employees. Large establishment: 100 or more employee. The threshold is defined so that the figure can correctly include statistics from the NCS (statistics on benefit availability are not available in the NCS for establishments with 50 or more employees. Though not shown in the figure, NLSY97 data also include establishments employing 50 to 100 employees.

I argue that this aspect is likely to reflect the reality of leave provision more than a residual problem with measurement error in NLSY97 data. It is worth reminding that small establishments are not required to comply with FMLA provisions. Thus, two things are possible. First, small establishments with a larger share of female employees may be more likely to offer some form of parental leave to their employees. Second, at small workplaces, employers may not offer parental leave equally to male and female employees. According to Klerman, Daley, and Pozniak (2012b), while 77.6% of worksites not covered by the FMLA allowed leave for the care of a newborn in 2012, the share of small workplaces offering leave to care for a newborn irrespective of the type of legal or biological relationship between the caregiver and the child was 64.5%. This evidence suggests that one cannot exclude that some worksites not covered by the FMLA may offer some form of parental leave to mothers but not to fathers.

The following figures report the actual and matching-imputed incidence of parental leave (either paid or unpaid) among college graduate workers in the final sample of interest in my paper, before imposing that all establishments comply with the FMLA and with California, New Jersey and Rhode Island paid family leave policies. The sample differs from the broader NLSY97 sample used in the previous part of this section. First, the following figures use restricted-access geocoded NLSY97 data, and exclude workers for whom information regarding state of residence is missing. Second, the following figures only include workers who enter the labor market with at least a college degree. The sample used in the following figures coincides with the sample in Table 1.

Interestingly, gender differences in workers' answers regarding parental leave are substantially smaller among college graduate workers than among the full sample of workers. Furthermore, the share of workers of both genders who report to be employed in FMLA-compliant firms is higher among college graduate workers than among the full sample of workers. This evidence is in line with the idea that employers are more likely to offer leave when they expect workers to make substantial human capital investments before childbirth (Goldin, Kerr, & Olivetti, 2020). Furthermore, the evidence is somewhat reassuring that gender differences in measurement error in variables concerning parental leave provision are less likely to affect estimates of preferences for these benefits among college graduate workers than among workers of all levels of education. Using imputed measures of parental leave provision, gender differences in the share of FMLA-eligible workers who are ever offered either paid or unpaid leave during their tenure at their current workplace are small and only marginally statistically significant.

Figure A28: Actual and imputed access to leave - Final NLSY97 sample



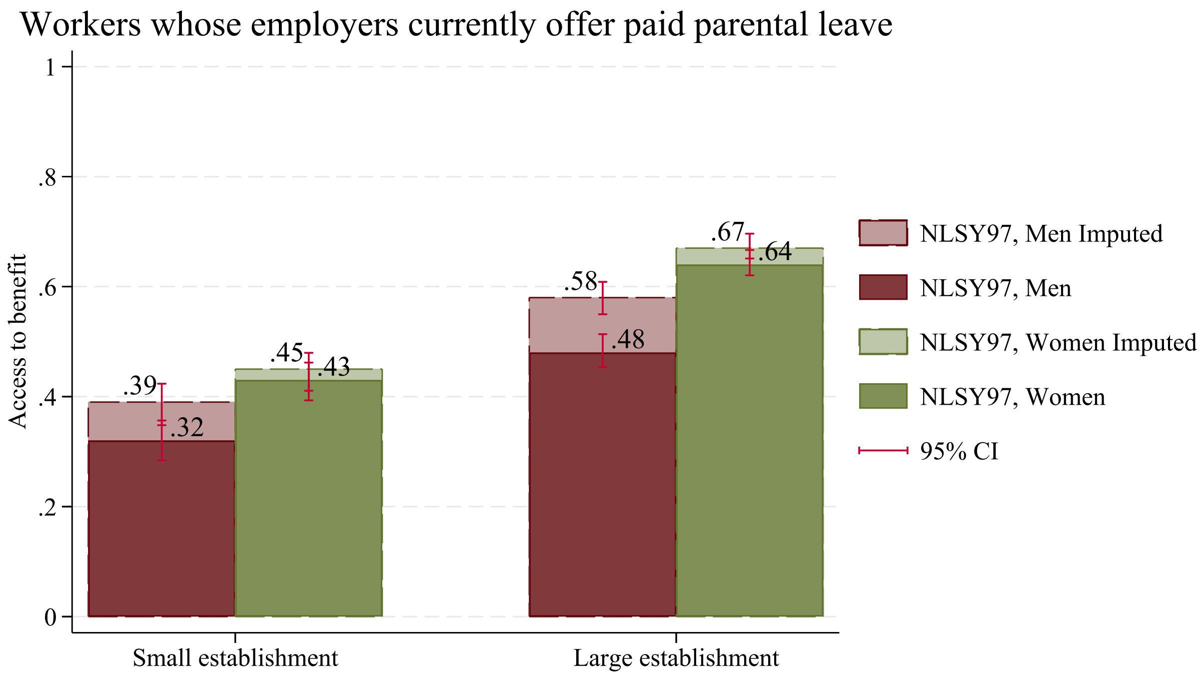
Notes: NLSY97. Samples as in Table 1. The threshold for large establishments is 100 employees.

The graphs below show the effect of the matching-estimator imputation that I use on the shares of male and female workers (at any level of tenure) who report to be *currently* offered, respectively, any parental leave, paid parental leave, and unpaid parental leave, in the final NLSY97 sample of interest in my paper.

Figure A29: Workers with college degree - FMLA compliance

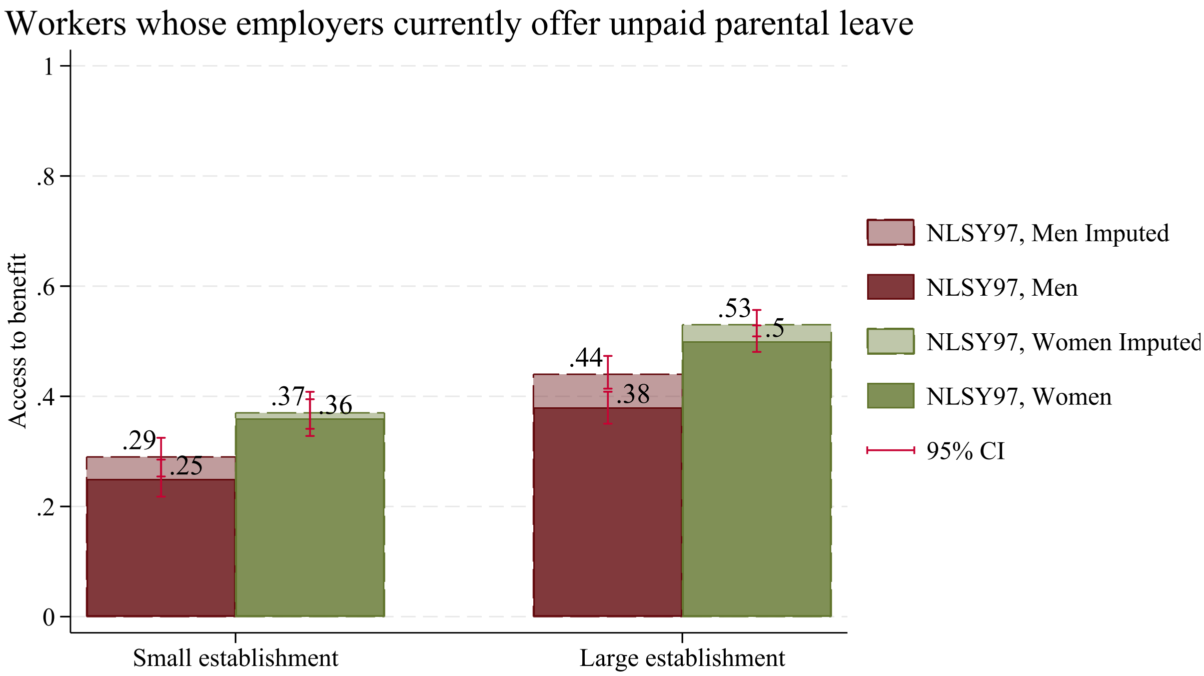
Notes: NLSY97. Samples as in Table 1.

Figure A30: Workers with college degree - Paid parental leave



Notes: NLSY97. Samples as in Table 1.

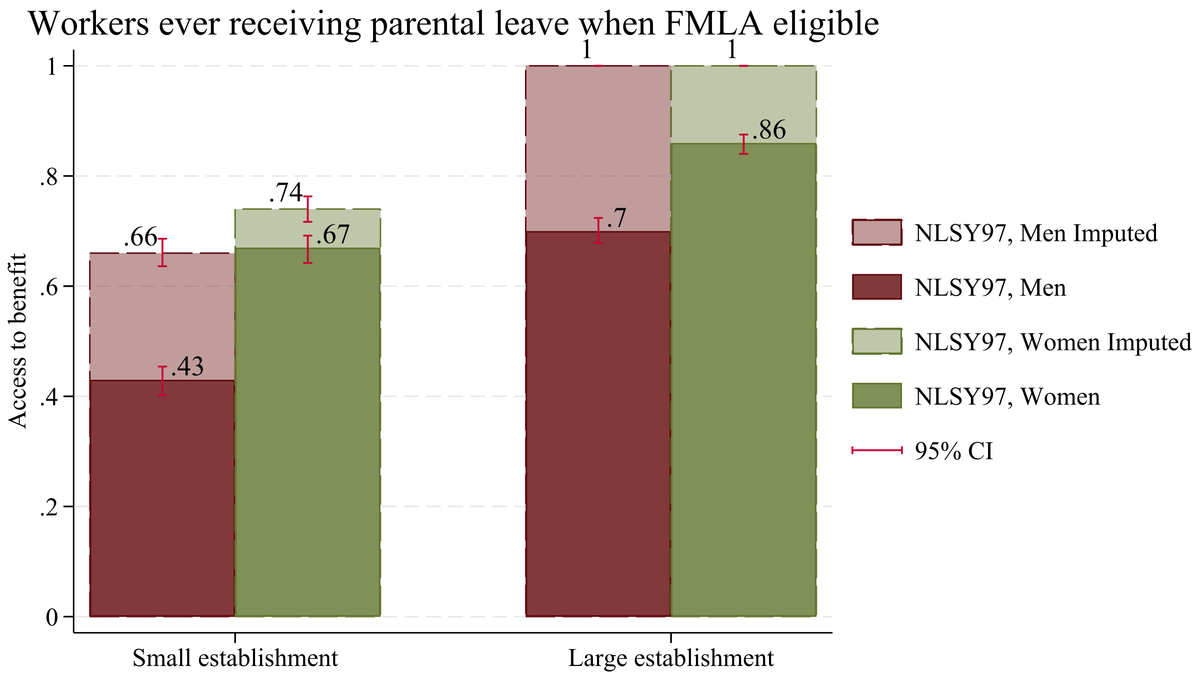
Figure A31: Workers with college degree - Unpaid parental leave



Notes: NLSY97. Samples as in Table 1.

For completeness, the graphs below show the impact of imputation of parental leave coverage on workers in the final geocoded NLSY97 sample of workers without a college degree

Figure A32: Workers without a college degree - FMLA compliance

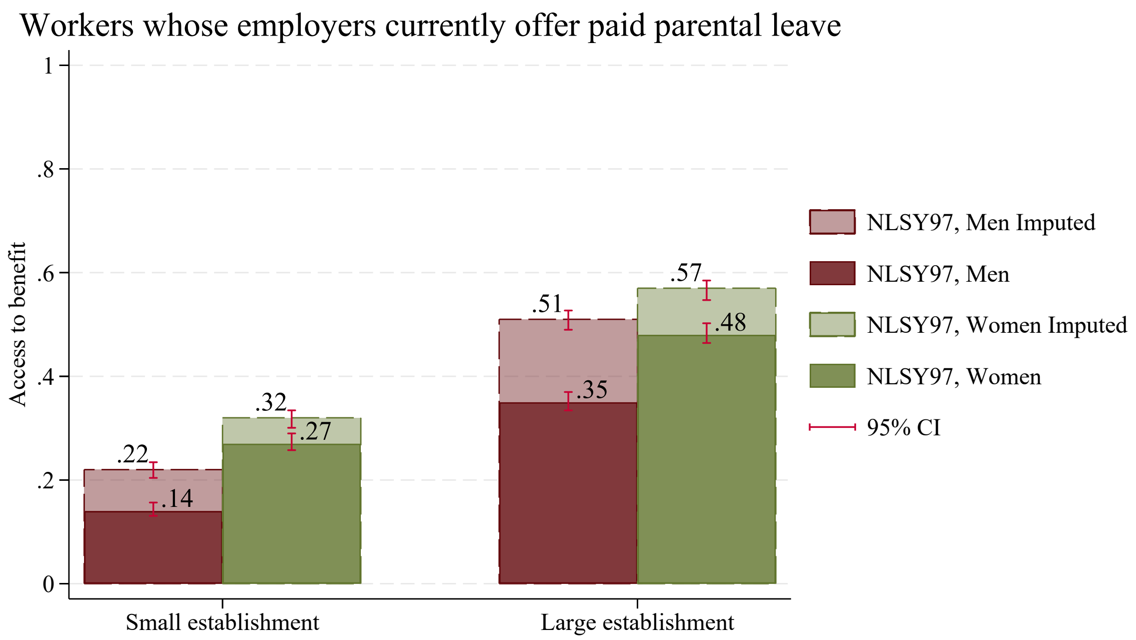


Notes: Samples as in Appendix Table A5.

Figure A33: Workers without a college degree - Any parental leave

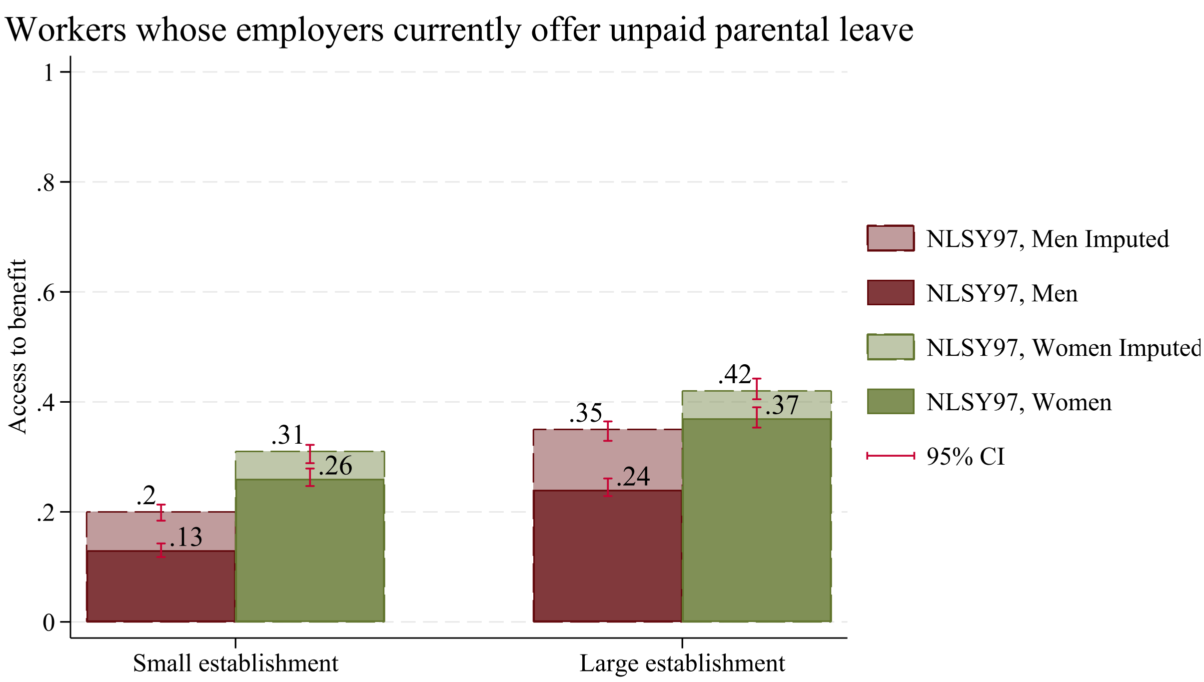
Notes: Samples as in Appendix Table A5.

Figure A34: Workers without a college degree - Paid parental leave



Notes: Samples as in Appendix Table A5.

Figure A35: Workers without a college degree - Unpaid parental leave



Notes: Samples as in Appendix A5.

The results of the full imputation is showed in Figure 4 in Section 2 of the paper for workers with a college degree and in Figure A16 in the Online Appendix for workers without a college degree. The imputation increases the share of workers who are assumed to be offered, respectively, paid and unpaid parental leave. As far as unpaid parental leave is concerned, the imputation substantially improves the similarity between leave access measured through NLSY97 data and leave access measured through NCS data.

Throughout my work, I use imputed measures of paid and unpaid leave availability. In Table A17 in the Online Appendix, I show that the main results included in my work are qualitatively unaffected when using raw measures of paid and unpaid parental leave availability.

G Model estimation and results

Table A9: Workers ever working in California, Rhode Island, New Jersey in jobs ending after paid parental leave laws implementation

	Share Pop. estimate using weights	Share Sample	N. Obs. men and women
Men	0.138	0.147	266
	(0.345)		
Women	0.170	0.172	379
	(0.376)		

Notes: NLSY97. Sample selection as in Table 1. Standard deviations in parentheses. Custom population weights applied to calculate the population estimates.

G.1 Characterization of the model steady state

Following Bonhomme and Jolivet (2009), the steady state distribution of wage-amenities bundles among employed workers can be found as follows. First, the steady-state probability that a worker leaves their job can be written as

$$P^g(\text{leave}|w_{i,j}, \mathbf{a}_{i,j}, b, c), pol = q^g + \lambda_2^g + \lambda_1^g \bar{F}_u^g(w_{i,j} + \delta^{g\prime}\mathbf{a}_{i,j}|b, c, pol) \qquad (2)$$

It is the sum of the employment loss probability, q^g, the constrained job-to-job transition probability, λ_2^g, and the probability that the worker receives a job offer yielding higher utility than the worker's current job, $\lambda_1^g \bar{F}_u^g(w_{i,j} + \delta^{g\prime}\mathbf{a}_{i,j})$.

Second, the steady-state flows of workers in and out of employment are equal, implying

$$\lambda_0^g U^g = q^g(1 - U^g) \tag{3}$$

Third, the steady-state flow of workers into jobs yielding utility at most as large as u must equal the flow of workers leaving these jobs. Hence, defining $G^g(.|b,c)$ the conditional distribution of jobs among employed workers of gender g given workers' ability and career, and $G_u^g(.|b,c)$ the observed distribution of utility levels among workers in the same group, the following equality must hold in steady state

$$\begin{aligned}\lambda_0 U F_u(u|.) + \lambda_2 F_u(u|.)(1-U)\bar{G}_u(u|.) = q(1-U)G_u(u|.) + \lambda_2 \bar{F}_u(u|.)(1-U)G_u(u|.) + \\ + \lambda_1 \bar{F}_u(u|.)(1-U)G_u(u|.)\end{aligned} \tag{4}$$

Where I dropped the superscript g to simplify notation. Equation (4) further implies that the steady-state cumulative distribution of utility levels among employed workers of gender g and ability b in career c is

$$G_u(u|.) = \frac{F_u(w_{i,j} + \delta' \mathbf{a}_{i,j}|.)}{1 + k\bar{F}_u(w_{i,j} + \delta' \mathbf{a}_{i,j}|.)} \tag{5}$$

Using (5), the density function of utility levels among employed workers is, thus,

$$g_u(u|.) = (1+k)\frac{f_u(u|.)}{[1 + k\bar{F}_u(w_{i,j} + \delta' \mathbf{a}_{i,j}|.)]^2} \tag{6}$$

Finally, using equation (6), the steady-state cross-sectional distribution of wages and amenities among employed workers is[11]

$$g(w, \mathbf{a}|.) = (1+k)\frac{f(w, \mathbf{a}|.)}{[1 + k\bar{F}_u(w + \delta' \mathbf{a}|.)]^2} \tag{7}$$

[11]As Bonhomme and Jolivet (2009) show based on previous results by Dey and Flinn (2005), equation (7) can be obtained using (6) and the equality in steady state between the distribution of wages and amenities offers conditional on utility, $f(w, \mathbf{a}|u)$, and the distribution of accepted job offers conditional on utility, $g(w, \mathbf{a}|u)$.

G.2 Functional forms for $f(w^*, \mathbf{a}^*|.)$ and $\bar{F}_u(u|.)$

In this section I show how to find the functional the functional forms for $f(w^*, \mathbf{a}^*|.)$ and $\bar{F}_u(u|.)$ needed to estimate the model.

First, the functional form for $f(w^*, \mathbf{a}^*|.)$ can be found as follows. Let $\varphi_0^w + \mu_1^w b_i + \mu_2^w pol_i + \sum_{c\in\{\text{ex, pr, ot}\}} \varphi_c^w \mathrm{c}_i = \mu^w(X)$, where $X = \{b, c, pol\}$. Notice that

$$f(w^*, \mathbf{a}^*|.) = f(w^*|\mathbf{a}^*, .)P(\mathbf{a}^*|.) = f(w^*|\mathbf{a}^*, .)\prod_{k=1}^{K} P(a_k^*|.) \tag{8}$$

To find an expression for $f(w^*|\mathbf{a}^*, .)$, notice that

$$\begin{aligned} F(w^*|.) &= P(\mu^w(X) + \rho'\mathbf{a} + \sigma_w\varepsilon_w \leq w^*) \\ &= P\left(\varepsilon_w \leq \frac{w^* - \mu^w(X) - \rho'\mathbf{a}}{\sigma_w}\right) \\ &= \Phi\left(\frac{w^* - \mu^w(X) - \rho'\mathbf{a}}{\sigma_w}\right) \end{aligned} \tag{9}$$

So that

$$f(w^*|.) = \frac{1}{\sigma_w}\phi\left(\frac{w^* - \mu^w(X) - \rho'\mathbf{a}}{\sigma_w}\right) \tag{10}$$

Where $\Phi(.)$ and $\phi(.)$ denote, respectively, the standard normal cumulative distribution function and the standard normal probability density function.

Regarding $P(\mathbf{a}^*|.)$, let $\mu_0^{a^k} + \mu_1^{a^k} b_i + \mu_2^{a^k} pol_i + \sum_{c\in\{\text{ex, pr, ot}\}} \varphi_c^{a^k} \mathrm{c}_i = \mu^{a_k}(X)$, where $X = \{b, c, pol\}$. Notice that every $a_k \in \{a^{fs}, a^{lh}, a^{pl}, a^{ul}\}$ takes value 1 if an employer offers amenity and 0 otherwise. Hence,

$$P(a_k^*|.) = p^{a_k^*}(1-p)^{1-a_k^*} \tag{11}$$

Where

$$\begin{aligned} p &= P(\mu^{a_k}(X) + \varepsilon_{a_k} > 0) \\ &= P(\varepsilon_{a_k} > -\mu^{a_k}(X)) \\ &= 1 - \Phi(-\mu^{a_k}(X)) = \Phi(\mu^{a_k}(X)) \end{aligned} \tag{12}$$

Consequently, for each amenity a_k

$$
\begin{aligned}
P(a_k^*|.) &= \Phi(\mu^{a_k}(X))^{a_k^*}(1-\Phi(\mu^{a_k}(X)))^{1-a_k^*} \\
&= \Phi\left(\mu^{a_k}(X)(-1)^{(1-a_k^*)}\right) \qquad (13)
\end{aligned}
$$

Substituting (10) and (13) in (8)

$$
f(w^*, \mathbf{a}^*|.) = \frac{1}{\sigma_w}\phi\left(\frac{w^* - \mu^w(X) - \rho'\mathbf{a}}{\sigma_w}\right)\prod_{k=1}^{K}\Phi\left(\mu^{a_k}(X)(-1)^{(1-a_k^*)}\right) \qquad (14)
$$

The functional form for $\bar{F}_u(u|.)$ can be found as follows. First, notice that

$$
\bar{F}_u(u|.) = \sum_{\mathbf{a}^*\in\{0,1\}^K}\bar{F}(u|\mathbf{a}^*,.)P(\mathbf{a}^*|.) \qquad (15)
$$

Where

$$
\begin{aligned}
\bar{F}(u|\mathbf{a}^*,.) &= 1 - P(w^* + \delta'\mathbf{a}^* \leq u|.) \\
&= 1 - P(\mu^w(X) + \rho'\mathbf{a}^* + \sigma_w\varepsilon_w + \delta'\mathbf{a}^* \leq u) \\
&= 1 - P\left(\varepsilon_w \leq \frac{-(\mu^w(X) + \rho'\mathbf{a}^* + \delta'\mathbf{a}^* - u)}{\sigma_w}\right) \\
&= 1 - \Phi\left(-\frac{(\mu^w(X) + \rho'\mathbf{a}^* + \delta'\mathbf{a}^* - u)}{\sigma_w}\right) \\
&= \Phi\left(\frac{(\mu^w(X) + \rho'\mathbf{a}^* + \delta'\mathbf{a}^* - u)}{\sigma_w}\right) \qquad (16)
\end{aligned}
$$

Substituting (16) and (13) into (15)

$$
\bar{F}_u(u|.) = \sum_{\mathbf{a}^*\in\{0,1\}^K}\Phi\left(\frac{(\mu^w(X) + \rho'\mathbf{a}^* + \delta'\mathbf{a}^* - u)}{\sigma_w}\right)\prod_{k=1}^{K}\Phi\left(\mu^{a_k}(X)(-1)^{(1-a_k^*)}\right) \qquad (17)
$$

G.3 The Bonhomme and Jolivet (2009) iterative estimation procedure - No unobserved heterogeneity

I explain here the sequential maximum likelihood estimation proposed by Bonhomme and Jolivet (2009). I implement the estimaton separately for male and female workers.

For every $t \in [0, T = 72]$, a worker's contribution to the likelihood in $(t+1)$ in equation 7 can be rewritten as

$$l_{t+1}(\theta,\lambda,\delta) = l_{1,t+1}(\theta) \times l_{2,t+1}(\theta,\lambda,\delta) \times l_{3,t+1}(\theta,\lambda,\delta) \tag{18}$$

Where

$$l_{1,t+1}(\theta) = f(w_{t+1}, a_{t+1}; \theta)^{uj_t} \tag{19}$$

$$l_{2,t+1}(\theta,\lambda,\delta) = [1-\lambda_1\bar{F}(w_t+\delta'\mathbf{a}_t;\theta)-\lambda_2-q]^{s_t}[\lambda_1\bar{F}(w_t+\delta'\mathbf{a}_t;\theta)+\lambda_2]^{jj_t} \tag{20}$$

$$l_{3,t+1}(\theta,\lambda,\delta) = q^{ju_t}[1-\lambda_0]^{uu_t}\lambda_0^{uj_t}\left[\frac{(\mathbf{1}\{w_{t+1}+\delta'\mathbf{a}_{t+1} > w_t+\delta'\mathbf{a}_t\}+\lambda_2)f(w_{t+1},a_{t+1};\theta)}{\lambda_1\bar{F}(w_t+\delta'\mathbf{a}_t;\theta)+\lambda_2}\right]^{jj_t} \tag{21}$$

The model parameters can be estimated as follows.

First, the wage-amenities outcomes of workers undergoing an unemployment-to-employment transition identify θ. Hence, the parameter vector describing the features of the job offers distribution is estimated as

$$\hat{\theta} = \text{argmax}_\theta \log L_1 = \text{argmax}_\theta \sum_{i=1}^{N}\sum_{t=t_0}^{T} \log l_{1,t+1} \tag{22}$$

Second, taking $\hat{\theta}$ as given, I guess an initial value $\tilde{\delta}$ for workers' preferences for amenities, and estimate

$$\begin{aligned}\hat{\lambda}^1 &= \text{argmax}_\lambda \log L_2 + \log L_3 = \\ &= \text{argmax}_\lambda \sum_{i=1}^{N}\sum_{t=t_0}^{T} \log l_{2,t+1}(\hat{\theta},\lambda,\tilde{\delta}) + \log l_{3,t+1}(\hat{\theta},\lambda,\tilde{\delta})\end{aligned} \tag{23}$$

Finally, taking $\hat{\theta}$ and $\hat{\lambda}^1$ as given, I estimate $\hat{\delta}^1$ as

$$\hat{\delta}^1 = \text{argmax}_\delta \log L_2 = \text{argmax}_\delta \sum_{i=1}^{N}\sum_{t=t_0}^{T} \log l_{2,t+1}(\hat{\theta},\hat{\lambda}^1,\delta) \tag{24}$$

I iterate the last two steps until convergence. In my estimation, five iterations are required to achieve convergence in the estimated δ and λ for both male and female workers. In the data I use, approximately 10 iterations are required for the estimation to converge, for both male and female workers. The likelihood function I estimate, includes all months

$t \in (1, 72)$.

G.4 Parameter identification using the Bonhomme and Jolivet (2009) method

As shown in the previous section, the identification of the parameters of interest comes from the additive separability of the log-likelihood function in the parameters of interest.

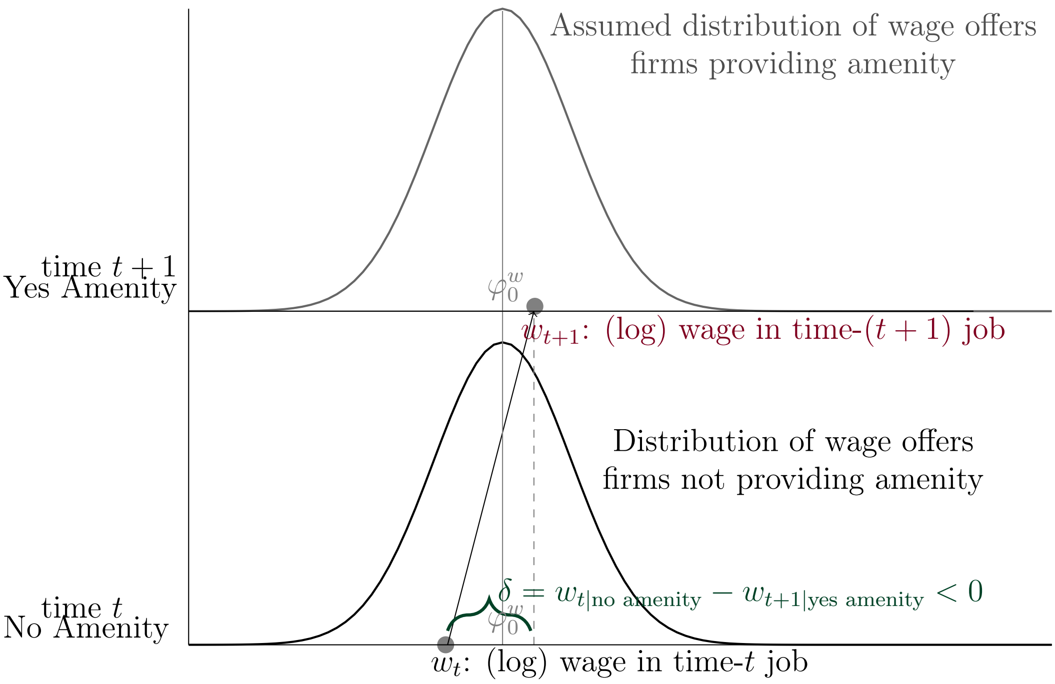
$$\log L(\theta, \lambda, \delta) = \log L_1(\theta) + \log L_2(\theta, \lambda, \delta) + \log L_3(\theta, \lambda, \delta) \tag{25}$$

In equation (25), $L_1(\theta)$ is the contribution to the likelihood function of the distribution of wages and benefits among workers who exited unemployment between t and $t + 1$. Under the assumption that the labor market is in equilibrium, so that all employers offer wage-benefits bundles whose utility is at least as large as workers' reservation utility, all unemployed workers accept any job offer they receive. Consequently, L_1 does not depend on workers' preferences (δ's), and its maximization allows to identify the features of the wage-benefits offers (job offers) that workers receive. Importantly, the maximization of L_1 identifies φ_0^w and φ_c^w, the gender-specific average wages offered to workers in different careers, and ρ, the parameter-vector measuring the gender-specific wage gains or losses that workers obtain when working for employers who provide benefits.

Due to the identification of θ through the wage-benefits outcomes of workers who accept a job offer when unemployed, the wage-benefits outcomes of workers who undergo a job-to-job transition (given θ) identify workers' preferences parameter-vector δ, and the frequency of different labor market transitions (given δ and θ) identifies the vector of search friction parameters λ.

Estimating whether gender differences exist in θ and λ is crucial to properly estimate gender differences in workers' willingness to pay for amenities, δ. To see this, consider the following argument, where I assume for simplicity that job offers consist of a bundle of hourly pay (w) and one valuable amenity (say, paid parental leave), that search frictions are such that $k = 1$, and that there is only one gender.

Figure A36: The estimation of δ without accounting for features of the job offer distribution



Notes: Hypothetical estimated value of δ under the assumption that the distribution of wages offered to workers is identical between amenity-providing firms and firms that do not provide the amenity.

Suppose that one estimates δ by comparing wage outcomes of workers who move from an employer who does not offer paid parental leave to an employer who offers this benefit. Suppose that, as shown in Figure A36, most workers' wages increase when experiencing this type of job-to-job transition. This implies that $\delta = w_{t|\text{no amenity}} - w_{t+1|\text{yes amenity}} < 0$, suggesting that workers require a higher wage when being offered parental leave.

When θ is not estimated, assuming employers who either offer or do not offer a certain benefit are identical in the average wage they offer is necessary to identify workers' preferences through the average wage change of employees undergoing job-to-job transitions.

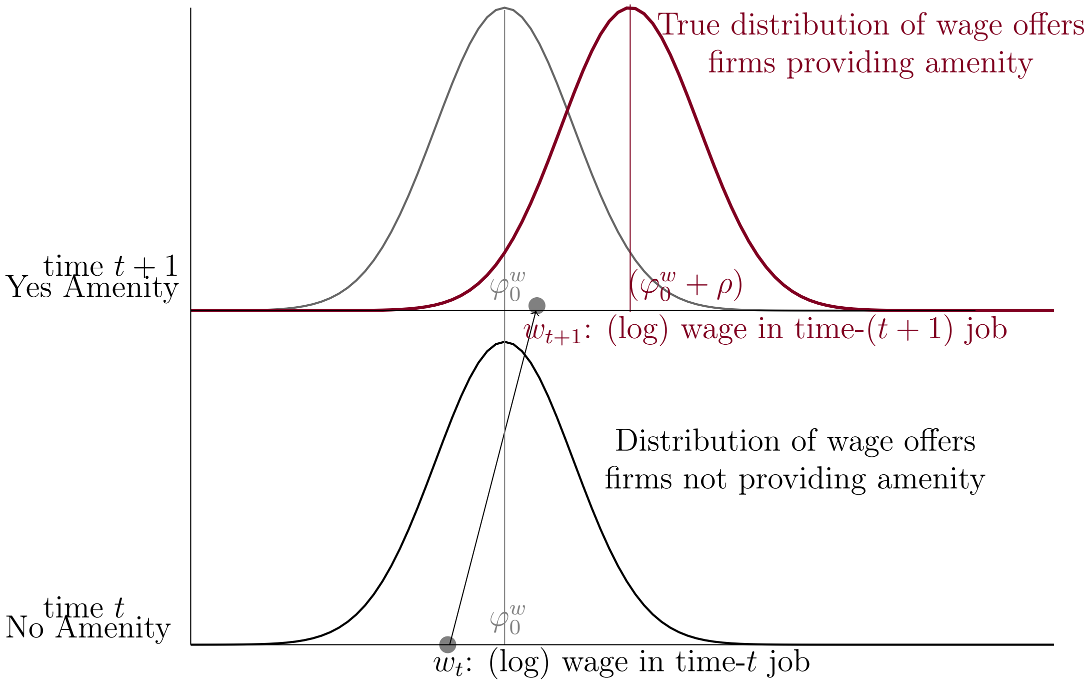
This statistic, however, is a biased estimator of δ, and the resulting estimated preferences may have counterintiutive values and signs, if firms offering valuable benefits pay higher wages. As illustrated in Figure A37, panel (a), if the true θ is such that firms providing a benefit such as paid parental leave are more productive than firms that do not offer it, the higher wage that workers get in $(t+1)$ upon moving into leave-providing firms is at least partly explained by firms heterogeneity rather than by workers' preferences.

Estimating θ, firm heterogeneity is properly taken into account when estimating preferences. Once θ is estimated, preferences for a certain benefit are identified by comparing workers' time-t rank in the distribution of wage offers among firms that do not provide the benefit with their rank in the time-$(t+1)$ distribution of wage offers among firms that do provide it, conditional on workers undergoing a job-to-job transition involving a change in the provision of the benefit of interest. Figure A37, panel (b) shows that, if most workers accept a shift-back in the conditional distribution of wages upon being offered of a benefit, the estimated δ will be positive, reflecting workers' willingness to pay for it.

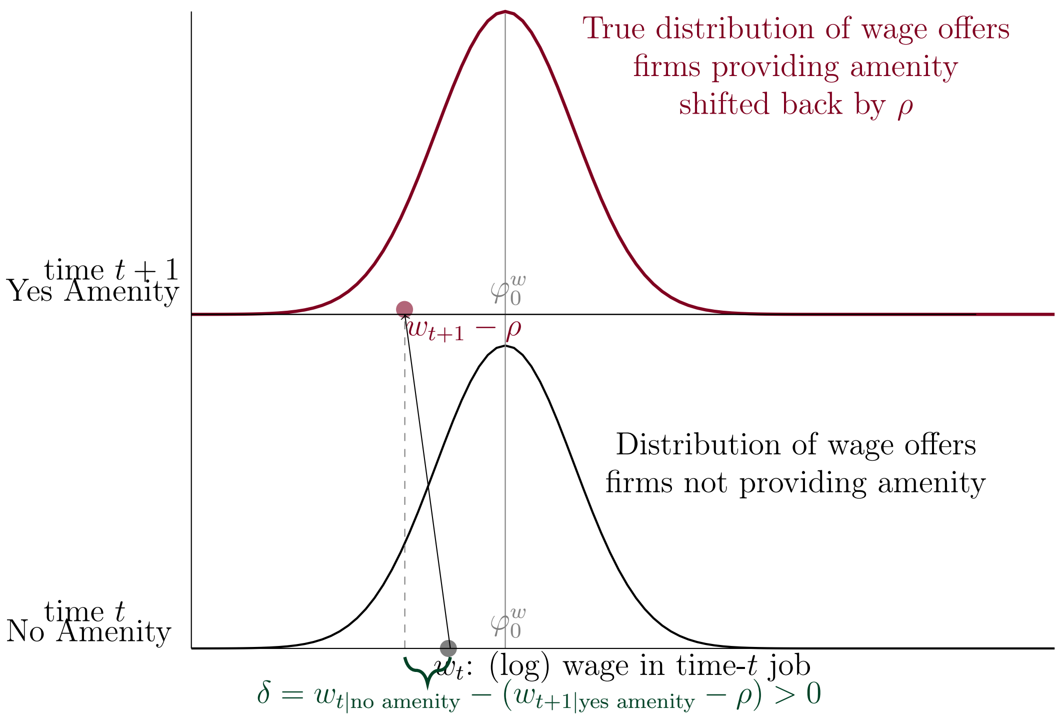
This argument shows that not estimating θ may cause the estimated gender differences in workers' willingness to pay for benefits to be biased if the difference in wages offered by employers who provide benefits compared to employers who do not is heterogeneous across genders. This is likely to occur when benefits are costly for firms,[12] and costs differ, or are perceived to differ, by workers' gender. Offering paid parental leave to women, for example, may be thought to be more costly than offering the same benefit to men, if women are expected to be more likely to use the benefit. This may foster statistical discrimination towards women in firms that provide paid leave (Olivetti & Petrongolo, 2017). If so, even if such firms are more productive than others, the wage premium for working in firms providing paid leave, ρ, will be higher for men than for women.

[12] Ginja, Karimi, and Xiao (2023) study the impact of a parental leave extension reform in Sweden and find that the most exposed firms faced costs comparable to up to 10 full-time equivalent months of wages.

Figure A37: The estimation of δ accounting for features of the job offer distribution



(a) The distribution of wage offers among firms that do or do not provide amenities



(b) δ-estimate accounting for θ

Notes. Panel (a), the time-t graph: distribution of wages offered by firms that do not provide the amenity. Time$(t+1)$ graph: hypothetical true distribution of wages offered by amenity-providing firms, assuming that the latter are more productive than firms that do not provide the amenity $\rho > 0$. Panel (b): identification of the δ parameter after correcting for eventual productivity-differences between amenity-providing firms and firms that do not provide the amenity.

G.5 Estimation results

Search friction parameters Table A10, reports the estimated search friction parameters. Gender differences exist in the rate at which workers undergo different labor market transitions, and young men climb the job-ladder faster than young women. In any given month, male employed workers have a 1% chance of receiving a utility-improving job offer (λ_1), while employed women receive utility-improving job offers at a monthly rate of 0.6%. This implies that, for young employed men, the annual probability of receiving at least one utility-improving job offer is 11.4%, while the probability equals 7% for women.[13] Conversely, constrained job-to-job transitions are slightly more likely among women than among men. Comparing the estimated λ_2 parameters, the annual probability of undergoing at least one job-to-job transition entailing an utility loss is 3.5% among women and 2.4% among men.[14] Young women are also more likely than men to undergo an employment-to-unemployment transition. The estimated q parameters imply that the annual probability of exiting employment at least once is 18.6% for women and 14.5% for men. The estimated λ_0, the arrival rate of job offers among unemployed workers, is the only parameter whose estimated value does not strongly differ across genders. However, this is partly due to the fact that unemployment statuses include one period of job search at labor market entry whose duration is assumed to be identical for men and women. Including the labor market entry search period into the estimation also causes the estimated λ_0 to be large for all workers.

[13]The annual arrival rate of at least one utility-improving job offer is computed as $P(\text{at least 1 utility-improving offer per year}) = 1 - P(\text{no utility-improving offer in one year}) = 1 - (1 - \lambda_1)^{12}$.

[14]This result can have several interpretations. First, some of the married or cohabiting women in my sample may undergo job-to-job transitions due to household migration and incur wage and utility losses as a consequence. Recent evidence shows that, under these circumstances, earnings losses occur for "trailing spouses", those who move following the primary earner, and who may end up either unemployed or in lower-paying jobs compared to their pre-migration labor market outcomes (Burke & Miller, 2017; Venator, 2023). Second, some women may switch job to decrease their commuting time (Le Barbanchon, Rathelot, & Roulet, 2021), and take low-pay jobs providing no amenities in order to work closer to home. In this event, the rate at which women undergo constrained job-to-job transitions also captrure the impact of willingness to pay in exchange for a decrease in commuting time.

Table A10: Estimated search friction parameters

	λ_0	λ_1	λ_2	q
(a) Women	.227	.006	.003	.017
	(.015)	(.001)	(.000)	(.001)
	[.000]	[.000]	[.000]	[.000]
(a) Men	.235	.010	.002	.013
	(.021)	(.002)	(.000)	(.001)
	[.000]	[.000]	[.000]	[.000]

Notes: NLSY97. Sequential maximum likelihood estimates of search-friction parameters defined in text. Block-bootstrapped standard errors in parentheses. P-values of the likelihood ratio tests for search friction parameters equal to 0 are in brackets. The number of observations used in the estimation are 2788 for women (379 women observed over 72 months since labor market entry) and 19152 men (266 men observed over 72 months since labor market entry).

Job offer distribution parameters Table A11 reports the estimated features of the distributions of wages offered to male and female workers. The parameters denoted φ^w indicate career-specific wage premia with respect to the average wage in the sales and clerical career (μ_0^w). μ_2^2 indicates the change in average wage offers in California, New Jersey and Rhode Island after the implementation of paid family leave laws. μ_1^w is the estimated ability wage premium. The last four columns of the table report the average difference in the log-value of received wage offers between workers employed in establishments where, respectively, schedule flexibility, long hours, paid parental leave and unpaid parental leave, are available, and other establishments.

The estimated parameters suggest that women are offered lower wages compared to men in most careers, and the gap in wage offers is especially large among professional workers. The μ_2^w also indicates that wage offers tend to be higher in establishments in states with implemented paid parental leave laws (California, New Jersey, Rhode Island) than elsewhere.

The estimated values of the ρ-parameters show that firms offering paid parental leave and unpaid parental leave offer to their employees wages at least as high as firms where these benefits are not available. Although female workers appear to be offered lower wages when entering firms that provide paid parental leave ($\rho_{pl} < 0$) the parameter is not statistically significant at conventional levels.

Since paid and unpaid parental leave are the most valuable amenities from workers' perspective, and their provision can be costly for employers, the absence of significant wage losses associated with firms providing such benefits suggest that both male and female workers are able to progressively select themselves into more productive firms offering higher wages and better working conditions (Hwang, Mortensen, & Reed, 1998). This interpretation is consistent with recent findings by Sockin (2022), who shows that American higher-pay firms also provide better amenities, thus improving workers' job satisfaction, and by Goldin, Kerr, and Olivetti (2020), who find that American firms offering parental leave are larger and tend to disproportionately employ workers who make pre-childbirth investments in firm-specific human capital.

Regarding work arrangements, both employers requiring employees to work more than 40 hours per week, and employers that allows employees to work on a flexible schedule, offer a pay premium. Interestingly, the long-hours pay premium is larger fore male workers, while the flexibility pay premium is smaller for them (and not statistically different from zero). This result is consistent with the theory proposed by Goldin (2014), according to which part of the gender wage gap is driven by men's selection into convex-pay jobs entailing strong wage premia for working long hours, and by women's selection into linear-pay jobs, that prize long work hours less.

The results in Table A11 clarify the importance of estimating the distribution of wages offered to men and women when studying gender differences in preferences for benefits and work arrangements. Not accounting for demand-side differences between firms offering amenities and firms not offering them would have led to estimate negative δ for unpaid parental leave among men and women alike. Both men and women, in fact, do experience wage increases, on average, upon entering an unpaid-leave-providing firm. Moreover, the wage premium associated with firms providing unpaid parental leave is twice as large for women than for men. Assuming that firms providing unpaid leave would offer identical average wages to men and women would have led to the misleading conclusion that men's preferences for this benefit are almost twice as large as women's preferences. Not estimating the parameters of the wage offer distribution would have also let to overestimate gender differences in willingness to pay for paid parental leave.

Table A11: Estimated parameters of the job offer distribution

	μ_0^w	μ_2^w	φ_{ex}^w	φ_{pr}^w	φ_{oth}^w	μ_1^w	ρ_{fs}	ρ_{lh}	ρ_{pl}	ρ_{ul}
(a) Women	1.723	.347	.163	.091	.024	.197	.215	.207	-.126	.135
	(.274)	(.200)	(.173)	(.146)	(.240)	(.079)	(.057)	(.108)	(.087)	(.067)
(b) Men	1.927	.145	.055	.125	.046	.069	.073	.486	.071	.061
	(.113)	(.134)	(.083)	(.078)	(.098)	(.033)	(.041)	(.099)	(.069)	(.049)

Notes: NLSY97. Sequential maximum likelihood estimates of wage offer parameters. Block-bootstrapped standard errors are in parentheses. The number of observations used in the estimation are 2788 for women (379 women observed over 72 months since labor market entry) and 19152 men (266 men observed over 72 months since labor market entry).

The following tables show the parameters that describe the inverse cumulative distribution function of each benefit and work arrangement. In other words, the parameters characterize the offer distribution of each benefit.

Table A12: Estimated parameters of the offer distribution of schedule flexibility

	μ_0^{fs}	μ_2^{fs}	φ_{ex}^{fs}	φ_{pr}^{fs}	$\varphi_{\text{oth}}^{fs}$	μ_1^{fs}
Women	-.045	-.198	.220	.510	-.118	-.032
	(.267)	(.653)	(.223)	(.218)	(.378)	(.071)
Men	-.200	.041	.190	.060	.012	.101
	(.195)	(.223)	(.132)	(.131)	(.152)	(.042)

Notes: NLSY97. Sequential maximum likelihood estimates of the parameters of the offer distribution of schedule flexibility. Block-bootstrapped standard errors are in parentheses. The number of observations used in the estimation are 2788 for women (379 women observed over 72 months since labor market entry) and 19152 men (266 men observed over 72 months since labor market entry).

Table A13: Estimated parameters of the offer distribution of long hours

	μ_0^{lh}	μ_2^{lh}	φ_{ex}^{lh}	φ_{pr}^{lh}	$\varphi_{\text{oth}}^{lh}$	μ_1^{lh}
Women	-.567	-.296	.064	.042	-.039	-.084
	(.263)	(.596)	(.288)	(.214)	(.493)	(.069)
Men	.061	.026	.433	.431	.023	.127
	(.248)	(.543)	(.216)	(.258)	(.147)	(.053)

Notes: NLSY97. Sequential maximum likelihood estimates of the parameters of the offer distribution of long hours. Block-bootstrapped standard errors are in parentheses. The number of observations used in the estimation are 2788 for women (379 women observed over 72 months since labor market entry) and 19152 men (266 men observed over 72 months since labor market entry).

Table A14: Estimated parameters of the offer distribution of paid parental leave

	μ_0^{pl}	μ_2^{pl}	φ_{ex}^{pl}	φ_{pr}^{pl}	$\varphi_{\text{oth}}^{pl}$	μ_1^{pl}
Women	.624	4.656	-.030	.295	.055	-.043
	(.381)	(.360)	(.208)	(.229)	(.372)	(.094)
Men	.048	4.71	.168	.325	-.030	.036
	(.200)	(.263)	(.154)	(.194)	(.175)	(.052)

Notes: NLSY97. Sequential maximum likelihood estimates of the parameters of the offer distribution of paid parental leave. Block-bootstrapped standard errors are in parentheses. The number of observations used in the estimation are 2788 for women (379 women observed over 72 months since labor market entry) and 19152 men (266 men observed over 72 months since labor market entry).

Table A15: Estimated parameters of the offer distribution of unpaid parental leave

	μ_0^{ul}	μ_2^{ul}	φ_{ex}^{ul}	φ_{pr}^{ul}	$\varphi_{\text{oth}}^{ul}$	μ_1^{ul}
Women	-.080	.335	-.310	.001	2.152	.127
	(.199)	(.450)	(.226)	(.233)	(1.016)	(.052)
Men	-.048	.081	.091	.200	.193	.055
	(.141)	(.191)	(.181)	(.201)	(200)	(.047)

Notes: NLSY97. Sequential maximum likelihood estimates of the parameters of the offer distribution of unpaid parental leave. Block-bootstrapped standard errors are in parentheses. The number of observations used in the estimation are 2788 for women (379 women observed over 72 months since labor market entry) and 19152 men (266 men observed over 72 months since labor market entry).

G.6 Robustness checks

The following table reports the willingness-to-pay parameters estimated in different models. The table shows that, changing the set of benefits included in the hedonic search model does not qualitatively change the estimated gender differences in workers' willingness to pay for paid and unpaid parental leave. The likelihood ratio test p-values reported in brackets also show that the model including paid and unpaid parental leave describes the data better than a model that excluded these two benefits would. Instead, it cannot be rejected that preferences of young male and female college graduates for other benefits and work arrangements, are statistically equal to zero.

Table A16: Willingness to pay for amenities - different models - workers with college degree

	Schedule flexibility	Long hours	Paid parental leave	Unpaid parental leave	Paid vacation
			(a) Women		
1. No long hours, yes paid vacation	1.002		1.797	1.482	1.272
	[1.000]		[.080]	[.072]	[.205]
1. No flexibility, yes paid vacation		-.642	2.000	1.287	.958
		[1.000]	[.129]	[.121]	[1.000]
			(b) Men		
1. No long hours, yes paid vacation	.586		.790	1.002	.888
	[1.000]		[.038]	[.005]	[.271]
1. No flexibility, yes paid vacation		.260	.702	.738	.425
		[1.000]	[.851]	[.054]	[1.000]

Notes: NLSY97, sample as in Table 1. Sequential maximum likelihood estimates of preference parameters. P-values for likelihood ratio tests (null hypothesis: $\delta^{a_k} = 0$) are reported in brackets. The table reports the preference parameters estimated in job search models that do not account for workers' preferences for long hours (line 1) and schedule flexibility (line 2) while account for workers' preferences for paid vacation. The choice to include paid vacation is motivated by the evidence in Figure A7, showing that the provision of paid vacation is associated with a decline in the probability that workers change employer between two consecutive years, and that the coefficients are statistically significant for men and only marginally not significant for women. The number of observations used in the estimation are 2788 for women (379 women observed over 72 months since labor market entry) and 19152 men (266 men observed over 72 months since labor market entry).

In the following table, lines 1 to 3 show the estimated preference parameters in the hedonic job search model estimated on the sample of all workers who enter the labor market between 2000 and 2011 (and satisfy all other sample-cleaning criteria) with any level of education. Line 1 replicates the baseline model estimated in the paper for the full sample of workers. Lines 2 and 3 progressively include additional benefits: paid vacation and retirement plan. The inclusion of paid vacation is motivated by the evidence in Figure A7. The figure shows that, among college graduate workers, the availability of paid vacation at workers' current workplace is associated with a statistically significant decline in the probability that male workers change employer by the following year, and with a marginally not significant decline in the probability that female workers change employer by the consecutive year. The inclusion of retirement plan is motivated by evidence in Figure A17. It shows that working for an employer that provides a retirement plan is associated with a significant decline in the probability that female workers without a college degree change employer by the consecutive year. This evidence suggests that workers may value the availability of paid vacation and of a retirement plan when making job-change decisions, and that these two benefits should be included in the model.

The choice to estimate the model with additional benefits on the larger sample of workers with any level of education is motivated by the need to increase the number of observation used to estimate the increasingly large number of parameters that such models have. The college graduate sample that I study throughout the paper is relatively small and the model is estimated separately by genders. In order to estimate preferences for benefits, it is necessary to observe a sufficient number of workers who, conditional on all other benefits included in the model, move at least once between an employer that does not offer a certain benefit and an employer who does. The variation in the provision of benefits is not sufficient in the college graduate sample to estimate models including too many benefits. It is also worth noting that the hedonic search model including four benefits (flexibility, long hours, paid parental leave and unpaid parental leave) in fact estimates 44 parameters: four preference parameters δ's, four search friction parameters λ's, and 36 parameters describing the distribution of wage offers and of benefit offers that workers receive. Every other benefit included in the model implies an addition of 8 more parameters to be estimated in the model: one parameter ρ^{a_k} describing the difference in offered wages between employers who offer amenity a_k and employers who do not; one parameter δ^{a_k} describing workers' willingness to pay for a_k, and six parameters describing the distribution of a_k-offers, that is, the probability that a_k is offered to workers of different ability, in different time-constant occupation classes (careers), in states with or without implemented paid parental leave policies. It implies that the five-benefit hedonic search model estimates 52 parameters, while the six-benefit hedonic search model estimates 60 parameters.

The results in lines 1 to 3 in Table A17 are, however, reassuring. First, estimating the baseline model on the full sample of workers of all levels of education does not qualitatively change the estimated gender difference in workers' willingness to pay for paid and unpaid parental leave. Second, gender differences in workers' willingness to pay for paid and unpaid parental leave remain when additional benefits are added to the model.

It is important to notice that the estimated preferences for valuable benefits appear stronger the more benefits are added to the model. This result is intuitive. As previously mentioned, preferences are estimated by comparing the wage outcomes of workers who, by undergoing a job-to-job transition, move from an employer that does not offer the benefit to an employer who does, conditional on the provision of all other benefits included in the model. Thus, in line-3 model, preferences for a certain benefit, say, schedule flexibility, are estimated on the subsample of workers who move across employers offering the same benefits but schedule flexibility. It is plausible that workers with especially strong preferences for benefits may decide to undergo this type of job-to-job transition.

Lines 4 to 6 in panel (a) (women) and 4 to 5 in panel (b) (men), estimate additional versions of the baseline model on the main sample of interest in the paper: workers who enter the labor market after college graduation. Line-4 estimated preferences result from a model estimated using raw measures of paid and unpaid parental leave. That is, the model underlying Line-4 results does not address potential gender differences in measurement error in workers' reporting of paid and unpaid parental leave availability at their current employer. Importantly, gender differences in workers' willingness to pay for paid parental leave are not affected. However, workers' preferences for unpaid parental leave become stronger when measurement error is not addressed, and especially so for men. This implies that male workers who report that unpaid leave is available to them at their current employers may be more likely to be aware of the availability of this benefit due to their especially strong preference for this benefit. Thus, addressing possible measurement error by using the details of the Family and Medical Leave Act of 1993 to impute whether workers are employed in unpaid-leave-providing establishments reduces the upward bias in men's (average) preferences for unpaid parental leave.

In Line 5, I estimate the baseline model without allowing the distribution of wages and amenities offered by employers by vary depending on whether states have paid parental leave laws implemented or not. In the time-frame that I study, these states are: California from 2004, New Jersey from 2009, and Rhode Island from 2014. This change in the model does not qualitatively affects gender differences in preferences for benefits.

Finally, in Line 6 in panel (a) I report the main results of the baseline model estimated on the subsample of college graduate women who do not have children throughout their early careers. Preferences for paid and unpaid leave are strong in this subgroup of women, and stronger than they are for men. This evidence is highly suggestive that, since labor market entry, a non-negligible share of women may account for future family-formation and fertility decisions when searching for jobs, thus selecting into firms that provide benefits that will be valuable to them in the future and accepting lower wages in exchange for the provision of these benefits.

Table A17: Willingness to pay for amenities - different models - all workers and workers with college degree

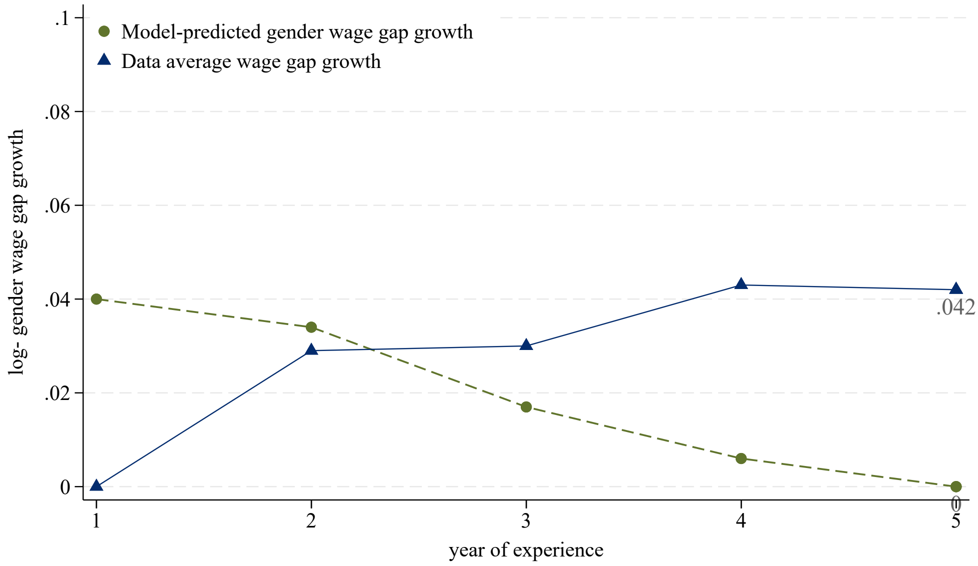
	Schedule flexi-bility	Long hours	Paid parental leave	Unpaid parental leave	Paid vaca-tion	Retire-ment plan
			(a) Women			
1. All: Baseline	.787	-.324	1.209	1.009		
2. All: Baseline, 5 benefits	.713	-.363	1.314	1 057	1.475	
3. All: Baseline, 6 benefits	1.338	-1.117	1.553	1.813	2.000	1.400
4. Col: No meas. error corr.	1.235	.288	1.564	1.534		
5. Col: No heterogeneity	.865	.206	1.480	1.208		
6. Col: No children	.934	.343	1.375	.975		
			(b) Men			
1. All: Baseline	.934	.993	.767	.683		
2. All: Baseline, 5 benefits	.939	-.146	.823	.669	1.134	
3. All: Baseline, 6 benefits	1.320	-.724	1.328	1.068	1.937	.715
4. Col: No meas. error corr.	.478	.423	.509	2.000		
5. Col: No heterogeneity	.554	.586	.696	.936		

Notes: NLSY97, sample in lines 1, 2, and 3 as in Table 1. The sample in lines 4, 5 and 6 includes workers of all levels of education. Preference parameters estimated through sequential maximum likelihood. The table reports the preference parameters estimated in different models. In lines 1 to 3, the model is estimated on workers who enter the labor market between 2000 and 2011 with any level of education. Each line corresponds to a model with one additional amenity included. The total number of parameters estimated in line-1 model (including the θ and λ parameters) is 44. The total number of parameters estimated in line-2 model is 52. The total number of parameters estimated in line-3 model is 60. Line 4 reports the estimates of the baseline model estimated on college graduate workers using raw variables capturing the availability of paid and unpaid parental leave rather than imputed (or measurement-error adjusted) variables. Line 5 reports the result of the model estimated without allowing for heterogeneity in the distribution job offers (θ parameter vector) between states with implemented paid parental leave policies (California from 2004, New Jersey from 2009, and Rhode Island from 2014) and other states. The number of observations used in lines 1 to 3 is 92952 for women (1291 women observed for 72 months since labor market entry) and 84456 for men (1173 men observed for 72 months since labor market entry). The number of observations used in the estimation in lines 4 and 5 are 2788 for women (379 women observed over 72 months since labor market entry) and 19152 men (266 men observed over 72 months since labor market entry). Line 6 in panel (a) restricts the sample of college graduate women to the 69% of women who do not have children throughout their early career.

G.7 Counterfactual analyses

Fit of models with different sets of benefits The following figure shows the model fit of the observed average early-career growth in the gender wage gap when the model is estimated excluding long work hours and including paid vacation as an alternative benefit. Ignoring long hours causes the model to mistakenly predict that the gender wage gap declines during early career. This model, in fact, does not account for the strong wage premia that men obtain (ρ_{lh}) upon entering firms where workers are required to work long hours.

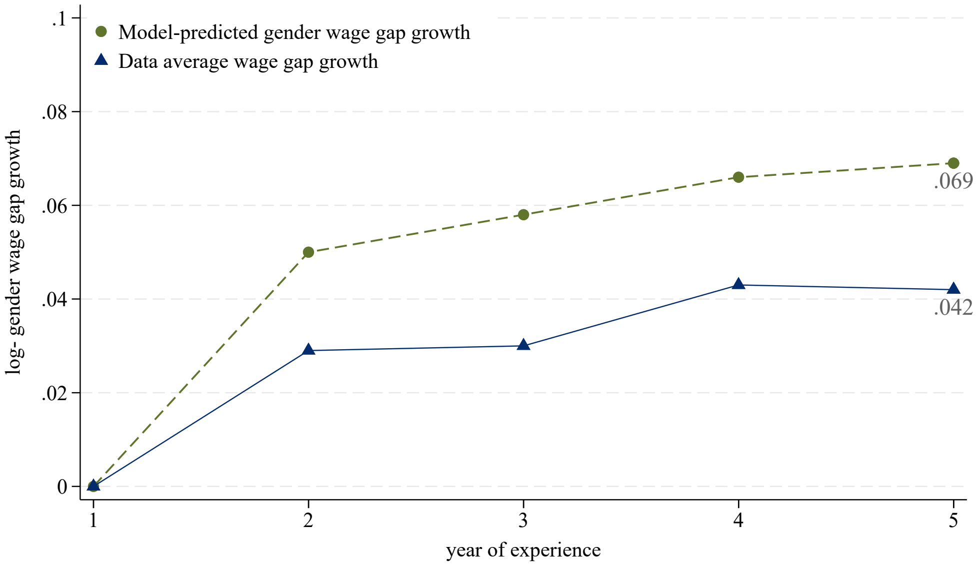
Figure A38: Predicted gender wage gap growth: excludes long hours, includes paid vacation



Notes: NLSY97. The green dotted line depicts the early-career growth in the gender wage gap predicted by a search model in which long hours are not included in the set of relevant benefits and work arrangements, while paid vacation is included.

The following figure shows the model fit of the observed average early-career growth in the gender wage gap when the model is estimated excluding schedule flexibility and including paid vacation as an alternative benefit. This model over-predicts the early-career growth in the gender wage gap by 2.7 log-points. It does not account for the fact that, while gender differences in preferences for flexibility exist but are not large, entering firms that provide flexible work schedules entails stronger wage gains among women than among men (ρ_{fs}).

Figure A39: Predicted gender wage gap growth: excludes schedule flexibility, includes paid vacation



Notes: NLSY97. The green dotted line depicts the early-career growth in the gender wage gap predicted by a search model in which schedule flexibility is not included in the set of relevant benefits and work arrangements, while paid vacation is included.

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