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

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Abstract

I show that highly educated millennial Americans search for employers that provide parental leave, and that women's stronger willingness to pay for this benefit contributes to the early-career growth in the gender wage gap. Using a hedonic job search model, I estimate that workers are offered higher wages when hired by employers providing paid and unpaid parental leave, but women are willing to pay, respectively, 40% more and 56% more than men for these benefits. While all workers search for jobs and experience wage growth by entering firms offering both high pay and valuable benefits, the gender wage gap increases as young women accept lower wages, compared to men, upon receiving job offers from employers who provide parental leave. While the early-career growth in the gender wage gap would decline by 75% if willingness to pay for parental leave did not differ across genders, a policy mandating and subsidizing parental leave provision could itself halve the early-career wage-gap growth. The widespread availability of parental leave would lessen workers' need to accept lower wages in exchange for its provision, reducing the gap in accepted wages between men and women entering leave-providing firms.

JEL Codes: J16, J31, J32, J64
Keywords: Gender wage gap, non-wage benefits, paid parental leave, unpaid parental leave, job search, early careers.

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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, the Federal Employee Paid Leave Act (FEPLA) of 2019 provides paid leave to eligible federal workers, and nine 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 2022, only 25% of civilian American workers had access to paid parental leave (BLS, 2022), while the FMLA requirements exclude from unpaid leave coverage at least 40% of the US workforce (S. Brown, Herr, Roy, & Klerman, 2020).[1]

Anecdotal evidence suggests that US employers may be filling gaps in parental leave coverage by providing this benefit in an effort to attract and retain employees (Cain Miller, 2018a; Michelson, 2021). Yet, the availability of employer-sponsored leave is limited to a small number of large employers (Cain Miller, 2018b), it can change with macroeconomic conditions and is currently appearing to decline (Dill & Yang, 2022).

In this paper I ask how the provision of employer-sponsored paid and unpaid parental leave, and lack thereof, affect the gender wage gap among college graduate, millennial American workers, and its rapid increase during workers' early careers.[2] In a scenario where the availability of parental leave is scarce and largely at employers' discretion,

[1]Concerning paid leave, as of 2023, 14 US states have passed paid parental leave laws: California, Colorado, Connecticut, Delaware, District of Columbia, Illinois, Maine, Massachusetts, Maryland, New Hampshire, New Jersey, New York, Oregon, Rhode Island, Vermont, Washington. The implementation is still pending in: Colorado, Illinois, Maine, Maryland, Oregon, and in Vermont for private sector employees. 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 month 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. In most states, paid leave is funded through a mixture of payroll deductions and employer contributions. Two exemptions are D.C., where paid leave is 100% employer-funded and New Hampshire, where the Paid Family and Medical Leave plan is a 100% payroll-deduction-funded voluntary insurance plan. The modal annual duration of paid family leave is up to 12 weeks. Regarding unpaid leave, the FMLA exempts firms with less than 50 employees in a 75-mile radius from the requirement of providing the benefit. Eligibility is limited to employees having worked for an employer for at least one year and for 1250 hours minimum in the previous year. 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).

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

young workers may search for employers offering this benefit and be willing to accept lower wages in exchange for its provision. If parental leave is more salient for young women than for young men, then women's 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 answer this research question, I use 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 triples by the sixth year of labor market experience, that at least 50% of such increase can be explained by the larger wage gains that male workers obtain when changing employer (job), that the likelihood of being offered valuable benefits such as paid and unpaid leave increases upon changing employer, and that the availability of parental leave significantly reduces the chances of undergoing a job change for women but not for men. These results hold prior to, and potentially irrespective of, marriage and childbirth.[3]

Second, I estimate men's and women's preferences 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 the provision of parental leave using a hedonic job search model.

The model builds on the Bonhomme and Jolivet (2009) seminal contribution. 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.[4] 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'

[3]While Angelov, Johansson, and Lindahl (2016) and Kleven, Landais, and Søgaard (2019) find that gender gaps in labor market outcomes expand after childbirth, the evidence I provide suggests that some of the roots of the divergence in labor market outcomes between men and women predate family formation decisions and are not a direct consequence of women's labor supply choices following childbirth.

[4]The monthly rate of arrival of job offers reflects search frictions. The more search frictions, the lower the arrival rate.

wage offers depend on whether benefits are provided, and on workers' skill level and occupation. All the parameters of the model are gender-specific.

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 approach estimates stronger willingness to pay for a benefit the lower is the average wage accepted by workers entering firms that provide it, compared to their previous wage.

The revealed-preferences approach used to identify preferences using 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.[5]

Second, it reduces concerns that unobserved gender differences in wage offers or in search frictions may affect the estimates of preferences. If women are offered lower wages or face stronger search frictions compared to men, and these factors are unaccounted for, women's willingness to pay for benefits may be misleadingly overestimated. 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). Gender differences in wage offers may increase in firms offering parental leave, if employers ex-

[5] 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. C. 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), Sullivan and To (2014), Sorkin (2018), Xiao (2021). Khandker (1988) was the first to introduce non-wage attributes in a search model.

pect women to take up leave more often and for longer periods (Olivetti & Petrongolo, 2017). Furthermore, 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, in press) which might trigger monopsonistic wage discrimination (Manning, 2003) and result in lower wage offers. If search behavior differs across genders (Bowlus, 1997; Cortes, 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.

Finally, the model allows to separately identify workers' preferences for different benefits and job characteristics. In this paper, 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 choice avoids that estimated gender differences in preferences for parental leave are biased due to women's possibly stronger preferences for employers offering part-time work (Bowlus & Grogan, 2009; Liu, 2016) and flexible work arrangements (Mas & Pallais, 2017; Xiao, 2021) who may be more likely to offer other family-friendly benefits, or due to men's selection in high-pay, long-hours jobs (Goldin, 2014) where benefits such as parental leave may not be provided.

The main estimation results show that both young women and men highly value the provision of paid and unpaid parental leave, but women's preferences for these benefits are, respectively, 40% and 56% stronger than men's. Even though workers, and most prominently women, are willing to pay for the provision of parental leave, I estimate that firms offering leave typically pay higher wages to both men and women. Furthermore, while I find that women are offered lower wages compared to men, this gap does not increase when employers offer paid or unpaid leave.[6]

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 and more valuable benefits (Hwang, Mortensen, & Reed, 1998; Sockin, 2022). Due to their stronger willingness to pay for paid and unpaid leave, however, young women accept lower wages

[6]Additional results show that preferences for flexibility and long hours are remarkably similar across genders at labor market entry, while frictions are slightly stronger for employed women than for employed men, but are similar across genders among unemployed workers.

compared to men, upon receiving job offers from employers who provide such benefits.

Coherently with this interpretation, counterfactual analyses show that the early-career growth in the gender wage gap would decline by 75% if willingness to pay for parental leave did not differ across genders. In this event, in fact, women's wages would grow as fast as men's wages as workers climb the job ladder to enter firms offering parental leave.

Further counterfactual exercises show that a policy mandating the provision of parental leave would mute the effect of preferences for this benefit on accepted wages, halving the early-career growth in the gender wage gap. The widespread availability of parental leave would lessen workers' need to accept lower wages in exchange for its provision, thus reducing the gap in accepted wages between men and women entering leave-providing firms.

This paper proposes a novel angle to understand parental leave. The vast literature on the topic highlighted the positive impacts of paid parental leave policies on parents' and children's health (Bartel, Rossin-Slater, Ruhm, Slopen, & Waldfogel, 2023b), and comprehensive reviews of the literature documented the positive effects on post-childbirth women's job continuity of policies granting relatively short leave periods (Olivetti & Petrongolo, 2017; Rossin-Slater, 2018). Some authors also found no evidence of adverse effects of leave-taking on wages and future labor market outcomes (Bana, Bedard, & Rossin-Slater, 2020), while it is still debated 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.

If the potential effects of parental leave policies are numerous, debated, 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 is also 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 paid and unpaid parental leave is left to employers and costs are not subsidized, only some firms will provide them. Consequently, workers for whom these benefits are more salient will disproportionately pay for them by accepting lower wages in exchange for their 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.

Studying the impact of workers' willingness to pay for parental leave on the gender wage gap, this paper also contributes to the growing literature analyzing the impact of preferences for non-wage job attributes (Flabbi & Moro, 2012; Hotz, Johansson, & Karimi, 2018; Liu, 2016; Mas & Pallais, 2017; Xiao, 2021), of location and commuting (Caldwell & Danieli, in press; Le Barbanchon, Rathelot, & Roulet, 2021), and of firm heterogeneity (Barth, Olivetti, & Kerr, 2021; Card, Cardoso, Heining, & Kline, 2018; Card, Cardoso, & Kline, 2016) on wages and on gender inequality in labor market outcomes.

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 wages 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.

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 explains 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, that the availability of valuable benefits affects workers' job-change decisions, and that workers' valuation of benefits such as paid and unpaid parental leave may differ by gender.

2.1 Features of the NLSY97 and sample selection

I use 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 for each year 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 and stock ownership, and of flexible work arrangements.

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.

The sample I study consists of workers who graduate from college by age 25, 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 35 hours per week on average (Blau & Kahn, 2017). I exclude individuals who are ever self-employed, who ever report hourly wages above $200 (in 2005 US dollars) or work hours above 112 per week, and who are ever employed in agriculture or in the military. I also exclude individuals for whom information about relevant job-specific information is ever missing. The final sample includes 319 male workers and 455 female workers, each observed weekly for the first six years in the labor market.[7]

[7]Female workers represent 56.4% of the final sample due to the focus on college graduate workers, the majority of whom are women among recent cohorts of Americans (Goldin, Katz, & Kuziemko, 2006). In the raw sample of NLSY97 individuals who obtain a bachelor degree by Round 17, 42% are males and

2.2 Sample characteristics and stylized facts

Table 1 reports the demographic characteristics of workers in the final sample.

Table 1: Time-invariant sample characteristics

	Men	Women	Diff.	Obs.
Age at labor market entry	23.68	23.83	-0.15	774
Bachelor degree by labor market entry	0.61	0.68	-0.07**	774
Master degree by age 26	0.07	0.10	-0.03	774
Prospective PhD graduate	0.02	0.02	0.00	774
African American	0.13	0.17	-0.05*	774
White	0.76	0.71	0.05*	774
Marries/cohabits by labor market entry	0.24	0.33	-0.09***	774
Marries/cohabits by 3rd yr in labor market	0.54	0.63	-0.09***	774
Marries/cohabits by 6th yr in labor market	0.71	0.74	-0.04	774
Marries by NLSY Round 17	0.75	0.75	-0.00	774
Has child by labor market entry	0.05	0.08	-0.03*	774
Has child by 3rd yr in labor market	0.15	0.18	-0.02	774
Has child by 6th yr in labor market	0.26	0.30	-0.04	774
Has child by NLSY Round 17	0.53	0.56	-0.03	774
Age at first childbirth	28.00	27.32	0.68*	418

Notes: National Longitudinal Survey of Youth 1997 (NLSY97), Rounds 1 to 15. The sample includes workers who graduate from college by age 25, with non-missing observations on demographic characteristics and on employer-employee-specific wages and job characteristics throughout the first six years of labor market experience (early careers). The number of observations in the table refers to the number of female (455) and male (319) workers in the final sample observed during their first week in employment at labor market entry. All workers are subsequently observed for six years.

Young men and women are approximately 24 years old at labor market entry and, while female workers are more likely to have completed their bachelor degree by that time, no significant gender differences emerge in the likelihood of pursuing and completing post-graduate education. Women do appear to anticipate family-formation decisions compared to men. While similar shares of men and women marry (75%), or have a child (above 50%), by the 2015 round of the NLSY97, a higher proportion of women are married or cohabit (33%) or have a child (8%) at labor market entry, compared to men. Among workers who have a child by 2015, women have their first child almost one year earlier than men, around three years since labor market entry.

58% 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 the selected workers are representative of the sex-composition of Millennial college graduates as a whole.

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. More than two-thirds of men and women in the final sample change at least one employer (job) by five years since labor market entry, the first job-change occurring around the third year of labor market experience for both young men and young women.

Table 2: Time-varying sample characteristics

	Men	Women	Diff.	Obs.
	(a) Labor market history			
Total n. of years employed	5.99	5.97	0.02	774
Tot n. spells out of work	2.03	2.40	-0.37*	774
Tot n. weeks out of work	14.45	17.83	-3.37*	774
Total n. of jobs held	2.39	2.58	-0.19	774
Changes employer by 6th year in labor market	0.67	0.69	-0.02	774
Year of experience first job change	2.98	3.08	-0.11	526
	(b) Outcomes - first yr of experience			
Average weekly hours worked	40.28	39.34	0.94	774
Weekly hours > 40	0.20	0.17	0.03	774
Total n. of weeks employed in t	46.87	48.06	-1.19**	774
Hourly rate of pay (in 2005 Dollars)	15.24	14.05	1.19**	774
Hourly pay - Executive/Managerial	14.74	14.83	-0.08	170
Hourly pay - Professional/Health Tech	17.64	16.95	0.69	195
Hourly pay - Social/Educ/Admin/Health Supp	13.76	12.98	0.78	270
Hourly pay - Other occupation	14.25	11.41	2.84***	139
Employer n. of employees	686.08	602.17	83.91	701
Employer with <50 employees	0.42	0.46	-0.04	701
	(c) Outcomes - sixth yr of experience			
Average weekly hours worked	44.09	40.75	3.35***	774
Weekly hours > 40	0.43	0.28	0.15***	774
Total n. of weeks employed in t	50.23	48.41	1.82***	774
Hourly rate of pay (in 2005 Dollars)	23.55	20.44	3.11***	774
Hourly pay - Executive/Managerial	23.68	22.33	1.35	170
Hourly pay - Professional/Health Tech	28.01	25.43	2.58	195
Hourly pay - Social/Educ/Admin/Health Supp	20.09	17.63	2.45	270
Hourly pay - Other occupation	21.34	18.24	3.10	139
Employer n. of employees	791.63	969.84	-178.21	705
Employer with <50 employees	0.42	0.35	0.07*	705

Notes: NLSY97, sample selection as in table 1. Panels (b) and (c) refer to workers observed during their first week in employment, respectively, at labor market entry and five years later. In both panels, workers' hourly rate of pay and work hours are employer-employee specific.

As shown in panels (b) and (c) of table 2, workers enter larger firms as they change job, suggesting that workers' aim to climb the job ladder by entering higher-pay firm should be a main determinant of job changes.[8] Nevertheless, more than one-third of both young men and young women in the final sample work for employers with less than 50 employees during their early careers.[9]

Concerning labor market attachment, work hours and job continuity, panel (a) of table 2 shows that women spend approximately 18 weeks out of work, overall, during the their early careers, while men's employment gaps duration sums up to 14 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 five years later women work approximately three hours less than men per week and almost two weeks 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, but are rather determined by men's faster rise in weeks and hours worked.

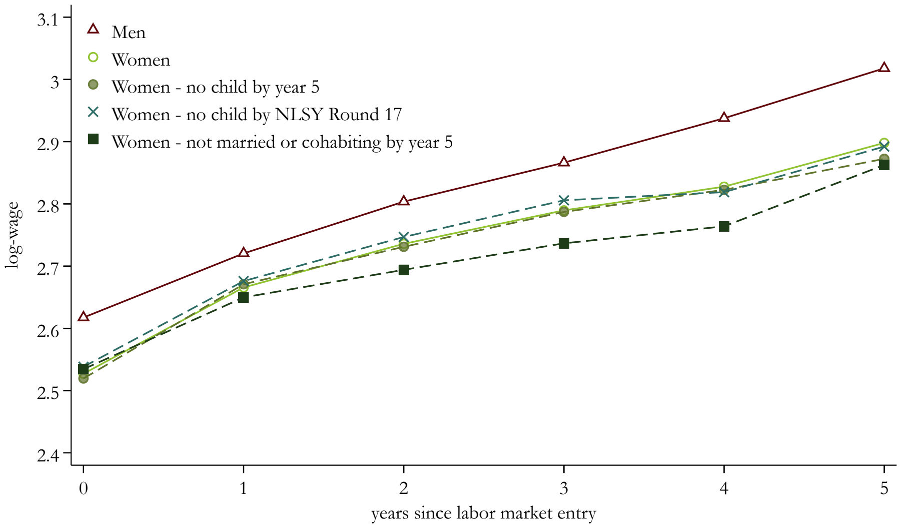
In line with the literature, panels (b) and (c) of table 2 also show that the gender pay gap expands over time in the labor market (Amano-Patiño, Baron, & Xiao, 2020; Barth, Olivetti, & Kerr, 2021; Loprest, 1992; Manning & Swaffield, 2009) due to the faster wage growth experienced by young men. The gap grows from roughly one dollar per hour at labor market entry to $3.11 five years later. The pay gap also expands with experience within occupation classes.

It is worth noting that the gender wage gap rises with experience irrespective of women's fertility status and family-formation decisions. Figure 1 compares the average experience (log) wage profiles of men to the experience wage profiles of all women, and of women who do not have children and do not marry during their early careers. The figure clearly depicts that the early-career expansion in the pay gap occurs prior to, and potentially irrespective of, childbirth and marriage. It implies that, while childbirth events can generate penalties that contribute to expand labor market gaps between men and women (Angelov, Johansson, & Lindahl, 2016; Kleven, Landais, & Søgaard, 2019), the roots of those gaps and of their growth over workers' careers exist since labor market entry.

[8]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).

[9]Small firms are exempt from compliance with several US labor market acts and regulations. For example, under the FMLA of 1993, firms with less than 50 employees are exempt from the requirement of offering up to 12 weeks of unpaid family leave to their employees.

Figure 1: Experience wage profiles



Notes: NLSY97, sample selection as in table 1. 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, women who do not have children by 2015 (Round 17 of the NLSY97), and women who are neither married nor cohabit by five years since labor market entry.

While the early-career increase in the gender wage gap does not appear to be purely an outcome of women's family formation decisions, it can be partly explained by the rising gender gap in labor market attachment and hours worked.

Table 3: Gender gaps in hours and weeks worked by year of experience

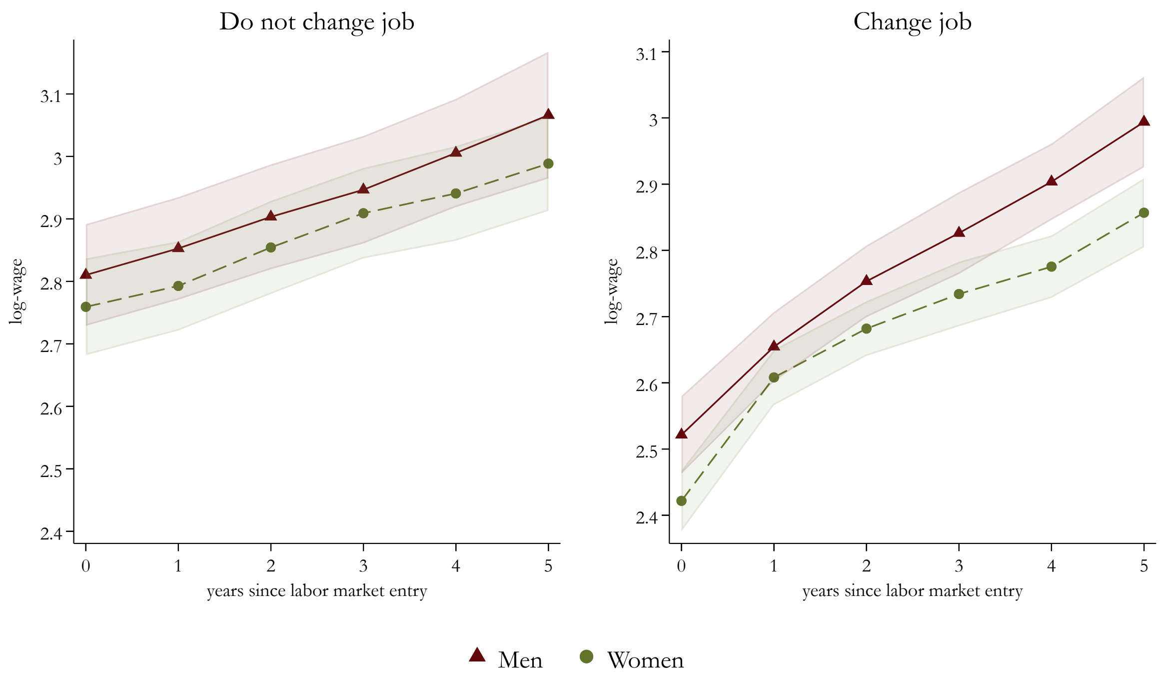
	Women no kids	Women not married	Do not change job	Change job
	(a) First year of experience			
Average weekly hours worked	1.05	-0.43	0.66	0.96
Weekly hours > 40	0.03	0.03	0.08	0.01
Total n. of weeks employed in t	-1.15**	-0.85**	-0.21**	-1.73**
	(b) Sixth year of experience			
Average weekly hours worked	2.82***	2.06***	1.57***	4.16***
Weekly hours > 40	0.14***	0.10***	0.20***	0.13***
Total n. of weeks employed in t	1.55**	-0.08**	1.76**	1.79**

Notes: NLSY97, sample selection as in table 1. 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.

As table 3 shows, even restricting the female sample to women who do not have children or do not marry during their early careers, gender gaps in work hours and workweeks do emerge over time in the labor market. As women's work hours and workweeks grow slowlier than men's with experience, men may enjoy stronger wage premia for long hours and work continuity. Such premia impact wages, predominantly among career-oriented workers in certain managerial and professional occupations (Bertrand, Goldin, & Katz, 2010; Cortes & Pan, 2019; Gicheva, 2013; Goldin, 2014) where college graduates represent the vast majority of the employed workforce. In light of this evidence, I will account for gender differences in work hours throughout this paper, defining long work hours as weekly work hours above 40.

Besides work hours and job continuity, job changes are a major determinant of the early-career expansion in the gender wage gap.

Figure 2: Experience wage profiles by job-change status



Notes: NLSY97, sample selection as in table 1. The left-hand panel shows the average log-wage profiles of workers who do not change employer during their early careers. The right-hand panel depicts the log-wage profiles of workers who do change at least one employer during the same time period. The areas depict 95% confidence intervals of the gender-specific mean (log) wages.

As shown in figure 2, the increase in the gender wage gap over years of experience is driven by workers who change at least one employer during their early careers. The gender pay gap among employees who work for the same employer throughout their entire early careers is small, not statistically significant, and roughly constant over time.[10]

2.3 Job changes, wage growth and the gender wage gap

To quantify the contribution of job changes to the growth in the early-career pay gap, in this section I estimate gender-specific fixed-effect wage regressions where returns to labor market experience vary between workers who change employer in a given year and workers who do not.

When workers voluntarily change job to enter a more profitable employment relationship, they should receive a wage increase compared to workers who, conditional on their previous-year experience and job characteristics, remain employed at the same firm. In other words, job changes lead to steep experience-wage profiles due to workers' accumulation of search capital (Burdett & Mortensen, 1998; Topel & Ward, 1992). Even if workers' job changes are driven by workers' desire 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 offering valuable benefits and work arrangements are more productive than firms who do not (Hwang, Mortensen, & Reed, 1998).

Between male and female workers, wage gains due to job changes may anyway differ. If women are less likely than men to receive lucrative job offers (stronger search frictions); if they are subject to some form of wage discrimination, face stronger mobility constraints or limit their search to lower-pay jobs (are offered lower wages); or if they are willing to pay more in exchange for the provision of valuable benefits (stronger preferences for benefits), job changes will lead to faster wage growth for men than for women.

The regressions that I estimate take the following form

$$
\begin{aligned}
w_{i,j,t} = \alpha &+ \beta_1 \mathrm{exp}_{i,t-1} + \beta_2 \mathrm{exp}^2_{i,t-1} + \gamma \mathrm{exp}_{i,t-1} \times \mathbb{1}(\text{voluntary job change [VJC]})_{i,t-1} + \\
&+ \delta \mathrm{exp}_{i,t-1} \times \mathbb{1}(\text{lost previous job [LPJ]})_{i,t-1} + \\
&+ \eta \mathrm{exp}_{i,t-1} \times \mathbb{1}(\text{other job change [OJC]})_{i,t-1} + x'_{i,j,t-1}\psi + \varepsilon_i + u_{i,j,t} \qquad (1)
\end{aligned}
$$

$w_{i,j,t}$ is the log-wage that employee i receives at firm (employer) j in the first week in

[10] Among job changers, the pay gap increases over time in the labor market also between men and women who do not have kids and do not marry during their early careers, as shown in figure A1 in the Online Appendix.

employment in year t. $\exp_{i,t-1}$ is worker i's actual aggregate experience up to $(t-1)$. Following Light and Ureta (1995), I calculate it using the annualized sum of weeks that i spent in employment between labor market entry and year $(t-1)$. Hence, the variable does not capture periods out of work, reducing concerns that gender differences in experience capture underlying differences in job continuity and human capital accumulation[11]

The experience variable is interacted with three dummy variables capturing, respectively, whether between $(t-1)$ and t employee i changed employer because they were willing to accept a job offer (voluntary job change), because they lost their previous job, their previous firm closed or there was a layoff (lost job), or due to other reasons (mobility constraints, medical reasons, family issues, others). $x_{i,j,t-1}$ is a vector of controls for worker i's and their employer j's characteristics in $(t-1)$, including the worker's occupation and industry, the set of benefits offered by their employer, work hours, and the total amount of time spent out of work until $(t-1)$.

The parameter of interest is γ, measuring the difference in returns to actual labor market experience between workers who change job between $(t-1)$ and t, and workers with equal $(t-1)$ experience who did not change employer.

Table 4 shows the results of the estimation for men (M) and women (F). Columns (1) to (3) progressively add control variables, while in column (4) the sample is restricted to workers who change at least one job during their early careers. This choice reduces concerns that the estimated γ solely reflects job changers' selection on unobservables which might differ between men and women.

The estimation results show that both young men and women obtain noticeable gains from actual labor market experience, although such gains are significantly higher for men than for women. Voluntary job transitions due to workers' willingness to accept a job offer are associated with significant increases in returns to experience which are approximately twice as large among men than among women.[12]

These results can be quantified noting that the gender wage gap increases by four log-points by six years since labor market entry. The most conservative estimates in table 4

[11] Actual aggregate experience is calculated as $\exp_{i,t-1} = (\sum_{k=1}^{K} \text{weeks worked in year } k)/52$ where $k=1$ is the year of labor market entry, and K is $(t-1)$.

[12] In line with the literature on the wage effects of job losses (Couch & Placzek, 1993; Jacobson, LaLonde, & Sullivan, 1993), constrained job transitions due to previous employment loss entail a decline in workers' returns to experience which are larger and significant for men.

show that the gender gap in returns to experience increases by two log-points following a job change. This suggests that gender differences in wage gains from job changes are at least as large as 50% of the early-career increase in the gender wage gap.

Table 4: Wage gains from job changes

	(1)		(2)		(3)		(4)	
	(M)	(F)	(M)	(F)	(M)	(F)	(M)	(F)
exp(t-1)	0.07***	0.07***	0.10***	0.05***	0.10***	0.05***	0.11***	0.04*
	(0.02)	(0.01)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
exp $\times \mathbb{1}$(LPJ)	-0.04**	-0.02	-0.03*	-0.02	-0.03*	-0.02	-0.03	-0.01
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
exp $\times \mathbb{1}$(VJC)	0.05**	0.03**	0.06***	0.03***	0.06***	0.03***	0.06***	0.04***
	(0.02)	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)
exp $\times \mathbb{1}$(OJC)	0.02	-0.01	0.02	-0.01	0.02	-0.01	0.02	-0.01
	(0.02)	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)
Adj. R^2	0.18	0.13	0.19	0.14	0.20	0.14	0.22	0.14
N	1587	2253	1587	2253	1587	2253	1057	1543
$x'_{i,j,(t-1)}$	N	N	Y	Y	Y	Y	Y	Y
$\text{Occ}_{i,j,(t-1)}$	N	N	N	N	Y	Y	Y	Y
$\text{Ind}_{i,j,(t-1)}$	N	N	N	N	Y	Y	Y	Y
All change j	N	N	N	N	N	N	Y	Y

Notes: NLSY97, sample selection as in table 1. The estimation uses annual observations for all workers in the final sample. Each observation represents the first week that a worker is employed in any given year. The job-specific information, including wages and work hours, refers to the working conditions and outcomes of each worker at the employer where they work in that specific week. The panel is unbalanced due to some workers being observed in employment for less than five years following labor market entry. Columns (1) to (3) include all workers, column (4) restricts the sample to employees who change at least one employer during the first six years in the labor market.Standard errors in parentheses are clustered at the individual level.

Table 5 verifies the robustness of previous results across different samples of women, to rule out that gender differences in wage gains from job changes are driven by within-household joint search dynamics or by changes in women's labor supply behavior following childbirth. Joint-search dynamics may affect married workers' choices and constraints (Guler, Guvenen, & Violante, 2012), the types of the job offers that workers receive (Flabbi & Mabli, 2018), and wage gains from job changes (Burke & Miller, 2017; Venator, 2023), while changes in women's labor supply behavior following childbirth may determine wage penalties (Angelov, Johansson, & Lindahl, 2016; Kleven, Landais, & Søgaard, 2019) and, possibly, a decline in returns from job changes.

Column (1) in table 5 reports the γ estimates for the whole sample of women. Columns (2) and (3) limit the sample, respectively, to women who do not have children and to women who do not marry during their early careers. Columns (4) and (5) restrict the sample to women with no children and to unmarried women who change their first job by three years since labor market entry. All coefficients are remarkably stable across women's samples.

Table 5: Wage gains from job changes - Women by family composition

	(1)	(2)	(3)	(4)	(5)
exp(t-1)	0.05***	0.06**	0.01	0.05*	0.00
	(0.02)	(0.02)	(0.03)	(0.03)	(0.03)
exp $\times \mathbb{1}$(LPJ)	-0.02	-0.02	-0.05	-0.02	-0.03
	(0.02)	(0.03)	(0.04)	(0.03)	(0.04)
exp $\times \mathbb{1}$(VJC)	0.03***	0.03**	0.03	0.04***	0.04*
	(0.01)	(0.01)	(0.02)	(0.01)	(0.02)
exp $\times \mathbb{1}$(OJC)	-0.01	-0.02	-0.01	-0.02	-0.00
	(0.01)	(0.01)	(0.02)	(0.01)	(0.02)
Adj. R^2	0.14	0.13	0.15	0.13	0.15
N	2253	1568	965	1108	620
$x'_{i,j,(t-1)}$	Y	Y	Y	Y	Y
$\text{Occ}_{i,j,(t-1)}$	Y	Y	Y	Y	Y
$\text{Ind}_{i,j,(t-1)}$	Y	Y	Y	Y	Y
All change j	N	N	N	N	N

Notes: NLSY97, sample selection as in tables 1 and 4. Samples of women differ by column as follows. (1): all women. (2): women with no children by the sixth year in the labor market. (3): women who are not married and do not cohabit by the sixth year on the labor market. (4): column (2) women who change their first job by the third year in the labor market. (5): column (3) women changing their first job by the third year in the labor market. Control variables as in table 4 column (3).

The similarities in individual characteristics, career paths and job-change outcomes between the average unmarried woman, the average woman without children, and the average woman in the final sample rule out that gender differences in early-career labor market outcomes and in wage gains from job changes are determined by married women's and mothers' search behavior, labor supply decisions, preferences and constraints. Throughout her early career, the average millennial woman appears to behave as an unmarried woman without children.

2.4 Benefits, work arrangements and job changes

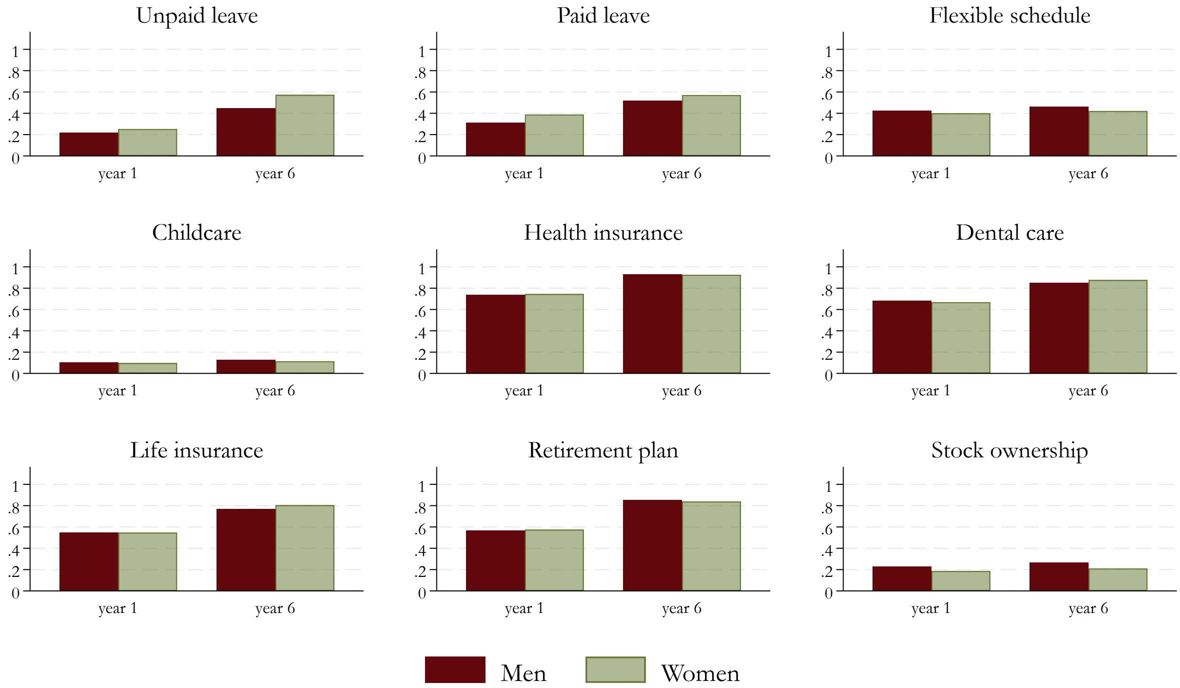
Although the increase in the gender wage gap and gender differences in wage gains from job changes predate, and possibly exist irrespective of, marriage and childbirth, the possibility of these life-events can nevertheless affect men and women differently by impacting their preferences for non-wage benefits and work arrangements. For example, the availability of paid or unpaid parental leave may affect young women's job search, as such benefits may represent a form of employment insurance for those who both anticipate the possibility of building a family in the future and are strongly attached to the labor market. While young men may also prefer jobs providing parental leave, the availability of such benefits may be less salient to them. It is also possible, although not obvious, that similar reasons may underlie eventual gender differences in the salience of work arrangements such as schedule flexibility and long work hours.

In this section I provide some reduced-form evidence suggesting that the availability of valuable benefits and work arrangements affects workers' job changes, and that paid and unpaid leave may be more relevant for young women than for young men.

Figure 3 shows that the share of employees working for employers who provide valuable benefits and work arrangements considerably rises in years of experience. This evidence suggests that job changes may be driven not only by workers' aim for higher wages, but also by their willingness to enter firms providing benefits that match their needs and preferences (Akerlof, Rose, & Yellen, 1988; Bonhomme & Jolivet, 2009; Card, Cardoso, Heining, & Kline, 2018; Flabbi & Moro, 2012; Gronberg & Reed, 1994; Hotz, Johansson, & Karimi, 2018; Hwang, Mortensen, & Reed, 1998; Liu, 2016; Sorkin, 2018; Sullivan & To, 2014; Xiao, 2021).

Furthermore, while men and women are equally likely to be offered several benefits throughout their early careers, gender differences exist in the availability of paid and unpaid leave. As a matter of fact, 22% of men and 25% of women are offered unpaid leave at labor market entry, while around 45% of men and 57% of women work for employers offering unpaid leave by the sixth year of labor market experience. Similarly, roughly 31% of men and 39% of women work for an employer offering paid parental leave at labor market entry, while 52% of men and 57% of women are offered this benefit five years later. This evidence hints that men's and women's valuation of paid and unpaid parental leave may differ.

Figure 3: Shares of employees working for amenity-providing employers



Notes: NLSY97, sample selection as in table 1. Each panel depicts the share of men (burgundy) and women (green) who report to work for an employer offering a certain benefit or work arrangement during the first week in employment, respectively, at labor market entry (year 1) and five years later (year 6).

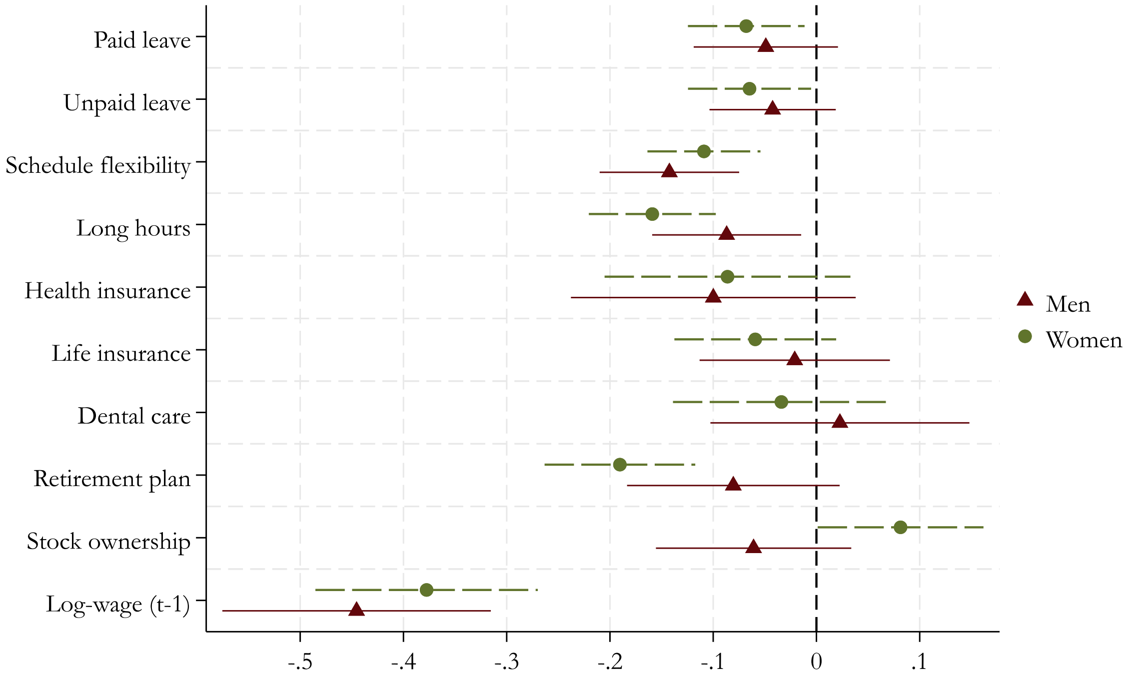
Supporting this hypothesis, figure 4 reports selected coefficient estimates of gender-specific fixed-effect linear probability models of job changes. They estimate the association between workers' probability of changing employer between two consecutive years and the characteristics of their previous job. The estimated regressions are

$$\mathbb{1}(J_{i,t} \neq J_{i,t-1})_{i,t} = \alpha + \beta_1 \exp_{i,t-1} + \beta_2 \exp^2_{i,t-1} + \\ + \sum_{k=1}^{K} \delta_k \mathbb{1}(\text{Benefit } k \text{ provided})_{i,j,t-1} + x'_{i,j,t-1}\delta + \varepsilon_i + u_{i,j,t} \qquad (2)$$

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 provided benefit k. $\exp_{i,t-1}$ is worker i's aggregate labor market experience up to $(t-1)$. $x'_{i,j,t-1}$ is a vector of $(t-1)$ characteristics of i and of their employer, j, ε_i is a worker fixed-effect and $u_{i,j,t}$ is an error term.

Results show that employees working in firms providing valuable 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, 7 percentage-points and 6 percentage-points in the probability of changing employer. The coefficients are negative and slightly smaller in magnitude for men, but not statistically significant at 10% significance level. Schedule flexibility and long work hours are associated with a significant reduction in the chances of changing employer for both men and women.[13]

Figure 4: Linear probability model of job changes - Selected coefficient estimates



Notes: NLSY97, sample selection as in tables 1, and 4 column (4). 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 $(t-1)$. Child care is not included in the estimation for insufficient within-individual over time variation in the provision of the benefit. Standard errors are clustered at the individual level.

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 affect the early-career path in the gender pay gap.

During workers early careers, most workers change at least one job and enter larger

[13]Table A3 in the Online Appendix reports the full set of coefficient estimates.

firms. As workers change job, they experience increases in work hours and in hourly wages that are larger among men, an increase in the likelihood of being offered valuable benefits such as paid or unpaid parental leave, and an increase in the likelihood of having flexible work arrangements.[14]

Consistently with hedonic job search theory (Hwang, Mortensen, & Reed, 1998), the steeper experience wage profiles of workers who change jobs, and the improvement in work arrangements and in benefits offered to workers, suggest that employees who change job 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 dynamic would then contribute to the increase in the gender wage gap in years of experience.

Yet, reduced-form evidence does not allow either to quantify men's and women's preferences (willingness to pay) for benefits and work arrangements, or to 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.

[14]Figure A4 in the Online Appendix shows that workers hired by larger employers are more likely to be offered paid and unpaid parental leave.

3 Hedonic Search Model

3.1 Model Setup

The set-up of the model is as follows. 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.[15] The monthly rates of job offer arrival and of job loss define search frictions.

A job is a 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 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 $F^g(w_{i,j}, \mathbf{a}_{i,j}|b, c)$. To control for within-gender heterogeneity in job offers and in workers' selection into jobs offering different benefits, I let $F(.)$ depend on workers' ability, denoted b, and on their career, c. Since $F(.)$ is taken as given, the model is in partial equilibrium.

When employed, a worker i obtains utility from their wage and from the benefits 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 minimum wage a worker would accept to work for an employer that provides benefit a^k relative to the hourly pay offered by an employer that does not provide a^k. The larger δ_k^g, the lower the wage that worker i accepts in exchange for the provision of a^k.

---

[15] The λ_2^g parameter that Bonhomme and Jolivet (2009) add to the basic 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.

The estimation of the model requires the characterization of the steady-state distribution of wages and amenities among employed workers. As in Bonhomme and Jolivet (2009), the steady state of the model can be found as follows. First, the steady-state probability that a worker leaves their job can be written as

$$P^g(\text{leave}|w_{i,j}, \mathbf{a}_{i,j}, b, c) = q^g + \lambda_2^g + \lambda_1^g \bar{F}_u^g(w_{i,j} + \delta^{g\prime}\mathbf{a}_{i,j}|b, c) \tag{4}$$

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

$$\lambda_0^g U^g = q^g(1 - U^g) \tag{5}$$

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

$$\begin{aligned} \lambda_0 U F_u(u|.) + \lambda_2 F_u(u|.)(1-U)\bar{G}_u(u|.) = q(1-U)G_u(u|.) + \lambda_2 \bar{F}_u(u|.)(1-U)G_u(u|.) + \\ + \lambda_1 \bar{F}_u(u|.)(1-U)G_u(u|.) \end{aligned} \tag{6}$$

Where I dropped the superscript g to simplify notation. Equation (6) further implies that the steady-state cumulative distribution of utility levels among employed workers of gender g and ability b in career c is

$$G_u(u|.) = \frac{F_u(w_{i,j} + \delta'\mathbf{a}_{i,j}|.)}{1 + k\bar{F}_u(w_{i,j} + \delta'\mathbf{a}_{i,j}|.)} \tag{7}$$

Using (7), the density function of utility levels among employed workers is, thus,

$$g_u(u|.) = (1 + k)\frac{f_u(u|.)}{[1 + k\bar{F}_u(w_{i,j} + \delta'\mathbf{a}_{i,j}|.)]^2} \tag{8}$$

Finally, using equation (8), the steady-state cross-sectional distribution of wages and amenities among employed workers is[16]

$$g(w, \mathbf{a}|.) = (1+k)\frac{f(w, \mathbf{a}|.)}{[1 + k\bar{F}_u(w + \delta'\mathbf{a}|.)]^2} \tag{9}$$

Where $k = \frac{\lambda_1}{q+\lambda_2}$ is a measure of gender-specific search frictions. 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.

Equation (9) shows 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 $\bar{F}$).

Equation (9) is especially relevant in the context of this paper, as it shows that potential gender differences in search frictions and in the wage-amenities bundles that employers offer to workers must be properly accounted for, to correctly estimate gender differences in workers' willingness to pay for non-wage benefits and work arrangements.

3.2 Model Estimation

I estimate the model using a 76-month panel dataset following the 774 workers in the final sample 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 is employed, I use the most frequent worker-specific employer identifier appearing in the weekly arrays to define the worker's firm and job. If an worker is out of employment, I assume they are unemployed.[17]

For employed workers, I retain information on the wage, benefits and work arrangements available at their current employer. The benefits and work arrangements of interest

[16] As Bonhomme and Jolivet (2009) show based on previous results by Dey and Flinn (2005), equation (9) can be obtained using (8) 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, $g(w, \mathbf{a}|u)$.

[17] Bowlus (1997) shows that part of the gender pay gap between US college graduate workers in the baby boom generation depended of the low serch intensity of women who temporarily exited the labor force rather than being unemployed. In my sample, however, the number of employment gaps is very small for both men and women, and their duration is too short, for me to be able to separately estimate heterogeneous search frictions depending on the nature of out-of-employment gaps.

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 more than 40 weekly hours (lh).

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

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 an indicator variable of the occupation class recurring most frequently during the worker's first six years of labor market experience. I define the following careers: administrative, executive, professional, other. The administrative career (ad) includes community and social services occupations, primary and secondary education teachers, librarians, administrative assistants and health care support workers. The executive career (ex) includes management, business and financial occupations. The professional career (pr) includes professional specialty occupations, post-secondary teachers and health care technicians. The remaining occupations are classified as "other" (ot).

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

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 allows to 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 allow to observe whether different employees work for the same employer and, consequently, 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-benefits offers received by male and female workers can only be modeled in reduced form. Specifically

$$w^*_{i,j}(b,c) = \varphi^w_0 + \mu^w_1 b_i + \rho' \mathbf{a}^*_{i,j} + \sum_{c \in \{\text{ex, pr, ot}\}} \varphi^w_c \mathrm{c}_i + \sigma_w \varepsilon^w_{i,j} \tag{10}$$

$$a^{k*}_{i,j}(b,c) = \mathbf{1}\{\mu^{a^k}_0 + \mu^{a^k}_1 b_i + \sum_{c \in \{\text{ex, pr, ot}\}} \varphi^{a^k}_c \mathrm{c}_i + \varepsilon^{a^k}_{i,j} > 0\} \quad \text{for} \quad a^k_{i,j} \in \{a^{fs}_{i,j}, a^{lh}_{i,j}, a^{pl}_{i,j}, a^{ul}_{i,j}\} \tag{11}$$

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 administrative career (the base group). The parameters φ^w_c and $\varphi^{a^k}_c$ represent, respectively, the difference between the average wage offered in career c (executive, professional, other) and the average wage offered in administrative careers, and the career-specific changes in the likelihood that amenity a^k is offered compared to the base group. Equation (10) shows that wage offers $w^*_{i,j}(b,c)$ depend on the amenities that employers offer through the $(K \times 1)$ coefficient vector ρ. For each amenity a^k, ρ^k represents the average difference in wages offered between employers providing benefit a^k and employers who do not provide it.

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 all employers offer the same utility in equilibrium. If so, the magnitude and sign of the ρ parameters closely correspond to the magnitude and sign of the δ parameters. 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, employees who work for the most productive employers, who offer valuable benefits, earn higher wages than employees working for employers who do not provide any benefit. Thus, if search frictions exist, the ρ-parameters corresponding to valuable benefits are positive and workers obtain a *de facto* wage premium for working in firms providing valuable benefits, even though they may be willing to accept lower wages to work in those firms.

Consistently with the impact of search frictions on the relation between wages and amenities that Hwang, Mortensen, and Reed (1998) theorized, equation (9) 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 likelihood function describing the distribution of wages and amenities among employed workers and workers' labor market transitions. 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 $t = 0$.[18] Following this initial period of unemployment, in any of the subsequent 76 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, which, as in Bonhomme and Jolivet (2009), is

$$\begin{aligned} l_{t+1} = & \, q^{ju_t}[1-\lambda_0]^{uu_t} \times \\ & \times \lambda_0^{uj_t} f_{t+1}(w_{t+1}, \mathbf{a}_{t+1}|.)^{uj_t} \times \\ & \times [1 - \lambda_1 \bar{F}(u_t|.) - \lambda_2 - q]^{s_t} \times \\ & \times [\lambda_1 \mathbf{1}\{w_{t+1} + \delta' \mathbf{a}_{t+1} > w_t + \delta' \mathbf{a}_t\} + \lambda_2]^{jj_t} f_{t+1}(w_{t+1}, \mathbf{a}_{t+1}|.)^{jj_t} \end{aligned} \tag{12}$$

Where $s_t, jj_t, ju_t, uj_t, uu_t$ are dummy variables indicating, respectively, workers who, between 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 $(t+1) \in \{1, 76\}$ of their early career, is

$$L(.) = \prod_{i=1}^{N^g} \prod_{t=0}^{75} l_{t+1}(e_{t+1}, w_{t+1}, \mathbf{a}_{t+1}, s_t, jj_t, ju_t, uj_t, uu_t | e_t, w_t, \mathbf{a}_t, b, c) \tag{13}$$

Following Bonhomme and Jolivet (2009), the functional forms for $f(w^*, \mathbf{a}^*|.)$ and

[18]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.

$\bar{F}_u(u|.)$ and, consequently, the functional form of the likelihood function (13) can be found by exploiting the assumptions of normality and independence of the random shocks in the wage and benefits offers.[19]

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

$$\underbrace{\Gamma}_{38\times 1} = [\underbrace{\theta}_{30\times 1}, \underbrace{\lambda}_{4\times 1}, \underbrace{\delta}_{4\times 1}]$$
$$\left[\left[\varphi_0^w, \underbrace{\rho'}_{[\rho^{fs}\ \rho^{lh}\ \rho^{pl}\ \rho^{ul}]}, \sigma_w, \mu_1^w, \underbrace{\varphi^{w\prime}}_{[\varphi_{ex}^w\ \varphi_{pr}^w\ \varphi_{ot}^w]}, [\varphi_0^{fs}\ \mu_1^{fs}\ \underbrace{\varphi^{fs\prime}}_{[\varphi_{ex}^{fs}\ \varphi_{pr}^{fs}\ \varphi_{ot}^{fs}]}], ..., [\varphi_0^{ul}\ \mu_1^{ul}\ \underbrace{\varphi^{ul\prime}}_{[\varphi_{ex}^{ul}\ \varphi_{pr}^{ul}\ \varphi_{ot}^{ul}]}]\right],\right.$$
$$[\lambda_0,\ \lambda_1,\ \lambda_2,\ q], \left[\delta^{fs},\ \delta^{lh},\ \delta^{pl},\ \delta^{ul}\right] \tag{14}$$

Where θ is the (30×1) 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).[20] The identification of the parameters of interest comes from the additive separability of the log-likelihood function in the parameters of interest.

$$\log L(\theta, \lambda, \delta) = \log L_1(\theta) + \log L_2(\theta, \lambda, \delta) + \log L_3(\theta, \lambda, \delta) \tag{15}$$

In equation (15), $L_1(\theta)$ 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 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

[19]Section B in the Online Appendix shows how to derive the functional forms for $f(w^*, \mathbf{a}^*|.)$ and $\bar{F}_u(u|.)$.

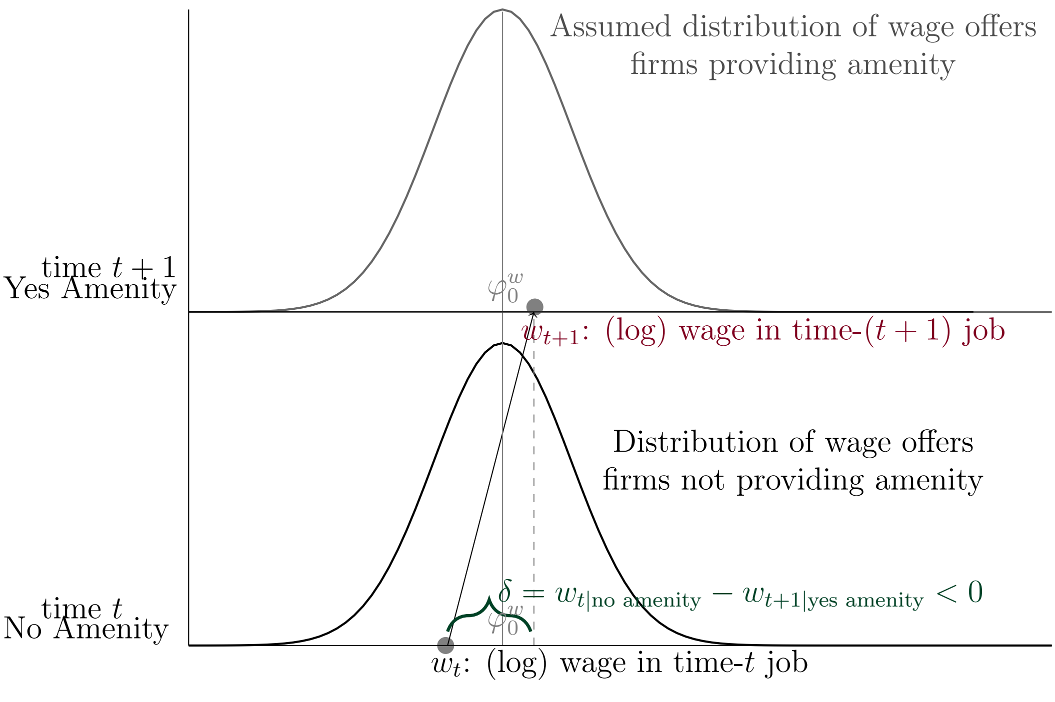
[20]I describe the algorithm in section B in the Online Appendix.

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.

As previously mentioned, 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 5: 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 5, 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.

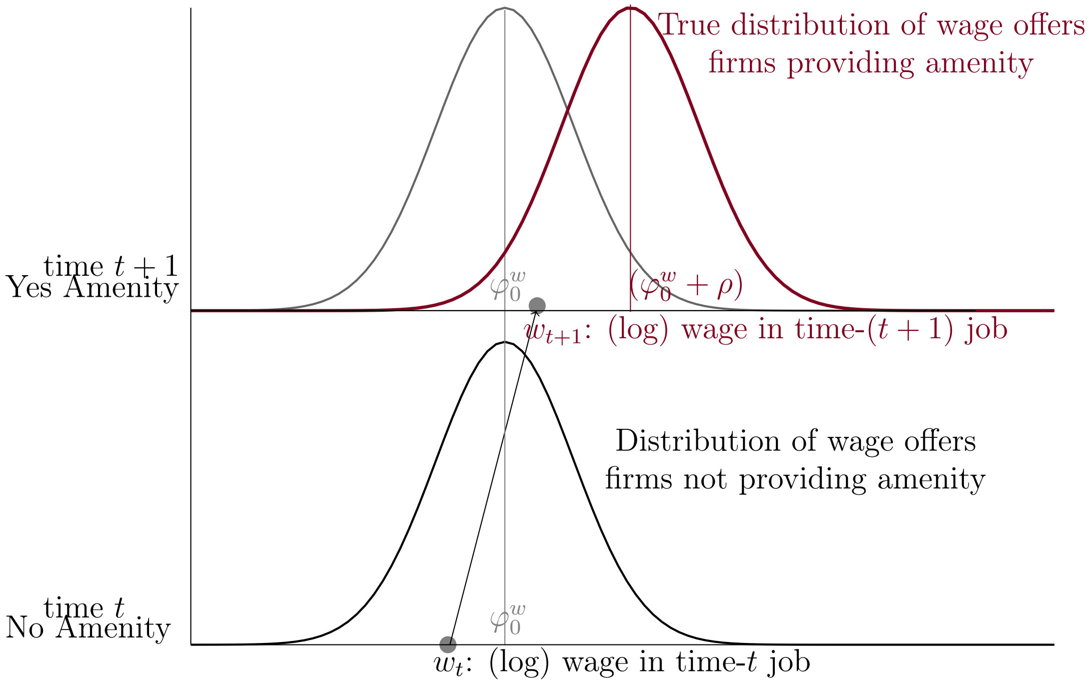
This statistic, however, is a biased estimator of δ, and the resulting estimated preferences may have counterintiutive values and signs, if firms offering valuable benefits pay higher wages. As illustrated in figure 6, panel (a), if the true θ is such that firms providing a benefit such as paid parental 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 allows to take firm heterogeneity 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 6, 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.

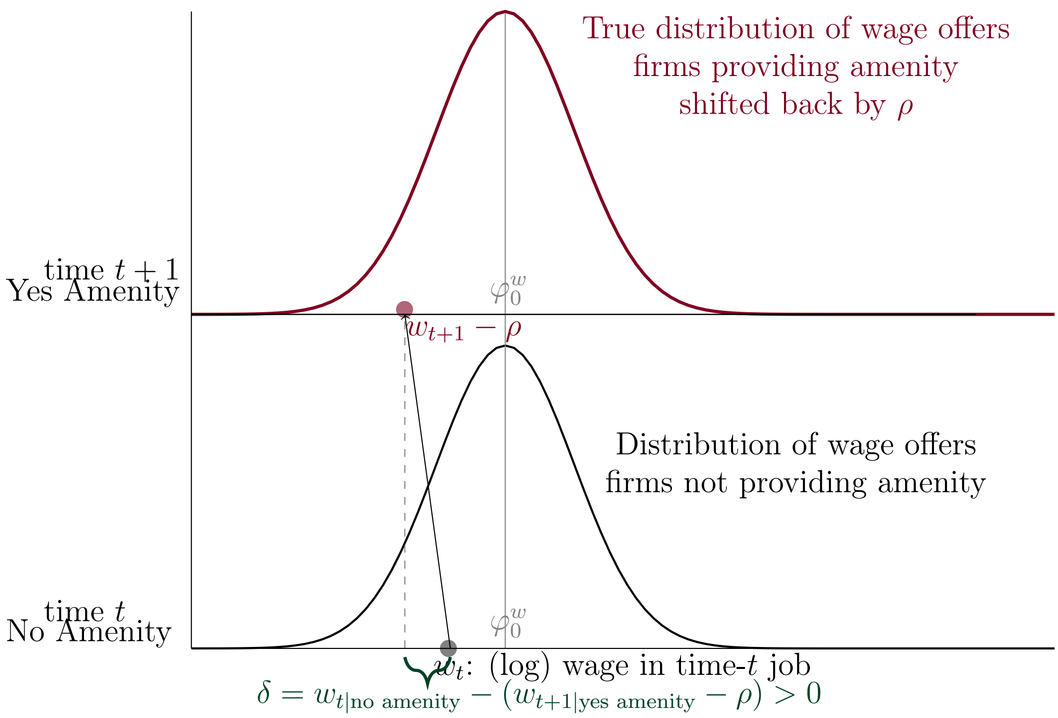
This argument shows that not estimating θ may cause the estimated gender differences in workers' willingness to pay for benefits to be biased if the difference in wages offered by employers who provide benefits compared to employers who do not is heterogeneous across genders. This is likely to occur when benefits are costly for firms,[21] 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.

[21] 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 6: 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.

3.3 Estimation Results

3.3.1 Parameter Estimates

Tables 6, 7 and 8 report the estimation results for the model parameters.[22] Regarding search frictions, table 6 shows that gender differences exist in the rate at which workers undergo different labor market transitions, and that young men climb the job-ladder faster than young women. In any given month, male employed workers have a 2.9% chance of receiving a utility-improving job offer (λ_1), while employed women receive utility-improving job offers at a monthly rate of 2%. This implies that, for young employed men, the annual probability of receiving at least one utility-improving job offer is 29.7%, while the probability equals 21.5% for women.[23] Conversely, constrained job-to-job transitions are 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 10.3% among women and 8.1% among men.[24] 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 15.6% for women and 10.3% 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.

[22] Table A4 in the Online Appendix reports the initial conditions I used in the sequential maximum likelihood estimation.

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

[24] 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 6: Estimated search friction parameters

	λ_0	λ_1	λ_2	q
(a): Women	.382	.020	.009	.014
	(.010)	(.002)	(.001)	(.001)
(b): Men	.380	.029	.007	.009
	(.013)	(.002)	(.001)	(.001)

Notes: NLSY97. Sequential maximum likelihood estimates of search-friction parameters defined in text. Asymptotic standard errors computed through the outer product of gradients method are in parentheses.

Regarding preferences, the estimated coefficients in table 7 panel (a) show that paid parental leave and unpaid parental leave are the most valuable benefits for both men and women, but women are willing to pay, respectively, 40% more and 56% more than men for their provision.

As shown in panel (b), compared to the average acceptable wage offered by employers who do not offer paid leave, young women would accept to be paid 67% less to work for an employer providing this benefit, while the maximum wage cut that young men would accept is 45%. Similarly, women would accept up to a 52% wage decline to work for an employer offering unpaid leave, while young men would not accept more than a 38% wage decline in exchange for the provision of this benefit.

Concerning schedule flexibility and long hours, both male and female workers appear to attach a positive value to these work arrangements, and the estimated preferences are very similar across genders, although the coefficients are noisily estimated and not statistically significant for women.[25] There are several reasons why workers' preferences for these work arrangements may not differ by gender. First, workers may account for the flexibility of their prospective job schedule and for required work hours when choosing their career, implying that gender differences in preferences for these work arrangements may reflect into gender-based occupational segregation, captured by workers' selection

[25] As far as long hours are concerned, the estimated parameters imply that both male and female workers are willing to accept wage cuts in order to work longer hours. While this may seem counterintuitive, working longer hours at labor market entry may unlock promotions and wage-growth possibilities throughout workers' careers (Bertrand, Goldin, & Katz, 2010; Booth, Francesconi, & Frank, 2003; Gicheva, 2013), especially among highly educated workers. If workers take these prospects into account, they may be willing to work long hours and renounce to a higher wage when young to enjoy faster wage-growth prospects throughout their life-cycle.

into different careers, rather than into within-career job-to-job transitions.[26] Second, the workers studied in this paper are at the very beginning of their labor market experience, they are on average 24 years-old at labor market entry, and more than 70% of both men and women in the sample do not have children throughout their early careers. It is possible that schedule flexibility and work hours are not differently salient between men and women at this stage of their work life, and that preferences may change and possibly start to differ by gender as workers age and form families. This interpretation is consistent with findings by Liu (2016), who shows that gender differences in preferences for part-time work increase after marriage and childbirth, and by Hotz, Johansson, and Karimi (2018), who use mothers' preferences for non-wage job characteristics to construct an index of firms' family friendliness.

Table 7: Estimated Marginal Willingness to Pay for Amenities

	(a) Estimated preferences $\hat{\delta}_k$		(b) Wage value of benefits, $e^{-\delta_k}$	
	Women	Men	Women	Men
Schedule flexibility	.363 (.276)	.434 (.216)	.696	.648
Long hours	.457 (.334)	.463 (.228)	.633	.629
Paid parental Leave	.843 (.385)	.601 (.268)	.430	.548
Unpaid parental leave	.739 (.352)	.473 (.211)	.478	.623

Notes: NLSY97. Sequential maximum likelihood estimates of preference parameters. Panel (a) reports the estimated parameters with asymptotic standard errors in parentheses. Panel (b) reports the wage value of benefits, or workers' marginal willingness to pay. The wage value of a benefit is the minimum wage that a worker would accept in the exchange for its provision, relative to the wage a worker would accept in a firm not offering the benefit. $e^{-\delta_k} = \frac{w(a_k=1,u)}{w,(a_k=0,u)}$.

Table 8 reports the estimated features of the distributions of wages offered to male and female workers.[27] The first five columns of the table report the career-specific wage-offer parameters φ^w, μ_1^w is the estimated ability wage premium, and the last four columns

[26]In her recent work, Xiao (2021) uses Finnish data to construct a detailed occupation-specific flexibility index to capture the impact of workers' preferences for this amenity on the early career gender wage gap in the northern european country.

[27]The structural parameters estimating amenity offers are reported in section B in the Online Appendix.

report the premia or losses in the wage offers received by workers whose employer provides, respectively, flexible schedule, long hours, paid parental leave, and unpaid parental leave.

The estimated average wage offer in administrative support, social services and lower education occupations, φ_0^w, and the ability premium imply that a worker in the 56^{th} percentile of the CAT-ASVAB test-score distribution, the average score among women in this career, receives a 2.3 log-points wage offer if female, and a 2.4 log-points wage offer if male. The wage-offer gap remains approximately equal to 10 log-points among executives and professionals, where the average woman is, respectively, in the 63^{rd} and 74^{th} percentile of the CAT-ASVAB.

The estimated values of the ρ-parameters show that firms offering either paid or unpaid parental leave do pay higher wages to their employees. Since paid and unpaid parental leave are the most valuable amenities from workers' perspective, and their provision can be costly for employers, the wage-premia 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, employers requiring employees to work more than 40 hours per week offer a pay premium, while firms offering schedule flexibility tend to pay lower hourly wages. Interestingly, both the long-hours pay premium and the flexibility pay penalty are larger, in magnitude, among male workers. 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, prizing long hours less and attaching smaller penalties to shorter work-hours.

The results in table 8 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 both paid and unpaid parental leave among men and women alike. Both men and women, in fact, do experience wage increases, on average, upon entering a 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.

Table 8: Estimated wage offer parameters

	φ_0^w	φ_{ex}^w	φ_{pr}^w	φ_{ot}^w	μ_1^w	ρ^{fs}	ρ^{lh}	ρ^{pl}	ρ^{ul}
(a): Women	2.296	.157	.281	-.046	.004	-.045	.045	.243	.200
	(.144)	(.042)	(.036)	(.043)	(.035)	(.035)	(.040)	(.037)	(.035)
(b): Men	1.757	.142	.255	-.004	.160	-.077	.070	.241	.111
	(.227)	(.062)	(.056)	(.063)	(.052)	(.048)	(.053)	(.049)	(.051)

Notes: NLSY97. Sequential maximum likelihood estimates of wage offer parameters. The asymptotic standard errors in parentheses are computed using the outer product of gradients method.

3.4 Counterfactual Analyses

I now 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 growth in the gender wage gap.

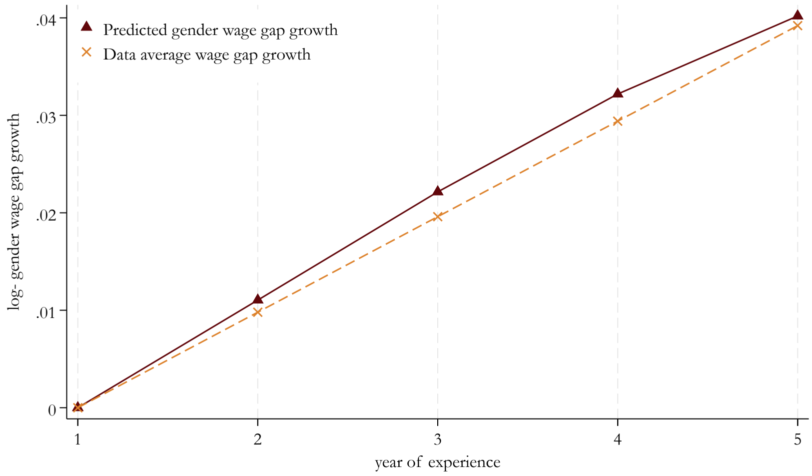
To compute workers' predicted and counterfactual wages, I simulate cross-sections of 1000 male labor market entrants and 1000 female labor market entrants. I then use the estimated parameters to model workers' yearly transitions across employment statuses and across jobs, and their wage-amenities outcomes. I perform the simulations separately by careers. Within careers, workers of the same gender in the simulated sample are homogeneous in ability. In each career and for each gender, workers' ability is measured by the gender-specific average log-percentile of the CAT-ASVAB test score observed in the data.[28] For each year in the labor market, the simulation generates a distribution of employed workers across jobs defined by pay level and amenities. I repeat each simulation

[28]The test-score percentiles used in the simulations are reported in table A6 in the Online Appendix.

100 times. The mean of the t^{th} year of experience average wage across the 100 career-specific simulations is the predicted wage in t for workers of a given gender in that career. The year-t average wage of workers of a given gender is the weighted average of the career-specific simulated wages, with weights equal to the share of workers of a given gender in each career.

Figure 7 verifies the fit of the model comparing the early-career growth in the gender (log) wage gap that the model predicts to the average wage-gap growth in the raw data.

Figure 7: Model Fit - Predicted and observed growth in the early-career gender wage gap

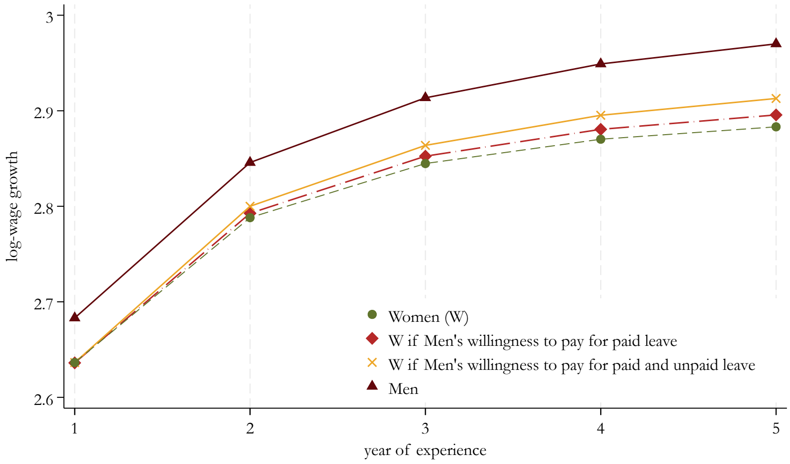


Notes: NLSY97. The dashed yellow line depicts the average growth path in the gender wage gap over workers' early careers. For every year of experience $t \in \{1, ..., 5\}$ it is computed as $\bar{g} \times (t-1)$ where $\bar{g}$ is the average of the year-by-year log-change in the mean gender wage gap observed in the data. The solid purple line depicts the growth path in the early-career wage gap predicted by the model simulations.

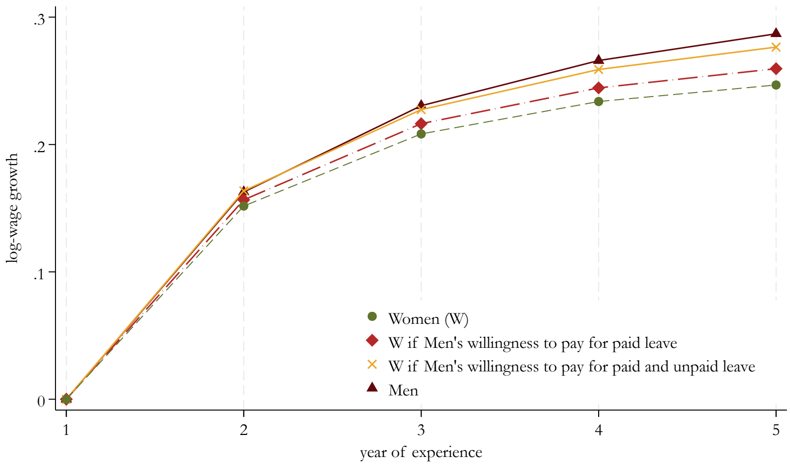
Figure 8 depicts the model-predicted wage profile of men (solid purple line) and women (dashed green line), and the counterfactual profiles of wages that women would earn if their willingness to pay for paid leave (dashed red line) and for both paid and unpaid leave (solid yellow line) were as high as men's. Panel (a) shows log-wages, while panel (b) shows log-wage growth.

Comparing men's and women's wage-growth profiles, the figure shows that the gender wage gap expands over workers' early careers due to the slow wage growth experienced by female workers. As women's counterfactual wage paths highlight, gender differences in willingness to pay for parental leave explain the bulk of the divergence in wage growth between young men and women.

Figure 8: Predicted and counterfactual log-wage growth



(a) Log wage paths



(b) Log wage growth paths

Notes: NLSY97. Predicted and counterfactual log wage paths, (a), and log wage growth paths (b).

Workers' preferences for parental leave affect earned wages by impacting the minimum wages that employees accept in order to enter a firm that provides such benefit. To the extent that women are willing to pay more than men in exchange for the provision of parental leave, they will accept lower wages, and experience lower wage growth, as they are hired by employers offering this benefit. Since the likelihood of being employed in leave-providing firms rises over time as workers search for better jobs, the gender wage gap due to willingness to pay for parental leave increases in years of experience.[29]

[29]Figure 8 shows that women's willingness to pay for the provision of parental leave does not explain the baseline gender wage gap observed in the data. This is due to the fact that the lower wages earned by women since labor market entry are mostly explained by gender differences in the distribution of workers

As figure 9 shows, while the gender wage gap increases by four log-points during workers' early careers, it would rise by less than one log-point over the same time-span if men's and women's preferences for paid and unpaid parental leave were identical. Hence, women's stronger willingness to pay for parental leave explains at least 75% of the early-career growth in the gender wage gap.

Figure 9: Predicted and counterfactual growth in gender wage gap

Notes: NLSY97. Predicted wage gap growth (purple) and counterfactual wage gap growth. Dashed red line: women assumed to have the same preferences for paid leave as men. Solid yellow line: women assumed to have the same preferences as men for both paid and unpaid leave.

In figure 10 I add the counterfactual early-career evolution in the gender wage gap that would be observed if all firms offered both paid and unpaid parental leave (dashed green line) to figure-9 results. To determine this counterfactual scenario, I predict the implications of the model assuming that the probability that firms provide paid and unpaid parental leave equals one, and that workers' preferences and search frictions, and the remaining parameters of the gender-specific job-offer distributions remain unaffected. This scenario corresponds to the partial-equilibrium impact of a law mandating and subsidizing the unconditional provision of parental leave.

As the figure shows, the early-career growth in the gender wage gap would halve if all firms offered parental leave. Assuming that this policy change would not impact male

between higher- and lower-pay careers, and by the lower wages offered to women with above-average CAT-ASVAB test scores by employers who do not provide benefits. Gender differences in offered wages may capture willingness to pay for unobserved amenities such as location, gender differences in workers' distribution across detailed occupation groups within careers, or residual discrimination.

and female workers' willingness to pay for parental leave, the pay gap growth would not disappear. In fact, stronger women's preferences for parental leave would still determine a gender gap in the average wage accepted by workers and, consequently, lower wage growth among women. To the extent that all firms offer parental leave, however, women's stronger willingness to pay for this benefit would impact accepted wages less, thus reducing the growth in the early career gender wage gap.

Figure 10: Policy-change implications

Notes: NLSY97. Predicted wage gap growth (purple) and counterfactual wage gap growth. Dashed red line: women assumed to have the same preferences for paid leave as men. Solid yellow line: women assumed to have the same preferences as men for both paid and unpaid leave. Dashed green line: wage gap if all firms provided paid and unpaid leave to all male and female workers.

The major limitation of this policy analysis is that it cannot account for the possibility that firms change their wage offers, or that workers modify their labor-supply behavior, 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, I proposed here a complementary angle from which to look at parental leave. It highlights that the scarcity of parental leave may itself be a significant institutional factor influencing workers' labor market outcomes and, potentially, gender inequality. To the extent that the choice and costs of providing parental leave are delegated to employers, only some firms will offer this benefit and workers for whom parental leave is more important will pay a higher price for it by accepting lower wages in exchange for its provision.

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 firms providing paid and unpaid parental leave offer higher wages, but women are willing to pay, respectively, 40% more and 56% more than men to be provided these benefits. The lower wages that women accept to work in firms offering parental leave explain 75% of the early-career growth in the gender wage gap.

I then showed that a policy mandating and subsidizing the unconditional provision of parental leave may halve the early-career growth in the gender wage gap. The widespread availability of parental leave would lessen workers' need to accept lower wages in exchange for it, thus reducing the gap in accepted wages between men and women entering leave-providing firms.

Though limited by 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 some employers will offer it, 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 especially detrimental to the labor market outcomes of young women with strong labor market attachment, for whom the availability of parental leave may be a crucial form of employment insurance, and foster job continuity, in the event of a childbirth.

References

Akerlof, G., Rose, A., & Yellen, a., J. L. (1988). Job Switching and Job Satisfaction in the U.S. Labor Market. *Brookings Papers on Economic Activity*, *1988*(2), 495-594.

Amano-Patiño, N., Baron, T., & Xiao, P. (2020). *Human Capital Accumulation, Equilibrium Wage-Setting and the Life-Cycle Gender Pay Gap* (Cambridge Working Papers in Economics No. 2010). Cambridge, UK: University of Cambridge, Faculty of Economics.

Angelov, N., Johansson, P., & Lindahl, E. (2016). Parenthood and the Gender Gap in Pay. *Journal of Labor Economics*, *34*(3), 545-579.

Bana, S., Bedard, K., & Rossin-Slater, M. (2020). The Impacts of Paid Family Leave Benefits: Regression Kink Evidence from California Administrative Data. *Journal of Policy Analysis and Management*, *39*(4), 888-929.

Bartel, A., Rossin-Slater, M., Ruhm, C., Slopen, M., & Waldfogel, J. (2023a). The Impact of Paid Family Leave on Employers: Evidence from New York. *Community, Work & Family*, 1-19.

Bartel, A., Rossin-Slater, M., Ruhm, C., Slopen, M., & Waldfogel, J. (2023b). The Impacts of Paid Family and Medical Leave on Worker Health, Family Well-Being, and Employer Outcomes. *Annual Review of Public Health*, *44*, 429-443.

Barth, E., & Dale-Olsen, H. (2009). Monopsonistic Discrimination, Worker Turnover, and the Gender Wage Gap. *Labour Economics*, *16*, 589-597.

Barth, E., Olivetti, C., & Kerr, S. (2021). The Dynamics of Gender Earnings Differentials: Evidence from Establishment Data. *European Economic Review*, *134*.

Bertrand, M., Goldin, C., & Katz, L. (2010). Dynamics of the Gender Gap for Young Professionals in the Financial and Corporate Sectors. *American Economic Journal: Applied Economic*, *July 2010*(2), 228-255.

Blau, F., & Kahn, L. M. (2013). Female Labor Supply: Why Is the United States Falling Behind? *American Economic Review: Papers & Proceedings*, *103*(3), 251-256.

Blau, F., & Kahn, L. M. (2017). The Gender Wage Gap: Extent, Trends, and Explanations. *Journal of Economic Literature*, *55*(3), 798-865.

Bloom, N., Guvenen, F., Smith, B., Song, J., & von Wachter, T. (2018). Is the Large Firm Wage Premium Dead or Merely Resting? *AEA Papers and Proceedings*, *108*, 317-322.

Bonhomme, S., & Jolivet, G. (2009). The Pervasive Absence of Compensating Differentials. *Journal of Applied Econometrics*, *24*, 763-795.

Booth, A., Francesconi, M., & Frank, J. (2003). A Sticky Floor Model of Promotion, Pay, and Gender. *European Economic Review*, *47*, 295-322.

Bowlus, A. (1997). A Search Interpretation of Male-Female Wage Differentials. *Journal of Labor Economics*, *15*(4), 295-322.

Bowlus, A., & Grogan, L. (2009). Gender Wage Differentials, Job Search and Part-Time Employment in the UK. *Oxford Economics Papers*, *61*(4), 275-303.

Brown, C. (1980). Equalizing Differences in the Labor Market. *The Quarterly Journal of Economics*, *94*(1), 113-134.

Brown, S., Herr, J., Roy, R., & Klerman, J. A. (2020). *Employee and Worksite Perspectives of the FMLA: Who is Eligible?* (Tech. Rep.). Rockville, MD, US: Produced for the U.S. Department of Labor, Chief Evaluation Office. Abt Associates Inc.

Burdett, K., & Mortensen, D. T. (1998). Equilibrium Wage Differentials and Employer Size. *International Economic Review*, *39*, 257-273.

Burke, J., & Miller, A. (2017). The Effects of Job Reloction on Spousal Careers: Evidence from Military Change of Station Moves. *Economic Inquiry*, *56*(2), 1261-1277.

Cain Miller, C. (2018a, February 1). Lowe's joins other big employers in offering paid parental leave. *The New York Times*. Retrieved 2023-08-15, from https://www.nytimes.com/2018/02/01/upshot/lowes-joins-other-big-employers-in-offering-paid-parental-leave.html

Cain Miller, C. (2018b, January 24). Walmart and now starbucks: Why more big companies are offering paid family leave. *The New York Times*. Retrieved 2023-08-15, from https://www.nytimes.com/2018/01/24/upshot/parental-leave-company-policy-salaried-hourly-gap.html

Caldwell, S., & Danieli, O. (in press). Outside Options in the Labor Market. *The Review of Economic Studies*.

Card, D., Cardoso, A., Heining, J., & Kline, P. (2018). Firms and Labor Market Inequality: Evidence and Some Theory. *Journal of Labor Economics*, *36*(S1), S13-S70.

Card, D., Cardoso, A., & Kline, P. (2016). Bargaining, Sorting, and the Gender Wage Gap: Quantifying the Impact of Firms on the Relative Pay of Women. *The Quarterly Journal of Economics*, *131*(2), 633-686.

Cortes, P., & Pan, J. (2019). When Time Binds: Substitutes for Household Production, Returns to Working Long Hours, and the Gender Wage Gap among the Highly Skilled. *Journal of Labor Economics*, *37*(2), 351-398.

Cortes, P., Pan, J., Pilossoph, L., Reuben, E., & Zafar, B. (2023). Gender Differences in Job Search and the Earnings Gap: Evidence from the Field and Lab. *The Quarterly Journal of Economics*, 1-58.

Couch, K., & Placzek, D. (1993). Earnings Losses of Displaced Workers Revisited. *American Economic Review*, *100*(1), 572-589.

Dale-Olsen, H. (2006). Estimating Workers' Marginal Willingness to Pay for Safety using Linked Employer-Employee Data. *Economica*, *73*, 99-127.

Del Bono, E., & Vuri, D. (2011). Job Mobility and the Gender Wage Gap in Italy. *Labour Economics*, *18*, 130-142.

Dey, M., & Flinn, C. (2005). An Equilibrium Model of Health Insurance Provision and Wage Determination. *Econometrica*, *73*, 571-627.

Dill, K., & Yang, A. (2022, August 22). Companies are cutting back on maternity and paternity leave. Retrieved 2023-08-15, from https://www.wsj.com/articles/the-surprising-benefit-some-companies-are-taking-awayparental-leave-11661125605

Egan, M., Matvos, G., & Seru, A. (2022). When Harry Fired Sally: The Double Standard in Punishing Misconduct. *Journal of Political Economy*, *130*(5), 1184-1248.

Flabbi, L., & Mabli, J. (2018). Household Search or Individual Search: Does It Matter? *Journal of Labor Economics*, *36*(1), 1-46.

Flabbi, L., & Moro, A. (2012). The Effect of Job Flexibility on Female Labor Market Outcomes: Estimates from a Search and Bargaining Model. *Journal of Econometrics*, *168*, 81-95.

Gicheva, D. (2013). Working Long Hours and Early Career Outcomes in the High-End Labor Market. *Journal of Labor Economics*, *31*(4), 785-824.

Ginja, R., Karimi, A., & Xiao, P. (2023). Employer Responses to Family Leave Programs. *American Economic Journal: Applied Economics*, *15*(1), 107-135.

Goldin, C. (2014). A Grand Gender Convergence: Its Last Chapter. *American Economic Review*, *104*(4), 1091-1119.

Goldin, C., Katz, L., & Kuziemko, I. (2006). The Homecoming of American College Women: The Reversal of the College Gender Gap. *Journal of Economic Perspectives*, *20*(4), 133-156.

Goldin, C., Kerr, S., & Olivetti, C. (2020). *Why firms offer paid parental leave: an explanatory study* (NBER Working Paper No. 26617). Cambridge, MA, United States: National Bureau of Economic Research.

Goldin, C., & Rouse, C. (2000). Orchestrating Impartiality: The Impact of "Blind" Auditions on Female Musicians. *American Economic Review*, *90*(4), 715-741.

Gronberg, T., & Reed, W. (1994). Estimating Workers' Marginal Willingness to Pay for Job Attributes using Duration Data. *The Journal of Human Resources*, *29*, 411-428.

Guler, B., Guvenen, F., & Violante, G. (2012). Joint-Search Theory: New Opportunities and New Frictions. *Journal of Monetary Economics*, *59*, 352-369.

Hotz, V., Johansson, P., & Karimi, A. (2018). *Parenthood, family friendly workplaces, and the gender gaps in early work careers* (NBER Working Paper No. 24173). Cambridge, MA, United States: National Bureau of Economic Research.

Hwang, H., Mortensen, D. T., & Reed, W. R. (1998). Hedonic Wages and Labor Market Search. *Journal of Labor Economics*, *16*(4), 815-847.

Hwang, H., Reed, W. R., & Hubbard, C. (1992). Compensating Wage Differentials with Unobserved Productivity. *Journal of Political Economy*, *100*(4), 835-858.

Jacobson, L., LaLonde, R., & Sullivan, D. (1993). Earnings Losses of Displaced Workers. *American Economic Review*, *83*(4), 685-709.

Janta, B., & Stewart, K. (2018). *Paternity and parental leave policies across the european union* (Tech. Rep.). Luxemburg: Publications Office of the European Union: RAND Europe, European Platform for Investing in Children (EPIC).

Keith, K., & McWilliams, A. (1999). The Returns to Job Mobility and Job Search by Gender. *Indutrial and Labor Relations Review*, *52*(3), 460-477.

Khandker, R. (1988). Offer Heterogeneity in a Two State Model of Sequential Search. *The Review of Economics and Statistics*, *70*(2), 259-265.

Kleven, H., Landais, C., & Søgaard, J. (2019). Children and Gender Inequality: Evidence from Denmark. *American Economic Journal: Applied Economics*, *11*(4), 181-209.

Lang, K., & Majumdar, S. (2004). The Pricing of Job Characteristics when Markets Do Not Clear: Theory and Policy Implications. *International Economics Review*, *45*(4), 1111-1128.

Lavetti, K., & Schmutte, I. (2018). *Estimating compensating wage differentials with endogenous job mobility* (Working Paper).

Le Barbanchon, T., Rathelot, R., & Roulet, A. (2021). Gender Differences in Job Search: Trading off Commute against Wage. *The Quarterly Journal of Economics*, *1*(136), 381-426.

Light, A., & Ureta, M. (1992). Panel Estimates of Male and Female Job Turnover Behavior: Can Female Nonquitters be Identified? *Journal of Labor Economics*, *10*(2), 156-181.

Light, A., & Ureta, M. (1995). Early-career Work Experience and Gender Wage Differentials. *Journal of Labor Economics*, *13*(1), 121-154.

Liu, K. (2016). Explaining the Gender Wage Gap: Estimates from a Dynamyc Model of Job Changes and Hours Changes. *Quantitative Economics*, *2016*(7), 411-447.

Loprest, P. (1992). Gender Differences in Wage Growth and Job Mobility. *AEA Papers and Proceedings*, *82*(2), 526-532.

Manning, A. (2003). *Monopsony in Motion: Imperfect Competition in Labor Markets.* Princeton and Oxford: Princeton University Press.

Manning, A. (2011). Imperfect competition in the labor market. In D. Card & O. Ashenfelter (Eds.), *Handbook of labor economics, volume 4b* (p. 973-1041).

Manning, A., & Swaffield, A. (2009). The Gender Gap in Early Career Wage Growth. *The Economic Journal*, *118*(*July*), 983-1024.

Mas, A., & Pallais, A. (2017). Valuing Alternative Working Arrangements. *American Economic Review*, *107*(12), 3722-3759.

Michelson, J. (2021, January 7). How small companies can offer great paid-leave programs. *Harvard Business Review*. Retrieved 2023-08-15, from https://hbr.org/2021/01/how-small-companies-can-offer-great-paid-leave-programs

Olivetti, C., & Petrongolo, B. (2016). The Evolution of Gender Gaps in Industrialized Countries. *Annual Review of Economics*, *8*, 405-434.

Olivetti, C., & Petrongolo, B. (2017). The Economic Consequences of Family Policies: Lessons from a Century of Legislation in High-Income Countries. *Journal of Economic Perspectives*, *31*(1), 205-230.

Robinson, J. (1933). *The Economics of Imperfect Competition.* London: Macmillan.

Rosen, S. (1974). Hedonic Prices and Implicit Markets: Product Differentiation in Pure Competition. *Journal of Political Economy*, *82*(1), 34-55.

Rosen, S. (1986). The theory of equalizing differences. In O. Ashenfelter & R. Layard (Eds.), *Handbook of labor economics, volume 1* (p. 641-692).

Rossin-Slater, M. (2018). Maternity and Family Leave Policy. In S. Averett, L. Argys, & S. Hoffman (Eds.), *The oxford handbook of women and the economy* (p. 323-344). Oxford, UK: Oxford University Press.

Royalty, A. (1998). Job-to-Job and Job-to-Nonemployment Turnover by Gender and Education Level. *Journal of Labor Economics*, *16*(2), 34-55.

Sockin, J. (2022). *Show Me the Amenity: Are Higher-Paying Firms Better All Around?* (CESIfo Working Paper No. 9842). Center for Economic Studies & Ifo Institute: CESIfo.

Sorkin, I. (2018). Ranking Firms using Revealed Preferences. *The Quarterly Journal of Economics*, *133*(3), 1331-1393.

Sullivan, P., & To, T. (2014). Search and Nonwage Job Characteristics. *The Journal of Human Resources*, *49*(2), 472-507.

Topel, R. (1991). Specific Capital, Mobility, and Wages: Wages Rise with Seniority. *Journal of Political Economy*, *99*(1), 145-176.

Topel, R., & Ward, M. (1992). Job Mobility and the Careers of Young Men. *The Quarterly Journal of Economics*, *107*(2), 439-479.

U.S. Bureau of Labor Statistics (BLS). (2022). *Bureau of Labor Statistics Employees Benefits.* Retrieved 2023-08-16, from https://www.bls.gov/ebs/home.htm

van Belle, J. (2016). *Paternity and parental leave policies across the european union* (Tech. Rep.). Cambridge, UK and Brussels, Belgium: RAND Europe.

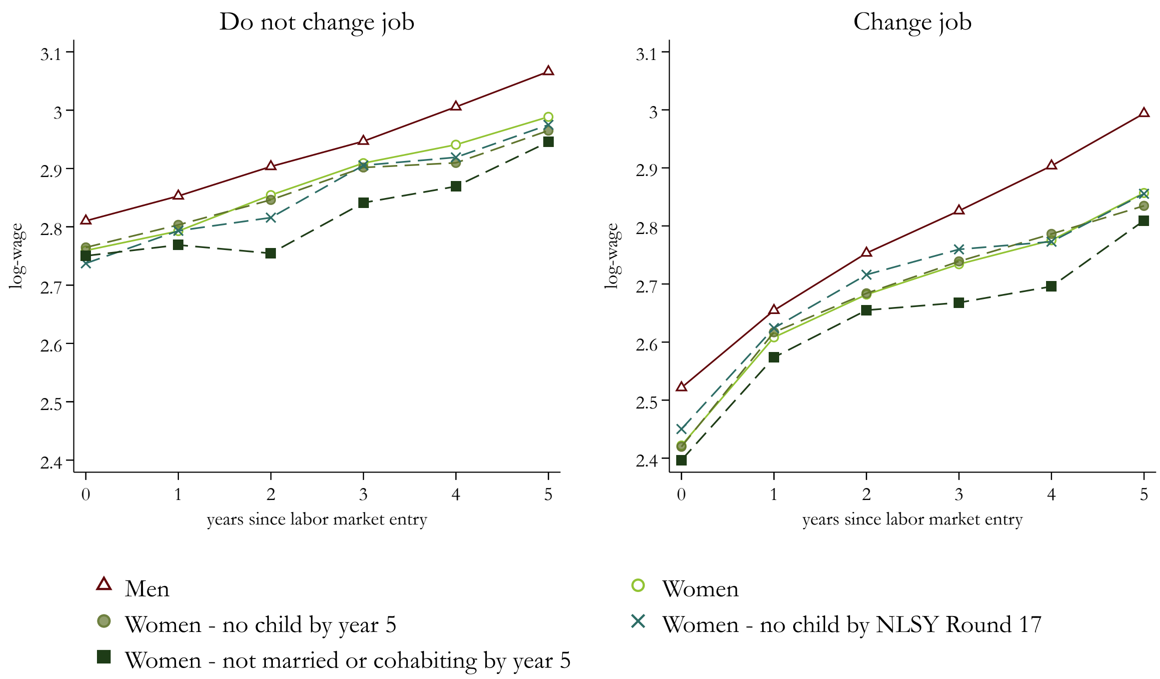
Venator, J. (2023). *Dual-earner migration decisions, earnings, and unemployment insurance* (Working Paper).

Xiao, P. (2021). *Equilibrium Sorting and the Gender Wage Gap* (VATT Working Papers No. 144). Helsinki, Finland: VATT Institute for Economic Research.

Online Appendix

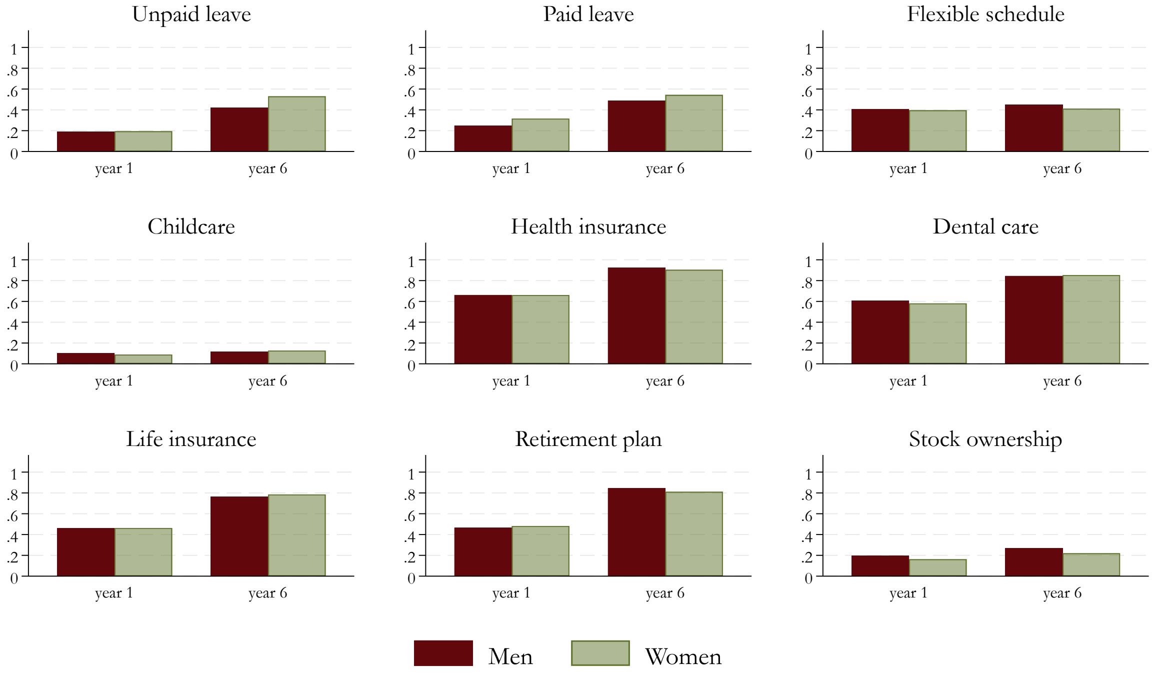
A Descriptive and Reduced Form Analyses

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



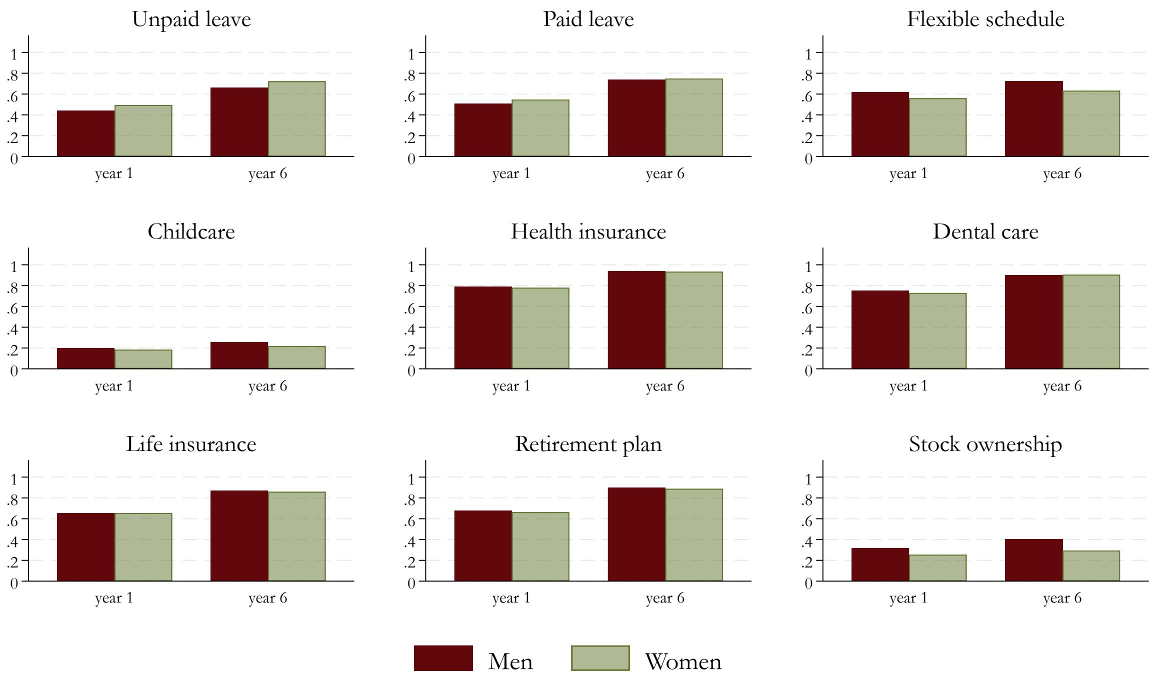
Notes: NLSY97, sample selection as in table 1. The figures depict the average log-wage profiles of men and women over years of experience. The left-hand panel refers to workers who remain with the same employer throughout their early careers, the right-hand panel refers to workers who change at least one employer during their early careers. Both figures compare men to, respectively, all women, women who do not have children by five years since labor market entry, women who do not have children by 2015 (Round 17 of the NLSY97), and women who are neither married nor cohabit by five years since labor market entry.

Figure A2: Shares of employees working for amenity-providing employers - Job changers



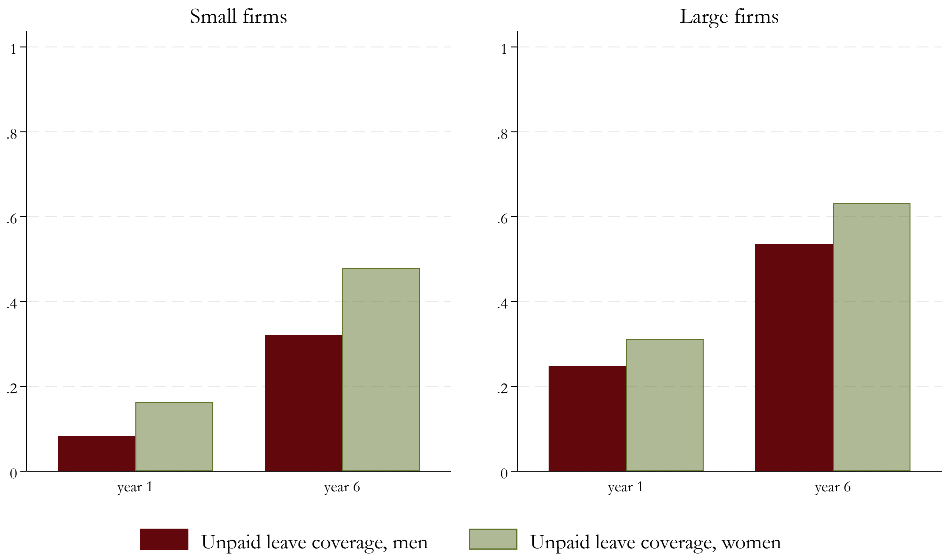
Notes: NLSY97, sample selection as in table 1. This figure only includes workers who change at least one employer during their early careers. Each panel depicts the share of men (burgundy) and women (green) who report to work for an employer offering a certain benefit or work arrangement during the first week in employment, respectively, at labor market entry (year 1) and five years later (year 6).

Figure A3: Shares of employees working for employers who will ever offer a certain benefit

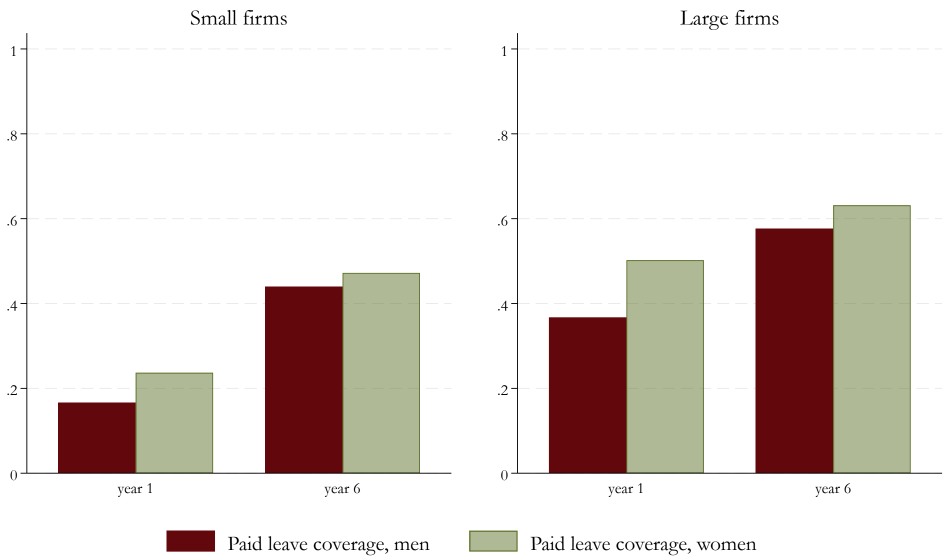


Notes: NLSY97, sample selection as in table 1. This figure only includes workers who change at least one employer during their early careers. Each panel depicts the share of men (burgundy) and women (green) whose year 1 or year 6 employer ever offers a certain benefit or work arrangement.

Figure A4: Parental leave coverage by employer size



(a) Unpaid leave



(b) Paid leave

Notes: NLSY97, sample selection as in table 1. The figures share of employees who currently work for employers providing unpaid leave (a) or paid leave (b) by employer size. Small employers are firms with less than 50 employees.

Table A1: Wage gains from job changes - All coefficients

	(1)		(2)		(3)		(4)	
	(M)	(F)	(M)	(F)	(M)	(F)	(M)	(F)
exp(t-1)	0.07***	0.07***	0.10***	0.05***	0.10***	0.05***	0.11***	0.04*
	(0.02)	(0.01)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
$\exp^2(t-1)$	0.00	-0.00	0.00	0.00	0.00	0.00	0.00	0.01
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)
exp $\times\mathbb{1}$(LPJ)	-0.04**	-0.02	-0.03*	-0.02	-0.03*	-0.02	-0.03	-0.01
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
exp $\times\mathbb{1}$(VJC)	0.05**	0.03**	0.06***	0.03***	0.06***	0.03***	0.06***	0.04***
	(0.02)	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)
exp $\times\mathbb{1}$(OJC)	0.02	-0.01	0.02	-0.01	0.02	-0.01	0.02	-0.01
	(0.02)	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)
tenure(t-1)			-0.02	0.00	-0.03	0.00	-0.01	0.01
			(0.02)	(0.02)	(0.02)	(0.02)	(0.03)	(0.02)
ten^2(t-1)			-0.01	-0.01	-0.01	-0.01	-0.01	-0.01
			(0.01)	(0.00)	(0.01)	(0.00)	(0.01)	(0.01)
Union(t-1)			-0.04	0.01	-0.04	0.01	-0.07	0.01
			(0.04)	(0.03)	(0.04)	(0.03)	(0.05)	(0.03)
U rate(t-1)			-0.01	-0.01	-0.01*	-0.00	-0.01**	-0.01
			(0.01)	(0.00)	(0.01)	(0.00)	(0.01)	(0.01)
Paid l(t-1)			-0.00	0.01	-0.00	0.01	-0.01	-0.00
			(0.02)	(0.01)	(0.02)	(0.01)	(0.03)	(0.02)
Unp l(t-1)			0.01	0.01	0.01	0.02	0.01	0.01
			(0.02)	(0.01)	(0.02)	(0.01)	(0.03)	(0.02)
Flex sch(t-1)			0.01	-0.02	0.01	-0.02	0.02	-0.01
			(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
Health in(t-1)			-0.02	0.08**	-0.03	0.08*	-0.02	0.09*
			(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.05)
Life in(t-1)			0.02	0.03	0.02	0.03	0.03	0.04
			(0.02)	(0.02)	(0.02)	(0.02)	(0.03)	(0.03)
Dental(t-1)			-0.01	-0.03	-0.00	-0.02	-0.01	-0.02
			(0.03)	(0.03)	(0.03)	(0.03)	(0.03)	(0.04)
Stock(t-1)			-0.03	-0.02	-0.03	-0.02	-0.02	-0.04
			(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.03)
Log hrs(t-1)			0.03	0.01	0.03	0.01	0.02	0.01
			(0.06)	(0.04)	(0.06)	(0.04)	(0.07)	(0.04)
Emp gaps(t-1)			-0.01	0.01	-0.00	0.00	-0.00	0.00
			(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Adj. R^2	0.18	0.13	0.19	0.14	0.20	0.14	0.22	0.14
N	1587	2253	1587	2253	1587	2253	1057	1543
$\text{Occ}_{i,j,(t-1)}$	N	N	N	N	Y	Y	Y	Y
$\text{Ind}_{i,j,(t-1)}$	N	N	N	N	Y	Y	Y	Y
All change j	N	N	N	N	N	N	Y	Y

Notes: NLSY97, sample selection as in table 1. Coefficient estimates of regression models in table 4. The estimation uses annual observations for all workers in the final sample. Standard errors in parentheses are clustered at the individual level.

Table A2: Wage gains from job changes - Women by family composition - All coefficients

	(1)	(2)	(3)	(4)	(5)
exp(t-1)	0.05***	0.06**	0.01	0.05*	0.00
	(0.02)	(0.02)	(0.03)	(0.03)	(0.03)
$\exp^2(t-1)$	0.00	0.00	0.01**	0.00	0.01*
	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)
exp $\times\mathbb{1}$(LPJ)	-0.02	-0.02	-0.05	-0.02	-0.03
	(0.02)	(0.03)	(0.04)	(0.03)	(0.04)
exp $\times\mathbb{1}$(VJC)	0.03***	0.03**	0.03	0.04***	0.04*
	(0.01)	(0.01)	(0.02)	(0.01)	(0.02)
exp $\times\mathbb{1}$(OJC)	-0.01	-0.02	-0.01	-0.02	-0.00
	(0.01)	(0.01)	(0.02)	(0.01)	(0.02)
tenure(t-1)	0.00	-0.01	0.02	0.01	0.07*
	(0.02)	(0.02)	(0.03)	(0.03)	(0.04)
ten^2(t-1)	-0.01	-0.00	-0.01*	-0.01	-0.03**
	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)
Union(t-1)	0.01	0.04	0.05	0.04	0.04
	(0.03)	(0.04)	(0.05)	(0.04)	(0.06)
U rate(t-1)	-0.00	-0.00	-0.00	-0.00	-0.01
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)
Paid l(t-1)	0.01	-0.01	0.05**	-0.01	0.05
	(0.01)	(0.02)	(0.02)	(0.02)	(0.03)
Unp l(t-1)	0.02	0.02	0.00	0.02	0.01
	(0.01)	(0.02)	(0.02)	(0.02)	(0.03)
Flex sch(t-1)	-0.02	-0.01	-0.02	0.01	-0.01
	(0.02)	(0.02)	(0.02)	(0.03)	(0.03)
Health in(t-1)	0.08*	0.11**	0.07	0.09	0.06
	(0.04)	(0.05)	(0.07)	(0.06)	(0.09)
Life in(t-1)	0.03	0.01	0.00	0.01	0.00
	(0.02)	(0.03)	(0.03)	(0.04)	(0.05)
Dental(t-1)	-0.02	-0.02	-0.05	0.01	-0.04
	(0.03)	(0.04)	(0.06)	(0.05)	(0.08)
Stock(t-1)	-0.02	-0.04	-0.03	-0.06**	-0.06
	(0.02)	(0.02)	(0.03)	(0.03)	(0.04)
Log hrs(t-1)	0.01	0.00	0.05	0.01	0.06
	(0.04)	(0.05)	(0.06)	(0.06)	(0.06)
Emp gaps(t-1)	0.00	0.00	0.03*	0.00	0.03**
	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)
Adj. R^2	0.14	0.13	0.15	0.13	0.15
N	2253	1568	965	1108	620
$\text{Occ}_{i,j,(t-1)}$	Y	Y	Y	Y	Y
$\text{Ind}_{i,j,(t-1)}$	Y	Y	Y	Y	Y
All change j	N	N	N	N	N

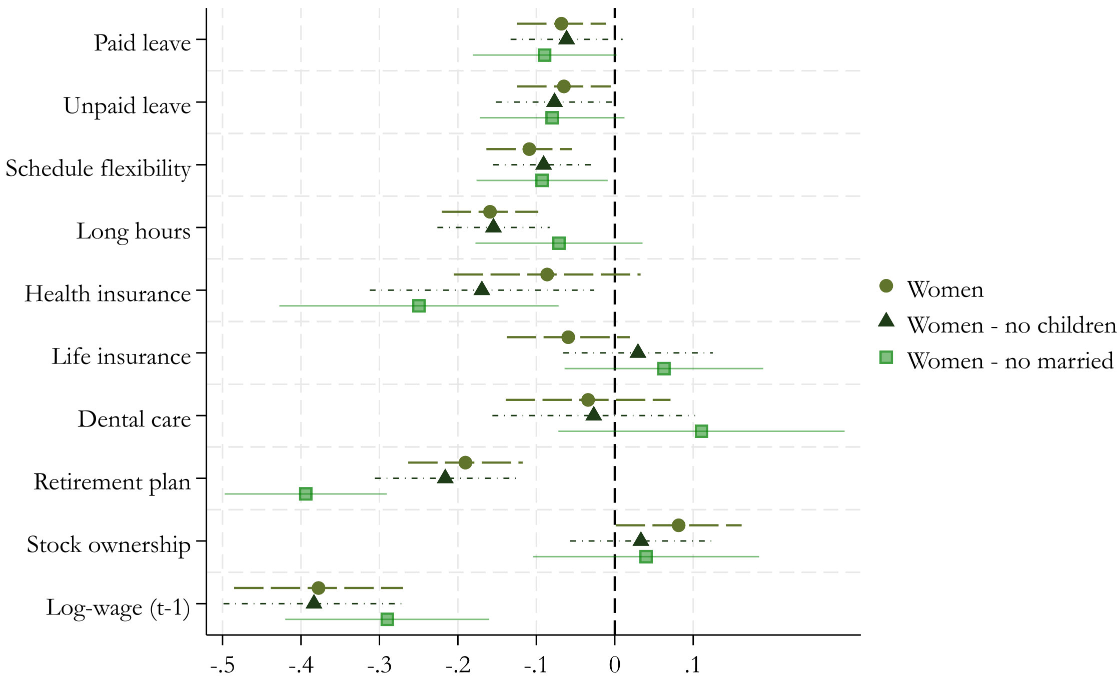
Notes: NLSY97, sample selection as in tables 1 and 4. Samples of women differ by column as in table 5.

Table A3: Linear probability model of job changes - Coefficient estimates

	Women A	Men B
Paid leave	-0.07**	-0.05
	(0.03)	(0.04)
Unpaid leave	-0.06*	-0.04
	(0.04)	(0.04)
Schedule flexibility	-0.11***	-0.14***
	(0.03)	(0.04)
Long hours	-0.16***	-0.09**
	(0.04)	(0.04)
Health insurance	-0.09	-0.10
	(0.07)	(0.08)
Life insurance	-0.06	-0.02
	(0.05)	(0.06)
Dental care	-0.03	0.02
	(0.06)	(0.08)
Retirement plan	-0.19***	-0.08
	(0.04)	(0.06)
Stock ownership	0.08*	-0.06
	(0.05)	(0.06)
exp(t-1)	0.16***	0.10**
	(0.03)	(0.04)
$\exp^2(t-1)$	-0.02**	-0.01
	(0.01)	(0.01)
Log-wage (t-1)	-0.38***	-0.45***
	(0.07)	(0.08)
Employment gaps up to (t-1)	-0.09***	-0.06***
	(0.01)	(0.02)
US Region unemployment rate (t-1)	-0.01	-0.02
	(0.01)	(0.01)
Log employer dimension (t-1)	0.00	-0.01
	(0.01)	(0.01)
Adj. R^2	0.25	0.23
N	1431	982
Occup. $(t-1)$	Y	Y
Industry $(t-1)$	Y	Y

Notes: NLSY97, sample selection as in tables 1, and 4 column (4). The table reports all coefficients of the fixed-effect linear probability model of job changes whose selected coefficients are depicted in figure 4. The sample only includes workers who change at least one employer throughout their early careers. Standard errors are clustered at the individual level.

Figure A5: Linear probability model of job changes - Women by family formation



Notes: NLSY97, sample selection as in table 1 and 4 column (4), 5 column (2) and 5 column (3). The figure reports selected coefficients of linear probability model 2 estimated for different groups of women. Controls as in figure 4.

B Model Estimation and Results

B.1 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 $\varphi_0^w + \mu_1^w b_i + \sum_{c\in\{\text{ex, pr, ot}\}} \varphi_c^w \text{c}_i = \mu^w(X)$, where $X = \{b, c\}$. Notice that

$$f(w^*, \mathbf{a}^*|.) = f(w^*|\mathbf{a}^*, .)P(\mathbf{a}^*|.) = f(w^*|\mathbf{a}^*, .)\prod_{k=1}^{K} P(a_k^*|.) \tag{16}$$

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

$$
\begin{aligned}
F(w^*|.) &= P(\mu^w(X) + \rho'\mathbf{a} + \sigma_w\varepsilon_w \leq w^*) \\
&= P\left(\varepsilon_w \leq \frac{w^* - \mu^w(X) - \rho'\mathbf{a}}{\sigma_w}\right) \\
&= \Phi\left(\frac{w^* - \mu^w(X) - \rho'\mathbf{a}}{\sigma_w}\right) \qquad (17)
\end{aligned}
$$

So that

$$
f(w^*|.) = \frac{1}{\sigma_w}\phi\left(\frac{w^* - \mu^w(X) - \rho'\mathbf{a}}{\sigma_w}\right) \qquad (18)
$$

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

Regarding $P(\mathbf{a}^*|.)$, let $\mu_0^{a^k} + \mu_1^{a^k} b_i + \sum_{c \in \{\text{ex, pr, ot}\}} \varphi_c^{a^k} \text{c}_i = \mu^{a_k}(X)$, where $X = \{b, c\}$. 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,

$$
P(a_k^*|.) = p^{a_k^*}(1-p)^{1-a_k^*} \qquad (19)
$$

Where

$$
\begin{aligned}
p &= P(\mu^{a_k}(X) + \varepsilon_{a_k} > 0) \\
&= P(\varepsilon_{a_k} > -\mu^{a_k}(X)) \\
&= 1 - \Phi(-\mu^{a_k}(X)) = \Phi(\mu^{a_k}(X)) \qquad (20)
\end{aligned}
$$

Consequently, for each amenity a_k

$$
\begin{aligned}
P(a_k^*|.) &= \Phi(\mu^{a_k}(X))^{a_k^*}(1 - \Phi(\mu^{a_k}(X)))^{1-a_k^*} \\
&= \Phi\left(\mu^{a_k}(X)(-1)^{(1-a_k^*)}\right) \qquad (21)
\end{aligned}
$$

Substituting (18) and (21) in (16)

$$
f(w^*, \mathbf{a}^*|.) = \frac{1}{\sigma_w}\phi\left(\frac{w^* - \mu^w(X) - \rho'\mathbf{a}}{\sigma_w}\right)\prod_{k=1}^{K}\Phi\left(\mu^{a_k}(X)(-1)^{(1-a_k^*)}\right) \qquad (22)
$$

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

$$\bar{F}_u(u|.) = \sum_{\mathbf{a}^* \in \{0,1\}^K} \bar{F}(u|\mathbf{a}^*, .)P(\mathbf{a}^*|.) \tag{23}$$

Where

$$\begin{aligned}
\bar{F}(u|\mathbf{a}^*, .) &= 1 - P(w^* + \delta'\mathbf{a}^* \leq u|.) \\
&= 1 - P(\mu^w(X) + \rho'\mathbf{a}^* + \sigma_w \varepsilon_w + \delta'\mathbf{a}^* \leq u) \\
&= 1 - P\left(\varepsilon_w \leq \frac{-(\mu^w(X) + \rho'\mathbf{a}^* + \delta'\mathbf{a}^* - u)}{\sigma_w}\right) \\
&= 1 - \Phi\left(-\frac{(\mu^w(X) + \rho'\mathbf{a}^* + \delta'\mathbf{a}^* - u)}{\sigma_w}\right) \\
&= \Phi\left(\frac{(\mu^w(X) + \rho'\mathbf{a}^* + \delta'\mathbf{a}^* - u)}{\sigma_w}\right)
\end{aligned} \tag{24}$$

Substituting (24) and (21) into (23)

$$\bar{F}_u(u|.) = \sum_{\mathbf{a}^* \in \{0,1\}^K} \Phi\left(\frac{(\mu^w(X) + \rho'\mathbf{a}^* + \delta'\mathbf{a}^* - u)}{\sigma_w}\right) \prod_{k=1}^{K} \Phi\left(\mu^{a_k}(X)(-1)^{(1-a_k^*)}\right) \tag{25}$$

B.2 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 $t \in [0, T = 75]$, a worker's contribution to the likelihood in $(t+1)$ in equation (12) can be rewritten as

$$l_{t+1}(\theta, \lambda, \delta) = l_{1,t+1}(\theta) \times l_{2,t+1}(\theta, \lambda, \delta) \times l_{3,t+1}(\theta, \lambda, \delta) \tag{26}$$

Where

$$l_{1,t+1}(\theta) = f(w_{t+1}, a_{t+1}; \theta)^{uj_t} \tag{27}$$

$$l_{2,t+1}(\theta, \lambda, \delta) = [1 - \lambda_1 \bar{F}(w_t + \delta' \mathbf{a}_t; \theta) - \lambda_2 - q]^{s_t} [\lambda_1 \bar{F}(w_t + \delta' \mathbf{a}_t; \theta) + \lambda_2]^{jj_t} \tag{28}$$

$$l_{3,t+1}(\theta, \lambda, \delta) = q^{ju_t} [1 - \lambda_0]^{uu_t} \lambda_0^{uj_t} \left[\frac{(\mathbf{1}\{w_{t+1} + \delta' \mathbf{a}_{t+1} > w_t + \delta' \mathbf{a}_t\} + \lambda_2) f(w_{t+1}, a_{t+1}; \theta)}{\lambda_1 \bar{F}(w_t + \delta' \mathbf{a}_t; \theta) + \lambda_2} \right]^{jj_t} \tag{29}$$

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

$$\hat{\theta} = \text{argmax}_\theta \log L_1 = \text{argmax}_\theta \sum_{i=1}^{N} \sum_{t=t_0}^{T} \log l_{1,t+1} \tag{30}$$

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

$$\begin{aligned} \hat{\lambda}^1 &= \text{argmax}_\lambda \log L_2 + \log L_3 = \\ &= \text{argmax}_\lambda \sum_{i=1}^{N} \sum_{t=t_0}^{T} \log l_{2,t+1}(\hat{\theta}, \lambda, \tilde{\delta}) + \log l_{3,t+1}(\hat{\theta}, \lambda, \tilde{\delta}) \end{aligned} \tag{31}$$

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

$$\hat{\delta}^1 = \text{argmax}_\delta \log L_2 = \text{argmax}_\delta \sum_{i=1}^{N} \sum_{t=t_0}^{T} \log l_{2,t+1}(\hat{\theta}, \hat{\lambda}^1, \delta) \tag{32}$$

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 $t \in (1, 76)$.

Table A4: Initial conditions in maximum likelihood estimation

δ_{fs}	0.5
δ_{lh}	0.5
δ_{pl}	0.5
δ_{ul}	0.5
λ_0	0.15
λ_1	0.15
λ_2	0.15
q	0.15
φ_0^w	2
φ_{ex}^w	0.1
φ_{pr}^w	0.1
φ_{ot}^w	0.1
ρ^{fs}	0.1
ρ^{lh}	0.1
ρ^{pl}	0.1
ρ^{ul}	0.1
σ^w	1
φ_0^{fs}	0.1
μ_1^{fs}	0.1
φ_{ex}^{fs}	0.1
φ_{pr}^{fs}	0.1
φ_{ot}^{fs}	0.1
φ_0^{lh}	0.1
μ_1^{lh}	0.1
φ_{ex}^{lh}	0.1
φ_{pr}^{lh}	0.1
φ_{ot}^{lh}	0.1
φ_0^{pl}	0.1
μ_1^{pl}	0.1
φ_{ex}^{pl}	0.1
φ_{pr}^{pl}	0.1
φ_{ot}^{pl}	0.1
φ_0^{ul}	0.1
μ_1^{ul}	0.1
φ_{ex}^{ul}	0.1
φ_{pr}^{ul}	0.1
φ_{ot}^{ul}	0.1

B.3 Estimated Parameters

Table A5: Estimated variance of wage offers and amenity offer parameters

	Women		Men	
	Coef.	Std.Err.	Coef.	Std.Err.
σ^w	0.406	0.007	0.439	0.013
φ_0^{fs}	0.414	0.444	1.316	0.818
μ_1^{fs}	-0.106	0.107	-0.295	0.190
φ_{ex}^{fs}	0.389	0.134	0.331	0.183
φ_{pr}^{fs}	0.376	0.136	0.204	0.183
φ_{ot}^{fs}	0.251	0.138	0.286	0.187
φ_0^{lh}	-3.332	0.687	-3.519	0.719
μ_1^{lh}	0.612	0.165	0.640	0.164
φ_{ex}^{lh}	0.047	0.141	0.171	0.191
φ_{pr}^{lh}	-0.243	0.151	0.110	0.193
φ_{ot}^{lh}	-0.356	0.174	-0.007	0.200
φ_0^{pl}	0.211	0.480	1.479	0.878
μ_1^{pl}	0.018	0.116	-0.360	0.207
φ_{ex}^{pl}	-0.054	0.140	0.123	0.207
φ_{pr}^{pl}	0.023	0.144	0.143	0.199
φ_{ot}^{pl}	-0.349	0.145	0.020	0.218
φ_0^{ul}	-0.617	0.440	0.396	0.898
μ_1^{ul}	0.195	0.106	-0.153	0.212
φ_{ex}^{ul}	-0.076	0.135	0.167	0.213
φ_{pr}^{ul}	-0.021	0.145	0.136	0.212
φ_{ot}^{ul}	-0.330	0.154	0.096	0.231

Notes: NLSY97. Sequential maximum likelihood estimates of wage offers variance and amenity offers parameters, with asymptotic standard errors calculated through the outer product of gradients.

B.4 Counterfactual Exercises

Table A6: Average CAT-ASVAB test score percentiles

	Men	Women
Admin/Educ/Health Support career	68	56
Exec career	71	63
Profess/Health techn career	77	74
Other career	68	56

Notes: NLSY97, sample selection as in table 1. Career and gender specific averages of the CAT-ASVAB test score percentiles are calculated in the first week in which an individual is observed.