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Gender Differences in Job Search Among Young Workers: A Study Using Displaced Workers in the United States

Astrid Kunze* and Kenneth R. Troske†

This article investigates gender differences in job search, job tenure, and wages, whether these differences vary over the early part of the life-cycle, and whether they are associated with fertility decisions. Using data from the National Longitudinal Survey of Youths on highly attached displaced workers aged 20 to 45, we find that 20- to 29-year-old women and women older than 40 experience longer spells of displacement than comparable men, but that time to a new job is similar by gender for those between 30 and 39 years of age. The age pattern in male–female wage differences in the post-displacement job is similar, with the largest differences occurring at ages 20 to 29 and over 40. We find no gender differences in tenure in the post-displacement job. We interpret the differences for the younger ages to be related to fertility and we provide evidence that supports this view.

JEL Classification: J31, J63, J64, J71

1. Introduction

Researchers have found a large and persistent gap in wages between men and women in the United States even after controlling for observable differences in characteristics (e.g., Altonji and Blank 1999). One explanation is that women face workplace discrimination (Becker 1971) which implies that the marginal employer must have a distaste for hiring women and that competition is not driving discriminatory employers out of the market (e.g. Black and Strahan 2001; Hellerstein, Neumark, and Troske 2002; and Ederington, Minier, and Troske 2009). An alternative explanation is that differences in the value men and women place on nonmarket time can lead to differences in search behavior (Mortensen 1986 and Mortensen and Pissarides 1999). In any basic search model (e.g., Burdett and Mortensen 1998), differential search behavior, and therefore differences in wages, can arise between groups of workers, if the groups differ in the value they place on nonmarket time. One possible source for variation in the value of nonmarket time between men and women is women's comparative advantage in childbearing and childrearing. If this is the case, then differences in search behavior should be most prevalent during women's prime child-bearing years and should be correlated with fertility.

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* NHH Norwegian School of Economics, Helleveien 30, N-5045 Bergen, Norway; E-mail: astrid.kunze@nhh.no.

† Department of Economics, University of Kentucky, 255G Gatton College of Business and Economics, Lexington, KY 40506, USA; E-mail: ktroske@uky.edu; corresponding author.

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In this article, we focus on this later hypothesis by examining gender differences in search duration, how these differences vary over the life-cycle, and whether these differences are associated with fertility. Previous empirical studies of gender differences in search behavior tend to find that women search longer for a new job. However, these studies do not examine whether gender differences in job search vary with age or whether they are related to fertility. The only study that we are aware of that does examine how gender differences in search vary over the life-cycle is a companion article by us (Kunze and Troske 2012), which uses panel data on German workers and finds that women 20 to 35 years old search longer for a new job than men, but men and women ages 36 to 50 experience similar lengths of displacement. While these results are consistent with the hypothesis that gender differences in search behavior are related to child bearing and rearing, the German data do not contain direct measures of fertility. In contrast, the data used in this current study do contain measures of fertility allowing us to investigate a more direct link between fertility, job search, and post-displacement wages.

We examine gender differences in the duration of job search, subsequent job tenure, and wages, focusing on highly attached younger workers searching for a job following displacement due to a plant closing. An ideal experiment to examine gender differences in job search would be to have data on a sample of workers who randomly switch jobs. Data on all unemployed are difficult to reconcile with such a setting as unemployment can result from both involuntary as well as voluntary separations. Assuming that a plant closing is independent of the behavior of workers our data will come closer to the ideal data than data including workers who chose to switch jobs.

Our main data on displaced workers come from the 1979 National Longitudinal Survey of Youths (NLSY) data. These data are useful for our study because they identify worker displacement relatively soon after it occurs, they cover a long time period allowing us to follow workers for a number of periods after they are displaced, and they contain a rich set of covariates including information on the number of children living in a household.

We find that the length of displacement between men and women is largest for the youngest workers, age 24 years or younger, smaller for workers age 25 to 29, and essentially zero for workers age 30 to 35 and 35 to 39. Women 40 to 45 years old have significantly longer periods of displacement than men of the same age. The age pattern for the 20 to 39 year olds closely matches fertility patterns for women. Once we limit our sample to women for whom the number of children in the household does not increase prior to their finding a new job, we find no significant difference in the length of displacement between men and women up to age 39. Gender differences among the oldest workers appear unrelated to our measure of fertility.

We also find that there is no statistically significant difference between men and women of any age in the length of tenure on their post-displacement job. Finally, while we find that women do earn lower wages than men in their post-displacement job, these differences are largest for the youngest women in our sample. Limiting our sample to workers whose spell of displacement is relatively short, and to women who find a job without experiencing an increase in the number of children in the household, the male/female wage gap falls for the youngest women in our sample. Finally, we present some weak evidence showing that young women in the labor market immediately after being displaced experience wage gains relative to young men.

While we find evidence of a consistent life-cycle pattern in search behavior and wages, there are potential concerns about whether our results are affected by nonrandom selection of who is displaced. We present some arguments and evidence suggesting that our results are robust to these forms of selection.

2. Previous Work on Gender Differences in Search and Wages

It is widely recognized that any basic equilibrium search model in which workers differ in their value of nonmarket time can generate predictions that capture gender differences in the labor market even in the absence of discriminatory behavior on the part of employers. In particular, Burdett and Mortensen (1998) show that the unemployment rate, the length of a job-less spell, and the wage distribution will all be a function of the worker's value of nonmarket time. When considering differences between men and women, one possibility is that men and women have different nonmarket opportunities because of women's comparative advantage in having and raising children. This hypothesis goes back to Becker (1981) and is based on empirical evidence showing that women specialize in child care and home production while men specialize in market work. As having and raising children primarily occurs at younger ages, it seems likely that the differences in search behavior due to relatively higher reservation utility should occur at the early stage of careers—younger women will have longer spells of displacement but, once women leave the prime child bearing ages, there will be no difference in the length of displacement between men and women.

The articles by Black (1995) and Bowlus and Eckstein (2002) show that it is possible to incorporate employer discrimination into an equilibrium search model and produce additional implications. In particular, these models demonstrate that in the presence of employers with discriminatory tastes toward women, women will search longer than men for jobs, will have lower wages after search, and will have a lower match quality, which leads to shorter tenure at the firm. We will examine the additional implications from these models in our empirical analysis.

Previous work examining gender differences in search behavior using U.S. data finds that women experience longer spells of displacement (Podgursky and Swaim 1987; Farber 1997; Abbring et al. 2002; Kletzer and Fairlie 2003; Hu and Taber 2011), but the evidence on gender differences in wage changes after displacement has been mixed with some studies showing greater wage loss for women after displacement (Madden 1987; Podgursky and Swaim 1987; and Crossley, Jones, and Kuhn 1994), some studies showing similar loss in wages after displacement (Jacobson, LaLonde, and Sullivan 1993; Hu and Taber 2011), and at least one study showing that women experience smaller wage losses after displacement (see Kletzer and Fairly 2003).[1]

It is an empirical question whether gender differences in job search vary over the life cycle and, to our knowledge, no evidence exists on this question with the exception of evidence on Germany presented in Kunze and Troske (2012). In this companion article, we present results from hazard models estimated using German register data consisting of highly attached workers who are displaced from their employer. We find that women tend to search longer than men for a job after displacement, but this difference is mainly driven by young women ages 20 to 35. Younger women search longer than younger men, but there is no difference in the length of search among older workers (workers older than 35 but younger than 50). We hypothesize that this difference in search behavior over the life-cycle is related to fertility behavior, but data limitation prevent us from directly examining the relationship between

[1] One possible source for the various findings is the different data used by these studies. Kletzer and Fairly (2003) used the NLSY data for the period 1984–1993, and compare young displaced workers to young nondisplaced workers. Podgursky and Swaim (1987), Farber (1997), Abbring et al. (2002), and Hu and Taber (2011), all used the Displaced Worker Survey but for different time periods. The study by Jacobsen et al. (1993) used employer–employee matched data for Pennsylvania. Crossley et al. (1994) used data from Canada.

fertility and search. We are unaware of any study examining whether gender differences in job search over the life cycle are related to fertility.

Our analysis provides several contributions to the literature on worker search and the potential causes of the male/female wage gap. First, this is one of the few studies to exploit the NLSY data with its long panel and rich set of covariates, enabling us to provide additional evidence on the possible sources of any gender differences in post-displacement job search. Second, as far as we are aware, ours is the first paper to use data on U.S. workers to examine changes over the life-cycle in search behavior and whether these changes are related to fertility decisions. Finally, as a minor extension, we add evidence on the length of the post-displacement job that theory predicts should be shorter for workers that experience discrimination (Bowlus and Eckstein 2002).

3. The Data

Our data on displaced workers come from the 1979 NLSY, which cover the period 1979–2002. The NLSY is a panel data set on individuals who were between 14 and 22 years old in 1979 when the survey began. They were interviewed every year between 1979 and 1994 and every two years after 1994.

The NLSY79 includes a subsample that is representative of the noninstitutionalized civilian population, a subsample that over samples Hispanics, blacks, and the poor, and a subsample of individuals serving in the military. To ensure that we have a sufficiently large sample of displaced workers, our initial sample includes all three subsamples. We do not use the population weights in any of our analyses.

Data from each interview contain information on up to five jobs that individuals have held since the last interview along with information that allows jobs to be linked across interviews. For every job, we know the date individuals start a job, as well as the number of weeks they have worked at a job up to the date of the interview. When workers leave a job there is information on why they left the job, the date they left the job, and the total number of weeks they worked at the job prior to leaving. To focus on the search behavior of individuals who have left school, we only consider workers who indicate they were not enrolled in school at any time since the last interview. We then identify workers as being displaced from a job when they indicate they are no longer working at a job because their plant closed and when the data indicate they are not working for the same employer in the subsequent interview. Since the information on plant closing is not available prior to 1984, we can only identify displacement starting with the 1984 interview.[2] To limit ourselves to workers with reasonably strong attachment to the labor market, we only consider workers who are displaced from a job where they usually worked 20 or more hours a week. Finally, we drop workers who have hourly wages below \$1.50 in 1993 dollars, or who have missing values for usual hours worked per day.[3]

We identify a worker's post-displacement job as the first job we find where they usually worked more than 20 hours a week and worked at the job for more than nine

[2] To reduce possible problems with recall bias we only keep workers who were interviewed in the wave prior to displacement.

[3] We drop workers with wages below \$1.50/hour because these are abnormally low wages representing less than one percent of workers, and this wage is below the minimum wage for all years in our sample. All dollar figures have been converted to 1993 dollars using the CPI-U.

weeks.[4] For our main analysis, we measure the length of displacement as the number of weeks between the date they stopped working at the displacement job and the date they started at their post-displacement job. We consider workers who never find a job as being censored.[5] We drop the few workers who have multiple jobs when they are displaced because they will have a negative length of displacement.

Our final data contains 1322 displacement events. As the same person can experience multiple displacements over the period of our data and appear in the data more than once, we adjust all of the standard errors in our subsequent analysis to reflect this clustering. Table 1 presents summary statistics for the main variables we use in the analysis. In these data age, weeks of tenure, years of schooling and wages at displacement are all measured in the interview year where displacement occurs. Because weeks of experience, marital status and location information could be affected by the timing of displacement relative to the timing of the interview, we measure these variables in the interview prior to displacement. We measure weeks of experience for workers by summing the variable measuring the number of weeks worked at the previous interview over all interviews.[6]

Looking at Table 1 we see that 43% of displacement events involve women. We also see that this is a fairly young sample as workers are only 31 years old at displacement on average. Seven percent of workers never find a post-displacement job. For those who do find a job after displacement, the average period between jobs is about 48 weeks.

Comparing the numbers in columns 3 and 5 in Table 1 we see that men and women are about the same age at displacement, but that women are much more likely to be married and less likely to be non-white. Women also have accumulated less experience, but have about the same tenure at the displacement firm. Women are more likely to have a censored spell than men (91% vs. 95%) and, conditional on finding a job, women have almost twice as many weeks between jobs as men but they experience a similar change in hourly wages between the displacement and post-displacement job.[7]

One question that arises in our analysis is whether our measure of fertility is simply capturing whether or not women leave the labor market after displacement. We examine this question by first creating a dummy variable that equals one if a worker leaves the labor market at some point during their displacement spell. In Table 1, we report the proportion of workers for whom this dummy variable equals one. Comparing this statistic for men and women shows that women are much more likely to exit the labor market during displacement, with 43% of women reporting being out of the labor market at some point compared to 27% of men. To further examine this hypothesis we

[4] Prior to 1988 the NLSY did not collect any wage, industry, or occupation information on jobs where an individual worked fewer than 20 hours a week or had fewer than 10 weeks of tenure. To be consistent, we imposed this restriction in every year.

[5] While there is significant attrition from the sample, we do not consider a worker as being censored if they do not appear in a subsequent interview. This is because workers frequently reappear in the survey after missing one or more interviews.

[6] This will measure actual weeks worked for workers who enter the labor market after 1978. For the handful of workers who start working prior to 1978, this variable will measure total experience in the market since 1978. In our sample, 54% of people were working in 1978. However, only 12% were working full time and the average hours worked in 1978 was 573.

[7] Appendix A columns 1 and 2 present the percentage of displacements in our sample by year for men and women. Comparing the distribution of displacements for men and women shows that both men and women are more likely to experience displacement earlier in the period and that there is not any notable difference in the timing of displacement by sex.

Table 1. Summary Statistics for the Sample of Displaced Workers, Main Sample and Limited Sample

	Full Sample—Means (1)	Full Sample—Standard Deviation (2)	Men—Means (3)	Men—Standard Deviation (4)	Women—Means (5)	Women—Standard Deviation (6)	Women Limited Sample—Means (7)	Women Limited Sample—Standard Deviation (8)	Women Excluded from Limited Sample—Mean (9)	Women Excluded from Limited Sample—Standard Deviation (10)
	1.00		0.57		0.43		0.37			
Proportion less than 24 years old	0.16		0.17		0.15		0.12		0.30	
Proportion 25 to 29 years old	0.30		0.29		0.30		0.29		0.34	
Proportion 30 to 34 years old	0.21		0.22		0.20		0.20		0.18	
Proportion 35 to 39 years old	0.21		0.20		0.23		0.24		0.09	
Proportion 40 or older	0.11		0.11		0.12		0.12		0.09	
Weeks of full time experience prior to displacement	451.66	274.55	470.13	277.39	427.203*	269.03	441.92	271.75	328.8**	228.33
Weeks of tenure in job displaced from	154.97	198.09	155.07	199.58	154.84	196.28	159.97	201.8	120.55	150.96
Highest grade completed at displacement in years	12.20	1.99	12.15	2.04	12.27	1.92	12.33	1.95	11.86**	1.66
Proportion post displacement job observed	0.93	0.25	0.95	0.23	0.91	0.29	0.93	0.25	0.76	0.43
Length of displacement in weeks[a]	47.85	119.65	38.94	105.35	59.653*	135.48	36.25	70.5	216.18**	283.97
Proportion full time job prior to displacement	0.87	0.34	0.92	0.27	0.80	0.40	0.80	0.4	0.74	0.44
Proportion full time job in first post displacement job[a]	0.75	0.43	0.84	0.36	0.62	0.49	0.67	0.47	0.30	0.46
Tenure in post displacement job in weeks	140.40	187.58	138.78	184.94	142.62	191.30	145.04	193.23	122.71	174.9
Proportion change industry	0.60	0.49	0.60	0.49	0.59	0.49	0.58	0.49	0.72**	0.45
Proportion change occupation	0.57	0.49	0.58	0.49	0.56	0.50	0.54	0.5	0.72**	0.45
Proportion leaving the labor market during displacement	0.34	0.47	0.27	0.44	0.43	0.50	0.32	0.47	0.69	0.47
Hourly wage on displacement job	9.26	6.45	10.16	6.87	8.07*	5.63	8.19	5.8	7.26	4.24
Change in log hourly wage	−0.01	0.45	−0.01	0.44	−0.01	0.48	−0.007	0.448	−0.07	0.664
Log hourly wage displacement job	2.07	0.52	2.17	0.51	1.94*	0.51	1.95	0.51	1.87	0.46

Table 1. (*Continued*)

	Full Sample—Means (1)	Full Sample—Standard Deviation (2)	Men—Means (3)	Men—Standard Deviation (4)	Women—Means (5)	Women—Standard Deviation (6)	Women Limited Sample—Means (7)	Women Limited Sample—Standard Deviation (8)	Women Excluded from Limited Sample—Mean (9)	Women Excluded from Limited Sample—Standard Deviation (10)
Hourly wage post displacement job	9.21	6.68	10.17	7.62	7.921*	4.86	8.03	4.92	7.00	4.25
Proportion married	0.62	0.49	0.57	0.50	0.68*	0.47	0.68	0.47	0.72	0.45
Non-white	0.48	0.5	0.50	0.50	0.44*	0.50	0.45	0.5	0.39	0.49
AFQT score	31.38	25.89	30.90	0.97	32.20	1.03	32.81	25.19	26.59**	21.19
Number of displacement observations	1322		753		569		495		74	

Note: The limited sample in columns (8) and (9) only includes women who are living in households where the number of children did not increase between displacement and the post-displacement job. The sample in columns (9) and (10) only includes women who are living in a household where the number of children increased between displacement and the post-displacement job.

[a]Indicates that these means only include noncensored observations. * Indicates that the difference in means between men and women is statistically significant at the 95% confidence level. ** Indicates that there is a significant difference in means between the limited sample of women and the excluded women at the 95% confidence level.

divide our sample of women into two groups—a limited sample of women that live in households where the number of children in the household remains constant or declines during their displacement spell and women who are excluded from this limited sample. The means for these two groups are presented in columns 7 and 9, respectively. Examining the numbers in these two column shows that 32% of women in the limited sample leave the labor market during displacement—which is comparable to the 27% of men who exit—while close to 70% of women in the excluded sample exit the labor market during displacement.[8] These results provide some support for our claim that our measure of fertility could have a differential impact on search than simply causing women to drop out of the labor market.

Table 2 presents more detailed information on the length of displacement. The numbers in this table include both censored and uncensored spells.[9] The numbers in Panel A show that about 20% of displaced workers get a new job within a week, but men are 5 percentage points more likely than women to obtain a job this quickly. Panel A also shows that approximately 10% of our sample does not obtain a new job within two years of being displaced and that women are about twice as likely as men to experience a gap of two years or longer between jobs.

In Panels B and C, we report spell lengths separately for workers who never leave the labor market during displacement and those who do leave the labor market. The numbers in Panel B show that, conditional on staying in the labor market, men and women experience similar spells of displacement, while the results in Panel C show that leaving the labor market results in much longer displacement spells for women relative to men.

These data have a number of characteristics that make them well suited for identifying differences in search behavior between men and women. First, the frequency and length of these data should minimize the effect of recall bias that is present in other data sources, such as the Displaced Workers Survey.[10] Second, the length of these data ensures that we are able to follow workers for an extended period after displacement; minimizing the effect of censoring. Third, they allow us to accurately measure search that occurs after an exogenous job loss. By focusing on differences in search behavior between men and women after an exogenous job loss our results should be less sensitive to unobserved factors, such as differential investment in human capital, that could affect the search behavior of individuals who are initially entering the labor market or who are returning to the labor market after a prolonged absence.[11] Our assumption is that unobserved differences will be captured in the characteristics of the previous job and by focusing on changes we will control for these unobserved differences. Fourth, these data also contain a rich set of covariates measured at the individual, household, and employer level that can be used to control for other factors

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[8] We have also estimated the probability that a worker leaves the labor market using a logit model controlling for demographic characteristics such as sex, age, experience, tenure, education, and whether the worker lives in a household where the number of children increased during displacement. The results from this estimation show that women were 12% more likely than men to exit the labor market, but women living in households where the number of children increased were 26% more likely to exit the market than comparable women in households where the number of children did not increase. These results are available on request.

[9] When constructing the means in this table, for censored spells we measure the length of displacement as the number of weeks between the date of displacement and December 31, 2002. In our hazard analysis, someone is considered censored if they have not found a job by the date of their last interview.

[10] These data are still based on individual responses and, therefore, still contain measurement error. Chesher, Dumangane, and Smith (2002) show that in duration models error in the measure of spell length leads to inefficient estimates and have a complex and difficult to sign effect on the coefficients.

[11] One possible concern with these data is that men and women have different attrition rates from the sample. However, data from the 2002 wave show that 65% of men and 68% of women remain in the sample—so it does not appear that attrition varies by sex.

Table 2. Distribution of Displacement by Length of Displacement, Percentages

	Entire Sample (1)	Men (2)	Women (3)
Panel A—All Displaced Workers			
Up to 1 week	20.50	22.71	17.57
2–20 weeks	40.24	41.83	38.14
21–32 weeks	10.29	9.96	10.72
33–52 weeks	9.08	8.90	9.31
53–104 weeks	10.21	9.69	10.90
More than 104 weeks	9.68	6.91	13.36
Panel B—Displaced Workers Who Never Leave the Labor Market			
Up to 1 week	30.82	31.09	30.37
2–20 weeks	44.18	43.82	44.76
21–32 weeks	9.47	8.73	10.74
33–52 weeks	6.96	7.27	6.44
53–104 weeks	5.48	5.64	5.21
More than 104 weeks	3.08	3.45	2.45
Panel C—Displaced Workers Who Leave the Labor Market During Displacement			
Up to 1 week	0.00	0.00	0.00
2–20 weeks	32.51	36.45	29.22
21–32 weeks	11.88	13.30	10.70
33–52 weeks	13.23	13.30	13.17
53–104 weeks	19.51	20.69	18.52
More than 104 weeks	22.65	16.26	27.98

Note: All columns within each panel sum to 100. As reported in Table 1, 5% of men and 9% of women are never observed in a post-displacement job.

that could affect the length of search. Finally, the fact that we are focusing on a single cohort of workers means we should minimize any differential change in the labor force participation of men and women.

One concern about our empirical approach is our assumption that job loss due to displacement is exogenous (Bowlus and Vilhuber 2002 and Lengermann and Vilhuber 2002). Two possibilities are that more able workers leave the plant prior to its closing or that plants containing the least able workers are more likely to close. However, as we are focusing on differences between men and women, this possibly endogenous relationship between displacement and worker ability should have a similar impact on men and women.[12]

Another concern is whether women and men react differently to plant closings, possibly due to different roles within the household. For example, high ability men might be more likely to leave prior to a plant closing than high ability women. We investigate this possibility in several ways. First, we compare scores from the Armed Forces Qualifying Test (AFQT) for both the entire sample of men and women in the NLSY as well as our sample of displaced workers.[13] In neither sample is there a significant difference in AFQT score between men and women (see Table 1). Next, in Appendix A columns 3 to 6 we present the proportion of individuals that are married and the proportion that are working full-time, by year, for the entire sample of men and women in the NLSY.

[12] One change over this period that could affect the type of workers displaced is the passage of the Worker Adjustment and Retraining Notification (WARN) Act, which requires employers with 100 or more employees to notify workers 60 days prior to a layoff or plant closing that will affect 50 or more employees. However, WARN became effective in February 1989, which is early in our sample. In addition, WARN should have a similar impact on men and women. For these reasons, we do not believe that this change impacted our results.

[13] The AFQT is a test that was administered to all respondents in the 1979 NLSY in 1980.

Comparing these numbers with the corresponding numbers in Table 1 shows that our sample of displaced workers are more likely to be married and more likely to work full-time than the typical worker in the NLSY. This is consistent with our focus on workers with strong labor market attachments. However, it does not appear that among our displacement sample women are more likely to be married and less likely to work full-time than men. These results, taken together, suggest that the selection into being displaced does not differ between men and women.

While these data have a number of strengths, they also have some obvious weaknesses; the main one being the variation in age. First, because these workers are between 14 and 22 years old in 1979, the range of age at displacement is limited, varying between 19 and 45 years old, so we have limited observations on workers beyond prime child bearing and child rearing ages. Second, because this is a cohort of workers of about the same age, the main source of variation in age is due to the timing of displacement—older displaced workers are those who are displaced later in the sample period. The fact that we are focusing on differences between men and women should again help mitigate this problem. To further control for possible year effects, we include dummy variables for the year of displacement in all of our regressions.

One additional potential weakness of our analysis is the fact that we only observe a post-tenure job and wages for those who find a job after displacement. As we are comparing differences between men and women who have strong attachments to the labor market, we expect there will be smaller differences in search behavior among men and women in our sample than would be found in a random sample of workers searching for a job. Therefore, we expect that selection-bias will have a smaller impact on our results. In addition, the fact that we have a long panel of data minimizes this problem. The statistics in Table 1 show that we observe post-displacement jobs for 91% to 95% of our sample.

4. Empirical Examination of Men's and Women's Search Behavior

Length of Displacement

To demonstrate the differences between men and women in the length of displacement, we start with Figure 1, which presents estimates of Kaplan–Meyer survival functions for men and women. In this figure, the estimated survival function for women lies completely above the survival function for men indicating that women experience longer spells between jobs. The 95% confidence bands in the figure show that the difference in spell length is significant for spell lengths of less than 200 weeks. Given the relatively few displacement spells lasting longer than 200 weeks, differences are no longer significant after this point.[14]

Figure 2 presents estimates of the survival function for men and women by age categories. Throughout this analysis, we divide workers into five age categories: 24 years old or younger; 25 to 29 years old; 30 to 34 years old; 35 to 39 years old; and 40 years old or older. The summary statistics in Table 1 include the distribution of observations by these age categories. Figure 2 shows that there is an observable difference between men and women in the estimated survival function for workers that are 24 years old or younger, that the difference appears smaller for workers 25 to 29 years old, that the two functions appear virtually identical for 30 to 34 year olds and 35 to

[14] A log-rank test for the differences in the survival functions fails to reject the hypothesis that the two functions are the same. We suspect this is because of the number of very short and long spell lengths.

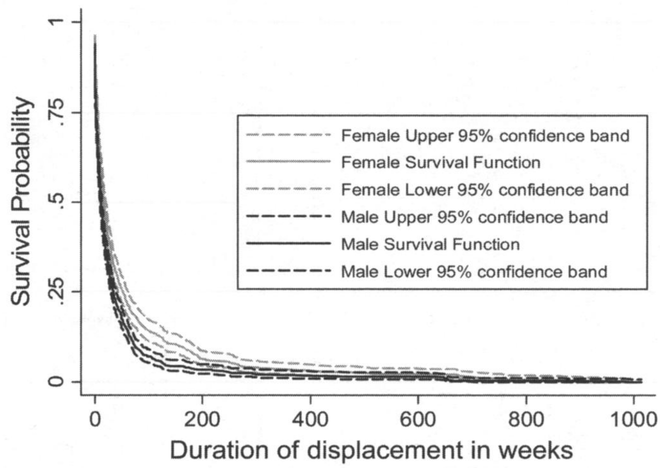

Figure 1. Survival Function of Length of Displacement by Gender

39 year olds, but that there is again a large difference in the functions for workers 40 years old or older. A test of the equality of the two functions in each figure rejects the hypothesis that the functions are equal for the younger than 24, the 25 to 29, and the 40 and over groups, but fails to reject the hypothesis for the two other age groups.

Next we present results from estimating a hazard model where the hazard is finding a job in period t. A hazard model seems like the obvious choice as a hazard model allows us to deal with censoring in a fairly straight forward fashion. The main issue when estimating a hazard model is specifying the distribution of the baseline hazard function. We have chosen to avoid this issue by estimating a Cox proportional hazard model which does not require us to specify the exact parametric form of the underlying hazard function, but does rely on the assumption that changes in the hazard function are proportional. Given the estimates of the survival function presented in Figures 1 and 2, this appears to be a reasonable assumption.[15]

The results from our estimations are presented in Table 3. In all our models, we control for age using the five age categories discussed previously.[16] We also include a quadratic in experience and tenure, years of schooling completed, a race dummy (white vs. non-white), a set of year dummies indicating the year a worker was displaced and controls for the worker's industry at displacement (nine categories). In this table, we present estimates of the hazard ratio which shows the proportional change in the hazard when the variable is increased by one unit. A ratio of less than one indicates the hazard declines as the variable increases while a hazard ratio of greater than one indicates a positive relationship between the hazard and the variable.

The results for our entire sample are presented in column 1. These results are fairly typical for the displaced worker literature (Carrington 1990; Fallick 1993, 1996; and Abbring et al. 2002). We

[15] We have estimated parametric hazard models assuming the underlying hazard function has a Weibull, log normal, and generalized gamma distribution, and the results from all of these models are similar. We have estimated parametric versions of the models controlling for unobserved heterogeneity. Both the likelihood ratio tests and graphical tests examining the fit of the models controlling and not controlling for unobserved heterogeneity show that the possible presence of unobserved heterogeneity is not affecting our estimates. Therefore, we focus on results from nonparametric models where we do not adjust for unobserved heterogeneity. These results are available from the authors on request.

[16] We have estimated all of our models using continuous measures of age and age squared and the results are nearly identical. We present the results using the age categories for illustrative purposes.

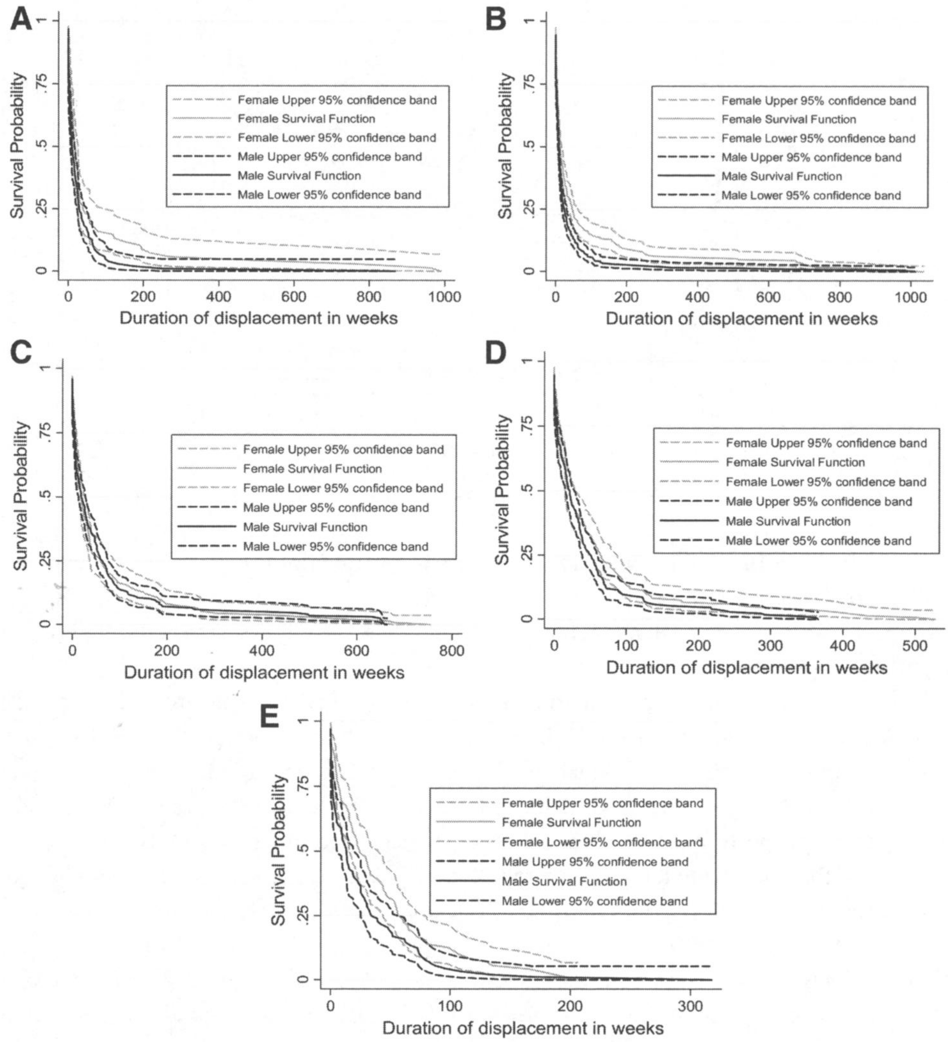

Figure 2. Survival Function of Length of Displacement by Age and Gender. Panel A: Under 24 years old, Panel B: Between 25 and 29 years old, Panel C: Between 30 and 34 years old, Panel D: Between 35 and 39 years old, and Panel E: 40 years old and older.

see that the length of displacement increases with age and tenure while it decreases with labor market experience. Also more educated workers, married workers and non-white workers have shorter spells of displacement.

In columns 2 and 3, we present results from samples of men and women, respectively. Comparing the results in these two columns we see that men and women have a similar estimated relationship between most of the control variables and the length of the displacement spell.[17] The one notable exception is the age variables. The hazard ratio estimates on the age categories show that

[17] Note that these results are conditional on industry in the predisplacement job. We have estimated the model on the pooled data including interactions between the female dummy variable and all of the other control variables. We fail to reject the hypothesis that the coefficients on the interaction between the female variable and all of the other variables, except the age variables, are zero.

Table 3. Proportional Hazard Estimation of the Length of Displacement

	Entire Sample (1)	Men (2)	Women (3)	Entire Sample (4)	Entire Sample (5)	Limited Sample (6)	Out of Labor Force Sample (7)	Limited Sample Out of Labor Force (8)
Age 25–29	0.89	0.87	0.96	0.88	0.87	0.87	1.18	1.21
	(−1.14)	(−1.02)	(−0.24)	(−1.05)	(−1.07)	(−1.08)	(0.66)	(0.75)
Age 30–34	0.75	0.59	1.04	0.63	0.63	0.61	0.82	0.88
	(−1.97)	(−2.55)	(0.18)	(−2.78)	(−2.66)	(−2.82)	(−0.60)	(−0.37)
Age 35–59	0.59	0.47	0.72	0.50	0.51	0.49	0.61	0.71
	(−2.64)	(−2.70)	(−1.10)	(−3.17)	(−3.06)	(−3.13)	(−1.11)	(−0.75)
Age 40+	0.54	0.62	0.47	0.62	0.65	0.67	0.084	1.19
	(−2.60)	(−1.42)	(−2.20)	(−1.80)	(−1.59)	(−1.48)	(−0.34)	−0.34
Female x								
Age 20–24	—	—	—	0.67	0.68	0.99	0.65	1.16
				(−2.82)	(−2.58)	(−0.07)	(−1.70)	(0.53)
Age 25–29	—	—	—	0.72	0.74	1.02	0.54	1.08
				(−3.14)	(−2.62)	(0.21)	(−2.87)	(0.35)
Age 30–34	—	—	—	1.02	1.05	1.18	1.25	1.64
				(0.20)	(0.39)	(1.35)	(0.91)	(1.99)
Age 35–39	—	—	—	1.02	1.04	1.18	1.19	1.74
				(0.11)	(0.25)	(1.18)	(0.61)	(1.91)
Age 40–45	—	—	—	0.53	0.52	0.54	0.73	0.71
				(−3.01)	(−3.16)	(−2.90)	(−0.77)	(−0.79)
Experience	1.002	1.003	1.003	1.002	1.002	1.003	1.001	1.002
	(4.44)	(3.34)	(3.22)	(4.46)	(4.43)	(4.64)	(1.71)	(2.48)
Experience Squared	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000
	(−2.38)	(−1.78)	(−1.97)	(−2.48)	(−2.50)	(−2.65)	(0.44)	(−0.25)
Tenure	0.9984	0.9988	0.9973	0.9984	0.9984	0.998	0.999	0.999
	(−3.4)	(−2.01)	(−3.25)	(−3.45)	(−3.46)	(−3.32)	(−1.22)	(−1.30)
Tenure Squared	1.000002	1.000002	1.000002	1.000002	1.000002	1.000002	1.000000	1.000000
	(2.69)	(2.26)	(2.05)	(2.80)	(2.79)	(2.65)	(−0.09)	(−0.03)
Non-white	1.05	1.10	0.94	1.03	1.04	1.03	0.995	0.95
	(0.82)	(1.30)	(−0.68)	(0.52)	(0.65)	(0.55)	(−0.04)	(−0.39)
Years of schooling completed	1.06	1.06	1.05	1.06	1.05	1.06	1.004	1.02
	(3.37)	(3.23)	(1.84)	(3.55)	(3.05)	(3.26)	(0.13)	(0.58)
Married	1.07	1.12	1.04	1.10	1.10	1.14	1.08	1.28
	(1.04)	(1.33)	(0.42)	(1.59)	(1.56)	(2.08)	(0.63)	(1.79)
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Occupation dummies	No	No	No	No	Yes	Yes	Yes	Yes
Number of observations	1322	753	569	1322	1321	1247	446	394
Log likelihood	−7735.06	−4056.17	−2820.93	−7721.55	−7710.66	−7267.52	−1989.97	−1738.92

Note: The table shows the hazard ratios along with *z*-scores in parentheses. All standard errors have been corrected for clustering. All variables are measured at the time of displacement. The limited sample in column 6 includes all men but only includes women who are living in households where the number of children did not increase between displacement and the post-displacement job. Column 7 reports results for the sample of workers who left the labor market at some point during their displacement spell, while column 8 reports results for the limited sample of workers who also left the labor market during their displacement spell.

for men the expected length of displacement is increasing with age until age 40. In contrast for women, the positive relationship between age and spell length only occurs in the 35–39 and 40 and older categories. These results suggest that we can pool the data on men and women when estimating the hazard model as long as we include an interaction between sex and age.

To see exactly how the length of the displacement spell varies by sex and age, in columns 4 and 5 we present results from models estimated on the entire sample including a complete interaction between a female dummy variable and the age categories. In column 5, we also include controls for a worker's occupation in the displacement job (nine categories). Focusing on the results in column 4, we see that women 24 years old or younger have a significantly longer expected spell length than men of similar age. The same is true for men and women 25 to 29, although the difference is smaller. Among workers age 30 to 35 and 35 to 39 there is no significant difference in the expected duration of nonemployment between men and women. Finally, among the oldest workers women again have a significantly longer displacement spell than men. We see the identical pattern in column 5 when we control for occupation. These results show that, while women have longer spells of displacement than men, this is due to differences in spell lengths among the youngest and oldest workers.[18]

Length of Displacement and Fertility

In the following, we investigate whether gender differences among the young are related to fertility. One possible explanation for the observed life-cycle patterns of relative spell length is that differences in search behavior are due to differences in nonmarket opportunities related to child bearing and raising. To examine this hypothesis, Appendix B presents age-specific fertility rates for all U.S. women for years corresponding to the time period of our data. This table shows that fertility rates initially rise as women age, reach a peak among 25- to 29-year-old women and then drop rather quickly after age 30. The ages in which women have longer displacement spells seem to correspond to the ages in which fertility rates are high and rising and the ages in which men's and women's length of search are similar corresponds to ages in which women's fertility rates are low and falling.

We see a similar fertility pattern among the women in our sample who have strong attachments to the labor market. Table 4 presents the number of children living in the household at displacement and the number of children living in the household at the start of the post-displacement job separately for women and men in our sample.[19] For censored spells, we have used the number of children reported living in the household in the 2002 survey. This table shows that among the youngest women in our sample there is a 50% increase in the number of children between the displacement and the post-displacement job. There is a 16% increase among women age 25 to 29, but among women 30 years old or older there is only a very small increase or a decrease in the number of children living in the household. When we focus on women who experience displacement spells

[18] To examine whether any of our results are being driven by differences in ability between men and women, for all of our basic regressions we have estimated two additional models. In the first, we have included a dummy variable indicating whether a worker was working part time at the time of displacement and a variable indicating the worker's AFQT score. In the second specification, we have estimated these models separately for workers with a high school degree or less and for workers with more than a high school degree. In all of these additional regressions, we see the same basic patterns in the coefficients on the female-age interactions. These results are available from the authors on request.

[19] As a measure of fertility, we use the change in the number of children in the household, that includes all children born to the mother and living in the household as well as the children of the partner and adopted children if they live in the household. We use this more expansive definition of fertility to capture any event related to raising children that could affect nonmarket value of time. For robustness checks, we constructed this measure both for women and men. The NLSY contains other quite detailed fertility information which is only available for women and would not always capture changes in household size.

Table 4. Number of Children Living in the Household with Displaced Workers

	Entire Sample			Displacement Spell 26 Weeks or Longer		
Age	Number of Children at Displacement (1)	Number of Children at Post-Displacement Job (2)	Proportional Change (3)	Number of Children at Displacement (4)	Number of Children at Post-Displacement Job (5)	Proportional Change (6)
Panel A—Women						
24 years old or younger	0.58	0.87	0.50	0.55	1.17	1.13
25–29 years old	0.98	1.14	0.16	1.03	1.42	0.38
30–34 years old	1.62	1.68	0.04	1.79	1.83	0.02
35–39 years old	1.80	1.70	−0.06	1.79	1.70	−0.05
40 years old or older	1.50	1.34	−0.11	1.30	1.16	−0.11
Panel B—Men						
24 years old or younger	0.28	0.41	0.46	0.30	0.35	0.17
25–29 years old	0.68	0.79	0.16	0.61	0.54	−0.11
30–34 years old	0.88	0.88	0.00	0.88	0.88	0.00
35–39 years old	1.26	1.27	0.01	1.42	1.38	−0.03
40 years old or older	1.18	1.14	−0.03	1.20	1.23	0.03

of 26 weeks or more the pattern is even more dramatic. Among the youngest age categories there is between a 38% to 113% increase in the number of children, while among women over 30 there is between a 2% increase and an 11% decrease in the number of children living in the household. In contrast, for men we see that there is very little difference in the relationship between changes in the number of children in the household and the length of the displacement spell.

To examine the hypothesis that gender differences in the length of displacement result from some women choosing to have children after being displaced, column 6 in Table 3 presents results from our hazard model estimated on the entire sample of male workers but only including women who have either no change or a decline in the number of children living in the household between the time of displacement and the start of the displacement job. We see that among this sample there is no significant difference in the estimated hazard ratio for any age category except for workers over age 40. The last four columns in Table 1 present summary statistics for women included in our limited sample (columns 7 and 8) and women excluded from our limited sample (columns 9 and 10). These data show that on average women who have children after being displaced have much longer spells of displacement and also are younger, have less labor market experience, have attended fewer years of school and have lower AFQT scores. These results taken together suggest that any difference between men and women in the length of search after displacement is due to younger, less skilled, women who are displaced choosing to have children prior to returning to the labor market.[20]

To further examine the question of whether our measure of fertility simply captures the effect of women being more likely to leave the labor market during displacement, column 7 in Table 3 presents estimation results for the sample of workers who leave the labor market at some point during their displacement spell, while column 8 contains results for our limited sample of workers who also leave the labor market during displacement. The results in column 7 are almost identical

[20] Estimating the hazard model on a sample that also excludes men who live in households where the number of children increases between the displacement and post-displacement job produces nearly identical results to those shown in column 4 of Table 3.

to the results for our full sample reported in column 4. In contrast, the results in column 8 show no significant gender difference in spell length for workers under age 30 and show that women between 30 and 39 years old have significantly shorter lengths of displacement than comparable men of the same age. The results in these two columns indicate that the life-cycle gender differences in spell length are coming from the longer spells of displacement experienced by women living in a household with an increase in the number of children during the displacement spell. The life cycle pattern is not simply the result of women being more likely to exit the labor market during displacement.

Tenure on the Post-Displacement Job

The Bowlus–Eckstein (2002) model predicts that if women experience discrimination they should have a shorter tenure on their post-displacement job than men. When we plot gender differences in the estimated survival function for remaining in the post-displacement job, we find no significant gender differences at the mean or within any of the five age groups.

In Table 5, we present estimation results from a proportional hazard model of the length of the post-displacement job. We include the same set of controls in these models as we did in the previous models except we have dropped the tenure variables from the model and we measure marital status at the start of the post-displacement job.[21] In these results, as the hazard is the probability that a worker stops working at the post-displacement job in period t, a hazard ratio below one indicates a lower probability of failure and, therefore, a longer tenure on the post-displacement job. Looking at the results for the entire sample, we see that more experienced workers and married workers experience longer tenure on the post-displacement job. The estimated effects on years of schooling and non-white are insignificant. Looking at the age effects we see that tenure on the post-displacement job is increasing with age, although only the estimates for the two oldest categories are significant at standard levels.

Comparing the results in column 2 for men with the results in column 3 for women, we see that the estimated effects are fairly similar.[22] In particular, for both men and women, duration on the job is an increasing function of age.

In columns 4 and 5, we present results from our estimation where we include interactions between the female dummy variable and the age categories. In neither of these regressions do we find any significant difference between men and women in the estimated tenure on the post-displacement job. Furthermore, it is only among the younger workers where the point estimate suggests that women have a shorter tenure than men, but these estimates are imprecisely estimated. The overall conclusion from the analyses in this section is that there does not appear to be any significant difference between men and women in the length of the post-displacement job.

Wages on Post-Displacement Job

In this section, we examine worker wages on the post-displacement job. Table 6 presents results from standard OLS wage regressions where the log of the hourly wage at the start of the

[21] We have estimated a version of the model including tenure in the previous job and its square and these variables were insignificant so we dropped them from the final model.

[22] Estimating the model on the pooled data and including a complete interaction between the female dummy variable and all of the other control variables shows that the only interactions that are significantly different from zero are the female-experienced squared and female-marital status interactions.

Table 5. Proportional Hazard Estimation of the Length of Post Displacement Job

	Entire Sample (1)	Men (2)	Women (3)	Entire Sample (4)	Entire Sample (5)
Age 25–29	0.85	0.93	0.76	0.93	0.94
	(−1.47)	(−0.52)	(−1.74)	(−0.51)	(−0.47)
Age 30–34	0.77	0.67	0.79	0.82	0.80
	(−1.54)	(−1.69)	(−0.93)	(−1.07)	(−1.16)
Age 35–59	0.70	0.70	0.62	0.76	0.75
	(−1.60)	(−1.16)	(−1.44)	(−1.12)	(−1.20)
Age 40+	0.61	0.65	0.45	0.64	0.60
	(−1.80)	(−1.21)	(−1.77)	(−1.50)	(−1.71)
Female x					
Age 20–24	—	—	—	1.17	1.22
				(1.09)	(1.42)
Age 25–29	—	—	—	0.94	0.96
				(−0.51)	(−0.35)
Age 30–34	—	—	—	1.00	1.01
				(0.01)	(0.06)
Age 35–39	—	—	—	0.95	0.96
				(−0.29)	(−0.25)
Age 40–45	—	—	—	1.04	1.17
				(0.13)	(0.55)
Experience	0.999	0.999	0.998	0.999	0.999
	(−1.54)	(−0.69)	(−2.32)	(−1.51)	(−1.71)
Experience squared	1.000001	1.000000	1.000002	1.000001	1.000001
	(1.03)	(−0.18)	(2.41)	(1.00)	(1.27)
Non-white	1.06	1.06	1.09	1.07	1.06
	(0.83)	(0.61)	(0.73)	(0.86)	(0.75)
Years of schooling completed	0.98	0.96	0.99	0.98	0.99
	(−1.31)	(−1.69)	(−0.22)	(−1.30)	(−0.34)
Married	0.82	0.73	1.04	0.82	0.82
	(−2.75)	(−3.33)	(0.32)	(−2.75)	(−2.69)
Year dummies	Yes	Yes	Yes	Yes	Yes
Industry dummies	Yes	Yes	Yes	Yes	Yes
Occupation dummies	No	No	No	No	Yes
Number of observations	1135	648	487	1135	1135
Log likelihood	−5577.02	−2913.61	−2031.63	−5576.27	−5568.55

Note: The table shows the hazard ratios along with z-scores in parentheses. All standard errors have been corrected for clustering.

post-displacement job is the dependent variable. In all of these regressions, we have included the same controls that we included in our previous regressions except for tenure in the previous job. In column 1, we present results where we have included a female dummy variable. The coefficient on this variable indicates that female workers earn 16% lower wages on average than similar male workers.

In columns 2 and 3, we present results where we have interacted the female dummy variable with the age dummies. In column 3, we drop the controls for occupation and industry as there is some question whether these variables are exogenous and should be included in the regression. As the basic pattern of the results is similar in the two columns we will focus on the results in column 2.

The coefficients on the age-female interaction terms follow a pattern that is strikingly similar to the pattern seen in our hazard estimation of the length of displacement. The youngest female

Table 6. Regression of the Log Hourly Wage in the Post-Displacement Job

	Entire Sample (1)	Entire Sample (2)	Entire Sample (3)	Limited Sample (4)
Age 25–29	0.04	0.07	0.06	0.09
	(0.05)	(0.06)	(0.06)	(0.06)
Age 30–34	0.03	0.00	−0.02	−0.05
	(0.09)	(0.10)	(0.10)	(0.09)
Age 35–59	0.18	0.13	0.11	−0.06
	(0.12)	(0.13)	(0.12)	(0.13)
Age 40+	0.20	0.25	0.20	0.01
	(0.14)	(0.14)	(0.14)	(0.14)
Female	−0.16	—	—	—
	(0.03)			
Female x				
Age 20–24	—	−0.16	−0.20	−0.06
		(0.07)	(0.07)	(0.08)
Age 25–29	—	−0.22	−0.27	−0.20
		(0.06)	(0.06)	(0.06)
Age 30–34	—	−0.10	−0.15	−0.09
		(0.07)	(0.07)	(0.07)
Age 35–39	—	−0.09	−0.12	−0.11
		(0.06)	(0.06)	(0.07)
Age 40–45	—	−0.32	−0.32	−0.33
		(0.10)	(0.10)	(0.11)
Experience × 1000	0.07	0.11	0.14	0.312
	(0.31)	(0.31)	(0.32)	(0.27)
Experience squared × 1000	0.0005	0.0004	0.0005	0.0003
	(0.0003)	(0.0003)	(0.0003)	(0.0002)
Years of schooling completed	0.06	0.06	0.08	0.06
	(0.01)	(0.01)	(0.01)	(0.01)
Married	0.09	0.10	0.12	0.09
	(0.03)	(0.03)	(0.03)	(0.04)
Non-white	−0.02	−0.02	−0.04	0.004
	(0.03)	(0.03)	(0.03)	(0.03)
Year dummies	Yes	Yes	Yes	Yes
Industry dummies	Yes	Yes	No	Yes
Occupation dummies	Yes	Yes	No	Yes
Number of observations	1119	1119	1119	834
R-Square	0.34	0.34	0.27	0.39

Note: The table shows the OLS regression coefficients along with standard errors in parentheses. All standard errors have been corrected for clustering. All variables are measured at the time of displacement. The limited sample in column 4 only includes workers who find a job within 32 weeks of being displaced and only includes women living in households where the number of children did not increase between displacement and the post-displacement job.

workers in our sample have wages that are 16% less than the youngest males in the sample. This difference grows to 22% in the 25–29 age categories. The wage differential between men and women then falls dramatically in the 30–34 and 35–39 age categories and the difference is insignificant for both of these groups. Finally, the wage differential is largest among the oldest workers in the sample.

To examine the role that displacement length and fertility have on post-displacement wages, column 4 contains results from a wage regression where we only include workers who obtain a job within 32 weeks of being displaced and only include women living in households where the

Table 7. Regression of the Change in Log Hourly Wage between the Displacement Job and the Post-Displacement Job

	Entire Sample (1)	Entire Sample (2)	Entire Sample (3)	Limited Sample (4)
Age 25–29	0.09	0.08	0.08	0.07
	(0.05)	(0.06)	(0.06)	(0.06)
Age 30–34	0.10	0.10	0.10	0.07
	(0.08)	(0.09)	(0.09)	(0.09)
Age 35–59	0.16	0.18	0.17	0.17
	(0.09)	(0.10)	(0.10)	(0.10)
Age 40+	0.18	0.22	0.22	0.202
	(0.10)	(0.11)	(0.10)	−0.11
Female	−0.01	—	—	—
	(0.03)			
Female x				
Age 20–24	—	0.01	0.01	0.09
		(0.07)	(0.07)	(0.07)
Age 25–29	—	0.03	0.04	0.06
		(0.05)	(0.05)	(0.05)
Age 30–34	—	−0.01	−0.004	−0.003
		(0.06)	(0.06)	(0.06)
Age 35–39	—	−0.03	−0.03	−0.04
		(0.05)	(0.05)	(0.06)
Age 40–45	—	−0.10	−0.10	−0.06
		(0.07)	(0.07)	(0.07)
Exerience × 1000	−0.56	0.54	−0.55	−0.58
	(0.27)	(0.28)	(0.27)	(0.27)
Experience Squared × 1000	0.0004	0.0004	0.0004	0.0004
	(0.0002)	(0.0002)	(0.0003)	(0.0002)
Tenure × 1000	−0.53	−0.54	−0.53	−0.33
	(0.22)	−0.2200	−0.2200	(0.24)
Tenure squared × 1000	0.0005	0.0005	0.0005	0.0004
	(0.0003)	(0.0003)	(0.0003)	(0.0003)
Years of schooling completed	0.001	0.0004	0.0003	0.0041
	(0.01)	(0.01)	(0.01)	(0.006)
Married	0.01	0.01	0.01	0.0055
	(0.03)	(0.01)	(0.03)	(0.03)
Non-white	0.04	0.04	0.04	0.05
	(0.03)	(0.03)	(0.03)	(0.03)
Year dummies	Yes	Yes	Yes	Yes
Change industry dunmmy	Yes	Yes	No	Yes
Change occupation dummy	Yes	Yes	No	Yes
Number of observations	1203	1203	1203	817
R-Square	0.08	0.08	0.08	0.07

Note: Table shows the OLS regression coefficients along with standard errors in parentheses. All standard errors have been corrected for clustering.

number of children does not increase between the displacement and post-displacement job. Changing the sample in this way primarily affects the estimated male–female differential among the young workers in our sample. The estimated differential among the youngest workers falls from 16% to 6% and becomes statistically insignificant. Among 25- to 29-year-old workers, the differential falls from 22% to 20%. Among the older workers, there is almost no change in the estimated wage differential. While there are still significant differences in wages between men and women, these results suggest that differential search behavior due to fertility decisions on the part of women accounts for at least some of the observed male/female wage differential.

If it is the case that a higher value of nonmarket time accounts for the observed differences in search between young men and women then presumably young women who chose to remain in the labor market and not have children after being displaced should experience a relative increase in their wage after displacement compared to similar men. To examine this hypothesis in Table 7, we present regression results that parallel the regressions presented in Table 6 except the dependent variable is now the change in the log hourly wage between the displacement and post-displacement job. The coefficient on the female dummy variable in column 1 shows that there is no significant gender difference in the change in wages after displacement among all workers. The coefficients on the age × female interaction terms in columns 2–4 do show the expected pattern—younger women who remain in the market experience wage increases relative to men of similar ages. In addition, the wage gains are even larger when we limit our sample to workers who find a job within 32 weeks and to women who do not have a child after becoming displaced. Of course, none of these coefficients are significant, so this is only weak evidence in support of the theory that it is life-cycle changes in nonmarket opportunities that is driving the male/female differences in search.

5. Conclusion

In this article, we have examined how job search behavior varies between men and women over the early part of the life-cycle and whether fertility decisions can account for some of the observed differences in search. We find that women experience longer spells of displacement than men, but that these differences are primarily the result of differences in job search between men and women during prime childbearing years for women (29 years or younger), and in the age group 40 to 45 years. However, once we restrict our sample to women who do not experience an increase in the number of children in the household prior to finding a post-displacement job there is no significant difference in the length of displacement between men and women younger than 40. We find no evidence that women have a longer tenure on their post-displacement job at any age. We find that women's wages in the post-displacement job are lower than the wages for similar men. However, once we focus on men and women who find a job within 32 weeks of being displaced and focus on women who live in households where the number of children does not increase during the displacement spell, we find smaller differences in wages between young men and women. We find no significant differences between men and women in post-displacement changes in wages.

These results are consistent with our findings in our previous paper exploiting worker data from Germany (Kunze and Troske 2012). Both studies show that job search behavior varies between men and women over the life-cycle. This similarity in findings is somewhat surprising given that the United States and Germany have considerably different employment protection laws and maternity/parental leave policies. Our findings are also consistent with the findings in Del Bono, Weber, and Winter-Ebmer (2012) who show that job displacement reduces the fertility of relatively high skilled women but has very little impact on the fertility of low skilled workers. They find that more skilled female workers postpone their fertility to more quickly move into a new job, which is consistent with our findings that more skilled women transit into a new job after displacement in a manner similar to male workers.

One interpretation of our results is that fertility is not exogenous to displacement—once they become displaced some women choose to have children instead of returning to the labor market. An alternative explanation is that younger women who suffer the greatest taste-based discrimination in the market are the ones who choose to have a child after being displaced from their previous job. However, for this model to match our main findings the discrimination needs to have a larger impact on women who would have experienced the longest spell of displacement even if they did not have a child. This is because men and women experience similar spells of displacement once we focus on women who do not have children after being displaced. It is also the case that the women who have children after being displaced tend to have less education, less experience and lower levels of labor market skill, which is consistent with them having the lowest opportunity cost of having children.

It is also possible that our results can be explained by statistical discrimination on the part of employers; that is, employers may be less likely to offer a job to women of prime child bearing age because they believe that some women will become pregnant in the future and either quit the firm or provide less effort when employed. Again, however, given that we only see longer spells of displacement among women who actually have children, this statistical discrimination must be focused on a subset of women within age categories who would have experienced the longest spells of displacement in the absence of discrimination.

We do not have a good explanation of the gender differences in displacement durations among 40 to 45 year olds. Note that we find no evidence of gender differences in wage changes after displacement for older workers. One possible explanation is that after age 40 women who are displaced are more likely than similar aged men to permanently leave the labor market. Unfortunately, we are unable to examine this possibility with our data, so this remains a topic for future research.

Appendix A: Supplementary Summary Statistics, across Years

	Distribution of Displacement by Year—Sample of Displaced Workers		Proportion Married in the NLSY		Proportion Full-Time in the NLSY	
Year	Men (1)	Women (2)	Men (3)	Women (4)	Men (5)	Women (6)
1983	5.4%	4.8%	24.5%	35.6%	43.6%	35.0%
1984	6.5%	7.9%	28.5%	39.5%	50.4%	40.5%
1985	7.0%	8.1%	30.1%	41.6%	58.9%	46.0%
1986	7.2%	7.2%	35.2%	45.7%	65.8%	48.6%
1987	6.6%	5.1%	39.7%	49.1%	71.5%	54.2%
1988	5.1%	5.6%	44.3%	52.0%	74.0%	56.9%
1989	6.6%	6.7%	47.1%	53.7%	76.5%	58.7%
1990	4.8%	5.6%	49.3%	55.0%	78.1%	58.9%
1991	4.4%	4.6%	50.6%	54.6%	78.8%	60.3%
1992	4.1%	3.9%	52.0%	55.3%	77.4%	60.7%
1993	3.6%	1.2%	53.1%	56.1%	79.7%	63.0%
1994	4.4%	3.0%	53.9%	56.7%	80.5%	62.7%
1995	4.8%	3.9%				
1996	3.9%	5.3%	55.1%	57.4%	83.6%	65.3%
1997	5.3%	5.3%				
1998	5.3%	4.4%	57.2%	58.6%	86.3%	66.7%

Appendix A. (*Continued*)

	Distribution of Displacement by Year—Sample of Displaced Workers		Proportion Married in the NLSY		Proportion Full-Time in the NLSY	
Year	Men (1)	Women (2)	Men (3)	Women (4)	Men (5)	Women (6)
1999	3.6%	3.9%				
2000	4.3%	5.5%	58.0%	58.2%	88.6%	70.4%
2001	5.2%	5.1%				
2002	2.0%	3.2%	59.2%	58.1%	87.5%	69.6%

Note: The original NLSY contains 12,686 individuals, the Sample of Displaced Workers contains 1322 spells.

Appendix B: Age Specific Fertility Rates for Selected Years (Live Births per 1000 Women)

	15–19 (1)	20–24 (2)	25–29 (3)	30–34 (4)	35–39 (5)	40–44 (6)
1980	53.0	115.1	112.9	61.9	19.8	3.9
1985	51.0	108.3	111.0	69.1	24.0	4.0
1990	59.9	116.5	120.2	80.8	31.7	5.5
1995	56.0	107.5	108.8	81.1	34.0	6.6
2000	47.7	109.7	113.5	91.2	39.7	8.0

Source: Nation Center for Health Statistics, http://www.cdc.gov/nchs/data/statab/t001x07.pdf

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References

Abbring, Jaap, Gerard J. van den Berg, Pieter A. Gautier, A. Gijsbert C. van Lomwel, Jan C. van Ours, and Christopher J. Ruhm. 2002. Displaced workers in the United States and the Netherlands. In *Losing work, moving on: International perspectives on worker displacement*, edited by Peter J. Kuhn. Kalamazoo, MI: W.E. Upjohn Institute for Employment Research, pp. 105–94.

Altonji, Joseph G., and Rebecca M. Blank. 1999. Race and gender in the labor market. In *Handbook of labor economics 3c*, edited by Orley Ashenfelter and David Card. New York, NY: Elsevier, pp. 3143–260.

Becker, Gary S. 1971. *The economics of discrimination*. Chicago, IL: The University of Chicago Press.

Becker, Gary S. 1981. *A treatise on the family*. Cambridge, MA: Harvard University Press.

Black, Dan A. 1995. Discrimination in an equilibrium search model. *Journal of Labor Economics* 13:309–34.

Black, Sandra, and Phillip Strahan. 2001. The division of spoils: Rent-sharing and discrimination in a regulated industry. *American Economics Review* 91:814–31.

Bowlus, A., and Zvi Eckstein. 2002. Discrimination and skill differences in an equilibrium search model. *International Economic Review* 43:1309–45.

Bowlus, Audra, and Lars Vilhuber. 2002. Displaced workers, early leavers, and re-employment wages. *Longitudinal Employer-Household Dynamics Technical Papers 2002-18*. Washington, DC: Center for Economic Studies, U.S. Census Bureau.

Burdett, Kenneth, and Dale Mortensen. 1998. Wage differentials, employer size and unemployment. *International Economic Review* 39:257–73.

Carrington, William J. 1990. *Specific human capital and worker displacement*. Ph.D., Dissertation, University of Chicago, Chicago, IL.

Chesher, Andrew, M. G. B. Dumangane, and R. J. Smith. 2002. Duration response measurement error. *Journal of Econometrics* 111:169–94.

Crossley, Thomas F., Stephen R. G. Jones, and Peter Kuhn. 1994. Gender differences in displacement cost: Evidence and implications. *Journal of Human Resources* 29:461–80.

Del Bono, Emilia, Andrea Weber, and Rudolf Winter-Ebmer. 2012. Clash of career and family: Fertility decisions after job displacement. *Journal of the European Economic Association* 10:659–83.

Ederington, Josh, Jenny Minier, and Kenneth R. Troske. 2009. *Where the girls are: Trade and labor market segregation in Columbia. Discussion Paper #4131*. Bonn, Gemany: IZA.

Fallick, Bruce C. 1993. The industrial mobility of displaced workers. *Journal of Labor Economics* 11:302–23.

Fallick, Bruce C. 1996. A review of the recent empirical literature on displaced workers. *Industrial and Labor Relations Review* 50:5–16.

Farber, Henry S. 1997. The changing face of job loss in the United States, 1981-1995. In *Brookings papers on economic activity: Microeconomics*, edited by Henry S. Farber, John Haltiwanger, and Katharine G. Abraham. Washington, DC: Brooking Institution, pp. 55–142.

Hellerstein, Judith, David Neumark, and Kenneth R. Troske. 2002. Market forces and sex discrimination. *Journal of Human Resources* 37:353–80.

Hu, Luojia, and Christopher Taber. 2011. Displacement, asymmetric information and heterogeneous human capital. *Journal of Labor Economics* 29:113–52.

Jacobson, Louis, Robert LaLonde, and Daniel G. Sullivan. 1993. Earnings losses of displaced workers. *American Economics Review* 83:685–709.

Kletzer, Lori, and Robert W. Fairlie. 2003. The long-term costs of job displacement for young adult workers. *Industrial and Labor Relations Review* 56:682–98.

Kunze, Astrid, and Kenneth R. Troske. 2012. Life-cycle patterns in male/female differences in job search. *Labour Economics* 19:176–85.

Lengermann, Paul, and Lars Vilhuber 2002. *Abandoning the sinking ship: The composition of worker flows prior to displacement. Longitudinal Employer-Household Dynamics Technical Papers 2002-11*. Washington, DC: Center for Economic Studies, U.S. Census Bureau.

Madden, Janice. 1987. Gender differences in the cost of displacement: An empirical test of discrimination in the labor market. *American Economic Review* 77:246–51.

Mortensen, Dale T. 1986. Job search and labor market analysis. In *Handbook of Labor Economics 2*, edited by Orley Ashenfelter and David Card. New York, NY: Elsevier: 849–920.

Mortensen, Dale T., and Christopher A. Pissarides. 1999. New developments in models of search in the labor market. In *Handbook of Labor Economics 3b*, edited by Orley Ashenfelter and David Card. New York, NY: Elsevier, pp. 2567–628.

Podgursky, Michael, and Paul Swaim. 1987. Job displacement and earnings loss: Evidence from displaced worker survey. *Industrial and Labor Relations Review* 41:17–29.