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Life-cycle patterns in male/female differences in job search

Astrid Kunze [a,*], Kenneth R. Troske [b]

[a] NHH Norwegian School of Economics, Norway
[b] University of Kentucky & IZA, Bonn, Germany

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ABSTRACT

We investigate whether women search longer for a job than men and whether these differences change over the life cycle. Our empirical analysis exploits German register data on highly attached displaced workers. We apply duration models to analyze gender differences in job search taking into account observed and unobserved worker heterogeneity and censoring. Simple survival functions show that displaced women take longer to find a new job than comparable men. Disaggregation by age groups reveals that these differences are driven by differential behavior of women in their prime-childbearing years. There is no significant difference in job search duration among the very young and older workers. These differential outcomes remain even after we control for differences in human capital and when unobserved heterogeneity is incorporated into the model.

1. Introduction

One of the persistent questions in economics is whether the differences in wages between men and women reflect observed differences in productivity, unobserved differences in productivity, or discrimination. One possible difference between men and women that has received relatively little attention is differences in job search and job mobility. Previous work has found that mobility among young workers is an important source of wage growth (Topel and Ward, 1992; von Wachter and Bender, 2006); however, evidence for the U.S. and Germany suggests that young women change jobs less often than men and experience smaller gains in wages when they do switch jobs (Loprest, 1992; Fitzenberger and Kunze, 2005). Unfortunately, these findings are difficult to interpret because job movers are a select sample of workers, where the selection is often based on worker characteristics that are unobservable to the econometrician but are correlated with outcomes (for a discussion see e.g. von Wachter and Bender, 2006).

In order to address some of the limitation with the previous research, in this paper we examine gender differences in the duration of job search and subsequent wages focusing on workers searching for a job following displacement due to a plant closing. Ideally, in order to examine gender differences in job search, we would like to have data on a random sample of workers who unexpectedly lost their job. Assuming that a plant closing is independent of the behavior of workers; then our data will come closer to the ideal data than data including workers who chose to switch jobs.[1]

As an extension to the previous literature we also will examine whether gender differences in search vary over the life-cycle. While there has been some theoretical work predicting differences in search behavior between men and women related to productivity differences or discrimination (e.g. Black, 1995; Bowlus and Eckstein, 2002), little attention has been paid to life-cycle variation in the search behavior of men and women. From the limited empirical evidence so far on gender differences in job search it is not clear whether differences exist across age groups. One intuitive reason why one may expect gender differences to vary with age is related to the comparative advantage of child bearing of women, which may generate life-cycle patterns in gender differences in job search.[2]

In our empirical analysis, we exploit administrative panel data drawn from the German social security insurance program. We follow displaced workers until they either obtain a new job or our data end. The data cover the period from 1975 through 2001. Our use of longi-

* Corresponding author at: NHH Norwegian School of Economics, Helleveien 30, N-5045 Bergen.
E-mail address: astrid.kunze@nhh.no (A. Kunze).

[1] Factors that could invalidate this assumption include workers receiving advanced warning of a plant closure. We discuss this issue more thoroughly below.
[2] Following theory a usual assumption is that women have a comparative advantage in child bearing that is constant. Empirically the advantage varies with age.

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tudinal administrative data ensures that we have an accurate measure of the length of displacement for all workers. In addition, since we use administrative panel data where spell length is measured directly from the receipt of unemployment benefits we avoid some of the problems with previous studies that have relied on cross-sectional data — e.g. not knowing the length of time a worker searches, or having search length self-reported by the worker several periods after the time of displacement. Since our data contain a large sample of workers age 20–60, we are able to examine how gender differences in search vary over the life cycle.

By applying duration models to analyze gender differences in job search taking into account observed and unobserved worker heterogeneity and censoring this study contributes new evidence on displaced workers in a European country to a literature that has been primarily shaped by studies on male displaced workers in the U.S.[3] In addition, as far as we are aware, ours is the only study to examine the job search behavior of European women who have been displaced and to compare the behavior of men and women using European data.[4] Finally, this study contributes to our understanding of the role that job mobility plays in producing the observed gender differences in labor market outcomes.

Our empirical results show that women do experience longer spells of displacement and a larger drop in wages after displacement than men. However, when we examine these differences over the life cycle we find that the differences in job search are concentrated among workers age 24 to 35, which are prime child bearing and child rearing ages for women. Among younger and older workers we find that men and women exhibit similar lengths of displacement and similar changes in wages. While not conclusive, these results do suggest that differences in job search and mobility are related to fertility decision. More particularly, a plausible interpretation is that the presence of young children reduces displaced women's job search intensity, though our data do not allow us to measure either search intensity or the presence of children directly. Alternatively, it might be that the presence of young children causes women to increase their reservation wages.

The remainder of the paper is as follows. In the next section we review the related work on displacement and job search. In Section 3 we describe our data and present some summary statistics. In Section 4 we present our empirical results on displacement durations and in Section 5 on wages. In Section 6 we discuss our results and present our conclusion.

2. Previous work examining male–female differences in job search

The basic theoretical arguments that have been offered to explain why women may search longer for a new job and receive lower wages operate through two main channels: differences in productivity and employer discrimination. One example is Bowlus (1997) who estimates a search model that allows for productivity differences between men and women, but does not explicitly allow for discriminatory behavior by firms. Another example is Black (1995) who shows that, if there exists discriminatory employers in the market, women will receive lower wages than men but the impact on the duration of search is ambiguous. Finally, the equilibrium search model of Bowlus and Eckstein (2002) allows for both productivity differences between men and women as well as discriminatory employers. In this model firms search over male and female workers but some fraction of employers are prejudiced against women; that is some firms experience a loss in utility from hiring women. Additionally, it is assumed that firms search less intensively for workers if they are less productive, and prejudiced firms also search less intensively for women even if they are as productive as men.[5] Under these conditions Bowlus and Eckstein show that women will be unemployed longer than men, even if they are as productive as men. They also show that wages will be lower for women because, in the presence of some prejudiced firms, in equilibrium all firms can exert monopsony power and hence offer all women relatively lower wages.[6]

Most previous empirical studies of displaced workers have focused on men or have pooled data for men and women. Simple comparisons of mean durations of displacement suggest that women take longer than men to find a new job after displacement (Podgursky and Swaim, 1987; Farber, 1997; Abbring et al., 2002; Kletzer and Fairlie, 2003; Hu and Taber, 2008). Hu and Taber (2011) find that women are displaced for a longer period than men after plant closure. With the exception of Hu and Taber (2008), none of these studies has analyzed in detail the gender differences in displacement durations.[7] One limitation of the previous empirical work which focuses on mean differences is that simple comparisons of mean duration among displaced workers can be misleading because durations are subject to censoring and are affected by worker heterogeneity.

The few studies that have examined gender differences in post-displacement wage outcomes have found mixed results. Early studies found that women experience larger wage losses after displacement (See Madden, 1987; Jacobson, et al., 1993; Crossley, et al., 1994) while later studies have found the opposite result (Kletzer and Fairlie, 2003). Hu and Taber (2011) find similar wage loss for men and women after plant closure. Further, there is no agreement on the mechanism that generates differential outcomes.[8] From this literature little is known about whether job search processes are different between men and women.[9]

Previous research on displacement that has compared data from North America with data from Europe has found striking cross-country differences.[10] North American studies find that displaced workers tend to experience large and fairly persistent wage losses after displacement. In contrast, European studies find relatively small declines in wages and that workers transit relatively quickly to a new job. However, to the best of our knowledge, no study compares the experience of European men and women who have been displaced.

3. The data

Our data on displaced workers come from the Institut fűr Arbeitsmarkt und Berufsforschung Sample (IABS) data for the period 1975–2001. We focus on West-Germany. The IABS is a two percent random sample of individuals drawn from the administrative data for the social security insurance program in Germany. The complete social security data are maintained by the German Federal Bureau of Labor and contain information for all workers who have at least one employment spell that is covered by the German social security system, which is approximately 80% of all workers in Germany in this

[3] See the articles in Kuhn (2002) for some exceptions to this statement.

[4] For a cross-gender comparison of the search behavior based on U.S. data see our companion paper Kunze and Troske (2009).

[5] Search intensity is exogenous.

[6] There are several papers that incorporate taste based employer discrimination into an equilibrium search model, e.g. Rosén, 2003 and Flabbi, 2010. Also Sasaki (1999) models job search in a world with co-workers discrimination. However, none of these papers models search as a function of worker demographics, such as age, which is one of the focuses of our empirical analyses.

[7] Azmat, et al. (2006) have looked at cross-country variation in gender gap in unemployment rates. Their conclusion is that differential outcomes are mainly driven by gender differences in human capital. Swaim and Podgursky (1994) have analyzed female labor supply employing a duration model.

[8] These conflicting results are somewhat puzzling because all of the studies use data for the U.S. with exception of one study which is based on Canadian data. All studies for the U.S. rely on the Displacement Worker Sample (DWS) supplements to the Current Population Survey (CPS) with the exception for Kletzer and Fairlie (2003) using the National Longitudinal Survey of Youth (NLSY), 1984–1993, and Jacobson et al. (1993) using employer-employee matched data for Pennsylvania.

[9] Crossley, et al. (1994) have suggested that gender differences in job search are important but have not empirically examined whether such difference exist. Hu and Taber (2011) focus on a model with asymmetric information and heterogeneous human capital.

[10] See the articles in Kuhn (2002).

period. Workers who are not included in these data are civil servants, self-employed workers and unpaid family workers.[11] In order to maintain the representativeness of the IABS data, workers who retire and/or leave the labor market are replaced by workers who enter the labor market.[12]

A major advantage of the administrative data is that the daily wage information reported in the data set is based on taxable income which makes this information highly reliable. Since we only have information on hours worked grouped into three categories, those working less than 17.5 h, between 17.5 and 35 h, and more than 35 h per week, we cannot calculate hourly wages. Wage information in these data are top coded with the top code values changing over time.[13] Finally, we drop all observations where a worker's wage is below the minimum social security level.

The other major strength of these data is that they include very detailed employment history information. These data contain the exact date of any change in a worker's labor market status. This includes any switch between full and part-time work, any interruption in work, any movement to unemployment, and any change in employer. A worker is reported as unemployed if he or she is in registered unemployment and is receiving unemployment insurance or unemployment assistance payments. Interruptions are reported by employers when the employer-employee relationship is on hold, but the contract is still valid. In this case no wage payments are made. Every other status results in a gap in an individual's record.

Individual records in the IABS are organized in spells by calendar date. In addition to any change in labor market status, establishments have to report information about each employee by the first of January each year, so individual records contain at least one spell per year if a worker is employed at some point in the year.

3.1. Definition and identification of closing establishments and displaced workers

We start the selection of the analysis sample by identifying displaced workers in the IABS data through establishment closures. Every June, the unique identifier for the establishment where a worker is currently employed, along with data on the total number of employees in the establishment, is added to a worker's record. An establishment is considered closed when an identifier appears in one year, but does not appear in the subsequent year. All workers who are identified as having worked in an establishment in the year prior to the establishment disappearing from the data are considered displaced.

In order to minimize the pre-displacement difference between men and women we focus on highly attached workers who transition out of employment into unemployment. As we will explain further below, given the German unemployment benefit system, we argue that these workers are all searching for a job after displacement. To ensure that we focus on workers with strong labor market attachments, we only keep workers who were between 20 and 60 years old at the time of displacement, have completed their education, and who were displaced from a job where they worked at least 17.5 h a week.[14] We also drop workers who switch jobs without experiencing a spell of unemployment. In our data 41% of men and 43% of women experience no unemployment following displacement.[15] One possibility is that these job-to-job transitions are the result of the reorganization of firms (changes in ownership or mergers), which is why we exclude these workers from our sample. Another possibility is that requirements that owners provide workers advanced notice prior to closing a plant enable some workers to find another job prior to their plant closing. If more able workers are better able to take advantage of the advanced warning, and if this selection varies by gender, then men and women in our sample may differ in ability, which in turn means that the observed cross-gender differences in search may be the result of unobserved cross-gender differences in ability and not a function of gender.

To further ensure that we are identifying true establishment closures we focus on workers who work in establishments that have at least six workers in the last year they appear in the data. This rule leads to the exclusion of 55% of person-plant spells for men and 65% for women. As has been shown previously, including establishments with fewer than six workers tends to overstate the number of closures (see Eliason and Storrie, 2006). This is because small establishments occasionally change identifiers when they change owners or for other unspecified reasons. In addition, we exclude workers in the construction and retail sectors, since these sectors have a large amount of seasonal variation in the fraction of new and closing establishments.

For each worker we keep up to four different displacement events. Around 20 (15) percent of male (female) workers in our sample were displaced more than once.[16] In the following analyses we adjust all standard errors to reflect the fact that the same worker can appear multiple times in the data.

Our measure of the length of displacement is the duration of registered unemployment from the end of the displacement job until a worker finds a new job, or the data end (censoring). In Germany, workers are part of registered unemployment whenever they are receiving unemployment compensation. Unemployment compensation consists of two parts: unemployment insurance (Arbeitslosengeld) and unemployment assistance (Arbeitslosenhilfe).[17] The length of time that a worker is eligible to receive unemployment insurance is a function of their age and the length of their previous job spell.[18] Typically, the length of time workers under the age of 42 can collect unemployment insurance is capped at one year. Once a worker exhausts their eligibility for unemployment insurance they are eligible for means-tested unemployment assistance. In principle, workers can receive unemployment assistance indefinitely. In the German unemployment compensation system workers are required to be searching for a job in order to receive unemployment compensation. By focusing on displaced workers who are reported as unemployed in the data we are able to measure the duration of displacement as the time workers are actively searching for a job.[19]

11 For more details, see Bender et al. (1996), Bender, et al. (2000).

12 It is important to note that women are underrepresented in the German workforce. For example, in 1991 41% of the Germany labour force were female. The fraction is lower when considering only full-time work. In addition, women's representation in the sectors that are excluded from the data varies. While approximately 10–15% of all workers are employed in the public sector, only a third of these are civil servants who are excluded from our data. The fraction of women in the public sector is 50%. Only 11% of self-employed workers are female.

13 Since less than 4% of men's wages and 2% of women's wages are top coded in our sample, this should not have any significant effect on our analysis. In order to test the sensitivity of our results to top coding we estimated both mean and median wage level and growth regressions and the two models produce similar estimates.

14 Since for German men part-time work is of negligible importance this restriction primarily affects women.

15 A large number of job to job changes among displaced workers is typical in European data.

16 One obvious question is whether men and women differ in the probability of being displaced. While it is possible to answer this question using the complete IABS data, these data are restricted and we were only given access to data on workers who experience some transition, from which we identified workers who were displaced as a result of a plant closing. Therefore, we cannot estimate the overall probability of being displaced from our data. However, the fact that men experience slightly more multiple spells of displacement is indirect evidence that men are slightly more likely to be displaced than women.

17 The main difference between the two programs is that unemployment assistance has a lower replacement rate than unemployment insurance. Additionally, unemployment assistance is means tested.

18 The replacement rate for unemployment insurance depends on family status and varies between 60 and 67% of the previous earnings.

19 Any time a worker does not receive unemployment compensation, and is therefore not considered unemployed, is not counted in our measure of the length of displacement

Table 1
Summary statistics — sample of displaced workers age 20–60 from IABS 1975–2001.

	All		Men		Women	
	Mean	Standard deviation	Mean	Standard deviation	Mean	Standard deviation
Proportion female (%)	37	–	0	–	100	–
Age prior to displacement	37.26	11.30	37.45	11.06	36.95	11.69
Weeks of experience prior to displacement	900.79	663.70	926.64	665.00	856.32	659.17
Weeks of tenure in displacement job	109.70	150.43	103.10	147.86	121.05	154.10
Proportion unskilled	0.37	0.48	0.34	0.47	0.42	0.49
Proportion skilled (vocational training)	0.61	0.49	0.63	0.48	0.56	0.50
Proportion graduate	0.02	0.15	0.03	0.16	0.01	0.12
Proportion post displacement job observed	0.66	0.47	0.69	0.46	0.63	0.48
Length of displacement/unemployment (in weeks)	230.93	342.68	215.26	332.44	257.88	358.08
Proportion full time in displacement job	0.91	0.28	0.99	0.08	0.78	0.42
Proportion full time job in first post displacement job[a]	0.86	0.34	0.97	0.16	0.66	0.47
Log real daily wage displacement job (only full-time workers)	4.73	0.46	4.83	0.41	4.49	0.47
Log real daily wage post displacement job (only full-time workers)[a]	4.68	0.40	4.76	0.35	4.46	0.42
Number of individuals	7212		4437		2775	
Number of displacement observations	8820		5578		3242	

[a] These means only include non-censored observations.

The rules governing unemployment insurance and unemployment assistance are gender neutral. The fact that unemployment assistance is means tested, and women are more likely than men to be married to a high-wage employed spouse, however, implies that women at the mean qualify for less unemployment assistance than men. This may lead to higher exit rates from unemployment for women at the expiration of unemployment insurance.

We identify a worker's post-displacement job as the first job we see where the worker works more than 17.5 h a week. We only keep workers for whom we have at least two years of data following the displacement event. Hence, in our data, 1999 is the last year a worker could be displaced.

We distinguish between three education groups: unskilled (10 or fewer years of compulsory schooling and less than 1.5 years of vocational training or college), skilled (10 years of schooling and an apprenticeship) and graduates (12 or 13 years of schooling and who have achieved a technical college degree or a university degree). Graduates are underrepresented in our sample, primarily because the IABS does not contain civil servants and self-employed.

Actual experience is calculated for every individual throughout the period 1975–2001. Around 50% of workers have entered the labor market before 1975 and for these workers accumulated labor market experience in 1975 is adjusted by potential experience. We assume that graduates are not older than 23 in 1975, and everybody else is not older than 16 in 1975.[20] Wages that are used are daily wages and are adjusted for inflation using the CPI for West Germany. The base year is 1995.[21]

3.2. The sample retained for analysis and summary statistics

Our final sample of displaced workers from the IABS data contains 8820 displacement events; 5578 events for men and 3242 for women.

In Table 1 summary statistics for our sample of displaced workers are shown. Displaced workers are on average approximately 37 years old with displaced women being 0.5 years younger than men. Once the gender difference in experience is adjusted for age, women have 55 fewer weeks of experience than men of the same age; this difference appears smaller than we would expect from population data. Recall we focus on workers who are highly attached to the labor market which accounts for why this difference is smaller than what would be found in data for the entire population. A large fraction of workers in these data find a job after displacement. Around 69% of men and 63% of women are observed in a post displacement job.

Women are less educated than men in our sample with relatively more women classified as unskilled. Displaced workers are primarily skilled, which is not surprising since skilled workers are the largest group in the German labor market. Immediately prior to displacement, the unadjusted difference in daily wage between men and women working full-time is 34%.

Including censored spells we see that on average women's displacement duration is 42 weeks longer than men's.[22] Comparing wages in the displacement and post-displacement jobs we see that men's wages fall by 7% while women's wages decline by 3%.[23] While this fall in wages is significant, it is smaller than the decline found by previous studies conducted using U.S. data (see Jacobson et al., 1993).[24] The data also capture typical differences between men and women in the probability of working full or part-time. Virtually all men work full- time, while 78% of women do so before displacement. A substantial fraction of women changes to part-time work after displacement.

Table 2 shows the distribution of displacement events by age and sex. Note the relatively high share of young displaced workers in our sample and that the probability of displacement falls with age.[25] Table 3 displays the length of displacement, in weeks, separately for men and women. Women are less likely than men to

[20] Potential experience is calculated as worker's age in the first spell observed, minus six minus years of education. We assume 9 years of schooling for the unskilled/low skilled workers, 11 years for skilled workers and 16 years for graduate workers.

[21] A complete list of selection rules and their effect on the size of the sample are available from the authors.

[22] One possibility is that more women are taking jobs where they are working less than 17.5 h a week. Given our focus on workers with a strong attachment to the labor market we treat these workers as still being displaced. In addition, while it is the case that women are more likely to work part time than men, very few workers in Germany work fewer than 17.5 h a week, so excluding these workers is unlikely to have any significant impact on our results.

[23] Conditional on re-employment in a full-time job women's wages drop by 2.7 percentage points more than men's wages.

[24] We acknowledge that these results are not directly comparable since Jacobson, et al. (1993) estimate the wage loss after displacement by comparing the actual wage to the expected wage. Expected wages are estimated from a control group (non-displaced workers).

[25] While we cannot test given the IABS data whether men and women in our sample are equally likely displaced, it appears that the distribution across age is very similar among displaced men and women.

Table 2
Distribution of age (%).

Age	All	Men	Women
20–25	18.28	16.37	21.56
26–30	16.89	17.00	16.72
31–35	14.57	15.72	12.58
36–40	12.09	13.28	10.02
41–45	10.91	10.94	10.86
46–50	9.91	9.65	10.36
51–55	9.74	9.27	10.55
56–60	7.62	7.78	7.34

Sample of displaced workers age 20–60 from IABS 1975–2001.
N = 5578 for men and N = 3242 for women.

Table 3
Distribution of individual durations of displacement events (%).

Weeks	All	Men	Women
Less than 14 weeks	31.0	34.0	25.6
15 to 20	7.0	7.3	6.3
21 to 32	8.4	8.6	7.8
33 to 52	8.7	7.6	10.6
53 and more	44.8	42.0	49.4

Sample of displaced workers age 20–60 from IABS 1975–2001.

find a job within 14 weeks of being displaced and more likely to wait over a year before finding a post-displacement job.

In order to describe male/female difference in durations of job search in more detail in Fig. 1 we plot Kaplan–Meier survival functions separately for men and women.[26] The survival function for women lies everywhere above the function for men. Hence, these unconditional estimates show that women tend to experience longer spells of displacement than men.[27] Somewhat puzzling is the kink at around 52 weeks, which is more pronounced for women than for men indicating that women experience a relatively larger jump in the probability of finding a job after one year. [28]

To investigate how gender differences in search duration vary by age in Fig. 2a and b we plot survival functions by gender and for eight different age groups: 20 to 25, 26 to 30, 31 to 35, 36 to 40, 41 to 45, 46 to 50, 51 to 55 and 56 to 60. The figures demonstrate that the gender differences in the length of displacement are most pronounced among those workers under age 35, but tend to shrink among older workers. This is supported by statistical tests for the difference in the functions which show that the functions are significantly different among the younger workers, but are not significantly different among the older workers. In addition, we see that the kink in the survival functions disappears for workers older than 45. Finally, looking at the functions for younger workers, we can see that much of the difference in search behavior is driven by the fact that men are more likely to find a job

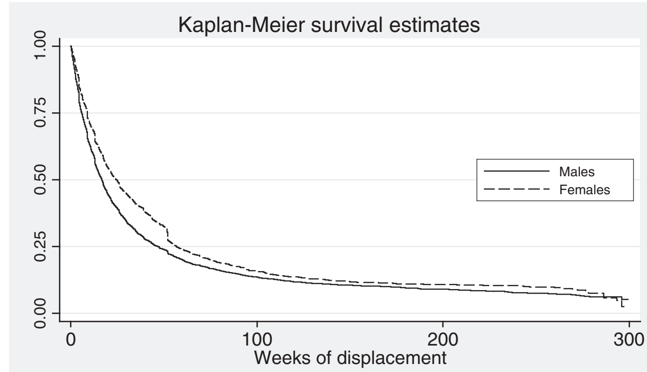


Fig. 1. Durations of unemployment.

relatively quickly. After two years the probability of finding a job appears to be quite similar for men and women in the younger age categories.

4. Analysis of the length of displacement

To investigate how worker characteristics affect the length of displacement, and to test whether gender differences in job search durations are varying by age we estimate a proportional hazard model which has the advantage of imposing few parametric assumptions. The main assumption in the model is that the baseline hazard is proportional and common to all groups (see Fig. 2a and b).

Following the proportional hazard model we specify the hazard rate as:

$$\lambda_u(t|x,\beta_u) = \lambda_0(t)^* \exp\Big(x_{pre}\beta_u + \delta(\alpha)^* female + \delta(\alpha)\Big) \quad (1)$$

where λ_u (t|.) is the transition rate from unemployment into employment at elapsed job search duration t conditional on control variables, x_{pre}, measured in the last spell before displacement and an indicator function in age that is interacted with a dummy variable for being female. The indicator function is defined as $\delta(\alpha) = \sum_k \delta_k I_k(\alpha)$ where the δ_ks are the key parameters and $I(\alpha)$ is a series of interval dummy variables with the intervals being: 20 to 25, 26 to 30, 31 to 35, 36 to 40, 41 to 45, 46 to 50, 51 to 55, and 56 to 60 years old.[29] The model is partially non-parametric since the baseline hazard function, $\lambda_0(t)$, is unspecified. The influence of the covariates is modeled parametrically as a multiplicative effect on the hazard. In this model estimation of the coefficients of the covariates does not depend on the specification of the baseline.

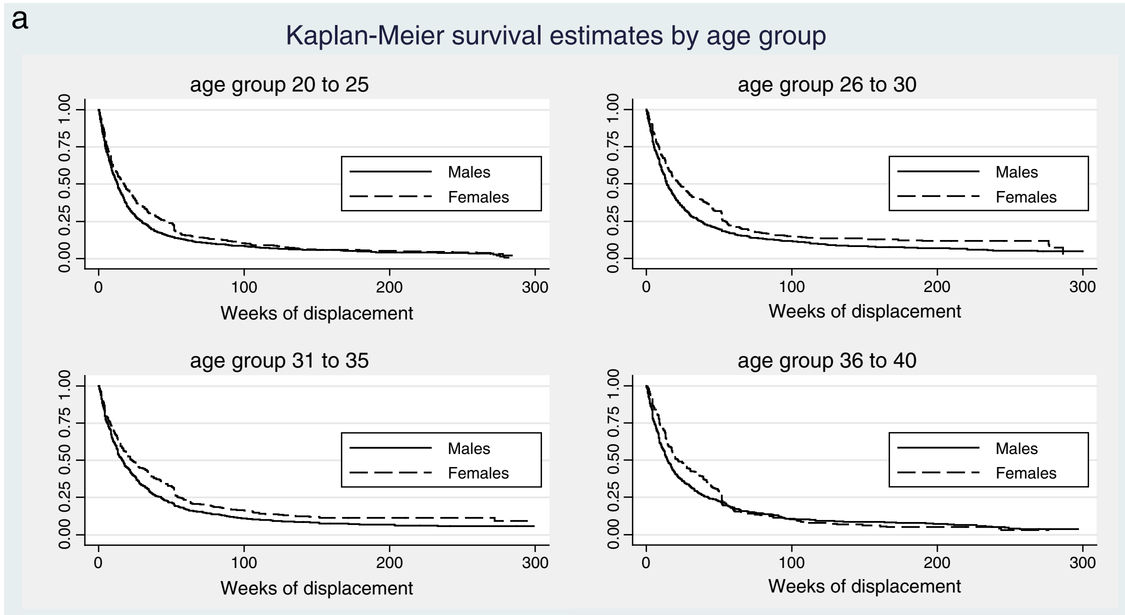
The vector of controls measured in the last spell before displacement, x_{pre}, in the basic specification includes experience, experience squared, tenure, tenure squared, as well as dummy variables for education group (three groups), industry (fifteen groups), full time/part time, and calendar year. In extensions we include controls for occupation (eight groups). In the estimation all displacement events are pooled. Hence, we assume that these are random independent draws. We adjust standard errors for clustering on individuals; that is the fact that a single individual can have multiple displacement events. In the tables we report estimates of the hazard ratios 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.

[26] We only plot spells that are less than 300 weeks since there are very few spells lasting more than 300 weeks.

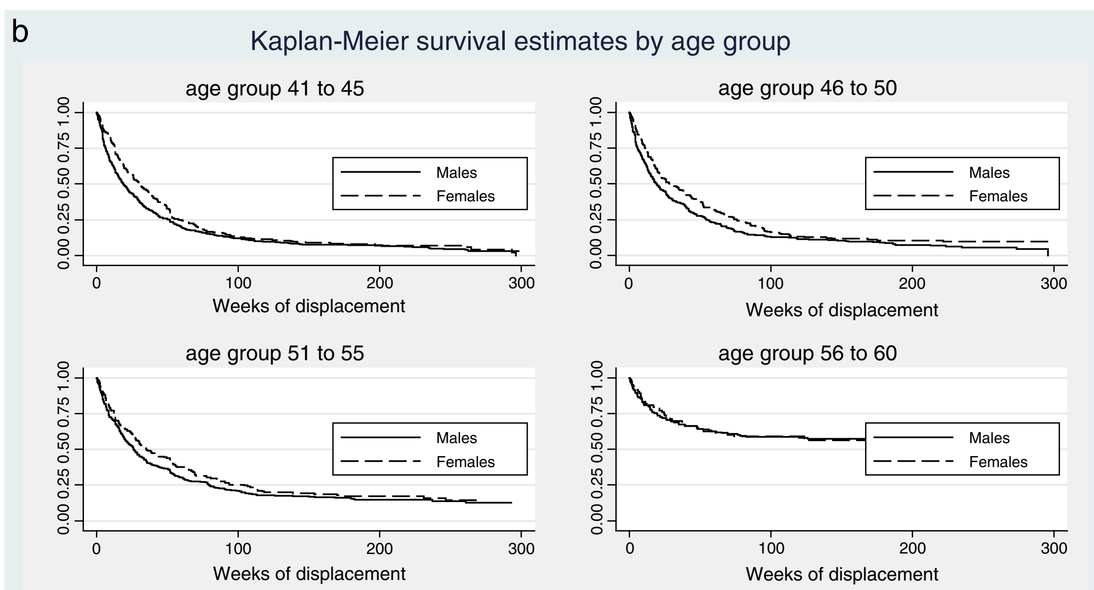
[27] The p-value of the rank test shows that the two functions are significantly different at standard levels of significance.

[28] Note that our measure of unemployment duration includes periods of receipt of unemployment insurance as well as unemployment assistance. Eligibility rules for the receipt as well as the duration of pay of unemployment insurance (Arbeitslosengeld) and unemployment assistance (Arbeitslosenhilfe) have been modified several times during our observation period. Changes are stated in the job employment act 1969 to 1997 and the social code (SGBIII) since 1997. To analyze how durations are affected by these changes goes beyond the scope of this paper. See e.g., Hunt (1995) for an analyses of changes on men's outcomes during the late 1980s and 1990s. Currently workers under the age of 42 can only receive unemployment insurance for one year and, because women are often not the primary wage earner in the household, they are likely to receive smaller unemployment assistant payments than men. These two facts may account for the larger jump in the probability of leaving unemployment for women around 52 weeks.

[29] We also estimate specifications using an indicator function with two year age brackets.



p-values for rank test (H0=no diff): age group 20-25: p=0, 26-30: p=0, 31-35: p=0, 36-40: p=0.2.



p-values for rank test (H0=no diff): age group 41-45: p=0.05, 46-50: p=0.04, 51-55: p=0, 56-60: p=.76.

Fig. 2. Durations of unemployment of displaced workers by Gender and Age Group. a: 20 to 40 years old by gender b: 41 to 60 years old by gender.

In Table 4 we report estimation results from the Cox proportional hazard model specified in Eq. (1). We see from the results in column 1 that the length of worker displacement rises fairly quickly with age. In addition, we see that more experienced workers tend to have shorter spells of displacement, workers who have longer tenure at the firm have longer spells of displacements, while more educated workers have shorter spells of displacement. All of these results are similar to results found by previous studies of displaced workers.

Focusing on the hazard ratios on the interaction between the age categories and the female dummy variable we see that there is a distinct life-cycle pattern in the differences between men and women in the length of displacement. The coefficients on the female/age interactions show that the increased length of displacement among women occurs exclusively among women who are 35 years old or younger. For women 36–55 there is no significant difference in the length of displacement, and women 56–60 experience shorter lengths of displacement than men of similar age. [30]

In column 2 we present results from a model where we include a worker's occupation in the displacement job. Since there is some dispute regarding whether occupation controls should be included in analysis such as we are conducting, we decided to estimate models both excluding and including occupational controls. Comparing the results in columns 1 and 2 shows that including occupational controls has very little effect on the results. The only important differences are that, in column 2 it appears that women age 51–55 experience significantly longer spells of displacement than similarly aged men, while

[30] This could reflect positive selection into work. Those with potentially long spells transition into early retirement.

Table 4
Proportional hazard estimation of the length of displacement.

	Entire sample (1)	Entire sample (2)	First displacement sample (3)
Age 26–30	.789 (.046)***	.788 (.046)***	.765 (.048)***
Age 31–35	.673 (.047)***	.674 (.047)***	.667 (.050)***
Age 36–40	.632 (.055)***	.632 (.055)***	.612 (.058)***
Age 41–45	.559 (.057)***	.553 (.057)***	.562 (.065)***
Age 46–50	.504 (.059)***	.502 (.059)***	.512 (.069)***
Age 51–55	.354 (.050)***	.350 (.049)***	.382 (.063)***
Age 56–60	.093 (.019)***	.092 (.019)***	.093 (.022)***
Female*age20–25	.887 (.051)**	.902 (.053)**	.922 (.057)
Female*age 26–30	.705 (.047)***	.720 (.049)***	.736 (.055)***
Female*age 31–35	.779 (.058)***	.781 (.058)***	.803 (.067)***
Female*age 36–40	.987 (.075)	.993 (.076)	1.001 (.091)
Female*age 41–45	.939 (.072)	.953 (.073)	.964 (.089)
Female*age 46–50	.928 (.079)	.939 (.081)	.908 (.091)
Female*age 51–55	.794 (.082)	.805 (.084)**	.793 (.092)**
Female*age 56–60	1.015 (.218)***	1.018 (.219)	.922 (.223)
Weeks of experience	1.000 (.0001)***	1.00 (.0001)***	1.000 (.0001)***
Weeks of experience squared	.999 (6.50e−08)***	.999 (6.51e−08)***	.999 (7.44 e−08)***
Weeks of tenure	.999 (.0002)***	.999 (.000)***	.999 (.0003)***
Weeks of tenure squared	1.00 (3.76e−07)	1.00 (3.77e−07)	1.00 (3.96e−07)
Unskilled worker	.970 (.027)	.973 (.027)	.980 (.032)
Skilled worker	–	–	–
Graduate worker	.679 (.074)***	.693 (.075)***	.684 (.079)***
Year dummies	Yes	Yes	Yes
Industry dummies	Yes	Yes	Yes
Occupation dummies	No	Yes	No
Log-likelihood	−49744.22	−49685.351	−39262.785
Number of individuals	8655	8648	7185
Number of observations	8797	8790	7185

Sample of displaced workers age 20 to 60 from IABS 1975–2001. Hazard ratios are reported together with robust standard errors in parenthesis.
*Significance at 10% level.
**Significance at 5% level.
***Significance at 1% level.

the coefficient on the female age 56–60 interaction becomes insignificant.

Our results may be sensitive to the inclusion of people who are displaced more than once if unemployment is persistent over time because of unobserved heterogeneity, or time dependence (see Van den Berg (2001) for an overview). In Table 4 column 3 results are presented from the model estimated on a sample that only includes a person's first displacement spell. Comparing the results in column 3 with those in column 1 shows that the results from the two samples are quite similar.[31]

Another possibility is that cohort effects may be producing the observed pattern. That is, it may be that there are smaller differences in the length of displacement among the older cohort of workers who will tend to dominate the older age categories. To examine this possibility in Table 5 we present the results from our basic model estimated separately on four different cohorts of workers: workers born between 1951 and 1955, workers born between 1956 and 1960, workers born between 1961 and 1965 and workers born between 1966 and 1970. We focus on these cohorts because other cohorts have too few observations. While less precisely estimated than our estimates based on the full sample, the coefficients on the age-female interactions show that same basic pattern as before. The difference in length of displacement primarily occurs among the younger workers with prime age men and women experiencing similar lengths of displacement.[32]

In summary, these results show that gender differences in job search duration are not constant but varying in age. They do suggest that gender differences in job search and mobility are related to fertility decision. More particularly, a plausible interpretation is that the presence of young children reduces displaced women's job search intensity, though our data do not allow us to measure either search intensity or the presence of children directly. Alternatively, it might be that the presence of young children causes women to increase their reservation wages.

5. Wages after displacement

Equilibrium search models with discrimination (e.g. Bowlus and Eckstein, 2002) can generate the outcome that on average women search longer for a new job, because of the presence of prejudiced firms in the market, and have lower wages than men, because all firms can exert monopsonic power. In this section we examine the gender wage gap in post-displacement wages. We follow the same strategy that we followed in the previous section, by examining how wage levels differ between men and women over the life-cycle. Note that we condition on re-employment which may introduce selection bias since workers who become re-employed may not be a random sample of all workers who lose their job. We will discuss implications of this possible nonrandom selection at the end of the section. Note that all regressions include dummy variables for year which capture aggregate time varying shocks.

The first two columns in Table 6 show the results from a regression where the log of daily wages in the post-displacement job is the dependent variable. Because we only have data on daily wages, and do not have detailed information on hours worked, we only include workers with full-time jobs in our analysis. All of our control variables are measured at the last spell before displacement. We include the same set of controls that we did in our hazard models, with the exceptions that we have dropped potentially endogenous variables, such as the tenure variables and the controls for industry.

The results in column 1 show that, when we include our basic set of controls, women's wages are approximately 27% less than men's wages. The results in column 2, where we have included an interaction between the female dummy and the age variables, show no distinct age pattern in the male/female wage gap. Of course, showing that women earn lower wages than men in their post-displacement job does not tell us much about the direct gender difference in wages following displacement, since women likely earned less than men in their pre-displacement job. In order to examine the differential impact of displacement on wages, in columns 3 and 4 we present regressions where the dependent variable is the change in log wages between the pre- and post-displacement job.[33] In column 3 we see that, conditional on education and experience, women experience approximately a 3% larger drop in wages after displacement than men.[34] The results in column 4 show that women age 20 to 25 and 46 to 50 do seem to experience larger wage losses after displacement than men, for the rest of the age groups there is no significant gender

[31] In unreported results we examined whether the observed pattern is being produced by the use of five year age categories. Corresponding regression results using two-year age categories show the same life-cycle patterns seen in Table 4. Results available upon request. We also estimated the exponential model allowing for individual heterogeneity where the unobserved heterogeneity component is uncorrelated with the explanatory variables. The estimates are robust to how we treat this type of heterogeneity once we allow for a flexible structure of the hazard. Since we observe too few individuals with multiple displacement spells we cannot exploit those to estimate the distribution of unobserved heterogeneity so we are forced to rely on distributional assumptions. Results are reported in the Appendix Table A.1.

[32] Note that effects for the age groups older than 51 years cannot be estimated since our panel data end in 2001.

[33] Again we focus on full-time workers and those re-employed.

[34] This may underestimate the population mean effect if those with relatively high pre-displacement wages return.

Table 5
Proportional hazard estimation of the length of displacement for selected birth cohorts.

	Cohort 1951–1955 (1)	Cohort 1956–1960 (2)	Cohort 1961–1965 (3)	Cohort 1966–1970 (4)
Age 26–30	1.03 (.303)	.900 (.141)	.952 (.181)	.834 (.107)
Age 31–35	1.04 (.351)	.900 (.254)	.999 (.227)	.727 (.206)
Age 36–40	.950 (.392)	1.05 (.348)	.874 (.279)	
Age 41–45	.800 (.366)	.770 (.345)		
Age 46–50	.510 (.256)			
Female*age20–25	.926 (.216)	.747 (.075)**	.989 (.118)*	.757 (.101))**
Female*age 26–30	.793 (.123)	.797 (.117)	.658 (.089)**	.673 (.096)**
Female*age 31–35	.860 (.140)	.774 (.126)	.769 (.118)*	.631 (.221)
Female*age 36–40	.904 (.156)	.948 (.145)	.854 (.260)	
Female*age 41–45	.855 (.135)	.944 (.344)		
Female*age 46–50	1.748 (.634)	.999 (.000)		
Weeks of experience	.999 (6.58e−04)	.999 (4.13e.04)	1.00 (6.05e.04)*	1.01 (.001)***
Weeks of experience squared	1.00 (3.50e−07)	1.00 (3.45e−07)**	.999 (7.53e−07)	.999 (1.89e−06)***
Weeks of tenure	.998 (6.03e.04)***	.999 (7.32e−04)	.997 (7.93e−04)***	.995 (.001)***
Weeks of tenure squared	1.00 (8.51e−07)**	.999 (1.31e−06)*	1.00 (1.49e−06)***	1.00 (21.04e−06)***
Unskilled worker	.927 (.069)	1.08 (.068)	.998 (.072)	.987 (.088)
Skilled worker	–	–	–	–
Graduate worker	.531 (.143)**	.646 (.159)*	1.23 (.299)	.644 (.299)
Dummy for part time	.882 (.149)	1.14 (.169)	1.08 (.185)	1.05 (.245)
Log-likelihood	−5893.84	−7102.24	−5838.79	−3784.53
No of individuals	1173	1368	1224	872
No of observations	1205	1393	1235	876

Sample of displaced workers age 20 to 60 from IABS 1975–2001. Hazard ratios are reported together with robust standard errors in parenthesis.
* Significance at 10% level.
** Significance at 5% level.
*** Significance at 1% level.

Table 6
Regression results of the log real daily wage in levels and first differences.

Variable	Log wage in post-displacement job		Difference in log wages pre and post displacement job	
	(1)	(2)	(3)	(4)
Female	−.277 (.012)***		−.027 (.013)**	
Age 26–30		.042 (.023)*		.007 (.024)
Age 31–34		.033 (.029)		.021 (.030)
Age 35–40		.019 (.037)		.064 (.039)***
Age 41–45		−.014 (.046)		.121 (.048)***
Age 46–50		−.016 (.058)		.190 (.061)***
Age 51–55		−.035 (.072)		.237 (.076)***
Age 56–60		−.062 (.099)		.224 (.104)**
Female*age 20–25		−.301 (.022)***		−.049 (.023)**
Female*age 26–30		−.182 (.029)***		−.032 (.030)
Female*age 31–35		−.275 (.033)***		−.021 (.035)
Female*age 36–40		−.206 (.037)***		.021 (.039)
Female*age 41–45		−.248 (.040)***		.034 (.042)
Female*age 46–50		−.426 (.043)***		-.079 (.045)*
Female*age 51–55		−.374 (.052)***		-.004 (.055)
Female*age 56–60		−.321 (.094)***		.044 (.099)
Adjusted R-squared	0.1945	0.2022	0.0184	0.0198
No. observations	4075	4075	4075	4075

Sample of displaced workers age 20 to 60 from IABS 1975–2001 who obtain a full time job after displacement. All regressions include control variables for experience (squared), education and year. Coefficients and robust standard errors in parenthesis are reported.
* Significance at 10% level.
** Significance at 5% level.
*** Significance at 1% level.

difference.[35] While it seems that the youngest women in our sample fall slightly behind comparable men in terms of wages, there is no significant difference between the age groups. We do not see the same life-cycle pattern in wages that we saw in job displacement durations. A plausible interpretation of this finding is that women of prime child-bearing age raise reservation wages relative to other women — meaning that the women who do find employment after being displaced receive relatively higher wages. One should keep in mind that these wage regressions are for full-time workers and are conditional on re-employment and do not account for the possible non-random selection into re-employment. Hence, if after displacement only the most able women search for a new job while a random sample of men search then these estimates would understate the effect of displacement on the gender wage gap. Conversely, if the opposite is true – only the most able men search for a new job after

[35] While we feel that the specification in Table 6 is the appropriate one for the question we are addressing, it does impose a specific relationship between the current wage and the wage prior to displacement. In order to relax this assumption we have estimated a model where the log daily wage at the current job is the dependent variable and the log daily wage at the pre-displacement job is the independent variable. When we estimate this model we see the same basic pattern in the coefficients on the age-female interaction terms that we see in Table 6. However, none of these coefficients is significant in this alternative model.

displacement while a random sample of women search – then these estimates overstate the impact of displacement on the gender wage gap.

6. Conclusions

In this study we use longitudinal register data for workers in West-Germany along with duration analysis to analyze gender differences in job search behavior. In order to examine differences in job search that results from potentially exogenous factors we focused on highly attached workers who were displaced through a plant closure. A particular strength of our data is that they cover workers of all ages so we can examine whether unemployment durations following displacement differ significantly across male and female workers by age groups. Our main empirical result shows that the gender differences in displacement durations vary across the life cycle with the largest differences occurring among workers age 26–35 — there is no significant cross-gender difference in the length of displacement among younger or older workers. This finding remains valid even after we control for a rich set of characteristics, as well as when we control for potential unobserved heterogeneity and duration dependence, so it does not appear that changes in the composition of displaced workers is driving our results.

This result seems at odds with the implications of existing equilibrium search models that incorporate taste based employer discrimination (e.g. Bowlus and Eckstein, 2002) which predict that all women should experience longer spells of displacement. This observed life-cycle pattern in gender differences in displacement durations is new evidence for a European country and it would be interesting to see whether these patterns are also observed in other countries. Interestingly, our finding for West-Germany is quite similar to the life-cycle pattern we find when we conducted a similar analysis using data on displaced workers in the U.S. (Kunze and Troske, 2009). This suggests that the pattern is not unique to Germany and may be driven by some general behavioral mechanisms. It also suggests that our finding is not the result of the relatively generous unemployment benefit system in Germany.

The obvious question is then, what mechanism could generate the observed life-cycle pattern in job search durations? In Kunze and Troske (2009) we present evidence supporting the hypothesis that the life cycle pattern is a function of fertility related opportunity costs. That is women in the prime childbearing and rearing years have relatively high opportunity cost of working compared to similarly aged men. In our companion paper we show that, once we focus on women who do not have a child after being displaced, men and women have similar lengths of displacement across all age groups. Unfortunately, we cannot perform a similar analysis using the German data. However, the results in both papers suggest that the fertility decisions of women have a significant impact on women's labor market mobility.

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Appendix A

Table A.1
Estimation of the length of displacement controlling for unobserved heterogeneity(*).

	Without duration dependence	With duration dependence
Age 26–130	.530 (.079)***	.728 (.057)***
Age 31–35	.335 (.058)***	.584 (.054)***
Age 36–40	.250 (.053)***	.551 (.064)***
Age 41–45	.176 (.045)***	.459 (.063)***
Age 46–50	.124 (.036)***	.406 (.064)***
Age 51–55	.051 (.017)***	.255 (.047)***
Age 56–60	.004 (.001)***	.056 (.014)***
Female*age 20–25	.858 (.130)	.848 (.066)**
Female*age 26–30	.482 (.078)***	.628 (.055)***
Female*age 31–35	.595 (.108)***	.725 (.071)***
Female*age 36–40	1.077 (.213)	.920 (.092)
Female*age 41–45	.953 (.190)	.888 (.089)
Female*age 46–50	.898 (.190)	.864 (.096)
Female*age 51–55	.664 (.150)*	.749 (.096)**
Female*age 56–60	1.037 (.378)	.994 (.229)
Weeks of experience	1.001 (.000)***	1.000 (.000)***
Weeks of experience squared	.999 (1.48e-07)***	.999 (8.28e-08)
Weeks of tenure	.997 (.000)	.998 (.000)***
Weeks of tenure squared	1.000 (8.01e-07)**	1.000 (4.65e-07)***
Education (omitted group:skilled)		
Unskilled	.936 (.065)	.959 (.035)
Graduate	.422 (.100)***	.606 (.082)***
Full-time job	1.1675 (.165)	1.103 (.085)
Duration		
27–52 weeks		.558 (.030)***
53–78 weeks		.302 (.022)***
79–104 weeks		.139 (.013)***
105–130 weeks		.080 (.009)***
131–156 weeks		.036 (.005)***
157–182 weeks		.019 (.003)***
183–208 weeks		.014 (.003)***
209–234 weeks		.010 (.002)***
235–260 weeks		.011 (.003)***
260 and more weeks		.001 (.000)***
Log-Likelihood	−17205.044	−14660.522
Number of individuals	8654	8654

Sample of displaced workers age 20 to 60 from IABS 1975–2001. All regression contain dummy variable for year and industry. Hazard ratios are reported together with robust standard errors in parenthesis. (*) We estimate a model with a constant hazard and inverse gaussian distribution for the unobserved heterogeneity component.

* Significance at 10% level.
** Significance at 5% level.
*** Significance at 1% level.

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